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Preface

Demographic changes are now at the heart of the development of the economy. Within the area of Population Economics, the interest in household and family issues has been steadily rising in the last few years, which is reflected in the number of papers published by scientific journals, and in the programs of economics conferences. While this interest was fostered by rapid and ongoing changes in family structures over the last decades, only the growing availability of high-quality micro datasets and the required computer capabilities made it possible to deal more intensively with this topic. The increasing empirical evidence also produced further challenges to develop the theoretical framework.

In the editorial process of the Journal of Population Economics many papers have passed our desks and only the best of them have been published after a refereeing process. Given the increasing interest in the matter and the high quality of the research published in our journal, we saw a substantial value-added to publish a selection of recent contributions to make the work more accessible to the scientific community. Hence, we are very glad to present the book herewith. It is divided in three parts, "Time use and non-market work", "Household and family development" and "Transition to work and younger employees". Of course, there are close interactions between these topics. The following editorial shall give a short overview.

We would like to take the opportunity to thank Springer-Verlag for the fruitful cooperation over the past years.

Klaus F. Zimmermann
Michael Vogler

Editorial

by *Klaus F. Zimmermann* and *Michael Vogler*

During the last decades, the appearance of a family has changed substantially. Not long ago a typical family consisted of a husband who left home in the morning to go to work, while his wife was tended to the housework during the day. They normally lived together for their whole life times, having one or more children, which primarily were raised by the wife. Although this model for most people might still be the optimal way of living together, in practice differing living models became much more common than before. Statistics provide clear observations of the trends: The number of single-person households increased. Age of marriage, as well as women's age at first motherhood became remarkably higher. Fertility has fallen rapidly. The number of divorces has steadily increased.

The ancestral role allocation became less relevant in practice, because today neither economic nor social constraints are as import for the decision to start a family as they were in the past. From an economic point of view, there are several reasons. Among them: In the course of increasing equality today the average woman is by far better educated than before. In some industrial countries she even has a school education superior to that of the average man. This led to more financial autonomy and higher professional ambitions. Furthermore, the expansion of social security systems decreased the need of family assistance in distress and old age. In the face of decreasing fertility in industrialized countries, it is often argued that legal and fiscal provisions even discriminate families.

The perception and valuation of housework and child care respectively, has changed. On the one hand this should to some degree be a result of changing family structures, while on the other hand it probably even contributed to the latter. With a growing number of single-person households and two-earner couples, men have taken over more responsibilities at home although it is still common that women stay at home after their children are born. Yet we are still witnessing a lack of knowledge in economics about the mechanisms of household time allocation. Little attention has been given to this topic, because housework normally is regarded as non-market work and time not spent on paid work simply has been defined as leisure time. Only in the last few years, there has been a rising interest in the 'time use' of people.

The first chapter ("Time use and non-market work") starts with a study of *Daniel S. Hamermesh,* who points out that the dissociation from the standard view on labor supply provides useful new insights in the topic of time use. *Sebastien Lecocq,* as well as *Steinar Vagstad,* and *Thomas Aronsson, Sven-Olov Daunfeldt* and *Magnus Wikström* use the 'household production model' as a starting point to empirically and theoretically analyze various aspects of the intra-household time allocation. *Alfonso Sousa-Poza, Hans Schmid* and *Rolf Widmer, as well as Michael Lundholm* and *Henry Ohlsson* take a closer look especially at the time parents dedicate to child care.

The ongoing changes in family structures generate people, whose notion of a family is completely different from that of their grandparents, not only regarding family size. Undoubtedly, it is social progress that childlessness, divorce, or single motherhood are no longer stigmatized by society. However, because less children experience the traditional family model, its assumed merits could get more and more forgotten over time. There is increasing evidence that disturbed family structures lead to unfavorable development of children. Divorces are often not only accompanied by psychological strains but also by financial problems. Apart from aspects of social behavior, from an economic point of view it is argued that affected children have lower achievement potential.

The second chapter ("Household and family development") provides special attention to the important influence of the household background, mainly in terms of income and household formation, on the development of the family and children respectively. The study by *Stephen Jenkins* deals with income dynamics, which not only are influenced by economic factors but also by changes in the family composition, *Alessandra Guariglia* takes a closer look at the relationship between income uncertainty and saving behavior of households. *John F. Ermisch* and *Marco Francesconi, Andrew McCulloch* and *Heather E. Joshi*, and *Martha S. Hill, Wei-Jun J. Jeung* and *Greg Duncan* analyze the effects of family structure and wealth on children's cognitive and educational development. *Maite Martinez-Granado* and *Javier Ruiz-Castillo*, as well as *Stephen Garasky, R. Jean Haurin* and *Donald R. Haurin* examine the decisions of adolescents whether to stay with or leave their parental home and their favorite living models.

Nearly all statistics show that young adults face an above-average risk of being unemployed. School-to-work transition and early years of labor market participation are subject to mechanisms, which are different than those for experienced workers. Entering the labor market, all adolescents inevitably belong to the unemployed first and have to search for a job. Being 'outsiders' they are especially affected by structural and cyclical labor market problems. Furthermore, depending on age and personal development, young people often do not have a profound knowledge of their abilities and options. Not surprisingly, the exits from employment are substantially higher than those for more experienced workers. Because young women have to decide whether and when they want to bear children, their job decisions are influenced by additional aspects. To the contrary, fertility decisions are affected by the labor market situation, too.

In the last chapter ("Transition to work and younger employees") the first set of papers deals with the determinants of youth's labor market success. *Regina T. Riphahn* analyses the determinants of school-to-work transition in general, *Oivind Anti Nilsen, Alf Erling Risa* and *Alf*

Torstensen deal with the exits of youths from employment, while *Paul Fronstin, David H. Greenberg* and *Philip K. Robins* especially concentrate on the identifying effects of parental disruption on the labor market performance of children. The second focus of this chapter is on young females. *Siv Gustafsson* gives an overlook of the economic view on timing of fertility. Focusing on different countries, *Adriaan S. Kalwij* (for the Netherlands), *Namkee Ahn* and *Pedro Mira* (for Spain), and *Linda Adair, David Guilkey, Eilene Bisgrove* and *Soccoro Gultiano* (for the Philippines) analyze the relationship between childbearing and the situation in the local labor markets.

Time use and non-market work

Most studies on time allocation are based on the so-called 'household production model', which was introduced by Gary Becker and radically widened the economic view of non-market activities at home. The basic notion is that households combine time and market goods to produce commodities that enter their utility function. Household members specialize according to their comparative advantages and also allocate investments according to this point of view. Most attention was given to the modeling of the household utility function, abandoning the original assumption of households as being single utility maximizing units. A popular model is the so-called 'collective model' by Pierre-Andre Chiappori. However, there still are a lot of theoretical questions that are open, and empirical evidence in the past often suffered from the scarcity of usable data sets.

Hamermesh stresses the fact that household structure not only accounts for people's supply of paid work in terms of hours, but also determines people's preferences on when to work. Pleasant working times can be seen as a non-monetary benefit, an aspect which is especially important to two earner couples and families that prefer to spend as much common time at home as possible. Using data from the U.S. Current Population Survey (CPS), he shows that evening and night work decreased since the early 70s. Rising real earnings power obviously has been used to shift away from unpleasant work time. At the same time, not only earnings inequality increased but also the distribution of unpleasant working times, with low wage worker having to accept a larger fraction of evening and night work. An analysis of spouses' decisions revealed that common leisure time actually is a determinant of individual labor supply, and that income increases are partly converted to the realization of togetherness.

The "household production model" is a common starting point for the study of time allocation. For econometric reasons, the absence of concrete commodities led researchers to construct a function which is weakly separable in goods and time used for the production of commodities in the sense of the household production model. Using French data, *Lecocq* tests this so-called 'weak separation hypothesis'. His results for instance show that meal preparation actually is separable from restaurant expenditures, market goods inputs and household leisure time. However, opposed to the hypothesis' assumption, it is not separable from time inputs devoted to other household activities. In his theoretical study, *Vagstad* shows the consequences of non common preferences of family members, and thus abandons an usual assumption. Although the mechanisms of specialization keep valid as suggested by the household production model, neither investments in specialization remain efficient nor the time allocations to it. *Aronsson, Daunfeldt* and *Wikström* use an extended version of the mentioned 'collective model' to estimate the intra-family distribution of income, leisure and household production from Swedish household data. In contradiction to other studies, their results confirm the importance of the 'pooling hypothesis', which states that only the aggregated income, but not income distribution, determines the intra-household allocation of time. Education and the number of children are the most important factors for the allocation of housework and leisure.

Sousa-Poza, Schmid and *Widmer* take a closer look at the allocation of time to housework and child care. Using Swiss data they can confirm that the presence of children primarily influences women's behavior. The time men invest in housework does not rise when children are present, and only little time is dedicated to child care. Furthermore, the results show that men with higher education allocate more time to housework and child care. *Lundholm* and *Ohlsson* extend the 'quality-quantity model', which says that increased income could not only increase demand for children, but also could be used for investments in the quality of children. The study shows that, if parents face restrictions in terms of time and the possibility to purchase child care, income increases still are ambiguous regarding fertility outcomes.

Family structure and development

While the model of an optimal family should not have changed so much in the mind of most people, in practice there have been rapid changes in the realized living models. In research there is a growing awareness concerning the consequences of these changes, not only in sociology but also in (population) economics. These new family backgrounds, which often entail economic and psychological problems, might produce children who are left aggrieved with regard to different aspects, e.g. in their cognitive development and educational attainment. Revolving around this focus, the chapter presents new insights concerning household income formation and its consequences, family structure and development, and the decisions of youths regarding their living arrangements towards autonomy.

Although, of course, labor earnings of the household head are crucial for family care, looking at the poverty dynamics reveals that over time other factors play an important role as well. In his study for Britain *Jenkins* shows that, overall, demographic events (e.g. partnership dissolution) are more important for poverty spell beginnings than changes in the household head's labor earnings. In the case of spell endings, demographic events do not have this same relevance. However, it is shown that other, additional money income, such as the spouse's labor earnings and benefits, overall are of higher importance for leaving poverty than the head's earnings. In her study, which also uses the British Household Panel Survey, *Guariglia* analyzes the influence of income uncertainty on household saving behavior. Her results show that there actually is a general component of precautionary savings. Furthermore, in accordance with the life cycle model, expected financial deteriorations let people accumulate reserves.

The development of a family and children respectively, is primarily an outcome of the underlying 'in-house' background although external influences, of course, also might play an important role. Most attention is given to the relevance of household wealth and family structure. *Ermisch* and *Francesconi*, using data from the British Household Panel Survey, show that children from a single-parent family not only are aggrieved in terms of education and have a higher risk of inactivity, but also more often suffer from health problems. *Hill, Yeung* and *Duncan,* using U.S. panel data (PSID), find that parental marital change has stronger influences when the event occurs during late childhood. In their study family income appears to be the most important factor for better educational attainment and lowers daughter's risk of a nonmarital birth. Based on British data, *McCulloch* and *Joshi* conclude that family poverty is associated with poorer average cognitive development of children. However, material disadvantages obviously can be overcome by positive parental care, which is mostly depending on the mother's education.

Martinez-Granado and *Ruiz-Castillo* analyze three import decisions of adolescents towards their autonomy: Whether to study, to work and to leave the parental home, or not. Their study for Spain explicitly considers the interdependencies of these decisions. Among other interesting results, they find that education has a positive influence on the probability of males leaving their homes, while this is not the case for females. Housing prices clearly matter, but living in metropolitan areas by itself, leads to a higher propensity to leave. Using a national American sample of adolescents aged between 16 and 30, *Garasky, Haurin* and *Haurin* look at the factors, which influence adolescents' choices of destination when exiting the parents' home. They also realize that the home-leaving decision is arrived at differently by males and females. Furthermore, while economic variables are relevant for the leave, the decision to move into large or small groups is solely influenced by socio-demographic factors.

Transition to work and young employees

Labor markets in industrial countries suffer from high youth unemployment. Many governments started policy measures to fight the threat that the persistence problem creates a generation with a substantial fraction of hopeless people. This would not only mean a waste of human resources, but also would cause substantial long-term economic and societal problems. There are no doubts that in a globalized world education is the key for future development, and that the course can only be set accordingly during childhood if there are to be reasonable perspectives. Furthermore, there is a clear relationship between the situation in labor markets and fertility decisions of young women. While causality in principle works both ways, for industrialized countries the effects of the labor market situation on the fertility decision is of greater interest in practice.

While *Riphahn* investigates the determinants of school-to-work transitions in Germany, using data from the German Socio-Economic-Panel (GSOEP), *Nilsen, Risa* and *Torstensen* analyze the exits from employment based on a large representative sample of young Norwegian workers. Both studies confirm the importance of human capital acquisition, especially measured by school type, experience and age. In general good educational attainment secures a lower risk of unemployment, as predicted by the human capital theory. *Nilsen, Risa* and *Torstensen* and *Riphahn* also show that, given the personal characteristics, the conditions of local labor markets and in industrial sections have a strong influence on the employment of young workers. Considering the institutional framework and past legal measures both studies provide useful insights into youth labor market policy in particular. Using British data *Fronstin, Greenberg* and *Robins* show that the accumulation of human capital is highly dependent on parental disruptions during early childhood. Divorce or parental death lead to lower educational attainment and worse labor market outcomes. The scope of these effects depend on the age of children at the time the disruption occurs, and surprisingly is different for male and females. The importance of the family background, too, is relevant in the study of *Riphahn,* who shows that parents' educational attainment positively influences the labor market success of their children.

In her study on the optimal age of motherhood, *Gustafsson* provides an overlook of economic determinants of fertility timing. While unemployment and income can always be found in the spotlight of studies, she realizes that consumption smoothing and individual career planning also play a major role in practice. *Kalwij* shows that in the Netherlands, controlling for other characteristics, employed women schedule the first children later in life and overall bear fewer children during their life spans than unemployed women. Educational attainment seems to have no direct effect on motherhood but works via the employment status. Looking at the 'fertility crisis' in Spain *Ahn* and *Mira*, explicitly focus on the importance of the male employment status. Their results show that spells of male unemployment have a negative effect on the timing of marriage, and subsequently on the decision to have the first child. Focusing on the effects of childbearing on women's earning, in their study of the Philippines *Adair, Guilkey, Bisgrove* and *Gultiano* not only concentrate on the type of work but also on supplied hours. Allocating restricted time between child care and work it might be optimal to have lower paid but more flexible work. In fact the results clearly show that a higher number of children lead to lower earnings. However, this effect is only remarkable in presence of babies, which shows that women adjust work time in favor of child care.

Contents

Timing, togetherness and time windfalls

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Abstract. With appropriate data the analysis of time use, labor supply and leisure can move beyond the standard questions of wage and income elasticities of hours supplied. I present four examples: 1) American data from 1973 through 1997 show that the amount of evening and night work in the U.S. has decreased. 2) The same data demonstrate that workers whose relative earnings increase experience a relative diminution of the burden of work at unpleasant times. 3) U.S. data for the 1970s and 1990s demonstrate that spouses' work schedules are more synchronized than would occur randomly; synchrony among working spouses diminished after the 1970s; and the full-income elasticity of demand for it was higher among wives than among husbands in the 1970s but equal in the 1990s. 4) Dutch time-budget data for 1990 show that the overwhelming majority of the windfall hour that occurred when standard time resumed was used for extra sleep.

JEL classification: J20

Key words: Leisure, time use, work amenities

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1. Introduction

For many years labor supply has been the single most heavily researched topic in the subfield of labor economics (Stafford, 1986). Nearly all of this research

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has been based on data derived from questions about how many hours, weeks or years people have been engaged in market-based activities. The focus has been on the integration of workers' time to derive the fraction of some particular interval that is spent in market work. Very little research has examined time use – how individuals spend their time at work and in other activities; and almost none has examined the economic implications of *when* people engage in work and nonwork activities.

These little-studied supply-related topics can provide insights into a variety of questions that have been addressed in other ways, and often not so successfully, using more standard approaches and more commonly used data. For examples, changes in the distribution of workers' well-being depend not only on the monetary returns to work, but also on the changing distribution of such nonmonetary returns as the timing of work. The issue of jointness in a married couple's supply of labor can only be addressed if we know when the couple is working. Simply examining how the total of one spouse's hours affects the other's is not informative about their decisions on supplying labor as affected by what is presumably their desire to be together, or by their possible need for childcare. As still another example, there is an immense literature attempting to estimate pure income effects on labor supply. Yet equally important, and for obvious reasons essentially unstudied, is the pure full-income effect of an increase in available time.

The purpose in this study is not to provide a definitive list of new ways of viewing time use that might be generally interesting to economists and to labor economists/demographers especially. Rather, it is to give what I believe are some novel and interesting examples that I hope might inspire others to approach these and similarly motivated issues using the many underutilized sets of data that are available for this purpose. This is a much more fruitful endeavor than the development of ever more complex econometric models of labor supply that focus on the same standard questions of measuring wage and income effects on hours/weeks worked using standard data sets. I hope to demonstrate that moving beyond refinements to the standard model and its estimation can be useful and interesting.

Accordingly, in Sect. 2 I examine the role of work timing – when people work – as an amenity of the employment relation and consider how changes in timing in the United States might be taken as reflecting changes in the well-being of the average worker. Section 3 uses this same idea to consider how our understanding of labor-market inequality is altered when we take nonmonetary characteristics of work, in this case the timing of work, into consideration. In Sect. 4 I study the demand for work timing in the context of the household, focusing particularly on whether spouses' "togetherness" is affected by their incomes and how this demand has changed over time. Section 5 focuses on examining responses to an exogenous increase in the time at their disposal by a random sample of households. These ideas and empirical analyses are tied together by the common themes that they illustrate new ways of thinking about labor supply and leisure and that they test how shocks to the economy alter outcomes along a variety of dimensions of time use and the timing of activities.

2. Work timing as a workplace amenity

The argumentation here and in the rest of the study compares outcomes across equilibria in the labor market. Unsurprisingly, very little can be inferred outside of equilibrium, especially if the burden of the disequilibria varies across workers. The value of this standard, neoclassical approach lies in its predictive ability, so that the contribution of these analyses must be measured by whether the facts that are uncovered accord with the theory that is outlined.

It is easy to see how changes in amenities are altered when the real earnings capacities of workers in different groups change. View workers as being able to obtain a combination of real earnings, other monetary benefits (which I henceforth subsume under earnings), and nonmonetary benefits from the jobs they occupy (as originally in Rosen 1974). Workers sort themselves among jobs that differ by the amenities that the jobs offer according to their preferences for nonmonetary amenities and earnings. Workers who especially prefer amenities (e.g., are extremely averse to working at night) will sort into jobs that avoid night work. Jobs that fail to offer the amenity of day work must compensate for its absence through higher wages in order to attract workers. We will observe that otherwise identical workers obtain higher wages in those jobs, so that they may be viewed as offering premium wages (see Kostiuk 1990, for evidence on this). Because workers whose overall earnings ability is low require earnings just to get by, they will be especially willing to accept unpleasant jobs that compensate for the unpleasantness by offering higher wages.

What will happen in such a labor market as full earnings rise generally? We will observe ever-fewer workers who are willing to accept work at undesirable times. This will induce employers to: 1) Offer higher premia to attract workers to such times; but 2) Price out of the market some employers who would otherwise have conducted their business at evening/night. We should observe the price (compensating wage differential) for such work rising, while the quantity of such work falls. Indeed, if we are uncertain about the path of real earnings (perhaps, as in the United States, because of difficulties measuring indexes of living costs, Boskin et al. 1998), a good indication that real earnings have risen is that the quantity of disamenities observed in the labor market, including work at undesirable times, has fallen (barring major changes in legal restrictions on the provision of amenities/disamenities, none of which occurred in the United States during this period).

This entire discussion is from the supply side of the labor market and entirely ignores the effects of possible shocks to employers' labor demand. If technical change makes evening/night work more expensive for employers at a given set of supply conditions, we would observe a decline in the quantity of such work performed even though workers' full earnings have not risen. While this is possible and is extremely difficult to contradict, most observers of the labor market argue that technology has shifted people toward a 24-hour economy, implying that the bias in technology has been toward an *increase* in the demand for evening/night work, other things equal. Thus if we find despite this that the amount of evening/night work has declined, we can reasonably assume that supply behavior has dominated this implicit market.

To illustrate this approach I take data from the United States Current Population Survey May Supplements for 1973, 1978, 1985, 1991 and 1997 (the earliest four of which are also used for a related purpose in Hamermesh 1999a). In these few surveys (and in the May Supplements from 1974 through 1977) respondents were questioned about the starting and ending times on their main jobs: "At what time of day did ... begin (end) work on this job most days last week?" Regrettably the questions are not specific to each day of the week, but rather talk about what the worker "usually" does. The ideal, a set of repeated cross-sections of large numbers of time diaries showing exactly when people are at work for each of a number of days, is simply unavailable in the United States or elsewhere. To ensure that the workers in the sample are at work at these hours on most days, only employees with at least 20 hours of work per week are included in the analysis in this section.

From the information the respondents provided I can construct a set of 24 indicators, L_{ist}, for each worker i interviewed in year t, with the indicator equaling 1 if the responses imply that the person worked in the market at hour s, 0 if not. This is different from identifying workers as being on shifts, as has been done by, for example, Mellor (1986). Because a majority of workers on the job at, for example, 3AM would not be classified as night-shift workers (Hamermesh 1996), this hour-by-hour approach gives a fuller picture of the distribution of work.

Before examining how the distribution has changed, we need to establish whether in fact there is a consistent pattern relating work at various times of the day to workers' demographic characteristics. To save space I define the variables $EVE = 1$ if the worker was on the job at any time between 7PM and 10PM, 0 otherwise, and $NIGHT = 1$ if he/she was on the job at any time between 10PM and 6AM, 0 otherwise. I relate these variables to workers' educational attainment, their age, ethnic/racial status and other controls available in the CPS. In addition, in the some of the estimates I hold constant for the workers' detailed industry affiliation (thus controlling for potential differences caused by employers' rather than the workers' behavior).

The top row of Table 1 presents for both genders the mean fractions of employees working evenings or nights. Unsurprisingly, men are more likely to be working during these unusual hours than are women. Below these means the Table lists the coefficients from linear-probability estimates of the determinants of EVE and $NIGHT$ for all workers in the May 1997 Supplement whose usual weekly hours were 20 or more. (Probits yield qualitatively similar conclusions.) For both EVE and $NIGHT$ the first column in each pair presents estimates that exclude industry indicators, while the second includes them. The results make it very clear that evening or night work disproportionately burdens those with lower educational attainment (since the excluded category is workers with less than a high-school diploma). Similarly, the U-shaped relationship between age and the incidence of evening or night work shows that such labor is disproportionately done by younger workers or those nearing retirement. Holding constant their total workhours, the lowest probability of work outside the standard workday is among workers around age 50, roughly the peak of age-earnings profiles. This negative relationship between the probability of working evening or night and a worker's earnings ability is changed only slightly even when we account for the worker's detailed industry affiliation.

The estimates in Table 1 also provide some evidence that evening and night work is performed disproportionately by minorities, especially by African-Americans, even after accounting for racial/ethnic differences in age and educational attainment. There are essentially no differences in the probabilities of evening and night work between nonhispanic whites (the excluded category) and Hispanics. The differences in the probabilities of working evenings/nights are consistent with the notion that workers whom the labor market rewards less are more likely to work evenings or nights. By inference, evening/night work is a disamenity.

Table 1. Means and regression estimates of the determinants of evening and night work, May 1997 CPS[a]

Work in:	Men				Women			
	Evening		Night		Evening		Night	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Fraction working:	0.168		0.117		0.129		0.079	
HS grad	−0.036	−0.027	0.004	−0.001	−0.123	−0.085	−0.046	−0.023
	(0.009)	(0.008)	(0.007)	(0.007)	(0.009)	(0.009)	(0.007)	(0.007)
Some college	−0.022	−0.023	−0.004	−0.012	−0.123	−0.076	−0.055	−0.028
	(0.009)	(0.009)	(0.008)	(0.008)	(0.009)	(0.009)	(0.007)	(0.007)
College degree	−0.082	−0.080	−0.089	−0.086	−0.159	−0.096	−0.095	−0.062
	(0.009)	(0.009)	(0.008)	(0.008)	(0.009)	(0.010)	(0.007)	(0.008)
Age	−0.023	−0.017	−0.005	−0.005	−0.024	−0.020	−0.006	−0.006
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
$Age^2/100$	0.024	0.018	0.005	0.005	0.025	0.021	0.006	0.006
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
African-American	0.032	0.012	0.053	0.021	0.023	0.024	0.033	0.026
	(0.010)	(0.009)	(0.008)	(0.008)	(0.008)	(0.007)	(0.006)	(0.006)
Hispanic	0.003	−0.005	0.004	−0.001	−0.009	−0.001	0.008	0.009
	(0.009)	(0.009)	(0.008)	(0.008)	(0.009)	(0.009)	(0.007)	(0.007)
Industry controls:	No	Yes	No	Yes	No	Yes	No	Yes
R^2	0.140	0.232	0.130	0.215	0.148	0.217	0.206	0.247
N	19520				17402			

[a] The equations also control for marital status, geographic location and total hours worked. Standard errors are in parentheses below the parameters estimates here and in Tables 2, 4–6.

For each hour of the day in each year t for which the data are available I calculate:

$$\Delta F_{st} = F_{st} - F_{s73}, \quad t = 1978, 1985, 1991, 1997, \tag{1}$$

where F_{st} is the fraction of employees at work at hour s in year t, adjusted so that the average daily hours worked are unchanged over the time period. These differences thus summarize what happened to work timing in the United States in the final quarter of the twentieth century. Figures 1a and 1b use these CPS Supplements to present for male and female workers the fractions that were at work at each hour of the day. (The patterns look quite similar if the few employees working less than 20 hours per work are added to the samples.) To get a feel for the magnitude of these changes, one should note that in 1973 the fraction of men at work at Noon was 0.88, while the fraction working at 3AM was 0.09. The fractions for women were slightly lower.

The figures show very clearly that the trend was toward less work being performed by men in the evening and at night, and some of these drops are substantial. For example, the drop of over 0.02 at 3AM represents a decline of over 30% (and of over five standard errors in these samples). In percentage terms the 0.07 rise in the fraction of men at work at 7AM is smaller (around 20%), but it is clear that more work is being accomplished in the early morning hours. The decline in evening and night work did not occur between 1973 and 1978, a time

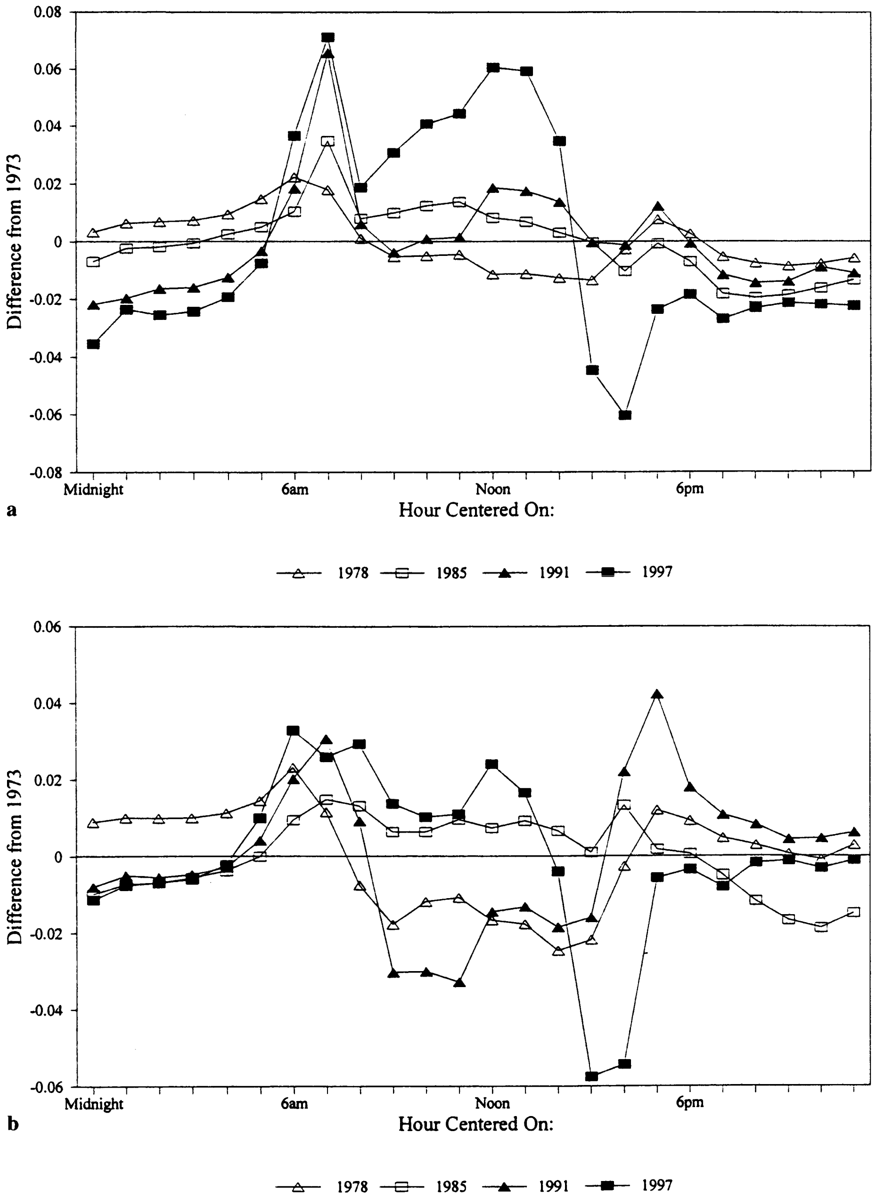

Fig. 1a,b. Differences over 1973 in fraction working. **a** Men; **b** Women

when it is quite clear that real earnings in the U.S. failed to increase; but it is fairly steady thereafter. This decline is fully consistent with rising full earnings.[1]

Among women the changes are less pronounced, with significant declines being observed in evening but not in night work. With women's wages surely increasing over this period relative to men's this one deviation in the results may be disturbing. One should remember, however, that the kinds of industries and occupations where technological change may have been most heavily biased toward night work are those that are especially female-intensive, particularly service and retail industries. Those occupations/industries may be sufficient in number that the bias of technology toward night work is sufficient to have outweighed the induced reduction in supply of night-time labor.

The main point of this section is that, almost certainly contrary to popular belief, the best evidence suggests that evening/night work in the United States has diminished in importance since the early 1970s. As shown in Hamermesh (1999a, Fig. 3), evening and night work among all workers decreased through 1991 among men in all major industries except the tiny (in the United States) agriculture sector. The same is true for evening work in all major industries among women. This is consistent with the view that workers' real earning power has increased and that they have used part of it to shift away from work at an unpleasant time. Whether this is true universally is unclear; but the approach taken here should be applicable in other economies. Examining secular changes in other labor economies would be a useful approach to understanding the changing well-being of their workers.

3. Work timing and economic inequality

In the past 20 years, whether because of increased international trade (Leamer 1996), technical change that is biased toward skilled workers (Berman et al. 1998), declines in institutions that protect low-skilled workers, or still other causes, shocks to the labor market have raised the earnings ability of skilled workers relative to unskilled workers essentially worldwide (Pereira and Martins 2000). These changes have implied a relative improvement in the prospects of those who would have earned more even without them. This should have caused those workers, even more than before, to shy away from jobs that lack such workplace amenities as desirable schedules, since their earning power has increased most. Obversely, low-skilled workers will be observed occupying an even greater fraction of the jobs that have undesirable characteristics: Because the supply of skilled workers to those jobs is reduced, employers offering them will pay higher wage premiums; and, with their earnings ability falling relative to other workers, the relative supply of lower-skilled workers to jobs offering these premiums will be higher than before.

Changes in the distribution of workplace amenities should thus mirror changes in the distribution of wages. We would expect that the widening distribution of earnings would have been accompanied by an increasingly unequal distribution of the burden of unpleasant workplace characteristics. This will be true so long as employers' ability to offer daytime jobs has not changed differentially by the skill of its workers. In other words, only skill-biased technical change in the provision of the amenity, working during the day, will cause this prediction to fail.

While it is clear that sorting in the changing implicit market for the amenity of desirable work timing will cause a change in the distribution of the amenity, the implications for inequality of full earnings – wages plus the value of the amenity – are unclear. Write full earnings E in logarithmic form as:

$$E = W + \theta D, \tag{2}$$

where θ is the premium for evening/night work, and D equals 1 if the worker works evenings/nights. Imagine a shock to the labor market that increases the variance of full earnings. Assume that the full-income elasticity of demand for the amenity exceeds unity by enough to offset the rise in θ (an assumption that is consistent with evidence showing very high income elasticities of demand for monetary benefits, Woodbury and Hamermesh 1992). We will then observe that an increase in the variance of log-wages (W) will be accompanied by an increase in the variance of E.

Having shown that workers with lower earnings potential have a greater likelihood of performing evening/night work, we can examine how patterns of work timing have changed in relation to changing earnings differences. As in the literature on earnings inequality (Juhn et al. 1993), I base the comparisons on the weekly earnings of full-time (35+ hours per week) workers. To verify that the earnings of full-time workers in these May CPS Supplements exhibit the same rise in inequality that has been noted more generally, Fig. 2 presents estimates of:

$$\Delta^2 W_t^q = [W_t^q - W_t^4] - [W_{73}^q - W_{73}^4],$$
$$q = 1, 2, 3, \quad t = 1978, 1985, 1991, 1997, \tag{3}$$

where W is the logarithm of average weekly earnings among workers in earnings quartile q in year t, and the superscript 4 refers to workers in the bottom quartile of earnings.[2] The measures $\Delta^2 W_t^q$ for men and women thus show percentage changes in average earnings within each of the three upper quartiles since 1973 compared to percentage changes in earnings in the lowest quartile.

The estimates of these double-differences in earnings are shown in Figs. 2a and 2b for men and women. The results parallel what has been demonstrated generally for the United States over this period. For both genders there has been a very sharp rise in earnings inequality since the early 1970s, with much of the increase coming between 1978 and 1985. The biggest relative increases have been in the top earnings quartile, with increases generally being somewhat larger among men than women. Similar patterns to these, and to the remaining results in this section, are shown if we disaggregate the full-time workforce by earnings decile.

The data are sorted by weekly earnings, and for each worker the fraction of his/her total workday accounted for by work at each hour s is calculated. These data were then averaged to give f_{st}^q, the fraction of all work by those in the q'th earnings quartile in year t that was performed at hour s. The measure f indicates the intensity of work at each hour by the average full-time worker in the earnings quartile. Relative changes since 1973, and thus in the burden of work at each hour of the day, can be summarized by the differences

$$\Delta^2 f_{st}^q = [f_{st}^q / f_{st}^4] - [f_{s73}^q / f_{s73}^4], \tag{4}$$

calculated as ratios to allow for convenient presentation. A ratio below one implies that workers in quartile q performed a smaller fraction of their total hours of work at hour s than did workers in the lowest earnings quartile. A negative difference means that since 1973 workers in quartile q became relatively less likely than workers in the bottom earnings quartile to work at time s.

Figure 3a shows these interquartile differences for men, while Fig. 3b presents the same information for women. To save space only the differences be-

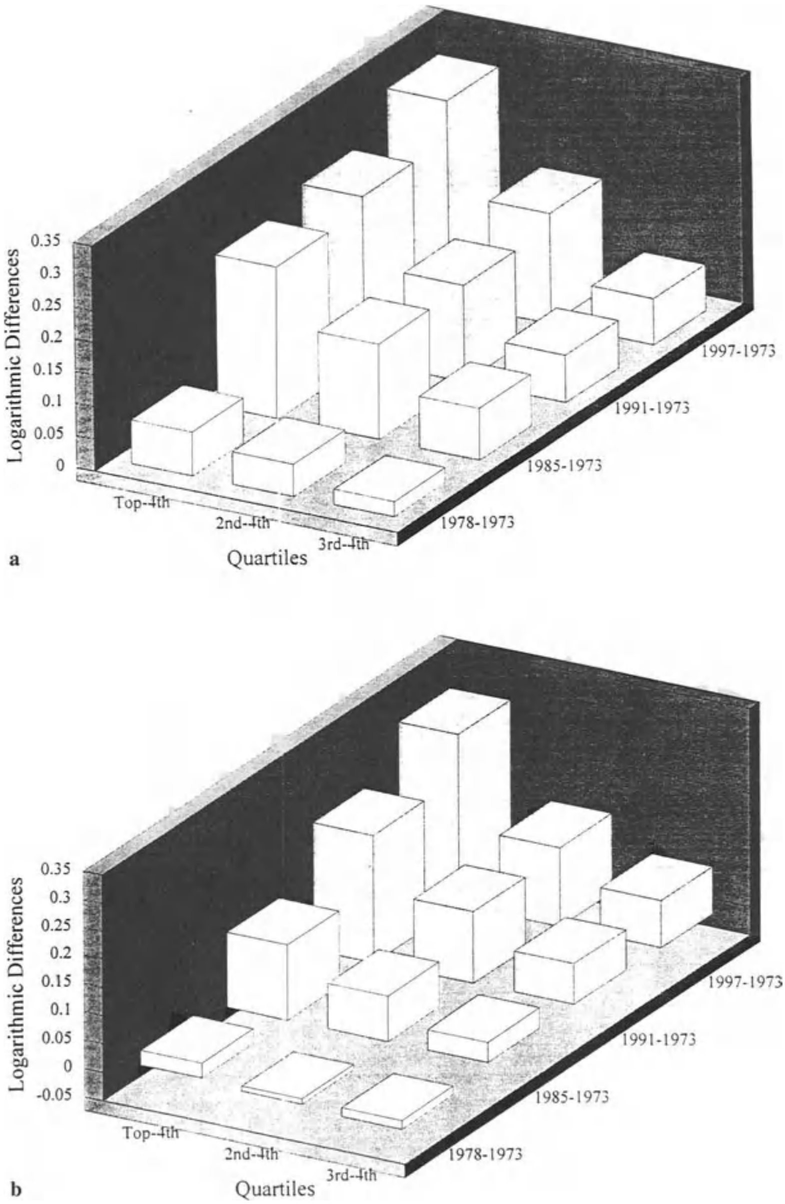

Fig. 2a,b. Double differences in weekly earnings. **a** Men 1973–1997; **b** Women 1973–1997

tween workers in the top and bottom earnings quartiles are shown. The results for workers in the second quartile (percentiles 75 to 50) look similar, while there are no major changes in timing between workers in the third and bottom quartiles (which is not surprising given the small relative changes in earnings shown in Fig. 2). While the differences are small in 1978, beginning in 1985 they started to depart from zero. In particular, for both men and women there

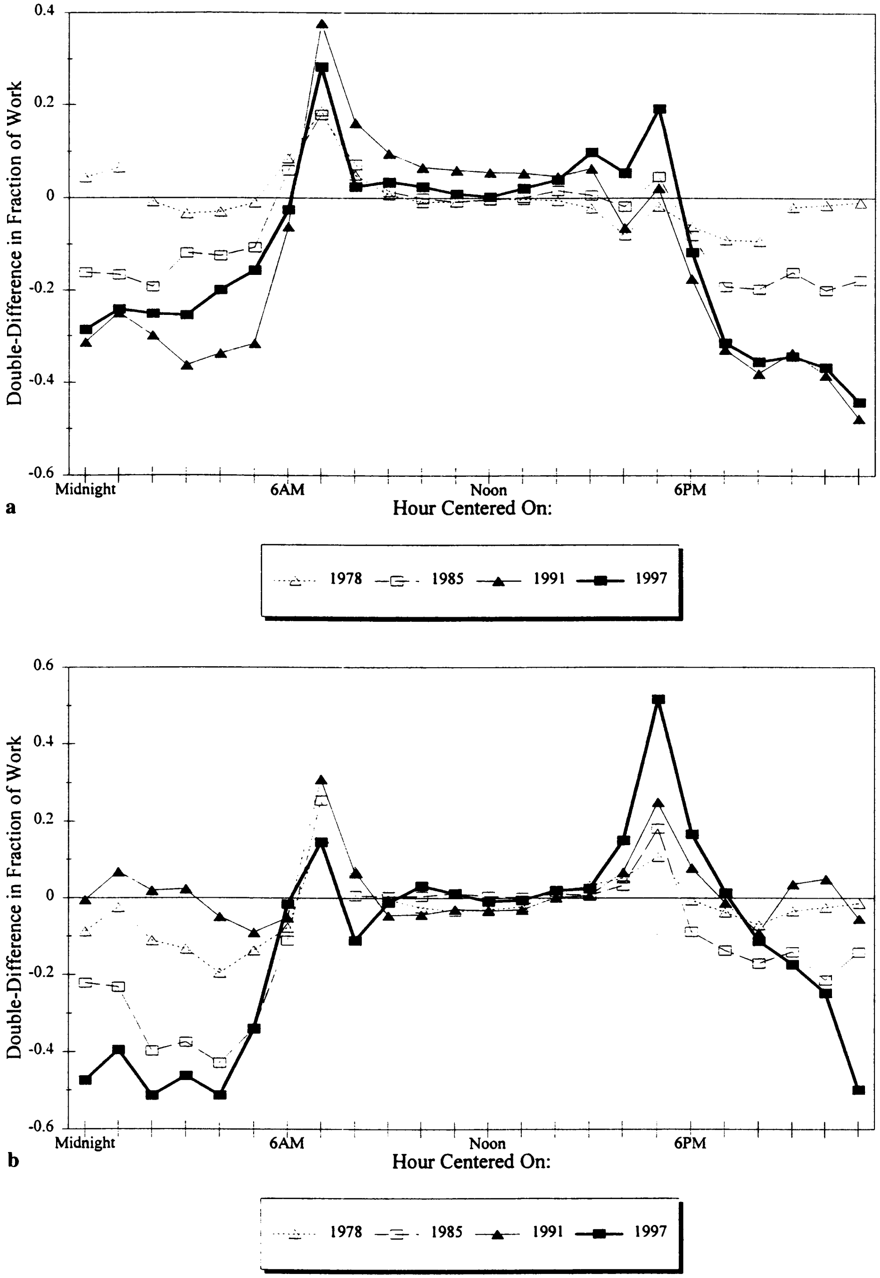

Fig. 3a,b. Top-bottom quartile differences in work timing over 1973. **a** Men; **b** Women

was a general, albeit unsteady decrease in the differences in the evening and night hours. The negative values of $\Delta^2 f_{st}^1$ between 8PM and 5AM show that the relative burden of evening and night work was increasingly borne over this quarter-century by workers in the bottom quartile of the earnings distribution.

The negative values of $\Delta^2 f_{st}^1$ between 8PM and 5AM must be offset by positive values at other times. These offsets occur especially at the fringes of the "normal" workday. Implicitly, higher-wage workers, whose total workhours have been increasing (see Juhn et al. 1991), have been spreading their workdays to early morning and late afternoon, at the same time that they have been cutting back from working in the evenings and at night (at least compared to lower-wage workers). The double differences for 1997 are quite similar for men and women; but for women the decline in evening/night work and the rise in work at the edges of the regular workday do not exhibit the same steady trend that they do among men. Since similar steady changes exist for men by major industry, but not for women, this gender difference is not a reflection of the sexes' different representations by industry.

I have demonstrated that there has been a relative decline in work at undesirable times of the day among precisely those workers whose earnings have risen relatively. To infer the strength of the relationship between changes in the incidence of evening and night work and changes in relative earnings I estimate:

$$F_{st}^q - F_{st}^4 = a + b[W_t^q - W_t^4], \quad s = 1, \ldots, 24, \tag{5}$$

where, as in Section 2, F is the fraction of employees in earnings quartile q who are at work at hour s in year t. Also included in the estimation is a pair of indicator variables for quartiles 1 and 2. Each regression is based on 15 observations, since each pools three differences (earnings quartiles 1, 2 and 3 compared to quartile 4) for each of the five years 1973, 1978, ..., 1997. Each is estimated over each hour s for men and women separately. They indicate a relationship between changes in earnings and changes in work timing, not causation: Both work timing and earnings are outcomes that are generated by a combination of workplace technologies and workers' earnings capacities and preferences.

The estimates of the slope parameters in (5) are shown for selected hours s in Table 2. They make it fairly clear that, as interquartile earnings differences have increased, in the upper earnings quartiles the probabilities of working at odd hours have decreased relative to those of workers in the lowest earnings quartiles. This is especially apparent for men (among whom interquartile earnings differences rose more rapidly than among women). Obversely, the relative probabilities of working during regular daytime hours have increased along with increases in interquartile differences in earnings.

The comparisons show clearly that widening earnings inequality has been associated with lower-wage workers bearing an increasing share of the burden of work at these times. I have explained this in terms of workers' choices of jobs and occupations. One might instead argue that it has become relatively easier for employers to schedule higher-skilled workers' jobs outside of evenings and nights. This explanation is inconsistent with the common observation that it is higher-paid managerial and clerical workers who must work unusual hours to remain part of the Internet-wired global economy. It is also inconsistent with the facts: Figures like Figures 3 calculated for managerial and clerical workers alone show the same increasing relative burden of evening/night work on low-wage workers as do graphs based only on blue-collar workers.

While it is clear that the distribution of the amenity, desirable work timing, has widened in the same direction as the distribution of earnings, it is unclear whether the distribution of full earnings has also widened – whether, as dis-

Table 2. The relation between interquartile differences in the fraction at work and interquartile differences in earnings, May CPS 1973, 1978, 1985, 1991, 1997

Work at:	Men (1)	Women (2)
Midnight	−0.105	−0.071
	(0.031)	(0.041)
3AM	−0.062	−0.033
	(0.021)	(0.037)
6AM	−0.014	0.016
	(0.036)	(0.043)
9AM	0.334	0.085
	(0.083)	(0.042)
Noon	0.241	0.209
	(0.059)	(0.061)
3PM	0.357	0.188
	(0.106)	(0.068)
6PM	−0.089	0.081
	(0.046)	(0.060)
9PM	−0.221	−0.069
	(0.037)	(0.036)

cussed above, the amenity is a luxury good. Under certain very restrictive assumptions about homotheticity of workers' preferences and employers' profit functions, the full-earnings elasticities of demand for desirable work timing far exceed unity (Hamermesh 1999b). This is consistent with evidence on the demand for monetary nonwage job characteristics such as pensions and health care (e.g., Woodbury and Hamermesh 1992). We can be quite sure that the distribution of the amenity has widened substantially in the U.S.: The burden of working at bad times has increasingly been borne by low-skilled workers. We cannot, however, be sure that price changes in this amenity have been sufficiently small to ensure that the distribution of full earnings (including this amenity) has widened more in percentage terms than the distribution of earnings.

4. Joint decision-making about the timing of leisure

The jointness of spouses' work/leisure choices cannot be inferred by concentrating on the quantities that they consume over some interval of time. Given the relatively small fractions of the week that people in developed economies typically work in the market, we could very easily find that husbands' longer weekly hours are associated with wives' longer weekly hours, holding their wage rates constant, although each one is at home while the other works. Understanding the extent of jointness in time use requires analyzing when each spouse works in the market, i.e., the extent of overlap in the spouses' use of time. A couple can consume more leisure jointly when the number of hours that both spouses are at home is greater, not when the partial correlations of their total work times are higher.

In order to analyze the instantaneous jointness of spouses' decision-making, we need to specify the household's utility in arguments defined over points in time. Let the basic unit of time be the day, divided arbitrarily into 24 hours. Then we can write the household's maximand as:

$$V(U^M([1-L_1^M],\ldots,[1-L_{24}^M]), U^F([1-L_1^F],\ldots,[1-L_{24}^F]),$$
$$U^J(Z_1^J,\ldots,Z_{24}^J), C), \tag{6}$$

where $Z_s^J = [1-L_s^M][1-L_s^F]$, C is the household's consumption, and M and F denote the husband and wife respectively. The household's monetary gains are implicitly spent entirely on the one composite (household) public good. I also assume that each hour is indivisible, with the individual either working the entire hour or enjoying leisure. Equation (6) is maximized subject to the spending constraint:

$$C = \sum_s [w_s^M L_s^M + w_s^F L_s^F], \tag{7}$$

where $w_s^j = w^j[1+\theta_s]$. Each spouse j faces an exogenous wage rate that varies over s around w^j by a percentage θ_s that is determined by the market supply and demand for labor at hour s and that faces all workers regardless of sex.

Maximizing (6) subject to (7) yields the couple's optimizing sequences of market work times, $\{L_s^M\}$ and $\{L_s^F\}$. If the sequences were integrated over the day, they would yield each spouse's daily hours supplied to the market, H^j. A spouse will be working at hour s if $w_s^j > w_s^{jr}$, the spouse's reservation wage for working at that hour. These reservation wages vary over s and may be determined jointly by the spouses' bargaining. The object of interest here is to infer whether or not the subfunction $U^J \equiv 0$, that is, whether the outcome of the spouses' bargaining reflects any interest they may have in being at home together, *conditional* on their working in the market for given numbers of hours. (Alternatively, one might observe that couples' behavior is joint but implies a preference for being apart.) Only through this approach can we examine whether consuming *synchronous* leisure matters to the couple.

To examine the possibility of jointness in the timing of potential leisure I combine all the available data from the May 1970s CPS Supplements into one data set, and combine the data from the 1991 and 1997 May Supplements into another. In each CPS Supplement I match husbands' and wives' records to create a record for the couple that generates the sequences $\{L_s^M\}$ and $\{L_s^F\}$ and uses them to create the sequences Z_s^J.[3] Any matched couple in which one spouse was age 60 or over was excluded from the sample, since the purpose is to focus on market work and its complement. Only couples with both spouses working are included in the samples, both to avoid problems with corner solutions to the maximization of (6) and because our definition of joint leisure is identically the inverse of the working spouse's market time if there is a non-working spouse. The usual CPS controls are used; and for each spouse I measure total daily hours of market work using the sequence $\{L_s^j\}$ and infer full hourly earnings using usual weekly earnings.

Whether the spouses are actually enjoying leisure jointly when $Z_s^J = 1$ is not clear. It might well be that one partner is out carousing while the other is at home; perhaps they are both home in separate parts of the house; or perhaps they are together physically but not engaged in the same activity (Larson and Richards 1994, Chapt. 5). The data do not allow us to distinguish these possi-

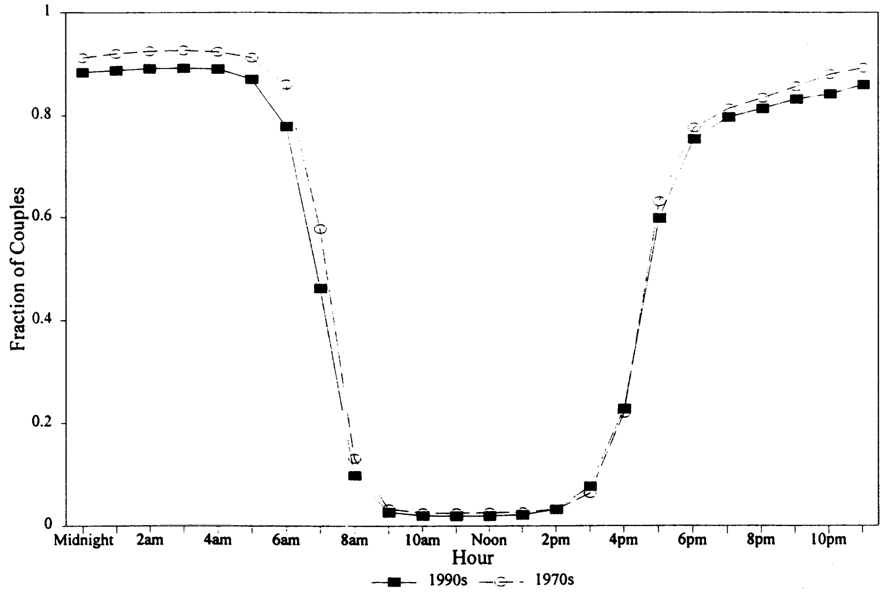

Fig. 4. Fraction of working couples at leisure, 1970s and 1990s

bilities. All we can do with these, the only available data that meet the criteria for sample size, is to examine the amount of time that the spouses could be together.

The indicator Z_s^J shows whether both partners are away from work (and thus have the possibility of consuming leisure jointly) at the same hour of the day. Figure 4 graphs this measure for the 1970s and 1990s samples. Not surprisingly, given the paucity of evening and night work, the average Z_s^J approaches one at those times of the day. What is interesting in the figure is how low Z_s^J is at the prime working times of the day among working couples. Very clearly, most members of such couples are either working at roughly the same time, or are away from work, and thus possibly consuming leisure jointly, at the same time. The figure also makes it clear that the possibility for joint leisure decreased over the two-decade interval between the samples, a substantial and statistically significant decline in the total amount of joint leisure of 0.67 hours (s.e. = 0.07). This change is the clear result of the increase in average hours of market work among working spouses.

While Fig. 4 is interesting, it merely shows that there is substantial overlap in men's and women's timing of leisure. It says nothing about whether the overlap in spouses' leisure is any different from what would be observed if we generated such measures artificially by creating pairs of randomly matched men and women. To test for the existence of jointness we need to show that the actual distribution of work timing is different from what it would be if spouses' work timing were independent. If it were independent at hour s, the fraction of couples with both spouses enjoying leisure would then just be the product of the mean fractions of husbands and wives not working, $\hat{Z}_s^J = [1 - L_s^M][1 - L_s^F]$.

To save space, in Table 3 I list the values of the differences $Z_s^J - \hat{Z}_s^J$ for selected hours only, for the 8353 couples in the 1970s sample and the 4003

Table 3. Differences, actual – predicted jointness of leisure timing among working couples, CPS samples 1973–1978, and 1991 and 1997[a]

Hour:	Year	
	1970s	1990s
Midnight	0.0016	0.0059
	(0.48)	(1.09)
3AM	0.0002	0.0051
	(0.08)	(0.97)
6AM	0.0060	0.0075
	(1.46)	(1.03)
9AM	0.0090	0.0065
	(1.62)	(0.84)
Noon	0.0100	0.0062
	(1.98)	(0.88)
3PM	0.0137	0.0169
	(2.14)	(1.76)
6PM	0.0064	0.0101
	(1.27)	(1.31)
9PM	0.0070	0.85
	(1.67)	(1.31)
N	8353	4003

[a] t-statistics in parentheses below the differences.

couples in the 1990s. As in Fig. 4, all the differences are shown as fractions. These are quite small, but all are positive and thus consistent with a demand for jointness of leisure. Moreover, and despite the relatively small samples, at many hours of the day the hypothesis that we can predict the fraction of couples in which both spouses are at work knowing only the fraction of men and women generally who are at work at that hour is rejected with at least some degree of confidence. At times when most market work is accomplished in the United States, if one spouse is at not at work the other spouse is disproportionately likely not to be at work too.

This evidence suggests that couples attempt to time their market work to provide themselves the opportunity to be together when they are not working. If, however, jointness is something that people desire, we should observe that couples with higher full incomes consume more of it – jointness should be a normal good. To examine this idea, for working couples with each spouse usually working at least 6 hours per day (implicitly at least 30 hours per week) I estimate the impact on Z_s^J of each spouse's earnings, holding constant each spouse's hours of market work and demographic characteristics.[4] I thus focus on the relative impacts of the full earnings of the husband and wife on their joint timing of work.

The estimates are presented in Table 4, with Columns (2) and (6) showing the basic results.[5] Before examining the impacts of earnings, consider the effects of extra hours of work on hours of joint leisure. The sum of the impacts of a one-unit increase in each of the spouses' workhours on their joint leisure time is $[\partial Z_s^J/\partial H^M + \partial Z_s^J/\partial H^F]$. This sum exceeds one in absolute value in both sets of data, suggesting that the spouses are unable to time marginal increases in market work in perfect synchrony. It is also interesting to note that in both samples $\partial Z_s^J/\partial H^M$ is essentially equal to $\partial Z_s^J/\partial H^F$ – jointness is reduced as

Table 4. Determinants of hours of joint leisure time, full-time working couples, 1973–1978, 1991 and 1997[a]

	1970s				1990s			
	Mean (s.d. of means)	(2)	(3)	(4)	Mean (s.d. of means)	(6)	(7)	(8)
H^M	9.107	−0.538	−0.544	−0.564	9.530	−0.698	−0.699	−0.695
	(0.016)	(0.022)	(0.022)	(0.022)	(0.042)	(0.018)	(0.018)	(0.018)
H^F	8.443	−0.597	−0.597	−0.601	8.851	−0.700	−0.699	−0.701
	(0.017)	(0.021)	(0.021)	(0.021)	(0.039)	(0.019)	(0.019)	(0.019)
$w^M/100$	2.689	0.064	0.132	0.076	6.720	0.055	0.064	0.058
	(0.016)	(0.023)	(0.018)	(0.023)	(0.072)	(0.011)	(0.008)	(0.011)
		[0.013]		[0.016]		[0.030]		[0.031]
$w^F/100$	1.462	0.276	0.132	0.237	4.487	0.079	0.064	0.064
	(0.010)	(0.037)	(0.018)	(0.037)	(0.049)	(0.016)	(0.008)	(0.016)
		[0.031]		[0.026]		[0.028]		[0.023]
$\tilde{Z}^J$	13.161				12.487			
	(0.033)				(0.061)			
Industry controls		No	No	Yes		No	No	Yes
Adjusted R^2		0.191	0.188	0.207		0.494	0.494	0.505
N			7129				3605	

[a] The samples include all couples with each spouse working at least 6 hours per day (implicitly at least 30 hours per week). Each equation also includes continuous measures of each spouse's age and indicators of each spouse's race and ethnicity, location, and calendar year. Estimated elasticities are in brackets.

much by an increase in the wife's market work as by an equal increase in the husband's, even though working wives spend fewer hours in the labor market than their husbands.

The most important result in this table is the estimated impact of each spouse's earnings, which, since the workhours of each are held constant, can be viewed as the spouses' full earnings. I thus interpret the coefficients (and the bracketed elasticities) on w^M and w^F in Columns (2) and (6) of Table 4 as partly reflecting income effects: With higher full earnings the spouses will be better able to indulge their desire for joint leisure. Jointness may also have a price in terms of a lower hourly wage that one spouse might receive because he/she chooses to consume leisure at the same time as his/her spouse, and this means that the estimated $\partial Z_s^J/\partial w^j$ also reflect a negative price effect. With this interpretation the parameter estimates imply that the income effect dominates any price effect. The elasticities are not large (0.013 for husbands, 0.031 for wives in the 1970s, 0.030 for husbands and 0.028 for wives in 1991), but they are significantly positive.[6]

There is no reason to believe that the price effects on the demand for jointness by the two spouses differ for equal increases in each w^j. We can interpret the relative magnitudes of the estimated $\partial Z_s^J/\partial w^j$ as reflecting how equal increases in each spouse's full earnings affect the couple's demand for jointness at constant prices. The equations presented in Columns (3) and (7) of Table 4 constrain the effects of the husband's and wife's earnings on their joint leisure to be identical. This constraint is soundly rejected for the 1970s sample: Raising the wife's earnings by one dollar has a larger effect on the jointness of their leisure than does raising the husband's. While evidence against the notion that couples pool their income is accumulating rapidly (Thomas 1994; Lundberg et al. 1997; Inchauste 1997), those studies all examine spending on items that might be viewed as specific to children. The evidence for the 1970s suggests that, even in their demand for an activity that is *ipso facto* joint, husbands and wives responded differently, so that a change in the relative earnings of the spouses affected the couple's consumption.

The result disappears in the data for the 1990s: The constraint implied in Column (7) cannot be rejected, and the elasticities in Column (6) are almost identical. Even though working wives in the 1990s sample still worked the same 0.7 hours less in the market per day than their husbands, an increase in their full earnings generated the same change in the couples' joint leisure as did an increase in their husbands' full earnings. Indeed, if we follow the literatures by assuming that hours supply elasticities are more positive for wives than for husbands (Pencavel 1986; Killingsworth and Heckman 1986), we can infer that by the 1990s the total effect (direct, and indirect through the spouses' total workhours) on Z_s^J of an increase in the wife's full earnings was less positive than that of an increase in her husband's full earnings.

Unless one believes that the relative price of jointness in response to higher women's wages fell over the twenty-year period, the equalization of the responses to husbands' and wives' full earnings might suggest that men's preferences for joint leisure rose to equal those of their working spouses. Alternatively, the extent of marital sorting along the dimension of preferences for jointness may have changed over this twenty-year period in such a way as to alter the mix of married couples in these CPS samples. Without much additional information we cannot distinguish between these possibilities, or between them and others.[7]

Columns (4) and (8) of Table 4 include one-digit indicators of industry affiliation for both husband and wife. Although their inclusion does not stem from the consumer model in (6) and (7), one might view them as testing whether any correlated demand-side constraints could be generating the results. Alternatively, their inclusion may allow us to account for possible discrimination in the kinds of work environments available to women. Regardless, the estimated effects of both the H^j and the w^j do not change qualitatively from the basic estimates in Columns (2) and (6).

The evidence in this section suggests strongly that the subfunction U^J in (6) is not identically zero. The most appropriate notion of complementarity in the context of time use is as an instantaneous phenomenon: Is spouses' time used in such a way as to indicate that they are better off having the opportunity to consume leisure together? Examining their instantaneous use of time, we can infer that their time use is complementary in this sense. A desire for togetherness is implicit in couples' decisions about the timing of each spouse's supply of effort to the labor market; and couples use some of their income to purchase the "good," synchronous leisure.

5. The longest day

An immense literature has tried to isolate the effects of exogenous increases in monetary wealth on consumption, labor supply and other life-cycle choices (e.g., efforts such as Holtz-Eakin et al. 1993, and Imbens et al. 1999). Exogenous increases in the other component of full income, the amount of time at the worker-consumer's disposal, have been investigated much more rarely. I am aware only of one attempt (Hamermesh 1984) that examined the responses of consumption and labor supply to exogenous differences in time endowments in the form of greater expected longevity, and one other (summarized in Biddle and Hamermesh 1990) in which people offered subjective responses about how they would spend a hypothetical increase in their endowment of time.

There is clearly room for interesting empirical research here. One can imagine, for examples, examining behavior after such unusual (and often depressing) cases as surprising cures from or diagnoses of usually fatal diseases, late-term miscarriages or stillbirths, prison early-release programs, and others. The difficulty, of course, is that data on these events and on the consumption-leisure choices of their victims or beneficiaries are difficult to come by.

There is one exogenous, albeit completely foreknown event that affects residents of most industrialized societies – the annual loss of one hour on a Sunday early in spring and the gain of one hour on a Sunday early in autumn. While this is not a perfect natural experiment – it is hardly unexpected – it provides a rare opportunity to examine how people respond to a truly exogenous change in their endowment of time. The data set that provides this opportunity is the Dutch *Tijdbestedingsonderzoek* of 1990, a time-budget study of over 3000 individuals ages 12 and up. Each respondent maintained a diary of his/her activities that he/she filled out for the previous day each morning. The diaries were kept for seven days, Sunday through Saturday. The list of activities was subsequently coded into over 200 categories, and the data are presented showing each person's activities for each of 96 quarter-hours on each of the seven sampled days.

Half the sample kept diaries for a week in early October of 1990; the other half sample kept diaries for the week before that, the Sunday of which included the day that the Netherlands went back on winter time. Thus for half the sample a diary is kept for a day on which each person's time endowment increased by 60 minutes. I include in the analysis all respondents ages 18 through 70; and, because the average respondent engaged in only 16 different activities per day (Gronau and Hamermesh 2001) and because of space constraints, I aggregate activities into twelve major categories.

For each activity a I estimate:

$$T_a = \alpha_{a1}X + \alpha_{a2}SUNDAY + \alpha_{a3}LONG + \alpha_{a4}SUNDAY \bullet LONG + \varepsilon_i,$$
$$a = 1, \ldots, 12, \tag{8}$$

where T_a is the time spent on that activity on a particular day; X is a vector of control variables including education indicators, a quadratic in age, the number of children and indicators of their ages; $LONG$ is an indicator equaling one for those respondents whose diaries cover the week including the return to winter time; and the α_{ai} are parameters to be estimated. Estimates of the α_{a4} show how the extra hour on that Sunday is spent and are essentially double differences in time use that compare behavior on Sunday by the half-sample interviewed during the long week to their behavior on other days, relative to the same difference in the other half-sample.

Table 5. Extra minutes by activity on winter-time day, the Netherlands, 1990[a]

	Men		Women	
	Married	Unmarried	Married	Unmarried
Cleaning and cooking	−0.65	−4.26	0.45	8.7
	(8.67)	(14.86)	(7.39)	(12.51)
Eating	1.04	9.97	−3.39	−4.50
	(4.80)	(8.57)	(3.60)	(7.70)
Family care	−0.49	−1.42	1.44	−0.14
	(3.44)	(3.26)	(4.61)	(5.83)
Other personal activities	5.09	−8.42	2.73	−1.55
	(2.84)	(8.19)	(2.21)	(4.43)
Organized activities	2.16	2.13	−6.23	6.65
	(5.72)	(8.91)	(3.87)	(8.71)
Radio and TV	11.42	36.82	−1.42	1.66
	(8.31)	(19.63)	(5.54)	(11.68)
Reading and writing	1.01	−9.05	−1.41	7.06
	(5.04)	(11.26)	(3.89)	(8.07)
Schooling and training	2.80	−11.33	−0.91	−0.19
	(4.38)	(10.34)	(3.20)	(7.23)
Shopping	3.45	−14.66	−2.13	−0.93
	(4.93)	(13.55)	(4.03)	(7.45)
Sleeping	43.32	27.61	52.92	41.38
	(7.48)	(16.15)	(6.39)	(12.78)
Sports and leisure	−25.64	45.15	18.76	11.24
	(12.67)	(28.51)	(10.23)	(19.53)
Work in the market	26.66	−12.54	−0.80	−9.39
	(18.79)	(38.61)	(10.33)	(22.51)
N	862	308	1315	494

[a] Each equation also includes measures of education, age, and the number of children and their ages.

Before considering the estimates of (8) it is crucial to be clear what the equation does not necessarily show. First, the results describe behavior on a Sunday, a day in which, especially in the Netherlands in 1990, market work was a quite unusual phenomenon even among workers classified as full-time. Also, and most important, while I do not expect that people adjusted to the impending "gain" of one hour over the entire six months since the country "lost" an hour in the previous spring, it is possible that some respondents adjusted by altering their behavior on the Saturday before the "gain" of one hour. (For example, some people, including this author, may set their clocks back on the Saturday before standard time begins.) Thus to the extent that people preadjust their behavior and take account of the future exogenous increase in their time endowments, our estimates will fail to depict the full set of responses to this temporary increase in full incomes.

With these caveats in mind, consider the estimates of (8) that are presented by marital status for men and women separately in Table 5. These are least-squares regression coefficients that do not account for the substantial left-censoring that occurs in many of activities. I present them for ease of interpretation and because the tobit estimates of the equations yield results that are qualitatively identical. The extra hour that is gained when the country went on standard time is used overwhelmingly for additional sleep.

Table 6. Extra minutes by activity on winter-time day, the Netherlands, 1990, married by presence of young children

	Men		Women	
	No children under 13	Children under 13	No Children under 13	Children under 13
Cleaning and cooking	4.00	−3.10	1.91	−1.88
	(10.88)	(13.15)	(9.73)	(10.48)
Eating	4.51	−2.60	−2.51	−4.83
	(6.03)	(7.29)	(4.74)	(5.11)
Family care	−2.48	6.05	0.82	1.42
	(4.30)	(5.20)	(6.08)	(6.55)
Other personal activities	−3.45	−7.21	3.76	1.66
	(3.57)	(4.31)	(2.91)	(3.13)
Organized activities	0.13	6.85	−5.20	−7.23
	(7.18)	(8.68)	(5.10)	(5.50)
Radio and TV	14.05	9.02	−17.57	17.28
	(10.44)	(12.63)	(7.30)	(7.86)
Reading and writing	4.33	−2.73	0.08	−4.16
	(6.33)	(7.66)	(5.12)	(5.51)
Schooling and training	6.23	−2.52	0.77	−3.06
	(5.51)	(6.66)	(4.22)	(7.31)
Shopping	3.64	4.98	−3.58	−0.88
	(6.19)	(7.49)	(5.30)	(5.71)
Sleeping	45.01	44.67	47.48	59.88
	(9.40)	(11.36)	(8.42)	(9.07)
Sports and leisure	−46.27	9.05	31.50	3.48
	(15.91)	(19.24)	(13.47)	(14.51)
Work in the market	30.30	−2.48	2.53	−1.68
	(23.49)	(28.40)	(13.60)	(14.65)
N	519	343	710	605

Indeed, the only group for which sleep accounts for less than half of the extra hour is unmarried men. Among this group the extra time is used for sports/leisure and radio/TV watching more than for additional sleep (on this Sunday).[8]

A bit more can be learned by disaggregating married men and women, the large majority of the sample, by the presence of young (under age 13) children. These results of this disaggregation are shown in Table 6. Quite remarkably, and unlike the other three groups, married women with young children "spend" the extra hour entirely on extra sleep, corroborating at the margin the results in Biddle and Hamermesh (1990) for this group on average. Very little else in these regressions is statistically significant, except for the decline in sports/leisure time among husbands without young children and the increase in this same category among wives without young children. The source of these latter effects is absolutely unclear.

The essential result of this little exercise is that the large majority of the exogenous increase in time that occurs every autumn is used for sleep. This marginal effect far exceeds the average propensity to spend time on sleep of roughly 1/3. The short-run full-income elasticity at the margin on Sundays through an increase in the endowment of time is very high for sleep, but quite low for all other activities.

6. Conclusions and new directions

The results make several new facts clear. Between 1973 and 1997 the burden of inequality in the job disamenity, working at a generally unpleasant time of day, appears to have shifted in the same direction as the burden of earnings inequality. This suggests that measures of changes in earnings inequality understate the extent of change in inequality in the overall returns to market work. There is clear evidence that couples arrange their work schedules to allow time for leisure that they consume jointly. Moreover, the demand for joint leisure is not inferior – all else equal, those couples with higher earning capacities enjoy more of it. Finally, an expected windfall of time, the one-hour gain that occurs every autumn when clocks are turned back to standard time, is consumed mainly as sleep, especially by married women with young children.

These few examples here have been designed to demonstrate that there are many potential avenues for learning about time use beyond the standard ones of examining weekly or other aggregations of reported hours of work. There are some cases, such as issues of spouses' togetherness, where going beyond standard analyses is the only way to understand the underlying behavior. There are others where this approach can generate tests of ideas examined in other contexts, such as markets for amenities, economic inequality, and power relationships within households, that may expand upon and possibly surpass conventional approaches in their ability to allow us to understand behavior.

Endnotes

[1] At least through 1991 these changes occurred independently of any changes in demographics or in the distribution of workers across one-digit industries or occupations (Hamermesh 1999b).

[2] I multiply top-coded earnings by 1.5, as is common in this literature. Unlike the literature on earnings inequality, which compares earnings across points in the distribution, these calculations are based on averages across workers in different quartiles. This is done for comparison purposes to the distribution of work timing: It would make no sense to compare work histories of those few individuals who happen to be at particular points in an earnings distribution.

[3] Individuals were matched as spouses based upon listing as household head or spouse and on line number (the person's position in the household) in the CPS records. Only people whose marital status was denoted as married, spouse present, are included in the match. Given the possibility that unmarried siblings are in the same residence, the data do not allow us to perform similar analyses for cohabiting unmarried couples.

[4] If we include all working couples the results are quite similar, except for a substantially lower coefficient on wife's hours. The same measures of the presence of children cannot be included for both samples, so that for purposes of comparability I exclude them in the Table. Nonetheless, when the equations for the 1970s and 1990s are expanded to include indicators of the presence of children and their ages, the conclusions about the impacts of hours and weekly earnings on jointness are unaffected.

[5] A regional difference exists in the demand for jointness. In the "Rust Belt," defined here as the New England, Mid-Atlantic and East North Central subregions, joint leisure was a significant 0.4 hours per day lower in both the 1970s and the 1990s than in the rest of the United States among otherwise identical couples. Whether this reflects differences in tastes or differences in the (unmeasured) constraints on couples' choices is not clear.

[6] An additional test of the validity of this approach asks how well one could have predicted Z^J_{90s} with knowledge of the structure of the relationship in the 1970s and the means of the determinants of Z^J_{90s}. Using the PCE deflator to adjust 1990s wages, and the coefficients in Columns (2) and (6), this decomposition shows that only 26% of the change in Z^J between the 1970s and the 1990s was explicable by changes in the independent variables. Most of the change resulted from changes in the coefficients.

[7] Among married couples under age 60, in the 1970s the likelihood that both spouses were working was lower than in the 1990s. This selectivity (of wives into the work force) may affect our results, but the direction of any bias is unclear, as it depends on the nature of the selectivity. This is not chiefly an issue of wives' market wages, since we hold those constant. One reasonable possibility is that wives who worked in the 1970s had relatively more power compared to their husbands than wives in the 1990s. If so, accounting for this change in selectivity would strengthen the inference that husbands' and wives' preferences for joint leisure converged over this time interval.

[8] The data set regrettably contains no information on individuals' wage rates, but it does have information on the household's net income. I included this variable in reestimates of the equations presented here, with little change in the estimated α_{ai}.

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The allocation of time and goods in household activities: A test of separability

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Abstract. In this paper, we test for the weak separability hypothesis imposed by the household production model between goods and time inputs used in the production of different commodities. Our data come from a French survey which reports both expenditures and time that households devote to some activities. The results allow us to show that the weak separability assumption cannot be rejected only when households are strongly time constrained. In the opposite case, home time uses are found to be nonseparable.

JEL classification: D13

Key words: Conditional demand functions, household production model, separability test

1. Introduction

By formulating the household production theory, Becker (1965) provided a new approach to the theory of household behaviour and laid the foundations of new home economics. This 'new' approach assumes that households combine market goods and time in household production functions to produce commodities that directly enter their utility function. This theory is very useful in the analysis of behaviours ignored by traditional theory and related to such diverse fields as the allocation of nonmarket time, education or fertility.

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For practical reasons, empirical studies devoted to household production theory are not based directly on Becker's model, but on a simplified version. The absence of a direct measure of commodities indeed lead researchers to focus on the allocation of time and goods between household activities (Pollak and Wachter 1975).[1] That is, the household production functions are introduced into the utility function to obtain what Michael and Becker (1973) call the 'derived utility function', in terms of market goods, time and environmental variables (durables, for example). In this case, estimation of the model only depends on the econometrician's ability to distinguish goods and time use in activity i from their use in other activities. The main characteristic of this function is to be weakly separable in the goods and time devoted to the production of a given commodity. For example, inputs in meal preparation are weakly separable from inputs in dressmaking. This restriction is extremely important since it partitions goods and time uses in groups of substitutes and complements. However, it has never been tested, for lack of satisfactory data.

The purpose of this paper is to provide such a test. We use a data set that is particularly suited for this purpose: the French *Modes de Vie* survey, led from November 1988 to November 1989 by the *Institut National de la Statistique et des Etudes Economiques* (INSEE). This survey contains information on several wide domains of household production: meal preparation; cleaning house; washing dishes, laundry; child care; dressmaking, seaming, knitting; gardening; pottering about; hunting, fishing and gathering. For each domain, the survey is interested in the way the household proceeds, the nature and the value of the ingredients it uses, the time it spends, the quantity and the nature of goods or services it produces. In addition to the questionnaire, a time diary records all activities by the head of the household and his or her conjoint, and an expenditures and fitting book allows the household to note all its expenditures for a week and to describe its fitting in durable goods. A last book records the contents of the deep freezer (if the family has one) and the stock of homemade tinned. Although this survey studies many household activities, we focus here on the 'meal preparation' activity, which seems to be the most appropriate given the problem we want to deal with.

Clearly, our aim in this paper is not to test for the empirical validity of alternative models of decision within the household (as in Fortin and Lacroix 1997). We are only interested here in the test of the separability hypothesis. This should allow us to answer the question: do households allocate time and goods in the way stated by the home production model?

The paper is organised as follows. In Sect. 2 the separability test principle is presented. In Sect. 3 we report the empirical results. Section 4 contains some concluding remarks.

2. Separability test principle

We consider a household composed of at least two adults, a male and a female, with a utility function of the form

$$U = U(z_1, z_2, x_1', l_m, l_f, d), \tag{1}$$

where z_1 and z_2 are the quantities of commodities resulting respectively from the meal preparation activity and other household activities, x_1' is the 'quantity' of meals taken outside the home (*a priori* substitutable for z_1), l_m and l_f

represent the pure leisure time of male and female, respectively, and d denotes a vector of socio-demographic variables. The household production obeys the following relations

$$z_i = z_i(x_i, t_{im}, t_{if}), \quad i = 1, 2, \tag{2}$$

where x_i is the quantity of market goods devoted to activity i, t_{im} and t_{if} are the time devoted to activity i by male and female, respectively.[2] The household faces a budget constraint

$$p_1 x_1 + p_2 x_2 + p_1' x_1' = w_m n_m + w_f n_f + v, \tag{3}$$

where p_1, p_2 and p_1' are the prices of x_1, x_2 and x_1', respectively,[3] w_m and w_f are respectively the male and female wages (which can also be viewed as the prices of time), n_m and n_f denote the time devoted to market work by male and female, respectively, and v is nonlabor income. In addition, the household faces time constraints

$$l_j + t_{1j} + t_{2j} + n_j = T, \quad j = m, f, \tag{4}$$

where T represents total time available for j. These three last constraints, (3) and (4), can be regrouped into the full income constraint

$$p_1 x_1 + p_2 x_2 + p_1' x_1' + \sum_{j=m,f} w_j (l_j + t_{1j} + t_{2j}) = \sum_{j=m,f} w_j T + v = S, \tag{5}$$

where S is full income, that is the income the household would earn if both male and female spent all their time on market work. Then, the household problem is to choose the quantities of goods and time maximising its utility function (1) subject to the technological (2) and the full income (5) constraints.

When the household production functions (2) are introduced into the utility function (1), one obtains the derived utility function, in terms of market goods and time

$$U = U(x_1, t_{1m}, t_{1f}, x_2, t_{2m}, t_{2f}, x_1', l_m, l_f, d). \tag{6}$$

According to the household production theory, the derived utility function is weakly separable in the goods and time used to produce a given commodity, that is inputs x_1, t_{1m} and t_{1f} are weakly separable from x_2, t_{2m} and t_{2f}. In other words, marginal rates of substitution for pairs of inputs in meal preparation are functionally independent of the quantities of inputs in other household activities. In this case, relation (6) becomes

$$U = U[u_1(x_1, t_{1m}, t_{1f}, x_1', l_m, l_f, d), x_2, t_{2m}, t_{2f}, x_1', l_m, l_f, d]. \tag{7}$$

Then, optimal quantities are obtained by maximising the derived utility function, (7) if it is consistent with the theory, (6) if it is not, subject to the full income constraint (5).

Now, assume that x_1', l_m, l_f, x_2, t_{2m} and t_{2f} are preallocated, in the sense that the household's consumption of these goods is determined before he enters the market (Pollak 1969, 1971). The household is then supposed to choose quantities of x_1, t_{1m} and t_{1f} so as to maximise (6) or (7) subject to the new full income constraint

$$p_1 x_1 + w_m t_{1m} + w_f t_{1f} = y, \tag{8}$$

where y is total expenditure devoted to meal preparation inputs, and the additional constraints

$$x_1' = \bar{x}_1', l_m = \bar{l}_m, l_f = \bar{l}_f, x_2 = \bar{x}_2, t_{2m} = \bar{t}_{2m}, \quad \text{and} \quad t_{2f} = \bar{t}_{2f}. \tag{9}$$

The maximisation of (6) or (7) subject to (8) and (9) yields two different conditional demand systems for meal preparation inputs (Pollak 1971 and Browning and Meghir 1991). The one related to the nonseparable derived utility function (6) is the following

$$q = g(p_1, w_m, w_f, y, x_1', l_m, l_f, x_2, t_{2m}, t_{2f}, d), \tag{10}$$

where q is the vector of inputs in meal preparation, while the one corresponding to the weakly separable derived utility function (7) is of the form

$$q = h(p_1, w_m, w_f, y, x_1', l_m, l_f, d). \tag{11}$$

Therefore, a simple way to know whether, in the derived utility function, inputs used in the production of z_1 are weakly separable from inputs used in the production of z_2, consists in testing whether the coefficients associated with x_2, t_{2m} and t_{2f} in (10) are not significantly different from zero.

From an empirical point of view, we have to mention the following problems related to our data set. First, the survey records no prices. Actually, x_1 and x_2 are market goods expenditures devoted to meal preparation and to other household activities, respectively, and x_1' is expenditures devoted to food-out. Moreover, hourly wages are not available in the data set. This latter provides annual wages, but not the number of hours worked during the year. As we shall see in the next section, only proxies can be constructed.

Without price indices or good measures of wages, we cannot estimate a sophisticated demand system. So, we consider the estimation of a simple Working-Leser model, without prices or wages.[4] In this case, the conditional demand functions (10) can be written as

$$s_i = \alpha_{1i} + \alpha_{2i} \ln y + \beta_{1i} x_1' + \beta_{2i} l_m + \beta_{3i} l_f + \gamma_{1i} x_2 + \gamma_{2i} t_{2m} + \gamma_{3i} t_{2f} + \delta_i d + \varepsilon_i, \tag{12}$$

for $i = 1, 2, 3$, where α_i, β_i, γ_i and δ_i are unknown parameters (the α_i's, β_i's and γ_i's are scalars, and δ_i is a vector), ε_i is an error term with the usual properties, and where s_1, s_2 and s_3 are the shares of expenditures on inputs x_1, t_{1m} and t_{1f}, respectively, in total expenditure devoted to meal preparation, $y = x_1 + w_m t_{1m} + w_f t_{1f}$.[5] Formally, $s_1 = x_1/y$, $s_2 = w_m t_{1m}/y$, and $s_3 = w_f t_{1f}/y$. Estimating (12) equation by equation by ordinary least squares (OLS) yields a set of parameters that satisfies adding-up automatically. Note that wages appear in the right hand side of each equation through total expenditure. This latter will thus have to be instrumented, not only because of its simultaneous determination with expenditure shares, but also because of the measurement errors on wages.

Given this parameterisation, the test to know whether x_1, t_{1m} and t_{1f} are weakly separable from x_2, t_{2m} and t_{2f} simply consists in testing whether the parameters γ_i are not significantly different from zero. Furthermore, if the form of the derived utility function is really given by (7), that is x_1', l_m and l_f are nonseparable from x_1, t_{1m} and t_{1f}, the parameters β_i in (12) have to be significantly different from zero.

3. Empirical results

The *Modes de vie* survey records household expenditures on a weekly basis. For each household, we constructed x_1 as the total amount spent on food products and drinks, x_1' as the amount spent on food-out, and x_2 as the difference between total expenditure and the sum of x_1 and x_1'.

Each time in the survey is reported for one day, which may be either a weekday or a weekend day. Unfortunately, we have no information about the day the household has answered. Time recordings are thus not representative of the household average timetable, and we shall have to be very cautious in interpreting the results below. We defined t_{1m} and t_{1f} as the time spent on meal preparation and food shopping by male and female, respectively, t_{2m} and t_{2f} as the difference between total time devoted to household activities by male and female and t_{1m} and t_{1f}, respectively, l_m and l_f as the pure leisure time of male and female, respectively. For convenience, expenditures and time were then expressed as monthly expenditures and monthly time. Note that these conversions cannot alter the results of the test.

Only male and female yearly wage incomes are available in the survey.[6] A monthly wage can however be computed by dividing yearly wage income by the number of months in the year on which the income has been earned. But we cannot compute hourly wages because the number of hours worked during the month is not recorded. We only have the time worked by male and female for the time recording day, which cannot be used to the extent that it includes commuting time and that it should be zero or close to zero for a large number of the households who have answered a weekend day. As an alternative, we exploited an information that is available in the data set and that gives the ratio, lower than or equal to one, of the time worked by each individual part-time or full-time employed to the time he would have worked if he had been full-time employed. For example, a ratio of 0.5 means that the individual have worked half-time. So, to obtain the monthly working time of each individual, we had to multiply that ratio by the number of hours worked by a full-time worker. Given that all individuals kept in our subsample are wage-earners, we assumed this number to be 39 hours a week (the legal working duration in France at the time of the survey), that is 169 hours per month. We finally determined proxies for hourly wages w_m and w_f by dividing monthly wages by the resulting working time. Of course, this calculation does not take into account the fact that some high-wage workers may work more than the legal duration. Hourly wages will in this case be overestimated and, since wages enter the right hand side of (12) through total expenditure, a bias in the parameter estimates may arise. As pointed out at the end of the previous section, standard instrumental variable techniques will therefore have to be applied to obtain consistent estimates.

In the vector d, we included male age, seven dummy variables for the degree obtained by male (from no degree to university degree, classified by increasing order of education level), the number of children in each of four age groups (0 to 1 year old, 2 to 7 years old, 8 to 14 years old, and 15 to 24 years old), and a dummy variable indicating whether the household is homeowner.[7]

From our initial data set, we selected a subsample of 1099 households composed of at least a male and a female who were both working as wage-earners. For these households, we observed a significant number of null observations for the dependent variable of the second equation: the share of male time expenditures in total expenditure devoted to meal preparation was equal to zero in 239 cases, while the two other dependent variables, the shares of market goods and female time expenditures in total expenditure devoted to meal preparation, were null in only 5 and 20 cases, respectively. To avoid corner solutions, we deleted observations for which at least one meal preparation input was zero, keeping then 840 households. Our aim being only to test for the separability on a part of the population, without extending the conclusions to the whole population, we do not think that the bias that may result from this selection has any importance. Moreover, although descriptive statistics show some observable differences between the sample of corner solutions and the sample of interior solutions,[8] none of them is found to be significant. Table 1, column (1), contains some descriptive statistics for the main variables used in the estimation.

As we previously noticed, total expenditure cannot be taken as exogenous in the conditional demand system, because it is simultaneously determined with market goods and time expenditures devoted to meal preparation, and because it is a function of wages that are not accurately measured. So, we instrumented it, not with income since this latter is used to determine wages, but with socio-demographic variables: male age and its square, degree obtained by male, two dummy variables indicating whether the male and the female have the french nationality, the number of children in each of the four age groups, a dummy variable equals to one if the household is homeowner and zero otherwise, a dummy variable indicating whether the male profession has a public status, and seven dummy variables for the household city size (classified by increasing order of number of inhabitants). Total expenditure is not the only variable that may be endogenous however. Individuals with long hours of work (and thus high hourly wages) will presumably have few hours to devote to leisure, to other household activities and/or to meal preparation, and will probably spend more on market goods. We would thus expect these variables to also be correlated to the error terms of (12). We performed Hausman (1978) exogeneity tests for food-out expenditures, inputs in other household activities, and male and female leisure time, together with Sargan tests for orthogonality of instruments. These latter were the same socio-demographic variables as those used in the instrumentation of total expenditure. This one was instrumented both under the null and under the alternative. Instruments required for identification are male age squared, and the dummy variables related to the male and female french nationality, to the male profession public status, and to the household city size. Sargan statistics show that variables used as instruments in the demand system have the required properties.[9] Yet, the endogeneity hypothesis is strongly rejected by Hausman statistics. It is also strongly rejected when we focus on inputs devoted to other household activities only. These results can be interpreted in two different ways. Either the instruments, which are not correlated with the residuals of the demand equations, are not correlated enough with the variables to instrument, or the previous hypothesis of preallocation is verified. Unfortunately, we have no mean to settle between these two interpretations. Estimation results are presented in Table 2.

Table 2, column (1), shows that, among the 19 variables considered, 6 are significant. The share of market goods expenditures in total expenditure devoted to meal preparation increases with the number of children, regardless of their age, and decreases with the time spent on other household activities by

Table 1. Descriptive statistics: means and standard deviations

Variables	(1)		(2)		(3)	
Market goods expenditures in francs per month						
Meal preparation (x_1)	2486.2728	(1577.1772)	2387.4318	(1538.5040)	2569.9075	(1606.1093)
Other household activities (x_2)	6519.5590	(18005.8972)	5284.7576	(14108.8877)	7564.3910	(20697.6929)
Food-out expenditures in francs per month (x_1')	213.9306	(606.4244)	179.4978	(507.7413)	243.0659	(678.0127)
Male time in hours per month						
Meal preparation (t_{1m})	21.3587	(23.8584)	16.6607	(17.2779)	25.3339	(27.6565)
Other household activities (t_{2m})	59.7060	(63.6101)	34.4972	(29.6031)	159.8527	(149.4683)
Leisure (l_m)	95.9543	(65.1100)	70.4047	(37.1310)	117.5732	(75.1412)
Female time in hours per month						
Meal preparation (t_{1f})	42.9147	(31.5967)	30.4470	(17.8428)	53.4643	(36.5011)
Other household activities (t_{2f})	94.3605	(63.6112)	62.7294	(34.0346)	121.1252	(70.2256)
Leisure (l_f)	73.5057	(54.3019)	48.4152	(31.6523)	94.7362	(60.1354)
Wages in francs per hour						
Male wage (w_m)	51.0040	(24.0932)	51.5819	(25.3330)	50.5149	(23.0089)
Female wage (w_f)	41.0593	(35.4567)	43.5025	(44.1942)	38.9921	(25.7285)
Total expenditure for meal preparation in francs per month (y)	5264.4297	(3921.3830)	4442.5778	(2178.2746)	5959.8429	(4831.8401)
Share in total expenditure for meal preparation						
Market goods expenditures (s_1)	0.4904	(0.1499)	0.5262	(0.1827)	0.4600	(0.1860)
Male time expenditures (s_2)	0.1933	(0.1632)	0.1861	(0.1468)	0.1993	(0.1524)
Female time expenditures (s_3)	0.3164	(0.1721)	0.2877	(0.1615)	0.3406	(0.1772)
Number of observations	840		385		455	

Notes: Standard deviations in parentheses.
(1) Means based on the complete subsample; (2) means based on the weekday subsample; (3) means based on the weekend day subsample.

Table 2. Estimation results based on the complete subsample

Independent variables	(1)		(2)		(3)	
Instrumented logarithm of total expenditure ($\ln y$)	−8409.6	(−0.785)	5104.0	(0.604)	33053.6	(0.340)
Food-out expenditures (x'_1)	0.6	(0.593)	−0.1	(−0.099)	−0.6	(−0.568)
Male leisure time (l_m)	−16.6	(−1.471)	11.2	(1.256)	5.4	(0.530)
Female leisure time (l_f)	−8.3	(−0.600)	6.2	(0.567)	2.1	(0.169)
Other household activities						
Market goods expenditures (x_2)	0.0	(0.895)	0.0	(0.048)	0.0	(−1.030)
Male time (t_{2m})	−28.9	(−2.723)[a]	39.3	(4.686)[a]	−10.4	(−1.076)
Female time (t_{2f})	−34.6	(−3.133)[a]	−32.9	(−3.780)[a]	67.5	(6.748)[a]
Male age	283.3	(1.381)	−42.1	(−0.260)	−241.3	(−1.297)
Degree obtained by male						
Certificat d'Etudes Primaires	−2484.0	(−0.946)	471.0	(0.227)	2013.0	(0.846)
Certificat d'Aptitude Professionnelle	856.7	(0.358)	−673.9	(−0.357)	−182.9	(−0.084)
Brevet des Collèges	−1984.0	(−0.709)	2695.3	(1.221)	−711.3	(−0.281)
Brevet d'Enseignement Professionnel	−3882.8	(−0.899)	1462.3	(0.429)	2420.5	(0.618)
Baccalauréat	778.1	(0.215)	−98.0	(−0.034)	−680.1	(−0.207)
University degree	−795.7	(−0.228)	3346.6	(1.214)	−2550.8	(−0.806)
Number of children						
0 to 1 year old	7820.0	(3.203)[a]	1039.1	(0.539)	−8859.1	(−4.004)[a]
2 to 7 years old	3384.6	(2.506)[b]	−690.9	(−0.648)	−2693.7	(−2.201)[b]
8 to 14 years old	4442.6	(2.509)[b]	−2617.2	(−1.873)	−1825.5	(−1.137)
15 to 24 years old	4692.7	(2.964)[a]	−1494.4	(−1.196)	−3198.4	(−2.229)[b]
Homeowner	−959.9	(−0.635)	−1553.8	(−1.303)	2513.7	(1.835)
Constant	111003.7	(1.379)	−20766.8	(−0.327)	9763.1	(0.134)
Number of observations	840		840		840	

Notes: No degree is the reference for degree obtained by male; coefficients are multiplied by 10^5; asymptotic t-statistics in parentheses.
[a] Significant at the 1% level; [b] significant at the 5% level.
(1) Dependent variable is the share of market goods expenditures in total expenditure devoted to meal preparation (s_1); (2) dependent variable is the share of male time expenditures in total expenditure devoted to meal preparation (s_2); (3) dependent variable is the share of female time expenditures in total expenditure devoted to meal preparation (s_3).

male and female. Only 2 variables are significant in column (2). The share of male time expenditures in total expenditure devoted to meal preparation increases with the time spent on other household activities by male, and decreases with female time spent on other household activities. Finally, 4 variables are significant in column (3). The share of female time expenditures in total expenditure devoted to meal preparation increases with the time spent on other household activities by female, and decreases with the number of children of 0 to 1 year old, of 2 to 7 years old, and of 15 to 24 years old.

Note that food-out expenditures, market goods expenditures devoted to other household activities, and male and female leisure time are never significant, hence suggesting that x_1, t_{1m} and t_{1f} are weakly separable from x_1', x_2, l_m and l_f. Furthermore, 5 variables are significant in at least two equations. These are the number of children of 0 to 1 year old, of 2 to 7 years old, of 15 to 24 years old, and especially the time spent on other household activities by male and female. The only variable that is significant in all three system equations is the time spent by female on other household activities. Hence, x_1 and t_{1m} are nonseparable from t_{2m} and t_{2f}, and t_{1f} is nonseparable from t_{2f}. This is clearly not consistent with the household production model.

We mentioned earlier that the day the household has reported his time uses might be either a weekday or a weekend day, and that we had not this information. Given that behaviours observed during a weekday are likely to be quite different from behaviours observed during a weekend day, we might wonder whether the previous results are not due to the lack of information about the time recording day. To answer this question, we performed OLS regressions of the three conditional demand functions, by adding in all equations a dummy variable constructed so as to indicate if the household has answered a weekday or a weekend day. Actually, more than a weekday or a weekend day answer, this dummy variable indicates the level of the time constraint faced by the household. It was constructed on the basis of the time worked by male and female during the time recording day. If both male and female had worked at least seven hours, they were assumed to have answered a weekday, and the dummy variable was set to 1. Conversely, if male or female had worked less than seven hours, they were assumed to have answered a weekend day, and the dummy variable was set to 0. We chose seven hours to be consistent with the hypothesis that a full-time wage-earner works 39 hours a week.

Estimation shows that the dummy variable is significant in the first and the third equations. Its coefficients take the expected sign: when the household is supposed to have answered a weekday, the share of market goods expenditures increases and the share of female time expenditures decreases. Concerning the other variables, there is only little change for the second and the third equations. But for the first equation, the time spent by male and female on other household activities is no longer significant. The effect of the dummy variable dominate the effects of these two quantities. Thus, the fact that the household has answered a weekday or a weekend day seems to have some importance.

So, we divided the previous subsample into two subsamples, according to the value of the dummy variable. The first subsample consists of the households assumed to have answered a weekday, and the second subsample consists of the households assumed to have answered a weekend day. Some descriptive statistics for the main variables used in the estimation are presented for these two subsamples in Table 1, columns (2) and (3), respectively. Then, we reestimated the conditional demand system on the two subsamples. Table 3 presents the results obtained on the weekday subsample, and Table 4 presents the results obtained on the weekend day subsample.

Table 3 shows that the only significant variable for all three system equations is the number of children of 15 to 24 years old in column (1). None of the preallocated quantities is significant, hence suggesting that the derived utility function is weakly separable for a weekday. Conversely, Table 4 gives results that are close to those obtained on the complete subsample. The only differences are that the time spent on other household activities by male and the number of children of 15 to 24 years old in column (1), and the number of children of 8 to 14 years old in column (3) are no longer significant. But as for the complete subsample, the time spent on other household activities by female is significant in all three system equations.

The difference between the results obtained on the weekday subsample and on the weekend day subsample could be explained in terms of time constraint. Households who belong to the weekday subsample are subject to a strong time constraint. Both male and female have worked at least seven hours, so they have little time left. They have however to spend a minimum of this time on vital activities. But the constraint they face is such that they may not be able to make time choices and to decide to substitute time from a given activity to another activity. The time they spend on each activity would be the minimum time required and thus could not be reduced. This may explain the weak separability of the derived utility function for a weekday.

In the weekend day subsample, households are not strongly time constrained. They can decide to spend more time on a given activity by reducing the time spent on other activities. This possibility of making time choices during a weekend day may explain the rejection of the weak separability assumption.

4. Conclusion

As far as we know, this paper is the first attempt to test for the weak separability hypothesis imposed by the household production model between inputs of different home activities. Using original data, we estimated a Working-Leser system of three conditional demand functions for meal preparation within the household. When households are not strongly time constrained, which is undoubtedly the most interesting case, our results show that inputs devoted to meal preparation are separable from restaurant expenditures, market good inputs devoted to other household activities, and male and female leisure time; but contrary to what is stated by the theory, they are nonseparable from male and especially female time inputs devoted to other household activities.

These results are interesting for several reasons. First, they show that food-out cannot be considered as a substitute for food-in. This could mean for example that food-out is constrained by professional activity and/or that this kind of expenditures belongs to another activity. Next, the weak separability observed between market good expenditures confirms that one can estimate a demand system for food products independently of other goods. Finally, these results show that activities that are weakly separable in goods may be nonseparable in time. This is what we observe here for meal preparation and other household activities, when households are not strongly time constrained.

Table 3. Estimation results based on the weekday subsample

Independent variables	(1)		(2)		(3)	
Instrumented logarithm of total expenditure ($\ln y$)	−12745.2	(−0.793)	−805.4	(−0.069)	13550.6	(0.923)
Food-out expenditures (x_1')	1.8	(0.881)	0.8	(0.553)	−2.6	(−1.402)
Male leisure time (l_m)	19.4	(0.658)	−25.5	(−1.191)	6.0	(0.223)
Female leisure time (l_f)	−31.7	(−0.872)	48.7	(1.856)	−17.1	(−0.515)
Other household activities						
Market goods expenditures (x_2)	0.1	(0.714)	0.0	(−0.412)	0.0	(−0.455)
Male time (t_{2m})	−64.1	(−1.748)	49.2	(1.853)	15.0	(0.447)
Female time (t_{2f})	33.0	(0.964)	−30.7	(−1.243)	−2.2	(−0.072)
Male age	335.8	(1.051)	42.8	(0.185)	−378.6	(−1.297)
Degree obtained by male						
Certificat d'Etudes Primaires	−4494.0	(−1.110)	4786.8	(1.635)	−292.8	(−0.079)
Certificat d'Aptitude Professionnelle	4145.9	(1.149)	−793.5	(−0.304)	−3352.4	(−1.017)
Brevet des Collèges	1369.5	(0.306)	−1052.0	(−0.325)	−317.5	(−0.078)
Brevet d'Enseignement Professionnel	1573.8	(0.241)	7442.7	(1.577)	−9016.5	(−1.513)
Baccalauréat	−590.4	(−0.105)	3096.5	(0.761)	−2506.1	(−0.488)
University degree	2386.5	(0.460)	4818.2	(1.285)	−7204.7	(−1.521)
Number of children						
0 to 1 year old	4232.1	(1.026)	324.7	(0.109)	−4556.8	(−1.209)
2 to 7 years old	1602.5	(0.785)	923.6	(0.626)	−2526.1	(−1.355)
8 to 14 years old	4429.2	(1.596)	−2554.1	(−1.272)	−1875.1	(−0.739)
15 to 24 years old	5676.7	(2.405)[b]	−1809.6	(−1.060)	−3867.2	(−1.793)
Homeowner	125.9	(0.053)	−2082.3	(−1.208)	1956.4	(0.898)
Constant	140150.2	(1.161)	24219.9	(0.277)	−64370.1	(−0.584)
Number of observations	385		385		385	

Notes: No degree is the reference for degree obtained by male; coefficients are multiplied by 10^5; asymptotic t-statistics in parentheses.

[a] Significant at the 1% level; [b] significant at the 5% level.

(1) Dependent variable is the share of market goods expenditures in total expenditure devoted to meal preparation (s_1); (2) dependent variable is the share of male time expenditures in total expenditure devoted to meal preparation (s_2); (3) dependent variable is the share of female time expenditures in total expenditure devoted to meal preparation (s_3).

Table 4. Estimation results based on the weekend day subsample

Independent variables	(1)		(2)		(3)	
Instrumented logarithm of total expenditure ($\ln y$)	−5237.8	(−0.347)	15309.5	(1.167)	−10071.7	(−0.674)
Food-out expenditures (x_1')	0.5	(0.371)	−0.7	(−0.603)	0.2	(0.155)
Male leisure time (l_m)	−18.3	(−1.443)	19.5	(1.774)	−1.3	(−0.101)
Female leisure time (l_f)	19.4	(1.214)	−9.7	(−0.700)	−9.7	(−0.611)
Other household activities						
Market goods expenditures (x_2)	0.0	(0.625)	0.0	(0.221)	0.0	(−0.824)
Male time (t_{2m})	−13.2	(−1.137)	32.8	(3.248)[a]	−19.6	(−1.702)
Female time (t_{2f})	−33.9	(−2.595)[a]	−34.6	(−3.046)[a]	68.4	(5.290)[a]
Male age	253.1	(0.908)	−179.6	(−0.741)	−73.5	(−0.266)
Degree obtained by male						
Certificat d'Etudes Primaires	−1181.7	(−0.331)	−2976.6	(−0.958)	4158.3	(1.174)
Certificat d'Aptitude Professionnelle	−3087.3	(−0.925)	−634.5	(−0.218)	3721.9	(1.124)
Brevet des Collèges	−5274.2	(−1.416)	4724.1	(1.458)	550.2	(0.149)
Brevet d'Enseignement Professionnel	−8833.9	(−1.464)	−4837.9	(−0.922)	13671.9	(2.286)[b]
Baccalauréat	−153.2	(−0.031)	−2939.7	(−0.690)	3092.9	(0.637)
University degree	−4271.8	(−0.862)	1226.9	(0.285)	3044.9	(0.620)
Number of children						
0 to 1 year old	9050.1	(2.916)[a]	702.4	(0.260)	−9752.5	(−3.170)[a]
2 to 7 years old	4762.2	(2.537)[b]	−2510.7	(−1.538)	−2251.5	(−1.210)
8 to 14 years old	4892.2	(2.018)[b]	−4067.8	(−1.929)	−824.3	(−0.343)
15 to 24 years old	3824.8	(1.701)	−2298.9	(−1.175)	−1525.9	(−0.684)
Homeowner	−2333.7	(−1.146)	−1029.0	(−0.581)	3362.8	(1.666)
Constant	83306.1	(0.735)	−97243.0	(−0.986)	113937.0	(1.014)
Number of observations	455		455		455	

Notes: No degree is the reference for degree obtained by male; coefficients are multiplied by 10^5; asymptotic t-statistics in parentheses.
[a] Significant at the 1% level; [b] significant at the 5% level.
(1) Dependent variable is the share of market goods expenditures in total expenditure devoted to meal preparation (s_1); (2) dependent variable is the share of male time expenditures in total expenditure devoted to meal preparation (s_2); (3) dependent variable is the share of female time expenditures in total expenditure devoted to meal preparation (s_3).

Endnotes

[1] Some references are Azzi and Ehrenberg (1975), Gronau (1977), Wales and Woodland (1977), Kooreman and Kapteyn (1987), and Biddle and Hamermesh (1990). Rosenzweig and Schultz (1983) use birth weight as an indicator of the output of the household health production function. But the model they estimate is still far from Becker's.

[2] The survey contains no information about these time uses for children (and other adults). However, to the extent that the value of time for parents is largely higher than the value of time for children, the contribution of the latter to the value of household production can be considered as negligeable.

[3] In the present work, we assume that these prices are identical through households. But it should be more realistic to consider that they differ with the structure of the household.

[4] However, to make sure of the robustness of the results, we also proceeded to the estimation of the system with the logarithm of wages (instrumented) included in the set of independent variables. The results of the separability test remained unchanged.

[5] We reason here in terms of opportunity cost by valuing an hour of male and female time at their market wage.

[6] For practical reasons that will become clear further on, only wage-earners are selected in our subsample.

[7] We have no information about female age and the degree she obtained. Note also that children are not considered here as household production, but only as control variables.

[8] For example, the average time spent at work and in transports during the time recording day is 60 min. higher for male and 41 min. lower for female when they belong to the sample of corner solutions than when they belong to the sample of interior solutions.

[9] We also added identifying instruments in the demand system. Estimation showed that none was significant.

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On private incentives to acquire household production skills

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Abstract. In non-cooperative family models, being good at contributing to family public goods like household production may reduce one's utility, since it tends to crowd out contributions from one's spouse. Similar effects also arise in cooperative models with non-cooperative threat point: improved contribution productivity entails loss of bargaining power. This strategic effect must be traded against the benefits of household production skills, in terms of increased consumption possibilities. Since cooperation involves extensive specialization, incentives to acquire household production skills are strikingly asymmetric, with the one not specializing in household production having strong disincentives for household skill acquisition.

JEL classification: D13, H41, J16, J22, J24

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1. Introduction

The nature of the sexual division of labor has undergone vast changes the last few decades. The traditional pattern of specialization, with a breadwinning father and a mother solely working at home, is fading in importance.

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However, also modern households practice a substantial degree of specialization, for instance with women choosing working arrangements that are compatible with having the main responsibility for children.[1] Also in families without small children we find a pattern of women taking more than half of the household work, while men typically spend more hours than women in paid work. The existence of comparative advantages is one explanation of this phenomenon. Gary Becker (1991) has pointed out that even small comparative advantages may lead to substantial specialization. Moreover, in order to maximize family output the family members should allocate their productive investments (their "education") according to the sector of specialization: the one specializing in household production should invest to improve his – or rather *her*, to frame the discussion using traditional gender stereotypes – household production skills, while the one specializing in market work should undertake investments that improve his labor market performance. This implies that even if "natural" differences between men and women may be rather small, these differences tend to increase as a consequence of investment decisions.[2] If there is learning-by-doing this would work in the same direction.

A problem with Becker's explanation of how comparative advantages may have evolved is that it cannot account for the fact that already at the date of marriage the family members have developed substantial comparative advantages, along traditional gender roles. It is well-known that women and men educate in different directions.[3] Perhaps even more striking are the differences between the sexes in what can be called domestic skills: an average woman about to be married is much better skilled to keep and maintain a house than is her coming husband, and this difference is found for a broad range of housekeeping activities: caring and nursing children, washing and mending clothes, shopping, house cleaning, cooking, baking, etc.[4] These are skills that are rarely acquired through formal educational, but rather passed on from parents to children or acquired by self studies.

Another problem with Becker's explanation is that what is in a family's joint interest is not necessarily in the interest of the individual family members.[5] In particular, the investments ("education") required to maximize family output is not necessarily serving the interests of the individuals who have to make those investments. Therefore, families and households should not be treated as single-person decision-makers, but rather as a collection of individuals with some degree of conflicting interests. Consequently, the fact that specialization maximizes family output is not a fully satisfactory explanation of the observed division of labor: it has to be verified that specialization is in the individual's interest. This applies in particular to decisions that are made before families are formed.

Often, skill acquisition, choice of education and many other decisions in life do not reflect rational decision-making but can rather be seen as responses to some social norms. This raises the question how these social norms have developed. The present paper also attempts to give an answer based on individual incentives – measured by private returns to improve one's household production skills.

Our point of departure is a non-cooperative model of family decisions, as laid out by Konrad and Lommerud (1995). On the topic of private provision of public goods, see Bergstrom et al. (1986). In their model, each of the two family members divide their time between market work and household work.

Household work produces a household good that is jointly consumed by both family members – i.e., it is a public good – while market work yields money income that is a private good for the individual worker. As is well-known from other private-provision-of-public-goods problems, there tends to be underprovision (relative to first-best) of the public good, due to the free rider problem. Moreover, the time spent in household production depends positively on one's productivity in the household production (hereafter called *contribution productivity*) and negatively on the spouse's contribution productivity. This implies that while there is an obvious positive direct effect of increased contribution productivity – the same amount of household goods can be produced with less effort – there is also a strategic effect working in the opposite direction. Thus, improved contribution productivity is not necessarily good for one's utility, even if the costs of improving one's household production skills are zero.

As a description of actual family life the non-cooperative model rests on quite pessimistic assumptions, and a natural way to proceed is to investigate the extent to which similar incentives can also arise in more cooperative families. Here I proceed along the lines of Konrad and Lommerud (2000), who assumes that i) education investments that determine individual wages are determined non-cooperatively; ii) time allocations are determined cooperatively; and iii) the outcome of the bargaining process can be described by the Nash bargaining solution, using equilibrium utilities in the non-cooperative model as fallback. While Konrad and Lommerud (2000) study incentives to improve one's wage, I study incentives to improve one's productivity in the household sector. A more important difference is that while Konrad and Lommerud focus on symmetric incentives to over- or underinvest in education, a key point of the present paper is that investment incentives may be strikingly asymmetric. There are also some technical differences that will be commented on in due course.

In short, also in the cooperative model there are direct and strategic effects, but there are important differences, too. First, in the cooperative model presented there will be full specialization, implying that one of the family members spends all time at home while her spouse devotes all his attention to market work.[6] This maximizes the joint surplus that will subsequently be divided according to the Nash bargaining solution. Since only the one with the highest contribution productivity works at home, there will be no direct effect of increasing her spouse's contribution productivity. In contrast, there is a strong direct effect of increasing her own contribution productivity. The strategic effect now works through the fallback payoffs, and it is clear that it will be negative for both parties: it can be shown that in the non-cooperative model it is always better to have one's spouse improve the contribution productivity than to improve one's own. (This holds for both family members.) Thus, the less productive will experience no direct effect and a negative strategic effect, so he will have disincentives to improve his domestic skills. Moreover, his spouse will experience a strong positive direct effect, and this effect will be only partially offset by the strategic effect, implying that she will still have incentives to improve her domestic skills.

The paper proceeds as follows. In the next section I present a cooperative family model with non-cooperation as the bargaining threat point. Then I derive the equilibrium payoffs to the family members for different combinations of contribution productivities. By performing comparative statics on the equilibrium payoffs, incentives to invest in household production skills are derived. Discussions are found in Sect. 3, while Sect. 4 concludes the paper.

2. Model

Consider the following simple family model with two persons, 1 and 2, which can be interpreted as a woman and her husband. Utility of person $i \in \{1, 2\}$ is given by a Cobb-Douglas function defined over one private and one public good;

$$u_i = x_i G, \quad i \in \{1, 2\}, \tag{1}$$

where x_i is individual consumption of a private good and G is the (common) consumption of the public good. The two goods are produced by devoting time to either of two tasks: paid work or (unpaid) household production. Let c_i denote the time person i spends on household production and the remaining time $1 - c_i$ in the labor market. The production function for the public good is given by

$$G = h_1 c_1 + h_2 c_2 \tag{2}$$

where the parameters h_1 and h_2 will be referred to as contribution productivities. In what follows we will assume that individual 1 is more productive in the household activity, that is, $h_1 > h_2$.

Individual income is given by

$$x_i = (1 - c_i) w_i, \quad i \in \{1, 2\}. \tag{3}$$

where w_i denotes a person's wage. Then individual utility can be written

$$u_i = (1 - c_i) w_i (h_1 c_1 + h_2 c_2), \quad i \in \{1, 2\} \tag{4}$$

Our description of the cooperative model contains three elements: i) A characterization of efficient allocations and the utility possibility set (i.e., the Pareto frontier); ii) A presentation of the threat point, which will be taken to be the utilities in the non-cooperative model (for a discussion of this approach, see Konrad and Lommerud 2000; and iii) A characterization of equilibrium allocations using the Nash bargaining solution. After the equilibrium is found, what remains is to perform comparative statics on equilibrium payoffs in order to assess the incentives to improve one's domestic skills.

The utility possibility set. By construction the cooperative outcome is efficient. Suppose that a transfer of size t goes from individual 1 to individual 2. Then we can write the sum of utilities as

$$u \equiv u_1 + u_2 = (x_1 - t)G + (x_2 + t)G = (x_1 + x_2)G \tag{5}$$

which is independent of t. Hence a necessary condition for an allocation to be efficient is that it maximizes xG, where $x \equiv x_1 + x_2$. In what follows we normalize both wages to 1. In the working paper version (Vagstad 1999) I provide the extension to cases in which wages may differ. Clearly, at least one of the family members will be fully specialized, implying that (since $h_1 > h_2$) either $c_1 = 1$ or $c_2 = 0$ or both.[7] Suppose that $c_1 = 1$. Then the optimal c_2 maximizes aggregate utility given by $u = (2 - c_1 - c_2)(h_1 c_1 + h_2 c_2) = (1 - c_2)(h_1 + h_2 c_2)$, and

$$\frac{\partial u}{\partial c_2} = h_2(1 - c_2) - (h_1 + h_2 c_2) = -(h_1 - h_2) - 2h_2 c_2 < 0 \tag{6}$$

yielding $c_2 = 0$. Conversely, suppose that $c_2 = 0$. Then the optimal c_1 maximizes $u = (2 - c_1 - c_2)(h_1 c_1 + h_2 c_2) = (2 - c_1)h_1 c_1$. Then

$$\frac{\partial u}{\partial c_1} = (2 - c_1)h_1 - h_1 c_1 = 2h_1(1 - c_1) > 0 \quad \text{for } c_1 < 1) \tag{7}$$

implying that c_1 should be set equal to 1. In conclusion, the unique efficient allocation of effort entails full specialization. The resulting aggregate utility to be shared is then given by $u = h_1$. The solution above implies that the family members' bargaining problem is reduced to bargaining over how to divide an aggregate amount $x = 1$ of the private good, both agreeing on a joint consumption of $G = h_1$ of the public good. Consequently, the utility possibility set is given by $\{(u_1, u_2) | u_1 + u_2 \leq h_1\}$.

Fallback utilities. We take non-cooperation as the alternative to reaching an agreement in the cooperative model. Therefore, the fallback utilities will be taken to be the payoffs in a non-cooperative model, in which the familiy members simultaneously and non-cooperatively decide how to allocate their time. An interior Nash equilibrium in pure strategies exists if the two contribution productivities do not differ too much. We therefore make the following assumption on productivity differences: $h_1 < 2h_2$.

Now the (interior) non-cooperative equilibrium is found by first solving the set of first order conditions for c_1 and c_2 and then substituting the solution back into the variable definitions above. The first order conditions ($\frac{\partial u_1}{\partial c_1} = (h_1 - 2h_1 c_1 - h_2 c_2)w_1 = 0$ and $\frac{\partial u_2}{\partial c_2} = (h_2 - h_1 c_1 - 2h_2 c_2)w_2 = 0$) yield $c_i = c_i^* \equiv \frac{2h_i - h_j}{3h_i}$, $G = G^* \equiv \frac{1}{3}(h_1 + h_2)$ and $x_i = x_i^* \equiv \frac{w_i}{3h_i}(h_1 + h_2)$ $(i, j \in \{1, 2\} \wedge i \neq j)$. Note that $\frac{\partial c_i^*}{\partial h_j} < 0$, telling us that improved productivity leads to a reduction in one's spouse's contribution to the public good. What we are most interested in is equilibrium utilities, since these will serve as the threat points in the bargaining game:

$$u_i = u_i^* \equiv \frac{w_i}{9h_i}(h_1 + h_2)^2, \quad i \in \{1, 2\}. \tag{8}$$

For use in the subsequent discussion, we will take a closer look at investment incentives in a purely non-cooperative model before we proceed. Differentiating individual utility with respect to the contribution productivities yields

$$\frac{\partial u_i^*}{\partial h_j} = \frac{2w_i}{9h_i}(h_1 + h_2) > 0, \quad \text{and} \tag{9}$$

$$\frac{\partial u_i^*}{\partial h_i} = \frac{w_i}{9}\left[1 - \left(\frac{h_j}{h_i}\right)^2\right] > 0 \quad \text{iff } h_i > h_j. \tag{10}$$

Both these partials are interesting in the sense that they say something about private incentives to influence the contribution productivities in a non-cooperative family. The first one, (9), tells us that it is in a family member's interest that the other family member improves his or her contribution productivity. The second one, (10), reveals that it is unclear whether it is an advantage or not to be productive in the household sector:

Proposition 1. *If time allocation is decided non-cooperatively, the individual having the highest contribution productivity would prefer to be even more productive, while the individual having the lowest contribution productivity would prefer to have even lower productivity.*

Note that Proposition 1 depends on our particular model formulation. Most important is the degree of substitution implicitly assumed by our choice of utility function. Clearly, as the strategic effect works mainly through substitution between the two goods, any utility function involving less substitutability will entail weaker strategic effects with the possibility of both individuals having incentives to improve their domestic skills,[8] while utility functions involve more substitutability may result in both partners having negative investment incentives.

Nash bargaining. It is somewhat dissatisfactory to base one's results on an assumption of families being unable to reach efficient decisions. From now on we will work with the opposite assumption: that decisions in the family are efficient. However, we maintain the assumption that decisions taken before families are formed are made non-cooperatively (and that such decisions may therefore not be efficient). To be more precise, suppose that the family members decide cooperatively how to allocate their time (i.e., effort), but that productivities are decided before the families are formed. To avoid the complications arising if the choice of productivity investments affects who marries whom, we will leave the discussion of possible marriage market effects until Sect. 3. That is, in this section we assume that men and women match randomly to form families.

Let Δ denote the gains from bargaining, defined as follows:

$$\Delta = u_1 + u_2 - (u_1^* + u_2^*) = h_1 - \left(\frac{1}{9h_1}(h_1 + h_2)^2 + \frac{1}{9h_2}(h_1 + h_2)^2\right) \tag{11}$$

In the Nash bargaining solution this gain will be split evenly, implying that individual utilities in the bargaining equilibrium equal

$$u_1 = \frac{1}{9h_1}(h_1 + h_2)^2 + \frac{\Delta}{2} = \frac{1}{2}h_1 - \frac{1}{18}\frac{h_1 - h_2}{h_1 h_2}(h_1 + h_2)^2 \tag{12}$$

$$u_2 = \frac{1}{9h_2}(h_1 + h_2)^2 + \frac{\Delta}{2} = \frac{1}{2}h_1 + \frac{1}{18}\frac{h_1 - h_2}{h_1 h_2}(h_1 + h_2)^2 \tag{13}$$

Incentives to improve contribution productivities. Differentiating these expressions for individual equilibrium payoffs yields

$$\frac{\partial u_1}{\partial h_1} = \frac{1}{18}\frac{8h_1^2h_2 - 2h_1^3 - h_2^3}{h_1^2h_2} > 0 \tag{14}$$

$$\frac{\partial u_2}{\partial h_1} = \frac{1}{18}\frac{2h_1^3 + 10h_1^2h_2 + h_2^3}{h_1^2h_2} > 0 \tag{15}$$

$$\frac{\partial u_1}{\partial h_2} = \frac{1}{18}\frac{h_1h_2^2 + 2h_2^3 + h_1^3}{h_1h_2^2} > 0 \tag{16}$$

$$\frac{\partial u_2}{\partial h_2} = -\frac{1}{18}\frac{h_1h_2^2 + 2h_2^3 + h_1^3}{h_1h_2^2} < 0 \tag{17}$$

Again both cross partials are positive, but for slightly different reasons than before. When the more productive family member improves – h_1 increases – his spouse gains twofold. First, the set of feasible allocations expands. Second, the spouse's relative bargaining position improves.[9] When it is the less productive who improves, the first effect is absent: due to complete specialization, it is only the most productive's productivity parameter that matters for the utility possibility set. However, the second effect – on bargaining position – is still there. In sum, also in the cooperative model an individual benefits from having a spouse who has a high contribution productivity.

More interesting is the own derivatives, as they determine the incentives to acquire household production skills. We see that – qualitatively speaking – the results from the non-cooperative model are valid also in the cooperative model:

Proposition 2. *Also in the cooperative model, the most productive family member has incentives to improve, while the less productive has disincentives to do so.*

The intuitions behind Propositions 1 and 2 are slightly different, however. Consider a change in the most productive person's contribution productivity. Such a change affects the utility possibility set: when h_1 increases, the Pareto frontier shifts outward. (Clearly, this has some of the same flavor as the direct effect in the non-cooperative model, but it should be noted that they are not the same.) Moreover, increasing h_1 reduces person 1's bargaining power, in a way that resembles the strategic effect in the non-cooperative model. But again they are different. One important difference becomes evident when the contribution productivities are almost equal. In the non-cooperative model we then know that incentives in either direction are weak ($\frac{\partial u_i^*}{\partial h_i} = \frac{1}{9}\left[1 - \left(\frac{h_j}{h_i}\right)^2\right] \approx 0$ when $h_i \approx h_j$). In contrast, in the cooperative model $\frac{\partial u_1}{\partial h_1} = \frac{1}{18}\frac{8h_1^2h_2 - 2h_1^3 - h_2^3}{h_1^2h_2} \approx \frac{5}{18}$ and $\frac{\partial u_2}{\partial h_2} = -\frac{1}{18}\frac{h_1h_2^2 + 2h_2^3 + h_1^3}{h_1h_2^2} \approx -\frac{4}{18}$ when $h_1 \approx h_2$, implying that even the smallest difference in contribution productivities produces distinct incentives to differentiate further.

Second, consider a change in the less productive's contribution productivity. Such a change does not affect the utility possibility set, implying that in the cooperative model there are no positive effects on u_2 from increasing h_2. The change affects the bargaining positions, however, and this effect is strong and clear even at the smallest possible differences.

3. Discussion

We have seen that wage differences do not affect contributions in the non-cooperative model. In the cooperative model things are more complicated. In the working paper version (Vagstad 1999) I allow for differences in labor market skills as well as household production skills. In what follows I briefly sketch the method and the main result, and refer the interested reader to Vagstad (1999) for details.

First, efficient allocations maximize aggregate utility given by $u = (x_1+x_2)G = ((1-c_1)w_1 + (1-c_2)w_2)(h_1c_1 + h_2c_2)$. Depending on absolute as well as comparative advantages, efficiency may entail full specialization (i.e., one family member works full time in the market while the other works full time in household production) or only partial specialization (i.e., only one member is fully specialized while the other split his or her attention between household production and market work).

After finding the efficient allocations, we employ the Nash bargaining solution to find how the gains from cooperation is split. This yields equilibrium payoff of each family member as explicit functions of all the parameters of the problem, that is, w_1, w_2, h_1 and h_2. Finally, investment incentives are found by differentiation of equilibrium payoffs. One robust result stands out from the crowd of partial derivatives (proof in Vagstad 1999):

Proposition 3. *The one with a comparative advantage for market work gets more than 50% of total output. Moreover, the same person would prefer to be even less domestically skilled, and would also prefer to have a spouse with even lower wage.*

Many matters of interest have been left out of the above analysis. In the cooperative model investigated above I have implicitly assumed that it is clear who will have the higher contribution productivity already at the time when the investments in household production skills are made. This is a strong assumption, if we are talking about the way children are brought up. The assumption may nevertheless be reasonable in a particular sense: even if a couple does not know the identity of their daughter's particular future husband, they may rationally believe that he will be brought up as a typical boy, that is, not being taught how to run a household. Conversely, parents of boys may rationally believe that his prospective wife will be brought up to have a rather high contribution productivity. Consequently, it appears that rational expectations may turn into self-fulfilling expectations.

This is not at all clear, however. Even if the average suitor has little household production skills, it is possible for a girl (or her parents) to undercut that level by appropriate (lack of) skill acquisition. How come parents do not engage in such undercutting? One possible explanation could be that it may be undesirable to have two persons without household production skills form a family – such families suffer in terms of a small utility possibility set. In this sense specialized upbringing can be seen as a coordination device – it assures that at least one of the family members is skilled in household production.[10]

Another explanation can be sought in marriage market forces. Suppose that instead of random matching, matching is based on utility-maximizing behavior – behavior that can be turned into forces of supply and demand. What follows is a brief discussion of the demand for partners with certain productivity parameters. I also discuss the demand for what can be seen as the alternative to domestic skills: labor market skills (measured by wage levels). Underlying the discussion is an assumption of highly incomplete marriage contracts. In particular, a marriage contract is taken to be only a commitment to play a particular game (either the non-cooperative or the cooperative game) with a specific partner – there are no up-front payments (i.e., no dowries or bride prices).

First, in the cooperative model as well as the non-cooperative model there will be a positive demand for spouses with high contribution productivities, suggesting that the negative incentives to acquire household production skills may be mitigated by forces in the marriage market: having high skills for household production may be bad for you once you are married, but it makes you more attractive as a marriage partner.

Perhaps more interesting is the demand for labor market characteristics: in the non-cooperative model, wages are irrelevant as a sorting criterion – they are neutral as regards contribution to the public good. This is not so in the cooperative model. The most plausible assumption is perhaps that women have an absolute advantage in household production while men have an absolute advantage in the labor market. If so, the strategic effect then dominates in two cases: the husband's utility is decreasing in his wife's wage as well as in his own contribution productivity, while all other derivatives (own and cross) are positive (see Vagstad 1999, for details). The intuition for this pattern is that since the wife will spend all her time in household production and the husband will spend all his time in market work, increasing the wife's wage marginally will not affect the utility possibility set at all, it will only make her demand a larger share of total output.

Finally, the demand for high-wage husbands is somewhat unclear a priori. The effect on the utility possibility set suggests that there will be a demand for high-wage men, but since high-wage men (just like high-wage women) will demand a larger share of total output, it is a priori unclear whether there will be demand for high-wage men. In this model, however, the positive effect always dominates, creating a positive demand for high-wage men.

To sum up the marriage market effects, the above discussion suggests that women improve their market positions by investing in household production skills and disinvest in improving their labor market performance. In contrast, for men both attributes are valuable in the marriage market. There will be a certain demand for men with high contribution productivities, because such men will demand a smaller fraction of output. But there will also be demand for high-wage men, in spite of such men being more demanding, because they expand the utility possibility set.[11] Consequently, optimizing women will concentrate on household production skills, while optimizing men will try to improve along two dimensions, which is to say that women have stronger incentives to become domestically skilled.

4. Concluding remarks

If family members have common preferences, Becker (1991) – among others – has argued that family members will specialize according to comparative advantages, not only regarding division of labor, but also when it comes to investments that improve one's skills. Therefore, investments tend to increase any productivity differences. We have seen that even if families are less harmonic, Becker is basically right on this point: individual investment incentives tend to increase specialization also when we allow for explicit conflicts of interest between family members.

Becker's second hypothesis concerns efficiency: with common preferences, investments as well as time will be allocated in a way that maximize joint surplus. This result does not hold when we relax the assumptions of common preferences. To be more specific: with truly non-cooperative families, neither investments nor time allocations will be efficient, for well-known reasons: there are strong positive externalities from contributing to household production as well as from investing in domestic skills, giving rise to a serious free-rider problem. In what I have called a cooperative family, time allocations will be efficient (by assumption), but this may give rise to even less efficient investments, in order to affect the outcome of the bargaining process.

I conclude with a short list of issues for further research. The first concerns policy issues. My analysis points at two problems that I believe deserve attention. First, from an efficiency point of view one should be concerned about the welfare effects of underinvestment in household production skills. Second, from a distributional point of view one should be concerned that investment incentives increase the gender differences, not only in time and investment allocations (which is hardly a problem in itself) but, more seriously, in equilibrium payoffs.

The fact that a decrease in one person's contribution to the public good increases the other person's contribution is what gives rise to underinvestment. Consequently, anything that can break the substitutability between the family members' contributions should in principle reduce the underinvestment problem. The government may affect this substitutability in many ways. First, the they may promote the provision of close substitutes to the household public good (publicly provided day care for children is only one example), implying that if one family member contributes less than his or her "just" share, he or she may be told to buy the rest in the market. Second, also taxation may affect specialization. In particular, progressive taxation based on individual labor income will discourage specialization and therefore reduce substitution. Third, implementing a cap on how many hours a week one is allowed to work will also limit specialization. It is more difficult to imagine explicit regulation of division of labor within the family, but social norms may have a similar effect: if the public goods to be provided can be divided in "male" and "female" tasks that are socially sanctioned, this should have a beneficial effect on investment incentives – although the benefits must of course be traded against the possible inefficiencies such rigidities may entail.

Finally, the marriage market forces discussed in the previous section warrants more analysis. In particular, important features of modern societies like the increase in divorce rates and the number of single-person households should be incorporated in future work.

Endnotes

[1] For some US evidence, see, e.g., Fuchs (1989) and Hersch (1991). Relevant facts from Britain and Germany are found in Joshi (1989) and Beblo (1998), respectively.

[2] From nature's side men and women's comparative advantages in market work and household production appear to be of limited importance today. Women's natural advantage in taking care of children should apply only for small babies, and physical strength is of little or no importance in most jobs today.

[3] Looking at education we see – roughly speaking – that women choose educations that are compatible with housekeeping: nurses and kindergarten teachers are prime examples. In contrast, men educate themselves to traditional bread-winning positions in the labor market: industrial workers, engineers and managers.

[4] There are of course exceptions to this general pattern, however: different types of light maintainance work – e.g., fixing the car when it is broken – are typically done by men. What should be noted, however, is that the activities in which the women are more skilled comprise the bulk of household production activities.

[5] This was first formally analyzed by Manser and Brown (1980) and McElroy and Horney (1981). More recent contributions include Lundberg and Pollak (1993) and Konrad and Lommerud (1995, 2000). For surveys, see Ott (1992) or Lommerud (1997).

[6] Clearly, this result depends upon the particular parameterization of the model. More generally one will expect cooperation to promote specialization, with similar qualitative effects.

[7] If none were fully specialized, then output could be increased with the same total effort by shifting attention in the direction of comparative advantages.

[8] See e.g. Konrad and Lommerud (2000), whose model features quasi-linear utility.

[9] For this to be the case, u_2^* must increase more than u_1^* when h_1 increases. Simple differentiaton reveals that $\frac{\partial u_2^*}{\partial h_1} - \frac{\partial u_1^*}{\partial h_1} = \frac{1}{9}(h_1 + h_2)\frac{2h_1^2 - h_2 h_1 + h_2^2}{h_1^2 h_2} > 0$ whether $h_1 > h_2$ or not. This implies that one's bargaining position always deteriorates with increasing contribution productivity.

[10] A full-fledged analysis of the interplay between beliefs and investment decisions is beyond the scope of this paper. See Francois (1998) and, in particular, Lommerud and Vagstad (2000) for a discussion of the considerations involved in such an analysis.

[11] Some limited support for these conjectures is found in data for firms doing professional marriage partner search in Germany (see Nitschke 1998). Such firms typically price discriminate by charging lower prices if you are attractive yourself, and by looking at prices, Nitschke finds that what makes a man attractive is a high ability to provide money, while what makes a woman attractive is being young and pretty. He also quotes figures telling that it is particularly difficult to find partners to older women with higher education (i.e., high potential wages).

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Estimating intrahousehold allocation in a collective model with household production

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Abstract. The purpose of this paper is to estimate the intra-family distribution of income and the individual demand for leisure and household production from Swedish cross-sectional household data. As a basis for the analysis, we use a collective model where each individual is characterized by his or her own utility function and divides total time between leisure, household production and market work. For the purpose of comparison, we also estimate a version that is consistent with a more traditional model of labor supply, the unitary model.

JEL classification: D13, J22

Key words: Time-use, household production, collective model

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1. Introduction

Traditionally, the household has been considered as a single utility maximizing agent. This so called unitary model has lately been criticized both from a theoretical and an empirical viewpoint. It has been argued that a multiperson household cannot be modeled as a single individual because it contradicts the neoclassical starting point that every individual should be characterized by his/her own preferences. Moreover, the unitary model only considers allocations between households and disregards questions concerning intra-household inequalities, which may lead to wrong welfare implications (see Haddad and

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Kanbur 1990, 1992). One testable restriction imposed by the unitary model is that the distribution of nonlabor income across spouses is not important for behavior: only the sum of the spouses' nonlabor income matters. This so called pooling restriction has also been rejected in several empirical studies, e.g., in Schultz (1990); Thomas (1990) and Kawaguchi (1994).

The theoretical and empirical criticism against the unitary model has lead to other models of household behavior, where individual preferences are recognized. One model that has received widespread attention is the collective model, developed by Chiappori (1988, 1992). In the collective model, the household is assumed to reach a Pareto-efficient outcome. Within this framework, testable restrictions can be derived and the intra-household distribution of income can be estimated. As pointed out by Becker (1965) a significant amount of time not used for market work is used for household production. Apps and Rees (1997) extend Chiappori's model by introducing a good produced by the households. They show that the sharing arrangement underlying the collective model cannot be retrieved unless specific assumptions are made. Chiappori (1997) shows that when the household good is tradable at a given price, the sharing arrangement may still be retrieved up to an additive constant.

There have been few empirical studies based on the collective model. Among the exceptions are Browning et al. (1994); Fortin and Lacroix (1997) and Chiappori et al. (1998). The results from their studies are generally supportive of the collective model. However, none of these studies have considered household production. The purpose of our study is to use the extended collective model to estimate simultaneously the intra-family distribution of income, individual leisure demands and the household production function from Swedish cross-sectional household data. Two versions of the model are estimated, one that is consistent with the household good being tradable on the market, and the other where households cannot buy or sell the household good. For purposes of comparison, we also estimate a version that is consistent with the unitary model.

The paper is organized as follows. In the next section, we present the theoretical background of our study based on Chiappori's (1997) work. Section 3 consists of a description of the data used in the empirical study. In Sect. 4, we present the empirical model and the estimation results. Section 5 concludes the paper.

2. Theory

We consider a two-member household where m denotes the male and f denotes the female. Assume that individual i's direct utility function can be written

$$u^i = u^i(l_i, c_i, x_i; \mathbf{z}_i), \quad i = m, f.$$

Utility is defined over three different goods: leisure, l_i, a market consumption good, c_i, and a consumption good, x_i, that is produced within the household. The vector of demographic characteristics of the individual is denoted as $\mathbf{z}_i$. Let us for the time being assume that x_i is marketable, i.e., it can be sold and bought on the market. The household production function, $h(t_m, t_f; \mathbf{a})$, is assumed to be characterized by constant returns to scale[1], where t_m and t_f are hours of household work for the male and the female respectively, and

$\mathbf{a} = (\mathbf{a}_m, \mathbf{a}_f)$ denote characteristics of the household members that are of importance for household production.

Suppose that the household decision process can be interpreted as a two-stage procedure where the household members first determine a production plan and how the resources are going to be shared within the household. Following the collective approach, we assume that the household decision process leads to a (within household) Pareto-efficient outcome. The domestic production plan can formally be written as a profit maximization problem

$$\max_{t_m, t_f} \pi = ph(t_m, t_f; \mathbf{a}) - w_m t_m - w_f t_f, \tag{1}$$

where w_m and w_f denote marginal (after tax) wage rates, and p is the price of the domestically produced good. The first order conditions for t_m and t_f can be combined to read

$$\frac{\partial h(\cdot)/\partial t_m}{\partial h(\cdot)/\partial t_f} = \frac{w_m}{w_f}.$$

Given the production plan and the sharing arrangement, the second stage of the decision problem will be

$$\max_{l_i, c_i, x_i} u^i(l_i, c_i, x_i; \mathbf{z}_i), \quad i = m, f, \tag{2}$$

$$\text{s t} \quad c_i + px_i + w_i l_i = s_i,$$

where s_i denotes member i's part of the household's full income.

We assume that the tax system is piecewise linear, so the marginal wages are defined conditional on the segment of the tax schedule where individual's labor supply is observed. In this case, the full income of the household can formally be written

$$s_F = s_m + s_f = (w_m + w_f)H + \tilde{y}_m + \tilde{y}_f,$$

where H are the maximum hours available, $\tilde{y}_m$ and $\tilde{y}_f$ are, respectively, the male's and the female's virtual income components. The virtual income is defined as the intercept income resulting from linearizing the individual's budget constraint around the tax segment where the observed hours of work are located.[2] Each member's share of the household's full income, s_i, can in general be seen as a reduced form function describing the determinants of the sharing arrangement made in the first stage of the decision process. We choose to write family member i:s part of the full income as $s_i = s_i(w_m, w_f, y_m, y_f, \mathbf{z}, \mathbf{EEP})$, where y_m and y_f denote nonlabor incomes, $\mathbf{z} = (\mathbf{z}_m, \mathbf{z}_f)$ the personal characteristics and **EEP** is a vector of so called extra-environmental parameters (EEP's) describing the opportunity cost of household membership.[3] The consumption of leisure determined by (2) can be written

$$l_i = l_i[w_i, p, s_i(w_m, w_f, y_m, y_f, \mathbf{z}, \mathbf{EEP}); \mathbf{z}_i], \quad i = m, f. \tag{3}$$

Within this framework it is possible to identify the intra-family distribution of income up to an additive constant (see Chiappori 1997).

If we assume that the household good, x_i, cannot be sold or bought on the market, the analysis becomes more complicated from the point of view of identification. In the maximization problem (2), p now depends on the marginal wages of the household members as well as household production characteristics, $\mathbf{a}$, and can be interpreted as the shadow price of the market consumption good (see Apps and Rees 1997). The budget constraints now read $c_i + p(w_m, w_f, \mathbf{a})x_i + w_i l_i = s_i$, $i = m, f$. The demand for leisure can be written

$$l_i = l_i[w_i, p(w_m, w_f, \mathbf{a}), s_i(w_m, w_f, y_m, y_f, \mathbf{z}, \mathbf{a}, \mathbf{EEP}); \mathbf{z}_i], \quad i = m, f. \tag{4}$$

where we have indicated that the sharing rule depends on all characteristics of the household. This implies that it is not possible to retrieve all the partial derivatives of the sharing rule. However, nonlabor incomes and the EEP's affect leisure demands only through their effect on the sharing rule. This means that we can at least obtain partial information of the sharing rule by estimating male and female leisure demands on the form of (4).[4]

3. Data

The data used in this study are based on the 1984 and 1993 Swedish Survey of Household Market and Nonmarket Activities (HUS)[5]. The 1984 (1993) HUS-Survey consists of 2619 (4137) randomly selected individuals aged 18 to 74. Besides the conventional survey, a selection of the respondents were subject to a time-use study. The interviews were performed using the yesterday 24 hour recall diary technique (see Juster and Stafford 1991), and each respondent was interviewed on at most two occasions. The sample size for the first and second time-use interview in 1984 (1993) was 2552 (3249) and 2468 (3218) individuals respectively.

One important characteristic of HUS is that both partners have been interviewed. This is a necessary condition for our empirical study, as we want to estimate the distribution of resources among the household members. Our main sample refers to two-adult households with and without children where both spouses are between 20 and 60 years of age. Including families with children may not be entirely unproblematic, however. Children may be seen as public goods within households. If these goods are nonseparable in the utility functions of their parents, the collective model portrayed above may not be valid. We will, therefore, compare the estimates based on our main sample with those obtained from restricting the sample by excluding families with children.

In the empirical analysis we include households where both adult members have participated in the main survey and at least in one time-use interview.[6] The number of hours used for household production is calculated from the time use data, and the sample is restricted to households were both members have stated a positive amount of household work. Only information on primary activities are used and household work is defined as the sum of: (i) traditional housework, i.e., food and drink preparation, dishwashing, cleaning-up, washing, ironing, clothes care and household management; (ii) active child care; (iii) purchase of everyday goods and clothing together with associated travel; and (iv) maintenance, repairs and improvement on one's home including yard work.

Information on hours worked on the market is collected from the conventional survey data and we only consider households where both partners have chosen to participate in the labor market. Households were at least one of the members has been on sick leave for more than three weeks during the year, or has provided inconsistent tax-return values are excluded. We also exclude households were individual wages are reported missing. Non-labor income is defined as the sum of interest incomes, interest subsidies, dividends and capital gains less capital losses, interest on debts and administrative expenses. To obtain a measure of non-labor income that is consistent with this definition, farmers and owners of more than one property (aside from vacation home) are excluded from the 1984 data.

Following Chiappori et al. (1998) we use measures of the sex ratio, i.e., the relative supplies of males and females in the marriage market, as EEP's describing the state of the marriage market. Using data from Statistics Sweden, we calculate sex ratios on the basis of age group, county (län) of residence, and sex. We have experimented with several different measures of the sex ratio. The final measure that is used in the empirical analysis is the female sex ratio defined as the number of females in an age group over the 'efficient' number of males supplied to that age group. We assume that females match with men that are 0–3 years older than themselves. The efficient number of males, however, is reduced by the fact that they can match with women other than those from the relevant age group.[7] In addition, a number of individual and household characteristics are used in the empirical study. These are described in the empirical section below. In total, the 1984 and 1993 main sample contains 326 and 338 households, respectively. Descriptive statistics are presented in Table 1.

4. Empirical model and estimation results

4.1. Empirical model

In this section we will estimate a collective model including household production. From the discussion in Sect. 2, it is clear that the assumption of constant returns to scale in home production is convenient from the viewpoint of identification, and this is also how we proceed. Household production is assumed to be of the constant elasticity variety. Specifically, we assume that the household production function take the CES-form;

$$f(t_m, t_f; a_m, a_f) = (a_m t_m^{-\rho} + a_f t_f^{-\rho})^{-1/\rho}$$

where a_m and a_f are productivity parameters that may be made dependent upon personal characteristics of each spouse. We will describe their content below. The first order conditions for t_m and t_f can be combined to read

$$\ln(t_m/t_f) = \lambda[\ln(w_m/w_f) - \ln a_m + \ln a_f] \tag{5}$$

where $\lambda = -\dfrac{1}{1+\rho}$.

Leisure demand functions emanate from Eqs. (3) and (4). In the latter case, the demand functions depend on the endogenously determined price of the household good. Since our main interest is to estimate the intrahousehold distribution of resources, and since there is no identification to be gained by

Table 1. Sample statistics

Variables	1984				1993			
	Men		Women		Men		Women	
	Mean	St.dev.	Mean	St.dev.	Mean	St.dev.	Mean	St.dev.
Age	40.84	(8.90)	38.30	(8.80)	43.77	(9.12)	41.55	(9.21)
Years of education	11.74	(3.63)	11.28	(3.11)	12.61	(3.60)	12.55	(3.32)
Gross wage rate	53.06	(18.74)	41.30	(10.96)	106.49	(78.20)	86.55	(52.99)
Marginal wage rate	21.85	(6.53)	22.89	(6.35)	57.79	(40.14)	57.78	(48.21)
Marginal tax rate	56.20	(13.24)	43.34	(12.27)	43.99	(11.02)	32.53	(36.83)
Nonlabor income	−817.47	(5572.8)	509.11	(3490.68)	−14756	(31017)	−3407.03	(17866)
Market work	2193.18	(320.46)	1651.83	(543.24)	2207.29	(329.76)	1787.63	(460.34)
Household work	811.31	(600.20)	1456.26	(800.43)	851.49	(612.24)	1336.55	(756.37)
Leisure	5731.51	(664.43)	5627.92	(830.08)	5677.22	(700.14)	5611.22	(810.10)
Sexratio	1.02	(0.05)	0.98	(0.05)	1.02	(0.06)	0.98	(0.06)
H-hold full income	433077	(80232)			1055946	(531660)		
Children/household	1.32	(1.00)			1.04	(1.08)		

explicitly modelling the price of the household good, we choose not to model the price function explicitly. Instead, we specify leisure demand functions on 'semi-reduced' forms, i.e. they depend on cross wages as well as spouse characteristics. We shall also assume that the leisure demand functions are linear;

$$l_m = \beta_m w_m + \delta_m w_f + \gamma_m s_m + \tilde{z}_m + \tilde{z}_f \tag{6}$$

$$l_f = \beta_f w_f + \delta_f w_m + \gamma_f s_f + \hat{z}_m + \hat{z}_f \tag{7}$$

where s_m and s_f denote, as before, the male and female share of the households full income. The scalars $\tilde{z}_m$, $\tilde{z}_f$, $\hat{z}_m$, and $\hat{z}_f$, should now be interpreted to contain characteristics originating from the utility function as well as from the production function, and β_i, δ_i, γ_i, $i = m, f$, are parameters to be estimated. Characteristics are assumed to include age, a dummy variable indicating the presence of children in a specific age bracket (0–6, 7–12 and 13–17 years of age) and a dummy variable reflecting the educational attainment of the individual. The educational dummy variable takes the value one if the respondent has a university or a university college degree. Equations (5), (6), and (7) correspond to our most general model specification, which is the case where the household good cannot be traded. We will below discuss how to impose restrictions in order to make the empirical model compatible with alternative theoretical interpretations.

Denote by θ_m the male's relative share of full income such that $s_m = \theta_m s_F$. Consequently, $\theta_f = 1 - \theta_m$ is the female's relative share. We assume that these relative shares are determined by marginal wages and nonlabor incomes, and, in addition, a number of other exogenous variables (to be described below) reflecting the income sharing arrangement. Following Browning et al. (1994) we assume that the relative shares can be modelled by a logistic function, i.e., $\theta_m = 1/(1 + \exp(d))$, and $\theta_f = 1 - 1/(1 + \exp(d))$, where d contains the variables affecting the income sharing arrangement.

In specifying the determinants of the sharing rule, the reader should first observe that when the household good is tradable, it can be seen from Eq. (3) that $\partial l_i/\partial \mathbf{z}_j = \partial l_i/\partial s_i \cdot \partial s_i/\partial \mathbf{z}_j$, meaning that the characteristics of the partner only affect the leisure demand via the sharing rule. Therefore, spouse characteristics can be used to obtain identification of the parameters of the sharing rule in this case. This could be done for instance by including differences in wage rates and other characteristics between the household members (see also Browning et al. 1994, for a similar argument). If, on the other hand, the household good cannot be traded, identification of the sharing rule originates from the nonlabor income of the spouses and the EEP's, since spouse characteristics will in this case affect l_i both via the sharing arrangement and via the price of the household good. We assume that the sharing rule is determined by the female sex ratio and by the differences in age, the number of years of schooling, marginal wages, and nonlabor incomes between the household members. Finally, since the sharing rule cannot be fully recovered in either of the two models set out in Sect. 2, the mean of the sharing rule is centered around one half.

In an economy without nonlinear taxation the wage rate is exogenous to hours of work. However, under progressive income taxation the marginal wage rate is endogenous. To address this problem, estimation is accomplished using an instrumental variables method. The instruments chosen for the marginal wage rate are the gross wage rate, the square of the gross wage rate, capital income and capital income squared. Similarly, full income of the household is also endogenous under nonlinear taxation and is instrumented as well. In this case, capital income, the gross wage rate for both spouses and nontaxable benefits are used as instruments. The final Eqs. (5), (6), and (7) are then estimated jointly using the nonlinear least-squares estimator, where cross equation restrictions pertaining to the sharing rule parameters are invoked and the stochastic specification allows for contemporaneous correlations of the error terms.

4.2. *Estimation results*

The estimation results are presented in Tables 2 and 3. The tables contain three different versions of the empirical model. Model I refers to the full model given by Eqs. (5), (6), and (7). Model II is a restricted version of model I, where cross-wage effects are set to zero in the leisure demand equations, and the spouse's characteristics affect leisure demands solely via the sharing rule. Model III, finally, is based on the assumption that the sharing rule is constant by imposing the restrictions that $\theta_f = \theta_m = 1/2$. With reference to the theoretical section, we can interpret these models as consistent with the non-marketable household good case (model I), the marketable household good case (model II), and the unitary model (model III). The restrictions imposed by models II and III can easily be tested against the more general alternative (model I).

We start by comparing the estimates of model I with those of models II and III by performing Wald tests. Comparing model I and model III, the latter is obtained as a special case of the former by setting the five parameters in the sharing rule to zero. The critical value for rejecting the null at the conventional 95% level of significance is $\chi^2(5) = 11.07$. The Wald-test statistic is 3.96 and 46.7 for the 1984 and 1993 data respectively. A similar test of model II yields the Wald-test statistics 6.59 and 17.59 (the critical value is in this case given by 12.59). Hence, we are not able to reject model II and model III using the 1984 data, while using the 1993 data both models II and III can be rejected. We have also tested the null hypothesis of constant parameters over time. This hypothesis was clearly rejected.

The reader should note that, since identification of the sharing rule in model I is obtained by assumptions regarding functional form, the test discussed above can also be interpreted as tests of functional form. It is, therefore, important to impose as few restrictions as possible during estimation. In addition to the estimates presented in Tables 2 and 3, we have tried other more complex functional forms of the leisure demand equations, for instance by allowing for quadratic wage effects. However, the additional parameters introduced did not contribute significantly to the value of the likelihood function. Therefore, our conclusions regarding model selection may not depend much on the choice of the functional form for the leisure demand functions.

Turning to the individual parameter estimates, note first that the (relative) amount of time spent working in the household is insensitive to the relative wage. Similarly, the wage and income effects are generally not significantly determined in the leisure demand equations, and the point estimates differ between the different versions of the collective model.

As can be seen from Tables 2 and 3, the estimates suggest that the presence of pre-school children (0–6 years of age) reduces leisure time for both spouses. In addition, younger school children (7–12 years of age) significantly reduces

Table 2. Estimation results 1984 sample

Equation/parameter	Model I		Model II		Model III	
	Estimate	t-value	Estimate	t-value	Estimate	t-value
Household work eq.						
Constant	−1.13	−3.28	−1.08	−3.15	−1.15	−3.33
Relative wage	−0.173	−0.70	−0.147	−0.61	−0.173	−0.70
Male age	0.008	0.47	0.010	0.59	0.007	0.41
Female age	−0.001	−0.07	−0.004	−0.26	$0.14 \cdot 10^{-3}$	−0.46
Male education	−0.072	−0.39	0.017	0.10	−0.071	−0.39
Female education	0.459	2.36	0.324	1.92	0.463	2.38
Child dummy (0–6)	0.247	1.66	0.233	1.56	0.238	1.60
Child dummy (7–12)	0.006	0.05	0.005	0.04	0.013	0.10
Child dummy (13–17)	0.003	0.02	$-0.61 \cdot 10^{-5}$	−0.18	0.003	0.02
Female demand eq.						
Constant	5073.4	10.70	5358.87	13.47	4965.15	10.83
Own marginal wage	8.93	0.55	7.90	0.42	15.53	0.86
Cross marginal wage	18.71	0.91	0^a		26.71	0.87
Full income	0.001	0.85	0.002	1.08	$0.2 \cdot 10^{-4}$	0.005
Female age	8.01	0.32	−6.58	−0.99	−9.21	−0.67
Male age	−14.40	−0.55	0^a		4.42	0.33
Female education	102.95	0.63	116.17	0.80	103.56	0.63
Male education	−105.74	−0.70	0^a		−87.21	−0.60
Child dummy (0–6)	−236.99	−1.93	−241.24	−1.97	−236.00	−1.89
Child dummy (7–12)	−105.11	−1.02	−121.37	−1.18	−94.67	−0.91
Child dummy (13–17)	−22.17	−0.20	−0.025	−18.56	−7.90	−0.07
Male demand eq.						
Constant	5840.8	16.93	5999.70	17.69	6047.03	18.71
Own marginal wage	−8.30	−0.31	−18.59	−0.83	−11.80	−0.57
Cross marginal wage	−3.89	−0.21	0^a		−11.42	−0.90
Full income	0.003	1.61	0.003	1.31	0.003	1.08
Male age	37.55	1.48	−4.73	−0.95	−1.62	−0.16
Female age	−42.73	−1.67	0^a		−5.97	−0.58
Male education	−173.01	−1.48	−216.06	−2.10	−203.87	−1.82
Female education	−140.22	−1.13	0^a		−124.93	−1.03
Child dummy (0–6)	−261.11	−2.80	−248.76	−2.67	−256.37	−2.76
Child dummy (7–12)	−156.37	−2.01	−163.63	−2.10	−163.58	−2.11
Child dummy (13–17)	22.56	0.28	29.81	0.37	5.22	0.06
Sharing rule param.						
Age difference	0.171	1.94	0.006	0.22	0^a	
Years of educ. diff.	0.057	1.18	0.041	1.03	0^a	
Marginal wage diff.	−0.030	−0.50	−0.028	−0.97	0^a	
Nonlabour income diff.	$-0.57 \cdot 10^{-4}$	−1.13	$-0.46 \cdot 10^{-4}$	−1.46	0^a	
Sexratio	1.89	0.77	1.34	0.63	0^a	
Log L	−5342.22		−5345.10		−5364.04	

Note: Standard errors are heteroscedastic consistent. a: parameter restricted.

leisure time for the male in the 1984 data. This effect is not present in the 1993 data, where the presence of older school children (13–17 years of age) instead significantly reduces leisure time for the female. The results also suggest that the individual education level is important for the allocation of time within the household. In the 1984 data, female education appears to be an important determinant of the relative amount of time spent in household work, while male education is not. In households where the woman is highly educated, she spends

Table 3. Estimation results 1993 sample

Equation/parameter	Model I		Model II		Model III	
	Estimate	t-value	Estimate	t-value	Estimate	t-value
Household work eq.						
Constant	−0.431	−1.20	−0.405	−1.14	−0.422	−1.18
Relative wage	0.085	0.72	0.113	1.13	0.180	1.69
Male age	−0.001	−0.08	−0.006	−0.42	−0.001	−0.08
Female age	−0.002	−0.13	0.002	0.14	−0.002	−0.15
Male education	−0.098	−0.62	0.065	0.45	−0.127	−0.81
Female education	0.171	1.16	0.028	0.24	0.194	1.31
Child dummy (0–6)	−0.057	−0.35	−0.045	−0.28	−0.058	−0.35
Child dummy (7–12)	−0.080	−0.55	−0.080	−0.55	−0.074	−0.51
Child dummy (13–17)	−0.057	−0.41	−0.057	−0.41	−0.060	−0.43
Female demand eq.						
Constant	5725.84	21.71	5718.85	22.74	5695.30	21.94
Own marginal wage	10.45	1.65	1.65	1.06	−2.47	−0.31
Cross marginal wage	−1.67	−1.46	0^a		−5.27	−0.53
Full income	−0.001	−1.50	$-0.27 \cdot 10^{-3}$	−1.87	$0.89 \cdot 10^{-3}$	0.44
Female age	−32.44	−2.11	1.23	0.22	−20.70	−1.81
Male age	33.94	2.21	0^a		22.09	1.94
Female education	−135.51	−1.29	−170.29	−1.76	−151.81	−1.46
Male education	−185.58	−1.65	0^a		−194.07	−1.71
Child dummy (0–6)	−265.42	−2.29	−240.36	−2.08	−255.01	−2.16
Child dummy (7–12)	97.96	0.96	76.47	0.74	87.22	0.86
Child dummy (13–17)	−195.08	−2.00	−190.61	−1.93	−205.27	−2.11
Male demand eq.						
Constant	5985.74	25.08	6114.27	27.21	6068.08	27.14
Own marginal wage	19.79	2.72	−2.24	−1.66	8.26	1.10
Cross marginal wage	−1.01	−0.74	0^a		9.40	1.57
Full income	−0.002	−2.86	$0.21 \cdot 10^{-3}$	1.88	−0.002	−1.25
Male age	−17.80	−0.95	−7.31	−1.54	5.60	0.57
Female age	12.88	0.69	0^a		−12.03	−1.22
Male education	58.79	0.59	−70.43	−0.80	29.28	0.30
Female education	−218.37	−2.34	0^a		−171.85	−1.92
Child dummy (0–6)	−285.46	−2.83	−334.46	−3.37	−320.89	−3.21
Child dummy (7–12)	−31.64	−0.36	−29.74	−0.34	−20.01	−0.23
Child dummy (13–17)	−93.44	−1.09	−67.59	−0.80	−74.03	−0.88
Sharing rule param.						
Age difference	0.056	1.46	−0.67	−0.83	0^a	
Years of educ. diff.	−0.046	−2.18	−0.75	−0.82	0^a	
Marginal wage diff.	−0.016	−6.46	0.004	0.61	0^a	
Nonlabour income diff.	$0.44 \cdot 10^{-7}$	0.03	$0.14 \cdot 10^{-4}$	0.43	0^a	
Sexratio	−1.37	−1.19	−5.63	−0.32	0^a	
Log L	−5777.73		−5792.62		−5788.73	

Note: Standard errors are heteroscedastic consistent. a: parameter restricted.

relatively less time in household work in comparison with households where the female has less education. The important insights with respect to the consequences of education appear to be that educated females work more in the market and enjoy about the same leisure as do uneducated females. For men, on the other hand, the results using the 1984 data provide a weak indication that educated men enjoy less leisure but work more in the labor market and/or in the household depending on the characteristics of the spouse. In the 1993

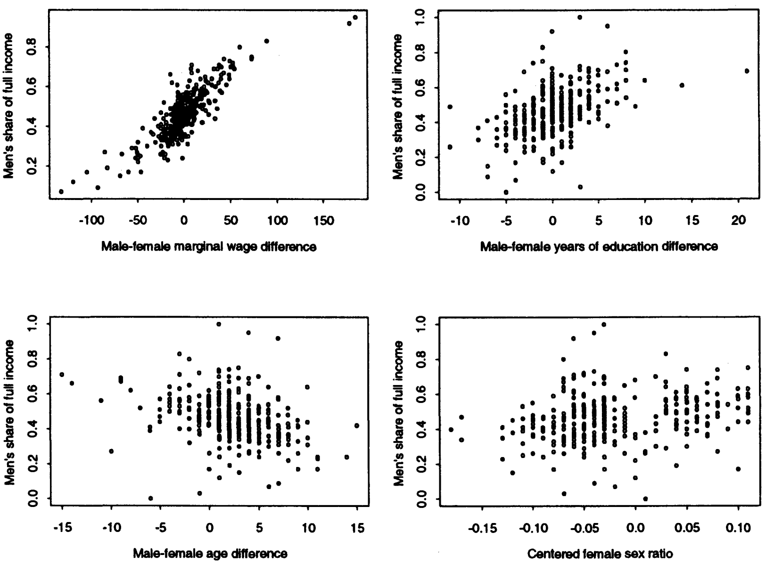

Fig. 1. Estimated sharing rule 1993

data, highly educated females seem to have a negative effect on the leisure consumed by the male, but there is no evidence that these men spend more time in household production.

Turning to the determinants of the sharing rule, in the 1984 data none of the determinants are statistically significant at the 5% level. By using the 1993 data, model I suggests that years of education differences appear to matter as well as wage differences, and to some extent age differences. In contrast to Chiappori et al. (1998), the sex ratio does not significantly influence the sharing rule and leisure consumption. Figure 1 plots the men's estimated share of full income for the year 1993 against four of the variables entering the sharing rule. As can be seen from the figure, male-female wage differences and years of education differences are positively related to the men's share of full income.

One important implication of the unitary model of labor supply is that the distribution of income within the household does not matter for the allocation of leisure in the household, i.e. only aggregate income matters. This so called pooling hypothesis has been tested in several earlier studies, and most studies find that the pooling hypothesis can be rejected. In the present framework we can test the hypothesis simply by checking the t-value for the parameter of nonlabor income differences. As can be seen from the tables, we are not able to reject the pooling hypothesis at the conventional 5% level. We would, nevertheless, like to exercise caution when interpreting this result. Personalized nonlabor incomes may be difficult to measure and there may be an element of choice involved when distributing nonlabor incomes between the spouses. Hence, although we cannot reject the hypothesis, this may be due to poor measurement and/or endogeneity problems.

The estimates based on our main sample is compared to the estimates from a restricted sample containing families without children. The results are presented in Tables A1 and A2 in the Appendix. The number of observations of this restricted sample is 81 in 1984 and 144 in 1993, and it may therefore be difficult to draw any strong conclusions based on such a small number of observations. Nevertheless, the results are similar to those obtained for the main sample. In the 1984 data, we are not able to reject models II and III, while both these models can be rejected using the 1993 data.

5. Conclusions

This paper analyzes leisure consumption and household production within the framework of a collective model. The paper should be viewed as a first attempt to empirically include household production in the collective framework. Three different models are estimated on Swedish data from 1984 and 1993. By comparing the results of the different models, we are able to reject the unitary model against a (statistically) more general alternative using the 1993 data. In the 1993 data, we are also able to reject the version of the collective model where the household good can be traded against the model where the household good is non-tradable. On the other hand, by using the 1984 data, we cannot distinguish between the three models. A formal test of the income pooling hypothesis indicates that pooling cannot be rejected. This result contradicts many earlier studies. The major determinants of leisure demands and household production appear to be household characteristics such as the presence of children and the education of the household members.

Note finally that the paper is based on a set of very restrictive assumptions regarding the technology and measurement of household production. First, the estimation rests on the assumption that household production can be characterized by constant returns to scale. One advantage of this assumption is that the system of equations to be estimated becomes recursive so that household work does not directly affect leisure demands. Future work should consider relaxing this assumption. Second, it may be fruitful to distinguish more carefully among different activities of home production in order to characterize the household technology more properly.

Appendix

Table A1. Estimation results 1984 sample (no children)

Equation/parameter	Model I		Model II		Model III	
	Estimate	t-value	Estimate	t-value	Estimate	t-value
Household work eq.						
Constant	−1.01	−2.26	−1.06	−2.43	−1.00	−2.26
Relative wage	−0.931	−1.61	−0.91	−1.78	−0.886	−1.55
Male age	−0.064	−1.73	−0.039	−1.15	−0.064	−1.72
Female age	0.072	2.00	0.046	1.41	0.072	1.99
Male education	−0.504	−1.16	−0.303	−0.74	−0.505	−1.16
Female education	0.782	1.82	0.674	1.79	0.789	1.84
Female demand eq.						
Constant	5643.87	9.33	5581.21	11.10	5631.57	9.51
Own marginal wage	95.57	1.30	10.88	0.51	26.37	0.91
Cross marginal wage	−39.47	−0.67	0[a]		31.43	0.74
Full income	−0.005	−1.05	−0.0003	−0.54	−0.005	−0.97
Female age	93.91	1.33	−1.46	1.48	26.21	1.04
Male age	−93.06	−1.33	0[a]		−26.76	−1.03
Female education	481.88	1.56	431.70	1.48	468.62	1.51
Male education	−359.85	−1.15	0[a]		−359.54	−1.16
Male demand eq.						
Constant	5679.84	12.39	5833.53	13.59	5687.19	12.59
Own marginal wage	102.01	1.37	26.30	1.32	48.25	1.49
Cross marginal wage	−43.75	−0.82	0[a]		7.99	0.35
Full income	−0.004	−1.10	−0.001	−2.57	−0.004	−1.02
Male age	102.21	2.23	−5.79	−1.04	52.62	2.53
Female age	−109.34	−2.43	0[a]		−58.74	−2.90
Male education	−61.26	−0.24	32.91	0.14	−60.88	−0.25
Female education	28.24	0.11	0[a]		36.55	0.14
Sharing rule param.						
Age difference	−0.14	−1.20	0.65	1.21	0[a]	
Years of educ. diff.	0.007	0.23	−0.12	−0.59	0[a]	
Marginal wage diff.	−0.15	−0.99	0.13	0.61	0[a]	
Nonlabour income diff.	$0.15 \cdot 10^{-4}$	0.67	0.0005	1.11	0[a]	
Sexratio	1.61	0.58	8.12	0.48	0[a]	
Log L	−1359.35		−1360.41		−1360.63	

Note: Standard errors are heteroscedastic consistent. [a] parameter restricted.

Table A2. Estimation results 1993 sample (no children)

Equation/parameter	Model I		Model II		Model III	
	Estimate	t-value	Estimate	t-value	Estimate	t-value
Household work eq.						
Constant	−0.308	−0.61	−0.169	−0.34	−0.310	−0.62
Relative wage	0.344	1.65	0.350	1.77	0.469	2.35
Male age	−0.031	−1.02	−0.378	−1.35	−0.033	−1.08
Female age	0.022	0.77	0.028	1.02	0.024	0.83
Male education	−0.015	−0.06	0.283	1.22	−0.051	−0.19
Female education	0.450	1.69	0.032	0.15	0.485	1.81
Female demand eq.						
Constant	5865.94	17.67	5679.80	21.01	5676.35	14.20
Own marginal wage	10.40	2.16	9.25	1.18	1.07	0.09
Cross marginal wage	0.71	0.16	0^a		1.04	0.052
Full income	−0.001	−2.03	−0.0009	−1.04	$-0.85 \cdot 10^{-4}$	−0.03
Female age	63.55	1.90	0.82	0.14	2.83	0.17
Male age	−65.78	−1.87	0^a		−1.44	−0.08
Female education	19.09	0.12	−123.39	−0.92	−63.86	−0.42
Male education	−260.45	−1.58	0^a		−204.40	−1.26
Male demand eq.						
Constant	5790.45	16.68	5634.40	16.51	6463.56	17.66
Own marginal wage	8.91	1.10	21.40	2.28	−30.33	−1.71
Cross marginal wage	2.21	1.69	0^a		−11.86	−1.14
Full income	−0.001	−2.10	−0.002	−2.67	0.004	1.43
Male age	79.99	2.48	1.15	0.18	19.48	1.18
Female age	−78.75	−2.58	0^a		−23.01	−1.44
Male education	46.10	0.30	−123.80	−0.88	−7.50	−0.05
Female education	−374.25	−2.57	0^a		−353.43	−2.43
Sharing rule param.						
Age difference	−0.358	−2.95	0.038	1.27	0^a	
Years of educ. diff.	−0.127	−1.82	−0.031	−1.34	0^a	
Marginal wage diff.	−0.018	−3.88	−0.019	−3.90	0^a	
Nonlabour income diff.	$-0.43 \cdot 10^{-5}$	−0.74	$0.15 \cdot 10^{-6}$	0.05	0^a	
Sexratio	−4.12	−1.36	−1.94	−1.10	0^a	
Log L	−2470.25		−2478.28		−2479.14	

Note: Standard errors are heteroscedastic consistent. [a] parameter restricted.

Endnotes

[1] As noted by Pollak and Wachter (1975), characteristics influencing preferences and household productivity cannot be distinguish from each other if the household production process is not characterized by constant returns to scale.

[2] Formally, let the tax system be described by J linear segments and denote by τ^k, $k = 1 \dots J$, the marginal tax rate corresponding to each segment. Further, let H^k, $k = 1, \dots, J-1$, be kink-points in the tax schedule in terms of hours of work, where the labor supply interval (H^{k-1}, H^k) corresponds to the marginal tax rate τ^k. Then, the virtual income of the individual observed on segment k can be calculated as $\tilde{y}^k = w^g(\tau^k H^{k-1} - \sum_{j=2}^{k-1} \tau^j (H^j - H^{j-1})) + y$, where w^g is the gross wage rate and y the nonlabor income.

[3] Extra-environmental parameters was first definied by McElroy (1990) as factors affecting the intra-household bargaining power, although they do not affect individual prices and nonlabor incomes. Examples of *EEPs* are the competitiveness in the marriage market, additional non-labor income if the household is dissolved, the elimination of the marriage tax and the legal structure within which household formation and seperation occur.

[4] To see this, note that the derivates with respect to nonlabor income are given by $\frac{\partial l_i}{\partial y_j} = \frac{\partial l_i}{\partial s_i}\frac{\partial s_i}{\partial y_j}$, $i, j = m, f$. Since $s_i = s_F - s_j$, where s_F is known, it holds that $\frac{\partial s_i}{\partial y_j} = \frac{-\partial s_j}{\partial y_j}$. This means that we have four unknown partial derivates on the right hand side of the above equation, $\frac{\partial l_m}{\partial s_m}, \frac{\partial l_f}{\partial s_f}, \frac{\partial s_m}{\partial y_m} = \frac{-\partial s_f}{\partial y_m}$, and $\frac{\partial s_f}{\partial y_f} = \frac{-\partial s_m}{\partial y_f}$. This is equal to the number of income effects that can be identified, i.e., the partials on the right hand side of the above equation are exactly identified. For discussions about identification in the extended collective model, see Apps and Ress (1997) and Chiappori (1997).

[5] For further details about HUS, see Klevmarken and Olovsson (1993) and Flood et al. (1997).

[6] For individuals that have been interviewed about their time allocation once on a weekend and once during the working week, time used for each household activity is computed as a weighted average with the weights 5/7 for week days and 2/7 for weekend days.

[7] The sex ratio for the female is computed as:

$$\frac{F_t}{M_t \cdot \left(\frac{F_t}{F_t + F_{t-1} + F_{t-2} + F_{t-3}}\right) + \cdots + M_{t+3} \cdot \left(\frac{F_t}{F_{t+3} + F_{t+2} + F_{t+1} + F_t}\right)}$$

where M_t and F_t are the number of males and females of age t in a specific region (county). In a similar vein, we have constructed a male sex ratio. The two measures, however, are highly correlated (the correlation coefficient is -0.73 in 1984 and -0.71 in 1993). Therefore we have chosen to include only the female sex ratio in the empirical analysis.

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The allocation and value of time assigned to housework and child-care: An analysis for Switzerland

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Abstract. In this paper, data from the 1997 Swiss Labour Force Survey are used to analyse the allocation and value of time assigned to housework and child-care. It is shown that men's allocation of time to housework and child-care is largely invariant to changes in socio-economic factors. Women's allocation of time to housework and child-care, on the other hand, is shown to depend on several social, economic, and demographic factors. The value of time assigned to housework and child-care is calculated with two market replacement cost methods and three opportunity cost methods. The results show that the value of time assigned to housework and child-care ranges from 27% to 39% and from 5% to 8% of GDP (in 1997), respectively. The value of time assigned to housework and child-care is also calculated for different household structures.

JEL classification: D13, J13, J22

Key words: Time allocation, child-care, household production

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1. Introduction

Time is one of society's most important economic resources, and yet only a fraction of households' time, namely that time spent on market activities, has received attention in economic teaching and research. Time spent on housework and child-care has gone largely unnoticed. Shelton reflects that this is primarily because "what goes on in the household is not intrinsically interesting since (1) women do it, (2) it is not in the public sphere, and (3) it is not subject to change through policy" (Shelton 1992, p. 64). Although these three reasons still (to a large extent) reflect our society at the end of the twentieth century, some change is taking place in most industrialised countries (see also Hewitt 1993, pp. 55–56). First, the traditional gender roles (i.e., the man at work, the woman at home) are being questioned, and more women are actively participating in the labour market. Although it would be naive to suggest that men take on just as much responsibility in the household as women, there does appear to be an increased willingness on behalf of men to participate in the running of the household. Second, because of the increased participation of women in the political debate, the public is being made aware of the fact that the household distribution of time and the value of unpaid labour is indeed a topic for the public sphere, and that it can (at least partially) be directed by public policies. Currently, in Switzerland, topics concerning unpaid labour are very often encountered in public debate and political action.[1] Yet, although the time spent on unpaid activities in the household has already entered the Swiss public sphere, very little is known about the way this time is allocated and what the value of this time is (see, however, Schellenbauer and Merk 1994; Sousa-Poza and Widmer 1998; Schmid et al. 1999).

This paper aims to, first, analyse the *determinants* of the allocation of time assigned to housework and child-care and, second, to calculate the monetary *value* of this time. Both of these topics have been extensively researched in other countries (for an overview see Gronau 1997; Juster and Stafford 1991; OECD 1996).[2] The structure of this paper is as follows: Sect. 2 presents the theoretical framework. Section 3 discusses the methodological issues and the empirical specification. Section 4 analyses the data. Section 5 presents the results, and Sect. 6 concludes.

2. Theoretical background

The conceptual framework of this study is based on the household theory developed by Becker (1965) and extended by, among others, Gronau (1977). It is known as the "New Home Economics" (NHE) theory. In traditional microeconomic theory, households maximise their utility by consuming a combination of goods and subject to a budget constraint. Becker (1965) extended and modified this traditional model in two ways (see also Gronau 1986): first, Becker argued that it was not the goods as such which render utility, but, instead, the "commodities" in which these goods serve as inputs. Second, Becker expanded the set of inputs. According to him, commodities are not only produced through the combination of market goods, but also with a certain amount of time. This implies that a household is restricted not only by a budget constraint, but also by a time constraint. Individuals therefore maximise a utility function which has commodities (which are produced by market goods and consumption time) as arguments and subject to a time and budget constraint. The result of this maximisation problem will give rise to demand functions for the arguments of the utility function and the inputs used in the production of home goods. The reduced form equation for the demand of unpaid labour time can be written as follows:

$$H = f(Z)$$

where H is the time spent doing unpaid work and Z is a vector of socio-economic variables such as age, education, presence of children, potential wage, etc. The comparative static properties of such a model are discussed in Gronau (1977).

Compared to the sociological theories underlying the division of the allocation of time within a household[3], NHE provides a framework which is parsimonious and which can be reasonably well tested empirically. Nevertheless, this framework has been criticised on a number of grounds (see, for example, Hannan 1982; Ben-Porath 1982; Pollak 1985; Ferber and Birnbaum 1977; Berk 1985). Perhaps the main criticism of the NHE model is aimed at its analogy between households and firms, where individuals within the household operate in an "implicit" rather than an explicit market (Folbre 1986). The NHE model then applies techniques developed for studying the firm. This analogy between households and firms is, however, often strained: transactions taking place within a household do not conform closely to those of a competitive market, demand and supply are often embedded in the same person, costs are seldom exogenous, and households may not maximise profits. Furthermore, since it is assumed that firms operate in frictionless and efficient markets, a similar assumption applies to households. A consequence thereof is that the internal structure and organisation of households play as minor a role as that in the neo-classical theory of the firm. Although the whole notion of well-behaved markets may appropriately be applied to the analysis of the firm, it makes little sense in the analysis of household behaviour. According to Ben-Porath, "once we deal with transactions within small groups, the nature of markets and equilibria, the meaning of prices and competition cannot be taken for granted, and strategic behaviour is likely to be relevant, and thus there is need for greater variety in the theoretical inputs" (Ben-Porath 1982, p. 61).[4]

NHE has, however, established itself as the leading economic theory with regard to the allocation of time. In this study, the factors which determine the allocation of time to housework and child-care are analysed. This analysis will primarily revolve around the insights gained from NHE. Nevertheless, the approach taken here is an eclectic one in the sense that a certain amount of theorising will be based on insights from other theoretical traditions, or simply through educated guesses. Considering the limitations of the NHE theory, such a broader approach is indispensable; without it, a more or less all-encompassing analysis of the determining factors of time assigned to unpaid labour becomes virtually impossible (see also Berk 1985, p. 34). The valuation of time is also strongly based on NHE theory. This applies especially to the opportunity cost methods, which build on the NHE notion that individuals maximise their utility by freely choosing between different time categories.

3. Methodological issues and empirical specification

3.1. The determinants of the allocation of time

In order to analyse the determinants of the allocation of time to unpaid labour, functions of the following form need to be estimated:

$$H_i = \beta_i' X_i + \varepsilon_i$$

where H is the number of hours spent doing a specific unpaid activity (here housework and child-care), X is a vector of explanatory variables, β is a vector of coefficients to be estimated, and ε is a stochastic disturbance term. As a large portion of the population did not spend any time on child-care (18% of females and 29% of males which have children under the age of 14 reported doing no child-care) and housework (6% of females and 16% of males reported doing no housework), one is confronted with a limited-dependent variable problem. The usual way of tackling this problem is by estimating a standard tobit model. The latent (dependent) variable of the tobit model is usually interpreted as being the propensity, capacity or desire to assign time to a certain activity. Thus, although a large portion of the population reported having spent the same amount of time on child-care or housework (i.e., 0 hours), certain individuals may differ in their desire to provide such time.[5] In this study, tobit models were estimated. A similar approach was taken by Gustafsson and Kjulin (1994) and Malathy (1994).

The *dependent variables* in the hour equations are the number of hours spent per week on housework and child-care. Housework includes the following activities: preparing meals, washing-up, setting the table, shopping, cleaning, vacuum-cleaning, tidying-up, making the beds, washing, ironing, repairing, renovating, sewing, knitting, taking care of the pets, gardening, and administrative work. Child-care includes: feeding, bathing, helping with homework, going for walks, and accompanying children.

The *explanatory variables* are the following: age, age squared, high education, low education[6], children aged 0 to 1, children aged 1 to 2, children aged 2 to 6, children aged 6 to 14, children aged 14 to 20, ownership of a large home (more than 4 rooms), ownership of a small home (4 rooms or less), large rented home, small rented home, marital status, Swiss nationality, German speaking, rural area, and the logarithm of the (imputed) hourly wage rate. Since the analysis of child-care is restricted to individuals with children aged 0 to 14, the variable "children aged 6 to 14" was used as the reference category in the child-care regressions. The imputed wage rates were computed, for women, with maximum-likelihood estimates of a sample-selection model[7], and, for men, by an OLS model. Such an approach is appropriate since males are to over 90% employed, i.e., the degree of censuring is small. The extended human-capital wage functions estimated in this study are of the following form:

$$W_i = \alpha_0 + \alpha_1 \cdot \mathrm{EDU}_i + \alpha_2 \cdot \mathrm{EXP}_i + \alpha_3 \cdot \mathrm{EXP2}_i + \alpha_4 \cdot \mathrm{MANAG}_i + \alpha_5 \cdot \mathrm{TEN}_i + \varepsilon_i$$

where ε is a stochastic disturbance term; the α_j's are the coefficients to be estimated; EDU characterises the number of years schooling; EXP captures the years of working experience; EXP2 is the corresponding squared term; MANAG is a dummy variable depicting whether or not the individual has a management position; TEN captures the years of tenure[8]. The wage rate (W) used here is the logarithm of a net hourly wage rate based on the number of paid hours usually worked per year.

3.2. The value of time

The two most common methods used to value the amount of time spent on unpaid labour are the *market replacement cost method* and the *opportunity cost method* (see, for example, Chadeau 1992). The market replacement cost method multiplies the number of hours by the wage rate of a market substitute. Usually, the wage rate of a professional housekeeper is used. In this study, the average wage rate of a so-called "hauswirtschaftliche Angestellte" (in essence, a housekeeper) is used (equal to Fr. 16.35 per hour, net of taxes and based on paid working time). This version of the market replacement cost method is often called the *generalist method.* The *specialist method* is another version of the market replacement cost method. With this method, the value of time spent on unpaid labour is obtained by multiplying the wage rate of a professional with the amount of time spent on the corresponding activity for which this professional is a specialist. Thus, for example, the number of hours spent cooking is multiplied by the average wage rate of a cook. A refinement of the specialist method can be obtained by creating *equivalence groups* (see, for example, Chandler 1994). An equivalence group is a collection of professions which can serve as equivalent substitutes for a specific unpaid labour activity. In this study, such equivalence groups are used.[9]

The opportunity cost method values the time spent on unpaid labour by multiplying the number of hours by a measure for the "forgone profits" incurred by not spending that time on another activity. These "forgone profits" can be quantified in a number of ways. The most common approach is to define them with the forgone earnings that an individual faces by not spending that time working in the market. For employed individuals, these forgone earnings are equal to the hourly wage rate they earn. For non-employed individuals, these forgone earnings can be estimated by calculating their *potential wages* with wage functions. A second way to quantify these "forgone profits" is by estimating an individual's *reservation wage*, i.e., that wage rate at which an individual is indifferent between a unit of time spent working in the market and a unit of time at home. According to the neo-classical household model, these reservation wages are equal to the market wage for employed individuals. Non-employed individuals, however, have reservation wages above their potential wages since they have decided not to participate in the market (see, for example, Gronau 1986 and Chiswick 1982). Since the reservation wage for non-employed individuals is unobservable, they have to be estimated with the aid of an econometric model. The model used in this study is that of Nelson (1977), which is a censored regression model with unobserved stochastic threshold. It can be formulated as follows (see Maddala 1983, pp. 174–178):

$$Y_{1i} = \beta_1' X_{1i} + u_{1i}$$

$$Y_{1i} \text{ observed only if } Y_{1i} \geq Y_{2i}$$

$$Y_{2i} = \beta_2' X_{2i} + u_{2i} \quad \text{and} \quad Y_{2i} \text{ is unobserved and stochastic}$$

where Y_1, the wage rate, is assumed to be a linear function of certain socio-economic variables X_1 (namely education and experience) and a stochastic disturbance term u_1. The wage rate is only observed when individuals participate in the paid labour market, and this is the case when the wage rate is at least greater than or equal to the reservation wage Y_2. The reservation wage cannot be observed, but it is assumed that it can be characterised as a linear function of certain socio-economic variables X_2 (such as experience, education, presence of children, and marital status). The corresponding stochastic disturbance term is u_2. The coefficients to be estimated are β_1 and β_2 for the wage and reservation functions respectively. This model has also been used by Homan (1988) to value the time spent on unpaid labour. In this study, the two-step estimation of this model as described in Maddala (1983) is used (see Appendix B).

A number of statistical bureaus use *gender-specific average wages* as a measure for the opportunity costs (see OECD 1996; Schmid et al. 1999). Furthermore, the "opportunity costs" are usually calculated for the whole population, i.e., including individuals who are not in the active labour force. Needless to say, using average wage rates gives rise to very rudimentary approximations for the value of time, and, furthermore, when applied to the whole population, characterising this approach as an opportunity cost method is somewhat of a misnomer. Nevertheless, for the sake of completeness, such average wage rates are also applied in this study.[10]

Two important questions arise when trying to calculate the hourly wage rate used to value the time spent on unpaid labour: (i) Should net or gross wages be used? (ii) Should actual or paid working time be used? Unfortunately, for most methods there is no precise prescription, and there is also very little consensus in the prevailing practice. Yet the choice of wage concept is important, since the calculated value is very sensitive to this choice. The *market replacement cost methods* are, strictly speaking, based on a market substitution perspective, i.e., unpaid labour is valued with the wage rate of some market substitute. The hypothetical situation is one in which all unpaid labour were to be replaced by the market sector. If this perspective is taken at face value, then the cost of employing a worker should be used in the calculation. A gross wage concept (i.e., before taxes and after employer contributions) based on paid working time would therefore seem the most appropriate. De facto, however, no taxes and no employer contributions arise in the household production process. If one were to accept this "institutional framework" (see Schäfer and Schwarz 1994, p. 603–604; Statistisches Bundesamt Deutschland 1993, p. 7–8), then a net wage concept based on the actual working time is the most suitable. In this study *net hourly wages* based on *actual working time* are used. In theory, the *opportunity cost method* should use wages, on which individuals base their time-allocation decisions. This means that net (disposable) wages are the most appropriate. Furthermore, the hourly wage rates should be calculated using paid working time. In this study *net hourly wages* based on *paid working time* are therefore used.

Net wages are calculated by deducting income taxes and compulsory social security contributions (made by the employee) from a gross wage. The social security contributions are: unemployment insurance ("Arbeitslosenversicherung"), accident insurance ("Nichtbetriebsunfallversicherung"), pension fund contributions ("Pensionskassenbeiträge"), and old-age insurance scheme payments ("AHV-, IV-, EO-Beiträge").[11] Income taxes are made-up of two parts: community taxes ("Kantons-, Gemeinde- und Kirchensteuer") and federal taxes ("Bundessteuer"). In Switzerland, community-level tax rates differ substantially from community to community, and therefore only an approximation of the individual tax burden (based on individual and household characteristics) can be made. The community tax rates were approximated by indices calculated by the Swiss Federal Tax Office.[12] Federal tax rates were calculated according to the relevant law ("Bundesgesetz über die direkte Bundessteuer").[13] In order to obtain hourly wage rates, one needs to divide the annual wage by the number of hours worked per year. In general, one can distinguish two hour concepts: (i) paid working hours, and (ii) actual hours worked. Paid working hours are determined by law or collective agreements, and they include paid holidays and sick leave. Actual hours worked are based on the usual hours worked in a year. This includes paid and unpaid overtime and excludes holidays and weekends.

4. Data

The data for this study are taken from the 1997 Swiss Labour Force Survey (SLFS). The SLFS is a nation-wide and representative survey conducted annually by the Swiss Federal Statistical Office. During telephone interviews lasting approximately 20 minutes, individuals are questioned on a number of labour-market related topics. The 1997 SLFS collected (for the first time) time data on unpaid labour activities. A total of 16,207 respondents were questioned on 14 individual unpaid labour activities which can be collected into 4 broad categories: household work, child-care, care of elderly household members, and community services. This study restricts the analysis to the first two mentioned categories. Besides containing time-use data, the SLFS data set also has a number of socio-demographic and earnings variables (see Bundesamt für Statistik, 1996).[14]

In analysing the determinants of the allocation of time, only individuals between the ages of 18 and 62 were considered. Unemployed individuals, military recruits, and apprentices were also ignored. In other words, the analysis is restricted to individuals who (to a greater or lesser extent) can freely choose between paid and unpaid labour time. The same applies to the valuation of unpaid labour with the opportunity cost methods. This assumption (namely that individuals can freely choose between time categories) does not, however, apply to the market replacement cost methods. Therefore, these methods were applied to the whole population, i.e., to all individuals older than 14 years of age.

5. Results

5.1. The determinants of the allocation of time

The results of the tobit regressions are depicted in table 1. Only the marginal effects at the sample means are discussed. On average, men spend about half as much time on housework and child-care as women (13.33 hours compared to 27.72 hours per week for housework and 9.34 and 15.10 hours per week for child-care).[15] A further interesting point to note is that the explanatory power of the males' model is very low in the case of housework (pseudo-R^2 value of 0.012) and quite low in the case of child-care (pseudo-R^2 value of 0.105)[16]. This is a standard result, and it highlights the fact that, for men, the amount of time spent on unpaid labour is in essence a value that remains largely unaffected by changes in socio-economic factors.[17] In both the male and female samples, the amount of time spent on housework increases with age (at a decreasing rate). High levels of education increase male's time spent on both housework and child-care, although the effect is marginal (men with a high education spend 0.78 hours and 1.30 hours per week more on housework and child-care, respectively, than men with a medium education). A similar result was obtained by Hill and Stafford (1985)

Women with a low education spend 1.63 hours less on child-care than women with a medium education. As can be expected, the presence of children has the largest effect on women's allocation of time to both housework and child-care. Depending on the age of the children in the household, the amount of time spent on housework increases by 3.82 to 7.10 hours per week.[18] Compared to the amount of time spent on child-care for children aged 6–14 (reference group), the amount of time spent on child-care for children aged 0–1, 1–2, and 2–6 increases by 12.13, 6.33, and 3.38 hours per week, respectively. Thus, the amount of time spent on child-care decreases with the age of the child. In the male sample, the presence of young children (aged 0–6) increases the amount of time spent on child-care by approximately 2 hours per week (compared to the reference group). An interesting point to note is that men's time spent on child-care does not vary substantially with the age of the child. Furthermore, the presence of children does not affect men's allocation of time to housework. For both men and women, the presence of an older brother or sister in the household (aged 14–20) significantly reduces the amount of time spent on child-care (compared to the reference group). We take a closer look at the effect that the age of a child has on the amount of time spent on housework and child-care below. Owning a home with more than 4 rooms increases the amount of time spent on housework in the female sample. In fact, the magnitude of this effect is quite large: it increases the amount of time spent on housework by 3.46 hours per week (compared to the reference category, namely individuals that rent a home with 4 rooms or less). There are two possible explanations for this observation. First, home ownership may induce individuals to spend more time on unpaid activities such as gardening or renovations. Second, owning a large home is (at least in Switzerland) a good indicator of a high household income.[19] Women in such households are less likely to be employed and therefore have more time for household activities. However, owning a home with more than 4 rooms decreases the amount of time spent on child-care in both the male and female samples. This result also applies to males who own a smaller home. One possible reason for this obseravtion is that home ownership correlates with age, i.e., older individuals are more likely to own a home. Such individuals are also less likely to have young children in the household.[20] Married women spend more time on housework than single women, and the opposite applies to men. This is a standard result and shows that a certain amount of specialisation takes place in households (see also Becker 1991, pp. 30ff.). It is interesting to

Table 1. The determining factors of the allocation of time to housework and child-care[a]

	Housework		Child-care	
	Males	Females	Males	Females
Constant	1.14	28.06***	14.42*	−6.59
	(2.48)	(6.01)	(7.47)	(10.65)
Age	0.33**	0.94***	−0.41	−0.38
	(0.14)	(0.23)	(0.37)	(0.48)
$Age^2 \times 10^{-2}$	−0.29*	−0.81***	0.23	−0.075
	(0.16)	(0.28)	(0.45)	(0.63)
High education[b]	0.78**	0.16	1.30**	−0.13
	(0.38)	(0.87)	(0.58)	(1.08)
Low education[b]	−0.65	0.06	−0.05	−1.63**
	(0.62)	(0.67)	(0.98)	(0.80)
Children aged 0–1[b]	0.97	5.90***	2.47***	12.13***
	(0.75)	(1.02)	(0.70)	(0.82)
Children aged 1–2[b]	−0.02	7.10***	1.35***	6.33***
	(0.89)	(1.23)	(0.76)	(0.89)
Children aged 2–6[b]	−0.72	6.00***	2.06***	3.38***
	(0.57)	(0.75)	(0.54)	(0.61)
Children aged 6–14[b,c]	−0.21	5.99***	–	–
	(0.53)	(0.66)		
Children aged 14–20[b,d]	−0.21	3.82***	−2.92***	−5.76***
	(0.67)	(0.83)	(0.86)	(0.96)
Own home and 4 rooms or	−0.26	0.64	−2.11**	−0.97
less[b,e]	(0.56)	(0.76)	(0.87)	(0.97)
Own home and more than	0.03	3.46***	−1.97**	−2.26***
4 rooms[b,e]	(0.48)	(0.65)	(0.68)	(0.75)
Rented home and more	0.14	−0.15	−0.51	−1.39
than 4 rooms[b,e]	(0.66)	(0.91)	(0.82)	(0.89)
Married[b]	−0.84*	5.27***	1.09	0.79
	(0.48)	(0.57)	(1.37)	(0.81)
Swiss[b]	0.27	−2.06***	0.44	1.34
	(0.54)	(0.77)	(0.79)	(0.91)
Germanic[b]	1.35***	1.18**	−0.08	−0.98
	(0.40)	(0.53)	(0.57)	(0.63)
Rural[b]	0.55	0.88*	0.30	−0.32
	(0.40)	(0.53)	(0.54)	(0.61)
(Imputed) hourly wage rate	−0.50*	−11.21***	−0.14	9.45***
(ln)	(0.27)	(2.77)	(0.69)	(3.51)
Mean of dependent variable	13.33	27.72	9.34	15.10
N	5369	5956	1819	2057
Sigma	16.37	19.58	15.66	15.31
Proportion $y = 0$	16%	6%	29%	18%
Log likelihood	−19861	−24998	−5820	−7311
Pseudo-R^2	0.012[f]	0.147[f]	0.105[f]	0.304[f]

[a] Tobit regression: marginal effects at sample mean and standard errors in parenthesis.
[b] Dummy variables.
[c] Reference category in the child-care equations (note that child-care equations are only estimated for households with children between 1 and 14 years of age).
[d] In the child-care equations, this coefficient shows the effect of the presence of an older brother or sister in the household (since child-care equations are only estimated for households with children between 1 and 14 years of age).
[e] Reference category: rented home and 4 rooms or less.
[f] The pseudo-R^2 measure is that of McKelvey and Zavoina (1975).
*/**/*** Significant at the 10%/5%/1% level, respectively.

Table 2. Hours per week unpaid labour and the presence of children in different age groups

	Females		Males	
	Housework	Child-care	Housework	Child-care
Households with no children	21.95 (3439)	–	13.32 (3422)	–
One child: age 0–6	32.15 (386)	24.02 (386)	13.67 (337)	13.96 (337)
One child: age 6–14	29.74 (230)	9.20 (230)	14.24 (156)	6.14 (156)
Two children: age 0–6	35.49 (356)	23.31 (356)	11.90 (340)	13.03 (340)
Two children: age 0–6 and 6–14	34.40 (204)	13.59 (204)	13.20 (206)	9.51 (206)
Two children: age 6–14	35.29 (466)	6.94 (466)	12.90 (405)	5.18 (405)
Three children: age 0–6	37.92 (56)	26.92 (56)	12.57 (48)	11.75 (48)
Three children: age 0–6 and 6–14	40.44 (160)	15.35 (160)	11.76 (148)	8.75 (148)
Three children: age 5–14	37.58 (177)	6.29 (177)	13.98 (141)	5.71 (141)

Note: Number of observations in parenthesis.

note that foreign women tend to spend more time on housework than Swiss women. This result may either be associated with the fact that many foreign women are not employed in Switzerland and/or that cultural differences exist between foreign and Swiss women. One could perhaps argue that the traditional gender roles are even more common among the foreigners living in Switzerland than among the Swiss. Germanic Swiss invest more time in housework than their Latin counterparts. Women living in rural areas spend slightly more time on housework than women living in urban areas (although this result is only significant at the 10% level). An increase in the imputed hourly wage rate reduces women's time spent on housework. Such a result can be explained by a standard labour supply model. An interesting result is that there is a positive correlation between the imputed hourly wage rate and women's time spent on child-care. A similar result was obtain by Gustafsson and Kjulin (1994) and Malathy (1994).

In Table 2, the average number of hours spent on housework and child-care per week are depicted for different household structures which depend on the number of children in the household and the age group of the children. There are a few interesting points to note: (i) As discussed above, men's time spent on housework does not change with the presence of children. This is not the case for women, for whom the amount of time spent on housework increase when they have children. (ii) Both men and women spend the most time on child-care with their first child. The older the child, the less time intensive it becomes. (iii) There are substantial economies of scale with regard to child-care. In fact, households with more than one young child do not significantly increase the time spent on child-care. A similar result was also obtained by Tiefenthaler (1997) for Filipino women. This observation could be attributed to the increased child-care experience of parents who have more than one child and also to the fact that, when children play together, often less passive child-care is needed. (iv) For women, housework increases with the number of children in the house. This is not the case for men who remain largely unaffected by the household structure. For males, the same applies for the amount of time spent on child-care.

Table 3. The value of time assigned to housework and child-care (Mill. Fr.) in 1997

Method	Males		Females		Total	
	House-work	Child-care	House-work	Child-care	House-work	Child-care
Market replacement cost methods[a]						
Generalist method	38,107	7,016	85,031	11,420	123,138	18,436
	(10.30)	(1.90)	(22.98)	(3.09)	(33.27)	(4.98)
Specialist method	44,154	11,150	95,315	17,530	139,467	28,680
	(11.93)	(3.01)	(25.76)	(4.74)	(37.69)	(7.75)
Opportunity cost methods[b]						
Gender-specific average wage: all individuals	51,483	9,479	92,762	12,458	144,245	21,937
	(13.91)	(2.56)	(25.07)	(3.37)	(38.98)	(5.93)
Gender-specific average wage: individuals aged 18–62	39,222	9,300	70,312	12,224	109,534	21,525
	(10.60)	(2.51)	(19.00)	(3.30)	(29.60)	(5.82)
Potential wage	37,623	9,932	66,366	13,186	103,989	23,118
	(10.17)	(2.68)	(17.93)	(3.56)	(28.10)	(6.25)
Reservation wage	37,623	9,932	63,938	13,061	101,561	22,993
	(10.17)	(2.68)	(17.28)	(3.53)	(27.44)	(6.21)

Note: Percentage of GDP in parenthesis.
[a] Calculated with net wage rates based on actual hours worked.
[b] Calculated with net wage rates based on paid working hours.

5.2. The value of time

The value of time assigned to housework and child-care is depicted in Table 3. The market replacement cost methods are calculated using net wage rates and based on actual hours worked by an individual. The opportunity cost methods are calculated using net wage rates and based on paid working time. Furthermore, the market replacement cost methods are calculated for the whole population (i.e., all individuals older than 15 years), whereas the opportunity cost methods are restricted to individuals between the age of 18 and 62 (although we also apply the opportunity cost method based on gender-specific average wages to the whole population).

Depending on the methodology used, the value assigned to housework can vary from approximately 27% to 39% of GDP. For child-care, the value can range from 5% to 8% of GDP. In the case of housework, the largest value is obtained with gender-specific average wages and applied to the whole population. The lowest value is obtained when using reservation wages. When measuring the value of time assigned to child-care, the largest result is obtained with the specialist method, and the lowest with the generalist method. As one may expect, the contribution made by women is substantially larger than that of men. Using the market replacement cost methods the contribution made by men is approximately half of that made by women. Since men on average earn higher wages than women, the opportunity cost measures increase men's proportion compared to the market replacement cost methods.

One anomaly to note is that, for the sample as a whole, the reservation wage is lower than the potential wage.[21] As was discussed above, for non-employed individuals the reservation wage should (from a theoretical perspective) lay above the potential wage. Thus, the value of time calculated with reservation wages should be higher that when potential wages are applied. As

Table 4. Reservation and potential wages for non-employed women

	Reservation wage (in Francs per hour)[a]	Potential wage (in Francs per hour)[a]
Whole sample	18.56	19.52
Married	19.79	20.19
Children aged 0–6	20.55	20.21
Children aged 6–14	20.51	20.96
Children aged 0–6 *and* 6–14	22.03	20.56

[a] Net wage rate based on paid working hours.

Table 5. Costs of children in different age groups (Fr. per year)

	Opportunity cost method[a]			Market replacement cost method[b]		
	House-work	Child-care	Total	House-work	Child-care	Total
One child: age 0–6	30,586	26,826	57,412	21,549	17,699	36,111
One child: age 6–14	32,239	11,548	43,787	21,494	7,290	28,784
Two children: age 0–6	32,242	27,389	59,631	21,946	16,751	38,697
Two children: age 0–6 and 6–14	35,430	18,419	53,849	21,741	10,567	32,308
Two children: age 6–14	33,384	8,678	42,062	22,778	5,604	28,382
Three children: age 0–6	34,432	27,977	62,409	24,007	18,108	42,115
Three children: age 0–6 and 6–14	36,845	17,082	53,927	24,409	11,154	35,563
Three children: age 6–14	35,685	9,325	45,010	24,827	5,509	30,336

[a] Based on imputed hourly wages (net of taxes and based on paid working time).
[b] Market replacement cost method calculated with the average wage rate of a housekeeper equal to Fr. 16.35 per hour (net of taxes and based on paid working time).

can be seen in Table 3, this is not the case. In Table 4, the potential and reservation wages for non-employed women are depicted for different household structures. The reservation wage, taken over all non-employed women, is lower than the potential wage. One possible reason for this counterintuitive result is that the neo-classical conception that individuals can freely choose between different time-allocation categories, based solely on a comparison of their potential and reservation wages (which, in most cases are not observable), is not very realistic. It may well be possible that this choice is primarily determined by cultural, institutional or social factors. If this is the case, then the reservation wage need not lie above the potential wage. Finally, it could also be that the reservation wage functions (and the wage functions) are not fully specified: unobservable traits such as abilities and motivations may play an important role with regard to the market participation decision. In Table 4 it can be seen, however, that the reservation wage appears to increase (relative to the potential wage) with an increase in the number of children in the household. Thus, the mean of the reservation and potential wage rates depends on the distribution of the explanatory variables (and on the distribution of the disturbance terms).

In Table 5, the indirect costs of children (i.e., the time-use costs) in different age groups and according to the number of children in a household are presented. For the sake of clarity, only the two most common valuation methods, namely the opportunity cost method (calculated with imputed wages) and the market replacement cost method (calculated with the average wage rate of a housekeeper) have been used. An only child between the age of 0 and 6 will give rise to time-use costs of Fr. 57,412 per year when using the opportunity cost method and Fr. 36,111 per year when using the market replacement cost method. This large difference is due to the fact that housekeepers in Switzerland earn a relatively low wage (Fr. 16.35 per hour, net of taxes). Depending on the household structure, the indirect costs of children rise to Fr. 62,409 if the opportunity cost method is applied and Fr. 42,115 if the market replacement cost method is used (for households with three children, all under the age of 6).

6. Conclusion

The aim of this study was to analyse the allocation and value of time assigned to housework and child-care in Switzerland. The analysis in this paper was based on data from the 1997 Swiss Labour Force Survey, which contains nation-wide and representative data on the amount of time spent on specific unpaid labour activities. The results show that men spend about half as much time on housework and child-care as women, and that men's time allocation to unpaid labour is, to a large extent, invariant to changes in socio-economic factors. For the female sample, it was shown that a number of socio-economic factors influence the decision to invest time in housework and child-care. Especially the presence of children, the marital status, and the imputed hourly wage rate play an important role. The value of time assigned to housework and child-care was calculated with two market replacement cost methods and three opportunity cost methods. In the former case, a generalist method and a specialist method were applied. In the latter case, gender-specific average wages, potential wages and reservation wages were used. The results show that, depending on the methodology used, the values range from approximately 27% to 39% of GDP in the case of housework and from approximately 5% to 8% of GDP in the case of child-care. The generalist method and the opportunity cost method based on potential wages were also used to calculate the time-use costs of children for different household structures. It was shown that a single child between the age of 0 and 6 gives rise to time-use costs (for housework and child-care) of Fr. 57,412 per year if the opportunity cost method is used, and Fr. 36,111 per year if the market replacement cost method is used.

Appendix A
Summary statistics

Table A. Summary statistics[a]

	Housework		Child-care	
	Males	Females	Males	Females
Age	40.01	40.55	38.84	36.16
	(11.30)	(11.59)	(6.68)	(6.00)
Age^2 $\times$ 10^{-2}	1728.93	1778.54	1553.39	1343.18
	(938.32)	(968.32)	(538.69)	(448.12)
High education[b]	0.37	0.21	0.40	0.21
	(0.48)	(0.41)	(0.49)	(0.41)
Low education[b]	0.11	0.23	0.10	0.20
	(0.31)	(0.42)	(0.29)	(0.40)
Children aged 0–1[b]	0.07	0.06	0.21	0.18
	(0.26)	(0.24)	(0.41)	(0.39)
Children aged 1–2[b]	0.04	0.04	0.13	0.12
	(0.21)	(0.20)	(0.34)	(0.33)
Children aged 2–6[b]	0.15	0.15	0.45	0.43
	(0.36)	(0.35)	(0.50)	(0.49)
Children aged 6–14[b]	0.21	0.22	0.61	0.63
	(0.41)	(0.41)	(0.49)	(0.48)
Children aged 14–20[b]	0.09	0.10	0.14	0.14
	(0.28)	(0.30)	(0.35)	(0.34)
Own home and 4 rooms or less[b]	0.13	0.13	0.12	0.11
	(0.33)	(0.33)	(0.32)	(0.32)
Own home and more than 4 rooms[b]	0.24	0.24	0.34	0.30
	(0.43)	(0.43)	(0.48)	(0.46)
Rented home and 4 rooms or less[b]	0.54	0.55	0.40	0.45
	(0.50)	(0.50)	(0.49)	(0.50)
Rented home and more than 4 rooms[b]	0.09	0.08	0.14	0.14
	(0.28)	(0.27)	(0.34)	(0.34)
Married[b]	0.58	0.57	0.96	0.84
	(0.49)	(0.50)	(0.19)	(0.37)
Swiss[b]	0.86	0.88	0.84	0.87
	(0.35)	(0.32)	(0.37)	(0.34)
Germanic[b]	0.71	0.70	0.68	0.68
	(0.45)	(0.46)	(0.47)	(0.47)
Rural[b]	0.31	0.31	0.38	0.38
	(0.46)	(0.46)	(0.49)	(0.49)
(Imputed) hourly wage rate (ln)	2.59	2.98	2.80	3.01
	(0.73)	(0.17)	(0.40)	(0.15)
Mean of dependent variable	13.33	27.72	9.34	15.10
	(14.41)	(20.19)	(12.74)	(15.72)
N	5369	5956	1819	2057

[a] Standard deviation in parenthesis.
[b] Dummy variables.

Appendix B
Estimating reservation wage functions

The model to be estimated is the following (see Maddala 1983, pp. 174–178):

$$Y_{1i} = \beta_1' X_{1i} + u_{1i}$$

$$Y_{1i} \text{ observed only if } Y_{1i} \geq Y_{2i}$$

$$Y_{2i} = \beta_2' X_{2i} + u_{2i} \quad \text{and} \quad Y_{2i} \text{ is unobserved and stochastic}$$

Y_1 is the wage rate and Y_2 is the reservation wage. X_1 and X_2 are vectors of regressors for the two functions respectively. The corresponding stochastic disturbance terms are u_1 and u_2. β_1 and β_2 are the coefficient vectors to be estimated. For employed individuals, Y_1 is observed, and, for non-employed individuals, Y_1 is not observed.

If one assumes that u_1 and u_2 are *IN*, with zero means and covariance matrix,

$$\Sigma = \begin{pmatrix} \sigma_1^2 & \sigma_{12} \\ \sigma_{12} & \sigma_2^2 \end{pmatrix}$$

then if follows that

$$(u_1 - u_2) \sim N(0, \sigma^2) \quad \text{and where } \sigma^2 = \sigma_1^2 + \sigma_2^2 - 2\sigma_{12}.$$

It can now be shown that the likelihood function to be maximised is the following:

$$\log L = -N_1 \log \sigma_1 - \frac{1}{2\sigma_1^2} \sum_1 (Y_1 - \beta_1' X_1)^2 + \sum_1 \log \Phi(W) + \sum_0 \log \Phi\left(\frac{\beta_2' X_2 - \beta_1' X_1}{\sigma}\right)$$

where

$$W = \frac{1}{\sqrt{(\sigma_2^2 - \sigma_{12}^2)/\sigma_1^2}} \left(Y_1 - \beta_2' X_2 - \frac{\sigma_{12}}{\sigma_1^2} (Y_1 - \beta_1' X_1) \right)$$

and where $\sum_1$ is the sum over the N_1 observations for which Y_1 is observed, and $\sum_0$ is the sum over the $(N - N_1)$ observations for which Y_1 is not observed. This function can be estimated with the aid of a numerical algorithm. Note that a sufficient condition for the identification of the model is that at least one variable in X_1 is not included in X_2.

Maddala (1983) describes a simpler two-stage estimation of this model (see

Table B. Reservation wage functions for women

	Market participation probit	Wage function	Reservation wage function[a]
Constant	−0.251**	2.058***	2.108
Years schooling	0.039***	0.056***	0.048
Experience	0.132***	0.033***	0.007
Experience^2 × 10^{-2}	−0.298***	−0.059***	
Children between 0 and 6 years[b]	−0.816***		0.163
Children between 7 and 14 years[b]	−0.519***		0.104
Is married[b]	−0.586***		0.117
Selectivity correction term		−0.035	
N	5956	3491	
Log likelihood	−3141.623		
Pseudo-R^2/Adj. R^2	0.148[c]	0.127	

[a] All underlying coefficients are significant at the 5% level.
[b] Dummy variables.
[c] The pseudo-R^2 measure is that of McFadden (1973).
*/**/*** Significant at the 10%/5%/1% level, respectively.

Maddala 1983, pp. 228–230). In the first stage, a probit model representing the (reduced form) market participation decision is estimated:

$$I_i = Y_{1i} - Y_{2i} = \beta' X_i - u$$

where I is a dichotomous variable, and $u = u_2 - u_1$. In the second stage, a selectivity-corrected wage function is estimated, i.e.,

$$Y_{1i} = \beta_1' X_{1i} - \sigma_{1u}\lambda_{1i} + \varepsilon_{1i}$$

where λ_1 is the inverse of Mill's ratio.

From the probit model, one gets consistent estimates of $\frac{\beta_{1j}}{\sigma}$ and $\frac{\beta_{2j}}{\sigma}$ for the non-overlapping variables in X_1 and X_2. For the corresponding elements in X_1 and X_2, one gets estimates of $\frac{\beta_{1k} - \beta_{2k}}{\sigma}$. From the estimation of the selectivity-corrected earnings functions, one can get consistent estimates of β_1 and $\sigma_{1u} = \frac{\sigma_{12} - \sigma_1^2}{\sigma}$. If there is at least one variable in X_1 not included in X_2, then the $\frac{\beta_{ij}}{\sigma}$ estimate of this variable can be used to get a consistent estimate of σ, and thus all the remaining elements of β_2. In this study, this two-stage approach is used. The estimated wage and reservation wage functions are depicted in table B. Note that the wage function follows the usual parabolic labour productivity pattern. The reservation wage function, on the other hand, is usually assumed to depend linearly on experience, i.e., the reservation wage will rise with experience (i.e., age) but not eventually fall (see Homan 1988, p. 119). Furthermore, the reservation wage is assumed to depend on the years of schooling, the presence of children, and marital status.

Endnotes

[1] A number of parliamentarians and the Swiss Federal Council have, for example, expressed the need to value unpaid labour. See the following inquiries: Bacciarini (18/03/81 and 31/01/83); Fankhauser (01/06/94); Goll (17/06/94); FDP (02/02/95); Roth-Bernasconi (03/10/96).

[2] The analysis of the determinants of the allocation of time to housework and child-care is interesting from a sociological perspective since it tells us something about the society we live in, especially with regard to the differing gender roles. There are in essence three reasons for wanting to ascribe a monetary value to the time spent on unpaid labour. First, unpaid labour generates wealth and contributes substantially to a society's welfare. As a number of authors have come to recognise, this welfare contribution goes unnoticed in the conventional GDP measure (see, for example, Eisner 1988 and Gronau 1986). Consequently, various national statistical bureaus have, in the past decade, institutionalised the monetary valuation of unpaid labour. One field of research which has arisen from the valuation of unpaid labour is the analysis of the distribution of extended income (see Bonke 1992 and Jenkins and O'Leary 1996). Certain authors have further argued that policymakers should consider this source of wealth in tax and insurance schemes in order to guarantee a more equitable treatment of families (see, for example, Peskin 1984). A second reason is that such a monetary value is often needed in litigation testimony, in which the value of unpaid labour is used to determine appropriate compensations in case of injury, death, and divorce (see, for example, Douglass et al. 1990). Especially in Swiss legal practice there is a strong need for reliable valuation data since the current practice is based not only on antiquated data but also on an unsound methodological framework. A third important reason for the valuation of unpaid labour is that, one could argue, only an explicit monetary valuation catches the public's attention.

[3] Mention should be made of the "conjugal power" school which tries to explain the allocation of time on the basis of role and authority patterns, exchange and choice, family role typologies, and family decision-making. See Berk (1985), pp. 10ff.

[4] Further criticisms of NHE theory include: (i) the NHE assumption that a single individual maximises a household utility function obscures the fact that often household members have differing interests, (ii) cultural and institutional aspects are given no consideration, and (iii) its usefulness for empirical research is, to some extent, limited.

[5] The fact that there is a large portion of individuals that do not spend any time on child-care may come as a surprise. It can, however, be explained by: (i) time spent on active child-care such as feeding or bathing diminishes with the age of the child, and (ii) the reference period of the questions in the SLFS is based on the previous day. Thus, individuals were questioned on how much time they spent doing a certain unpaid activity on the previous day. Needless to say, such a short reference period may give rise to numerous "zero observations" (especially for those individuals with older children in the household). Other studies that analyse the allocation of time to child-care face similar levels of "zero observations" (see, for example, Gustafsson and Kjulin 1994; Malathy 1994).

[6] In this study two dummy variables are used to identify individuals with a high or low education. Degrees from the following institutions were considered as "high education": university, technical college ("höhere Berufsausbildung", "Technikum", "höhere Fachschule"), and high school ("Matura"). The following categories were considered as "low education": no degree, only compulsory schooling, one-year housekeeper apprenticeship ("Haushaltslehrjahr"), and lower apprenticeship schemes ("Anlehre"). The reference category was primarily made up of apprenticeships ("Berufslehre") and similar qualifications ("Vollzeitberufsschule").

[7] The selection equation depicts a woman's choice to participate in the market or not. The following explanatory variables were used: age, age squared, years of schooling, children aged 0–6, children aged 6–14, and marital status.

[8] The TEN and the MANAGE variables were set to zero for persons not employed. Furthermore, it should be noted that there is a potential endogeneity problem in including variables such as experience and tenure when calculating imputed wages for the hour functions since these variables reflect past time in paid employment which, in turn, is jointly determined with unpaid labour time.

[9] We have refrained from presenting the exact structure of the equivalence groups. These are available from the authors upon request.

[10] It is beyond the scope of this paper to comprehensively discuss all the methods used to value the time spent on unpaid activities. The interested reader is referred to Goldschmidt-Clermont (1993), Chadeau (1992), and Schmid et al. (1999).

[11] These contributions were calculated according to the regulations laid down by the relevant laws. Due to the lack of information, in certain cases the deductions had to be approximated. For a more precise description, see Schmid et al. (1999).

[12] See Eidgenössische Steuerverwaltung (1997a, 1997b).

[13] Due to the lack of information, a number of possible tax deductions could not be considered, i.e., only those tax deductions were considered which could convincingly be made based on the information available in the Swiss Labour Force Survey. See Schmid et al. (1999).

[14] It is often argued that telephone time-use surveys are less reliable that those conducted by the diary method (see, for example, Robinson 1985 and Niemi 1993). Robinson (1986) argues that the stylised questions used in telephonic surveys (for example, "How much time did you spend cooking yesterday?") will tend to overestimate the true value. Unfortunately, there is no reference data set (i.e., a time-use survey conducted by the diary method) in Switzerland which can be used to judge the reliability and validity of the SLFS data set. In Schmid et al. (1999), however, it is shown that, at an aggregate level, the amount of time spent on unpaid labour activities – as derived from the SLFS data set – is similar to the amount of time spent on unpaid labour in other OECD countries (which use the diary method).

[15] These averages apply to the *observed* dependent variable, i.e., zero values were treated as such.

[16] The pseudo-R^2 value is that of McKelvey and Zavoina (1975). Although McKelvey and Zavoina (1975) did not explicitly consider the tobit model with their pseudo-R^2 expression, Veall and Zimmermann (1994) show that it is also valid for the tobit model. Furthermore, they show that, compared to a number of other pseudo-R^2 expressions, this pseudo-R^2 performs the best as a "closeness to OLS-R^2" criterion. See also Veall and Zimmermann (1996).

[17] According to Jenkins and O'Leary, "the smaller extent of unexplained variation in domestic work time amongst females than amongst males might be explained by greater perpetuation of traditional home-making roles amongst women ('new men' increase the variance)" (Jenkins and O'Leary 1995, p. 274).

[18] As has been pointed out in other studies (see, for example, Malathy 1994), there is a potential endogeneity problem in incorporating these children variables. The authors have, however, estimated the reduced form equations (i.e., excluding the children variables), and the results obtained are quite similar.

[19] The SLFS data set does contain a variable for the non-labour income earned in a household. Unfortunately, this variable is most probably quite unreliable, and the large number of missing values restricts the sample size considerably. The situation is aggravated by the fact that individuals were allowed to make gross or net replies. One therefore further restricts the data set by having to decide on one of the two measures. The authors of this study did try to estimate (selectivity-corrected) imputed values. The underlying regressions were, however, very poor, thus giving rise to a variable with a very low variance. It was not possible to incorporate such a variable in the hour functions. Despite the important role non-labour income plays (especially with regard to the market-participation decision of women), a variable characterising such non-labour income has therefore been omitted in the analysis.

[20] We do not, however, observe a serious multicollinearity problem in our regressions.

[21] The reservation and potential wages are calculated with the estimates presented in table B1 (in the appendix).

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Who takes care of the children? The quantity-quality model revisited

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Abstract. We study the Becker and Lewis (1973) quantity-quality model of children adding an explicit child care time constraint for parents. Parents can take care of the children themselves or purchase day care. Our results are: (i) If there only is own care, a quantity-quality trade-off, different from that of Becker and Lewis (1973), arises. The income effect on fertility is positive if child quantity is a closer complement than child quality to the consumption of goods. (ii) If, instead, there is a combination of purchased and own care, the effect of income on fertility is ambiguous, even if quantity of children is a normal good in the standard sense. This is the Becker and Lewis (1973) result extended to a situation with a binding child care time constraint. The conclusion is that the Becker and Lewis (1973) result holds as long as at least some child care is purchased.

JEL classification: D1, J13

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1. Introduction

How does increased family income affect fertility? The standard answer is that fertility increases with income if the quantity of children is a normal good. The seminal contribution of Becker and Lewis (1973) (henceforth denoted BL) shows that this answer is seriously misleading.[1] A *ceteris paribus* increase in quality implies an increase in the marginal rate of substitution between quantity and quality, if quantity is a normal good. But such an increase in quality also increases the relative price of quantity in the BL model unlike standard models where prices are constant. The direction of the change in quantity when income increases is, therefore, indeterminate.

BL use a single period model that includes all phases of life for parents. Sometimes, see Hotz et al. (1997), the perspective of a newly married couple is emphasised. Recent empirical studies, e.g., Connelly (1992), Powell (1997), and Blau and Hagy (1998), adopt this perspective. They also recognise that small children require child care, which BL do not.[2]

Our purpose is to study how changes in income affect fertility in the quantity-quality model when parents face an explicit child care time constraint. We assume that the quality of children depends on the type of child care provided. In addition to taking care of the children themselves (own care) parents can also purchase care (day care).[3]

In some cases we replicate the BL results, in other cases we do not. Our main results are: If parents exclusively take care of the children themselves, a quantity-quality trade-off, of a different kind than that of Becker and Lewis (1973), arises. The income effect on fertility is positive if the quantity of children is a closer complement to consumption goods than the quality of children.

If there is a combination of own and purchased care, we find that the effect of income on fertility still is ambiguous when the quantity of children is a normal good. Necessary conditions for a solution with both own and purchased care are, however, that the marginal utility of spending time with the children is low and that the marginal utility of an additional child is high. This combination of conditions is not impossible but is somewhat odd.[4] But this is the Becker and Lewis (1973) result extended to a situation with a binding child care time constraint. The main conclusion is, therefore, that the Becker and Lewis (1973) result holds as long as at least some child care is purchased.

In Sect. 2 we describe our generalisation of the BL model. Section 3 derives the results and Sect. 4 concludes the paper.

2. Model

A parent reproduces asexually and chooses the quantity of children $n \in N = \{n \in \mathbb{R}_+ : n \geq 1\}$. Child quality $q \in \mathbb{R}_+$ is only acquired through child care, which is produced by the parent herself or purchased. Producing the care herself, the parent spends $c \in \mathbb{R}_+$ of her own time to take care of her n children. The number of purchased hours of day care are $d \in \mathbb{R}_+$ during which the child gets the full attention of a care taker. The total care time during the childhood of each child is $d + c$.[5] This must not be less than the total childhood time D during which *each* child needs care; $D \leq c + d$.

The quality of own care for each child equals the average time during which a child gets the full attention of the care taker; i.e., $\frac{c}{n}$. Purchased quality per child is proportional to the number of purchased hours of day care. As a matter of convenience we choose units so that an hour of day care yields a unit of quality. Own and purchased quality are perfect substitutes and the parent treats all children identically.

Average quality is $q = \frac{c}{n} + d$. We assume that the child care time constraint is binding, i.e., $D = c + d$. This implies that $q = D - \frac{(n-1)}{n}c$ and $\partial q/\partial c = n^{-1} - 1 \leq 0$. Substitution of day care for own care reduces average quality when the child care time constraint is binding.[6]

Working hours $\bar{h} \in \mathbb{R}_{++}$ are fixed and paid the wage rate w. Lifetime income is spent on own lifetime consumption ($x \in \mathbb{R}_{+}$), the price of which is numeraire and normalised to unity, or on purchased quality *pnd*, where p is the unit price of purchased day care, n is the number of children, and d is the quantity of purchased care per child. The parent's budget constraint is $w\bar{h} = x + pnd$.

The total time endowment during the lifetime T is spent on market work, taking care of one's children and leisure time ($\ell \in \mathbb{R}_{+}$): $T - \bar{h} = c + \ell$. Naturally, total childhood time is less than total time, $D \leq T$. The assumption that time in market work is fixed, leaves the parent with an own child care-leisure choice in the time dimension.

Parents have preferences represented by the quasi-concave utility function $U^* : \mathbb{R}^4_+ \times \mathbb{R}_{++} \times N \to \mathbb{R}$ defined by $U^*(x, \bar{h}, c, \ell, q, n)$. We use the notation $U_1^* = \frac{\partial U^*(\cdot)}{\partial x}$ etc to denote the partial derivatives and assume that $U_1^* > 0$, $U_2^* < 0$, $U_3^* \gtreqless 0$, $U_4^* > 0$, $U_5^* > 0$ and $U_6^* > 0$. Hence, we do not make any particular assumption about how the parent values the time spent with her own children.

Combining the time constraints yields

$$T - \bar{h} - D = \ell - d. \tag{1}$$

Substituting for c, ℓ, and q, the problem of a parent can be written as

$$\max_{x,d,n} U(x, d, n) \quad \text{s.t.} \quad w\bar{h} = x + pnd \quad \text{and} \quad d \geq 0, \tag{2}$$

where $U(x, d, n) = U^*\left(x, \bar{h}, D - d, T - \bar{h} - D + d, d + \frac{D-d}{n}, n\right)$. This problem has the following first order conditions for $x^* > 0$, $d^* \geq 0$ and $n^* \geq 1$

$$U_x - \lambda^* = 0, \tag{3a}$$

$$U_d - \lambda^* pn^* - \mu^* \leq 0 \quad d^* \geq 0 \quad \mu^* d^* = 0, \tag{3b}$$

$$U_n - \lambda^* pd^* = 0, \quad \text{and} \tag{3c}$$

$$w\bar{h} - x^* - pn^* d^* = 0 \tag{3d}$$

where $\lambda^* > 0$ and μ^* are the Lagrange multipliers in the optimal point associated with the constraints. Alternatively, we can express the derivatives of U in terms of derivatives of U^*, i.e.,

$$U_x = U_1^*, \tag{4a}$$

$$U_d = -U_3^* + U_4^* + \frac{n-1}{n} U_5^* \quad \text{and} \tag{4b}$$

$$U_n = -\frac{D-d}{n^2} U_5^* + U_6^*. \tag{4c}$$

3. Quality vs. quantity

We now revisit the problem of quantity versus quality of children and ask how fertility is affected by income changes when there is an explicit child care time constraint. We make the analysis in two steps: In Subsect. 3.1 we consider the corner solution where the parent produces all child care herself ($d^* = 0$). The interior solution when child care is arranged through a combination of purchased care ($d^* > 0$) and own care is discussed in Subsect. 3.2.

3.1. Only own care

Suppose that there is only own care; i.e., $d^* = 0$. Then (3b) implies $U_d - \lambda^* pn^* \leq 0$, possibly with a strict inequality. This situation may occur when the parent loves staying home to take care of the children and, therefore, $U_3^* > 0$ and also relatively high. Consumption is given by equation (3d) and equals income, which can be defined as $\bar{y} = w\bar{h}$. The main issue is how fertility is affected by income changes; i.e., what is the sign of $\frac{\partial n}{\partial \bar{y}}$?

When we consider the effect on an income increase the first order condition describing individual behaviour simplifies to $U_n = 0$, with the second order condition $\Delta_2 := U_{nn} < 0$. In the notation of the general model we have

$$U_n = -\frac{D}{n^2} U_5^* + U_6^* = 0 \quad \text{and} \tag{5a}$$

$$U_{nn} = \frac{D}{n^3} U_5^* + \frac{D^2}{n^4} U_{55}^* - \frac{2D}{n^2} U_{56}^* + U_{66}^* < 0. \tag{5b}$$

If $U_{nn} < 0$ we can continue with the comparative static analysis. However, this is not necessarily the case. Because quality is non-linear in quantity a solution satisfying the first order condition may be a local optimum only, giving lower or the same utility as the global optimum. Also, the global optimum may be the corner solution $n = 1$. In the following we disregard these problems and assume that the second order condition is satisfied so that there exists a unique interior solution ($n^* > 1$) for the quantity of children.

Total differentiation of the quantity of children n with respect to exogenous income $\bar{y}$ yields

$$\frac{\partial n}{\partial \bar{y}} = \frac{1}{\Delta_2}\left[\frac{D}{n^2}U_{51}^* - U_{61}^*\right], \tag{6}$$

where the denominator is negative by the second order condition.

The Eq. (6) shows the quantity-quality trade-off when there is no purchased care. This condition states that if the quantity of children is a sufficiently closer complement to the consumption of goods than quality in the sense of $\frac{D}{n^2}U_{51}^* - U_{61}^* < 0$, then increased exogenous income will increase the quantity of children. This would be the case, for example, if quality is a substitute and quantity is a complement to the consumption of goods.[7] Since $\frac{\partial q}{\partial n} = -Dn^{-2} < 0$ the quality of children will be reduced. However, if the quality of children is a closer complement to consumption of goods than quantity in the sense of $\frac{D}{n^2}U_{51}^* - U_{61}^* > 0$, then increased exogenous income will reduce the quantity of children and also increase the quality of children.

3.2. *Purchased and own care*

In a completely interior solution the parent is using purchased as well as own care. Then (3b) is strictly binding and $\mu^* = 0$ so that $d^* > 0$. Let the (assumed) unique solution satisfying these first order conditions (3a)–(3d) be denoted $(x(\bar{y}), d(\bar{y}), n(\bar{y}))$. Once again the issue is how fertility is affected by income changes; i.e., what is the sign of $\frac{\partial n}{\partial \bar{y}}$?

Consider the optimal non-linear solution evaluated in a linear model. In such a model we can write the budget as $I = p_n\tilde{n} + p_d\tilde{d} + \tilde{x}$, where $I = \bar{y} + pd(\bar{y})n(\bar{y})$ is full income, $p_n = pd(\bar{y})$ and $p_d = pn(\bar{y})$. Let $\tilde{S}_{ij}$ denote the substitution effect in the linear model where $i, j = p, d$. Standard symmetry gives $\tilde{S}_{ij} = \tilde{S}_{ji}$. Then we have

$$\frac{\partial n}{\partial \bar{y}} = \frac{\frac{\partial \tilde{n}(I)}{\partial I}(p\tilde{S}_{dn} - 1) - \frac{\partial \tilde{d}(I)}{\partial I}p\tilde{S}_{nn}}{p^2\tilde{S}_{dd}\tilde{S}_{nn} - (p\tilde{S}_{dn} - 1)^2}, \tag{7}$$

where $\frac{\partial \tilde{n}(I)}{\partial I}$ and $\frac{\partial \tilde{d}(I)}{\partial I}$ are standard income effects. Normality of the quantity of children in the standard sense implies $\frac{\partial \tilde{n}(I)}{\partial I} > 0$, but this is not sufficient to sign Eq. (7); see Razin and Sadka (1995, p. 20f) for a discussion about various conditions signing (7). This is essentially the BL result. The difference is that 'total quality' in their model corresponds to 'purchased quality', i.e., day care, in our model.

Necessary conditions for an interior solution are that the marginal utilities of purchased care and quantity are positive; i.e., $U_d > 0$ and $U_n > 0$. Although these marginal utilities look similar to the marginal utilities in BL we see the difference clearly if we study the signs of (4b) and (4c) rather than simply the signs of the derivatives of U. Then we see that given our assumptions $U_d \gtreqless 0$ and $U_n \gtreqless 0$.

Utility is affected through three different channels when the parent purchases an additional unit of day care. *First*, by the child care time constraint, the amount of time spent with children is reduced. This reduces utility if the parent likes to be with the children. *Second*, by the time constraint, more leisure time becomes available since working hours are fixed, which increases utility. *Third*, the quality per child is affected. As a direct effect, quality increases with one unit while the reduction of own care with one unit only reduces quality with n^{-1} units. Therefore, the quality effect is non-negative since we assume that $n^* \geq 1$. This means that the marginal utility of purchased care is positive if the second and third effects dominate the first effect.

When the quantity of children is increased marginally there will be a direct positive effect on utility and an indirect negative effect through reduced quality. If the direct effect dominates the indirect effect, then additional children will increase utility.

We can note that the necessary condition for purchased day care is met if the parent dislikes spending own time to take care of the children. In other words, $U_3^* < 0$ in the optimal point is sufficient for $U_d > 0$. Moreover, the necessary condition for the quantity of children is met if all care is purchased. In other words, $d^* = D$ is sufficient for $U_n \geq 0$. But this would move us from an interior solution to a corner solution or even beyond that. When the parent purchases more day care than necessary the child care time constraint is no longer binding and we are back to the model and the results of BL.

4. Conclusions

Becker and Lewis (1973) show that the effect of income on fertility is ambiguous, even if the quantity of children is a normal good in the standard sense. In this paper we have shown that this result extends to a situation where parents face an explicit child care time constraint and choose a combination of purchased day care and child care produced by themselves.[8]

On the other hand, if parents exclusively care for the children themselves, a different kind of quantity-quality trade-off arises. More children reduce the quality of an hour of the parent's time spent on child care. The income effect of fertility now is positive if the quantity of children is a closer complement than quality to the consumption of goods.

Consequently, the conclusion is that the Becker and Lewis (1973) result extends to a situation with a binding child care time constraint as long as at least some child care is purchased.

Endnotes

[1] For an early discussion see Becker (1960) and for further development Becker and Tomes (1976). For policy discussions see, e.g., Batina (1986), Cigno (1983, 1986), Ermisch (1989), and Nerlove et al. (1984, 1986).

[2] See also Lundholm and Ohlsson (1998) who apply the same perspective to analyse wage determination and female labour force participation.

[3] We also assume that parents without constraints can choose how much day care they want to purchase.
[4] Odd since parents on the margin like having more children but have, in a sense, a low preference for spending more time with them.
[5] We abstract from parents' infrequent and short-time purchases of other peoples time to take care of children, e.g., baby sitting.
[6] If the child care time constraint is not binding, $D < c + d$, which occurs if the parent purchases a lot of quality, then the model becomes analogous to the BL-model. Their results are also replicated.
[7] Note that this definition of complementarity, the Pareto-Georgescu criterion, may deviate from the standard definition that the compensated cross elasticity should be positive. See e.g., Samuelson (1974).
[8] We assume that the parent has a time choice between own care and leisure while time in market work is exogenous. A natural extension of this analysis is to instead allow for an own care – labor supply choice.

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Modelling household income dynamics

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Abstract. This paper is about income and poverty dynamics and their socio-economic correlates. The first half of the paper aims to establish some of the salient facts for Britain, applying the pioneering methods of Bane and Ellwood (1986). Important for poverty dynamics are changes in labour earnings from persons other than the household head, changes in non-labour income (including benefits), and changes in household composition, in addition to changes in the heads' labour earnings. The second half of the paper is a review and critique of the multivariate modelling frameworks which might be used to explain and forecast these salient facts for Britain or elsewhere.

JEL classification: C23, C41, D31, I32

Key words: Income dynamics, poverty dynamics, income distribution

1. Introduction

This lecture is about the longitudinal dynamics of personal economic well-being, i.e. the patterns of change, from one year to the next, of needs-adjusted

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Table 1. Cross-section perspective on the British income distribution 1991–1996

	1991	1992	1993	1994	1995	1996
Mean	259	269	272	274	288	290
Gini coefficient	0.31	0.31	0.31	0.31	0.32	0.32
Percentage below half contemporary mean	17.8	16.6	17.3	16.6	17.1	16.4
Percentage below half 1991 mean	17.8	15.3	15.1	14.1	12.4	12.0
Number of persons (unweighted)	11634	11001	10473	10476	10119	10511

Source: BHPS waves 1–6, data weighted using cross-section enumerated individual weights. Income is needs-adjusted household net income per person in January 1997 pounds per week (see Table 3 for details).

household net income for each person in the population. My aims are, first, to establish some of the salient facts for Britain about income dynamics in general and poverty dynamics in particular, and their socio-economic correlates, drawing on new evidence for the 1990s, and second, to review the multivariate modelling frameworks which might be used to explain and forecast these patterns for Britain and those for other countries. I offer a guide to the progress made and to the questions outstanding, and issue some challenges for future research in the hope that it will lead others to work in the area and take it forward. There is much to be done.

We know much less about income mobility and poverty dynamics of income than we do about secular trends in inequality and poverty.[1] Since I am going to ask you to take a longitudinal perspective rather than a (time series of) cross-sections one, I shall take a minute first to demonstrate that there is a substantial amount of income mobility to be explained, and that this longitudinal flux exists even when there is cross-sectional stability in income inequality.

Table 1 provides a standard cross-sectional perspective on changes in the distribution of needs-adjusted household income in Britain during the 1990s, derived from the British Household Panel Survey. (The data set and definitions are discussed in more detail later.) Over this period, average income rose by about 12% and reflecting this the fraction of the population with incomes below half 1991 average income fell. Meanwhile, however, income inequality and the proportion of persons with incomes below half contemporary mean income hardly changed at all, a sharp contrast with the large increases during 1980s (Jenkins 1996).

1.1. Longitudinal flux coexists with cross-sectional stability in income inequality

This picture of stability disappears if one examines year-to-year income mobility instead. Table 2 shows average annual transition rates between six income groups where group membership depends on the size of a person's needs-adjusted household income relative to five fixed real income thresholds. The pattern revealed is one of much mobility, but most of it short-range (Jarvis and Jenkins 1998). Fewer than 60% of the persons in any one group remain in the same group from one year to next (with the exception of the

Table 2. Longitudinal perspective on the British income distribution 1991–1996: Outflow rates (%) from wave $t-1$ income group origins to wave t income group destinations

Income group*, wave $t-1$	Income group*, wave t <0.5	0.5–0.75	0.75–1.0	1.0–1.25	1.25–1.5	≥1.5	All	(col. %)
<0.5	54	30	9	4	2	2	100	(13)
0.5–0.75	15	56	21	5	1	2	100	(22)
0.75–1.0	5	19	48	20	5	3	100	(21)
1.0–1.25	3	6	20	44	20	7	100	(16)
1.25–1.5	2	3	8	25	35	27	100	(10)
≥1.5	1	2	4	6	12	75	100	(18)
All	12	22	20	17	10	19	100	(100)

* Income is needs-adjusted household net income per person in January 1997 pounds per week (see Table 3 for details). Persons classified into income groups according to the size of their income relative to fixed real income cut-offs equal to 0.5, 0.75, 1.0, 1.25, and 1.5 times mean Wave 1 income = £259 per week. Transition rates are average rates from pooled BHPS data, waves 1–6, subsample of 6821 persons present at each wave.

richest group, for which the figure is 75%), but the vast majority of those who move end up in the adjacent income group. If half 1991 average income is taken as the poverty line, then of those poor who are poor in one year, almost one half are not poor the following year. (There is a correspondingly high inflow into poverty as well – not shown.)

The amount of mobility can be summarised by the extent to which longitudinal averaging of each person's income reduces the degree of measured inequality (Shorrocks 1978). The averaging smoothes out variability due to income mobility. If each person's income is averaged over the full six year period, then the Gini coefficient falls to about 0.27, some three percentage points lower than the Gini for incomes in 1991. This is approximately equal to the equalising impact of direct taxation (more precisely, it is the difference between the Gini coefficients for gross and disposable income for a given year). Put another way, 'permanent' (six year) inequality is about 88% of (averaged) one year cross-section inequality.[2]

Income mobility also means that the proportion of the population who are touched by poverty over the six-year period is substantially larger than the proportion who are poor in any one year, almost twice as much, in fact, if half 1991 average income is taken as the poverty line (32% compared with 18%: see Table 1). Almost one-fifth (19%) of the BHPS sample were poor at least twice in the six year period. Just under 2% were poor at all six interviews, and about two-thirds were not poor at any of the six.

In sum, the data show that income mobility is a significant empirical phenomenon. The challenge for us is to unravel its causes, separating out the role of various systematic factors from transitory variations and measurement error.

1.2. Motivation

There are several reasons why this task is interesting and important. First, income and poverty dynamics have intrinsic social relevance and policy significance. The extent of mobility and poverty persistence are important social indicators to be placed alongside information about the income distribution at a point in time. For example a former British Minister of Social Security, Peter Lilley, recently discounted the rising incidence of low income during the 1980s with reference to new evidence about longitudinal income mobility:

Social mobility is considerable. Discussion about poverty is often based on the assumption that figures for households on low incomes describe a static group of people trapped in poverty unable to escape and getting poorer. However, this picture has been blown apart by recent studies. They show that the people in the lowest income category are not the same individuals as were in last year, still less fifteen years ago. (P. Lilley, speech in Southwark Cathedral on 13 June 1996, quoted by Hills 1998, p. 52.)

Regardless of whether the conclusions Mr Lilley drew from the evidence were correct (they are debatable), my point is that the evidence is seen as important by many. That it is of interest to a wide range of people is underlined by the fact that mobility findings by Sarah Jarvis and myself (Jarvis and Jenkins 1996) were reported not only in *The Financial Times*, but also on the front page of *Socialist Worker*.

Moreover, longitudinal analysis is an essential ingredient in policy formulation. Researchers in the US and UK have long drawn attention to the differences between the poverty experience of the population over a period of time and the poverty at a one particular time, and emphasised that the design of anti-poverty policy measures should depend on whether poverty is a short-duration event which most people experience at one time or a long-duration event concentrated amongst particular identifiable groups in the population.[3] Indeed a dynamic perspective leads to different anti-poverty strategies, as David Ellwood, a leading researcher recruited as welfare reform advisor by President Clinton, has pointed out:

[D]ynamic analysis gets us closer to treating causes, where static analysis often leads us towards treating symptoms.... If, for example, we ask who are the poor today, we are led to questions about the socioeconomic identity of the existing poverty population. Looking to policy, we then typically emphasise income supplementation strategies. The obvious static solution to poverty is to give the poor more money. If instead, we ask what leads people into poverty, we are drawn to events and structures, and our focus shifts to looking for ways to ensure people escape poverty. (Ellwood 1998, p. 49.)

The New Deal policies for the unemployed and lone parents which have been introduced in the UK by Tony Blair's Labour government are an example of this change in focus.

In this lecture I take it for granted that we are concerned with doing something about reducing poverty by raising exit rates and lowering entry rates. To do this we need empirical models in order to engage with and influence policy-makers and their advisers. Achieving this policy relevance might conflict with the imperative of producing papers rated highly by our fellow academics, a trade-off I return to later.

A second reason for discussing income and poverty dynamics is that virtually all the widely-varying interests and disciplinary affiliations of the ESPE membership are relevant. To study income dynamics one must draw on aspects of household and labour economics, economic demography, public economics, and econometrics and statistics. There is something for almost everyone.

A third reason for the topic is that, even for those who have little interest in engaging with 'real world' of policy, there are many academic challenges raised by income dynamics, for both theoretical and applied researchers. I shall be emphasising the roles played by changes in labour earnings from persons other than the household head, changes in non-labour income (including benefits), and by changes in household composition generally, whereas our current analytical frameworks for income dynamics are best suited to characterising the employment earnings dynamics of prime-age male household heads. There is much work yet to be done on developing more comprehensive models.

Income and poverty dynamics are of interest, fourth, because remarkably little research has been done on them. Although there has been much work on the dynamics of specific income sources for particular population subgroups (wage dynamics and benefit dynamics especially), this has not been matched by analysis of total (needs-adjusted) income for the whole population.[4] Most of the work which has been done refers to the USA, and many of my themes have been inspired by US researchers. However I make little apology for emphasising their points anew, especially since even in the US there has been surprisingly little research and, in any case, it is important to investigate whether US findings are applicable in Europe. My British examples, to follow, are a contribution to that task.

The availability of suitable longitudinal data has been a constraint on European research, but new opportunities are opening up. This is the fifth reason for addressing my topic. Major household panel surveys began in 1984 in Germany, the Netherlands and Sweden, and in a number of other countries subsequently. These surveys are becoming increasing useful as they mature. One of the most important developments in European data is the European Community Household Panel, established in 1995, which aims to provide comparable panel data for more than ten European Union member countries.[5] I shall be illustrating my arguments with analysis of data from interview waves 1–6 of the British Household Panel Survey (covering 1991–1996), of which more below.

1.3. Outline

The remainder of the lecture is organised as follows. In Sect. 2, I offer a working definition of personal economic well-being and its constituent components, and then use this to raise some methodological issues and organise my discussion of the socio-economic correlates of income and poverty dynamics. I also provide some introductory summary statistics. In Sect. 3 I establish some salient facts about British poverty dynamics and its socio-economic correlates. Arnold Zellner's (1992) Presidential Address to the American Statistical Association enjoined us to always ensure we 'GET THE FACTS', and that is my aim. I apply the pioneering methods of Bane and Ellwood (1986). Although I shall be examining Britain in the first half of the 1990s rather than the US in the 1970s, and utilising six years of data rather than twelve, my findings echo theirs. I shall demonstrate that analysis which concentrates on the earnings dynamics of continuously-working households is likely to miss a great deal of the dynamics of poverty for the population as whole.

I turn from description to modelling, from cross-tabulations to multivariate regression, in Sect. 4. I shall not present any of my own estimates;

rather, I shall present an overview of the multivariate approaches which have been used in the literature, and attempt to clarify their strengths and weaknesses. This is a manifesto for a future research programme for myself and – I hope – others. The final section contains some brief concluding comments. Overall my aim is to draw attention to issues and stimulate interest in them, rather than to present particular solutions or findings.

2. A definition of economic well-being to organise the analysis

Suppose economic well-being for each person in each household $h = 1, 2, \ldots, H$, at each time period $t = 1, 2, \ldots, T$, is summarised by their 'personal income-equivalent' (PI), which is the needs-adjusted household net income of the household to which s/he belongs, Π_t^h:[6]

$$\Pi_t^h = \frac{\sum_{j=1}^{n^h} \sum_{k=1}^{K} x_{jkt}^h}{m(a^h, n^h)} \tag{1}$$

The numerator is a double summation: over all persons in the household $j = 1, 2, \ldots, n^h$, and over each money income source x_{jkt}^h, $k = 1, 2, \ldots, K$, where these include net transfer income from the government (benefits less taxes). The denominator is a household equivalence scale factor which depends on household size n^h and household composition and other characteristics summarised by the vector a^h.

This definition prompts several observations. First, PI depends both on household *money income* and on household *demographic composition.* And therefore changes in PI may arise through changes in money income (of one's own or of other household members, via the numerator), or changes in household composition (via the numerator and the denominator), or both. To put things another way, if everyone lived alone, then there would be a direct link between one's labour market status and one's economic well-being. However the majority of the population live together in household units with the opportunity to pool and share resources. More persons means (potentially) more income earners, for example from labour earnings or from social security benefits. But more persons also means that a given amount of money income is worth less in per capita terms.

This discussion leads naturally to a distinction between *income events* and *demographic events* when examining the correlates of income dynamics. Income events are those associated with changes in different types of income (labour earnings of different household members, investment income, public and private transfers net of taxes, etc.). Demographic events include joining events, such as the arrival of a new baby or partnership formation, and leaving events, such as death of a partner, marital dissolution, or a child leaving home.

My second observation is that the population of interest includes every individual in the population, adults and children, those in work and out of work – this follows directly from the policy relevance constraint stated earlier. Of course different income and demographic events may have different impacts amongst different subgroups within the population. This raises the methodological question of whether it is best to derive the required picture for everyone via separate models for each of the most important subgroups or income components, and then reassemble the material later, or to have an all-encompassing model. Most economists' reaction has been to focus on particular subgroups or income sources (working households and earnings, benefit recipients and welfare benefits), but the overall picture is rarely drawn. This strategy is not defensible on the grounds that income mobility for non-working households is negligible relative to that for working households, as I show below (also see Jarvis and Jenkins 1998). And it is important to incorporate groups such as the elderly and lone parents into the analysis, since they have notably high risks of poverty at a point in time and attract much policy attention. I return later to the 'unit of analysis' issue.

My third observation concerns the question of what dimension(s) of income dynamics one should focus on, supposing we have available a sequence of observations on PI for a sample of individuals over a number of years. The first important distinction to be made here is between income dynamics and poverty dynamics – a focus on mobility throughout the whole distribution of income (changes in a continuous variable) or simply transitions above or below some low income cut-off (movements between discrete states).[7] There is no right answer here of course. I simply note that most research to date on household income has focused on poverty dynamics, most probably reflecting greater social and policy concern about low income transitions rather than income mobility in general. This paper will focus on poverty dynamics as well.

The second important distinction is between analysis in which the dependent variable is the longitudinal sequence as a whole and analysis of changes within the sequence. Much of the poverty dynamics literature, following Bane and Ellwood (1986) takes a spell-based perspective, in which the focus is on *consecutive* observations within a given state (single spell 'poverty duration' analysis). By contrast several authors have made a good case for distinguishing between different longitudinal *patterns* of poverty experience, thus incorporating incidence, prevalence and spell repetition elements. Early US studies of poverty dynamics defined the persistently poor as those experiencing more than some large number of years of poverty within some observation window (e.g. 8 years out of 10).[8] This approach is open to criticism because it fails to take account of left- and right-censoring of poverty spells at the boundaries of the observation window. However the approach is valid when the observation window covers a complete life stage such as 'childhood': see for example Ashworth et al. (1994) who differentiate between transient, persistent, permanent, occasional, recurrent, and chronic child poverty patterns. A different, but related, tradition distinguishes between observed and chronic poverty where the latter is defined to be when a person's longer-period income (derived by some type of longitudinal averaging) falls below the poverty line.[9] Obviously it would be nice to be able to build up descriptions of whole sequences of poverty 'experience' from characterisations of its constituent components, and there has been some research doing this, as I discuss later. My empirical illustration takes a spell based approach, following Bane and Ellwood (1986), in part because one needs relatively mature panel surveys to differentiate sequence patterns (only 6 waves of BHPS data were available at the time of writing). But following Stevens (1994, 1995, 1999), I provide some description of poverty re-entry rates in addition to poverty exit rates.

My final observation is that PI is measured here in terms of income rather than consumption, reflecting that fact that all existing annual household panel surveys contain measures of income rather than consumption expenditures. This is a potential problem because temporarily high or low income values may not reflect a person's true economic well-being because borrowing or saving allows consumption smoothing. One must therefore take transitory income variations into account.

2.1. *The data set (British Household Panel Survey waves 1–6) and definitions used in the paper*

Before turning to these illustrations, I need to be more specific about my data set, the British Household Panel Survey waves 1–6 (1991–1996), and the precise definitions of variables such as PI. The first wave of the British Household Panel Survey (BHPS) was designed as a nationally representative sample of the population of Great Britain living in private households in 1991. Original sample respondents have been followed and they, and their co-residents, interviewed at approximately one year intervals subsequently.[10] Children of sample members begin to be interviewed as sample members in their own right when they reach age 16. Most of my analysis is based on an unbalanced panel subsample of more than 10,000 persons (adults and children) in complete respondent households for all waves for which they are in the panel.

All analyses of income distribution, whether cross-sectional or longitudinal, have to make assumptions about the definition of PI (the components of money income and the equivalence scale), and the income accounting unit and measurement period. The choices made for this paper, summarised in Table 3, are a conventional set of assumptions, at least in the context of British research, and match those used to derive the official British low income statistics (Department of Social Security 1997).

The definitions are somewhat different from those used in much income dynamics research based on e.g. the US PSID. In that literature, a pre-tax post-transfer (rather than post-tax post-transfer) income definition is more common, and the sharing unit is typically the family (a single person or persons related by blood or marriage living together) rather than the household. The British McClements equivalence scale corresponds to a Buhmann et al. (1988) parametric equivalence scale with household size elasticity of about 0.6–0.7 (Coulter et al. 1992; Jenkins and Cowell 1994), whereas the needs relativities most commonly used in the US (those implicit in the official poverty line) correspond to an household size elasticity of 0.56 (Burkhauser et al. 1996).

Perhaps the major difference between my standard British definition and the standard US one is the time period over which money incomes are measured. In Britain it is (broadly speaking) the month prior to the interview, in the US, the year. An annual income definition is often judged to be superior on the grounds that a longer period measure is less likely to reflect transitory variations. In the absence of annual income data, I am forced to assume that income and poverty status round about the time of the interview proxy annual income and poverty status. This should be kept in mind when, for brevity's sake, I refer later in the paper to movements in and out of poverty from one year to the next.

Table 3. Methodological issues and definitions used in this paper

Issues	Definitions and assumptions used in this paper
Data set, subsample, and coverage	British Household Panel Survey, waves 1–6 (1991–6). Analysis in Tables 4–12 based on all persons in each complete respondent household (i.e. those for whom net income estimates available) while in panel.
Income sources included in definition of household income	Net household income = labour earnings from employment and self-employment + returns from savings and investment + returns from private and occupational pensions + all public cash transfers (cash benefits) + private transfers – national income taxes and social security contributions – local taxes
Time period over which income measured	'Usual' employment earnings; most recent pay period preceding interview for other income sources (except investment income – annual). Taxes and social security contributions estimated on a pro rata basis All income sources converted to a pounds per week basis and expressed in January 1997 prices.
Income sharing and the income unit	Equal pooling and sharing of income within households, where a household is one person living alone or a group who either share living accommodation or one meal a day and who have the address as their only or main residence. The household head is defined to be the owner or renter of the property, and where the ownership or tenancy is jointly held, the eldest owner or renter is defined to be the head.
Equivalence scale used for 'needs adjustment' of money incomes	'McClements Before Housing Costs' scale. Scale rate for childless married couple = 1.0; single householder 0.61; rates also vary by children's age (see Department of Social Security, 1997, for details).
Poverty definition	Needs-adjusted household net income of person's household (PI) is less than half average wave 1 (1991) needs-adjusted net income, i.e. the poverty line equals £129.74 per week.

There is one particular advantage of the British definition for this paper, however, in which household demographic change plays an important role. In both the PSID and the BHPS household composition is measured at the time of the interview. Thus in the British data the contributions to numerator and denominator elements of PI are more likely to be consistent with each other (each depends on household composition, which may change over the year).

To study poverty dynamics we need a poverty line. By contrast with the US, in Britain there is no official poverty line but half average income is the most commonly used cut-off used in public discourse, and half wave 1 (1991) average income the most commonly used line for analyses based on the BHPS. I use it too. Its level corresponds, broadly speaking, to social assistance benefit levels (Jarvis and Jenkins 1997). Given the economic recovery over the period 1991–1996, this 'absolute' poverty line, fixed in real income terms, implies a declining cross-sectional poverty rate: see Table 1. (Preliminary analysis suggests that changing the generosity of the poverty line does not change the general tenor of my conclusions.) The Bane and Ellwood (1986) analysis referred to later uses the official US poverty line, and covers the 1970s, a period during which the poverty rate was 11%–12% (Triest 1998) and the poverty line was about 40% of median income.

2.2. Longitudinal summary statistics about PI and its components

Before turning to analysis of poverty dynamics analysis itself, I provide some longitudinal summary statistics about PI and changes in its constituent numerator (money income) and denominator (demographic) elements. Table 4 provides information about the share of each of the nine different income components in household money income packages. The statistics are based on longitudinally averaged incomes for each person. The first column shows the average across all persons of these six-wave averages for each person; the remainder of the columns show averages amongst subgroups of persons classified according to their wave 1 household type (which may of course subsequently change). The sum of the shares of each income component in total household money income is 100%, but observe that some shares are negative: this is because taxes are treated as deductions from income.

Average household net money income amongst all persons is £350 per week, compared with an average PI of £280 per week and average household equivalence scale rate of 1.27. Table 4 column 1 shows that labour earnings are by far the largest income component in household income packages. Notice the importance of labour earnings of household members other than the household head: their combined share is some 45% compared to 60% for head's earnings. The two other most important elements of household income packages are income taxes, with a share of −27%, and benefit income with a share of 15%.

The other columns of Table 4 reveal the variation in income packaging across household types. For example, amongst those in elderly households at wave 1, benefit and pension income are the predominant income sources, as expected. Amongst persons in households with the head aged less than 60 years at wave 1, labour market earnings are, of course, much more important, but observe that it is other labour earnings, and not only the head's earnings which are important. For example amongst those in non-elderly childless couple households at wave 1, the share of spouse's plus others' earnings is as large as the head's share.[11]

Although Table 4 gives us some clues about which income sources are likely to be most relevant to explaining the dynamics of PI, it is not informative about longitudinal variability in incomes. Table 5 provides two types of summary information about this, for all persons and broken down by wave 1 household type. The first type of information concerns longitudinal variability itself, here characterised using the coefficients of variation for PI, money income and household composition. The second type of information summarises the contribution of each income component to the total variability of each person's household (money) income package. The variability contribution of each income component depends on the component's share in total income, its own longitudinal variability, and its covariance with other income sources.[12] The statistics have the same form as the 'β coefficients' used by finance economists to summarise the contribution of a stock to the riskiness of a stock portfolio.

Table 5 shows, first, that longitudinal variability in PI is only slightly less than longitudinal variability in household net money income. Second, longitudinal variation in both variables is quite similar across all (wave 1) household types, despite their very different income packages.

The middle rows of the table display the β coefficient estimates. As it

Table 4. Six-wave-average incomes and their composition, by person's wave 1 (1991) household type

	All persons	Head aged 60+ (at wave 1)		Head aged <60 (at wave 1)				Other
		Single	Couple	Single	Couple, no kids*	Couple and kid(s)	Lone Parent	
Needs-adjusted household income, PI (6-wave mean, £ per week)	280	201	252	359	367	262	197	279
Household net income (6-wave mean, £ per week)	350	125	262	253	421	405	243	309
Income source as % of household net income:								
Head's labour earnings	59.7	6.2	10.0	100.6	58.8	71.0	40.9	45.1
Spouse's labour earnings	33.9	0.0	10.7	13.8	45.5	39.0	11.4	10.2
Other labour earnings	10.5	1.9	6.6	2.4	12.1	8.8	13.4	39.5
Investment income	4.8	10.5	13.0	6.0	5.3	3.1	2.6	5.2
Private & occupational pension income	5.0	23.1	30.6	2.7	5.0	0.6	0.8	4.2
Benefit income	14.8	65.2	38.3	10.3	6.0	10.4	41.7	19.0
Private transfer income	1.4	0.1	0.1	0.9	0.7	1.2	6.7	4.6
Income taxes	−27.0	−1.5	−5.2	−33.3	−30.5	−31.0	−14.0	−23.9
Local taxes	−3.3	−6.2	−4.6	−3.3	−3.0	−3.0	−3.7	−3.9
Household size: 6-wave mean	3.08	1.02	2.11	1.24	2.45	4.17	3.36	2.28
Equivalence scale rate (McClements): 6-wave mean	1.27	0.62	1.05	0.70	1.16	1.56	1.26	1.13
Number of persons	6821	484	770	329	1410	3063	441	324
(As percentage of all persons)	(100.0)	(7.1)	(11.3)	(4.8)	(20.7)	(44.9)	(6.5)	(4.8)

BHPS subsample is all persons present all 6 waves. Income, income components, household size and household equivalence scale rate longitudinally averaged for each person, and then averaged across persons by subgroup.
* Children ('kids') are defined as aged 0–16 years.

Table 5. Longitudinal variability of income and household size, and the proportionate contribution of income components to longitudinal income variability, by person's wave 1 household type

	All persons	Head aged 60+ (at wave 1)		Head aged <60 (at wave 1)				Other
		Single	Couple	Single	Couple, no kids	Couple and kid(s)	Lone Parent	
CV (needs-adjusted household income, PI)*	0.25	0.24	0.23	0.27	0.25	0.25	0.28	0.28
CV (household net income)*	0.27	0.25	0.25	0.30	0.27	0.26	0.33	0.33
*Proportionate contribution of income component to longitudinal income variability ('β coefficient')***								
Head's labour earnings	0.49	0.06	0.13	0.81	0.55	0.64	0.33	0.22
Spouse's labour earnings	0.31	0.00	0.14	0.23	0.42	0.41	0.21	0.11
Other labour earnings	0.20	0.01	0.11	0.03	0.25	0.20	0.26	0.63
Investment income	0.09	0.22	0.23	0.09	0.09	0.04	0.04	0.09
Private & occupational pension income	0.04	0.17	0.15	0.01	0.00	0.01	0.01	0.05
Benefit income	0.13	0.52	0.28	0.14	0.04	0.05	0.29	0.10
Private transfer income	0.01	0.01	0.02	0.01	0.01	0.00	0.04	0.07
Income taxes	−0.28	−0.01	−0.07	−0.31	−0.35	−0.36	−0.18	−0.25
Local taxes	−0.00	0.01	0.01	−0.01	−0.00	0.01	−0.01	−0.02
CV (household size)*	0.09	0.01	0.05	0.13	0.13	0.08	0.13	0.22
CV (McClements equivalence scale rate)*	0.09	0.01	0.05	0.09	0.10	0.09	0.14	0.18
Number of persons	6821	484	770	329	1410	3063	441	324
(As percentage of all persons)	(100.0)	(7.1)	(11.3)	(4.8)	(20.7)	(44.9)	(6.5)	(4.8)

BHPS subsample is all persons present all 6 waves.

* Coefficients of variation for income, household size and household equivalence scale rate calculated longitudinally for each person, and then averaged across persons by subgroup.

** The slope coefficient from a six-observation regression, for each person, of each income component on total net income, averaged across persons by subgroup (see text).

happens, patterns correspond closely to those shown by the income shares in Table 4. (The most notable exception is the β coefficient for benefit income which is lower than its share.) In other words, the greatest contribution to longitudinal variability in household net money income appears to be the labour earnings of the household head for the majority of households. However the contributions of other labour earnings are also relatively large. Indeed for persons in non-elderly couple households, the combined contribution of secondary earnings is greater than for head's labour earnings.

The final rows of Table 5 suggest that variability in household net money income (the numerator of PI) is greater than variability in household needs summarised by either household size or equivalence scale rate (the denominator of PI). However it is arguable whether the measures are fully comparable given the contrasting metrics for the variables (continuous versus intrinsically discrete). Let us then examine the extent of household demographic change directly: see Table 6.

The most commonly used indicator of demographic change is a change in a person's household head. The top panel of Table 6 shows the cumulative proportion of persons with a change in their household head, by wave, first for all persons, and then broken down by the person's wave 1 household type. Between waves 1 and 2, almost one tenth of all persons had experienced a change in household head, but by wave 6, the figure was more than one fifth. The experience of demographic change varies substantially by household type. The greatest contrast is between single elderly persons (for whom the event is virtually non-existent by definition – it would require a new partnership and change in responsibility for housing costs) to non-elderly 'other' persons, mostly unrelated adults, amongst whom 38% experienced a household change between waves 1 and 6.

When the definition of demographic change is extended to include all types of events in which people either join or leave the household, or both, many more persons are counted as experiencing it. Between waves 1 and 2, almost one fifth experienced some kind of demographic change; by wave 6, that figure had more than doubled, to almost one half (47%).

These statistics demonstrate clearly that the incidence of demographic events is substantial, and therefore cannot be ignored in any study of the correlates of income dynamics. There is an important corollary: if one restricts analysis to persons and households who do not experience compositional change, one will be omitting a significant fraction of the population and introducing a form of selection bias. (See Duncan and Hill 1985, for the authoritative statement of this case.) The results also raise questions about the use of the household as the unit of analysis when estimating life cycle consumption expenditure and saving models.[13]

3. Some salient facts about British poverty dynamics and its socio-economic correlates

The aim of this section is to establish the main socio-economic correlates of transitions into and out of poverty in Britain. I shall first provide some facts about poverty dynamics, and then examine the relative roles played by income and demographic events using methods pioneered by Bane and Ellwood (1986).

Table 6. The cumulative experience of household demographic change, by person's wave 1 household type

		All persons	Head aged 60+ (at wave 1)		Head aged <60 (at wave 1)				Other
			Single	Couple	Single	Couple, no kids*	Couple and kid(s)	Lone Parent	
Percentage with a change in household head, by wave:									
	2	9.8	0.0	8.4	4.1	13.0	10.8	4.1	19.1
	3	15.5	0.2	15.1	6.9	19.7	17.1	8.5	28.8
	4	18.5	0.2	16.1	8.9	23.5	20.2	11.8	35.6
	5	20.6	0.2	18.8	10.6	24.7	22.8	17.6	34.5
	6	22.5	0.8	20.8	12.5	26.8	24.4	19.5	37.8
Percentage with any household demographic change, by wave:									
	2	19.3	0.8	11.4	14.5	26.2	20.2	20.7	33.7
	3	30.2	1.1	19.7	20.3	41.1	31.8	31.8	48.1
	4	38.4	2.7	22.7	24.5	49.9	41.0	49.0	58.4
	5	43.6	2.7	27.2	29.5	54.9	47.9	56.8	55.7
	6	47.4	3.7	39.8	34.4	59.1	51.5	62.6	61.9

BHPS subsample at each wave comprises all persons in complete respondent households present at the given wave and all previous waves. Calculations based on data for 9824 persons (wave 2); 8579 persons (wave 3); 7894 persons (wave 4); 7197 persons (wave 5); 6821 persons (wave 6).

* Children ('kids') are defined as aged 0–16 years.

3.1. Poverty exit rates and re-entry rates

Table 7 shows Kaplan-Meier product-limit estimates of poverty exit rates (and their standard errors) for a cohort of persons starting a poverty spell, together with estimates of the proportions remaining poor after given lengths of time. Table 8 provides similar information, but about re-entry rates to poverty for those people who end a poverty spell.[14]

The tables immediately reveal some of the problems which arise in empirical implementation. First, the amount of information is relatively limited. Because there are only six waves of data, exit rates at long durations cannot be estimated. There is typically little exogenous time series variation in the data, which hinders identification of some effects. And relatively small sub-sample numbers constrain breakdowns by population subgroups.

Second there are potential measurement error issues. My poverty line delineating the states of 'poor' and 'not poor' is arbitrarily defined, like virtually all low income cut offs. It is implausible to treat small income changes – for example from one pound below the line to one pound above or *vice versa* – as a genuine transition out of or into poverty, when it is as likely due to transitory variation or measurement error. To avoid these threshold effects, I count a rise in PI as a poverty exit only if the post-transition PI value is at least 10% higher than the poverty line. Similarly I require PI to fall below 90% of the poverty line to count as a transition into poverty. Adjustments such as these have been implemented in most previous studies: see for example Bane and Ellwood (1986) and Duncan et al. (1993). I have also made a further measurement error adjustment to the data. Preliminary analysis revealed that a non-trivial number of poverty transitions were accounted for by implausible changes in benefit income from one year to the next. Transitions have been censored in these cases.[15]

Consider now the substantive estimates, beginning with poverty exits. By construction (the exclusion of left-censored spells), all persons starting a poverty spell are poor for at least one year. However almost one half (47%) of this cohort leave poverty the following year, and the exit rate falls further to about one third and one fifth for the subsequent two years, raising the issue of duration dependence. The exit rate for the fifth year is not lower still, but higher (0.32), albeit with a larger standard error. (The secular growth in average incomes between 1991–1996, combined with the fixed poverty line, is one potential reason for the rise.)

The exit rates imply a median poverty spell duration for a cohort beginning a spell of between two and three years and after five years, almost four-fifths of an entry cohort would have escaped poverty. Equivalently – more pessimistically and emphasising poverty persistence – about one fifth of the entry cohort are still poor after five years. Without a longer panel, and thence estimates of exit rates at longer durations, we can only speculate about the incidence of very long poverty spells. If we assume that the exit rates were 0.25 for all years after the fifth, then just over one-twentieth (0.06) of those beginning a spell would be poor at least ten years and the average duration about 3.8 years.[16]

This picture of poverty persistence describes the experience of those beginning a poverty spell. But, as many have emphasised, the length of completed poverty spells for those who are currently poor is rather different. Although only a small fraction of people entering poverty have long spells, the stock of

Table 7. Proportion remaining poor, and exit rates from poverty, by duration, for all persons beginning a poverty spell

Number of interviews since start of poverty spell	Total number of persons at risk of poverty exit at start of period	Cumulative proportion remaining poor (%)	(s.e.)	Annual exit rate from poverty	(s.e.)
1	2067	100.0	(–)	0	(–)
2	1514	53.2	(1.3)	0.47	(0.02)
3	436	34.9	(1.5)	0.34	(0.03)
4	181	27.0	(1.6)	0.22	(0.04)
5	44	18.4	(2.2)	0.32	(0.09)

Kaplan-Meier product-limit estimates based on all non-left censored poverty spells, pooled from BHPS waves 1–6. 272 exits have been recoded as censored if needs-adjusted income rises to not more than 10% above poverty line. 375 spells with poverty exits apparently due to benefit income measurement error have been excluded from the calculations (see text for further details).

Table 8. Proportion remaining non-poor, and poverty re-entry rates, by duration, for all persons ending a poverty spell

Number of interviews since start of non-poverty spell	Total number of persons at risk of poverty re-entry at start of period	Cumulative proportion remaining non-poor (%)	(s.e.)	Annual re-entry rate to poverty	(s.e.)
1	2834	100.0	(–)	0	(–)
2	2187	88.9	(0.7)	0.11	(0.01)
3	1159	84.3	(0.8)	0.05	(0.01)
4	616	81.7	(1.0)	0.03	(0.01)
5	277	79.3	(1.3)	0.03	(0.01)

Kaplan-Meier product-limit estimates based on all non-left censored non-poverty spells, pooled from BHPS waves 1–6. 308 re-entries have been recoded as censored if needs-adjusted income falls to not more than 10% below poverty line. 147 poverty re-entries apparently due to benefit income measurement error have been recoded as censored (see text for further details).

poverty is dominated by those with long spells because those with short spells leave. It is straightforward to illustrate this, assuming a no-growth steady-state in which the poverty inflow rate is constant. In this case, and assuming the exit rate is 0.25 for all years after the fifth, then, of those who are poor in a given year, the fraction with a poverty spell length of at least 10 years is almost one-fifth (0.18) and the median completed spell length between 4 and 5 years, and the average completed spell length, 5.9 years.[17]

These calculations underestimate people's total experience of poverty over a given period because they ignore the fact that a significant fraction of people experience multiple spells of poverty. Stevens (1995, 1999), using US PSID data, has shown most effectively that combining information on poverty re-entry rates with poverty exit rates provides much better predictions of poverty experience than does relying on single spell estimates (as above).[18]

Table 8 provides information about poverty re-entry rates for all persons ending a poverty spell (again left-censored spells have been excluded from the calculations). Re-entry rates fall from 0.11 in the second year after leaving poverty to less than one third of that rate after five years, 0.03. (The number of persons 'at risk' at the start of the period is larger than in Table 7 because of the high prevalence of left-censored poverty spells in this short panel.) The re-entry rates imply that about one-fifth of those leaving a poverty spell will have experienced another spell within the subsequent five years. This reiterates the point made by Jarvis and Jenkins, using BHPS waves 1–4, that 'the path out of low income is not a one-way up-escalator:... there is a not insignificant chance of finding oneself on the down escalator to low income again' (1997, p. 131).

The different pictures about persistence provide different impressions about the concentration amongst the poor of receipts of social assistance and other benefits for poverty alleviation. A focus on the poverty stock tells us that the persistently poor receive most of the total resources devoted to poverty alleviation at a point in time. But a focus on flows, both out of and (back) into poverty, reminds us that the number of people who are ever helped by poverty alleviation measures is many more than those currently poor.

This is of course the same message as provided by the US literature. In this connection it is interesting to note that the exit rates shown in Table 7 are broadly similar to the estimates reported by Bane and Ellwood (1986) for the US in the 1970s. On the other hand, the poverty re-entry rates shown in Table 8 are noticeably lower than those reported by Stevens (1994) using PSID data for 1970–1987. Taken at face value this cross-national comparison suggests greater poverty turnover in the US than in Britain. This conclusion must remain tentative however, given the differences in periods covered, definition of income, equivalence scale and the poverty line, and in the population sub-samples examined.

3.2. The definition of income events and demographic events

It is now time to put the information about income and demographic events together with the data about spells. I use the decomposition methods pioneered by Bane and Ellwood (1986) to determine the main events associated with poverty spell endings and beginnings.[19] I will return later to evaluate the advantages and shortcomings of this approach.

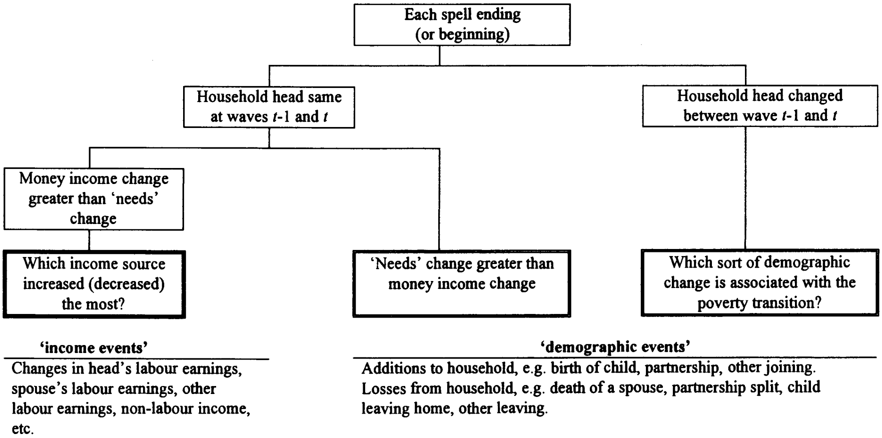

Fig. 1. Classification of 'income events' and 'demographic events' associated with a poverty spell ending (or beginning) between waves $t-1$ and t (after Bane and Ellwood 1986)

The first step in the analysis is to derive a mutually exclusive hierarchical categorisation of event types for each person experiencing a poverty spell ending and for each person experiencing a poverty spell beginning (left-censored spells are now included in the analysis). This procedure is summarised in Fig. 1. In each case, one determines first whether there was a change in household headship concurrently with the poverty transition.[20] Amongst those with a change in household head, one then determines what type of demographic event was involved. Examples include a child leaving the family home and becoming a household head, partnership dissolution where a married woman and her children become a lone parent family, and unrelated adults changing their living arrangements. Amongst those with no change in household head, one checks whether the change in household 'needs' (as summarised by the household equivalence scale rate) is proportionately greater than the concurrent change in household net money income. Examples might include the birth of a child or death of a spouse. All the events identified so far are labelled demographic events. All remaining poverty transitions are classified as income events, and further sub-divided by type. Amongst the persons with an unchanged household head and for whom household income changed by more than 'needs', one determines for spell endings (beginnings) which income component increased (decreased) the most. I distinguish nine types of income event, ranging from a change in household head's labour earnings through to a change (in the opposite direction) in household local tax payments. Let us consider first the correlates of poverty spell endings.

3.3. *The correlates of poverty spell endings*

Table 9 summarises the classification of spell endings by type. (By contrast with Table 7, the analysis includes all spell endings, whether their start is censored or not, subject to the caveat mentioned in the Table notes.) Just over four-fifths (82%) of exit transitions were associated with favourable income events, and just under one-fifth (18%) with demographic events. Changes in

Table 9. Poverty spell ending types

Main event associated with spell ending	Percentage of all spell endings	Cumulative percentage
Rise in money income from:		
Head's labour earnings	33.6	33.6
Spouse's labour earnings	15.5	49.1
Other labour earnings	13.0	62.1
Investment income	5.7	67.8
Private & occupational pension income	4.9	72.7
Benefit income	6.5	79.2
Private transfer income	3.0	82.2
Income taxes (fall)	0.0	82.2
Local taxes (fall)	0.0	82.2
Demographic event:		
'Needs' fall (same household head)	3.7	86.0
Child became head or spouse	1.4	87.4
Spouse became female head	1.7	89.3
Female head became spouse	1.3	90.6
Child of male head became child of female head	1.4	91.9
Child of female head became child of male head	2.2	94.1
Other change (other relatives or unrelated persons)	6.1	100.0
All spell endings	100.0	
Number of spell endings	1684	

Analysis based on all persons with poverty spell endings observed in BHPS waves 1–6 regardless of whether spell beginning censored or not, except that 272 endings for which needs-adjusted income rose to less than 10% above the poverty line, and 532 endings in the benefit income rise category apparently due to benefit income measurement error, have been excluded from the calculations (see text for further details).

labour earnings account for three-quarters of all the income events (62% of all endings). Interestingly, although increases in the earnings of the household head are the most common event, changes in others' labour earnings are almost as prevalent: 29% of all endings compared to 34%.

These statistics for all persons disguise substantial heterogeneity. Table 10 breaks down the events according to each person's household type at the interview prior to the poverty transition (i.e. the last year of the poverty spell). Decomposition detail is constrained by cell size, but even when only a sevenfold household type partition is used and events are aggregated into four main types, there are some clear cut patterns. Amongst elderly households, increases in non-labour income dominate. The incidence of demographic events is above average amongst persons in non-elderly childless households and 'other' households (mostly unrelated adults), and mainly involves others leaving the person's household.

The breakdowns for multi-adult households provide some lessons about the relative importance of increases in the labour earnings of the household head. Amongst married-couple-with-children households, even though the main event is an increase in household head's labour earnings for 45% of all endings, for more than a third (34%) it is increases in others' labour earnings. Results are even more striking for childless couple households for whom the main ending event is a not an increase in the labour earnings of the household head but the labour earnings of others (37% of all endings compared with

Table 10. Poverty spell ending types, by person's household type in last year of poverty spell (column percentages)

Main event associated with spell ending	All	Head aged 60+		Head aged <60				Other
	persons	Single	Couple	Single	Couple, no kids	Couple and kid(s)	Lone Parent	
Household head's labour earnings rose	33.6	7.2*	7.3*	40.3	25.5	44.9	36.2	11.5*
Spouse's or other labour earnings rose	28.5	1.2*	14.6	8.9*	36.7	34.4	29.1	38.5
Non-labour income rose	20.2	84.4	61.4	16.9	15.5	7.3	16.7	13.5*
Demographic event	17.7	7.2*	16.8	33.9	22.4	13.4	18.1	35.9
All spell endings	100.0	100.0	100.0	100.0	100.0	100.0	100.0	100.0
Number of persons	1684	83	191	124	161	819	210	96
(row %)	(100.0)	(4.9)	(11.3)	(7.4)	(9.6)	(48.6)	(12.5)	(5.7)

Notes: as for Table 9.
* Calculation based on fewer than 20 persons.

26%). Others' labour earnings are also important for persons in lone parent households prior to the poverty transition. In this case, the extra earnings will typically be those of other adults sharing the household (but not a spouse).[21] The results suggest that increases in a lone parents' own labour earnings are more important than re-partnering for getting out of poverty – at least in the short-term. It should be remembered that the method of analysis used here focuses on contemporaneous changes; the correlates of long-term sustained escapes from poverty may be different from the short-term correlates.

These results underline the importance for poverty dynamics of changes in 'secondary' labour earnings rather than head's earnings amongst working age households. In each case increases in labour earnings may arise for a variety of reasons: for example, an unemployed person taking a job, or someone already with a job working more hours or being promoted, etc. Amongst the persons for whom a rise in household head's labour earnings was the most important event associated with a poverty spell ending, the household head changed from 'not working' to 'working' in about one-half (51%) of the cases. And amongst the persons for whom a rise in the others' labour earnings was the most important event associated with a poverty spell ending, there was a change from neither the spouse nor others (besides the head) working to at least one earning in one-half (50%) of the cases.

3.4. The correlates of poverty spell beginnings

Consider now the main events associated with poverty spell beginnings. Table 11 displays the breakdown for all persons. One notable finding is that demographic events account for a greater proportion (38%) of the spell beginnings than of spell endings (18%: see Table 9). Income events account for 62% of beginnings (cf. 82% of endings). Although most types of demographic event are relatively more numerous, what is driving the results is the 'new entrants' category, accounting for some 15% of all spell beginnings. These refer to persons who are present in the household currently (but who are not the household head) and who were not present prior to the poverty transition (i.e. at the last wave). Many of these individuals are children born into poverty. Other persons under this heading are new partners of the household head or other adults in a household which is poor when they are present. The figure for this group is an over-estimate because, when constructing the table, I assumed that these persons were not poor prior to joining their current household: some such assumption has to be made because, by definition, the income of their previous household is not observed in the panel. However, even if one took the opposite view, and assumed that they were all previously poor (and thus excluded from the table), the general conclusion about the relative incidence of income and demographic events for the spell beginning case compared to the spell ending one, would not change.

The poverty spell beginnings are broken down in Table 12 by event and household type in the first year of the poverty spell. Broadly speaking, the diversity of patterns by household type is similar to that in Table 10, but with a shift within each household type subgroup towards a higher incidence of demographic events. These events are particularly important amongst people belonging to non-elderly single person or lone parent households and 'other' households (mostly unrelated adults sharing). For the first of these groups,

Table 11. Poverty spell beginning types

Main event associated with spell beginning	Percentage of all spell beginnings	Cumulative percentage
Fall in money income from:		
Head's labour earnings	31.0	31.0
Spouse's labour earnings	11.6	42.5
Other labour earnings	4.3	46.9
Investment income	4.3	51.2
Private & occupational pension income	4.7	55.9
Benefit income	4.1	60.0
Private transfer income	2.3	62.3
Income taxes (rise)	0.0	62.3
Local taxes (rise)	0.0	62.3
Demographic event:		
'Needs' rise (same household head)	4.8	67.1
Child became head or spouse	5.6	72.9
Spouse became female head	2.4	75.1
Female head became spouse	1.0*	76.1
Child of male head became child of female head	1.2*	77.3
Child of female head became child of male head	3.7	81.0
Other change (other relatives or unrelated persons)	5.2	86.2
New entrant to household: baby	5.9	91.4
New entrant to household: partner	3.3	94.7
New entrant to household: other	5.3	100.0
All spell beginnings	100.0	
Number of spell beginnings	1475	

Analysis based on all persons with poverty spell beginnings observed in BHPS waves 1–6, except that 592 beginnings for which needs-adjusted income did not fall to more than 10% below the poverty line, and 375 poverty beginnings in the benefit income fall category apparently due to benefit income measurement error, have been excluded from the calculations (see text for further details).
* Calculation based on fewer than 20 persons.

most of the changes refer to children leaving their parents' household to become heads of their own households. Amongst persons in lone parent households, the most common demographic event is the birth of a child into poverty (this is also true for the couple with children group). Amongst 'other' households, there is a diversity of changes.

Amongst households with above-average incidence of labour earnings events – primarily non-elderly couple households with and without children – earnings falls for secondary earners are less important than earnings rises were for poverty endings. However it remains the case that work transitions are as important as pure earnings changes. Amongst the persons for whom a fall in household head's labour earnings was the most important event associated with a poverty spell beginning, the household head also changed from 'working' to 'not working' in 56% of the cases. And amongst the persons for whom a fall in other labour earnings was the most important event associated with a poverty spell beginning, there was a concurrent change from the spouse or other person besides the head working to no one in this category working in 51% of the cases.

Table 12. Poverty spell beginning types, by person's household type in first year of poverty spell (column percentages)

Main event associated with spell ending	All persons	Head aged 60+ Single	Head aged 60+ Couple	Head aged <60 Single	Head aged <60 Couple, no kids	Head aged <60 Couple and kid(s)	Head aged <60 Lone Parent	Other
Household head's labour earnings fell	31.0	8.9*	29.9	27.1	45.2	39.9	16.3	6.5*
Spouse's or other labour earnings fell	15.9	5.9*	18.9	4.5*	9.6*	20.0	13.1	23.9
Non-labour income fell	15.5	66.3	42.5	10.3*	7.8*	6.0*	17.4	12.0*
Demographic event	37.7	18.9	8.7	58.1	37.4	34.1	53.2	57.6
[of which new entrants]	[14.5]	[2.0*]	[0.0*]	[1.0*]	[14.8*]	[19.1]	[14.7]	[35.9]
All spell endings	100.0	100.0	100.0	100.0	100.0	100.0	100.0	100.0
Number of persons	1475	101	127	155	115	701	184	92
(row %)	(100.0)	(6.9)	(8.6)	(10.5)	(7.8)	(47.5)	(12.5)	(6.2)

Notes: as for Table 11.
* Calculation based on fewer than 20 persons. Children ('kids') defined as aged 0–16 years.

3.5. *Summary: events and poverty transitions*

In sum, even where income dynamics are more closely associated with labour earnings dynamics, we need to recognise that earnings dynamics are often a mixture of the earnings dynamics of several persons, not only the household head, and for each of these persons, a mixture of job dynamics and earnings dynamics for those continuously in work.[22] Moreover there are some households especially the elderly for whom the main events are changes in non-labour income. And the incidence of demographic change as a main event is not insignificant for large numbers of persons in the population. It is also important to account for measurement error.

These findings for Britain echo those of Bane and Ellwood (1986) for the USA. They reported that:

> less than 40 percent of poverty spells begin because of a drop of head's earnings, while 60 percent of the spells end when the heads' earnings increase. Thus, researchers must focus on household formation decisions and on the behavior of secondary family members. (Bane and Ellwood 1986, p. 1.)

My findings have given a greater role to secondary earners and a lesser one to demographic events than Bane and Ellwood. However more substantive comparisons of Bane and Ellwood's findings and my own are necessarily constrained by important differences – e.g. they use a different income definition, equivalence scale and poverty line, and focus on non-elderly households only. Also, Britain in the early 1990s is different from the USA in the 1970s: dual-earner households are more prevalent now than then, in the USA as well as Britain (Gregg and Wadsworth 1996). Moreover I have used the household as the income unit – reflecting British conventions – rather than the narrower concept of the family as they did, so allowing greater scope for household members other than the head and spouse to play a role.

An interesting issue for further research is whether the results generalize to other European countries. Some existing research suggests that they may. For example, Fritzell (1990) found that family composition changes were a major cause of income changes in Sweden as well as the USA.

The general message, notwithstanding various definitional questions (and further checks of the sensitivity of conclusions to choice of poverty line), remains the same. Analysis which concentrates on the earnings dynamics of continuously-working household heads is likely to miss a much of the dynamics of poverty even for working households let alone the population as whole.

The Bane and Ellwood approach, on which I have relied heavily thus far, provides a particularly useful framework for isolating the 'salient facts' about poverty dynamics and its socioeconomic correlates, as well as raising other issues such as measurement error. But this social arithmetic is not modelling. More particularly, first, the approach does not provide a means for simulating future poverty experience. Second, the differentiation between income and demographic events using a mutually exclusive hierarchy does not allow one to unravel the separate effects of events which occur simultaneously. A lone mother may both repartner and take up a job, but the Bane and Ellwood approach attributes importance to just one of these events. Third, and a related point, the approach does not provide a clear cut link with structural models of the labour market and household formation processes which lie behind the poverty transitions and PI changes more generally.

Most applied economists would probably think some form of multivariate regression model is a better approach for addressing these issues. To what extent is this so? To answer this question, we need to consider what sorts of dynamic models there are and we also require a set of evaluation criteria against which to judge them.

4. Multivariate models of income and poverty dynamics

There are four main types of multivariate dynamic model which have been applied in the income and poverty dynamics literature to date:

- longitudinal poverty pattern models,
- transition probability models,
- variance components models, and
- structural models.

I shall discuss the first of these rather briefly and then focus on the others in more detail. The number of applications of these multivariate models to income dynamics is actually very small, at least by comparison with models of the dynamics of wages, welfare benefit receipt, (un)employment, and household formation. Since the models share the same technical structures (which are relatively well-known), I can concentrate on the features of particular relevance to income dynamics.

Before proceeding to the models, I wish to set out three criteria for evaluating them. (These are ideals: inevitably achievement in some dimensions will lead to sacrifices in other dimensions.) I believe models should be practical, fit the past and be able to provide forecasts about the future, and be structural. Let me elaborate.

- *Be practical.* We need empirical models which can provide useful results in reasonable time. Practicality and feasibility are natural goals given the desire for policy relevance I expressed earlier. This is not to dismiss theoretical models – indeed theoretical models addressing all the various dynamic processes would be valuable, though undoubtedly hard to produce (a challenge to theorists!).
- *Fit the past and provide forecasts about future poverty experience.* Here the issues concern whether a model satisfactorily characterises the salient facts about income and poverty dynamics, and not only 'goodness of fit' and other econometric specification tests. At one level, there are questions such as: is the type of model suitable and is the specification of covariates appropriate, given the main patterns of poverty dynamics? For example, if poverty spell repetition is empirically important, is this modelled? Is a model based on the behaviour of heads of households appropriate when behaviour of other household members is also significant? At another level, the issue is whether the full potential of estimated models has been realised: it is insufficient to simply estimate models and discuss the signs of coefficients. Given the policy relevance constraint which I have insisted on, one needs to draw out the implications of the estimates for individuals' future poverty experience under different scenarios. At its simplest, this might be involve simple extrapolations using the fitted model; more sophisticated forecasting may involve complex micro-simulation of poverty experience under different

policy scenarios.[23] The ability to do this depends partly on the extent to which my third criterion is achieved.

- *Be structural.* As economists, we take it for granted that understanding of income and poverty dynamics would be advanced if there were a direct connection between our empirical models and structural models of the underlying dynamic processes of market and household formation. Recognition of these processes immediately raises questions about simultaneity and endogeneity biases in empirical work too. The appropriate balance between structural sophistication and feasibility is an issue I shall return to.

4.1. Longitudinal poverty pattern models

These models differ from the others because the dependent variable is based on complete longitudinal sequence of income for each person (see the discussion in Sect. 1). Examples include models of the probability of being 'chronically poor' where a person is defined as 'chronically poor' if her longitudinally-averaged income falls below the poverty line. A variant on this is a Tobit regression model of individuals' 'chronic poverty gaps', or some function of the gaps, where these gaps take into account the amount by which longitudinally-averaged income falls below the poverty line. Regressors include a variety of personal and household characteristics, but the nature of the model requires all covariates to be either fixed (for example sex, race), or else fixed at their values at the start of the sequence. See for example, Jalan and Ravallion (1997) and Hill and Jenkins (1999), who use the models to compare the characteristics of persons who are 'chronically poor' and those who are counted as poor using standard cross-section poverty indices. The results are relevant to assessing the targeting of poverty alleviation measures.

The models have a rather different focus of these models from the others I discuss: they do not typically look at income and poverty dynamics per se. Instead longitudinal data is used to derive (fixed) measures of 'permanent' income or 'chronic' poverty. Given this different focus, I shall not discuss them further. The distinction between transitory and 'permanent' differences is an important one nonetheless, and a fundamental part of the income variable components models discussed shortly.

4.2. Transition probability models

This class of models is perhaps the one which most applied economists would immediately think of. The most commonly estimated models are of poverty *exit* transition probabilities, of the form

$$\text{prob}(\text{person } i \text{ is not poor in year } t \,|\, \text{person } i \text{ is poor in year } t-1; Z_i, X_{it}, \theta) \tag{2}$$

but there are also models of poverty *entry*, or *re-entry* probabilities, of the form

$$\text{prob}(\text{person } i \text{ is poor in year } t \,|\, \text{person } i \text{ is not poor in year } t-1; Z_i, X_{it}, \theta) \tag{3}$$

where Z_i is a vector of fixed covariates, X_{it} is a vector of time-varying covariates (which may include duration), and θ is a vector of parameters. The covariates might also include lagged values of covariates (e.g. X_{it-1}), and thence events (ΔX_{it}). These are of course examples of event history (or hazard rate or duration) models. From the fitted conditional probabilities and thence the survivor function, one can derive the predicted spell length distribution for each person with a given set of characteristics. These are of course direct extensions, with covariates, of the Kaplan-Meier transition model estimates in Tables 7 and 8.

Let me focus on a few selected examples. Hill et al. (1998) estimate poverty entry and exit rates for young adults using US PSID data. Poverty exit rate regressions are estimated by Cantó Sanchez (1998) for persons in households responding to the Spanish quarterly labour force panel survey (ECPF), and by Muffels et al. (1990) and van Leeuwen and Pannekoek (1999) using Dutch panel survey data. Schluter (1997) and Van Kerm (1998) estimate poverty exit and re-entry rates using German and Wallonian panel data respectively. Stevens (1995, 1999) estimates exit and re-entry rates, for all persons and by race, using the US PSID. Her model is by far the most sophisticated one econometrically. She allows for unobserved heterogeneity in both exit and re-entry rates using bivariate hazard models, checks for the potential bias introduced by excluding left-censored spells, and provides standard error estimates which account for the occurrence of repeated observations from a given household in her sample. How do these models compare with each other and against my evaluation criteria, and what issues do they raise?

One contrast between the models is the types of covariates used, and income and demographic event variables in particular. Hill et al. (1998) for example, include marriage and divorce and child birth events as time-varying covariates (ΔX_{it}). Canto Sanchez's covariates include (un)employment events and changes in household size and number of income recipients. In both papers, these variables are found to be statistically significant alongside other personal characteristics. Interestingly, Stevens incorporated income and demographic events, defined in a Bane-Ellwood hierarchical fashion, in one version of her model but found relatively few significant effects. Her explanation was that she had controlled 'for several personal and household characteristics in addition to incorporating the event indicators. In particular inclusion of age and female headship controls reduces the estimated effects of many of the beginning and ending event variables' (1995, p. 21).

The different strategies raise several as-yet unresolved issues concerning whether such event variables should be used as covariates. One view is that they should, on the grounds that it facilitates Fit: after all, the earlier descriptive analysis assumes that income and demographic events are important socioeconomic correlates of transitions. Moreover the inclusion of events provides a direct link between poverty transitions and the underlying labour market and household formation processes, an advantage with reference to the Be Structural criterion. There are several contrary positions however. The first is that inclusion of event variables is likely to lead to econometric problems of endogeneity and simultaneity, since the underlying behavioural processes are likely to be jointly determined. To me the real issue is not whether this problem exists (for it undoubtedly does), but how large the biases are which are introduced. We do not know whether the bias in our estimates is 5% or 105%. If the biases are large, there is clearly great scope for econometricians to develop suitable methods to handle them (this is a challenge!)

A second problem is that we may not be able to identify separately the effects of labour market and demographic event variables once one has already controlled for labour market status and demographic status at the point in time (the point raised by Stevens: see above). The issue essentially concerns the incidence of events amongst subgroups of individuals at risk of the event: for example, amongst unemployed poor men included in a poverty exit rate regression in which current employment status is a covariate, how many end poverty spells by getting jobs? Answers presumably depend on the particular data set and context.

A third point is whether inclusion of event variables is akin to 'over-fitting' the model, i.e. whether (after controlling for personal characteristics) experience of specific events is *synonymous* with a poverty transition. For example, it has been put to me that it is hardly surprising or indeed interesting that job loss is strongly associated with entry to poverty. I am not wholly persuaded by this argument, primarily because poverty status depends on the incomes of all household members and the job loss (or other event) of one member need not make household income fall below the critical cut-off with a probability equal to one. The size of that probability does deserve investigation.

A fourth point concerns the construction of the covariates used to encapsulate the impact of each event variable. I have been implicitly assuming that they would be summarised by a binary variable which is equal to one in the interval the event occurred and equal to zero for all other intervals. This supposes that effects are entirely contemporaneous. But what if the impact of an event persists over time (in which case the event variable needs to be 'turned on' for several periods, but how many?), or what if events are anticipated by individuals, leading them to change their behaviour (a question hard to answer without some structural model)?

Further issues arise with the use of event variables, or in fact any time-varying covariate, when assessing model fit and preparing forecasts and simulations. The reason is that one has to specify a longitudinal sequence of values for such variables, conditioned on other characteristics, in order to derive predicted spell length probabilities for each individual.[24] For example consider not only event variables, but also, say, a covariate summarising the age of the youngest child in each interval at risk of a transition, or a covariates summarising the state of the macro-economic environment. The most common practice in event history modelling is simply to use some fixed value for these covariates in a simulation, but this is rather unsatisfactory. Hill et al. (1998) grasp the nettle, and specify a set of temporal sequences of covariate patterns, including assumptions about marriage and birth events and their timing, and predict sequences of poverty risks using these.

Arguably this issue is less of a problem for some sorts of policy advice. Consider for example the case where transition rate models are to be used to target those most at risk of long poverty spells, in the knowledge that the research users will only have at their disposal information about the current characteristics of the relevant population at risk. In this case there is some sense in estimating models with only fixed personal covariates. However I would not wish to push this argument too far: for example if the state of the macro-economy has an significant impact on poverty transition rates, then presumably this information should be incorporated into the predictions, especially if the impact differs between subgroups within the population. Although these covariate specification issues arise in all types of socioeconomic event history modelling – not simply poverty dynamics – it is remarkable how little explicit attention they have received.

Another issue concerns the unit of analysis or subsample used in the model (this point also arises for variance components models). For example, should the models be estimated using a sample of only adults or also include children. There is a reasonable *a priori* case for the former practice as it is the choices of adults which determine household income. However the argument is not decisive – it depends in part on how structural a model one wants. For purely descriptive purposes there is a case for using all observations, adults and children.

My final remarks under this heading are about fitting and prediction of individuals' *total experience* of poverty over a period of time. Much analysis suggests that spell repetition is a significant empirical phenomenon, and yet most models either examine single poverty spells or, if they have estimated models of poverty re-entry probabilities as well, have not combined the model outputs to examine the implications for total poverty experienced over some interval. Stevens (1995, 1999) is a notable exception. One explanation for the omission is that the derivations are technically demanding. It would therefore be interesting to know the pay-off to simpler models from which it is much easier to derive multi-spell predictions.

A two-state first order Markov transition model (see Boskin and Nold 1975; Amemiya 1986, Chapt. 10) is an obvious example of a simpler model. This can be interpreted as discrete-time hazard rate regression model in which, crucially, it is assumed that there is no duration dependence in either the exit or entry transition rate. Because of this, the model is clearly likely to be a misspecification but the assumption brings advantages. The model is straightforward to estimate and there are very simple expressions for expected poverty spell durations, the total proportion of time spent poor, and mean poverty recurrence times for persons with different characteristics. See Van Kerm (1998) for an example.[25] Such models are a prime example of where the appropriate balance between Practicality (in particular the feasibility of simulation) and Goodness of Fit deserves further investigation.

4.3. Income variance components models

Under this heading I group models used to describe the longitudinal covariance structure of PI (rather than poverty itself), but from which results about poverty dynamics can and have been derived. The basic methods were first developed by Lillard and Willis (1978), albeit with an application to men's labour earnings and the persistence of low pay. There have been many subsequent developments of the model and further applications to earnings: see *inter alia* Hause (1980) and Abowd and Card (1989). See also the earlier literature on Galtonian regression models of income dynamics, reviewed by for example Creedy (1985).

The first application of variance components models to household income and poverty dynamics was by Duncan (1983), and later work by Duncan and Rodgers (1991) and Stevens (1995, 1999). A prototypical model is

$$\log(PI_{it}) = Z_i\alpha + X_{it}\beta + \varepsilon_{it} \tag{4}$$

where the error structure of the model takes the form

$$\varepsilon_{it} = \delta_i + v_{it} \tag{5}$$

$$v_{it} = \gamma v_{it-1} + \eta_{it}$$

and where δ_i is a random individual component with mean zero and variance σ_δ^2 and η_{it} is a purely random i.i.d. component with mean zero and variance σ_η^2. The γ is a serial correlation coefficient common to all persons. As Lillard and Willis (1978, p. 988–989) explain, the δ_i terms encapsulate individual heterogeneity in average (log) PI, whereas the serial correlation coefficient γ may be interpreted as reflecting either the effects of random shocks which persist but whose effects deteriorate over time, or serially correlated unobserved individual-specific variables. Later applications have used increasingly more sophisticated specifications of the error structure than this simple one, for example heterogeneity in the component variances across population subgroups, a variety of higher-order autocorrelated moving average (ARIMA) error structures rather than the simple AR(1) one shown here, and calendar-time specific parameters and component variances to account for observed non-stationarity in covariances.[26] There is trade-off between model complexity and data availability: with short panels the number of time-varying parameters which can be estimated is relatively small.

Several implications for poverty dynamics can be derived. For example, assuming the distributions of the error components takes a particular form, such as Normal, one may calculate the proportion of population whose expected (or 'permanent') income is below the poverty line (see Lillard and Willis 1978, Eq. 3.12). One may also calculate the probabilities of observing specific T-year poverty sequences (for example T consecutive years of poverty), though the length of the sequence is currently constrained by the need to calculate the values of a T-variate normal distribution. (Deriving such predictions is harder in these models than in the transition probability models, for which such survivor rates are integral outputs). The more widespread availability of simulation-based estimation methods may alleviate this problem: see the survey of these by Stern (1997). Alternatively, Stevens (1995, 1999) simulates the distribution of PI over a T-year period for her sample using random draws from a (bivariate) Normal distribution calibrated from the variance component estimates, and calculates the distribution of years poor over a given period across the subsample of individuals who are estimated to start a poverty spell. These simulation results are in a form which can be compared with those from her hazard rate models and from tabulations of the actual data.

The appeal of the variance components models is threefold. First, there are attractions to analysing income itself, rather than discretising a continuous variable *ab initio* using a poverty line which is arbitrary and also thereby throwing a lot of information away (Ravallion 1996). In particular, one could have estimates describing whether people moving out of poverty move just above the poverty line or become well-off (cf. the transition probability models).

Second, there is the long standing appeal to economists of the fundamental decomposition of income and income changes into 'permanent' and 'transitory' components: controlling for systematic observed differences (via Z_i and X_{it}), each person is assumed to have some latent level of PI which is permanently fixed (or evolving very slowly), about which there may be temporary variations. As Duncan and Rodgers have pointed out,

> [this] seems reasonable in the face of unfavorable events like short-term unemployment and illness or beneficial events like the overtime hours provided by a temporary increase in labor demand. The strength of this approach lies in its quantification of the components in the overall distribution of poverty. (Duncan and Rodgers 1991, p. 540–541.)

However there are disadvantages too, as they also point out:

> [The approach] does not, however, provide an individual-level measure of permanent poverty, and it may not be as well-suited as other measures, such as spell-based measures, for taking into account the permanent changes in economic status which accompany events like divorce, remarriage, widowhood, or a long-term disability. (Duncan and Rodgers 1991, p. 540–541.)

To their list of permanent changes, one might also add: demographic events such as the birth of a child (perhaps combined with a mother's withdrawal from the labour market), or departure of an adult child from the household, or labour market events such as long-term unemployed household head getting a job, or a mother returning to work. A further disadvantage of the variance components models is that the dynamic processes are assumed to the same for all income groups, rich and poor, which is unlikely.

My view in the light of such comments has been that variance-components models are likely best suited for the phenomena and subgroups for which they were originally developed (such as men's earnings dynamics), rather than household income and poverty dynamics. This is because, first, focusing on homogeneous subgroups (e.g. prime age men) makes the assumption of a simple covariance structure more plausible. Second, modelling a single income source (earnings), the effects of accounting for the combination of different income sources and household composition change are obviated. Moreover there is no explicit or obvious link between the variance component specifications and the underlying labour market and household formation processes (a deficiency with reference to the Be Structural goal). And with an eye to the Practicality goal, it is worth mentioning that there are as yet no software packages available with canned routines for estimation of variance components models in the same way as there is for hazard models (though non-linear GLS modules, which can be used to derive GMM estimates of variance components models, are becoming available.)

Stevens's (1995, 1999) valuable work allows us to assess my opinions about the relative merits of variance components and hazard rate models in a more systematic manner, because she has estimated various versions of both classes of model on the same data set, and compared their predictions of poverty persistence over a fixed time interval for her subsample with direct tabulations from the data. Stevens favours the hazard models rather than the variance components ones, especially for describing dynamics for male-headed households. Her conclusion was that

> these comparisons suggest that the hazard model developed here comes close to replicating the distributions of time in poverty from a relatively simple method of directly tabulating years in poverty from the panel data.... The components-of-variance approach seems to over-estimate time in poverty among male household heads, and may under-estimate time poor among female household heads. While the three methods yield similar results, these discrepancies suggest that attention to the accuracy of variance-components models in predicting dynamic patterns near the

bottom of the income distribution may be an important area for further research. (Stevens 1995, p. 36.)

Further evaluative work of this kind, especially research based on data sets for countries other than the US, would be particularly valuable (though it is not work for the technically faint-hearted).

4.4. Structural models

The final model type on my list refers to disaggregate *structural models*, though I am aware of only one example of these. This is Burgess and Propper's (1998) innovative model describing poverty dynamics amongst a sample of American women aged 20–35 years from the National Longitudinal Survey of Youth (NLSY).

Rather than relating poverty transitions directly to explanatory variables, as in the other approaches described so far, the authors model the underlying dynamic processes which determine earnings – marriage, fertility, and labour force participation – and the earnings associated with the outcomes of these processes. From these, income and poverty status are calculated. The model is estimated separately for black and white women.

More specifically, Burgess and Propper first estimate hazard models for the probability of marital partnership formation and the probability of partnership dissolution, and a bivariate probit model of the probability of having a child during the relevant year and of working in the same year. Second, they model the distribution of labour market earnings separately for each combination of outcomes of the marriage, fertility, and labour force participation choices, controlling for sample selection into each state. Third, a model of husband's earnings is estimated using data about male NLSY respondents. Fourth, each woman's (expected) family income is calculated as a mixture distribution, the sum of the probabilities of being in each {marriage, fertility, work} state times state-conditional earnings where, for the relevant states, earnings includes estimated earnings of a spouse as well as the woman's earnings. Finally, each woman's poverty status for the year is determined by comparing estimated family income to the poverty line for her family type.

Burgess and Propper compare poverty rates fitted using their model with the actual poverty rates. They produce a close fit for the women when aged 25–30, but at ages 20–25, poverty rates are over-predicted by up to 50%. As Burgess and Propper remark, this over-prediction 'probably arises because [they] have not modelled income from adults other than partners, but in the under-25 age group there are a significant minority of individuals who still live in the parental home' (1998, pp. 40–41). Fit is also examined in terms of the fraction of time individuals spend poor over the 15 year period. The authors' assessment is that 'the approach does a reasonable job of separating people likely to spend a long time poor from those likely to be never poor' (1998, p. 41), though there is some systematic bias: the number of women who are frequently poor is under-predicted and the number of women rarely poor is over-predicted. Overall I am not as sanguine as the authors are about the goodness of fit of the model, in part because their sample (described above) is relatively homogeneous compared to those for whom poverty experience is estimated by, say, Stevens (1995, 1999). But, as the authors point out (p. 40), none of their estimation and predictions used the data on poverty status at all; recognising this, the fit is more impressive.

There is no doubt that this model comes far closer to satisfying the Be Structural criterion than any of other modelling approaches considered so far. This is demonstrated by the way in which Burgess and Propper employ it to unravel the causes of poverty in general and the sources of the differences in poverty rates between black and white women in particular. Using counterfactual simulations, they consider the impact of, first, differences in personal characteristics such as years of education and family background, and differences in socioeconomic origins at age 19 (for example, being a non-working lone mother versus being a childless single working woman). Education is found to be the variable with the largest impact on poverty rates. Burgess and Propper are able to analyse the mechanism by which this occurs: the effect of having more education comes about more by increasing the likelihood of being in states which are associated with higher incomes (such as being married with children and working) than by increasing earnings per se. Second, Burgess and Propper examine the effects on poverty rates of differences in black and white women's behaviour, as summarised by the estimated coefficients in the various behavioural process equations. These are found to have make a major contribution to inter-racial differences in poverty compared to differences in socio-economic origins at age 19 and family background: 'While ... all transition rates matter, rates of marriage appear to the single most important factor, as marriage gives access to another income stream' (1998, p. 50). Differences in short-term and long-term impacts are also revealed.

This sort of detailed unravelling of causes plus simulation is not possible within the reduced form models of dynamics discussed earlier. And observe too that the issue for those models about having to specify longitudinal sequences for time-varying covariates when doing simulations also does not arise: the relevant values are generated within the system.

Are structural models the future for income and poverty dynamics analysis then? I believe that they are an exciting and innovative approach, and strongly support further development of them. Several directions for this suggest themselves to me. One of the weaker links in the Burgess-Propper model is the treatment of other adults in the household and non-labour income. Their model of husbands' earnings is less developed than for women (in part reflecting the nature of the NLSY data), and they simply ignore all other income in household (which largely explains the worse fit for the 20–25 year olds). For the reasons given in their paper, omission of income from investments and savings and benefits such as AFDC is not likely to be an important source of bias. But this is not necessarily so in applications of such a model to countries other than the USA (for example Europe, where there are typically more benefits available, including for working households). And in analysis of populations of persons representing a more heterogeneous mixture of life-cycle stages than the NLSY subsample of young women, the incidence of multi-adult households is likely to be much greater.

One might also criticise some of the econometric methods and question the robustness of the identification assumptions. Certainly one valuable service done by Burgess and Propper is to reveal the sorts of assumptions which need to made in order to implement a structural approach, in particular those about correlations of unobservables across processes and across time. (I challenge econometricians to develop more sophisticated estimation methods for simultaneous dynamic processes.) Nonetheless there is a more fundamental question of whether a reliable structural model is a Holy Grail since in many plausible specifications of the underlying equations, arguably 'everything depends on everything else'. Moreover one may make a good case that for many of processes concerned, pure randomness is intrinsic. If one takes on board these various arguments, then structural models and the more sophisticated of the reduced form models discussed earlier are more similar than at first appears. Both strands could draw more on the literature on dynamic microsimulation: see example, Harding (1993, 1996) and references therein.

Overall I believe that despite their attractions, structural models are likely to remain relatively rare, if only for the simple reason that they are immensely complex and very time-consuming to develop. This is another example of the Be Practical criterion in head-on conflict with the Be Structural one.

5. Concluding comments

In his recent Presidential Address to the Royal Economic Society, AB Atkinson stated that his 'principal purpose ... has been to argue that the economic analysis of the distribution is in need of further development before we can hope to give a definitive answer to the questions in which the ordinary person is interested – such as what determines the extent of inequality and why has inequality increased?' (1997, p. 317). My aim, like Atkinson's, has been 'Bringing income distribution analysis in from the cold' (the title of his lecture), though I have taken a rather different, albeit complementary, perspective – the longitudinal one. I have directed attention more at questions such as 'how long do the poor stay poor?' and 'does getting a job get someone and their household out of poverty?'.

These questions are of widespread interest amongst the general public and amongst policy-makers. To answer them we analysts need to develop modelling approaches which better incorporate the impact of changes in individuals' household contexts – changes in the incomes contributed by others (especially 'secondary' labour earnings, but also non-labour income) and changes in household composition. This much is clear from both descriptive decomposition analysis of the sort pioneered by Bane and Ellwood (1986) and applied to British data in Sect. 3, and also from the review of multivariate models of income and poverty dynamics in Sect. 4.

I have drawn attention to a tension between the goals of practicality, fit, and being structural in developing models, and I hope that in future we can improve models according to all three criteria. I acknowledge that this may be difficult. It is easier to focus on dynamics for specific income sources or particular subgroups, and on estimation without simulation. This perhaps explains why the literature on income dynamics is relatively small, even in the USA where household panel data have been available for the longest. Incentive structures in the profession may exacerbate this problem. For academics in today's increasingly 'publish or perish' environment, it may be more rewarding to work on models which focus on particular aspects of the income determination process rather than attempting to characterise the 'big picture' for income (or rather PI) itself. I hope nonetheless that ESPE members and others will take up the many interesting theoretical and empirical challenges which the study of income and poverty dynamics offers and exploit the new longitudinal data sources now becoming available.

Endnotes

[1] See *inter alia* Atkinson et al. (1995) and Gottschalk and Smeeding (1997) for reviews of the substantial increase in inequality during the 1980s in many (but not all) Western developed nations, and Schultz (1998) about trends in world income inequality.

[2] This is the estimated value of a Shorrocks (1978) immobility index, which is equal to the Gini coefficient for cumulated six-wave incomes divided by a weighted average of the Gini coefficients for the income distribution in each of the six years. All calculations cited in this paragraph and the next are based on the balanced subsample of 6821 persons present at all six BHPS waves.

[3] See *inter alia* Duncan et al. (1984), and Bane and Ellwood (1986) for the USA, and Walker with Ashworth (1994) for the UK.

[4] For a survey of earnings mobility research, see Atkinson et al. (1992). For recent UK studies of the dynamics of men's labour earnings, see Dickens (1997), Gosling et al. (1997), Ramos (1997), and Stewart and Swaffield (1997). On the dynamics of social assistance (Income Support) receipt, see Shaw et al. (1996) and Noble et al. (1998).

[5] Cross-nationally comparable household panel data sets are also being constructed: cf. the Syracuse University PSID/GSOEP Equivalent File (Burkhauser et al. 1995), now being extended to include the BHPS, and the PACO project (Schmaus and Riebschläger 1995). Panels derived by record linkage of administrative registers on income, as in the Nordic countries, are promising sources too. By contrast, the possibility of using retrospective survey information to analyse income dynamics is severely limited.

[6] This definition incorporates the almost universally made assumption in the income distribution literature that all incomes are pooled within the income unit (assumed here to be the household) and equally shared out amongst household members. See Jenkins (1991) and Lazear and Michael (1988) for critiques and alternative strategies. For economic models of within-household distribution see *inter alia* Apps and Rees (1996) and Chiappori (1992).

[7] There is also the methodological issue of whether to model poverty or, instead, to model PI and derive the implications for poverty from this. I return to this later.

[8] See for example Duncan et al. (1984) and Hill (1981). Similar methods have been applied to German data by Headey et al. (1991) and Krause (1998), to British data by Jarvis and Jenkins (1997), to Hungarian data by Spéder (1998), and to Wallonia (Belgium) by Van Kerm (1998).

[9] See for example Duncan and Rodgers (1991), Rodgers and Rodgers (1992), Jalan and Ravallion (1997), and Hill and Jenkins (1999).

[10] For a detailed discussion of BHPS methodology, see Taylor (1994) and Taylor (1998). The derived net income variables are a publicly-available supplement to the main BHPS data set. For a detailed discussion of their creation, see Jarvis and Jenkins (1995, Appendix) and Redmond (1997).

[11] According to my household type definition, couple households may, in principle, include adults in addition to the household head and spouse (if present). A finer partition to distinguish these cases would results in cell sizes which were too small.

[12] More formally, for each person, the β coefficients satisfy the relationship $\Sigma_k \beta_k = 1$, where $\beta_k = \rho_k \sigma_k / \sigma$, ρ_k is the correlation between component k and the person's total net income (summed over six years), σ_k is the longitudinal standard deviation of component k, and σ is the longitudinal standard deviation of total income. For each person, this is the same as the slope coefficient from a six-observation regression of the given income component on total net income. Equivalently, $\beta_k = \rho_k (\mu_k / \mu) \sqrt{I_{2k} / I_2}$, where μ_k and μ are the longitudinal means of component k and total net income, and I_{2k} and I_2 are half the squared longitudinal coefficients of variation of component k and total net income. Hence the remark in the text linking β coefficients with covariances, income shares, and longitudinal variability. Income distribution specialists will recognise the β coefficient as a longitudinal version of Shorrocks's (1982) measure of the proportionate contribution of an income component to total inequality in a cross-section of persons.

[13] Cf. Miles (1997) who estimated dynamic models for households from cross-section data, and Banks, Blundell and Preston (1994) who use pseudo-cohort data.

[14] For more extensive analysis, based on BHPS waves 1–4, see Jarvis and Jenkins (1997).

[15] Reported household benefit income for these persons changed substantially from one year to the next even though there was typically no concurrent change in household composition or work pattern. (I.e. there were poverty spell exits associated with large rises in benefit income and poverty spell entries associated with large falls in benefits. Detailed case-by-case examination revealed no obvious reason for these changes. I suspect they are due to some form of transitory recall measurement error rather than genuine changes arising, say, from changes in benefit take-up or exhaustion. For the present, a straightforward rule of thumb was used to identify these unreliable cases: they were defined to be those persons for whom a change in their household's benefit income was the most important event associated with a poverty transition (as defined below) and where the benefit income change comprised more than 75% of the total household net money income change.

[16] These estimates are not very sensitive to the assumption about the exit rate. If it were 0.3 rather than 0.25, the proportion of those beginning a spell with a spell length of 10 or more years is 0.04 and the mean duration, 3.7 years. If the exit rate were 0.02, the corresponding estimates are 0.07 and 3.9 years.

[17] The calculations use the exit rates in Table 7 and Bane and Ellwood (1986, Eq. 2). Answers are little different with the alternative exit rate assumptions of the previous footnote. The importance of the distinction between the average of all completed spell lengths and the average completed spell length for those in the stock has also been stressed in the unemployment literature: see for example Akerlof and Main (1980).

[18] The importance of spell repetition has been stressed in the welfare benefit dynamics literature: see for example Gottschalk and Moffitt (1994) and Blank and Ruggles (1994), and Shaw et al. (1996).

[19] Other poverty dynamics applications to US data have been by Bane (1986) and Duncan and Rodgers (1988). A seven country cross-national study using similar methods is by Duncan et al. (1993). Bane and Ellwood (1994, Chapt. 2) applied the methods to AFDC benefit dynamics. A pioneering British poverty dynamics study is by Hancock (1985), with notable attention given to potential problems arising from data unreliability. Jarvis and Jenkins (1997) were the first to apply similar methods to BHPS data (for waves 1–4).

[20] See Bane and Ellwood's (1986, p. 9) discussion of the importance of the event and thence justification for giving it special treatment.

[21] Recall that re-partnership of a lone mother concurrent with a poverty exit would be counted as a demographic event, since the new male partner would virtually always be labelled the new household head.

[22] This point about earnings dynamics is also made by Swaffield and Stewart (1997) in their analysis of low pay dynamics using BHPS waves 1–4.

[23] All the comparisons of fitted and actual patterns which I know of are based on within-sample comparisons. It would be interesting to have some out-of-sample predictions and assessments.

[24] This issue arises even if the time-varying covariate is exogenous or deterministic rather than endogenous: the problem is that they are determined outside the system. There is an analogous problem in simulation using macro-economic models, except that the magnitude of the problem is much greater here: conditional predictions have to be made for a large number of individuals than a single economy (or small number of sectors).

[25] Schluter (1997) also estimated such a model, but did not draw out the implications of the parameter estimates for poverty spell durations and spell repetition. Nor did Muffels et al. (1992).

[26] If time-varying covariance parameters are used, they raise the same issues for simulation as time-varying covariates do in the transition probability models.

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Saving behaviour and earnings uncertainty: Evidence from the British Household Panel Survey

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Abstract. For the first time, this paper uses a panel data set, the British Household Panel Survey, to analyse saving behaviour in Britain. One objective is to test the precautionary saving hypothesis, according to which households save to self-insure against uncertainty. Our results show that in accordance with this hypothesis, various measures of uncertainty based on earnings variability have a statistically significant effect on households' saving decisions. Moreover, in accordance with the life cycle model, households save more if they expect their financial situation to deteriorate.

JEL classification: D12, D91, E21

Key words: Precautionary saving, uncertainty, earnings variability

1. Introduction

Saving rates differ significantly across countries. At one extreme, during the period 1984–1993, East Asian and Pacific countries witnessed average saving ratios of 27.6%. At the other extreme, the corresponding rates for Sub-Saharian African countries were of 6.4% (Schmidt-Hebbel and Serven 1997). These strong differences in saving ratios between countries are definitely due to the very particular historical, economical, demographic, and institutional

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characteristics of the countries. However, differences in saving ratios also exist within countries, across different groups of the population. For instance, Banks et al. (1994) used the UK Financial Research Survey for the years 1991–1992 and found a strong concentration of financial wealth at the top of the distribution. They also found that almost half the households in the survey had financial wealth of £455 or less, and that over one-tenth had no financial wealth at all. This indicates that some groups of the UK population have very high saving ratios, while other groups save very little or not at all. It is therefore an interesting exercise to try and analyse the factors that determine households' saving decisions. The aim of this paper is to gain a better understanding of saving behaviour in Britain over time, trying, in particular, to assess how much of this saving occurs for precautionary reasons.

From a theoretical point of view, the model most frequently used to analyse consumption and saving issues is the life-cycle/permanent income model, originated by Modigliani and Brumberg (1954), and Friedman (1957). One prediction of this model is that people save "for a rainy day". Saving should therefore be equal to the expected present value of the future declines in income to maintain a smooth consumption path throughout the life cycle.

The life-cycle/permanent income model, in its certainty-equivalence version, is based on a number of restrictive assumptions, such as the existence of quadratic preferences, the additivity of the utility function over time, and the absence of liquidity constraints. Over the years, the model has been amended in several ways, and the restrictive assumptions have been relaxed (see Browning and Lusardi 1996, for an overview of the literature). Allowing for the presence of uncertainty in the model, in the form of a non-quadratic utility function, was at the heart of the research on the precautionary motive for saving. If the marginal utility of consumption is convex, then increases in uncertainty, which raise the expected variance of consumption, lead to an increase in saving.

The empirical research based on quantifying the importance of the precautionary motive for saving has either focused on equations of wealth, saving, or consumption; or on Euler equations. Some measure of uncertainty was included in these equations, and a test for its significance was performed. The results of this research have been highly inconclusive, with some papers finding a strong precautionary saving motive, and others finding almost no evidence for it (see Browning and Lusardi 1996, for a survey). Most of the studies in this literature have either used cross-sectional data (Guiso et al. 1992; Lusardi 1998 etc.), or time series of repeated cross-sections (Merrigan and Normandin 1996; Banks et al. 1999 etc.). To our knowledge, only Kuehlwein (1991), Dynan (1993), Carroll and Samwick (1997, 1998), and Kazarosian (1997) have used panel data to evaluate whether US households' saving decisions react to several indicators of earnings uncertainty. The wealth equations that Carroll and Samwick (1997, 1998) and Kazarosian (1997) estimated to assess the existence of a precautionary motive for saving were, however, only based on a cross-section of their data. These authors only used the panel dimension of their data set to calculate their proxies for uncertainty and permanent income.

This paper gives an original contribution to the existing literature in the field by using for the first time a panel data set to analyse saving behaviour in Britain. Based on the first eight waves of the British Household Panel Survey (BHPS), corresponding to the period 1991–1998, we construct various measures of earnings uncertainty. As in Lusardi (1998), our first measure is a function of the subjective probability that household heads attribute to losing their job. We then construct three panel-based measures of uncertainty, which focus on the variability of the eight time-series observations on earnings available for each household. These household-specific measures of uncertainty are then used as explanatory variables in our saving equations. Their signs and significance levels will allow us to assess the importance of a precautionary saving motive in Britain.

The remainder of the paper is organised as follows. In Sect. 2, we summarise the existing literature on precautionary saving in the UK. Section 3 describes the BHPS data set and the measures of earnings uncertainty that are used. It also provides some descriptive statistics of households' saving behaviour and earnings uncertainty. In Sect. 4, we present the results of the estimation of our cross-sectional and random-effects Tobit regressions for saving. We find that all our measures of uncertainty significantly affect households' saving decisions, supporting the precautionary saving hypothesis. Moreover, if individuals expect their financial situation to deteriorate, they tend to save more, in accordance with the predictions of the life-cycle model. Section 5 concludes the paper.

2. Existing evidence on precautionary saving in the UK

A number of studies have focused on saving behaviour in the UK to quantify the importance of the precautionary motive. Among these, we can mention Dardanoni (1991), Miles (1997), Merrigan and Normandin (1996), and Banks et al. (1999). All these studies used the Family Expenditure Survey (FES), which is a time-series of repeated cross-sections.

Dardanoni (1991) only based his analysis on the 1984 cross-section of the FES. Within this cross-section, he used the variability of earnings in different occupations as a measure of uncertainty. He regressed average total expenditure in each occupational group on average disposable income and uncertainty. His results suggest that more than 60% of saving in the sample arise for precautionary motives. Miles (1997) used the 1968, 1977, 1983, 1986, and 1990 cross-sections of the FES. He first regressed household disposable income on age and age squared of the household head, and other demographic variables, separately for each of the cross-sections. He then used the fitted values from these regressions as a proxy for permanent income, and the square of the residuals as a measure of uncertainty. Regressing consumption on his proxies for permanent income and uncertainty, he found that, for each cross-section, the latter variable played a statistically significant role in determining consumption. Although Dardanoni (1991) and Miles (1997) both used the FES, their studies are only based on separate cross-sections of the data set.

Merrigan and Normandin (1996) and Banks et al. (1999) went one step further in the sense that they exploited not only the cross-sectional, but also the time-series dimension of their data set. Merrigan and Normandin (1996) estimated a model where expected consumption growth is a function of expected squared consumption growth and demographic variables. A larger expected squared consumption growth reflects greater uncertainty, and in the presence of a precautionary saving motive should be associated with larger saving. Their results, based on the period 1968–1986, suggest that precautionary saving is a non-negligible part of household behaviour. Using a few more cross-sections of the same data set, Banks et al. (1999) estimated an autoregressive moving-average process for cohort income. From the time-series innovation of this process, they obtained two components of income uncertainty: one which is common to all cohorts, and another which is specific to particular cohorts. Including these measures of uncertainty as separate regressors in a consumption growth equation, together with labour market and demographic variables, they found that it is essentially the latter component of uncertainty that plays an important role in determining changes in consumption.

The present paper improves on the existing studies on precautionary saving in the UK because it uses a panel data set, rather than time series of repeated cross-sections. Within a panel, the same households are interviewed every year. This allows us to calculate time-varying household-specific measures of uncertainty and to take into account unobserved household heterogeneity.

3. Main features of the data and descriptive statistics

3.1. The data set

We use waves one to eight of the BHPS, covering the years 1991 to 1998. A representative sample of 10,000 individuals living in Britain was interviewed in 1991. These individuals, together with their co-residents, were interviewed again each year thereafter. The BHPS provides information on respondents' demographic, occupational, educational, and income characteristics.[1] In each wave, individuals are asked the following questions on their saving behaviour:

"Do you save any amount of your income, for example by putting something away now and then in a bank, building society, or Post Office account other than to meet regular bills? Please include share purchase schemes and Personal Equity Plan schemes."

If the answer to this question is positive, then respondents (the savers) are asked the following:

"About how much on average do you personally manage to save a month?"

The information that is provided in these questions only refers to positive saving. Dissaving in the form of decumulation of financial assets is not considered, which makes the saving variable that we use in our analysis censored at zero. All the relevant income and saving variables are expressed in 1995 pounds.[2]

Since saving decisions are likely to be taken at the household level, we aggregate the individual saving observations accordingly. We thus consider the household as our unit of observation.

3.2. Measuring earnings uncertainty

We use four measures of earnings uncertainty. The first one is similar to that used by Lusardi (1998). In waves 6 and 7 of the survey, respondents are asked the following question:

"In the next twelve months, how likely do you think it is that you will become unemployed?"

The possible answers that can be given are: very likely, likely, unlikely, and very unlikely. After rescaling these responses to 0-1, we can interpret them as

Table 1. Ordered Probit regression for the probability of job loss of household heads

Dependent variable: Probability of job loss of the household head	
Age	0.013
	(5.39)
Past unemployment	1.103
	(3.19)
Tenure at current employer	−0.004
	(−1.06)
Full-time	−0.115
	(−1.34)
College education	−0.167
	(−2.50)
Private sector	0.09
	(1.82)
Sample size	2,647
Log likelihood	−2,575.84

Note: T-statistics are in parenthesis.
Source: BHPS, waves 6 and 7.

a subjective probability distribution of the relevant event. In Table 1, we report an Ordered Probit regression of these probabilities on a set of individual characteristics. The results are as one would expect: the probability of losing one's job is an increasing function of age and past unemployment, and is lower for college graduates, and for respondents who work in the public sector.

In this framework, an individual loses his job with a subjectively evaluated probability of p: in such case, he/she earns 0. With a probability of $(1-p)$, the individual does not lose his/her job and earns E. The individual's earnings can thus be seen as a random variable, with expected value equal to $(1-p)E$, and variance given by $p(1-p)E^2$. As in Lusardi (1998), we use the latter variable for the household head,[3] as our first measure of household earnings uncertainty, which we denote with *VAR1*.[4] We will estimate both a cross-sectional saving equation for wave 6 (Table 4), and a panel equation for waves 6 and 7 (Table 6), which include *VAR1*.

Furthermore, we construct three additional household-specific measures of earnings uncertainty, which make use of the panel dimension of our data set. The first one, *VAR2*, is an overall measure of uncertainty, obtained for each household by taking the square of the difference between detrended household earnings (Y) in 1998 and in 1991, divided by seven to have an annual rate (see Carroll and Samwick 1998, for a similar approach).[5] The second one, *VAR3*, is simply the variance of Y_t, where the subscript t represents our wave indicator, over the eight available waves. This measure of uncertainty assumes that all income shocks are transitory. Since this assumption is subject to criticism, we consider a final measure of earnings variability, *VAR4*, according to which all income shocks are fully permanent: this measure is given by the variance of $(Y_t - Y_{t-1})$ calculated over waves 2 to 8.[6]

We will estimate both cross-sectional and panel saving equations, which will be respectively based on wave 8 (Tables 4 and 5), and waves 6 to 8 of the survey (Table 6), and will include in turn *VAR2*, *VAR3*, and *VAR4*. In the panel regressions, we will allow our measures of earnings uncertainty to vary across waves. In general terms, $VAR2_t$ $(t = 6, 7, 8)$ will be defined as the difference between Y_t and Y_1, divided by $(t - 1)$ to have an annual rate. Similarly, $VAR3_t(VAR4_t)$ will be defined as the variance of $Y_t(Y_t - Y_{t-1})$ in the $t(t - 1)$ waves of the survey preceding and including year t.[7] This procedure assumes that households update their perceived earnings uncertainty in each wave, using the information on earnings that becomes available.

3.3. Sample restrictions and descriptive statistics

As in Carroll and Samwick (1997), we restrict our sample to households where the same person is head in each year, and where the spouse/partner, if present, remains the same. We exclude those households whose head is younger than 25 or older than 65, and those who do not have valid data on saving and net earnings.

In those specifications that use *VAR1* as a measure of earnings uncertainty, we further restrict the sample to those households whose head is in employment.[8] Since information on the subjective probabilities of job loss is only provided in waves 6 and 7 of the survey, the sample that can be used in estimation is limited to those two waves and consists of 2,750 household-years.

In those specifications which use *VAR2*, *VAR3*, or *VAR4* as measures of earnings uncertainty, the data set is restricted to those households who have been present in all the waves (balanced panel). We want in fact our measures of earnings variability to be calculated over the same number of years for each household. We also limit the sample to those households where at least one member was in paid employment in one or more of the eight waves. Moreover, as in Carroll and Samwick (1997), we exclude those households whose earnings in any year were less than 20% of the average over the period. If these households were included, our measures of earnings variability would be dominated by these few observations. Finally, since we calculate the variances over periods of at least six years, the waves that can be used in estimation in this case are waves 6 to 8, corresponding to 1,785 household-years.

In Table 2, we present descriptive statistics for the two samples described above. The Table reports the percentage of households where at least one member saved in the relevant period, and the average monthly saving of these households. We distinguish households by demographic characteristics, age, education and occupation of the head. About 65% of the households save when we consider the first sub-sample. The corresponding figure for the second sub-sample is 73%. As expected, the highest proportions of households who save can be found among married or cohabiting couples with no dependent children, among the more educated groups, and among the managers and administrators, and the professional, associate professional, and technical occupational categories. The non-zero average monthly saving is also generally higher for the above mentioned categories. Finally, looking at the differences in saving behaviour across age groups, we can see that in both sub-samples, those households whose head is aged between 35 and 44 are characterised by lower non-zero average monthly saving.

Table 3 describes how our measures of uncertainty differ across age, edu-

Table 2. Household saving by demographic characteristics, age, education, and occupation of the household head

	% of households who save (1)	Non-zero average monthly saving (£) (2)	% of households who save (3)	Non-zero average monthly saving (£) (4)
	Sub-sample 1 (2,750 observations)		Sub-sample 2 (1,785 observations)	
All	65.20	193.94	73.05	189.05
Demographic variables				
Married/cohabiting	70.07	207.20	75.15	195.51
Not married/cohabiting	52.72	148.77	63.14	152.78
No dependent children	68.67	207.05	74.81	208.98
One dependent child or more	60.73	174.83	71.11	165.89
Age				
25–34	66.45	202.23	77.49	179.59
35–44	62.01	168.21	71.47	174.44
45–54	67.78	204.27	74.34	203.84
55–65	65.06	214.43	70.48	196.27
Education				
Less than A levels	60.22	162.57	64.93	147.47
A levels	68.18	191.36	81.90	186.71
Some college	66.16	183.88	77.37	197.23
College	71.00	273.30	76.82	251.49
More than college	75.89	256.65	83.58	281.60
Occupation				
Managers & administrators	70.03	241.95	83.01	216.54
Professional	70.77	236.26	76.68	235.53
Associate professional & technical	71.04	221.27	78.11	226.35
Clerical & secretarial	57.48	126.43	62.91	126.03
Craft related	66.19	155.69	73.15	159.44
Personal & protective services	61.88	129.03	74.62	108.05
Sales	59.26	192.96	56.25	209.82
Plant & machine operatives	58.55	157.40	65.88	176.25
Others	58.49	191.84	63.64	113.41

Source: Column (1): BHPS, waves 6 and 7. Columns (2) to (4): BHPS, waves 6 to 8.

cation and occupation categories, all relative to the household head. From column (1), we can see that when *VAR1* is used as a measure of uncertainty, uncertainty gradually increases with the age and the educational qualifications of the household head. Moreover, *VAR1* is higher for the managers, as well as for those household heads working in professional, associate professional, and technical occupations. Columns (2) to (4) show that the age categories 25–34 and 45–54 generally have the highest earnings variability, measured by *VAR2*, *VAR3*, and *VAR4*, respectively. Uncertainty tends to be lowest for households whose head has an A level as the maximum educational qualification, and highest for respondents whose head has a college degree. As for the occupational categories, the personal and protective services, and other occupations tend to be characterised by particularly low uncertainty, while the professional and associate professional occupations generally have high uncertainty.

Table 3. Mean household earnings uncertainty by age, education, and occupation of the household head

	$VAR1_t$ (1)	$VAR2_t$ (2)	$VAR3_t$ (3)	$VAR4_t$ (4)
All	3.16	0.38	0.63	0.98
Age				
25–34	2.67	0.35	0.76	1.31
35–44	2.98	0.27	0.55	0.91
45–54	3.36	0.53	0.76	1.05
55–65	4.22	0.31	0.46	0.70
Education				
Less than A levels	1.85	0.25	0.42	0.66
A levels	2.83	0.18	0.40	0.60
Some college	3.47	0.34	0.65	1.04
College	4.61	0.97	1.36	2.08
More than college	9.82	0.64	0.99	1.10
Occupation				
Managers & administrators	4.66	0.28	0.54	0.78
Professional	5.46	0.84	1.22	2.01
Associate professional & technical	4.10	0.48	0.82	1.18
Clerical & secretarial	1.44	0.25	0.51	0.92
Craft related	2.13	0.27	0.50	0.79
Personal & protective service	1.58	0.18	0.39	0.60
Sales	2.94	0.45	0.61	0.65
Plant & machine operatives	1.91	0.36	0.57	0.80
Others	1.07	0.16	0.30	0.46

Note: See text for variable definitions.
Source: Column (1): BHPS, waves 6 and 7. Columns (2) to (4): BHPS, waves 6 to 8.

4. Empirical results

4.1. General specification

In our empirical specifications, we report Tobit regressions to analyse the determinants of household saving decisions, and assess the extent to which uncertainty affects these decisions. We use a Tobit estimation technique because, as mentioned in the previous section, the question that households are asked in the BHPS only allows for positive or 0 saving as a response. Saving could in principle take negative values, but these negative values are not observed due to censoring. The regressions are first based on single cross-sections of the survey, and then on a panel made up of two/three waves. In the former case, the equations that we estimate are of the following form, where the subscript i indicates the household:

$$\frac{S_i}{Y_i^P} = \alpha_0 + \alpha_1 Y_i^P + \alpha_2 VARj_i + X_i'\beta + e_i \quad (j = 1, 2, 3, 4) \tag{1}$$

S_i represents the average monthly amount saved by household i in the relevant cross-section. Y_i^P is permanent income for household i. It is included on the right hand side of our specification because there is evidence that saving varies across levels of permanent income, due to the non-homotheticity of preferences (Carroll and Samwick 1997, 1998).[9] We obtained permanent income by taking the fitted values from a random-effects regression of household earnings on household characteristics, gender, age, age squared, education dummies, occupational dummies, and interactions of the latter two groups of dummies with age and age squared, all relative to the household head (see Kazarosian 1997, for a similar approach).

$VARj_i$ represents one of the four household-specific measures of earnings uncertainty defined in the previous section. X_i includes a set of characteristics of household i or its head, which is assumed to affect saving. It includes a quadratic in the age of the head, various demographic and educational variables, regional and cohort dummies, and the household head's health status. These variables are generally aimed at capturing differences in preferences. X_i also includes the household's subjectively evaluated financial situation, and expectations about next year's financial situation. The expectations variables are included to see whether respondents save to offset future expected declines in income, in accordance with the life cycle model. Finally, e_i is an idiosyncratic error term.

4.2. Cross-sectional Tobit regressions

Table 4 gives the parameter estimates of Eq. (1). Column (1) refers to the case in which a Tobit saving equation is estimated on the wave 6 cross-section of the BHPS, in the case in which *VAR1* is used as an explanatory variable.[10] Columns (2) to (4) refer to the cases in which a similar equation is estimated on the wave 8 cross-section, in the cases in which *VAR2*, *VAR3*, and *VAR4* are respectively used.

In all regressions, we can see that the coefficient on the uncertainty variables is positive and statistically significant, ranging from 0.05, when *VAR1* is used to 0.27, when *VAR3* is used. A doubling in uncertainty suggests that the 'true' saving ratio increases by values ranging from 1.8% when *VAR2* is used to 5.7% when *VAR1* is used.[11] These results support the precautionary saving hypothesis.

The coefficients on the earnings uncertainty terms subsume two components: the effect due to those households who were not saving before, and start saving as a consequence of the increase in uncertainty; and the effect from the increased savings of those households who were already saving. It can be shown that the proportion of the total change due to the latter component is given by the following formula:

$$A = 1 - \frac{Zf(Z)}{F(Z)} - \frac{f(Z)^2}{F(Z)^2}, \tag{2}$$

where $f(Z)$ represents the unit normal density, and $F(Z)$ the cumulative normal density functions (McDonald and Moffitt 1980). As an estimate of $F(Z)$, one can use the proportion of households in the relevant sample who save. When *VAR1* is used, the proportion A is equivalent to 47.2%, whereas when *VAR2*, *VAR3*, or *VAR4* are used it is equivalent to 52.7%. This suggests that an increase in earnings uncertainty not only induces more households to save, but also induces those households who were already saving to save more.

Table 4. Cross-sectional Tobit estimates for household saving

	(1)	(2)	(3)	(4)
Constant	−8.787	−3.199	−21.713	−8.398
	(−0.26)	(−0.06)	(−0.41)	(−0.16)
Demographic variables				
Age	0.271	−1.346	−0.404	−1.129
	(0.156)	(−0.53)	(−0.16)	(−0.45)
Age^2	−0.005	0.022	0.012	0.020
	(−0.23)	(0.80)	(0.44)	(0.73)
Male	1.557	7.159	6.715	6.843
	(0.94)	(2.52)	(2.40)	(2.44)
Number of adults in household	−0.764	−0.581	−0.614	−0.572
	(−0.85)	(−0.54)	(−0.58)	(−0.54)
Number of dependent children in household	−2.786	−1.181	−1.232	−1.280
	(−4.37)	(−1.43)	(−1.53)	(−1.59)
Married/cohabiting	5.209	0.756	0.986	0.970
	(2.93)	(0.27)	(0.36)	(0.36)
Education				
Post-graduate degree	1.414	10.106	9.307	9.385
	(0.49)	(2.89)	(2.83)	(2.85)
College degree	4.281	5.813	5.699	5.853
	(2.29)	(2.54)	(2.53)	(2.58)
Some college	0.042	2.455	2.597	2.513
	(0.03)	(1.54)	(1.67)	(1.61)
A levels	2.523	4.928	5.369	5.304
	(1.55)	(2.42)	(2.68)	(2.64)
Financial variables				
Financial situation expected to deteriorate	4.381	5.589	5.568	5.639
	(2.51)	(2.67)	(2.69)	(2.72)
Financial situation expected to improve	−0.222	−0.843	−0.921	−0.818
	(−0.19)	(−0.59)	(−0.65)	(−0.58)
Financial situation worse than expected	−3.019	−1.276	−1.081	−1.168
	(−2.08)	(−0.71)	(−0.61)	(−0.66)
Financial situation better than expected	1.635	2.636	2.305	2.494
	(1.38)	(1.88)	(1.67)	(1.80)
Financial situation: good	9.476	8.082	8.390	8.206
	(7.31)	(4.63)	(4.93)	(4.79)
Financial situation: bad	−0.918	−1.855	−1.896	−2.173
	(−0.36)	(−0.51)	(−0.53)	(−0.60)
Other variables				
Health status	−0.995	−0.863	−0.811	−0.821
	(−1.49)	(−1.07)	(−1.02)	(−1.03)
Permanent income	0.002	−0.002	−0.001	−0.001
	(1.104)	(−0.87)	(−0.79)	(−0.80)
VARj	0.053	0.224	0.267	0.200
	(1.73)	(2.27)	(2.31)	(1.99)
Sample size	1,303	541	553	551
Number of censored observations	483	134	136	135
Log likelihood	−3,783.3	−1,722.7	−1,761.2	−1,757.0

Notes: Asymptotic *t*-ratios are in parenthesis. Regional dummies and cohort dummies were included in all specifications. *VARj* stands for *VAR1* in column (1), for *VAR2* in column (2), for *VAR3* in column (3), and for *VAR4* in column (4). See text for the definitions of these variables.
Source: Column (1): BHPS, wave 6. Columns (2) to (4): BHPS, wave 8.

The signs and coefficients of the other independent variables generally make sense. When *VAR1* is used as our measure for uncertainty, we can see from column (1) that those households with more dependent children, who currently perceive their financial situation to be worse than expected tend to save less. Those households whose head is either married or cohabits, and has a college degree, and who perceive their financial situation as good tend to have higher saving rates. In the specifications where *VAR2*, *VAR3*, and *VAR4* are used (columns 2 to 4), the gender of the household head and his/her education, play an important role in determining saving rates, with those households with the most educated heads saving a higher proportion of their permanent income. In accordance with the "saving for a rainy day" hypothesis, expectations of a deteriorating financial situation lead to higher saving rates in all specifications. Finally, the household head's health status, and the household's permanent income do not have a statistically significant effect on saving.

4.3. Instrumental Variable (IV) cross-sectional Tobit regressions

One problem with the specifications in columns (2) to (4) of Table 4 is that *VAR2*, *VAR3*, and *VAR4* are likely to be inconsistent estimators of the true underlying variances of earnings, due to the small number of waves available for their calculation. This introduces measurement error in the equations, which can lead to biased estimates of the parameters of interest. We address this problem by presenting, in Table 5, the results of the estimation of the relevant equations, where *VAR2*, *VAR3*, and *VAR4* are instrumented. We use the household head's tenure at his/her present job, dummies for his/her age, for whether he/she works in the private sector, and for whether he/she holds an occupational pension as instruments. The estimates are obtained using the procedure illustrated in Newey (1987). We can see that the coefficients on our proxies for uncertainty are still statistically significant and positive. However, they are generally larger in size compared to the coefficients reported in Table 4, suggesting a downward bias of the latter. According to the results presented in Table 5, a doubling in uncertainty makes the 'true' saving rate rise by values ranging from 18.2% to 41.1%, supporting once again the precautionary saving hypothesis.[12]

4.4. Random-effects Tobit regressions

The estimates reported in both Tables 4 and 5 can be criticised on the grounds of the fact that they do not take into account household unobserved heterogeneity. This particular heterogeneity may be thought of as household differences in some unobserved or unobservable attribute (like tastes), that might affect saving. Failure to control for unobserved heterogeneity may lead to biased coefficients. In Table 6, we report the results obtained from the estimation of random-effects Tobit regressions for saving rates, which exploit the panel dimension of our data set to control for household unobserved heterogeneity. Column (1) refers to the case in which *VAR1* is used as an explanatory variable, for waves 6 and 7 of the BHPS. Columns (2) to (4) refer to the cases in which *VAR2*, *VAR3*, and *VAR4* are respectively used, and they all refer to

Table 5. IV cross-sectional Tobit estimates for household saving

	(1)	(2)	(3)
Constant	−40.903	−38.44	−27.626
	(−0.71)	(−0.70)	(−0.50)
Demographic variables			
Age	3.970	4.172	3.424
	(1.16)	(1.25)	(1.02)
Age^2	−0.050	−0.054	−0.044
	(−1.21)	(−1.32)	(−1.07)
Male	7.142	6.794	6.387
	(2.49)	(2.41)	(2.27)
Number of adults in household	−0.954	−1.176	−1.659
	(−0.83)	(−1.05)	(−1.47)
Number of dependent children in household	−1.787	−1.919	−1.889
	(−2.13)	(−2.35)	(−2.30)
Married/cohabiting	3.242	3.365	3.496
	(1.08)	(1.16)	(1.16)
Education			
Post-graduate degree	13.167	11.382	11.167
	(3.52)	(3.38)	(3.30)
College degree	5.424	4.600	3.831
	(2.33)	(1.92)	(1.49)
Some college	5.084	4.924	4.411
	(2.71)	(2.75)	(2.53)
A levels	7.420	8.210	8.002
	(3.26)	(3.48)	(3.33)
Financial variables			
Financial situation expected to deteriorate	4.844	4.579	4.712
	(2.32)	(2.22)	(2.28)
Financial situation expected to improve	−0.971	−1.362	−1.733
	(−0.65)	(−0.94)	(−1.20)
Financial situation worse than expected	0.809	1.096	1.185
	(0.42)	(0.56)	(0.58)
Financial situation better than expected	3.822	4.228	4.881
	(2.42)	(2.42)	(2.38)
Financial situation: good	7.979	7.517	7.148
	(4.50)	(4.31)	(3.98)
Financial situation: bad	−3.507	−4.043	−5.436
	(−0.93)	(−1.08)	(−1.43)
Other variables			
Health status	−0.447	−0.497	−0.733
	(−0.54)	(−0.61)	(−0.87)
Permanent income	−0.007	−0.007	−0.006
	(−1.67)	(−1.65)	(−1.44)
VARj	2.212	2.784	2.414
	(2.32)	(2.21)	(1.96)
Sample size	499	509	508
Number of censored observations	126	126	126
Log likelihood	−1,565.25	−1,601.75	−1,599.11

Notes: The estimates were obtained using the method illustrated in Newey (1987). *VARj* stands for *VAR2* in column (1), for *VAR3* in column (2), for *VAR4* in column (3). Also see notes to Table 4.
Source: BHPS, wave 8.

Table 6. Random-effects Tobit estimates for household saving

	(1)	(2)	(3)	(4)
Constant	−7.715	−10.78	−16.98	−11.28
	(−0.36)	(−0.31)	(−0.48)	(−0.32)
Demographic variables				
Age	0.212	−1.507	−1.214	−1.517
	(0.20)	(−1.07)	(−0.87)	(−1.08)
Age^2	−0.003	0.025	0.022	0.026
	(−0.23)	(1.63)	(1.44)	(1.64)
Male	1.243	7.117	6.884	6.914
	(1.17)	(2.67)	(2.63)	(2.63)
Number of adults in household	−0.204	0.163	0.237	0.279
	(−0.35)	(0.21)	(0.30)	(0.354)
Number of dependent children in household	−2.418	−0.639	−0.739	−0.747
	(−5.87)	(−0.94)	(−1.11)	(−1.11)
Married/cohabiting	4.615	0.363	0.415	0.392
	(3.95)	(0.14)	(0.17)	(0.16)
Education				
Post-graduate degree	2.602	8.845	8.374	8.424
	(1.44)	(2.87)	(2.82)	(2.83)
College degree	3.396	5.092	4.868	5.024
	(2.80)	(2.58)	(2.51)	(2.57)
Some college	0.175	3.066	3.148	3.108
	(0.20)	(2.25)	(2.35)	(2.31)
A levels	1.700	4.271	4.441	4.376
	(1.59)	(2.38)	(2.50)	(2.46)
Financial variables				
Financial situation expected to deteriorate	3.554	4.492	4.309	4.271
	(3.29)	(3.95)	(3.82)	(3.78)
Financial situation expected to improve	−0.096	−0.127	−0.137	−0.101
	(−0.133)	(−0.16)	(−0.175)	(−0.13)
Financial situation worse than expected	−2.248	−1.253	−1.10	−1.150
	(−2.48)	(−1.33)	(−1.18)	(−1.23)
Financial situation better than expected	1.546	2.459	2.273	2.264
	(2.11)	(3.17)	(2.96)	(2.93)
Financial situation: good	8.699	5.041	5.314	5.221
	(10.62)	(5.01)	(5.32)	(5.20)
Financial situation: bad	−1.823	−4.070	−3.147	−3.067
	(−1.08)	(−1.93)	(−1.54)	(−1.49)
Other variables				
Health status	−1.14	−0.142	−0.125	−0.128
	(−2.78)	(−0.27)	(−0.24)	(−0.25)
Permanent income	0.0002	−0.002	−0.002	−0.002
	(0.26)	(−1.56)	(−1.43)	(−1.37)
VARj	0.086	0.231	0.235	0.124
	(3.47)	(2.77)	(2.42)	(1.52)
Sample size	2,623	1,606	1,636	1,630
Number of censored observations	911	422	426	424
Log likelihood	−8,406.3	−4,943.9	−5,049.6	−5,035.6
ρ	0.11	0.58	0.57	0.58

Note: See notes to Table 4. The term ρ represents the fraction of total variance attributable to the unobserved random-effects.
Source: Column (1): BHPS, waves 6 and 7. Columns (2) to (4): BHPS, wave 6 to 8.

waves 6 to 8 of the survey. Our estimating equation in both cases takes the following form, where t indexes waves:

$$\frac{S_{it}}{Y_{it}^{p}} = \alpha_0 + \alpha_1 Y_{it}^{p} + \alpha_2 VARj_{it} + X_{it}'\beta + v_i + e_{it} \quad (j = 1,2,3,4; t = 6,7,8) \qquad (3)$$

The error term is now made up of an unobservable household-specific time-invariant effect (v_i), which we assume to be random and captures the unobserved household heterogeneity, and of an idiosyncratic error term (e_{it}). Another advantage of the random-effects estimation technique is that it allows us to deal with time-invariant measurement error (which will be encompassed in v_i) without using instrumental variables (Baltagi 1995).

In all the saving equations that we estimated, the panel-variance component, ρ, is significantly different from 0 and relatively large. In particular, when *VAR2*, *VAR3*, or *VAR4* are used as explanatory variables, the ρ coefficient tells us that about 58% of all the variance in household saving ratios can be attributed to unobserved household-specific characteristics. This shows the importance of following the same households over time when trying to understand their saving behaviour.

We can see from column (1) of Table 6 that the coefficient on *VAR1* is once more positive and statistically significant at conventional levels. Similar results hold when *VAR2* and *VAR3* are used as measures of uncertainty. Column (4) however shows that *VAR4* is only significant at the 13% level. In all specifications, those households with a male head and those who currently perceive their financial situation as good or better than they had expected in the previous year generally save more. Saving rates are also higher for those households whose head has a college or a post-graduate degree. Finally, as in the previous specifications, if the financial situation of the household is expected to deteriorate, then saving rates are higher.

5. Conclusions

In this paper, we have analysed saving behaviour in Britain using for the first time a panel data set: the BHPS. We have constructed various measures of earnings variability, which we have used as proxies for uncertainty. Our results based both on cross-sectional and panel Tobit regressions have shown that households' saving decisions are, in all cases, significantly affected by these measures of uncertainty. Confirming therefore the conclusions of the previous FES-based studies on the UK, we conclude that there exists a significant precautionary component in saving behaviour. In line with the predictions of the life-cycle/permanent income model, it also appears from our regression results that those respondents who expect their financial situation to deteriorate in the year to come tend to save more (saving for a "rainy day").

The BHPS generally asks respondents what their first reason for saving is. The possible answers considered are the following: holidays, old age, car purchase, children, house purchase, home improvement, household bills, special events, no special reason, share purchase schemes, own education, and other motives. Saving for no special reason can be broadly considered as precautionary saving. In wave 8 of the survey, about 43% of the savers claim that they save for no special reason. This tends to confirm our results of a strong precautionary motive for saving.

Earnings variability is however not the only possible source of uncertainty. Health risk and longevity risk might also play a significant part in determining individuals' saving behaviour. These alternative types of uncertainty, need further investigation on British data. This will be the object of future research.

Endnotes

1 Information is provided both at the individual and the household level. For more details on the BHPS, see Taylor (1994) and Taylor (1999).

2 The variables are deflated using the Retail Price Index.

3 Following Guiso et al. (1992) and Lusardi (1997, 1998), we can justify this procedure by the underlying assumption that the variance of household earnings can be reasonably proxied by the variance of the earnings of the head of the household.

4 Note that we do not use the respondents' subjective probability of job loss (p) itself as a proxy for uncertainty for the following reason. Suppose that an individual is certain that he/she will lose his/her job in the near future, due to an announced close down of the firm where he/she works. In this case, if the individual saves more, his/her extra saving cannot be classified as precautionary: Increased saving in response to an expected drop in the mean of future income arises in fact in a standard certainty equivalence model (see Alessie and Lusardi 1997, for an investigation of this issue). One limitation of our approach is that it only takes into account the probability of job loss in the next twelve months, and not afterwards.

5 Since we are trying to measure uncertainty, we are not interested in that part of the variability of earnings, which is due to predictable life-cycle changes in earnings. In our formulas for *VAR2*, *VAR3*, and *VAR4*, we therefore use detrended net monthly household earnings (Y). We obtain Y by taking the residuals from a random-effects regression of net monthly household earnings on household characteristics, gender, age, age squared, educational and occupational characteristics, and interactions of the latter two groups of dummy variables with age and age squared (all relative to the household head).

6 We thank an anonymous referee for suggesting this alternative measure of uncertainty.

7 This framework assumes that households consider past earnings volatility as a measure of likely future volatility. See Guariglia and Rossi (2001) for evidence on the fact that in the BHPS, there appears to be some form of ARCH behaviour in earnings, which would justify the above assumption. Also see Banks et al. (1999) who make a similar assumption.

8 We exclude the self-employed because a measure of net earnings is not provided for them.

9 Y_i^p can also be seen as a proxy for household wealth.

10 Similar results were obtained for wave 7. We do not report them for brevity.

11 These numbers are elasticities, evaluated at sample means. By 'true' saving rate, we mean the latent variable underlying the Tobit model. We thank an anonymous referee for suggesting this calculation.

12 The R^2s in the first stage regressions are respectively 0.104, 0.112, and 0.111, when *VAR2*, *VAR3*, and *VAR4* are the variables instrumented for.

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Family structure and children's achievements

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Abstract. In this paper we estimate the relationships between several outcomes in early adulthood (educational attainment, economic inactivity, early childbearing, distress and smoking) and experience of life in a single-parent family during childhood. The analysis is performed using a special sample of young adults, who are selected from the first five waves of the British Household Panel Survey (1991–95) and can be matched with at least one sibling over the same period. We also perform level (logit) estimation using another sample of young adults from the BHPS. We find that: (i) experience of life in a single-parent family is usually associated with disadvantageous outcomes for young adults; (ii) most of the unfavourable outcomes are linked to an early family disruption, when the child was aged 0–5; and (iii) level estimates, whose causal interpretation relies on stronger assumptions, confirm the previous results and show that, for most outcomes, the adverse family structure effect persists even after controlling for the economic conditions of the family of origin.

JEL classification: J12, I20, I30

Key words: Family structure, intergenerational links, siblings estimators

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1. Introduction

This paper asks whether a number of outcomes in early adulthood are associated with experience of life in a single-parent family during childhood. Economists have long identified the family as one the most important institutions in a market economy that fosters the transmission of income inequality over time (Knight 1935; Becker 1981) and may act as a potential income equaliser across family members (Griliches 1979). A large literature has extensively studied the effect of children on marital instability and other household behaviour (e.g., Becker et al. 1977; Browning 1992 for a survey). Only in recent years, however, has the intergenerational transmission of human and social capital received empirical attention, with studies that examine the correlation between children's outcomes as young adults and a variety of parental circumstances and events during childhood, including family structure.[1]

A legitimate concern with much of this new literature is that the estimated effect of childhood family structure on children's achievements might be spurious (Manski et al. 1992; Mayer 1997). This is due to the mutual association that family structure and children's outcomes share with some unmeasured true causal factor. For example, the association between having experienced life in a single-parent family and, say, experiencing difficulties in the labour market may not be necessarily the result of family structure during childhood. Rather, differences in labour market success may simply reflect the characteristics of families in which the children of single mothers are brought up. For a broad set of young people's outcomes (education, inactivity, early childbearing and health), we estimate both level (cross-sectional) models and family-specific fixed-effects models.[2]

Although the use of sibling data has become increasingly common in economics, most of the empirical studies linking parents' behaviour and children's attainments have not addressed the problem of unmeasured heterogeneity with sibling estimators.[3] We show that the effect of family structure on outcomes can be identified with sibling differences if family structure does not respond to "idiosyncratic endowments" of children. On this assumption, our sibling-difference estimates would measure the causal impact of childhood family structure on young adults' achievements. But note that, in addition to inherent differences between siblings (e.g., one born with a disability), differences between siblings in their idiosyncratic endowments include differences over time in parental attitudes and behaviour which may affect both family structure and children's outcomes. For example, the father may develop an alcohol addiction, giving rise to a situation in which an elder sibling spends only a small part of his childhood with an alcoholic father while the youngest has one for most of her childhood. The father's alcohol problem may directly affect investment in the youngest child, and her parents may also divorce because of it, thereby causing correlation between idiosyncratic endowments and family structure. Thus, while the assumption of no such correlation is weaker than the assumptions needed in the cross-sectional model (see the Appendix), it is still a strong one. A causal interpretation of the level estimates relies on even stronger assumptions, but such estimates are useful for comparing our results to those provided in other studies. It is, therefore, safest to interpret both sets of estimates as suggestive associations, with the sibling-difference estimates controlling for more aspects of family background than the level estimates, making them less contaminated by unmeasured factors associated with both family structure and children's outcomes.

Most of what we know of the relationship between family background and children's outcomes is from the United States. Little is known about British cohorts born since 1958. In a recent study, Kiernan (1997) finds that, among members of the 1958 birth cohort, divorce during childhood is associated with outcomes – for educational attainment, economic situation, partnership formation and dissolution and parenthood behaviour in adulthood – which we would generally interpret to be worse than the same outcomes for young adults from intact families. Similarly, the study by Fronstin et al. (forthcoming) suggests that a parental disruption adversely affects labour market outcomes of children at age 33 (in 1991). Their results indicate that the effect occurs primarily through decreased employment for men and decreased wage rates for women. These studies build on earlier research by Elliot and Richards (1991), Ní Bhrolcháin et al. (1994) and Kiernan (1996) on schooling and socio-economic performance, and Kiernan (1992) and Cherlin et al. (1995) on demographic outcomes. All use samples from the National Child Development Study (NCDS) for the 1958 birth cohort.

Our analysis uses, instead, a special sample of young adults from the first five waves (1991–1995) of the British Household Panel Survey (BHPS), who can be matched with at least one sibling over the same period.[4] These young adults are then linked to their mothers' family history collected in the 1992 wave, as well as to other information about the mother from the mother's interviews during the panel,[5] and they are followed over the panel years. We also analyse another sample, in which young adults need not be matched with a sibling, but, as in the Sibling Sample, they must live with their mothers for at least one of the five panel years. The BHPS data have some clear advantages over the NCDS data. They are a better reflection of contemporary impacts of family background, which may be particularly important for that of family structure, because the incidence of single parenthood is now much higher than among the parents of the 1958 cohort. Moreover, they allow for better (although not ideal) measurement of family economic circumstances.[6] More importantly, with the information on siblings, they allow us to control for unobserved additive effects that are shared by children who belong to the same family.[7]

We find that experience of life in a single-parent family is typically associated with unfavourable outcomes for young adults. They achieve lower educational levels, face higher risks of economic inactivity and early birth, and they also have higher chances of smoking and feeling more distressed. Family structure in early childhood (ages 0–5 of the child) has the strongest associations with all the outcomes under analysis (similar to the impact of income on outcomes found by Duncan et al. 1997).[8] All these results also emerge with level estimates, whose causal interpretation relies on even stronger identifying assumptions than in the case of sibling differences. Our level estimates are consistent with most of the available evidence in this literature. As in McLanahan (1997), they confirm that, for almost all the outcomes under analysis, experience in a single-parent family during childhood still matters after taking the economic circumstances of the family of origin into account.

The rest of the paper is organised as follows. Section 2 describes the data and the estimation procedures, Sect. 3 presents our main results, and Sect. 4 concludes. The Appendix outlines the identification problem that is common to all studies of the relationship between family structure during childhood and children's outcomes.

2. Data

2.1. Estimating samples and family background measures

The data come from a special sample selected using the first five waves of the British Household Panel Survey (BHPS). In Autumn 1991, the BHPS interviewed a representative sample of 5,500 households, containing about 10,000 persons. The same individuals are reinterviewed each successive year, and if they leave their original households to form new households, all adult members of these new households are also interviewed. Similarly, children in original households are interviewed when they reach the age of 16. Thus, the sample remains broadly representative of the population of Britain as it changes through the 1990s.[9]

The second wave (1992) of the BHPS contains retrospective information on complete fertility, marital and cohabitation histories for all adult panel members in that year. Our analysis proceeds as if all children lived with their mothers throughout their years of dependency, which we assume to be until their sixteenth birthday.[10] This information provides the basis for our family structure measure: whether or not the young adult spent time in a single-parent family during his/her childhood. A child is defined as being brought up in an intact family if he/she lived continuously with both biological (or adoptive) parents, up to his/her 16-th birthday. Thus, according to our definition, a child would have spent some time in a single-parent family if he/she ever lived with a biological or adoptive mother who was not cohabiting nor married before his/her 16-th birthday, either because of a partnership dissolution or because he/she was born outside of a live-in partnership and the mother did not cohabit or marry within one year of the birth.[11] This measure is also broken down by the timing of the start of a spell in a lone parent family, distinguishing between three different child developmental stages, ages 0–5, 6–10, and 11–16.[12]

By matching young adults with their mothers, we are also able to measure other family background characteristics that would be unavailable otherwise, such as age of the mother at the young person's birth, mother's education, parents' real income when the child was aged 16 (or the youngest age at which we observe the child living with parents), whether or not parents are owner-occupiers, and if so, house value when the child was aged 16, years spent at the address occupied by the child at age 16 and current age of the child. We also obtain information on the smoking behaviour of the mother at the time the mother and her young-adult child lived together, which will be used for one of the health-related outcomes defined below (smoking). In addition, the third wave (1993) of the BHPS contains retrospective information on job histories for all adult panel members.[13] With this information we determine the proportion of months that mothers worked in each of the three developmental stages and over the entire childhood of the young adults.[14]

In our analysis, we use two samples. The first sample (labelled as "Individual Sample") consists of 764 individuals who: (i) were observed living with their mother when aged 16 or 17 during any of the first five waves (1991–

1995) of the BHPS; and (ii) report full information on outcomes and family background measures. The coresidence condition (condition (i)) is imposed in order to match data on family background from the mother's record to her child. Because 95% of the panel members live with their parents when aged 16–17 (Ermisch 1996), our sample is likely to be random. For some of the outcomes (e.g., schooling and early childbearing), we impose additional restrictions that we will describe below.

The second sample (labelled as "Sibling Sample") consists of 411 individuals with full information on outcomes and background measures who: (i) coresided with their mother for at least one year during the first five waves; (ii) were born between 1965 and 1979; and (iii) can be matched with at least one sibling (or half-sibling). Imposing condition (iii) – which allows us to control for unobserved heterogeneity in terms of family- or mother-specific fixed effects – on the Individual Sample would leave us with only 74 sibling pairs, arguably too few to draw reliable inference. At the cost of introducing some selection bias (i.e., correlation of the unobservables associated with the dependent variable(s) and the sample selection process), we impose conditions (i) and (ii), thereby increasing the number of sibling pairs to 252.

Table 1 shows the distribution by age of the young adults in the two samples. All individuals in the Individual Sample are 21 years-old or less (because of our selection, they were born between 1974 and 1979), and evenly distributed by age. About three-fourths of the individuals in the Sibling Sample are 22 or less and only 8% are more than 25 years-old. Table 2 presents the summary statistics of the variables for all individuals in the two samples. These statistics are computed for the last available year in which the indivi-

Table 1. Distribution of individuals by age in the two estimating samples

Age	Individual sample			Sibling sample		
	N	Prop.	Cum. Prop.	*N*	Prop.	Cum. Prop.
16	125	0.164	0.164	55	0.134	0.134
17	137	0.179	0.343	46	0.112	0.246
18	138	0.181	0.524	52	0.127	0.372
19	123	0.161	0.685	36	0.088	0.460
20	125	0.164	0.848	54	0.131	0.591
21	116	0.152	1.000	46	0.112	0.703
22				29	0.071	0.774
23				25	0.061	0.835
24				17	0.041	0.876
25				12	0.029	0.905
26				8	0.020	0.925
27				8	0.020	0.944
28				11	0.027	0.971
29				6	0.015	0.985
30				6	0.015	1.000
All ages	764			411	1.000	

Note: Individual Sample refers to individuals who lived with their mother at least at one interview date when aged 16 or 17 during the first five waves of the BHPS. Observations are at the last available period. Sibling Sample refers to individuals who coresided with their mother at least at one interview date during the first five waves of the BHPS, were born between 1965 and 1979, and are observed in any of the sample years with at least one sibling. For each sibling, observations are at the last available period.

Table 2. Summary statistics of variables used in analysis

Variable	Individual sample		Sibling sample	
	Mean	Std. Dev.	Mean	Std. Dev.
Age 16	0.164			
Age 17	0.179			
Age 18	0.181			
Age 19	0.161			
Age 20	0.164			
Age 21	0.152			
Age	18.437	1.682	20.304	3.543
Year of Birth	1976.4	1.703	1974.6	3.669
Female	0.476		0.418	
Ever in single-parent family	0.325		0.212	
Ever in single-parent family:				
child's age 0–5	0.200		0.090	
child's age 6–10	0.072		0.071	
child's age 11–16	0.052		0.051	
Mother has O level	0.346		0.350	
Mother has A level	0.153		0.124	
Mother has higher qualification	0.109		0.066	
Prop. of mother's time worked	0.434		0.375	
Prop. of mother's time worked:				
child's age 0–5	0.266		0.210	
child's age 6–10	0.422		0.359	
child's age 11–16	0.583		0.522	
Mother's age at birth ≤ 21	0.160		0.151	
Mother's age at birth ≥ 34	0.042		0.054	
Mother's age at birth	26.536	4.659	26.546	4.518
Mother smokes[a]	0.251			
Annual parents' real income (£10,000)	2.523	1.557		
Parents are house owners	0.775			
Current value of parents' house (£10,000)	7.482	7.134		
Years spent at current address	2.602	5.463		
Number of observations (individuals)		764		411

[a] Used in smoking regressions only.

duals are observed in the survey period under analysis. The Table indicates that the average year of birth of the young adults in the Individual Sample is 1976, with a mean age of 18 years. Nearly 48% of the sample are women. About 40% of the mothers of these young adults have no academic qualification, over three-fourths of parents were homeowners by the end of their offspring's childhood, and had an average real (1995) income of £25,000. Almost one-third of the sample experienced life in a single-parent family; either their mother's partnership dissolved before they reached age 16, or they were born outside of a live-in partnership. Of the children who spend some time in a single-parent family, 60% had this family experience below the age of 6.[15] On average, mothers gave birth at ages 26–27: 16% of the young adults in the sample were born when their mother was aged less than 22, and just over 4% of them have mothers aged 34 or more at their birth. Mothers worked almost 84 months, that is, 43% of the first sixteen years of life of their children. Maternal labour supply and child's age are clearly positively related. Approximately 1 in 4 of the young people in the sample have a mother who smokes.

Table 3a. Distribution of siblings (individuals) and sibling pairs in the sibling sample

	Number of:			
	Siblings per household	Households	Individuals	Comparisons (sibling pairs)
	2	165	330	165
	3	23	69	69
	4	3	12	18
Total		191	411	252

Table 3b. Sibling pairs with different experiences of family structure

All childhood		Ages 0–5		Ages 6–10		Ages 11–16	
Freq.	%	Freq.	%	Freq.	%	Freq.	%
30	11.9	31	12.3	22	8.7	15	60

Note: All percentages are computed in terms of total number of comparisons (see Table 3a).

The average figures for the slightly older Sibling Sample are similar, but here we have a smaller proportion of women and a smaller fraction of people that ever lived in a single-parent family. To ease the interpretation of the estimates, Tables 3a and 3b present this sample in greater detail. The 411 young adults come from 191 households: 165 of these households have 2 siblings in our sample, 23 have 3 siblings, and 3 have 4 siblings.[16] A total of 252 comparisons is then obtained from this sample. To identify an association between any variable x and any outcome y, the siblings estimator would require sibling variations in both x and y. Variations across siblings, say, in the proportion of months which the mother worked during each child's childhood are straightforward, simply because of birth order. It is only when the mother never/always worked over both children's childhood that there is no variation. But because our family structure measure is an incidence measure (over either the entire childhood or the three different developmental stages), the nature of sibling variations may be less clear. For example, in a two-child family, one of the half-siblings may have experienced a family break-up while aged 0–5, while the other child, born within a subsequent union of the mother, would never experience a family break-up if the mother and her new partner do not dissolve their union. Or, comparing two full siblings, one may be aged 0–5 when the parents' union dissolves while the other is aged 6–10. Table 3b shows that, of the 252 sibling pairs, 30 of them have a different experience of family structure over the entire childhood. Most of the action occurs at the early stages of child development: 31 sibling pairs live in different family structures when aged 0–5, while only 15 have a different experience when aged 11–16.

2.2. Outcomes

Education. Our measure of educational attainment is achieving an A-level qualification or higher qualification.[17] For each young person, we take the

Table 4. Mean outcomes by sample

	Individual sample	Sibling sample
Education	0.4765	0.4742
N	489	310
Inactivity	0.0781	0.0975
N	2388	1652
Early childbearing	0.0150	0.0195
N	1070	257
Distress	0.2038	0.2008
N	2388	1652
Smoking 10 or more cigarettes a day	0.1792	0.1660
N	2388	1652

Note: *N* is the number of observations (individuals or person-periods) used in estimation.

highest education level as that in the latest year in which we observe him/her in the panel. As it is rare to obtain A levels before the age of 18, we further limit the estimating sample to people who are in the panel at ages 18 or above. Thus, we perform our analysis on 489 and 310 individuals in the Individual Sample and the Sibling Sample, respectively. Table 4 indicates that the percentage of individuals who have achieved at least a highest qualification of A level is similar across samples, and around 47.5%.

Inactivity. This outcome is defined as neither working nor being in school nor looking after children, nor being in government training schemes. The analysis is based on 2,388 and 1,652 person-periods in the Individual Sample and Sibling Sample, respectively. This last sample matches siblings on the year of observation (thus avoiding comparisons at different points of the business cycle). As Table 4 shows, the inactivity rate is slightly larger for individuals in the Sibling Sample: 9.8%, versus 7.8% in the Individual Sample.

Early childbearing. This outcome is defined as having had a first birth at age 21 or less. For the young women in our samples, we estimate the association of the family background measures with the probability of becoming a mother in a given year, conditional on remaining childless up to that point and censoring women when they reach their 21st birthday. We have 1,070 and 257 person-periods in the Individual Sample and Sibling Sample, respectively. Because having a child is inherently age-dependent, the sisters' comparisons are made at common ages. On average, 2% of childless women aged 16–21 have a child each year, but the first birth rate increases with age. Lifetable estimates based on the Individual Sample imply that 13% of young women would become mothers by their 21st birthday, which is less than the one-fifth of women born during 1974–1975 who had a first birth by their 21st birthday indicated by registration statistics (Office for National Statistics (ONS), 1997, Table 10.3). This difference is likely to reflect our sample selection criteria based on coresidence with parents at age 16–17; that is, women who became mothers early are less likely to be observed living with their parents in the BHPS.[18]

Health. We analyse two measures of health-related outcomes. The first measure is defined as having a high level of distress, and it is derived from a set of subjective indicators of well-being.[19] The second measure takes the value of one if an individual smokes more than 10 cigarettes a day, and zero otherwise.[20] The analysis is conducted on 2,387 and 1,652 person-periods in the Individual Sample and Sibling Sample, respectively. In the latter, age enters parametrically, i.e., no age matching is imposed on sibling comparisons. Approximately 1 in 5 young adults reports a high level of distress, and 1 in 6 smokes 10 or more cigarettes a day.

2.3. *Estimation*

We estimate "level" logit regressions with the Individual Sample, and sibling fixed-effects (FE) linear probability models with the Sibling Sample.[21] The coefficient of the family structure variable can be interpreted as the average association of the outcome with family structure in a population in which the family structure impact varies in a random way.[22] Throughout the analysis, we compute robust standard errors that are consistent even if the residuals are not identically and independently distributed, that is, the standard errors are robust to arbitrary forms of heteroskedasticity for individuals over time. In the case of the education outcome, when all variables are measured at the last available year for each individual, the standard errors of the estimates obtained with the Individual Sample are robust to any form of correlation between siblings.

We consider young people's outcomes and their experiences of life in a single-parent family both during their entire childhood (Table 5) and at three developmental stages, ages 0–5, 6–10 and 11–16 (Table 6). All regressions include age, gender, and family structure during childhood. In the Individual Sample we further control for year of birth and mother's education, plus an indicator of whether or not the mother smokes for the smoking outcome only[23]. In an alternative specification of the logit regressions with the Individual Sample, we also include mother's employment patterns (proportion of months worked) during childhood, mother's age at child's birth,[24] and a set of variables that control for the economic circumstances of the family of origin. They are family income, whether parents are homeowners, value of the house if owners, and length of time spent at current (parental) address. These "economic" variables can only be measured when the child is aged 16, or in the first year in which we observe the child living with parent(s), and not during childhood. To the extent that there are persistent components of income or wealth, however, these variables should be indicative of the financial and economic environment of the family of origin.

The identification of an "effect" of childhood family structure on children's later achievements relies on assumptions about parents' (or mother's) and individual's behaviour as well as about the processes that determine cultural and genetic transmission of endowments across generations. In the Appendix we present a simple empirical model to clarify the identification problem inherent in all models of intergenerational links. The within-mother (of family) FE estimator applied to the Sibling Sample identifies the family structure effect under the assumption that family structure does not respond to, and is not correlated with, children's "idiosyncratic endowments". For the reasons discussed in the Introduction and the Appendix, this is a strong assumption, and we do not know the degree to which it is violated. If it is not true, the sibling-difference estimator only indicates suggestive associations between children's outcomes and family structure.

With the level (logit) estimator, this condition is necessary but not sufficient. Identification can be achieved only if one of the three following restrictions is further imposed: (a) there is no degree of "inheritability" of endowments across generations; (b) all the family background variables are independent of family endowments; (c) parents do not respond to child endowments or they only respond to differences in siblings' endowments. Dearden et al. (1997) find evidence of large and significant intergenerational correlations in earnings and years of schooling, thereby making it hard to accept condition (a). As pointed out above, we estimate two specifications of the logit regressions. The variables included in the main specification are likely to be independent of the children's idiosyncratic endowments, but mother's education and family structure may be correlated with the family endowment, thus violating condition (b). The variables in the alternative specification (such as mother's employment and family income) are likely to be correlated with both family and children's idiosyncratic endowments. Nevertheless, they are used in most of the studies in this literature, and our estimates can then be more easily compared to those currently available. Finally, if we believe that parents respond to child endowments, then also condition (c) is untenable. Behrman et al. (1982) and Ermisch and Francesconi (2000) formulate an optimising model of educational choice in which parents only respond to differences between their children's endowments, but identification with the level estimators still requires additional orthogonality assumptions (see Appendix).

3. Results

The estimation results are reported in Tables 5 and 6. A causal interpretation can only be given to the sibling-difference estimates if family structure does not respond to, and is not correlated with, idiosyncratic children's endowments. Both tables also contain the estimates from the logit regressions, whose causal interpretation relies on even stronger identifying assumptions. These estimates, expressed in the form of marginal effects for a young adult with average characteristics, offer however a useful benchmark for comparison.

3.1. Education

Having spent time with a single mother during childhood is associated with a significantly lower probability of achieving A level or more: 13.7% and 14.6% lower in the Individual Sample (main specification (i)) and the Sibling Sample, respectively (Table 5).[25] This association becomes weaker and less precisely estimated in the alternative specification (ii), thereby suggesting that the negative impact of single parenthood on young people's schooling operates partly through lower incomes and wealth in single-parent families (for a similar finding for the US, see Boggess 1998). While we cannot reject the hypothesis that the three stage-specific coefficients are equal (see *p*-values in Table 6), it is noteworthy that the FE estimates from the Sibling Sample exhibit their strongest negative association between schooling and experience in a single-parent family when the young adult was aged 0–5 (Table 6). An early

Table 5. Family structure during childhood and young adults' outcomes

	Individual sample					
	(i)		(ii)		Sibling Sample	
Outcome	Marg. eff.	|*t*-ratio|	Marg. eff.	|*t*-ratio|	Coeff.	|*t*-ratio|
Education	−0.137	2.913	−0.083	1.749	−0.146	1.896
Inactivity	0.056	3.901	0.037	2.577	0.018	0.917
Early childbearing[a]	0.018	2.250	0.012	1.531	0.024	1.688
Distress	0.055	2.453	0.056	2.455	0.036	1.412
Smoking 10+ a day[b]	0.073	2.790	0.063	2.325	0.068	2.670

Note: Estimates for the Individual Sample are marginal effects from logit regressions computed at the average values of all variables used. Regressions under specification (i) include: age, gender, year of birth, mother's education, and a constant. Regressions under specification (ii) include those in (i) plus: proportion of mother's time worked, mother's age at child's birth, family income at child's age 16 (or youngest age when child coresided with his/her mother), whether parents are house owners, value of the house if owners, and length of time spent at current address. Estimates in the Sibling Sample are obtained from within-mother fixed-effects (FE) model. All FE regressions include age. Sisters' differences are taken at the same age in the case of the early childbearing outcome; in all other cases, age enters parametrically.

[a] Women only.

[b] Controls for mother's smoking (Individual Sample only).

experience of parental loss, therefore, is more likely to jeopardise the child's subsequent educational career. Contrary to traditional stress theory, which predicts that the impact of a family split is strongest immediately after it occurred (Conger et al. 1993), our estimates suggest that it is a family disruption in *early* childhood (or being born outside of a live-in partnership) that has the most pronounced consequences on later educational achievements, possibly through its effects on salient aspects of the child's cognitive, cultural and social development. This result is in line with some of the US evidence (see Duncan et al. 1997).[26] The level estimates reveal that a family disruption in adolescence exhibits a large and significant negative correlation with educational attainment. Such a correlation becomes weaker after controlling for the family economic environment (specification (ii)).

3.2. *Inactivity*

A family breakdown during childhood is associated with a 5.6% higher probability of economic inactivity (specification (i), Table 5). Controlling for the economic circumstances of the family of origin reduces it only slightly, but the correlation is always positive and strongly significant. This finding is, however, not robust to the presence of mother's fixed effects. Interestingly, as we sort out the associations by developmental stage (Tables 6), the Sibling Sample estimates show that individuals who spent time in a single-parent family in their early childhood (when aged 0–5), have a 14% higher probability of being inactive in their young adulthood than those who did not experience that family structure. Notice also that the hypothesis of equality of the effects by developmental stage is rejected at any conventional level. The two specifications estimated with the Individual Sample detect the same pat-

Table 6. Family structure during childhood and young adults' outcomes by developmental stage

	Individual sample				Sibling Sample	
	(i)		(ii)			
Outcome	Marg. eff.	|t-ratio|	Marg. eff.	|t-ratio|	Coeff.	|t-ratio|
Education						
child's age 0–5	−0.119	2.141	−0.070	1.266	−0.264	2.561
child's age 6–10	−0.127	1.671	−0.062	0.806	−0.113	0.762
child's age 11–16	−0.228	2.291	−0.171	1.620	−0.102	0.611
equality p-value[a]	0.565		0.615		0.310	
Inactivity						
child's age 0–5	0.067	3.976	0.051	3.187	0.140	4.207
child's age 6–10	0.049	2.195	0.027	1.193	−0.078	1.610
child's age 11–16	0.009	0.330	−0.009	0.634	0.010	0.536
equality p-value[a]	0.105		0.069		0.001	
Early childbearing						
child's age 0–5	0.022	2.542	0.016	1.786	0.024	1.683
child's age 6–10	0.007	0.399	0.003	0.223	0.022	1.672
child's age 11–16	0.014	0.859	0.009	0.655	0.018	1.594
equality p-value[a]	0.543		0.632		0.245	
Distress						
child's age 0–5	0.052	2.004	0.055	2.059	0.044	1.309
child's age 6–10	0.072	1.836	0.076	1.919	0.068	1.402
child's age 11–16	0.036	0.778	0.031	0.660	0.040	1.243
equality p-value[a]	0.812		0.729		0.094	
Smoking 10+ a day						
child's age 0–5	0.070	2.282	0.054	1.745	0.049	2.185
child's age 6–10	0.039	0.814	0.037	0.751	0.021	1.478
child's age 11–16	0.134	3.210	0.138	3.183	0.046	0.847
equality p-value[a]	0.226		0.131		0.042	

Note: Estimates for both samples are obtained as explained in the footnotes of Table 5, except that proportion of mother's time worked during childhood (used in specification (ii) only) and family structure during childhood are broken down by the three developmental stages of the child.
[a] Figures are p-values of the test that the estimated coefficients are equal by developmental stage. The p-values are obtained from χ^2-statistic in individual sample and F-statistic in the siblings sample.

tern, but the impact is lower and around 5–7%. Again, although the parental loss occurred at early stages of life, it appears to have long-term consequences on young people's chances of being economically active (either in the labour market or in school).[27] Thus growing up in a disrupted family affects individuals' later success in at least two different ways, which may be related: it reduces young people's chances of obtaining higher levels of education and it increases their risk of inactivity (see also McLanahan and Sandefur, 1994, Chap. 3).

3.3. *Early childbearing*

Experience of life in a single-parent family during childhood is associated with significantly higher chances of an early birth: a young woman who had such an experience has a 1.8% per annum higher chance of early childbearing than a woman who did not (specification (i), Table 5). This association diminishes when the economic variables are included in estimation, but it is still quite large. Better economic circumstances appear, therefore, to play a role in reducing the risk of an early birth for women who grew up in a single-parent family but do not eliminate it.[28] A young woman's lower education and employment expectations (two characteristics that are common to women with experience in a single-parent family during their childhood) are likely to reduce her perception of the costs of early childbearing, and a birth may reduce her educational attainments. This illustrates the value of analysing young people's outcomes jointly. The estimates from the Sibling Sample are large and confirm that the positive association between family structure and early birth persists even when we control for mother's fixed effects. Having experienced a family breakdown increases the probability of early motherhood by 2.4% per annum (t-ratio = 1.69). While an early family disruption (when the girl was in pre-school years) exhibits a stronger association (both samples, Tables 6), the three stage-specific estimates from either sample are not statistically different from each other. While restating the importance of an early parental loss, this finding also suggests that family disruptions during school years and adolescence are likely to lessen monitoring of children's activities, and those are the times when parental supervision could prevent behaviour that leads to early childbearing (Thornton and Camburn 1987; Hill et al. 1998).

3.4. *Distress*

The level estimates show a positive and significant association between family structure and level of distress before and after including the economic variables in estimation: having spent time in a single-parent family during childhood increases the probability of having a high level of distress by 5.5%. This result is consistent with the large psychological literature in this area, which has shown that children and adults in non-intact families are at greater risk for psychological adjustment problems compared to those in families with both biological parents (e.g., Amato and Keith 1991; Bruce and Kim 1992). Family disruptions that occurred either in pre-school (ages 0–5) or in primary school years (ages 6–10) exhibit the strongest associations: for example, young adults whose parents separated when they were between 6 and 10 years of age are about 7% more likely to report high levels of distress than young adults who lived in an intact family during their entire childhood. After controlling for mother's fixed effects, however, the relationship between distress and family structure (at any developmental stage) is not statistically significant. The fact that this correlation is imprecisely measured when we account for unobserved heterogeneity may, however, reflect measurement error in our measure of distress rather than the absence of a robust relationship. It is well known that the presence of modest errors in variables can wipe out most of the associations of interest and that sibling differences exacerbate this effect of measurement error (Griliches 1979). The information on distress is, in fact, elicited in the self-completion questionnaire of the BHPS by a series of questions regarding the way the respondent has been feeling over the last few weeks. The exact phrasing is: "Have you recently ... (felt under strain, depressed, etc.)?". Given the subjective measure of the answer and the relatively short reference period (the last few weeks), this measure is likely to pick up a lot of noise and other (possible temporary) aspects of psychological well-being.[29]

3.5. *Smoking*

The association between family structure and smoking 10 or more cigarettes a day is always strong and well determined: having spent time with a single mother during childhood increases the probability of heavy smoking by about 7%. This association persists even after controlling for the economic conditions of the family of origin and falls only to 6.3%. Accounting for unobservable factors that are shared by children who belong to the same family (born to the same mother) confirms this finding, with the effect being 6.8%. Family disruptions that occurred either during pre-school years (both samples) or during adolescence (Individual Sample only) seem to have a stronger relationship with heavy smoking.[30] In the case of the Sibling Sample, we reject the hypothesis that the estimated coefficients are equal across the three developmental stages at conventional levels of significance. But this hypothesis cannot be rejected for the estimates obtained with the Individual Sample. In general, our findings confirm the evidence, documented in many social medicine studies, that children of single mothers have an increased risk of being smokers regardless of whether or not the mother smokes (Green et al. 1990; Turner-Warwick 1992). This suggests that, beside mother's smoking behaviour, other characteristics of single-parent families (such as lower parental control or lower expected human capital) foster children's smoking.

4. Conclusions

In this paper we estimate the relationship between several outcomes in early adulthood (educational attainment, economic inactivity, early childbearing, distress and smoking) and the experience of life in a single-parent family during childhood. We use a sample of young adults, selected from the first five waves of the BHPS (1991–1995), who can be matched with at least one sibling over the same period. This sample allows us to estimate the relationships of interest by sibling differences. We also perform our analysis on another sample which we estimate using a level (logit) model. These estimates are useful for comparison with the existing evidence.

We draw attention to four aspects of our findings. First, we show that sibling differences require a weaker assumption (as compared to the assumptions imposed by standard level estimators) for the identification of the family structure effect, namely that family structure is not correlated with children's idiosyncratic endowments, but it is still a strong one. Second, using such sibling differences, we find that experience of life in a single-parent family during childhood is usually associated with negative outcomes for children as young adults: lower educational attainments, higher risks of inactivity and early birth, and higher chances of being a heavy smoker and experiencing higher levels of distress in early adulthood. Third, family structure in early childhood (when the child was between the ages of 0 and 5) appears to be more important for shaping achievement, behaviour and mental health than does family structure during primary school years or adolescence. Fourth, the level estimates, whose causal interpretation relies on even stronger assumptions, confirm our results and are consistent with much of the evidence available in this literature. In addition, they allow us to show that the adverse family structure association generally persists even after controlling for the economic circumstances of the family of origin.[31]

Although we are only able to interpret these results as suggestive associations, the sibling-difference estimates control for more family background factors than has been usual in the literature. Identification of a causal impact must, however, await data which contain sufficiently convincing instruments that allow family structure to be modelled as an endogenous variable.

Appendix: The identification problem

Let j index family and i index individuals (or, interchangeably, young adults and children). For convenience, assume that the relationship that we estimate is given by the following linear probability model (see Angrist and Lavy 1996):

$$p_{ij} = \beta X_{ij} + u_{ij}, \tag{1}$$

where p_{ij} is a dichotomous variable indicating one of the outcomes, taking the value of 1 if the outcome under study occurs and 0 otherwise; X_{ij} is a vector of explanatory variables, such as age, family structure during childhood, mother's education and parental income (Sect. 2 gives a complete list of the variables used in estimation); u_{ij} is a random shock with zero mean. In this formulation, the parameters β are assumed to be the same for all individuals. Arguably, the effect of family structure is heterogeneous (i.e., some children might be better off in a non-intact family, while others might be worse off). Notice, however, that the methodology proposed here would apply even if one specifies a random-coefficients model in which $\beta_j = \beta + \phi_j$, and $E(\phi_j u_{ij}) = E(\phi_j X_{ij}) = 0$.

Our objective is to provide consistent estimates of the "effect" of family structure (contained in X) on the probability of various children's outcomes, p. Consistent estimation of β in (1) requires that the variables measuring mother's (or parents') behaviour during childhood contained in X be uncorrelated with the disturbance term u. We investigate this issue using a framework suggested by Behrman et al. (1994) and Rosenzweig and Wolpin (1995). Consider a two-child family. For the i-th child in family j with sibling k,

$$u_{ij} = \delta_1 \varepsilon_{ij} + \delta_2 \varepsilon_{kj} + \alpha_j + \mu_{ij} \tag{2}$$

$$\varepsilon_{ij} = \rho \varepsilon_j + v_{ij} \tag{3}$$

$$X_{ij} = \Pi \varepsilon_j + \gamma_1 \eta_{ij} + \gamma_2 \eta_{kj} + \theta_j \tag{4}$$

Equation (2) decomposes u_{ij} into four elements: a mother-specific fixed effect common to both siblings, α_j; two distinct stochastic components that depend on the endowments of each sibling, ε_{ij} and ε_{kj}; and measurement error, μ_{ij}. The parameters δ_1 and δ_2 capture the parental (or own) response to child endowments that are observable to all family members but are not observed by the econometrician. We assume that $E(\varepsilon_{ij}) = E(\mu_{ij}) = E(\alpha_j) = E(\varepsilon_{ij}\mu_{ij}) = E(\varepsilon_{kj}\mu_{ij}) = E(\alpha_j\mu_{ij}) = 0$, for all i, k, and j.

Equation (3) is a type of Galton's law of heritability of endowments (see Becker and Tomes 1986), with regression to the mean across generations ($0 \leq \rho < 1$); ε_j is the zero-mean parents' (or mother's) endowment and v_{ij} is an idiosyncratic disturbance with zero mean, and uncorrelated with ε_j and with v_{kj} (the analogous disturbance for sibling k).

Equation (4) relates the variables in X_{ij} to the parental endowment, ε_j, a mother-specific fixed effect, θ_j, and the idiosyncratic endowments of the children, η_{ij} and η_{kj}, where Π, γ_1 and γ_2 are conformable vectors of parameters. The parameters in Π capture the mother's (or parents') response to her (their) own endowment, while γ_1 and γ_2 measure the parental response to child-specific idiosyncratic endowments. Equation (4) allows for the possibility that aspects of family environment, including family structure, may be influenced by the family's as well as the children's endowments. We assume that $E(\eta_{ij}) = E(\eta_{kj}) = E(\theta_j) = E(\eta_{ij}\eta_{kj}) = E(\eta_{ij}\theta_j) = E(\eta_{ij}\varepsilon_j) = 0$, for all i, k, and j. We further assume that $E(\eta_{ij}\alpha_j) = E(\eta_{kj}\alpha_j) = E(\eta_{ij}\mu_{ij}) = E(\eta_{kj}\mu_{ij}) = E(v_{ij}\theta_j) = E(v_{kj}\theta_j) = E(v_{ij}v_{kj}) = E(\mu_{ij}\theta_j) = 0$.

This formulation introduces three different sources of family-specific heterogeneity: in Eq. (3), ε_j is transmitted through the endowments, ε_{ij} and ε_{kj}, to which parents' or individuals' behaviour (via δ_1 and δ_2) can respond; another component, θ_j in Eq. (4), affects children's outcomes indirectly through the parental behaviour measured by X_{ij}; the last source of heterogeneity, α_j, affects the outcome p_{ij} directly (through Eq. (2)), regardless of individual earnings endowments.

Substituting (3) in (2) yields $u_{ij} = \delta_1 v_{ij} + \delta_2 v_{kj} + (\delta_1 + \delta_2)\rho\varepsilon_j + \mu_{ij} + \alpha_j$. The level estimates of β in (1) are consistent if the covariance (vector) between X_{ij} and the error term u_{ij} is zero. However, from (4),

$$\operatorname{cov}(X_{ij}, u_{ij}) = \Pi(\delta_1 + \delta_2)\rho\sigma_\varepsilon^2 + \Pi\sigma_{\alpha\varepsilon} + (\delta_1 + \delta_2)\rho\sigma_{\varepsilon\theta} + \sigma_{\alpha\theta} + (\gamma_1\delta_1 + \gamma_2\delta_2)\sigma_{\eta v}, \tag{5}$$

where $\sigma_\varepsilon^2 = \operatorname{var}(\varepsilon_j)$, $\sigma_{yz} = \operatorname{cov}(y, z)$, for $y, z = \alpha_j, \varepsilon_j, \theta_j, \mu_{ij}, \eta_{ij}$, and v_{ij}. In general (5) is not zero, in which case estimates of β are inconsistent. Even after introducing other orthogonality assumptions, i.e., $\sigma_{\alpha\varepsilon} = \sigma_{\varepsilon\theta} = \sigma_{\alpha\theta}$, we still find that $\operatorname{cov}(X_{ij}, u_{ij}) = \Pi(\delta_1 + \delta_2)\rho\sigma_\varepsilon^2 + (\gamma_1\delta_1 + \gamma_2\delta_2)\sigma_{\eta v}$, and this covariance disappears only if: a) either $\Pi = 0$, or $\delta_1 = -\delta_2$, or $\rho = 0$; *and* b) $\sigma_{\eta v} = 0$, or $\delta_1 = -\delta_2$ and $\gamma_1 = \gamma_2$, or $\gamma_1 = \gamma_2 = 0$. It is not implausible that at least some of the variables in X_{ij} depend on the family endowment ε_j, that is, $\Pi \neq 0$. For example, mother's education depends on ε_j. If we believe that there exists a positive degree of "inheritability", through genetic and cultural transmission of endowments, then also ρ is non-zero. Furthermore, if we believe that parents respond to child earnings endowments, then $\delta_1 + \delta_2 \neq 0$. Behrman et al. (1982) and Ermisch and Francesconi (2000) do, however, present a family model of endogenous education in which $\delta_1 = -\delta_2$. For consistent estimates of β in this model, however, it is also necessary to assume that the variables in

X do not respond to child idiosyncratic endowments, so that $\gamma_1 = \gamma_2$.[32] That is, when $\delta_1 = -\delta_2$ is coupled with the orthogonality assumption $\sigma_{\alpha\varepsilon} = 0$, $\gamma_1 = \gamma_2$ is sufficient for identification of the parameter β.

Many of the problems of the level estimates have to do with the presence of mother-specific fixed effects.[33] A siblings estimator of the components of β for which the elements of X_{ij} differ between siblings is based on the differences between siblings, and such fixed effects would be eliminated via differencing. In our two-child family case, this is

$$\begin{aligned} \Delta p &= \beta \Delta X + (\delta_1 - \delta_2)\Delta\varepsilon + \Delta\mu \\ &= \beta \Delta X + \Delta\xi, \end{aligned} \tag{6}$$

where $\Delta z = z_{ij} - z_{kj}$, for any term z, and $\Delta\xi = (\delta_1 - \delta_2)\Delta\varepsilon + \Delta\mu$. From Eq. (4), it follows that $\Delta X = (\gamma_1 - \gamma_2)\Delta\eta$. Thus, the covariance between ΔX and the disturbance term in (6) is given by

$$\operatorname{cov}(\Delta X, \Delta\xi) = (\gamma_1 - \gamma_2)(\delta_1 - \delta_2)E(\Delta\eta\Delta v) + (\gamma_1 - \gamma_2)E(\Delta\eta\Delta\mu). \tag{7}$$

Our previous assumptions that $E(\mu_{ij}) = E(\mu_{ik}) = E(\eta_{ij}) = E(\eta_{kj}) = E(\eta_{ij}\mu_{ij}) = E(\eta_{kj}\mu_{ij}) = 0$ guarantee that the second term of (7) is always zero. We then only need to assume that either $\sigma_{\eta v} = 0$, or $\delta_1 = \delta_2$, or $\gamma_1 = \gamma_2$ to identify β. The conditions $\sigma_{\eta v} = 0$ (or, if $\eta_{ij} = v_{ij}$ as in Rosenzweig and Wolpin (1995) and Ermisch and Francesconi (1997), $\sigma_\eta^2 = 0$) and $\delta_1 = \delta_2$ are difficult to justify by theoretical arguments. It may, however, be plausible that many aspects of the family environment do not respond to children's idiosyncratic attributes, so that $\gamma_1 = \gamma_2$. This latter condition may be met by some family background variables used in the empirical analysis, like mother's education, because the children's idiosyncratic endowments are only apparent at their birth or afterwards. But other family behaviour, like mother's work patterns, may instead be the *result* of child-specific idiosyncratic attributes rather than the *cause* of the young adult's achievements. For this reason, with the Sibling Sample we estimate a model that contains only (sibling differences in) age, gender, and family structure during childhood. In other words, identification of the family structure "effect" with our data requires that family structure does not respond to, and is not correlated with, children's idiosyncratic endowments. The level (logit) regressions performed with the Individual Sample in specification (i) also contain year of birth and mother's education, variables that are likely to be insensitive to the child's idiosyncratic attributes. As most of the studies in this literature, in specification (ii) we add mother's employment patterns during childhood, mother's age at child's birth, family income, and other "economic" variables that control for the economic circumstances of the family of origin. This allows for a comparison of our estimates with those currently available in the literature, but imposes stronger identifying assumptions.

In sum the identification of the family structure effect on children's achievements rests both on the availability of "prior information" about the process generating parental behaviour and children's outcomes, and on the researchers' willingness to make specific assumptions on such a process. We share Manski et al. (1992) view that "as long as social scientists are heterogeneous in their beliefs about this process, their estimates of family structure effects may vary" (p. 36).

Endnotes

[1] For a detailed overview of existing studies, see McLanahan and Sandefur (1994), Haveman and Wolfe (1995) and Mayer (1997).

[2] Throughout the paper, the terms "mother-specific fixed effects" and "family-specific fixed effects" are used interchangeably, because of the data that we analyse. But it should be emphasized that our fixed-effects estimator eliminates the influence of all unmeasured persistent mother, family and community characteristics that do not differ by siblings.

[3] Recent exceptions are the longitudinal studies by Duncan et al. (1997) and Blau (1999) on the effect of family income on schooling and child development, and by Grogger and Ronan (1996) on the effect of fatherlessness on education and wages.

[4] Recent studies using BHPS data similar to those employed here include Ermisch and Francesconi (1997), who investigate the association of several family structure measures on educational attainments, and Ermisch and Francesconi (2000), who analyse the relationship between family background and young people's earnings.

[5] For data on household income, we use information from fathers, stepfathers or other adults, if present in the household.

[6] On the other hand, the NCDS does have larger sample sizes and more measures of non-economic background factors.

[7] By collecting information on individuals born in the first week of March 1958, the NCDS data are instead not suitable for this estimation procedure. By 1991 in the NCDS sample used in Dearden (1998), there are only 27 pairs of twins.

[8] This finding has never been emphasised before in the British context, partly because the NCDS data could not reliably identify family breakdowns from birth up until age 5. As a result, most of the studies with NCDS data use samples that are typically restricted to individuals whose parents were in an intact family at age 7. An exception is Frostin et al. (forthcoming). Some of their regressions are estimated for a sample of NCDS children aged 33 in 1991, which included those whose parents were not in an intact family either at birth or at age 7 of the child.

[9] Of those interviewed in wave 1 (1991), 88% were re-interviewed in wave 2. The wave-on-wave response rates from the third wave onwards have been consistently above 95%. The BHPS data are therefore unlikely to suffer from serious attrition bias.

[10] The first five waves of the BHPS indicate that 93% of single-parent families are headed by the mother and that 86% of dependent children living with a step-parent lived with their natural mother.

[11] If the birth occurred outside of a partnership and the mother partnered within one year, we assumed that the mother had moved in with the biological father (as assumed in Bumpass et al. (1995) and Ermisch and Francesconi (forthcoming)). For adopted children, we use information on the year in which they were adopted to match in the mother's family history appropriately. In 96% of the cases, the children are natural children.

[12] Ermisch and Francesconi (1997) experiment with other, more detailed measures of family structure, e.g., step-families and durations of different family structures. But this simple dichotomy by developmental stage performs as well in predicting educational attainment as more complex measures. In addition, Ermisch and Francesconi (forthcoming) find that only 242 women had a pre-partnership birth as of 1992 (wave 2), representing 0.05% of all women in that survey year. Since we cannot determine whether or not they subsequently lived with the child's father, we assume that the women who formed a union within one year of the birth did so with the father. Of the 242 women who had a pre-partnership birth, 77 (32%) lived with the child's father. Because of the small sample sizes, therefore, we cannot explore the distinction between children who experienced a family disruption when age 0–5 and children born into a single-parent family.

[13] That is, the third wave contains retrospective information on jobs held by all adults since they left full-time education, including their current work if it started before September 1990, up to

September 1990. For jobs started after that date, we use information collected in the panel wave-on-wave work history.

[14] We performed a similar analysis using more complex measures of mother's employment patterns over her child's childhood, including the proportions of months worked in full-time and part-time jobs and the proportion of months worked in broad occupation groups. Our main results are unchanged.

[15] While the overall figure may appear high, it is consistent with life table estimates by Ermisch and Francesconi (forthcoming), which indicate that 40% of mothers in the BHPS will spend some time as the only parent. The stage-specific figures may also appear large in the light of divorce registration statistics which indicate that about 40% of dependent children of divorcing parents are aged under 6. Recall, however, that single-parent families are also formed by the dissolution of cohabitations and births outside of a partnership. In the BHPS data, these last two categories account for 35% of instances of single parenthood. If, in conjunction with the divorce registration statistics, we assume that all dissolving cohabitations involve children under 6, then 60% of instances of single parenthood would start when the child is aged under 6.

[16] The 165 two-sibling households give rise to 165 comparisons, one per siblings' pair; the 23 three-sibling households produce 69 comparisons, 3 in each household; and the 3 four-sibling households give rise to 18 comparisons, 6 in each household.

[17] For readers unfamiliar with the British education system, "A(Advanced) level" corresponds to education beyond high school, but short of a university degree. At least one A level is necessary to be admitted to a university.

[18] In line with official statistics, data from the original representative BHPS sample show that approximately 20% of women have a baby by age 21. Registration statistics also indicate that, for the most recent cohorts of women for which we have data (women born in the late 1960s), the median age at first birth was 27 (ONS 1997, Table 10.3). The original BHPS data produce a similar figure.

[19] Distress is measured in comparison with "usual" conditions. The subjective indicators are: (i) loss of concentration; (ii) loss of sleep; (iii) playing a useful role; (iv) capable of making decisions; (v) constantly under strain; (vi) problem overcoming difficulties; (vii) enjoy day-to-day activities; (viii) ability to face problems; (ix) unhappy or depressed; (x) losing confidence; (xi) believe in self-worth; (xii) general happiness. Each indicator is measured over a scale that runs from 1 to 4. Recording 1 and 2 values on individual indicators to 0, and 3 and 4 values to 1, and then summing over all indicators gives a new scale running from 0 (the least distressed) to 12 (the most distressed). The scale is known in the health literature as *caseness.* For each young adult, our mental health measure takes the value of one if his/her caseness variable has a value of 3 or more, and zero otherwise. See Cox et al. (1994). Because collapsing the distress scale to a dichotomous variable may reduce the outcome variability, we also perform our analysis treating it as a continuous measure.

[20] The cut-off choice for the number of cigarettes smoked in a day is arbitrary. We also performed the analysis with a variable taking value of one if the respondent smokes and zero otherwise. The results are qualitatively identical to those reported below.

[21] A well-known problem inherent to all linear probability models is that the predicted outcomes are not constrained to lie between zero and one. This may obviously compromise the interpretation of the estimates as the probability that the event under study will occur. To gauge how serious this problem can be with our data, we first estimate our five outcomes with the Individual Sample using a linear probability model. We then use the estimated coefficients to compute the conditional expectation $E(y|x)$, where y is an outcome and x is the appropriate vector of explanatory variables. We predict, at most, 2 cases outside the unit interval for the education outcome (that is, 0.41% of the 489 observations used in this estimation), and 6 cases outside the unit interval for the inactivity outcome (0.25% of the observations). The maximum number of predictions that do not lie between zero and one is always lower for the other three outcomes. Therefore, the prediction problem of the linear probability model is arguably inconsequential for the samples used in this study.

[22] For expositional purposes, the model in the Appendix does not allow for heterogeneity in the family structure effect.

[23] There are six age dummies (age 16 is the base) in the Individual Sample. Mother's education is grouped in four categories: no qualification (base), O level, A level, and higher qualification. In the Sibling Sample, age enters non-parametrically only in the case of early childbearing, i.e., when sisters' comparisons are made at a common age.

[24] We distinguish three stages of mother's age at birth: young (maternal age less than or equal to 21), middle (maternal age between 22 and 33, base); and old (maternal age greater than or equal to 34).

[25] As a way of checking whether the differences between individual-based and sibling-based estimates arise from the different age distributions in the two samples (see Table 2), we have also estimated level models using the Sibling Sample for all outcomes. In general the point estimates are very close to those found with the Individual Sample while the standard errors are somewhat larger (presumably due to the smaller sample sizes). This suggests that the differences in results across samples do not systematically depend on their different age distributions. For example, the level estimates (in terms of marginal effects) of the education outcome found with the Sibling Sample are -0.138 (t-ratio $= -2.187$) and -0.122 (t-ratio $= 2.042$) for specification (i) and (ii), respectively.

[26] Duncan et al. (1998) find that "early childhood appears to be the stage in which family economic conditions matter most" (p. 420). Our findings are consistent with their result, given that the impact of family structure on educational attainment may operate partly through family income.

[27] For the estimates obtained from specification (ii) of the Individual Sample, we can reject the hypothesis of equality of the estimated stage-specific coefficients at the 10% level but not at the 5% level. For the estimates from specification (i), we instead never reject the equality hypothesis.

[28] The results from specification (ii) also reveal that if the mother herself was aged 21 or less when her daughter was born, then the odds that the daughter has an early birth are higher. Thus, there is evidence of a recurrence of early motherhood across generations. This association becomes slightly weaker when the economic circumstances of the family of origin are controlled for. On the other hand, if mothers were aged 34 or more at birth, the chance of their daughters having an early birth are lower.

[29] When the distress measure is treated as a continuous variable, we again find positive and well-measured estimates of family structure using the Individual Sample. For example, in comparison with the results reported in Table 5, having spent some time in a single-parent family during childhood increases the distress level (which on average equals 1.364) by 0.384 points (t-ratio $= 2.864$) in specification (i) and by 0.388 points (t-ratio $= 2.843$) in specification (ii). The estimates obtained from the Sibling Sample are always positive, but somewhat smaller and never statistically significant. In sum, these results provide evidence comparable to that found with our dichotomous measure.

[30] We also find a high persistence in smoking behaviour across generations. In the Individual Sample, a young adult whose mother smokes has approximately 15% higher chances of being a heavy smoker than a young adult (with similar characteristics) whose mother does not smoke. This association is always significant at any conventional level. The inclusion of mother's smoking behaviour, however, may be problematic for identification of the family structure effect (see Sect. 2.3 and the Appendix). Its exclusion from the level regressions does not alter our results.

[31] Discussing the findings from twelve different papers, McLanahan (1997) argues that growing up in a single-parent family has negative consequences for children's well-being across several domains (e.g., test scores, education, behavioural and psychological problems, jobs and income) even after taking family income into account.

[32] We discount the unlikely case that $\sigma_{\eta v} = 0$.

[33] Given the data analysed in this study and the setup of the model presented here, the definition of "mother-specific fixed effect" is equivalent to that of "family-specific fixed effect". In particular, our fixed-effects estimator captures all persistent mother, family and community effects that do not differ by siblings. The two labels are therefore equivalent and used interchangeably throughout the paper.

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Childhood family structure and young adult behaviors

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Abstract. This paper examines a wide variety of forms, and full histories, of family structure to test existing theories of family influences and identify needs for new theories. The focus is on links between childhood family structure and both completed schooling and risk of a nonmarital birth. Using a 27-year span of panel (PSID) data for U.S. children, we find that: (a) change is stressful, (b) timing during childhood is relevant, (c) adults other than parents are important, and (d) two more recently studied family structures (mother-with-grandparent(s) and mother-with-stepfather) do not fit the molds of existing theories. The findings suggest that new theories should consider allocation of resources and reasons people group into family structures.

JEL classification: J12, J13, J16

Key words: Demographic economics, marital dissolution, family structure

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1. Introduction

In recent decades, new forms of family structure have assumed prominence in the lives of U.S. children. Aspects of these changes have been documented repeatedly in cross-sectional studies but less fully explored in a panel context.

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Much remains to be learned about what they mean for children. Fortunately, panel studies are now yielding data tracing children and their family life from birth into young adulthood, all during this historical period of greater diversity in family structure. This affords a unique opportunity for a better understanding of how family life operates, the relevance of existing theories of family influences, and new directions for identifying underlying mechanisms.

A particular focus of this paper is the importance of childhood stage to the influence of family structure on children. Taking account of possible variation in effects of family structure by timing of occurrence allows for the possibility that a particular family structure mechanism, be it stress from disruption in family routines, change in the amount or quality of parental supervision, or low levels of economic resources, has stronger effects at some stages of childhood than others because of differences in children's developmental needs or susceptibility to problem behavior. Caring and nurturing, along with nutrition and a good learning environment, are viewed by developmentalists as crucial in early childhood, whereas parental supervision and emotional support may be important in late childhood when risks of behaviors such as dropping out of high school or having a nonmarital birth come into play. Information on the full history of the family structures children encounter while growing up allows us to test the sensitivity of the influence of the family structure to the ages when children encounter it.

The paper also provides a novel perspective on mother-only and non-intact structures by highlighting living arrangements that combine a single mother with grandparent(s) or combine a mother and stepfather. Though still relatively rare, these structures appear to be on the rise. Casper and Bryson (1998), for example, report a 118% increase between 1970 and 1997 in the number of children living in mother-with-grandparent(s) households. However, time series estimates of the proportion of children in mother-with-grandparent(s) and in mother-with-stepfather families are difficult to locate, and, as far as we can tell, there are no published figures regarding proportions of children ever in such structures, much less what parts of their childhood they spent that way. Research attention is beginning to be directed both to the factors that encourage these living arrangements (e.g., Mauldon and Maestas 1998) and to ways these family structures affect children and adults.

There is clear interest in mother-with-stepfather families because of high divorce and remarriage rates. Interest in mother-with-grandparent(s) families is piqued by the beginnings of a complicated picture of social and economic differences associated with this family structure. Research suggests that grandparent-grandchild coresidence tends to be associated with higher satisfaction on the part of adolescents with their parental relationships, less delinquent activity, but lower grades (Kirby and Uhlenberg 1999).

Observational and ethnographic studies suggest that the added adults in multigenerational families do not necessarily enhance the supervision of youth and can create confusion over who is in charge (Chase-Lansdale et al. 1994; Pattillo-McCoy 1999). Such studies also suggest that children in multigenerational families tend to take on adult roles more quickly because the generations are close in age and this blurs developmental role boundaries to the point that children and parents may behave as siblings (Burton, Obeidallah and Allison, 1996). Analysis of Current Population Survey data indicates that children coresiding with grandparent(s) tend to fare worse economically (Casper and Bryson 1998). All told, however, surprisingly little is known about mother-with-grandparent(s) structures. Although our results regarding these noteworthy but rare structures should be viewed cautiously owing to modest sample sizes, they are provocative and call for further research.

Our investigation emphasizes the timing of membership in each family structure in two regards: childhood stage and length of elapsed time. We also examine transitions between structures. Other researchers (e.g., Wu and Martinson 1993; Martinson and Bumpass 1990; and Bumpass and McLanahan 1989), using different data and specifications, have investigated aspects of non-intact or mother-only structures, but they have not specifically examined the mother-with-grandparent(s) structure, nor have they examined the mother-with-stepfather structure at different childhood stages.

Data from the Panel Study of Income Dynamics (PSID) are the basis for our empirical investigation. These data span all years from birth through young adulthood for a representative sample of cohorts born in the late 1960s and early 1970s. With these data, and a focus on family influences on children's educational attainment and daughter's risk of non-marital childbirth, we find some support for existing theories about the ways families influence children. But much of what we find is at odds with these theories.

Our analysis indicates that the stress and economic hardship associated with other-than-two-parent family structures are more important than social control. However, some of our evidence is not well accounted for in any of these theories: (1) parental remarriage and having a stepfather present tend to have beneficial consequences for daughters if these things occurred during adolescence, (2) parental divorce or separation experienced in adolescence tends to have a positive association with sons' education, and (3) the family structure with the greatest detrimental consequences for children, particularly if experienced during adolescence, appears to be one with grandparent(s) present along with a single mother. In addition, variations in consequences by timing of the family structure over the course of childhood do not always fit well with the implications of existing theories.

There is a clear need for further research and for further development of theories about family mechanisms. Our findings suggest that new theories need to take into account the timing of family structure and events, as well as integrate a variety of dimensions, including factors associated with the reasons people group into particular family structures. We leave the task of formulating new theories to future research, but sketch out the terrain they should cover.

The paper begins with a review of existing theories and evidence on the mechanisms by which family structure affects children's characteristics and behaviors. We then focus on our analysis of the PSID data. To set the stage we first examine patterns of childhood family structure experiences. Next we describe the approach and measures used for determining the relevance of existing theories. Estimates for the models are then presented, with discussion highlighting both results that fit and results that do not fit the theories. We end with a discussion of the implications of the findings and further needs for research, with a focus on developing theories and some important aspects to consider when formulating such theories.

2. Existing theories and relevant literature

The social science literature posits a number of different causal mechanisms relating family structure to children's outcomes. Our analysis focuses on two of the major theories – stress and social control, and takes into account another possible mechanism: family income as emphasized by economic resource theory.

Stress theory posits *change* in family life as the central cause of family structure effects on children, the idea being that change in family structure prompts reorganization of the roles of family members and adversely affects the nurturing and support provided by parents. (For relevant discussion, see McLoyd et al. 1994; Conger et al. 1993, 1992; Wojtkiewicz 1993; Cherlin et al. 1991; McLanahan 1988; Booth et al. 1984; and Elder 1974). Key aspects of family life for this theory are parental marital events. Family reorganization prompted by parental divorce or (re)marriage is viewed as stressful to parents and children, and the resulting weakening of emotional security and bonds is thought to encourage problem behaviors in children.

Social control theory views adult supervision and monitoring of children's behaviors as important means by which children are kept from engaging in problem behaviors. (For relevant discussion see Brooks-Gunn, forthcoming; Chase-Lansdale et al. 1994; Thornton 1991; McLanahan and Bumpass 1988; Steinberg 1987; Thornton and Camburn 1987; Dornbusch et al. 1985; Hogan and Kitagawa 1985; Hetherington 1979; Cherlin 1978; and Mueller and Pope 1977). Key aspects, according to this theory, are number and types of adults overseeing children. Social control is thought to increase with the number of adults present in the child's home. The more distant the relationship of the adult to the child, though, the weaker the social control. Some types of parents (stepparents) or substitute parents (grandparents) are likely to exert less authority and social control than biological parents because of their more tenuous relationship to the child and because the parenting roles of such relatives are ill-defined (Brooks-Gunn forthcoming; Cherlin 1978). In childhood homes containing both biological parents and grandparents, disagreements between the two regarding parenting style may undermine the social control exerted by both (Chase-Lansdale et al. 1994).

Family income is the operative mechanism in economic hardship theory. (For relevant discussion, see Dodge et al. 1994; Zill and Nord 1994; McLeod and Shanahan 1993; Axinn and Thornton 1992; DaVanzo and Goldscheider 1990; Goldscheider and DaVanzo 1989; Goldscheider and Goldscheider 1987; Weiss 1979; Rubin 1976; and Elder 1974.) Family income is likely to vary with family structure (mother-only families tend to have lower income than two-parent families) and to change with its changes (children's family income tends to drop substantially after parental marital disruption and rise with parental (re)marriage). It can be difficult to distinguish effects of income from other influences of family structure without precise and comprehensive measures of both over the entire course of childhood, and past research has rarely had such measures.

Correlational evidence repeatedly indicates that children fare much worse when raised in non-intact homes (see McLanahan and Sandefur 1994 and Seltzer 1994 for reviews of past research, and, in McLanahan and Sandefur 1994, a comprehensive multi-dataset analysis). Evidence suggests that these correlations are not merely due to lack of control for measured and unmeasured family characteristics (Sandefur and Wells 1997). That children from non-intact homes fare worse could be consistent with social control, stress, and economic hardship theories.

Evidence relating more directly to social control theory indicates that children in single-parent families are more susceptible to peer pressure than children in two-parent families (Steinberg 1987; Dornbusch et al. 1985). Lack of social control and more emphasis on peers could result in children having more disciplinary problems and, in adolescence, more intensive dating and sexual involvement. These behaviors, in turn, may encourage young people to leave school early and to form families or unions prematurely, and the effects may be strongest in adolescence. Studies have shown that parental marital disruption leads to early and more frequent sexual activity, premarital pregnancy and births and to the early formation of unions, and some of the evidence suggests that lack of parent supervision is one of the operative mechanisms (see, for example, Thornton 1991; McLanahan and Bumpass 1988; Thornton and Camburn 1987; and Hogan and Kitagawa 1985).

Research regarding stress theory shows that, at least among adolescents, there are linkages between disruptive family events, parents' depression, impaired parenting behavior, and children's impaired school performance, social behavior, and self-esteem (McLoyd et al. 1994; Conger et al. 1992, 1993). Furthermore, Wu and Martinson (1993) and Wu (1996) find that it is change in childhood family structure rather than a prolonged period living in a mother-only family that is most strongly linked to young women's chances of having a premarital birth. Wojtkiewicz (1993) also finds that transition into a mother-only family is more important to chances of high school graduation than is duration of time spent in a mother-only family. Some of Wojtkiewicz's findings about other non-intact structures, however, are not entirely consistent with the theory that changes are more important than the length of time spent in a non-intact family.

Few studies of the effects of growing up in single-parent households have had access to sufficiently reliable measures of family income and other family-process measures to provide a complete accounting of why children raised in single-parent families do so much worse than children from two-parent families. The handful that have, however, tend to indicate that economic differences account for a sizable portion, but not all, of the adverse effects of being raised in a single-parent household. For example, McLanahan (1985), Hill and Duncan (1987) and McLanahan and Sandefur (1994) all find that parental-income differences account for between one-third and two-thirds of the estimated impact on completed schooling of living in a single-parent family. Other studies find that income differences play a less important role (e.g., Sandefur et al. 1992) or a complex role that varies by type of non-intact family, accounting for single-parent influences but not influences of mother-with-stepfather (Boggess 1998).

The literature indicates differential influences of family structure by sex and age of the child, as well as race (e.g., Boggess 1998). Married parents treat sons and daughters differently, and this may be a factor in the sex differences in effects of changes in family structure (see Seltzer's, 1994, review). Hetherington (1979, 1987) found sex differences in reaction to both parents' marital disruption and remarriage. Soon after parental divorce, as well as several years later, sons in families where the divorced mother did not remarry displayed a number of problematic behaviors, including noncompliant behavior. Sons were better adjusted if the custodial mother did remarry; however, daughters were better adjusted if the mother did not remarry. Findings such as these, however, are not entirely consistent with other research (see Seltzer 1994 for a review).

Findings are mixed regarding variation in the influence of family structure by child's age (childhood stage). As summarized in Wojtkiewicz (1993), both the amount of time young children spend in the home and their inability to understand and cope with marital disruptions lead us to expect family composition changes occurring early in a child's life to be most harmful, and some research bears this out (Krein 1986; Krein and Beller 1988). However, Wojtkiewicz finds that years spent in mother-only families between the ages of 11 and 15 as opposed to younger ages were associated with substantially reduced chances of graduating from high school. McLanahan and Sandefur (1994) report an insignificant six-percentage-point increase in the risk of dropping out of high school if a marital disruption occurred before age 6 as opposed to after age 12. Haveman and Wolfe (1994) find no significant effect on completed schooling of the timing of years spent living with one parent (although their analysis omits the very early years of childhood).

With such diverse findings in the literature, there is no clear picture of the mechanisms involved and policy implications of the detrimental effects of shifts away from intact families. As some researchers (e.g., Seltzer 1994; Wu and Martinson 1993; and Wojtkiewicz 1993) are beginning to show, our understanding of the underlying processes may be clouded by the failure of empirical measures to reflect the dynamics and many-faceted nature of family structure. Much of the research on childhood family structure has approached the issues as if children experienced the same family structure throughout childhood. Wu (1996) and Wu and Martinson (1993) in their dynamic analysis of young women's premarital births and Wojtkiewicz (1993) in his dynamic analysis of high school graduation help to correct some of these problems. However, this research does not include recent cohorts of young adults and lacks comprehensive control for income.

Wu and Martinson investigate cohorts born 1938–1969 and control for possible cohort differences using only an additive control distinguishing three subgroups – those born 1938–1947 vs. 1948–1957 vs. 1958–1969. Wu (1996) and Wojtkiewicz (1993) examine cohorts born 1958–1965. Given the dramatic increase in non-intact families, effects for earlier cohorts may well differ from those for more recent ones. Wu and Martinson and Wojtkiewicz include no measures of childhood family income, and Wu (1996) includes only measures of family income during adolescence.

The research presented in this paper extends the boundaries of the territory charted by these authors by: (a) examining more recent cohorts of children, (b) measuring family structure with shorter recall, (c) considering structures not yet investigated or little researched (mother-with-grandparent(s) and mother-with-stepfather), and (d) incorporating more comprehensive measures of family income that are well matched to the family structure measures. In extending this work we pay particular attention to: (1) the stage of childhood when a family structure or change in family structure was experienced, (2) presence of adults other than parents (e.g., step-parents or grandparents), (3) parental marital changes as well as associated family structure, (4) family income throughout the course of childhood, and (5) other types of changes that accompany family structure changes (e.g., movement to a different residence, which McLanahan and Sandefur 1994 find important, or changes in mothers' work commitment and consequently in the amount of time spent with their children, which Seltzer, et al. 1989, document as important). We also focus on two types of outcomes for children as young adults – educational attainment and, for daughters, nonmarital childbirth.

3. Family structure experiences

The dramatic growth in mother-only families has fueled much of the research on family structure in recent decades (see, for example, Duncan et al. 1994; Hernandez 1993; Moffitt and Rendall 1993; Duncan and Rodgers 1990; Hofferth 1985; and Bumpass 1984). Yet some of this growth may reflect shifts to family structures that are more complex than a single-parent situation. The presence in single-parent families of adults other than parents rarely has been taken into account, and a family structure consisting of grandparents along with the mother has often been classified as mother-only by researchers.

3.1. Data

The Panel Study of Income Dynamics (PSID) data allow us to investigate a wide range of possible family structures and track children's experience of family structures from birth through late childhood. For observing the complete childhood experiences, as well as for the tests of family structure mechanisms reported later in this paper, we rely on 27 years of PSID data. Since 1968 the PSID has followed and annually interviewed a representative sample of about five thousand families. (See Hill 1992 for a full description of the dataset.) Splitoff families are formed when children leave home, when couples divorce, and when more complicated changes break families apart. This procedure produces an unbiased sample of families each year, as well as a continuously representative sample of children born into families. The survey's original design focused on poverty by oversampling lower-income and minority households. Our sample consists of the 1,325 PSID individuals born between 1967 and 1973 and present[1] in the PSID every wave from birth to age 20. Since this includes individuals from the oversampled households, the data are weighted to adjust for this feature of the sample design as well as differential nonresponse. Barring any nonresponse bias remaining uncorrected by the weighting adjustments we employ, the experiences of this group of children are nationally representative of the cohorts from which they were sampled.[2]

3.2. Family structure measures

Our family structure measures are constructed to facilitate investigation of structures that have received relatively little attention in the literature. We distinguish between different types of structures with only one parent, and we separate structures containing a stepparent from other two-parent structures. Our measures are based on demographic information provided in the main data files from the 1968–1991 interviews, plus data from the 1968–1985 Relationship File, which consolidates many years of data to determine all possible pairwise relationships.

Our child-based family structure measures are constructed from the PSID's annual (time-of-interview) categorization of the following family types:

a) "two-parent family": child living with both biological or adoptive parents;[3]
b) "mother-only family": child living with the biological or adoptive mother and no other person older than age 21 other than a sibling;

c) "mother-with-stepfather family": child living with the biological or adoptive mother and her husband or cohabiting partner who is not himself the biological or adoptive father;[4]
d) "mother-with-grandparent(s) family": child living with the biological or adoptive mother and at least one grandparent but not with the biological, adoptive, or stepfather; other adults may be present in the household; and
e) "other living arrangement": consisting primarily of child living with father only, relatives other than parents or grandparents, or other nonrelatives.

These year-by-year distinctions are used to construct both measures of childhood family status and dynamic formulations of the sequence of family statuses experienced over childhood.

Our most basic status measure is "whether ever in a non-intact family," a crude assessment of childhood family structure frequently used by researchers. A child is classified as being in a non-intact family at some time during childhood if at the time of interview in at least one year from birth to age 15 the child was living in any type of family structure other than a "two-parent family." While this identification procedure misses experiences of non-intact structures in place less than one year and between interviews, it does capture most children's exposure to non-intact family structures.

Because non-intact family types may vary in their influence on children, we provide a finer breakdown by type of non-intact family. Our more detailed set of "whether ever in various family types" measures makes distinctions among the four types of non-intact families listed above, assessing whether a given type of family structure occurred at any time over the entire 15-year period of childhood.

Our set of measures labeled "whether ever in various family types in each developmental stage" preserves the distinctions among the four types of non-intact families but provides a breakdown by childhood stage. This set of measures assesses incidence within a single developmental stage, distinguishing between three different stages – early childhood (birth to age 5), middle childhood (ages 6–10) and late childhood (ages 11–15).

We also develop a "sequence" measure of family structure based on the basic set of annual assessments of family status. This dynamic measure abandons the distinction of developmental stages and categorizes the entire 15-year period of childhood according to the flow among different types of family status. The categories for the "sequence" measure are:

a) two-parent all 15 years;
b) mother-only all 15 years;
c) mother-only to 2-parent (counting stepparents), continuing until age 15;
d) mother-only to 2-parent and back to mother-only;
e) 2-parent to mother-only, continuing until age 15;
f) 2-parent to mother-only and back to 2-parent, and
g) other sequences, including ones with more than two transitions and ones involving relatives other than the biological father and mother.

These distinctions emphasize transitions between, as opposed to status in, different types of family structures. They capture transitions to and from non-intact families as well as permanent residence in a non-intact family.

Finally, our dynamic measures of family structure also include direct measures of potentially disruptive parental marital events. We focus on two types of events – parental divorce or separation and parental marriage or remarriage. These events entail the loss or gain of parent figures, and such events can lead to considerable disorganization and major alterations in family roles. We make distinctions about the timing of these events by specifying the childhood stage – ages 0–5, 6–10, and 11–15 – when the events occurred. Both events could occur in a single childhood stage, and each event could reoccur in different childhood stages. Our specification makes allowance for these possibilities. We also include a variable representing the type of family into which the child was born – whether a two-parent family (as the excluded category), a mother-only family, or some other form of non-intact family. This variable establishes the initial conditions for family structure.

3.3. Patterns of family structure

We begin with a comparison of these PSID-based data to independent sources. Weighted descriptive statistics on the PSID's patterns of childhood family structure are presented in Tables 1 and 2. Our own search of the literature and correspondence with prominent scholars in the field provided surprisingly little in the way of independent estimates of these kinds of data. Few sources of published U.S. data include distributions of children across finely detailed categories of family structure. We were unable to locate any source that provided comparable information by single year of age of child or that tracked structures longitudinally for children. All sources with fine delineation of family structure involved cross-sectional estimates with children aggregated into a group aged 1–17. Hence, our reliability checks involved using the PSID to construct simulated cross-sectional estimates applicable to children ages 1–17 in 1980 from the distributions by single year of age and race provided in our Table 1.[5]

Other sources provide data on the percentage of children not living with both parents, on the percentage living in 'mother-only' homes (an aggregation of our 'mother-only' and 'mother-with-grandparent(s)' categories), and on the percentage living as stepchildren. We were unable to find an independent source of national figures for frequencies of children in mother-with-grandparent(s) arrangements. In addition, we encountered difficulties finding reliable sources for frequencies of children in stepchildren arrangements.

For the percentage of children not living with both parents, we find roughly similar percentages with our simulated cross-sectional estimates compared with estimates from other sources. Our estimate of 16.2% for non-blacks is comparable to Cherlin's (1988) [CPS-based] estimates of 14.6% (1975), 17.3% (1980) and 20.0% (1985) for whites. Similarly, our estimate of 53.5% for blacks is roughly comparable to his estimates of 50.6% (1975), 57.8% (1980) and 60.5% (1985) for blacks. There is similar correspondence between our overall estimate of 21.7% for children as a whole and the Hernandez (1993) [Census and CPS-based] figures of 17.5% (1970), 23.4% (1980) and 28.8% (1988).

Our simulated estimate of the percentage of children in a broad category of 'mother only' family (a category combining our 'mother-only' and 'mother-with-grandparents' types) is somewhat lower than that estimated in Hernandez (1993). We estimate 13.5% of children to be in this broad category of

Table 1. Percentage distribution of family structure by age during entire childhood, by race of head

Nonblacks (n = 787)

Age	1	2	3	4	5	6	7	8	9	10	11	12	13	14	15
Two parents*	91.2	91.4	90.2	89.6	88.7	88.1	87.1	86.2	84.3	83.3	83.0	81.1	79.5	78.4	76.6
Mother only	4.0	4.5	6.0	5.9	6.6	6.8	6.9	8.1	9.0	9.5	8.3	8.2	9.3	8.9	8.6
Mother w/ stepfather	0.3	0.0	0.4	0.5	1.2	1.6	2.3	2.9	3.6	4.1	4.9	5.6	6.3	6.8	7.6
Mother w/ grandparents	2.3	1.7	1.5	1.1	0.7	0.6	0.6	0.3	0.3	0.3	0.3	0.3	0.3	0.8	1.0
Father only	0.0	0.0	0.0	0.3	0.5	0.6	0.9	0.9	0.9	0.9	0.9	1.7	1.4	1.5	2.1
Father w/ stepmother	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.1	0.1	0.3	0.6	0.8	1.1	1.4	1.7
Other arrangements	2.2	2.4	1.9	2.5	2.1	2.2	2.2	1.4	1.7	1.7	2.0	2.2	2.1	2.3	2.3
Total	100%	100%	100%	100%	100%	100%	100%	100%	100%	100%	100%	100%	100%	100%	100%

Blacks (n = 538)

Age	1	2	3	4	5	6	7	8	9	10	11	12	13	14	15
Two parents*	51.4	54.7	56.9	56.9	54.4	53.2	52.0	51.7	49.0	46.2	42.5	40.3	39.9	39.7	36.9
Mother only	18.5	21.9	23.4	24.7	29.2	31.7	33.6	34.5	36.8	39.2	40.9	43.1	42.8	46.9	48.0
Mother w/ stepfather	1.3	1.3	0.0	0.8	1.5	2.9	3.4	4.0	3.9	4.0	4.8	4.8	4.9	4.4	4.8
Mother w/ grandparents	17.0	11.8	9.7	8.8	5.8	6.3	4.1	3.8	4.2	4.1	5.3	4.3	4.7	2.7	3.0
Father only	0.0	0.0	0.0	0.0	0.1	0.7	0.7	1.1	1.1	0.7	0.7	0.7	0.7	0.8	0.8
Father w/ stepmother	0.1	0.1	0.2	0.2	0.2	0.2	0.2	0.2	0.2	0.2	0.2	0.2	0.2	0.2	0.4
Other arrangements	11.7	10.3	9.8	8.6	8.7	5.0	6.0	4.7	4.8	5.5	5.6	6.6	6.8	5.3	6.0
Total	100%	100%	100%	100%	100%	100%	100%	100%	100%	100%	100%	100%	100%	100%	100%

* Includes biological and adoptive parents.
Source: Panel Study of Income Dynamics.

Table 2. Proportion of children with various family structures ages 0–15, for children b tween 1967 and 1973, by race and sex

	Sons			Daughters		
	All (n = 699)	Nonblack (n = 421)	Black (n = 278)	All (n = 626)	Nonblack (n = 366)	Black (n = 260)
	Whether ever in nonintact family					
Yes	0.32	0.25	0.71	0.36	0.30	0.71
	Whether ever in various family types					
Two-parent	0.90	0.95	0.62	0.89	0.93	0.65
Mother-only	0.25	0.19	0.58	0.30	0.25	0.57
Mother with stepfather	0.10	0.09	0.17	0.13	0.13	0.12
Mother with grandparents	0.07	0.04	0.21	0.05	0.03	0.16
Other living arrangements	0.13	0.11	0.27	0.15	0.12	0.32
	Whether ever in various family types in each developmental stage					
2-parent						
age 0–5	0.89	0.94	0.58	0.89	0.93	0.64
age 6–10	0.82	0.88	0.45	0.80	0.84	0.57
age 11–15	0.78	0.84	0.41	0.74	0.79	0.42
Mother-only						
age 0–5	0.15	0.11	0.37	0.15	0.13	0.30
age 6–10	0.17	0.11	0.49	0.21	0.16	0.46
age 11–15	0.17	0.12	0.52	0.21	0.15	0.54
Mother w/ stepfa						
age 0–10	0.07	0.06	0.10	0.08	0.08	0.09
age 11–15	0.09	0.08	0.11	0.12	0.12	0.09
Mother w/ grandp						
age 0–5	0.05	0.03	0.20	0.04	0.03	0.14
age 6–10	0.02	0.01	0.10	0.01	0.00	0.06
age 11–15	0.03	0.02	0.09	0.02	0.01	0.07
Other						
age 0–5	0.05	0.04	0.16	0.08	0.06	0.19
age 6–10	0.05	0.04	0.10	0.07	0.05	0.14
age 11–15	0.09	0.07	0.15	0.08	0.07	0.12
	Event-based family structure					
Born into 2-parent	0.86	0.91	0.51	0.86	0.91	0.54
Born into mother-only	0.11	0.07	0.38	0.10	0.06	0.33
Born into other non-two-parent	0.03	0.02	0.11	0.04	0.03	0.13
Parental div/sep						
age 0–5	0.09	0.07	0.18	0.09	0.10	0.07
age 6–10	0.07	0.07	0.11	0.10	0.10	0.10
age 11–15	0.07	0.07	0.08	0.09	0.08	0.12
Parental (re)mar						
age 0–5	0.05	0.05	0.06	0.06	0.07	0.04
age 6–10	0.06	0.06	0.06	0.06	0.06	0.07
age 11–15	0.06	0.07	0.05	0.08	0.08	0.08
	Sequence-based family structure					
Mother-only all 15 years	0.03	0.01	0.13	0.03	0.01	0.14
Mother-only to 2-parent & remained 2-parent	0.04	0.04	0.07	0.03	0.03	0.03
Mother-only to 2-parent and back to mother-only	0.02	0.01	0.13	0.03	0.02	0.11
2-parent all 15 years	0.71	0.78	0.31	0.66	0.73	0.31
2-parent to mother-only & remained mother-only	0.07	0.05	0.19	0.10	0.08	0.19
2-parent to mother-only and back to 2-parent	0.07	0.08	0.01	0.09	0.10	0.04
Other sequences	0.06	0.04	0.15	0.06	0.04	0.18

mother-only family compared with Hernandez' estimates of 11.8% (1970), 16.2% (1980) and 21.0% (1988). Here the difference in estimates may in part reflect our categorization of some complex living arrangements with mother but not father present in the 'other arrangements' category, whereas Hernandez may have counted them as 'mother-only.'

Our PSID estimates for stepchildren run low relative to what we could find elsewhere, but the comparison is not entirely direct and confounding elements may distort the comparisons. U.S. Census data from Don Hernandez (correspondence in October 1999) indicate the following percentages of children in a home with mother and stepfather: 6.5% (1970), 8.4% (1980) and 10.5% (1990). This compares with our simulated estimate of 3.7% of children living as stepchildren. The two types of estimates are not identical; we would expect our figure to be lower since there can be a mixture of biological and stepchildren in the same home. It is exceedingly difficult to gauge how much lower the actual figure might be. If we apply Hernandez' (1993) estimates of percentage of children living in two-parent families to the family structure distributions for children living in two-parent families given in Moorman and Hernandez (1989), we obtain an estimate of the percentage of children who are themselves stepchildren in mother-plus-stepfather homes as follows: 7.0% (1981) from NHIS data and 8.6% (1980) from CPS data, compared with our 3.7%.

Bearing in mind these data quality issues, we turn our attention to what the PSID data tell us about the living arrangements of children. The panel data are treated like pooled cross-sections to provide Table 1's view on family structure by single year of age over childhood. Although cross-sectional views can be deceptive about the underlying dynamics, this table does reveal some interesting age patterns. The table is disaggregated by race to show the striking differences for black vs. non-black children in chances of living in non-intact families at every single age during childhood. The age patterns, however, tend to hold for both subgroups.

As children age, the general tendency is for the proportion living in two-parent (non-stepparent) families to fall substantially and the proportion living in mother-only families or in mother-and-stepfather families to rise. Children are most likely to be living in mother-and-grandparent(s) families when they are young: this form of family structure is most common when children are between the ages of one and four. Relative to nonblacks, black children were especially likely to spend some time in a mother-with-grandparent arrangement.

The sizable variation by age in the types of family structure children experience calls into question reliance on static assessments of childhood family structure. A considerable amount of past research has used age 14 as the anchoring point for a childhood family structure measure, but a striking finding in Table 1 is that age 14 is very unrepresentative of family structure experiences in childhood.

The second panel of Table 2 shows in a more summary form the variety in the types of non-intact families children experience. Most children from non-intact families spent some time in a mother-only family, but many spent at least part of their childhood in more complex arrangements. About 10% experienced life in a mother-with-stepfather family, and roughly half that number were in a mother-with-grandparent(s) family at some time.

A notable fraction of children spent at least part of their childhood in what we categorize as "other living arrangements." This group consists of a diverse set of nontraditional family structures containing fathers only or relatives other than parents or grandparents. Each of these component structures tends to have an associated sample size too small for separate analysis. Black children were especially likely to experience these nontraditional structures, with about 30% in such a family at some point during childhood.

The event-based measures of family structure show that at each childhood stage roughly 10% of the children experienced a parental divorce or separation and about 5% experienced parental marriage or remarriage. The sequence-based measures of family structure show that only 3% of the children spent their entire childhood in a mother-only family, whereas 6% were born to a mother-only family and later switched either temporarily (especially likely for black children) or permanently to a two-parent family. Of the children in a mother-only family at some point, blacks were much more likely than non-blacks to have that type of structure as their only childhood family structure. About 15% were born to a two-parent family and later experienced life in a mother-only family either in transition to a new two-parent family or for the remainder of their childhood, and those situations were equally likely.

4. How well do the theories fit the data?

Our investigation of the relevance of existing theories focuses on two outcomes for children as young adults – educational attainment and, for females, risk of a first premarital birth. We analyze sons and daughters separately since the child development and sociological literatures frequently find differing influences of family structure by sex of the child. We do not estimate separate models for blacks and non-blacks, owing to the small sample sizes when the sample is divided into race-sex subgroups. Our investigation of possible race interactions yielded few notable instances of them.

Our models assume individual decision-making on the part of the young adult is influenced by exogenous childhood family history. Implicit in this model is the strong assumption that there are no unobserved processes jointly affecting both family background and children's attainments. Yet one hypothesis noted in the economic literature on intergenerational influences is that parents respond to children's mental and physical endowments in their financial support of children's education. To the extent that there is intergenerational correlation in these endowments and children's endowments are unobserved in the data, family income and parents' education may not be exogenous to children's educational attainment (Ermisch and Francesconi 2000, 2001). Sibling models, for example, could identify the influences of family background under weaker assumptions; however, modest sample sizes for sibling pairs in unusual family structures at different childhood stages restrict our ability to estimate sibling models for the wide variation in family structure central to our research. Because of such limitations, we retain our approach of modeling family background as exogenous but are cautious in drawing conclusions about what the results reveal about causal linkages.

4.1. Outcome measures

The most recent report of completed schooling, typically that given in the 1995 interview, forms the basis of the years of schooling measure. The age at

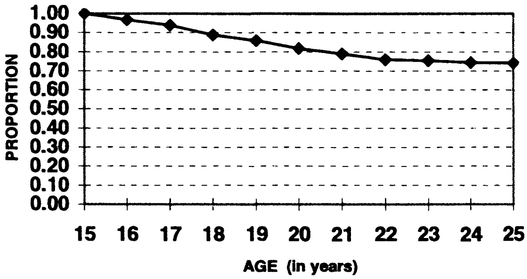

Fig 1. Proportion of daughters with no nonmarital birth by age

which the schooling was reported was between 20 and 25, depending on the point at which data for the individual became censored. To minimize potential biases from any systematic components in the differential censoring, our regression analyses of educational attainment include a control variable reflecting age at censoring. Educational attainment measured in this way was, on average, about one half year beyond high school, with little difference in mean level for young men versus young women (12.7 years for sons and 12.8 years for daughters, with standard deviations of 1.6 for each).

Risk of a first nonmarital birth is examined only for daughters because of concerns about reliability for a comparable measure for sons.[6] This focuses on the age at which a first nonmarital birth occurred, if ever, during the observation period. This event-history analysis of nonmarital fertility begins at age 16 and tracks on a monthly basis birth and marital events up through the 1992 interview. Since the majority of sample daughters will never have a nonmarital birth, there is a great deal of right-censorship in these data. In addition, a case could be censored either by attrition from the study or first marriage. The maximum age at the time of right-censorship is age 25.

Figure 1 traces, by age, the fraction of daughters having no nonmarital birth. It shows that nonmarital first births were most likely to occur to young women between the ages of 18 and 22: about 10% had a nonmarital first birth by age 18, compared to about 25% by age 22; relatively few had a nonmarital first birth after age 22.

4.2. Models estimated

In preliminary work we estimated several distinct models of family structure influences. To conserve on space we focus on just two, one that omits family income and one that includes family income. Both control for time-invariant characteristics of the individual and family. The first model facilitates comparisons of findings with prior research having no access to an income measure. The second model helps determine the relevance of the economic hardship theory and provides a base for assessing the relevance of other theories independent of the influence of economic hardship.

All models include time-invariant controls for: (a) race of the head of the family in the year of the child's birth (black = 1, non-black = 0); (b) mother's education (years of completed schooling); (c) age of the mother at the time of the child's birth, with missing data set to zero; (d) a dichotomous indicator of whether there is non-missing data on the mother's age at the time of the child's birth (yes = 1, no = 0); (e) total number of the child's siblings; (f) residence in the South (ever yes = 1, never = 0); (g) the average unemployment rate in the child's county of residence between ages 11 and 15; and (h) the average AFDC benefit in the child's state of residence at age 14, expressed in thousands of 1993 dollars. Means and standard deviations for these variables are provided in the Appendix Table 1.

Two distinct measures of childhood family income are used in our estimation process. When a measure of the family structure itself is the focus, family income is measured as average annual income (inflation-adjusted to units of 10,000 1993 dollars).[7] When parental marital events are the focus, family income is measured in terms of change, specifying whether or not there was a drop in family income of 50% or more during the specified period.[8] Family income is adjusted for family size with the inclusion of controls for family structure (which assesses presence of adults) and number of siblings.

4.3. *How results compare to findings in the literature*

Comparisons with prior research findings are complicated by differences in specification. These differences include: sample construction (other researchers have often combined sons and daughters, disaggregated by race, or both); disaggregation of family structure by types (e.g., other researchers have classified mother-with-grandparent(s) as part of mother-only); specification of the outcome (e.g., other researchers sometimes use high school completion rather than years of education as their education outcome); and choice of control variables (e.g., other researchers have tended to omit factors such as childhood family income, area-based measures, or age of mother at the birth of the child).

For tightest comparability to prior research, we look to findings with family structure designated in terms of the dichotomy of 'whether ever in a non-intact family' during childhood and with childhood family income omitted from the regressions (see first panel of estimates in Table 3). Coefficients are in the directions expected from comparable prior research, showing time spent in a non-intact family to be associated with lower levels of educational attainment and a higher risk of non-marital birth.[9]

Disaggregating non-intact family by type (second panel in Table 3) reveals similarities as well with earlier findings in terms of the direction of associations for more specific types of non-intact families. In our findings for both mother-only and mother-with-grandparent(s) structures we see negative associations with educational attainment and positive associations with risk of a non-marital birth. Wojtkiewicz's (1993) and Wu and Martinson's (1993) findings using an aggregated form of mother-only structure that include the mother-with-grandparent(s) type yielded coefficients in the same directions.[10] Results for mother-with-stepfather structures also bear similarities to prior research: like Wojtkiewicz (1993) we find no evidence of a relationship to educational attainment, and like Wu and Martinson's (1993) results for blacks, though not for whites, we find no evidence of a relationship to risk of a non-intact birth.[11]

Regarding structures rarely viewed in the literature, we find weak evidence for the importance of mother-and-grandparent(s) structure, at least with regard to children's education. This type of non-intact structure displays large

Table 3. Relationship of whole-childhood measures of family structure to children's completed schooling and risk of having a first nonmarital birth, by sex

	Years of completed schooling		Nonmarital birth	
	Sons coeff. (std.err.)	Daughters coeff (std.err.)	Daughters coeff (std.err)	Risk
Whether ever in non-intact family	-0.23^{+}	-0.27^{**}	0.53^{***}	1.70
	(0.13)	(0.12)	(0.11)	
Adjusted R-squared/log likelihood	0.175	0.275	−646.4	
Whether ever in various family types				
Mother-only	−0.25	0.01	0.45	1.57
	(0.16)	(0.15)	(0.24)	
Mother with grandparent(s)	−0.34	-0.49^{+}	0.17^{+}	1.19
	(0.23)	(0.26)	(0.33)	
Mother with stepfather	0.15	−0.03	0.02	1.02
	(0.21)	(0.19)	(0.28)	
Other living arrangements	−0.28	-0.36^{*}	0.81^{***}	2.25
	(0.18)	(0.17)	(0.24)	
Adjusted R-squared/log likelihood	0.181	0.279	−647.2	
Sequence-based family structure				
2-parent all 15 years (omitted)				
Mother-only all 15 years	-0.72^{*}	−0.01	−0.60	0.55
	(0.36)	(0.36)	(0.45)	
Mother-only to 2-parent	−0.08	−0.28	0.30	1.35
	(0.28)	(0.33)	(0.62)	
Mother-only to 2-parent to Mother-only	-1.36^{**}	−0.47	1.48^{***}	4.39
	(0.37)	(0.30)	(0.38)	
2-Parent to Mother-only	0.04	−0.28	0.98^{**}	2.66
	(0.22)	(0.19)	(0.31)	
2-Parent to Mother-only to 2-parent	0.03	−0.02	0.98^{**}	2.66
	(0.23)	(0.19)	(0.32)	
Other sequences	-0.78^{**}	−0.32	0.98^{**}	2.66
	(0.24)	(0.24)	(0.32)	
Adjusted R-squared/log likelihood	0.194	0.270	−646.7	

$^{+}$, *, **, *** Denotes estimate is statistically significant at 0.10, 0.05, 0.01 and 0.001 level respectively.
Other predictors include age at censoring, ethnicity, mother's education, age of mother at birth of child, nonmissing data on age of mother at birth, total number of siblings, ever lived in south, average county unemployment rate at age 11–15, and average state AFDC benefit at age 14.

negative, though insignificant, coefficients in the education equations of both sons and daughters.

Switching to sequences of family structure (bottom panel in Table 3) we see evidence of the mother-only structure throughout childhood being less important than some changes in family structure. The stable mother-only structure registers a statistically significant difference from the stable two-parent structure only in the sons' educational attainment regression. A transition structure – mother-only to two-parent back to mother-only – stands out with a stronger link to children's outcomes. Sons experiencing this structure have considerably lower educational attainment than sons from two-parent or mother-only structures. Daughters from this transition structure have a higher risk of a non-marital birth than daughters from stable two-parent or mother-only structures. Indeed, daughters from most any transition structure run a higher risk of a non-marital birth than daughters from the stable structures. Yet family structure sequences matter little for daughters' educational attainment, with transition structures showing negative but insignificant coefficients and the stable mother-only structure having an essentially zero coefficient. The tenor of these findings is roughly consistent with that of Wojtkiewicz (1993) and Wu and Martinson (1993). Given the differentials we find for sons' and daughters' educational attainment, collectively our results are in line with Wojtkiewicz's major conclusions regarding children's odds of finishing high school: a negative association with change in family structure and no association with prolonged stay in a mother-only family. Our results showing the heightened linkage of daughters' risk of a non-marital birth to transition structures fit with Wu and Martinson's finding that, for both white and black daughters, risk of a non-marital birth increases with number of changes in family structure.

Overall, our results tend to show similar but weaker associations than those of earlier researchers, with many of our coefficients failing to achieve significance at conventional levels. It is possible that these differences in findings are attributable to cohort differences since the other researchers investigated earlier cohorts of children, who were less likely to experience a mother-only structure and hence may have experienced greater differences in family life. The influence of non-intact family structures may well have diminished as they have become more common, but tests of such an assertion await further research.

4.4. Timing during childhood

One of the unique contributions of this paper is being able to observe variation in the influence of family structure over the course of childhood. To assess the role of timing of family structure influences, we turn to stage-specific measures constructed separately for early childhood (ages 0–5), middle childhood (ages 6–10) and late childhood (ages 11–15). Tables 4 and 5 present regression results regarding these stage-specific measures. Table 4 focuses on the type of family structure experienced at each stage of childhood. Table 5 focuses on change in family structure, with the change assessed in terms of parental marital events at different stages of childhood (controlling for the family structure children are born into).

These tables, each with precise timing of family structure experiences, emphasize the importance of: (1) timing, (2) new types of family structure, and (3) change as opposed to extended stay in non-intact families. Predictive power improves when we switch from regressions using 'whether ever in a non-intact family' as the sole indicator of childhood family structure to specifications differentiated by childhood stage and type of non-intact family (not shown).[12] The enhanced predictive power appears to be concentrated in distinguishing two family structures, mother-with-stepfather and mother-with-grandparent(s), at particular childhood stages (see Table 4).

There is suggestive evidence that the mother-with-stepfather arrangement has some importance for daughters' subsequent educational attainment with the direction of the association differing by childhood stage. Daughters in

Table 4. Relationship of family structure at different childhood stages to children's completed schooling and risk of having a premarital birth, by sex

	Years of completed schooling		Nonmarital birth	
	Sons coeff (std.err.)	Daughters coeff. (std.err.)	Daughters coeff (std.err.)	Risk
Age 0–5				
Family structure:				
Mother-only	−0.20	0.32+	−0.35	0.70
	(0.21)	(0.19)	(0.28)	
Mother with grandparents	−0.02	−0.02	−0.32	0.73
	(0.32)	(0.33)	(0.46)	
Mother w/ stepfather (age 0–10)	−0.21	−0.52+	−0.02	0.98
	(0.33)	(0.29)	(0.52)	
Other nonintact family	−0.17	−0.45+	0.34	1.40
	(0.26)	(0.23)	(0.33)	
Average family income[a]	0.17**	0.12**	−0.04	0.96
	(0.06)	(0.04)	(0.31)	
Age 6–10				
Family structure:				
Mother-only	−0.17	0.19	0.36	1.43
	(0.26)	(0.20)	(0.30)	
Mother with grandparents	−0.23	0.36	0.54	1.72
	(0.52)	(0.59)	(0.65)	
Mother w/ stepfather (see above)	–	–	–	
	–	–	–	
Other nonintact family	0.29	−0.06	0.48	1.61
	(0.30)	(0.25)	(0.37)	
Average family income[a]	−0.02	0.02	−0.01	0.99
	(0.04)	(0.04)	(0.12)	
Age 11–15				
Family structure:				
Mother-only	0.02	−0.34+	0.18	1.20
	(0.23)	(0.19)	(0.33)	
Mother with grandparents	−0.84*	−1.07*	0.84	2.32
	(0.40)	(0.44)	(0.55)	
Mother w/ stepfather (age 11–15)	0.34	0.19	0.40	1.49
	(0.30)	(0.24)	(0.46)	
Other nonintact family	−0.21	0.08	0.04	1.04
	(0.24)	(0.23)	(0.36)	
Average family income[a]	−0.02	−0.00	−0.25**	0.78
	(0.04)	(0.02)	(0.08)	
Adjusted R-squared/log likelihood	0.211	0.312	−630.5	

[a] Family income is measured in $10,000 dollars and inflated to the 1993 price levels using Consumer Price Index CPI-UX1.

+, *, **, *** denotes estimate is statistically significant at 0.10, 0.05, 0.01 and 0.001 level respectively.

Other predictors include age at censoring, ethnicity, mother's education, age of mother at birth of child, nonmissing data on age of mother at birth, total number of siblings, ever lived in south, average county unemployment rate at age 11–15, and average state AFDC benefit at age 14.

Table 5. Relationship of parental marital events at different childhood stages to children's completed schooling and risk of having a premarital birth, by sex

	Years of completed schooling		Nonmarital birth	
	Sons coeff (std.err.)	Daughters coeff. (std.err.)	Daughters coeff (std.err.)	Risk
Whether born into mother-only family	−0.73*** (0.21)	−0.21 (0.20)	−0.19 (0.31)	0.83
Whether born into other nonintact family	−0.83* (0.33)	−0.14 (0.28)	0.90* (0.34)	2.46
Age 0–5				
Marital events:				
Parental divorce/separation	−0.20 (0.23)	−0.21 (0.23)	1.08** (0.35)	2.94
Parental (re)marriage	0.53+ (0.28)	−0.13 (0.26)	−0.65 (0.48)	0.52
50%+ Family income loss	0.50** (0.08)	−0.11 (0.18)	−0.32 (0.26)	0.73
Age 6–10				
Marital events:				
Parental divorce/separation	−0.01 (0.24)	−0.33 (0.24)	0.87** (0.33)	2.39
Parental (re)marriage	0.09 (0.21)	0.43 (0.29)	−1.05* (0.48)	0.35
50%+ Family income loss	−0.16 (0.18)	−0.07 (0.15)	0.10 (0.25)	1.11
Age 11–15				
Marital events:				
Parental divorce/separation	0.50* (0.23)	−0.49* (0.22)	0.96** (0.36)	2.61
Parental (re)marriage	−0.84** (0.26)	0.69** (0.25)	−0.47 (0.39)	0.63
50%+ Family income loss	−0.37* (0.16)	−0.21 (0.15)	−0.01 (0.04)	0.99
Adjusted R-squared/log likelihood	0.208	0.278	−640.8	

+, *, **, *** denotes estimate is statistically significant at 0.10, 0.05, 0.01 and 0.001 level respectively.

Other predictors include age at censoring, ethnicity, mother's education, age of mother at birth of child, nonmissing data on age of mother at birth, total number of siblings, ever lived in south, average county unemployment rate at age 11–15, and average state AFDC benefit at age 14.

mother-with-stepfather families in early to mid childhood appear to complete less education than their counterparts in intact or mother-only families at that time; however, the evidence is not strong (significant at the 0.1 level). Daughters in mother-with-stepfather families in adolescence appear to complete more education than their counterparts in mother-only families; again the evidence is not strong (0.1 level). Both mother-with-stepfather and mother-only structures seem to be behaving differently at the two ends of daughters' childhood; standard errors are, however, quite large, suggesting caution about this conclusion. There is no evidence of the mother-with-stepfather structure being of importance to sons' subsequent educational attainment.

It is in late childhood that the mother-with-grandparent(s) structure shows its strongest associations with subsequent educational attainment, and here the associations are similar for sons and daughters. Young adults, sons as well as daughters, who lived in a mother-with-grandparent(s) structure during their adolescent years tended to have about one year less completed schooling than otherwise similar young adults who lived in two-parent families at that stage of childhood.

From the size of the coefficients alone, it would appear that the mother-with-grandparent(s) structure in late childhood has the strongest negative association with the outcomes we examine of any non-intact structure at any childhood stage. However, this conclusion is tempered by results of statistical tests for differences in coefficients both across the late-childhood structures and for the mother-with-grandparent(s) structure across the different childhood stages. Whether the mother-with-grandparent(s) structure is, of all the non-intact structures in late childhood, the most negative in its association with subsequent educational attainment is unclear. The coefficient on the mother-with-grandparent(s) variable is (at 0.05 significance level) more negative than that of all other non-intact structures in late childhood except, for daughters, the mother-only structure and, for sons, the 'other' set of structures. More lenient significance thresholds (0.15 significance level) support the difference between the mother-with-grandparent(s) structure and the mother-only structure in late childhood. Tests for differences in coefficients for the mother-with-grandparent(s) structure at different childhood stages also yield results highly sensitive to the significance threshold. The coefficient on the mother-with-grandparent(s) variable in late childhood is not perceived as larger than the coefficients on that same structure at earlier childhood stages when the significance threshold is set at the 0.05 levels. A more lenient significance level of 0.10 yields significant differences for daughters, though still not sons. On the whole, the evidence suggests the cautious observation that there is a good chance that the mother-with-grandparent(s) structure in late childhood is especially detrimental to children's education. The mother-with-grandparent(s) structure clearly merits further research as something distinct from the mother-only structure.

The evidence provides little support for social control theory's contention that additional adults in non-intact families lessen detrimental influences of the structure. Children spending time in mother-with-stepfather families or mother-with-grandparent(s) families do not tend to have higher levels of subsequent educational attainment or lower risks of a nonmarital birth than those spending time in mother-only families. The one exception is daughters in mother-with-stepfather families late in childhood.

Changes in parents' marital arrangements, especially those occurring in late childhood, show several strong associations with children's adult behaviors (see Table 5). The pattern, though, varies with the type of parental event, its timing in terms of childhood stage, and the sex of the child. One pattern that holds for both sons and daughters is a tendency for parental marital disruption and parental (re)marriage to show opposite associations with the children's young adulthood outcomes. Each of these changes, however, seems to influence sons and daughters differently.

Relative to otherwise similar counterparts, daughters who experienced parental marital disruption at any childhood stage tended to have a higher risk of a nonmarital birth (two and one-half to three times as high). Daughters who experienced parental marital disruption in late childhood also tended to have less subsequent completed schooling (about one half year less). Sons, on the other hand, who experienced parental marital disruption in adolescence tended to have more, rather than less, subsequent educational attainment (about one-half year more) than their otherwise similar counterparts.

Parental (re)marriage experienced in middle childhood is associated with a lower risk for daughters of a nonmarital birth (one-third as high). Parental (re)marriage experienced in adolescence is associated, for daughters, with higher (about two-thirds of a year more) educational attainment. For sons, on the other hand, parental (re)marriage at that late stage of childhood is associated with less (about four-fifth of a year less) educational attainment.

The strength of the relationship between parental (re)marriage and subsequent educational attainment differs depending on the timing of this event. Much of this difference is in terms of whether there is or is not a notable association. The difference is more striking, though, in the case of early versus late childhood parental re(marriage) experienced by sons. There is weak evidence of parental (re)marriage early in childhood being associated with more (by about one-half year), rather than less (compared with, in late childhood, one year less), subsequent educational attainment. This suggests that the timing of parental marital events is especially important for sons' educational attainment.

5. Discussion

With data on full childhood family experiences, we estimate models of educational attainment and nonmarital births that provide a variety of views of the role of family structure in shaping children's lives. To distinguish the ways in which family structure operates, we introduce into our models various sets of control variables. We cannot entirely rule out possibilities of endogeneity for all control variables; hence our findings are tempered with some caution.

In general, our findings are more consistent with the underlying family mechanisms posited by stress theory (stress caused by disruptions in family structure) than those posited by social control theory (tighter supervision through larger numbers of parents or substitute parents). Change rather than type of structure predominates in the strength of association with children's adult behaviors, and there is no evidence that the number of adults in the child's home, per se, reduces detrimental influences of exposure to a nonintact family. We find that parental marital change has a more pronounced association with the outcomes we examine if the events occurred in late childhood. This is consistent with stress theory also, which posits that change in family structure will have its greatest effects close to the time of the change. Because educational attainment and nonmarital births are observed in the years most immediately following late childhood, it is that stage of childhood that is closest in timing.

Yet some of the findings are also at odds with stress theory. Most especially, change is not always tied to the outcomes in a detrimental way. Parental marital change is not always associated with lower educational attainment or a higher risk of nonmarital birth; it sometimes has an association in the opposite direction, and many times has no association that is statistically different from zero at conventional levels. In addition, the young adult outcomes we examined appear to be unrelated to several sequences of transitions in family structure.

Our analysis also allows us to investigate the economic hardship theory, matching the measure of childhood family income to that of family structure. From Tables 4 and 5 we see that childhood family income is linked to children's educational attainment and daughters' risk of a nonmarital birth, though not at all childhood stages. Average childhood family income relates to educational attainment and risk of a nonmarital birth in the expected direction: when average childhood income is included in combination with stage-specific family structure measures (not shown), income is positively related to children's education and negatively related to daughters' risk of a nonmarital birth.[13] Predictive power improves, however, when the match in timing of income and family structure is tighter.[14] Our disaggregation of family income allows us to see what appears to be differential concentration of family income linkages by childhood stage. It is the income during early childhood that is most strongly linked to sons' and daughters' subsequent educational attainment. On the other hand, it is the income during late childhood that most strongly links level of income to daughters' risk of a nonmarital birth. Late childhood also appears to be the time of strongest linkage between income loss and subsequent lower levels of educational attainment, as the results in Table 5 indicate. Oddly enough, for sons family income loss in early childhood is associated with higher, rather than, lower, completed schooling. Especially since no similar findings appear for daughters, it is difficult to know what to make of this. Taken as a whole, childhood family income appears to play an important role in the outcomes examined, though the role is not entirely straightforward and shows strong evidence of notable differences by childhood stage.

Childhood family income does not appear to account for linkages between childhood family structure and the outcomes we examine. Those linkages appear to be largely independent of family income (a finding consistent with Wu 1996). When we estimate the models presented in Tables 4 and 5 omitting income as a control, the coefficients on the family-structure and parental-marital event variables are hardly affected.

To assess other possible mechanisms, we also investigated three other ways in which childhood family structure might influence educational attainment and nonmarital birth: via intergenerational transmission of reliance on public support, reduced attention from the mother because of her greater work commitment, and disruptive influences of residential relocation.[15] Of the three, residential relocation exhibits the strongest relationship to the outcomes we examine. Such moves, for daughters, are associated with reduced educational attainment if the move occurred in middle or late childhood (see Appendix Table 1). In analysis not shown here, residential moves were also found to account for part of the negative association between daughters' educational attainment and parental marital events or living in a mother-with-stepfather family during adolescence. Mother's market work commitment as well as receipt of AFDC showed no notable relationships to the outcomes we examined.

Some of the findings simply do not fit the major existing theories and call out for further theoretical development. The mother-with-grandparent structure is not well understood; our findings suggest that it merits more attention in both empirical research and theories of family structure influences. Though our modest sample sizes preclude strong conclusions, our results suggest this type of structure experienced late in childhood could be strongly tied to reduced educational attainment and enhanced risks of a nonmarital birth. Because it is a type of structure rather than a change in structure, this finding does not fit with stress theory. Because it involves more adults than a mother-only structure, it does not fit with social control theory either. Because the strength of the relationship holds even with controls for family income, it does not fit with economic hardship theory.

Why would the mother-with-grandparent(s) structure during late childhood be so negatively related to young adult outcomes, and why would that same structure early in childhood have little association with children's long-run outcomes? One possibility is that conflicts in rearing practices may be especially important in adolescence. The two generations of adults – parent and grandparent – may have conflicting ideas about how to raise teenagers, and inconsistency in their supervisory activities undermines their ability to steer adolescents away from destructive behaviors. Another possibility is that the observed associations reflect something about the reason the structure was formed. In late childhood such a structure probably means one of two things – that the child's mother has recently had a marital disruption and is getting help from the child's grandparents or that the child's grandmother is in need of financial assistance or help with activities and has moved into the child's family to receive such help. Both possibilities can mean, relative to intact families, that less of the total family income is allocated to children, that children have less privacy, and that children are expected to devote more time to responsibilities at home. This points to differences in within-family resource allocation as a possible underlying mechanism, with family structure as a determinant of within-family resource allocation patterns (Chase-Lansdale et al. 1994).

Findings for some types of parental marital events at certain stages of childhood also merit further thought to suggest theories about the underlying mechanisms. Parental (re)marriage in late childhood is positively related to daughters' educational attainment, and in middle childhood this event is negatively related to their risk of a nonmarital birth. In addition, parental divorce or separation in late childhood is positively related to son's educational attainment. Stress theory predicts the opposite because parental marital change means change in organization. The findings for parental (re)marriage might fit with social control theory (because the change adds a substitute parent). However, why is the association of this event with subsequent educational attainment so different (and strongly negative) for sons? A possible explanation for the pattern of results regarding parental (re)marriage may lie in differences between fathers and stepfathers in their commitment to adolescents. Men who marry into families with children in or near adolescence may tend to be more supportive because they self-select themselves into that position and that responsibility. Their support may benefit girls more than boys because girls may tend to accept their support whereas boys may tend to resist it, in part because boys may resent the stepfather stepping into the father role at a time when they feel they are old enough to be 'man of the family.' It is possible that something of a mirror image accounts for the positive association of parental divorce or separation late in childhood with sons' subsequent educational attainment: boys may take adult roles more seriously when they feel they are the 'man of the family.'

Overall, the results indicate that theories about effects of childhood family structure need to recognize that: (a) change in family structure is important, (b) the timing of family structure experiences over the course of childhood is relevant, (c) other adults in children's homes, besides just the parents, can be important, and (d) the influence of childhood family structure can vary by type of outcome and sex of the child. New theories should consider the allocation, as well as the level, of economic resources within children's homes and the reasons why people form different kinds of family structures. The development and testing of theories along these lines could call upon a variety of analytical approaches, including the use of sibling models and other techniques to control for unmeasured heterogeneity (see, for example, Sandefur and Wells 1997; Teachman et al. 1995; Griliches 1979) and consideration of techniques that account more fully for the detailed sequencing patterns of family structure over a given childhood life stage (see, for example, Rohwer and Trappe 1997, and Rohwer 1996). We encourage development and testing of new theories in these directions.

Appendix

Table A1. Means and regression coefficient estimates for predictors other than family structure or income

	Means (std. dev.)		Completed schooling		Nonmarital childbearing
	Sons	Daughter	Sons	Daughters	Daughters
Age at censoring	21.95	21.78	0.25***	0.36***	–
	(1.52)	(1.50)	(0.04)	(0.04)	
Ethnicity (black = 1)	0.14	0.15	0.67***	0.30+	0.60**
	(0.36)	(0.37)	(0.19)	(0.18)	(0.27)
Mother's education	12.17	12.6	0.13***	0.14***	−0.04
	(2.3)	(2.2)	(0.03)	(0.02)	(0.04)
Age of mother at birth of child	25.0	24.0	0.00	0.01	−0.07**
	(7.7)	(8.0)	(0.01)	(0.01)	(0.02)
Non-missing data on age of mother at birth	0.97	0.95	0.16	0.44	0.96
	(0.18)	(0.22)	(0.45)	(0.39)	(0.74)
Number of siblings	2.3	2.3	−0.08*	−0.13***	0.30***
	(1.9)	(1.9)	(0.03)	(0.04)	(0.06)
Ever lived in south	0.39	0.36	−0.03	−0.20	−0.06**
	(0.50)	(0.49)	(0.17)	(0.16)	(0.02)
County unemployment rate, age 11–15	7.5	7.8	−0.07**	−0.03	−0.05
	(2.3)	(2.6)	(0.02)	(0.02)	(0.04)
State AFDC benefit, age 14 (in $1,000)	0.37	0.38	0.10	0.00	0.31**
	(0.15)	(0.15)	(0.56)	(0.52)	(0.12)
Ever on AFDC					
age 0–5	0.14	0.16	−0.33	−0.21	0.06
	(0.36)	(0.37)	(0.25)	(0.20)	(0.28)
age 6–10	0.16	0.19	−0.01	−0.34	−0.01
	(0.37)	(0.40)	(0.24)	(0.21)	(0.32)
age 11–15	0.14	0.16	−0.40+	−0.08	0.11
	(0.36)	(0.37)	(0.23)	(0.21)	(0.29)

Table A1. (continued)

	Means (std. dev.)		Completed schooling		Nonmarital childbearing
	Sons	Daughter	Sons	Daughters	Daughters
Mother ever worked 1,000+ hours annually					
age 0–5	0.41	0.43	0.01	−0.01	−0.25
	(0.50)	(0.50)	(0.13)	(0.12)	(0.23)
age 6–10	0.50	0.52	−0.22+	−0.04	0.17
	(0.51)	(0.51)	(0.13)	(0.13)	(0.28)
age 11–15	0.68	0.68	0.20	0.11	0.19
	(0.48)	(0.47)	(0.14)	(0.14)	(0.26)
Ever moved					
age 0–5	0.66	0.67	0.01	0.05	−0.07
	(0.48)	(0.48)	(0.13)	(0.12)	(0.23)
age 6–10	0.52	0.49	−0.21+	−0.44***	−0.13
	(0.51)	(0.51)	(0.12)	(0.11)	(0.21)
age 11–15	0.38	0.39	0.18	−0.29*	0.06
	(0.50)	(0.50)	(0.13)	(0.12)	(0.22)

+, *, **, *** denotes estimate is statistically significant at 0.10, 0.05, 0.01 and 0.001 level respectively.
Other predictors include 'whether ever in various family types in each developmental stage' and stage-specific family income.

Endnotes

[1] By "present" we mean either in an interviewed household or associated with an interviewed household but living in an institution. For PSID aficionados, this translates into a requirement that Sequence Numbers are always in the 1–59 range.

[2] Since the first two years of the study, nonresponse losses to the Panel have been small, and checks against other data indicate no appreciable sample biases. Weights adjust for the original oversampling of the poor and for differential nonresponse and losses to the Panel. However, attrition is likely to be tied to parental marital dissolution, and those who attrite may well be more adversely affected by the event than those who remain in the study. In addition, during the time span of PSID data used in this analysis the PSID systematically dropped from the study children living with a non-sample parent after a parental marital breakup. A PSID tracking rule specified that only sample adults were to be followed. This meant that after the breakup of parents, children were followed only if they remained with a sample parent. To the extent that the PSID weights properly correct for differential nonresponse, problems of bias created by this procedure should be minimal. It is still possible, though, that differential attrition accompanied family structure change and was of a sort that biases estimates of the relationship between childhood family structure and young adult outcomes. To allow for this possibility it would be useful to have weights adjusted for family structure changes or to model attrition as part of the analysis procedure.

[3] Two-parent structures with grandparent(s) present are included in this category. We make no distinction regarding presence or absence of grandparent(s) in the two-parent structure in our research.

[4] A person is classified as a "stepparent" if he/she reports in the retrospective substitute-parenting history collected in 1985 that he/she raised the child for at least one year and was not the biological or adoptive parent of the child, and if we were able to confirm with annual interview information that the child co-resided with that person in a parent-child relationship.

However, limiting identification of stepparent situations to this definition alone misses some stepparent situations. This is because fertility and marital histories were not gathered until 1985 and not all adults who had co-resided with children at some point prior to 1985 survived in the panel until 1985. In particular, if a man was living in the family of the child as a husband or permanent partner of the mother prior to 1985 but was lost to attrition before 1985, we could not establish a definitive relationship between that man and the child. In such cases, we assumed that the man was the biological or adoptive father if the mother reported being married to the man and the man was co-residing with the child at the time of the child's birth; otherwise, we assumed he was a stepfather.

[5] The PSID estimates average across several calendar years and apply to single year of age for ages 1–15. To simulate estimates for a cross-section of children of all ages, we weighted the figures given in Table 1 according to the age distribution of children in 1980 (approximately the midpoint for the PSID data). The age distribution figures were those for July 1, 1980 obtained from Bureau of the Census (1990), Table 1. We also assumed the PSID family structure distribution for age 15 applied to ages 16 and 17 as well.

[6] Rendall et al. (1997) find, in both British and U.S. panel data, that reports of men's fertility show an overall deficit of between one-third and one-half of non-marital births, with the deficit especially high for U.S. black men. Non-reporting accounts for most of the deficit in the non-marital births.

[7] The means (standard deviations) for Average Annual Total Family Income (in 1993 $10,000) are: 4.49 (2.9) over all 15 years; 3.7 (2.1) during ages 0–5; 4.6 (3.1) during ages 6–10; and 5.2 (4.1) during ages 11–15. For the measure of whether ever in a non-intact family and for the sequence of family structure types, annual income is averaged over the entire childhood.

[8] The means (standard deviations) for whether there was a 50% or more Income Loss are: 0.12 (0.33) for ages 0–5; 0.15 (0.37) for ages 6–10; and 0.15 (0.36) for ages 11–15.

[9] Regarding the association of ever in a non-intact family to risk of a non-marital birth, Wu and Martinson (1993) estimate coefficients of 0.50 for whites and 0.28 for blacks, both significant at the 0.01 level, and Bumpass and McLanahan's (1989) estimates are similar.

[10] The coefficient (standard error) on the more aggregated version of mother-only structure is −0.406 (0.070) in Wojtkiewicz's (1993) odds of high school graduation regression for sons and daughters combined. The coefficients are 0.56 and 0.32, both significant at the 0.01 level, in Wu and Martinsons' (1993) risk of premarital birth regressions for white and black daughters, respectively.

[11] The coefficient (standard error) for mother-with-stepfather in Wojtkiewicz's (1993) odds of high school graduation regression is −0.075 (0.105). The coefficients in Wu and Martinson's (1993) risk of a non-marital birth regression are 0.47 for whites, significant at the 0.05 level, and 0.23 for blacks, not significant at the 0.05 level.

[12] Moving from the single non-intact family indicator to our stage-specific, type-specific characterization and continuing to omit childhood family income from the predictors, adjusted R-squareds increase from 0.175 to 0.189 in the education regression for boys and from 0.275 to 0.287 in the education regression for girls. The log likelihood in the nonmarital birth equation decreases from −646.4 to −642.6.

[13] The following are the coefficients (standard errors) on average childhood family income when it is substituted for stage-specific family income in the Table 4 regressions: 0.11 (0.03) for sons's education; 0.08 (0.02) for daughters' education; −0.28 (0.08) for daughters' risk of a non-marital birth.

[14] The R-squareds with income averaged over all of childhood are 0.206 for sons and 0.307 for daughters. The log likelihood is −634.6. This compares to R-squareds of 0.211 and 0.312, respectively, and a log likelihood of −630.5 with the stage-specific average family income measures of Table 4.

[15] These additional predictor variables are specified as childhood-stage specific measures of: (a) family receipt of income from the Aid to Families With Dependent Children program (coded 1 if the child's family ever received income from AFDC during the specified childhood stage and zero otherwise); (b) whether the mother devoted substantial time to market work (coded 1 if the child's mother ever worked an average of 1000 or more annual hours during the specified childhood stage and zero otherwise); and (c) whether the child's family moved (coded 1 if the family ever changed residences during the specified childhood stage). Each of these variables is based on information obtained in annual interviews. Appendix Table 1 shows the means and standard deviations.

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Child development and family resources: Evidence from the second generation of the 1958 British birth cohort

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Abstract. Studies of American and recently British children suggest that there is a link between family income and child development, in particular that one consequence of child poverty is to hold back cognitive development. This paper investigates the impact of family income, material deprivation, maternal education and child-rearing behaviour on an indicator of cognitive functioning, using British data on children aged 6 to 17 whose mothers are members of the 1958 Birth Cohort Study. The poorer average cognitive functioning among children from the lowest income groups could largely be accounted for, statistically, by the greater material disadvantage of these groups. These analyses provide evidence to suggest that low income has detrimental effects on children's cognitive functioning through the operation of longer-term material disadvantage, and that these effects may be mitigated by positive parental behaviours.

JEL classification: J13, I10, I30

Key words: Poverty, deprivation, child development

1. Introduction

Poverty among children, as of anyone else, is of concern from the point of view of equity in immediate living standards, but it is also of concern in the longer term, from the point of view of efficiency. Empirical studies in many

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countries show that children who are worst off in terms of their socioeconomic position, are also worst off when it comes to their health and cognitive development (McLanahan and Sandefur 1994; Duncan and Brooks-Gunn 1997; Gregg et al. 1999). During the 1980's poverty among children increased faster in the UK than in any other member state of the European Community (O'Higgins and Jenkins 1990). By the end of the decade nearly one child in three was living below a relative poverty line drawn at half average equivalised income (Gregg et al. 1999). At least in part this reflected an increase in children living in families with an unemployed head, or a lone parent. These families are more likely to live in poorer quality housing and to depend on state benefits. Does this lack of financial and material resources lead, through reduced parents' investment in the 'quality' of their children, to poorer outcomes for children? Can such pathways be identified and can the outcomes be quantified?

Researchers have assessed the impact of poverty and socioeconomic status on various indicators of cognitive functioning during early childhood. A number of US studies that controlled for mother's education, and a range of other maternal characteristics, have reported significant effects of poor circumstance on children's cognitive and verbal skills (for example, Korenman et al. 1995, Chase-Lansdale et al. 1997). Duncan et al. (1994) found that family income and poverty status were significant predictors of IQ scores in five-year-olds, even after accounting for maternal education, family structure and ethnicity and other differences between low and high-income families. Duncan and Brooks-Gunn (1997) assemble a collection of such studies which all include family structure among the controls. Although these studies also indicate that income may not be so powerful a predictor of behavioural outcomes and of achievements by adolescents and young adults, among children, income is at least as important as the absence of a father in predicting cognitive outcomes (McLanahan 1997).

Our own previous study, focussed on family structure, of two cognitive and one behavioural outcome, in samples of British as well as American school-age children, found income to be among the factors which reduced the size and significance of family structure as a predictor (Joshi et al. 1999). However, Lefebvre and Merrigan (1998) point out that income effects, though significant may not be large. The multiple stresses of living in poverty increase children's susceptibility to other problems, for example delinquency and teen pregnancy, and are often manifested as chronic disorders with an early onset (Kratzer and Hodgins 1997). These behaviours do not usually occur in isolation: in fact, they often occur together and have common risk factors including large family size, family discord and parental mental illness. This paper is intended to contribute new evidence to an understanding of the link between family income and child development in Britain.

Although disadvantage goes hand in hand with low income, there is no perfect association (Davies et al. 1997). People observed once in lower income groups are not necessarily deprived of material resources, whereas the absence of resources indicates deprivation only when it is enforced by income. Low income in itself is not an adequate proxy for deprivation, particularly if it is temporary. Hill and Jenkins (1999) examine the extent to which child poverty in Britain is chronic or transitory in nature. They find that children, especially very young children, have high long-term poverty risks compared to other groups in the population.

The causal mechanisms linking relative deprivation and children's outcomes are not well understood. Factors that are probably involved in any direct effect include adverse housing conditions (Macintyre et al. 1998) and a lack of cognitively stimulating resources, such as books, in the home (Baharudin and Luster 1998). Second, the low income family may suffer disadvantages from living in a disadvantaged locality with poor services and low social capital (Brooks-Gunn et al. 1997). Third, relative deprivation might translate into poor cognitive outcomes indirectly, involving among others behavioural factors. Living in a family with a lack of material or social resources might, for example, affect parental feelings of self esteem which is likely to have a negative effect on parenting and mental or emotional well being. Fourth, the child's development and the family's circumstances may both be the joint outcome of other factors, notably parental human capital, a cultural or biological inheritance (Duncan and Brooks-Gunn 1997; Haveman and Wolfe 1995; Hobcraft 1998).

The last decades of the Twentieth Century saw substantial changes in family structure in Britain. Lone parents with dependent children constitute an increasing proportion of households with children: 19% of children lived with a lone mother in 1994–1995 (Church 1996). The increase in the number of lone parent families since 1972 can be attributed mostly to a rising rate of partnership dissolution and also some growth in single motherhood. Over one third (34%) of births were outside of marriage in 1995, although roughly half of these were registered by two parents living at the same address (Office of National Statistics 1997). Currently 41% of marriages are projected to end in divorce in England and Wales, if hitherto rising divorce rates remain at their 1993–1994 levels (Haskey 1994). Similarly, children born to married parents face a 28% risk that their parents would divorce before they reach age 16 (Haskey 1997).

There have also been major structural changes in the labour market. The number of households without an employed member rose sharply in the recession of the 1980's but nearly all the subsequent recovery in employment occurred in households with one person already in work (Gregg and Wadsworth 1996). Women with a pre-school child were least likely to be in paid work in 1991 (Church 1997). Although economic activity has increased among married women (Office of National Statistics 1998), less than one in five lone mothers with a child under five years was in paid employment. As a consequence of the changes in labour market activity and family structure, the share of men's earnings in total family income fell dramatically over the 1980's (Harkness et al. 1996). Families with children were over-represented at the bottom of the income distribution, especially where there was no father in employment. (Gregg et al. 1999). Low income families are more likely to live in deprived physical environments than those higher on the income scale, with the majority of families receiving state benefits living in social rented properties (Marsh and McKay 1993; Macintyre et al. 1998).

In this paper, we examine the association of a range of indicators of socioeconomic position including family income and other indicators of material disadvantage with one indicator of children's cognitive functioning. The aim of the paper is, firstly, to examine whether income is independently associated with children's cognitive outcomes; secondly, to assess how far such association reflects high levels of disadvantages in low income groups; and thirdly, to examine which specific disadvantage contributes most to the link between income and child outcomes. This might elicit indications of any mechanism by which deprivation affects children's outcomes, though it cannot prove causation.

2. Data and methods

The National Child Development Study (NCDS) is a study of over 17,000 people in Britain born between the 3rd and 9th of March in 1958. Follow-up sweeps took place in 1965, 1969, 1974, 1981 and 1991. The 1991 NCDS follow-up obtained information not only from the cohort member, but from the children of 1 in 3 randomly sampled cohort members (Ferri 1993). In this paper we restrict the analyses to children who were aged 6 or older in 1991 and who had a mother who was an NCDS cohort member, and who were resident with her in 1991. This excludes only a small number of children who were not living with their cohort member mothers (see Joshi et al. 1999). Children with fathers in the cohort were not included because the information on them was in various ways less complete. Any children over 13 would have been born to a teenage mother. Thus the sample design imposes an artificial inverse correlation of the ages of children and the age of the mother at birth. The data are more representative of children with teenaged, and hence less educated, mothers, than a full cross-section of children. As shown in Table 1, the family structures in which the children were living are strongly related to the age of the child and therefore, to the age of the mother at the time the child was born. Younger women are more likely to have births outside partnerships; and older children have had more chance of experiencing a change in the family situation (Clarke et al. 1997).

2.1. Cognitive functioning

In this paper we use performance on the Peabody Picture Vocabulary Test (PPVT) as an indicator of children's cognitive functioning. The PPVT is an individually administered test of hearing vocabulary, with 175 test items arranged in order of increasing difficulty. Each item presents a multiple choice between four illustrations, the aim being to select the picture best illustrating the meaning of a stimulus word, presented orally by the examiner.

We selected the PPVT score as a broad single-variable measure of children's cognitive ability, after earlier work with other outcome measures, such as maths and reading and behavioural adjustment scores. One of the most basic skills that children need in order to succeed in school is the ability to use language. When children do not learn to use vocabulary, their general knowledge, their spelling, writing and reading abilities suffer. Children's vocabulary development serves as a major foundation for all school-based learning, and without it, the chances for academic and occupational success are limited. Vocabulary subscales of tests of children's cognitive abilities correlate more highly with full scale IQ scores than any other subscales (Wechsler 1974), and because the median correlation between the PPVT and the full scale Stanford-Binet (IQ test) across a number of tests is 0.71 (Dunn and Dunn 1981), we can assume that the PPVT scores reported in this paper are related to the scores

Table 1. Means of variables by family structure and mothers current employment status

	Complete sample	Intact family, non working mother	Intact family, working mother	Step family, non working mother	Step family, working mother	Lone mother, non working	Lone mother, working
PPVT test score	0.59	0.56	0.58	0.65	0.62	0.57	0.60
Child age (months)	118.59	112.21	115.18	140.60	135.65	121.53	126.04
Girl	0.51	0.46	0.53	0.44	0.59	0.53	0.51
[a]Imputed household income	288.18	259.09	347.18	213.52	355.62	99.01	166.81
[b]Household income	292.33	254.32	346.71	209.14	348.94	101.21	155.34
Father not employed	0.08	0.16	0.05	0.35	0.04	0.02	0.00
Number of children	2.39	2.54	2.25	2.81	2.31	2.66	2.39
Teen mother	0.19	0.16	0.14	0.40	0.30	0.33	0.24
[c]Mother's qualifications	1.87	1.82	2.10	1.32	1.65	1.52	1.54
Social housing	0.29	0.23	0.21	0.65	0.27	0.55	0.61
Car access	0.85	0.87	0.95	0.84	0.91	0.40	0.38
Holiday less than once a year	0.21	0.28	0.14	0.35	0.27	0.40	0.27
Eat out 2 or 3 times a month	0.39	0.32	0.43	0.30	0.32	0.29	0.52
Current benefit receipt	0.13	0.13	0.01	0.19	0.04	0.93	0.41
Ever benefit receipt	0.39	0.35	0.23	0.81	0.53	0.95	0.65
Past benefit receipt	0.33	0.27	0.23	0.72	0.53	0.43	0.56
Spend evening together once a week	0.94	0.95	0.94	0.88	0.90	0.97	1.00
Home: Emotional support	10.45	10.38	10.91	10.74	10.29	8.65	8.98
Home: Cognitive stimulation	10.53	10.65	10.87	9.72	10.12	9.28	9.87
Number of cases	1041	262	499	56	95	58	71
Number of cases with reported income	725	178	347	46	64	43	50

[a] Household income for reported cases.
[b] Household income for reported and imputed cases.
[c] Mother's highest educational qualification: 0 = none to 5 = degree.

these children would obtain on an IQ test. Verbal ability as measured by the PPVT is predictive of literacy scores as much as fifteen year later, even after controlling for the effects of educational, social and economic well-being (Baydar et al. 1993).

The PPVT was standardized nationally (in USA) on a sample of 4200 children and adolescents and 828 adults. Raw scores are usually converted to age-referenced norms. We have however used raw scores in our analyses because of concern about the suitability of the available norms for our 'abnormal' sample of older children. We divide the test scores for each child by the total number of items in the test. The test scores range between a minimum of 0.19 and a maximum of 0.98 and have a standard deviation of 0.12. We include age and age squared in the regressions to control for age, and we always also include a dummy variable for child gender.

2.2. *Income*

Although recent research emphasizes the life-course dynamics of poverty and low-income we have had to use current 'household' income in this analysis. 'Household' income is computed by summing across the income of the cohort member and, if present, partner. If there were any other income recipients in the household, their income was not included. However, at age 33, cohort members who had their own children were extremely unlikely to be living with their own parents, and very few of their children would have been old enough to earn. Unfortunately income data collection in surveys presents a number of problems. It is subject to high levels of item non-response both because respondents regard it as a sensitive subject and because they may not immediately know the answers. It is also likely that income data are subject to measurement error but this issue is not dealt with in this paper. In any case there is commonly a substantial item non-response, and this should be of concern to the data analyst.

Where income was missing, data were imputed. This was done to reduce potential bias arising from exclusion of 30% of cases with missing data. We adopted a multiple imputation technique following Schafer (1997; Appendix). The missing income values were imputed as a linear function of the following variables: the presence of a partner, his school leaving age, the mother's highest educational qualification, number of earners in the family, whether the family had access to a car and whether they live in social housing. Cases where the mother is currently a lone parent form the base category to age partner left full-time education. The impact of the method of imputation on our results was examined by imputing a sub-sample of NCDS data for which we have recorded responses, so that we do have a measure of the 'true' value for each case. A random sample of approximately 20% of cases are reset to missing and their data imputed using multiple imputation methods. Figure 1 shows the income values for the complete cases and their imputed values. The imputed values are close to the recorded values indicating that the imputation method is effective in reproducing the underlying pattern of relationships in the data.

Our approach assumes that the missing income values are missing at random (MAR), conditional on the imputation model. For processes that are

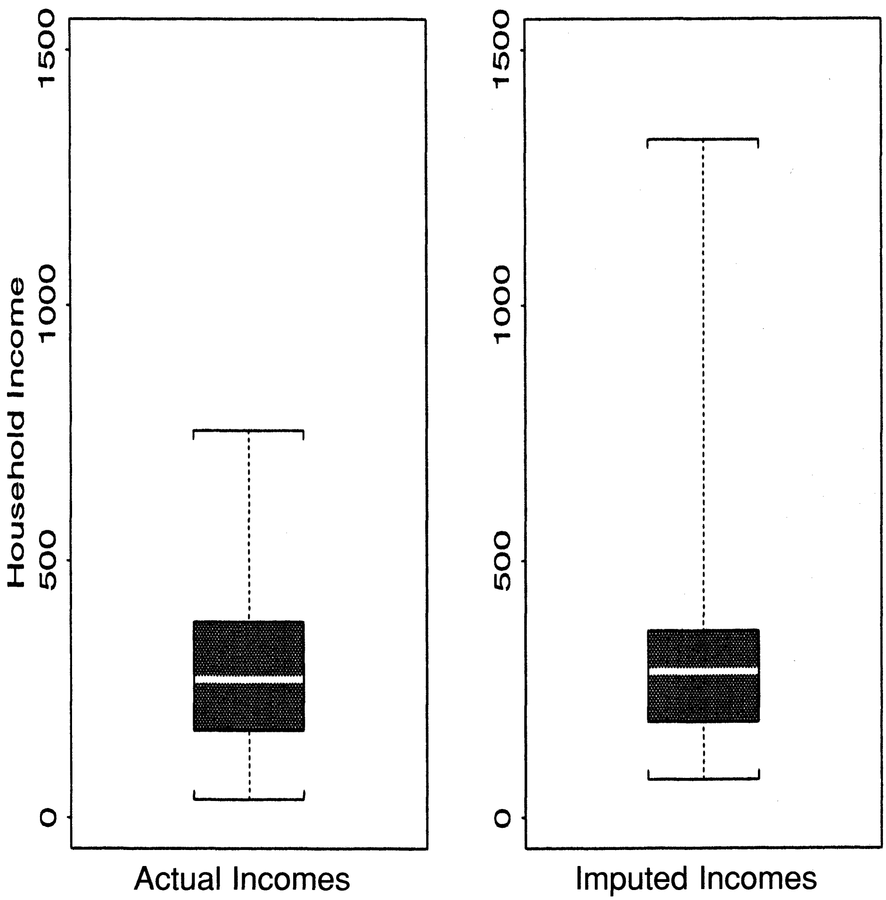

Fig. 1. Boxplot of actual income and imputed income. The box shows the limits of the middle half of the data; the line inside the box is the median. The vertical line shows the full data range

MAR, the probability that a value is missing may depend on the observed data, but it may not depend on values of the data that are unobserved. MAR assumptions can be made more applicable and more powerful by including more variables in the imputation process to help predict the pattern of missingness. If the probability that a value is missing depends in part on the unobserved value of the missing response, the process is said to be non-ignorable. Table 2 contains sample means for a range of variables, separately for the households with reported and missing income values and two-sample t-statistics. As can be seen, the households with missing income values, were less likely than the households with reported income values to contain an unemployed father or to be resident in social housing but were more likely than the households reporting income to have access to a car. The estimates of mean imputed income for those 316 children where income is unknown exceed the mean of income in observed cases. This implies that the average measured attributes of those with unmeasured income were characteristic of above-

Table 2. Comparison of sample means for a range of covariates for cases with reported and missing income values

	Income		2 sample t-statistic
	Present	Missing	
PPVT test score	0.59	0.59	0.18
Child age (months)	119.51	116.47	0.15
Girl	0.51	0.50	0.79
Father not employed	0.11	0.04	0.001
Number of children	2.44	2.29	0.06
Teen mother	0.20	0.16	0.15
Mums qualifications	1.90	1.82	0.42
Social housing	0.32	0.23	0.01
Car access	0.83	0.89	0.001
Holiday less than once a year	0.19	0.25	0.05
Eat out 2 or 3 times a month	0.39	0.37	0.51
Current benefit receipt	0.14	0.11	0.11
Ever benefit receipt	0.39	0.38	0.82
Past benefit receipt	0.32	0.34	0.52
Spend evening together once a week	0.93	0.96	0.07
Intact family + non working mum	0.24	0.27	0.45
Intact family + working mum	0.48	0.49	0.78
Step family + non working mum	0.06	0.03	0.06
Step family + working mum	0.09	0.10	0.52
Lone mother + non working	0.06	0.05	0.43
Lone mother + working	0.07	0.07	0.87
Home: Emotional support	10.38	10.62	0.04
Home: Cognitive stimulation	10.49	10.61	0.38
Number of cases	725	316	

average income, implying a lower response rate to income questions among respondents of high socio-economic status. Income is therefore not missing *completely* at random (MCAR), but this is not problematic as long as the variables included in the imputation procedure can predict which households have high income.

It is also necessary to adjust reported income for family size, which we do by dividing by the number of co-resident family members. We have thus given every person in the family the same equivalence scale. This is supposed to be transparently arbitrary, as it is not agreed which of other scales on offer is appropriate; we do not know that the implicit pooling of income takes place; and the conventional expenditure approach to the costs of children omits indirect, or forgone earnings costs.

We categorised income into 5 quintile groups. As children in low income families are probably over-represented in our sample (because they have relatively young mothers), it is likely that most of the bottom two quintiles would fall below the de-facto relative poverty line, 'below half average income' used in national statistics, although we have not derived exactly the same measure of equivalized income used in the latter. Since we only have one cross-section of income, moves between income quintiles represent both absolute and relative changes in income.

2.3. *Family structure*

In all but the first model we include a set of variables which describe the number of people in the child's family, whether the mother and father, if present, have paid work and whether the mother started child-bearing as a teenager. The latter term is intended to allow for the peculiarities of a sample of the children of one cohort, and appears as a dichotomy because age of mother at first birth is highly collinear with other terms such as child age and maternal qualifications. The number of children gives some idea of the competition for parental time. The presence of a father is included on the presumption that children thrive when they have the attention of more than one adult, particularly if the father is the child's natural father. Children living with step fathers are likely to have had some experience of change and or lone parent living, and the step father may not devote as much attention as a natural father to the index child. The employment of the mother is interacted with family structure, because there can be no more than one earner in a lone parent family. The employment of the mother may be expected to have mixed effects on a child's development. It may deprive a child of maternal time, though this depends on the time they do spend together; on the other hand, it may, depending on the quality of alternative care, hasten the child's social and psychological development. Since these effect are likely to be lagged, we might not expect to find much of an association of vocabulary score with current employment, but perhaps there is a weak presumption that children whose mothers are in employment might do somewhat worse on their test score.

Table 1 shows that family disruption is highly associated with early motherhood, low maternal educational qualifications and social housing (i.e. rented from council or housing association). There are differences in economic activity between family types: of these school age children living with lone mothers, 54% had an employed mother compared to 58% in step families and 66% in intact families. In accordance with the differences in employment status by family structure, lone parent and step families where the mother is not earning have the lowest mean family incomes. The polarisation of unemployment is apparent in the concentration of non-earning fathers in families with a non-earning mother (Gregg and Wadsworth 1996; Davies et al. 1998), particularly here in the step families. 35% of children living in step families with a non-employed mother also have an unemployed step father. In comparison, 16% of children living in intact families with a non-employed mother have an unemployed father. A larger proportion of children living in lone and step parent families than intact families have a parent with a history of drawing means-tested benefit – 81% and 95% of children who have a non-employed mother living with step fathers or with no partner, respectively, compared with 23% in intact families with an employed mother.

2.4. *Family resources*

Current income may not reflect lifetime income or assets such as saving or earning power. As an indicator of maternal resources we include the mother's educational attainment. Preliminary analyses suggested a linear trend in PPVT scores across the categories of educational attainment so this variable was entered in the model as a linear term. We examine a number of other

Table 3. Means of variables by quintile of family income

	[a]Income quintile				
	[b]I	II	III	IV	V
PPVT test score	0.58	0.60	0.59	0.58	0.59
Child age (months)	124.65	122.05	120.57	113.28	112.63
Girl	0.49	0.53	0.53	0.48	0.52
Father not employed	0.33	0.06	0.02	0.00	0.00
Number of children	2.57	2.48	2.40	2.26	2.23
Teen mother	0.27	0.22	0.21	0.11	0.12
Mums qualifications	1.38	1.62	1.83	2.12	2.50
Social housing	0.57	0.35	0.23	0.18	0.07
Car access	0.53	0.86	0.98	0.95	0.96
Holiday less than once a year	0.32	0.25	0.24	0.14	0.14
Eat out 2 or 3 times a month	0.36	0.30	0.36	0.40	0.52
Current benefit receipt	0.52	0.07	0.03	0.00	0.00
Ever benefit receipt	0.77	0.36	0.33	0.25	0.18
Past benefit receipt	0.52	0.33	0.33	0.25	0.18
Spend evening together once a week	0.99	0.94	0.94	0.91	0.91
Intact family + non working mum	0.29	0.39	0.27	0.16	0.14
Intact family + working mum	0.08	0.43	0.54	0.71	0.67
Step family + non working mum	0.12	0.07	0.04	0.03	0.00
Step family + working mum	0.03	0.04	0.14	0.09	0.18
Lone mother + non working	0.25	0.00	0.00	0.00	0.00
Lone mother + working	0.23	0.06	0.02	0.00	0.00
Home: Emotional support	9.64	10.66	10.55	10.78	10.74
Home: Cognitive stimulation	9.61	10.47	10.46	10.80	11.44

[a] Mean of 5 income imputations for missing cases
[b] Income quintile cut points are 176.0, 244.5, 314.6 and 379.2 pounds.

indicators of material deprivation which may also influence child development. These include social housing, car access and whether the family has an annual holiday or goes out to eat at least two to three times a month, and previous receipt of an income related state benefit either by the mother or partner since age 16 (but not currently receiving a means tested benefit as this is too closely linked to income quintile). As shown in Table 3, all indicators of material comfort increased in prevalence from the lowest to the highest income quintile. PPVT (unstandardised) changes little because the children in the higher income brackets are on average younger.

To assess the impact of parenting behaviour as an independent or mediating influence on child development, we look at whether the family spend an evening together at least once a week, and two assessments of the quality of children's home environments: the provision of cognitive stimulation and warmth of parental emotional support. These are available in our data from the interviewer's assessment of cognitive stimulation in the home and of emotional support provided by the mother during the interview (assessed on the short form of the HOME inventory, Sugland et al. 1995). This score has been found to account for a substantial portion of the effects of family income on children's PPVT test scores (Smith et al. 1997). In preliminary analyses we also found a significant association between the mother's stated aspiration for the child to stay on at school after 16 and the PPVT score, but do not include it in the final model because of possible reverse causation, perhaps through the downward revision of aspirations as children grow older and the limits of their capabilities become apparent. All other regressors are assumed to be exogenous.

2.5. Statistical models

To model the cognitive functioning of children within families we use the framework of the hierarchical linear model (Goldstein 1995). This is a variant of the multiple linear regression model for data with a hierarchical nesting structure. First consider a two-level multi-level model of children nested in families. Children (level-1 units) are indicated by i and families by j. The dependent variable must be defined at the lowest level, that of the individual, and it is denoted by Y_{ij}. A simple two level model can be formulated as:

$$Y_{ij} = \beta_{0j} + \beta_1 x_{ij} + e_{ij} \tag{1}$$

where Y_{ij} is the value of the dependent variable, β_{0j} is the family-specific intercept, β_1 is the fixed regression slope, x_{ij} is the value of the explanatory variable, and e_{ij} is the unexplained part of the dependent variable Y_{ij}. It is convenient to separate the coefficient β_{0j} in (1) into a fixed part (the mean) and a random part (with mean 0):

$$\beta_{0j} = \gamma_{00} + U_{0j} \tag{2}$$

where γ_{00} is the population mean of the intercepts and U_{0j} is the family-specific part of the intercept. Substitution of the models describing the variation of the coefficients between families into (1) then yields the combined model formula:

$$Y_{ij} = \gamma_{00} + \beta_1 x_{ij} + U_{0j} + e_{ij} \tag{3}$$

This is often referred to as a variance components model. The model contains two random effects: U_{0j} and e_{ij} which are both assumed to have $N(0, \sigma^2)$ normal distributions. Each of these indicates a different source of unexplained variation. The random intercept U_{0j} indicates unexplained differences between families in the average Y-values (controlling for the effect of x_{ij}). The random residual e_{ij}, indicates unexplained variation among the individuals, relative to their families. In the present exercise we analyze 1041 children (level 1) of 729 mothers (level 2).

It is appropriate to treat U_{0j} as a random effect because the objective is to use the individual observations to make inferences about the population of children. A fixed effects model would treat U_{0j} as a fixed variable. While appealing in that few assumptions need to be imposed on U_{0j}, this procedure has the drawback that it implies that out interest centres on the outcome of an individual child. Furthermore, and perhaps more restrictive, only the effects of covariates that change between siblings can be estimated.

3. Results

The results from estimating variance components models for PPVT test scores are presented in Table 4. The income coefficients are averages of the coefficients from separate analyses using 5 imputation estimates of family income (Schafer 1997). The P values for the income coefficients are computed using the percentiles of the t-distribution (Rubin 1987). The analyses are organised as follows: Model A contains age terms, the child's gender and income quintiles. Model B adds a summary of the child's family structure and current employment status of the parents. Model C adds controls for maternal qualifications. In comparison to Model C, Model D substitutes indicators of material deprivation for maternal qualifications and Model E substitutes indicators of parenting and the home environment for maternal qualifications. Model F examines the effect of all variables simultaneously. The comparison of results from this sequence of models allows us to see how far PPVT scores are associated with family income. It also allows us to see how far any association might be accounted for, or mediated by the social and human capital available to the child, the material disadvantage associated with the child's family circumstances, and the mode of parenting such resources may permit.

In Model A mean PPVT test scores increased with increasing income in comparison to the lowest income quintile controlling for child age and gender. The difference of test scores from the 1st quintile is statistically significant from the 3rd quintile upwards but the coefficients are of a modest order of magnitude: the top fifth of children by family income have PPVT scores 3.7% points higher on average than children in the lowest income bracket. This is not large in comparison with the age-adjusted standard deviation of PPVT at 7.5 percentage points. The estimates of the random part of the model suggests that families do differ in their average PPVT scores and that there is even more variation among children within families. The variance component between children (e_{ij}) in Model A is around one and a half times that between families (U_{0j}).

Introduction of information on the structure and number of earners in the family confines the statistically significant difference from the bottom income quintile to the top 2 quintiles (Model B). Given the control for income level, the family structure terms themselves do not predict much systematic advantage to having two natural parents, nor in general, much disadvantage to having a mother in employment. Children in step families appear at a disadvantage if the mother is employed. Unemployment of the father has a negative, but non-significant coefficient. Cognitive outcomes are lower in families with large numbers of children and in families where the mother had her first child when aged under 20. This term does not affect the estimated age coefficients, as it might have done, had the latter contained a strong age-at-parenthood element.

Introduction of information on the mother's qualifications (Model C) removes the remaining significant income coefficients (given the critical value for the t statistic on the imputed income variables is 3.1). Information about the maternal endowment seems to act as an alternative signal to current income of the resources available to the child. It also attenuates the teenage mother effect. As an indicator of human capital transmitted by at least one parent, maternal qualifications are a strong predictor of high test scores and teen motherhood a strong predictor of low test scores (McCloyd 1998; Joshi

Table 4. Multilevel regression results for 1041 children with complete data on background and family characteristics

	[a]Model A		Model B		Model C		Model D		Model E		Model F	
	β	t	β	t	β	t	β	t	β	t	β	t
Fixed effects												
Constant	0.18	0.53	0.59	1.68	−0.02	−0.06	0.24	0.67	−0.42	−1.07	−0.82	−2.10
Age (months)	0.06	11.48	0.06	10.09	0.06	11.16	0.06	10.42	0.06	9.84	0.06	10.85
Age-squared/100	−0.01	−5.86	−0.01	−4.39	−0.01	−5.22	−0.01	−4.59	−0.01	−4.10	−0.01	−4.91
Girl	−0.05	−1.07	−0.05	−1.03	−0.03	−0.74	−0.04	−0.85	−0.07	−1.58	−0.04	−1.02
[b]Income II	0.10	2.99	0.01	0.32	0.01	0.24	−0.01	−0.24	0.00	−0.08	−0.03	−0.44
Income III	0.19	3.84	0.10	1.64	0.08	1.58	0.05	0.76	0.05	0.87	−0.01	−0.16
Income IV	0.31	8.69	0.21	3.80	0.11	2.31	0.13	2.24	0.13	2.15	0.05	0.73
Income V	0.37	10.05	0.28	4.75	0.18	2.87	0.19	2.99	0.20	4.06	0.09	1.43
Intact family + working mum			0.01	0.12	0.00	−0.05	0.01	0.18	0.01	0.16	0.00	0.04
Step family + non working mum			0.16	1.33	0.18	1.53	0.22	1.80	0.17	1.49	0.22	1.90
Step family + working mum			−0.22	−2.21	−0.20	−2.10	−0.18	−1.80	−0.15	−1.58	−0.13	−1.41
Lone mother + non working			−0.17	−1.34	−0.16	−1.31	0.03	0.27	−0.07	−0.52	0.05	0.36
Lone mother + working			−0.03	−0.29	0.00	0.02	0.21	1.77	0.07	0.63	0.26	2.24
Father not employed			−0.18	−1.72	−0.14	−1.37	−0.07	−0.67	−0.19	−1.84	−0.09	−0.85
Number of children			−0.05	−2.33	−0.05	−2.51	−0.05	−2.06	−0.04	−1.74	−0.04	−1.91
Teen mother			−0.24	−2.82	−0.17	−1.99	−0.23	−2.74	−0.19	−2.24	−0.15	−1.84
Mums qualifications					0.15	7.80					0.11	5.71
Social housing							−0.19	−2.99			−0.12	−2.02
Car access							0.32	3.90			0.25	3.10
Holiday less than once a year							0.04	0.65			0.08	1.28
Eat out 2 or 3 times a month							0.07	1.35			−0.01	−0.21
Past benefit receipt							−0.06	−1.07			−0.05	−0.92
Spend evening together once a week									0.08	0.77	0.06	0.58
Home: Emotional support									0.04	2.75	0.03	2.44
Home: Cognitive stimulation									0.06	4.70	0.04	2.96
Random effects												
U_{0j}	0.41	10.71	0.39	10.23	0.35	8.82	0.37	9.48	0.37	9.53	0.32	7.80
E_{ij}	0.63	26.70	0.62	26.73	0.62	26.94	0.62	26.91	0.62	26.98	0.62	27.14

[a] All coefficients are ×10.

[b] *P* values for income coefficients are computed using the percentiles of the *t*-distribution (Schafer 1997). 5% significance levels for income coefficients are represented by a critical value of approximately 3.10.

et al. 1999). The significant (adverse) term for an employed mother living with a step family remains.

In Model D, indicators of material deprivation rather than mother's qualifications also remove the statistically significant impact of income on children's test scores (and the special case of the step families with employed mothers). The indicators of material deprivation which account for most variation in the test scores, (and also for the association between test scores and family status), are social housing and lack of a car. Children who are growing up in such a materially deprived family are disadvantaged in their cognitive development at all levels of current family income, though as Table 3 shows, most of the families without cars or in social housing do have low income. As Table 3 also shows, not all of the families with low current income are social tenants or without cars. The finding of this model suggests that it is in those low income families who are in these categories that children are falling behind.

Model E substitutes parenting behaviour for the material deprivation terms. The level of maternal emotional support and the level of cognitive stimulation in the home (e.g. academic and language) both show statistically significant positive associations with children's test scores. Whether the family spends time together (at least one evening a week) does not show much predictive power. Introduction of these terms (on top of those in Model B) confines the statistically significant impact of income on test scores to the top income quintile. Our measure of a stimulating home environment is almost as successful in explaining the association of children's test scores with family income as the measures of more permanent living standards and social status in Model D.

In our model containing all predictors (Model F), all the income terms become insignificant. We have thus succeeded in our search for factors which lie behind the association we reported in Model A. Other coefficients from previous models show little change in sign or statistical significance. One exception is that the teenage mother term becomes insignificant. The distinguishing characteristics of teenaged mothers have been captured elsewhere. Another distinction of the full model is the statistically significant positive estimate that children living with an employed lone mother do better, for a given level of maternal and family resources, than children in intact families with non-employed mothers.

4. Discussion

We started with a model of vocabulary attainment (PPVT) where the only explanatory variable, apart from basic controls for age and sex, was a five-fold grouping of income. This showed a significant relationship with test scores. We then introduced various combinations of other variables through which an income effect might work, or which might in themselves offer superior explanatory power. The introduction of measures of more permanent living standards and social status, particularly car access and housing tenure, eliminates the income differentials in test scores. We conclude that current income alone is not a complete yardstick for children at risk of impaired development. Our indicators of more long-term deprivation (car access and housing tenure) suggest that a more sustained experience of poverty is more damaging. The indicators of income and deprivation taken together show a more robust relationship with the PPVT than does, for the most part, family structure and current parental employment. The relative importance of material resources we find here is consistent with the analyses of the NCDS first generation offered by Gregg et al. (1999) and the American findings reported in the studies assembled by Duncan and Brooks-Gunn (1997) with respect to academic and cognitive outcomes of children of school age. It should also be noted that these estimates of income effects on PPVT scores of children aged 6–17, are modest, like those of Lefebvre and Merrigan (1998) for Canadian children aged 4–5. It would take very large increases in income to produce a great effect on children's proficiency at the vocabulary test. Model A says that a child from the bottom income quintile has a score 3.7 points below an otherwise identical child in the top quintile. This extreme move in income is equivalent to half an age standardized standard deviation in PPVT. The parameters of Model F imply that to move a child out a situation where the father was unemployed, the family lived in council housing, took a holiday less than once a year and ate out less than 2 or 3 times per month to the opposite of all these, would imply an increased PPVT score of 5 points. We do not directly replicate the finding of Lefebvre and Merrigan (1998) that income effects are strongest at very low levels of income.

Other variables contributed, at least partially independent, explanatory power. Our measure of the endowment of parental human capital in terms of mother's educational qualifications, was uniformly associated with higher PPVT test scores, and this relationship was independent of the household's economic status. To the extent that this endowment can be acquired through education, an across-the-board improvement of mothers' education by one grade would raise the average PPVT score by 1.1 percentage points. Although this would not be cheap or immediate, this long term effect might be equivalent to that of raising into home and car ownership all families currently without them (Model D parameters). Turning to variables about parenting, the findings on the home environment, as assessed on the HOME scale, indicate that children with better home environments appeared able to overcome the barriers set by low material circumstances. Of the two types of environment, the provision of stimulating and learning oriented experiences appeared to exert a greater influence on children than the degree of parental encouragement. Although our data are limited in the extent to which we can measure family processes, there is some evidence that parental 'competence' plays an important role which is independent of income. The measures of the home environment are not as successful in explaining the association of children's test scores with family income as the measures of more permanent living standards and social status. It appears that parenting differences are not the principal route through which children in more advantaged families benefit from parents' cash resources. A similar conclusion is reached by Hanson et al. (1997) although other US studies they cite had found a connection from poverty, or perceived financial stress, to poor parenting. Smith et al. (1997) found that the inclusion of Home scores in the analysis of PPVT scores of US children aged 5 mediated the mother's education term more than the income coefficient.

We found that family income (as we have been able to measure it) is only weakly related to effective parenting. Consequently, differences in parenting do not account for much of the association between economic resources and children's test scores. We examined the extent to which the effect of mother's education worked through income rather than income and maternal education having independent effects. Mother's education was found to contribute to child test scores even after controlling for family income, in contrast to our earlier results for family structure (Joshi et al. 1999). Improved mother's education raises children's emotional adjustment through complex mechanisms. Education may provide mothers with the knowledge, skills and self-confidence necessary for happy marital relationships and successful child rearing. On the other hand, lack of educational qualifications may also be a more general indicator of inherited social disadvantage. Perhaps our most surprising finding is that we find little evidence that a father's absence is related to children's test scores. Indeed living with a working single mother increased the test scores of children. This could be due to selection into employment, but it echoes a finding of Kiernan (1996) on the previous generation of NCDS. The daughters of lone mothers who were employed when the girls were aged 16 showed no disadvantage vis a vis their counterparts in intact families. Kiernan suggests that a positive role model may be part of the explanation. Our results may also be due to the fact that the children in the NCDS were born to younger than average mothers. Whatever the reason, the overall pattern of results obtained on children in single-parent families suggests that their outcomes are not significantly different from those exhibited by their two-parent counterparts, controlling for levels of economic and parenting resources.

The implication that material disadvantage can, at least partly, be overcome by positive parental behaviour offers one ray of hope for the future prospects of children growing up in persistent poverty. A substantial body of research has shown that parental warmth, involvement, and moderate control facilitate children's adjustment and achievement (McLoyd 1998). Direct observations of family interactions also suggest a link between parenting behaviour and economic stress. Economic pressure and marital conflict may lead to financial conflict and hostility between parents and adolescents. Economic deprivation also appears to be the best predicted concomitant of divorce, consequently placing a strain on the family's lifestyle, relationships and opportunities (Conger et al. 1990). Children may perceive remarriage as compounding these emotional stresses. Nevertheless, it can improve the possibility of financial support. Our finding of poor scores in stepfamilies with an employed mother (Models B and C) is a bit of a puzzle, because these should be the stepfamilies with the most cash resources, but their association with the other deprivation indicators (see Table 1) suggests that they are not particularly comfortably off.

The unobserved family effects in our analyses are substantial and point to major sources of unobserved variation between families independent of household income, material disadvantage, maternal education and measured parenting. Finally the unobserved variation among children points to the big role of chance (or other unobserved factors like individual motivation) in distributing success and failure in dealing with a vocabulary test, if not the rest of life. More recent evidence suggests that the level of income in neighbourhoods, over and above family income, is also associated with early school age developmental outcomes (Brooks-Gunn et al. 1993). We cannot begin however to allow for any teacher, school, or neighbourhood effect with the present dataset, although the public housing indicator may contain some information about the neighbourhood.

While it is customary to bemoan the lack of income data in many British data sets, this evidence suggest that current income is not the best indicator of families at risk, if evidence of more sustained long-term deprivation is available. A general picture emerges of an accumulation of risk factors. These factors include a lack of economic resources and having mothers with low educational attainment and little involvement in their children's development. Each factor is associated with a gradual increase in the risk of a poor outcome rather than a threshold precipitating poor outcomes. Poor cognitive outcomes during childhood increase the risk of many other negative outcomes during childhood, such as poor school performance and weak social skills, and also predicts an accumulation of economic and social problems over children's subsequent lives (Kratzer and Hodgins 1997; Hobcraft 1998). Although our evidence of a link from economic to educational disadvantage is systematic and statistically significant, it is not spectacular – small changes in families economic circumstance could not be expected to yield more than small changes in cognitive attainment. Our results are consistent with the growing realisation that causal structure of cognitive outcomes in children is likely to be complex. There are few interventions that will single-handedly yield significant improvements in children's abilities, but reducing child poverty should help.

Appendix: Income imputation

There are a number of methods for dealing with missing data. The most common approach is simply to eliminate entire observations when any one variable is missing (listwise deletion). Other general purpose methods include mean substitution (imputing the univariate mean of the observed observations), imputing a zero and then adding an additional dummy variable to control for the imputed value, and hot deck imputation (imputing a complete observation that is similar in as many observed ways as possible to the observation that has a missing value). However, even if the answers we make up for nonresponding respondents are right on average, these procedures considerably overestimate the certainty with which we know those answers. Consequently estimated standard errors will be too small.

Statisticians and methodologists have agreed on a widely applicable approach to many missing data problems based on the concept of multiple imputation. Implementing multiple imputation requires a statistical model from which to compute the imputations for each missing value in a data set. One model that has proven to be useful for missing data problems in a surprisingly wide variety of data types, assumes that the variables are jointly multivariate normal. Let Y denote an $n \times r$ matrix of multivariate data where each row of Y is a joint realization of variables $Y_1, \ldots, Y_r$. Denote the complete data by $Y = (Y_{\text{obs}}, Y_{\text{mis}})$ where Y_{obs} and Y_{mis} are the observed and missing portions of the matrix respectively. We assume that $Y_1, \ldots, Y_r$ have a multivariate normal distribution with mean vector μ and covariance matrix Σ; that is:

$$y_1, y_2, \ldots, y_r | \theta \sim \text{iid}\, N(\mu, \Sigma)$$

where $\theta = (\mu, \Sigma)$ is the unknown parameter. In multiple imputation, one must generate M independent draws $Y^{(1)}_{\text{mis}}, \dots, Y^{(M)}_{\text{mis}}$ from a posterior predictive distribution of the missing data:

$$P(Y_{\text{mis}}|Y_{\text{obs}}) = \int P(Y_{\text{mis}}|Y_{\text{obs}}, \theta)P(\theta|Y_{\text{obs}})\,d\theta$$

where $P(\theta|Y_{\text{obs}})$ is proportional to the product of the observed-data likelihood function:

$$P(\theta|Y_{\text{obs}}) = \int L(\theta|Y)\,dY_{\text{mis}}$$

and a prior density function $\pi(\theta)$. After imputation, the resulting M versions of the complete data are separately analyzed using complete-data methods, and the results are combined to obtain inferences that effectively incorporate uncertainty due to missing data. The difficulty in using this model is taking random draws from the posterior distribution $P(Y_{\text{mis}}|Y_{\text{obs}})$.

It is possible to create random draws of Y_{mis} from $P(Y_{\text{mis}}|Y_{\text{obs}})$ using techniques of Markov Chain Monte Carlo (MCMC). Consider an iterative simulation algorithm in which the current version of the unknown parameter $\theta^{(t)} = (\mu^{(t)}, \Sigma^{(t)})$ and the missing data $Y_{\text{mis}}(t)$ are updated in two steps:

$$\theta^{(t+1)} \sim P(\theta|Y_{\text{obs}}, Y^{(t)}_{\text{mis}})$$

$$y^{(t+1)}_{i(\text{mis})} \sim P(y_{i(\text{mis})}|Y_{\text{obs}}, \theta^{(t+1)})$$

Given starting values $\theta_{(0)}$ and $Y_{(0)\text{mis}}$, these two steps define a Gibbs sampler in which the sequences $\theta(t)$ and $Y_{(t)\text{mis}}$ converge in distribution to $P(\theta|Y_{\text{obs}})$ and $P(Y_{\text{mis}}|Y_{\text{obs}})$, respectively. Implementation of the Gibbs sampler requires us to specify a prior distribution for θ. For simplicity we apply an inverse-Wishart distribution $\Sigma^{-1} \sim W(v, \Lambda)$ where $W(v, \Lambda)$ denotes a Wishart distribution with $v > 0$ degrees of freedom and mean $v\Lambda > 0$. Small values for v makes the prior density relatively diffuse, reducing its impact on the final inferences. For μ, we use a conditionally multivariate normal density $\mu|\Sigma \sim N(\mu_0|\tau^{-1}\Sigma)$ where the hyperparameters are fixed and known. Under these priors, the conditional distribution of $y_{i(mis)}$ given $y_{i(obs)}$ and θ is multivariate normal. The complete data posterior $P(\theta|Y_{\text{mis}}, Y_{\text{obs}})$ is a normal inverted-Wishart distribution

$$\mu|\Sigma \sim N(\mu_0, \tau^{-1}\Sigma)$$

$$\Sigma \sim W^{-1}(v, \Lambda)$$

for some $(\tau, v, \mu_0, \Lambda)$ determined by the prior, the observed data Y_{obs} and the imputed missing data $Y^{(t)}_{\text{mis}}$ (Schafer 1997, p. 181–184).

For multiple imputation problems, we have the additional requirement that the draws we use for imputations must be statistically independent, which is not a characteristic of successive draws from Markov chain methods. Some researchers reduce dependence by using every nth random draw (where n is determined by examining the autocorrelation function of each of the parameters), but Schafer (1997), following Gelman and Rubin (1992), recommends addressing both problems by creating one independent Markov chain for each of the desired m imputations, with starting values drawn randomly from an overdispersed (normal-inverted Wishart distribution) approximation distribution. The difficult with taking every rth draw from one chain is the interpretation of autocorrelation functions (requiring analysts of cross-sectional data to be familiar with time series methods); whereas the difficulty of running separate chains is that the run time is increased by a factor of m.

The M completed data sets are then analysed using complete data methods giving M sets of completed-data statistics, for example point estimates θ_k and variances $U_k, k = 1, \ldots M$. The M sets of completed-data statistics are combined to create one multiple-imputation inference as follows (Rubin 1987). The estimate of θ is θ_{bar}:

$$\theta_{\text{bar}} = \frac{1}{M}\sum_{k=1}^{M}\hat{\theta}_k$$

the average of the M completed-data estimates of θ. Also, let U_{bar} be the average of the M completed-data variances, and:

$$B = \frac{1}{M-1}\sum_{k=1}^{M}(\hat{\theta}_k - \theta_{\text{bar}})^2$$

be the between imputation variance of the completed data estimates of θ. Then the total variance of $(\theta - \theta_{\text{bar}})$ is given by the sum of the within imputation component (U_{bar}) and the between imputation component (B) multiplied by a finite M correction $(1 + M^{-1})$, that is:

$$V = \overline{U} + \left(1 + \frac{1}{M}\right)B.$$

Interval estimates and significance levels are obtained using a t-distribution with center θ_{bar}, scale $V^{-1/2}$, and degrees of freedom $v = (M-1)(1+q^{-1})^2$, where $q = (1 + M^{-1})B/U_{\text{bar}}$, is the ratio of the between-imputation component of variance to the within-imputation component. The two-sided p-value for the null hypothesis H_0: $\theta = 0$ is computed by comparing $\theta_{\text{bar}}/V^{1/2}$ with a t-distribution with v degrees of freedom.

Some variables in our imputation are clearly nonnormal (ie. discrete) but are completely observed, therefore the multivariate normal model may still be used for inference given that (a) it is plausible to model the incomplete variables as conditionally normal given a linear function of the complete ones, and (b) the parameters of inferential interest pertain only to this conditional distribution. Convergence of the Gibbs sampler was assessed by examining plots of the sample trace for draws of missing income values. A burn-in of 5000 iterations was found to be adequate to achieve convergence. Figure 2 shows the simulated income values from one household where income was not reported using a further 1000 iterations of the Gibbs sampler.

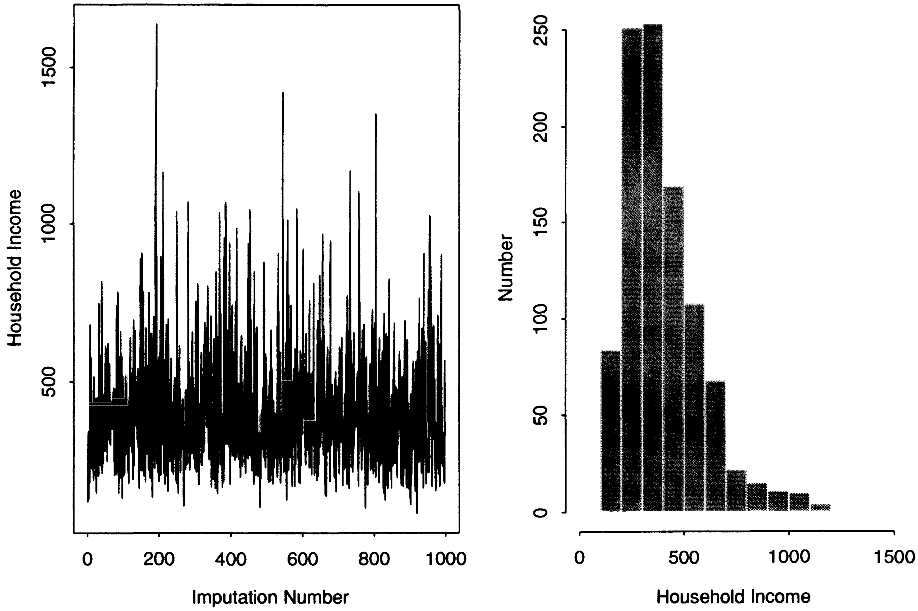

Fig. 2. 1000 simulated income values for one household

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The decisions of Spanish youth: A cross-section study

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Abstract. This paper presents a simultaneous model for the joint decisions of working, studying and leaving the parental household by young people in Spain. Using cross-section data from the 1990–1991 *Encuesta de Presupuestos Familiares*, the model is estimated by a two stage estimation method. Endogeneity of the three decisions proves to be important in order to understand the dynamics of household formation. Our results also confirm a number of plausible intuitions about the effect of individual characteristics and economic variables on these decisions, and provide some new insights into the reasons for young people in Spain remaining in large numbers in the parental home. Most of the results are gender independent.

JEL classification: J12, J21

Key words: Household formation, working and studying decisions, two stage estimation

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1. Introduction

Contrary to Anglo-Saxon and central European countries but in line with other Southern European nations, in Spain the proportion of young people

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living with their parents is very high. Del Río and Ruiz-Castillo (1997) describes the evolution of living arrangements and living standards during the 1980's in Spain. They find that the young (defined as individuals from 16 to 30 years of age) who live as dependents in the parental home, is one of the population subgroups with the highest social welfare indices in the distribution of household expenditures adjusted for household size. Jurado (1997) provides an interesting analysis of how regional variation in housing and labor market conditions, as well as education enrolment rates, are associated with regional variation in the rate of coresidence in a selected number of Spanish regions. However, we do not know of any econometric work that attempts an explanation of this important phenomenon using multivariate techniques.

A natural starting point for an economic analysis of household formation decisions by young adults is a utility comparison framework. Using a Nash bargaining model of family behavior, McElroy (1985) is the first paper to generate the indirect utility functions for this framework. The model is used to examine the joint determination of market work and family status of young men in the U.S. Among other things, the results indicate that the option to live in the parental household serves as "unemployment insurance". This provides a valuable insight into the Spanish case, where the unemployment rate among the young is one of the highest in the EU.[1]

However, it should be noted that more than 50% of all young people coresiding in the parental home[2] have a job. It is true that the vast majority of this employment is temporary and not well paid. From this perspective, parents might be providing means to compensate for job insecurity, low wages and/or high housing costs. However, it is less well known that the family as a safety net in Spain also works the other direction. Sastre (1999) shows that, under the assumption that all household members pool their economic resources, young dependents' income contributes to a decrease in household income inequality. On the other hand, Cantó and Mercader (1999) show that the employed young dependents reduce the risk of poverty for the rest of the household members, particularly when the household head is out of work. In brief, there are many reasons why the living arrangements and the labor participation decisions in Spain should be treated simultaneously.

A general model would have to take into account the interaction between parents and their descendants' decisions in a dynamic context. The more complete dynamic model of household formation is the one by Rosenzweig and Wolpin (1993) (or RW for short). The authors formulate an altruistic, imperfect-foresight, overlapping generations model, incorporating human capital investments, interhousehold transfers, and household formation. The presence of the human capital investment dimension is important in our case, because Spain has one of the largest enrolment rates in higher education among the OECD countries. From this perspective, parents are helping to finance their sons and daughters' investment in human capital by providing them with shelter, and possibly other goods and services, while coresiding, or with direct transfers when their offspring are living apart. However, the investment in school in RW is simply modeled as an all-or-nothing option. Moreover, the labor/leisure margin is not included. Accordingly, all young people are assigned some gross potential earnings (in Mincer's sense), depending on their own accumulated human capital in school or at work. In a similar spirit but in a much simpler context, Ermish (1996) (or E for short) explicitly takes into account the public good aspect of housing, and incorporates the human capital investment decision in a two period dynamic model in which the young adults face a binding borrowing constraint. Unfortunately, like RW, this model does not include the labor/leisure choice. Instead, in this work all young people are assigned some potential earnings, which varies with individual ability.[3]

Ideally, young people's decisions should be analyzed with panel data.[4] Moreover, the theories just reviewed demand a rich data set. As RW point out, their model indicates that information on the contemporary life-cycle earnings of both generations and their components, the human capital investments of the second generation, the household living arrangements, and the parental transfers are, at a minimum, required. Unfortunately, our data are far from these standards. In this paper we use the 1990–1991 *Encuesta de Presupuestos Familiares* (EPF), a large household budget survey gathered by the *Instituto Nacional de Estadística* (INE) with the main purpose of estimating the aggregate weights of the Consumer Price Index. Like other similar surveys, the main shortcoming of the 1990–1991 EPF is that it lacks data on wealth and permanent income for all households and on parental income and parental characteristics for the young people living independently. On the bright side, our survey's information on individual and household characteristics is completed with a variety of variables that reflect potentially important regional differences in housing and labor market conditions.

Combining the residence and the parental transfer decisions, both RW and E lead to a number of mutually exclusive states, which can be analyzed by means of a multinomial empirical model. We could always attempt to extend this utility maximizing framework in order to make the two additional decisions we are interested in truly endogeneous: the labor participation decision, and the decision to pursue some education. This approach has the disadvantage of imposing a particular form of endogeneity and simultaneity on the decisions involved. Moreover, the small sample size of some of the categories does not allow the identification of all the parameters of interest with our data set.

Alternatively, in this paper we use a multivariate probit model in which it is the *propensity* to select a state, rather than its occupancy, which determines the probability that a state is actually occupied.[5] In particular, we study the joint decision by young people aged 19–35 on whether to remain in the parental household, whether to work, and whether to keep on studying. Parental transfers and other non labor income are treated as exogenous variables conditioning these three simultaneous decisions. In principle, one could also attempt to endogenize the marriage decision. But we present evidence showing that, in the Spanish case, this is not necessary: the decision to marry or to leave the parental home almost always takes place simultaneously. Our probability model is estimated adapting a two-stage method proposed by Arellano and Bover (1997) in a related context.

Our results indicate that a rich pattern of interdependencies exists between the three decisions, and that both individual characteristics and economic variables play a significant explanatory role in the three propensities modeled in the paper. In particular, we establish that the following factors increase the propensity of young people to coreside with their parents: to be unemployed or inactive; to study; to have no income different from labor earnings or parental transfers; to have achieved a higher educational level; to live in a small village; to live in a region with high housing prices or little availability of rental housing, and to live in a region with a higher unemployment rate or a higher illiteracy rate. Most of these factors are common to male and female.

The rest of the paper is organized in four sections. Section 2 discusses in more detail the existing theory before presenting the empirical model. Section 3 is devoted to the data, Sect. 4 contains the results, and Sect. 5 concludes.

2. The empirical model

2.1. *The existing theoretical literature*

In order to structure the empirical analysis, our starting point is the single period framework of McElroy (1985) in which young people make decisions about living arrangements and labor force participation.[6] In another strand of the literature during the late 1980s, Cox (1987) presents a utility maximizing model containing both altruistic and exchange motives for private *inter vivos* transfers. In Cox (1990), a simple two period model in which the parent has access to capital markets but the child does not, is used to analyze whether private intergenerational transfers function as loans or as subsidies that are used to help family units to overcome liquidity constraints. However, Cox's models ignore the decision concerning the coresidence of family members. But to the extent that households contain public goods that can be jointly consumed, residence sharing is cheaper for the parent than providing an almost equivalent service without coresidence.

These different strands of the literature converge in the type of full-blown dynamic theory of household formation due to RW. Combining the residence and transfer decisions, this model leads to three mutually exclusive states: (i) living apart-receiving parental transfers; (ii) coresiding, and (iii), living apart-not receiving parental transfers.[7] As pointed out in the Introduction, we could always extend this framework to treat the decisions about labor participation and human capital investment endogenously. Consider, for example, a simple two period dynamic model in which all households consist of a one-child family. In this case, the three mutually exclusive states in RW and E become 12, depending on whether the child studies or not and, in addition, on whether s/he works or not in the first period. Even if we consider that parental transfers are predetermined variables for the young offspring, as we do in our empirical analysis, the number of states would still be 8.

What in RW, or in E is a reasonable specification, becomes an intractable one with our data when, together with the residence decision, we want to cover the labor participation and the human capital decisions as well. Moreover, the multinomial logit framework precludes us from asking whether the propensity (and/or the fact) that an adult child is employed has some explanatory role in the probability that this person lives independently. Analogously, this way of setting up the estimation problem forces us to ignore the possibility that the propensity (and/or the decision) to coreside with one's parents, enjoying the corresponding material advantages, might influence the probability that one finds a job in the labor market.

Our conclusion from the above discussion is that the multinomial logit empirical model, which can be rationalized quite directly with the help of the existing optimization theory, has certain disadvantages.[8] Alternatively, we believe that it is fruitful to take as our starting point the idea that household formation by young people is intrinsically linked to other decisions with regard to job or education status. Given this idea, the existing literature will be systematically used in the selection of explanatory variables and the interpretation of our empirical results.

2.2. The empirical model

We assume that the young decide simultaneously whether to remain in the parental household, whether to work, and whether to keep on studying. Empirically, there are different ways of expressing the interrelation among the three decisions. Our basic assumption is that it is the propensity to select a state, rather than its occupancy, which determines the probability that a state is actually occupied. In order to analyze the mutual influences among the three propensities, we consider the following simultaneous equations model in which each of the three propensities is defined in terms of the other propensities and a set of exogenous variables:

$$I_{1i}^* = X_{1i}\beta_1 + \delta_{12}I_{2i}^* + \delta_{13}I_{3i}^* + u_{1i} \tag{1}$$

$$I_{2i}^* = X_{2i}\beta_2 + \delta_{21}I_{1i}^* + \delta_{23}I_{3i}^* + u_{2i} \tag{2}$$

$$I_{3i}^* = X_{3i}\beta_3 + \delta_{31}I_{1i}^* + \delta_{32}I_{2i}^* + u_{3i}. \tag{3}$$

The variable I_1^* is the underlying individual propensity to leave the parental house, I_2^* represents the propensity to work, and I_3^* is the propensity to study. X_1, X_2 and X_3 are sets of exogenous demographic and economic variables that condition each equation. The β and δ vectors are the parameters of interest, and the error terms u_1, u_2 and u_3 are assumed to be jointly normally distributed. This approach has the empirical attractiveness of allowing us to test the endogeneity and simultaneity of the three decisions.[9]

In the data set we observe the outcomes of the choices, not the underlying propensities. That is, we observe whether an individual is independent from his/her parents, whether s/he is working and whether s/he is studying. The connection between our observations and the corresponding latent variables is given by the following three dichotomous variables:

$$\begin{aligned} I_{1i} &= 1 \quad \textit{if } I_{1i}^* > 0 \\ I_{1i} &= 0 \quad \textit{otherwise} \end{aligned} \tag{4}$$

$$\begin{aligned} I_{2i} &= 1 \quad \textit{if } I_{2i}^* > 0 \\ I_{2i} &= 0 \quad \textit{otherwise} \end{aligned} \tag{5}$$

$$\begin{aligned} I_{3i} &= 1 \quad \textit{if } I_{3i}^* > 0 \\ I_{3i} &= 0 \quad \textit{otherwise} \end{aligned} \tag{6}$$

We are interested in the estimation of the set of parameters $\Theta = \{\beta_1, \beta_2, \beta_3,$

$\delta_{12},\delta_{13},\delta_{21},\delta_{23},\delta_{31},\delta_{32}\}$ from the simultaneous probability model consisting of Eqs. (1) to (3) and the observability conditions (4) to (6). Given the interdependence among the unobserved latent variables, we face a trivariate probit. Although estimation by numerical methods in a Simulated Maximum Likelihood routine could be used to achieve full efficiency, two-stage methods provide consistent estimates of the parameters of interest and they are easy to implement in spite of some efficiency loss.[10] As discussed below, the method we use minimizes this efficiency loss.

Arellano and Bover (1997) propose a two-stage estimator for limited dependent variable models from panel data. In the first stage, reduced form equations for the endogenous variables are derived and estimated as independent probit equations;[11] in the second stage, the reduced form linear predictions replace all the unobservable latent variables.

This methodology can be readily extended to our case in which a simultaneous probability model must be estimated using a cross-section. This approach has two advantages over other methods that also estimate a probit on the second stage. First, given the assumption of unitary variance of the disturbances on the reduced form equations in these methods, the parameters of interest can only be recovered up to scale. However, because OLS in the second stage does not impose such identification conditions, the Arellano and Bover strategy allows us to recover the actual parameters without the scale restriction. Moreover, in an additional step, the Arellano and Bover method, allows for (i) the computation of a linear GMM estimator that is asymptotically efficient (relative to the first stage estimation), and (ii) the construction of a specification test for the overidentifying restrictions.

More explicitly, in the first stage, we consider the reduced form equations for the three endogenous variables,

$$I_{1i}^{*} = \pi_1 X + v_{1i} \tag{7}$$

$$I_{2i}^{*} = \pi_2 X + v_{2i} \tag{8}$$

$$I_{3i}^{*} = \pi_3 X + v_{3i}, \tag{9}$$

where X includes all variables in X_1, X_2 and X_3. The error terms, v_{1i}, v_{2i} and v_{3i}, are assumed to be jointly normally distributed with variance equal to 1. The parameters in Eqs. (7) to (9) are estimated by separate probit maximum likelihood, and the predictions for the unobserved latent variables, $\hat{I}_{1i}^{*} = \hat{\pi}_1' X_i$, $\hat{I}_{2i}^{*} = \hat{\pi}_2' X_i$ *and* $\hat{I}_{3i}^{*} = \hat{\pi}_3' X_i$, are then computed.

In the second stage, we use these predictions to replace both types of unobservable latent variables: the endogenous explanatory variables and the dependent ones. Then, the parameters can be consistently recovered by applying OLS to the following equations:

$$\hat{I}_{1i}^{*} = X_{1i}\beta_1 + \delta_{12}\hat{I}_{2i}^{*} + \delta_{13}\hat{I}_{3i}^{*} + \varepsilon_{1i} \tag{10}$$

$$\hat{I}_{2i}^{*} = X_{2i}\beta_2 + \delta_{21}\hat{I}_{1i}^{*} + \delta_{23}\hat{I}_{3i}^{*} + \varepsilon_{2i} \tag{11}$$

$$\hat{I}_{3i}^{*} = X_{3i}\beta_3 + \delta_{31}\hat{I}_{1i}^{*} + \delta_{32}\hat{I}_{2i}^{*} + \varepsilon_{3i}. \tag{12}$$

Given the consistency and normality of the reduced form parameters, the set of estimates $\hat{\Theta}$ is also consistent and asymptotically normal. However, since the dependent and the endogenous explanatory variables have been replaced by their predicted values, the asymptotic variance matrix of the estimates is *not* the traditional one for OLS estimators.

This second stage OLS estimator $\hat{\Theta}$ can be interpreted as a GMM estimator in which the weighting matrix has not been chosen optimally. An efficient estimation relative to the first stage estimates, $\hat{\Pi} = \begin{bmatrix} \hat{\Pi}_1' \\ \hat{\Pi}_2' \\ \hat{\Pi}_3' \end{bmatrix}$, can be obtained in a third stage by choosing optimally the weighting matrix as a consistent estimate of the inverse of the covariance matrix of the orthogonality conditions. A discussion of the consistency and normality of the estimates, the variance matrix and the election of the optimal weighting matrix can be found in Appendix A.[12]

This procedure enables us to address the principal technical issue of this study, that is, to estimate the coefficients of the endogenous variables as a means of inferring the interdependence among the three decisions considered. In that sense, we will refer to the model as "structural", as opposed to the reduced form equations that are estimated in the first stage.

3. Data

In Tables 1 to 3, we present some basic statistics illustrating the differences between some Northern and Southern European countries (including Spain), plus the United States, in the three dimensions we are concerned with.[13] Table 1 refers to the differences in living arrangements. We observe that in all countries a significant proportion of females leave the parental home before males of the same age. However, while in the U.S. or Germany, for instance, approximately 90% of the young live on their own at the age of 29, in the three Southern countries selected about 50% of the males and 25–35% of the females at that age are still living as dependants in their parents' house. In all countries,

Table 1. Proportion of men and women still living with parents by age group in six European countries – in 1986 – and the United States, in 1987 (in % of age group total)

	Men			Women		
	15–19	20–24	25–29	15–19	20–24	25–29
Spain	95.6	88.1	53.2	93.9	76.1	35.3
Italy	97.4	87.8	49.6	95.7	70.4	25.5
Greece	94.6	76.5	53.8	89.2	52.3	23.8
France	94.8	56.9	19.3	89.8	36.4	8.4
United Kingdom	93.6	57.2	21.9	87.8	33.8	8.6
Germany	94.8	64.8	27.4	92.0	42.8	11.0
United States		Up to 24: 27.4	Up to 29: 13.0		Up to 24: 22.2	Up to 29: 8.6

Source: For the six European countries, Tables 1 and 2 in Fernández Cordón (1997); for the U.S., Table 1 in Haurin et al. (1993).

Table 2. Unemployment rates by age groups in 1989

	14–24	25–29	30–34	Total rate
Spain	34.3	23.4	14.7	17.3
Italy	31.9	16.9	9.4	11.1
Greece	24.8	7.0	4.3	7.5
France	19.6	11.3	8.4	9.6
United Kingdom	10.3	8.9	7.1	7.4
Germany	5.5	6.2	6.4	5.7
United States	10.9(*)			5.3

Source: For the six European countries, Labour Force Survey; for the U.S., Current Population Survey. (*) U.S. data is for individuals from 16 to 24 years of age.

Table 3. Net enrolment in full-time public and private university education, by age group in 1991. Selected countries

	18–21	22–25	26–29
Spain	21.3	14.2	5.3
France	18.5	10.6	3.7
United Kingdom	12.4	3.0	0.9
Germany	6.8	14.7	9.3
United States	22.8	8.5	2.5

Source: Education at a glance, OECD, 1993. Data for Italy and Greece, not available.

the unemployment rate is higher for young people up to 29 years of age than for the population as a whole (Table 2), but this effect is larger in the Southern countries. Finally, Table 3 shows that as far as the net enrolment rate in university education is concerned, in 1991 Spain compares very favorably with the Northern countries, all of which are at a considerably higher level of economic development.

As we said in the Introduction, the data used in this paper comes from the 1990–1991 EPF. This is a household budget survey collected during 52 consecutive weeks, from April of 1990 to March of 1991, with the main purpose of estimating the weights of the Consumer Price Index. It is a representative sample consisting of 21,155 observations for a population of approximately 11 million households living in residential housing throughout Spain. There are 72,123 individuals in the sample, representative of a population of 38.5 million people.

A household is defined as "the person or set of persons who jointly occupy a residential family dwelling, or part of it, and consume or share food and other commodities under a common budget." Therefore, people living in collective housing – residences for College students or the elderly, hospitals, hotels, prisons and the like – are not directly interviewed. However, expenditures and characteristics of household members who are entirely dependant on household resources but who live elsewhere at the time of the interview, are recorded in our data – for more details on the 1990–1991 EPF, see INE (1992).

In view of Table 1, we choose 35 years of age as the upper bound in our definition of the young.[14] On the other hand, since we are interested in the education decision, we choose 18 years of age as the lower bound, the earliest age at which people in Spain are supposed to decide whether to continue their studies beyond secondary education. This gives us a sample of 9,741 males and 9,534 females. Ceuta and Melilla residents are not considered since some regional variables are not available for them.

All the endogenous and exogenous variables are described in the Data Appendix B. As far as the endogenous variables, we observe in Table B.1 that the proportion of young females living on their own is 44% versus only 33% among the males, while the proportion of females studying is almost 5 percentage points above the males. On the other hand, the female employment rate is only 37% as opposed to 63% for the males; however, although not shown here, the female unemployment rate is 28%, which is around ten percentage points higher than the male rate. Table B.2 shows the sample distribution according to the three dependent variables. With reference only to the minority groups, around 5% of males and females work and follow some type of studies at the same time, while only 2% of females and 0.5% of males study and are independent.

The exogenous variables entering Eqs. (1) to (3) are of two types: individual characteristics, and economic variables. We have information on the following individual characteristics: education, age, whether residing in a large city or in a small village, parental transfers, and other non labor income. We should point out that, in addition to coresidence, there are two ways parents can help their offspring: by a cash transfer, and by financing all or part of the housing services consumed by their descendants when they live independently. In fact, more than 80% of those independent males and females receiving any kind of transfer live in dwellings subsidized by some family member. Both types of parental transfers are assumed to be optimally selected by the parents. Hence, they are a predetermined variable for the young individuals. Only 2.19% of young people coresiding with their parents receive a cash transfer. Therefore, we are forced to add up both types of transfers, forfeiting the possibility of identifying the role of parental cash transfers. The final individual characteristic is the remaining non labor income. Unfortunately, it has been impossible to identify the effect of both variables separately. Therefore, in the final specification non labor income is the sum of parental transfers and other non labor income. In principle, both the parental transfers and the educational attainment could be considered endogenous variables. However, their exogeneity, as well as the exogeneity of the rest of explanatory variables is tested making use of the Sargan test of overidentifying restrictions proposed in Appendix A.

Our data source lacks information on three potentially important individual characteristics: the type of contract – temporary or indefinite – of those young people holding a job, the marital status and the fertility behavior of anyone different from the household head (defined as the household member with the highest earnings) and his/her spouse. The majority of the employed young people in Spain has a temporary job.[15] Our lack of data in this respect precludes a study of the interaction between the decision to leave one's parents house made by those who are employed and the type of contract they have. However, it should be noted that even having the data, we could not simply include the type of contract as an exogenous variable in the present framework. The recognition of its endogenous nature would possibly call for an independent analysis among the employed.

On the other hand, there is some empirical evidence that marriage is an

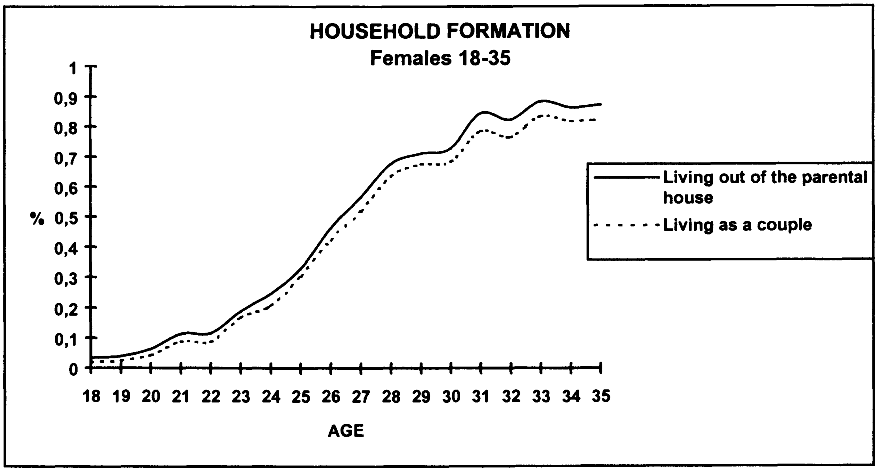

Fig. 1. Proportion of females living out of the parental house, by marital status

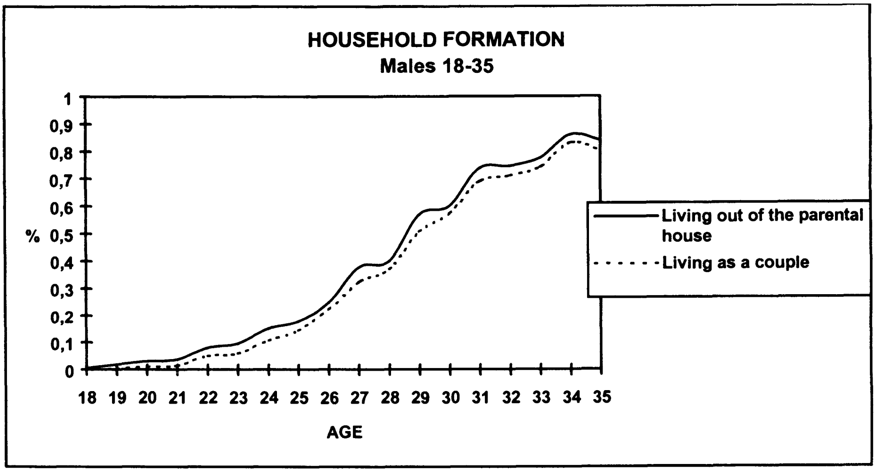

Fig. 2. Proportion of males living out of the parental house, by marital status

important explanatory variable of household formation – see, for example, Haurin et al. (1993). However, Figs. 1 and 2 indicate that in Spain almost all individuals that live outside their parents household are married, which implies that the decision of marrying or leaving the parents' home almost always takes place simultaneously.[16] Therefore, the only implication of not including marital status as an explanatory variable is that our results would refer, not only to the propensity to leave the parental house, but also to the propensity for getting married.

Finally the absence of fertility variables for those living with their parents could potentially be a shortcoming for interpreting the results, especially in the case of women's decisions. As in the case of the type of contract, fertility decisions are likely to be endogenous and influenced by the same exogenous variables we use in our analysis.

We have two sets of economic variables, both of which are supposed to allow us to identify the equations of interest. In the first place, we have selected three variables which refer to housing conditions and capture the spatial variability of: (i) rental-equivalent values of the housing services provided by the whole stock, (ii) housing prices in the owner-occupied sector, and (iii) the relative importance of the rental housing sector, which is the one typically frequented by young people. We assume that these variables only affect directly the propensity for leaving the parental house. In the second place, we have considered the following three variables which capture the spatial variability in labor market conditions: (i) the unemployment rate for the population as a whole, (ii) the unemployment rate disaggregated by sex and age, and (iii) the illiteracy rate. We assume that these variables only affect the propensity for leaving the parental house through their effect on the propensities for working or studying. Descriptive statistics for the individual characteristics and economic variables can be found in Table B.1.

4. Results

4.1. Simultaneity and identification

As pointed out in Section 2.1, the existing theory suggests the interdependence among the decisions of working, leaving the parental house and studying. But one of the advantages of our approach is that we can evaluate whether a simultaneous model of the three decisions is called for. We do this by testing for pairwise independence of the equations using bivariate probits.[17] Table B.3 in Appendix B reflects that for both males and females, we always find a significant correlation between the error terms of any two equations. Therefore, we conclude that the simultaneous model proposed in Section 2 has a sound empirical base.

We estimate separately each of the structural equations for males and females to reflect differences by gender. The reduced form estimates are presented in Tables B.3 and B.4 in Appendix B, but they are not discussed here because they do not provide any additional insight on the topic. Table 4 shows the optimal GMM estimates for our simultaneous equation system. In this step all endogenous variables have been replaced by their linear predicted values (see Appendix A). As previously stated, the economic variables allow us to identify the parameters of interest. In particular, the two variables on housing values and their interactions with age, as well as the variable reflecting the relative size of the rental housing market, only directly affect the propensity for living independently. Conversely, the regional unemployment, the provincial unemployment by age and sex, and the regional illiteracy rate only influence the decision of leaving the parental household through the effect of working and studying. These exclusion restrictions are not rejected by the Sargan test defined in Appendix A, whose value appears in the bottom line of Table 4. Moreover, the Sargan test does not allow us to reject the exogeneity of any of the variables since it implies that the orthogonality conditions imposed on the estimation are jointly not significantly different from zero.

Table 4. GMM optimal estimates

	Males		Females	
	Coeff	t-ratio	Coeff	t-ratio
Work equation				
Independence (P)	0.383	1.491	0.395	1.772
Studying (P)	−1.464	−6.443	−1.454	−8.598
Age	−1.479	−2.213	−2.576	−4.764
Age^2	0.426	2.128	1.087	4.836
Primary School	0.276	1.236	0.756	3.192
Secondary School	1.835	5.393	2.714	7.797
Higher Education	2.390	5.839	3.490	9.150
City	0.267	2.965	0.390	5.048
Village	−0.189	−1.657	−0.062	−0.605
Non labor Income	−0.039	−1.685	−0.064	−1.754
General UR	−2.524	−2.869	−1.739	−2.087
UR by sex and age group	0.132	0.321	0.435	0.883
Illiteracy rate	−4.544	−2.260	−3.379	−2.242
Intercept	−1.468	−2.806	−3.243	−7.412
Independence equation				
Working (P)	0.462	2.558	0.507	2.564
Studying (P)	−1.589	−3.084	−0.998	−2.453
Age	−0.492	−0.588	0.451	0.764
Age^2	1.203	2.669	0.545	1.444
Primary School	0.002	0.006	0.363	1.432
Secondary School	2.119	3.183	1.163	1.656
Higher Education	2.596	3.030	0.818	1.016
City	0.637	3.302	0.322	2.609
Village	−0.427	−3.590	−0.268	−3.153
Non labor Income	0.141	3.769	0.108	1.775
Owning Costs	−0.385	−2.134	−0.406	−3.312
Age x Owning Costs	0.355	1.190	0.903	3.655
Rental Values	−0.093	−0.341	0.004	0.015
Age x Rental Values	−1.218	−1.861	−2.058	−3.425
Rental Accommodation	2.016	2.999	1.056	1.620
Intercept	−4.133	−4.114	−1.819	−1.841
Studying equation				
Working (P)	−0.686	−7.486	−0.679	−11.039
Independence (P)	0.259	1.444	0.274	1.876
Age	−1.012	−2.494	−1.783	−5.854
Age^2	0.296	2.926	0.754	7.017
Primary School	0.189	1.160	0.514	3.173
Secondary School	1.253	8.287	1.861	11.463
Higher Education	1.631	10.897	2.391	12.955
City	0.181	4.235	0.268	6.731
Village	−0.130	−1.690	−0.042	−0.615
Non labor Income	−0.027	−1.661	−0.044	−1.977
General UR	−1.640	−2.522	−1.109	−2.050
UR by sex and age group	0.018	0.073	0.296	0.958
Illiteracy Rate	−3.172	−2.367	−2.508	−2.669
Intercept	−1.003	−3.876	−2.227	−12.941
Number Observations	9741		9535	
Sargan Test χ^2 (7)				
(Overidentifying restrictions)	5.839		6.892	
(p-value)	(0.559)		(0.440)	

(P): Predicted value; Age = (age-25)/10; UR: Unemployment Rate;

4.2. Interdependence

As expected, working increases the propensity to leave on one's own since individuals who work have access to the necessary funds that allow them to be independent from their family. By the same token, not working – i.e., being unemployed or inactive – increases the propensity to coreside. This is the important "unemployment insurance" effect discovered by McElroy (1985) for the 19–24, never married, out of school, U.S. males in 1971. On the other hand, living independently has a positive effect, although only marginally significant, on the propensity to work; that is to say, the increased costs that the individuals face when living on their own act as an incentive to work. Equivalently, it appears that coresiding is weakly associated with a greater propensity for being unemployed or out of the labor force.

Studying strongly increases the propensity to coreside.[18] This result reflects the conditions of the publicly dominated Spanish university system: as in Greece, Italy, France or Portugal, tuition costs in public Universities are nominal – less than 500 Ecus per year; moreover, the spatial dissemination of College centers in Spain reduces the costs of studying through coresidence. The opposite result in Ermish (1996) reflects the British University system, which strongly encourages people to study in a different locality from their parents.[19] On the other hand, although not very significant, the positive effect of living apart on the propensity to study is the only puzzling result in this subsection. We have to bear in mind that we are considering every type of course not only official studies. In particular, among the women living independently, 47% report following "other studies" different from primary, secondary, or College studies.

Unsurprisingly, the propensity to work strongly reduces the propensity to study. Conversely, to follow any type of studies significantly reduces the propensity to work.

It is important to emphasize that the nature of the interaction between the three pairs of decisions is similar for both men and women.

4.3. Individual characteristics

Age, although significant in most equations, does not have a clear interpretation since it enters the equations through many channels (age, age squared, the group specific unemployment rate, and interactions with housing prices).

The role of the education variables is important. In the first place, the higher the level of education attained by the individual, the higher the propensity to work – a plausible result. In the second place, given that the level of education can be considered a predetermined variable in the equation for studying, it is also plausible that the higher the level of education attained the higher the propensity to continue studying. In the third place, the effect of education on the propensity to leave the parental house has a differentiated effect for males and females. For males, education has a positive independent effect on the probability of forming a household, which reflects the fact that the more educated individuals are also those with higher earnings and, presumably, the more attractive partners in the marriage market. There is also a strong indirect effect through the increase on the propensity of working, offset by the increase in the propensity to study. For women, the indirect effects work in the same direction and are even stronger, but education does not have an independent effect on its own: only if the woman has finished secondary school is she more likely to be living on her own. Since most individuals living apart from their parents are married, this decreasing effect of education may merely reflect the postponement of marriage decisions by more educated women.

Parental transfers and other non labor income have a positive effect on the probability of leaving the parental home: access to more economic resources helps individuals to afford the expenses of living on their own and, therefore, increases the probability of leaving the parental house. On the other hand, non labor income has the usual negative effect on the propensity to work (and to study). This effect is stronger for females, possibly because a larger non labor income allows them to spend their time in other activities such as taking care of the family or having children.

Living in a large city increases the propensity to form a household, while living in a small village decreases it. This effect is especially strong for males. It probably reflects the fact that the traditional pattern of "extended families", where several generations cohabit under the same roof and cooperate in the same productive activities, is more prevalent in rural areas. On the other hand, the effect of the municipal size on the propensity to work and to study reveals that there are more jobs and opportunities to study as the municipal size increases. However, the effect pattern differs by gender. For males, relative to medium sizes municipalities, there is a negative effect in small villages, and a positive effect in big cities. For females, there is a clear discontinuity: there is only a significant strong positive effect in big cities.

4.4. Economic variables

All the variables reflecting the spatial variation in housing conditions have the expected influence on the propensity to live apart. In the first place, the higher the rental equivalence value of the whole stock and, above all, the higher the owner-occupied housing prices are, the lower the propensity to live independently. The difference between these two variables is that the effect of the second one decreases with age. In the second place, the variable reflecting the availability of rental housing has a strong and significant positive effect on the probability of living independently. These results are in agreement with those obtained in Börsch-Supan (1985), Haurin et al. (1993, 1994), Ermish and Di Salvo (1997) and Ermish (1999), emphasizing the role of housing costs – and not only own income – as a basic determinant of the household formation and related demographic decisions.

The gender and age group specific unemployment rate does not have any significant effect on the propensity to work: it appears that individuals perceive the general unemployment rate as the relevant variable. Of course, the general unemployment rate has a significant negative effect on the propensity to work, revealing the lower probability of receiving a job offer, as well as the discouraging worker effect that reduces the effort of looking for a job. Note that the indirect effect of unemployment on coresidence by lowering the propensity to work reinforces the role of the family in Spain as a cushion in the face of unfavorable labor market conditions.[20] On the other hand, given job status, the unemployment rate has a negative effect on the probability of studying for both men and women. Therefore, unemployment works in two ways: it reduces the probability of finding a job, favoring the studying option (opportunity cost effect), but it *directly* reduces the probability of studying (discouraging effect due to the poorer job perspectives).

The illiteracy rate has a strong negative effect on the propensity to study and to work. The first effect is consistent with different explanations. It can reflect a peer effect, so that young individuals in areas where not many people have studied in the past also tend not to study. It could also be reflecting how regions with a more educated stock of human capital have developed a wider net of possibilities for the young people to continue their studies after the compulsory age. Finally, the negative effect on the propensity to work can reflect the fact that areas with a more educated workforce may have better job opportunities through higher investment (and job creation) by the firms settled there – for a theoretical exposition of this idea, see Acemoglu (1996).

5. Conclusions

As far as we know, this paper constitutes the first econometric attempt to explain why parents and young descendants between 18 and 35 years of age decide to live together in Spain in rather large proportions. Lacking longitudinal data, we have found it interesting to work with a sufficiently rich, large and readily available household budget survey – the 1990–1991 EPF – collected by the Spanish INE with completely different aims in mind. The reason is that, even with cross-section data one can start addressing the issues involved in joint decision making. The major novelty in the paper is that, in addition to the joint decision of whether to remain in the parental house and whether to work, we have been able to add the decision of whether to continue studying.

The analysis is implemented through a two-stage method developed by Arellano and Bover (1997) for limited dependent variable models from panel data, which has been adapted here to the case of simultaneous probability models using cross-section data. Our results indicate that the behavior of Spanish youth is amenable to careful empirical analysis with standard tools.

From a methodological point of view, we have shown that the endogenous and simultaneous treatment of the three decisions should occupy the core of any attempt to understand the issues involved. Otherwise, seriously misleading results could be obtained, a possibility we have illustrated when the propensities to work or to study are treated as mere exogenous variables influencing the decision to abandon the parental household.

As pointed out in the Introduction, we have confirmed that there exists a rich pattern of interdependencies between the three decisions, and that both individual characteristics and economic variables have a significant explanatory role in the three propensities modeled in the paper. More specifically, our results show that the pattern of interdependencies among the three decisions, as well as the effect of the economic variables, are qualitatively the same for males and females. In particular, we have found support for the following regularities: (i) parents help their young offspring through coresidence when the latter do not have a job or are studying; (ii) living independently has a positive effect on the propensity to work, (iii) housing conditions significantly affect the living arrangements of the young in a direct way, while (iv) unemployment exerts its influence indirectly through its negative effect on the propensities to work and to study.

However, there are also subtle differences in gender behavior. (i) More educated women tend to postpone longer than men the decision to marry and form a new household; (ii) increases in non labor income have a stronger negative effect on the propensities to work and to study for women, and (iii) the decrease in the propensities to work, to study and to form a new household in a small village is stronger for men, while the increase in the first two propensities in a large city is greater for women.

This paper has several obvious shortcomings. In the first place, as we have pointed out from the beginning, even in a static framework the data we have used is rather incomplete. It lacks information on potentially important individual characteristics of young people, as well as on a host of fundamental variables reflecting the family background of the individuals living apart from their parents. More importantly, perhaps, is that all of our cross-section results are rather suspect because they reflect the combined impacts of both inflows and outflows from a certain work, study or living arrangement situation. In this respect, the present study can only be taken as a first step towards an understanding of the decisions of the Spanish youth that we have focused on here.

In the second place, we have been unable to provide any explanation for the role of the Spanish family, mentioned in the Introduction, as a safety net in the opposite direction of the one usually stressed. We refer to the evidence indicating that young people with a job in poorer households are making a decisive contribution in raising the standard of living of the remaining household members. To study this question, one would certainly need data on such "reverse transfers" – as well as data on the parental socioeconomic characteristics of all the young people in the population. Furthermore, to rationalize these transfers in a theoretical model, one could follow one of two routes: to postulate strong altruistic motives on the part of the young offspring; or to extend the existing exchange models substituting transfers from the young for parental loans. As a matter of fact, this line of thought might lead to a better understanding of the fact revealed by Rosenzweig and Wolpin (1993) and Ermish (1996), that, contrary to existing theory, higher parental income increases the probability of their offspring living independently.

In the third place, only an appropriate comparative study including non Southern European countries, would allow us to estimate the effect that institutions and public policies may have on the behavior of the young. Meanwhile, we have already mentioned that public policy in Spain favors high enrolment rates in public universities. To this we may add that, as pointed out in Cantó and Mercader (1999), social protection in Spain has developed during the 1980s, maintaining the pension and the unemployment subsidy programs at the center of the system. Thus, public protection for unemployed youth is either not available at all in the case of the "first-job-seekers" (35 and 28% of the unemployed males and females in our sample, respectively), or very limited indeed for "early-age-unemployed" holding mere temporary jobs. Moreover, general family support systems and child-care policies in particular are underdeveloped in the Spanish welfare state. Finally, housing policies tend to favor owner-occupied housing, which is the most inaccessible tenure choice for the young. One may assume, of course, that all these policies tend to reinforce the importance of coresidence in Spain.

Appendix A: The Arellano and Bover (1997) estimator

Consider Eqs. (1) to (3) from Section 2:

$$I_{1i}^* = X_{1i}'\beta_1 + \delta_{12}I_{2i}^* + \delta_{13}I_{3i}^* + u_{1i} \tag{1}$$

$$I_{2i}^* = X_{2i}'\beta_2 + \delta_{21}I_{1i}^* + \delta_{23}I_{3i}^* + u_{2i} \tag{2}$$

$$I_{3i}^* = X_{3i}'\beta_3 + \delta_{31}I_{1i}^* + \delta_{32}I_{2i}^* + u_{3i}. \tag{3}$$

They can be rewritten as

$$\begin{bmatrix} I_{1i}^* \\ I_{2i}^* \\ I_{3i}^* \end{bmatrix} = \begin{bmatrix} W_{1i} & O & O \\ O & W_{2i} & O \\ O & O & W_{3i} \end{bmatrix} \times \begin{bmatrix} \delta_1 \\ \delta_2 \\ \delta_3 \end{bmatrix} + \begin{bmatrix} u_{1i} \\ u_{2i} \\ u_{3i} \end{bmatrix}, \tag{4}$$

where $W_{1i} = (X_{1i}'\ I_{2i}^*\ I_{3i}^*)$ is a $1 \times (k_1 + 2)$ vector, $W_{2i} = (X_{2i}'\ I_{1i}^*\ I_{3i}^*)$ is a $1 \times (k_2 + 2)$ vector and $W_{3i} = (X_{3i}'\ I_{1i}^*\ I_{2i}^*)$ is a $1 \times (k_3 + 2)$; O are vectors of zeros that conform the W_{ji}. More compactly, (4) can be written as

$$I_i^* = \mathbf{W}_i\delta + u_i. \tag{5}$$

Let $\mathbf{X}_i = (I_3 \otimes X_i)$, where I_3 is a (3×3) identity matrix, and X is the set of all different exogenous variables in X_{1i}, X_{2i} and X_3. Since the error term in expression (5) is uncorrelated with the exogenous variables, we can write

$$E\{\mathbf{X}_i'(I_i^* - \mathbf{W}_i\delta)\} = 0. \tag{6}$$

Using the law of iterated expectations,

$$E\{\mathbf{X}_i'[E(I_i^*|\mathbf{X}_i) - E(\mathbf{W}_i|\mathbf{X}_i)\delta]\} = 0, \tag{7}$$

where

$$E(\mathbf{W}_i|\mathbf{X}_i) = \begin{bmatrix} X_{1i}' & \pi_2 X_i & \pi_3 X_i & O & 0 & 0 & O & 0 & 0 \\ O & 0 & 0 & X_{2i}' & \pi_1 X_i & \pi_3 X_i & O & 0 & 0 \\ O & 0 & 0 & O & 0 & 0 & X_{3i}' & \pi_1 X_i & \pi_2 X_i \end{bmatrix},$$

is a $(3 \times (k_1 + k_2 + k_3 + 6))$ matrix, and $E(I_i^*|\mathbf{X}_i) = \begin{bmatrix} \pi_1' X_i \\ \pi_2' X_i \\ \pi_3' X_i \end{bmatrix} = \Pi X_i.$

This suggests to consider GMM estimators of δ based on the following sample orthogonality conditions:

$$b_N(\delta) = 1/N \sum_{i=1}^{N} \mathbf{X}_i'(\hat{I}_i - \hat{\mathbf{W}}_i\delta), \tag{8}$$

where I_{1i}^*, I_{2i}^* and I_{3i}^* are replaced by their linear predictions from the first stage independent probit estimates. A GMM estimator of δ based on (8) takes the form

$$\tilde{\delta}_A = \left[\left(\sum_i \hat{\mathbf{W}}_i'\mathbf{X}_i\right) A_N \left(\sum_i \mathbf{X}_i'\hat{\mathbf{W}}_i\right)\right]^{-1} \left[\left(\sum_i \hat{\mathbf{W}}_i'\mathbf{X}_i\right) A_N \left(\sum_i \mathbf{X}_t'\hat{I}_i\right)\right], \tag{9}$$

where A_N is a weighting matrix. When $A_N = (\sum_i \mathbf{X}_i'\mathbf{X}_i)^1$, $\tilde{\delta}_A$ coincides with an OLS estimator, $\hat{\delta}$, applied to Eq. (5), where the endogenous variables have been replaced by their linear predictions from the independent probit estimates, that is,

$$\hat{\delta} = \left(\sum_i \hat{\mathbf{W}}_i'\hat{\mathbf{W}}_i\right)^{-1} \left(\sum_i \hat{\mathbf{W}}_i'\hat{I}_i\right). \tag{10}$$

Before discussing the consistency and asymptotic normality of $\tilde{\delta}_A$, we need an expression for the asymptotic variance of $b_N(\delta)$. Notice that (8) can be written as

$$\begin{aligned} b_N(\delta) &= 1/N \sum_{i=1}^{N} \mathbf{X}_i'(\hat{\Pi} X_i - \hat{\mathbf{W}}_i'\delta) \\ &= 1/N \sum_{i=1}^{N} \mathbf{X}_i'(\Gamma\hat{\Pi} X_i - X_i^*\beta)' \end{aligned} \tag{11}$$

with $\Gamma = \begin{bmatrix} 1 & -\delta_{12} & -\delta_{13} \\ -\delta_{21} & 1 & -\delta_{23} \\ -\delta_{31} & -\delta_{32} & 1 \end{bmatrix}$, $X_i^* = \begin{bmatrix} X_{1i}' \\ X_{2i}' \\ X_{3i}' \end{bmatrix}$, and $\beta = \begin{bmatrix} \beta_1 \\ \beta_2 \\ \beta_3 \end{bmatrix}$. Equations (1) to (3) can be conveniently rewritten as

$$\Gamma I_i^* = X_i^*\beta + u_i. \tag{12}$$

A comparison of expression (12) with the reduced form $I_i^* = \Pi X_i + \varepsilon_i$, premultiplied by Γ, proves that the following restriction holds:

$$X_i^*\beta = \Gamma \Pi X_i. \tag{13}$$

Using this restriction in expression (11), we obtain

$$\begin{aligned} b_N(\delta) &= 1/N \sum_{i=1}^{N} (I_3 \otimes X_i)(\Gamma\hat{\Pi} X_i - \Gamma \Pi X_i) \\ &= 1/N \left(I_3 \otimes \sum_{i=1}^{N} X_i\right)(\Gamma \otimes I_m)\,\mathrm{vec}(\hat{\Pi} - \Pi), \end{aligned} \tag{14}$$

[1] The numerical equivalence follows from the fact that the columns in $\hat{\mathbf{W}}_i$ are linear combinations of those in $\mathbf{X}_i$

and consequently

$$b_N(\delta) \sim_a N(0, E(\mathbf{X}_i'\mathbf{X}_i)V^*E(\mathbf{X}_i'\mathbf{X}_i)), \tag{15}$$

with $V^* = (\Gamma \otimes I_m)\,\mathrm{var}(\mathrm{vec}(\hat{\Pi}))(\Gamma' \otimes I_m)$. Hence, it can be proven that

$$\sqrt{N}(\tilde{\delta}_A - \delta) \sim_a N(0, (M_{\mathbf{XW}}'A_N M_{\mathbf{XW}})^{-1} M_{\mathbf{XW}}'A_N V_b A_N M_{\mathbf{XW}} (M_{\mathbf{XW}}'A_N M_{\mathbf{XW}})^{-1}), \tag{16}$$

where $M_{\mathbf{XW}} = p\lim\left(\frac{\mathbf{X}_i'\mathbf{W}_i}{N}\right)$, and $V_b = \mathrm{var}(b_N)$. A consistent estimate of the asymptotic variance of $\tilde{\delta}_A$ is given by

$$AVAR(\tilde{\delta}_A) = (\hat{M}_{\mathbf{XW}}'A_N \hat{M}_{\mathbf{XW}})^{-1} \hat{M}_{\mathbf{XW}}'A_N \hat{V}_b A_N \hat{M}_{\mathbf{XW}} (\hat{M}_{\mathbf{XW}}'A_N \hat{M}_{\mathbf{XW}})^{-1}, \tag{17}$$

where $\hat{M}_{\mathbf{XW}}$ and $\hat{V}_b$ are consistent estimates of $M_{\mathbf{XW}}$ and V_b respectively. The most efficient estimator relative to $\hat{\Pi}$ is obtained by choosing optimally A_N as a consistent estimate of the inverse of the covariance matrix of the orthogonality conditions

$$A_n = \hat{V}_b^{-1} = \left(M_{\mathbf{XX}}\,\mathrm{var}(\mathrm{vec}(\hat{\Pi}))M_{\mathbf{XX}}\right)^{-1}. \tag{18}$$

Now we only need an estimate for $\mathrm{var}(\mathrm{vec}(\hat{\Pi}))$. Let us consider

$$\mathrm{vec}(\hat{\Pi}) = \begin{pmatrix} \hat{\pi}_1 \\ \hat{\pi}_2 \\ \hat{\pi}_3 \end{pmatrix} = \arg\max(L) = \arg\max(L_1 + L_2 + L_3), \tag{19}$$

where L_j, $j = 1, 2, 3$, is the corresponding likelihood function. Subject to suitable regularity conditions, a first order expansion of $\partial L(\hat{\Pi})/\partial\Pi$ around the true value of Π gives

$$\left(-1/N\,\mathrm{diag}\left(\frac{\partial^2 L_j}{\partial\pi_j\partial\pi_j'}\right)\right)\sqrt{N}\,\mathrm{vec}(\hat{\Pi} - \Pi) = 1/\sqrt{N}\sum_i \begin{pmatrix} \partial L_{i1}/\partial\pi_1 \\ \partial L_{i2}/\partial\pi_2 \\ \partial L_{i3}/\partial\pi_3 \end{pmatrix} + O_p(1), \tag{20}$$

which suggests an estimate for the variance of the form

$$\mathrm{var}(\mathrm{vec}(\hat{\Pi})) = \hat{H}^{-1}\hat{\Psi}\hat{H}^{-1}, \tag{21}$$

where $\hat{H} = \text{diag}(N^{-1}\partial^2\hat{L}_j/\partial\pi_j\partial\pi_j')$ and $\hat{\Psi} = N^{-1}\sum_i\{\partial\hat{L}_{ij}/\partial\pi_j\}\cdot\{\partial\hat{L}_{ih}/\partial\pi_h'\}$

Finally, under the null hypothesis of lack of misspecification, a test statistic of the overidentifying restrictions can be derived for the optimal GMM estimator,

$$S = N\left(\sum_i^N \hat{u}_i' X_i\right)\hat{V}_b^{-1}\left(\sum_i^N X_i' u_i\right) \sim \chi^2_{3m-(k_1+k_2+K_3+6)}, \tag{22}$$

where $\hat{u}_i = \hat{I}_i^* - \hat{\mathbf{W}}_i\tilde{\delta}_A$.

In brief, the method discussed above works as follows:

1. First, the reduced form estimates are obtained.
2. Second, using these reduced form estimates, a sub-optimal GMM estimator is implemented that allows us to compute a consistent estimate of V_b.
3. The consistent estimate of V_b and the reduced form estimates are used to obtain the most efficient estimates relative to the first-stage reduced form ones.

Appendix B: Data

Endogenous variables

Independent: Dummy variable that equals one if the individual does not live in the parental home.
Work: Dummy variable that equals one if the individual is working (full or part time) at the interview date.
Studying: Dummy variable that equals one if the individual is carrying on any type of education. It is worth noting that 21.9% of female students are said to attend "other type of education", different from primary, secondary, and College education, while only 14.9% of male students declare to do so.

Exogenous variables

- *Individual characteristics:*
  Educational: We define three dummy variables reflecting the highest degree completed by the individual. *Educ2* equals one if the individual has finished primary school, *Educ3* equals one if s/he has finished secondary school and *Educ4* equals one if some College degree has been attained.
  City: Dummy variable that equals one if the individual lives in a large city (more than 500,000 inhabitants).
  Village: Dummy variable that equals one if the individual lives in a small village (less than 2,000 inhabitants).
  Parental transfers: Regular or occasional cash transfers, plus housing subsidies received by the young offspring living apart.
  Other non labor income: This is the summation of all types of current income sources, different from parental transfers and labor earnings or public sub-

sidies related to the economic activity (like the unemployment compensation). It includes lotteries, returns on capital (positive and negative when borrowing), school grants, and other public transfers.
Non labor income: Parental transfers plus other non labor income.

- *Economic variables:*
Rental values: Regional average across the 50 Spanish provinces (excluding Ceuta and Melilla) of: (i) annual rents by square meter actually paid in rental housing, and (ii) self-imputed annual rents by square meter for owner-occupied and other non-rental housing. Source: 1990–1991 EPF.
Owning costs: Regional average across the 17 Spanish *Comunidades Autónomas* (excluding Ceuta y Melilla) of house prices by square meter. Source: "*Precio medio del m² de las viviendas*", Ministerio de Fomento.
Rental accommodation: Regional percentage (across the 50 Spanish provinces) of rental accommodation. Source: 1990–1991 EPF
Unemployment: Regional unemployment rates (across the 17 Spanish Comunidades Autónomas) for the population as a whole and disaggregated by sex and age. Source: *Encuesta de Población Activa.*
Illiteracy rate: Number of illiterate individuals older than 10 per thousand inhabitants in every province. Source: *Encuesta de Población Activa.*

Table B.1. Characteristics for young individuals between 18 and 35 years old

	Males		Females	
	Mean	Standard error	Mean	Standard error
Endogenous variables				
Independent	0.334	0.472	0.446	0.497
Work	0.635	0.481	0.372	0.483
Studying	0.224	0.417	0.271	0.444
Exogenous variables				
Individual characteristics				
Age	25.8	5.2	26.0	5.2
Educ2	0.499	0.500	0.474	0.499
Educ3	0.355	0.479	0.350	0.477
Educ4	0.105	0.306	0.133	0.339
City	0.532	0.499	0.557	0.497
Village	0.145	0.352	0.129	0.335
Non labor income (N.L.I.)	103,607	282,892	168,636	344,843
• Parental transfers only	300,538	257,160	334,997	423,849
(Observations %)	(4.54%)		(5.86%)	
• Other N.L.I. only	29,723	242,582	57,822	217,738
(Observations %)	(17.16%)		(12.42%)	
• Both	328,842	353,161	445,495	441,016
(Observations %)	(1.66%)		(1.45%)	
Economic variables				
Rental values	3,287	1,344	3,345	1,369
Owning costs	86,987	22,426	87,476	22,987
Rental accommodation	0.117	0.041	0.117	0.041
Regional unemployment (RU)	16.927	5.468	16.796	5.459
RU by age-males	17.533	11.187		
RU by age-females			30.918	12.027
Illiteracy rate	0.041	0.026	0.041	0.026
Sample size	9,741		9,535	

Table B.2. Individual distribution according to dependent variables

	Males		Females	
	Number of observations	%	Number of observations	%
Working = 1, studying = 0, independent = 0	2984	30.63	1,675	17.57
Working = 0, studying = 1, independent = 0	1,631	16.74	1,931	20.25
Working = 0, studying = 0, independent = 0	1,609	16.52	1,385	14.53
Working = 1, studying = 0, independent = 1	2,695	27.67	1,417	14.86
Working = 0, studying = 0, independent = 1	268	2.75	2,475	25.96
Working = 1, studying = 1, independent = 0	268	2.75	295	3.09
Working = 1, studying = 1, independent = 1	242	2.48	161	1.69
Working = 0, studying = 1, independent = 1	44	0.45	196	2.06

Table B.3. Bivariate probit correlation coefficients

	Males		Females	
	ρ	LR test(*)	ρ	LR test(*)
Work – independence	0.448	390.239	−0.157	59.324
Work – studying	−0.575	726.757	−0.466	524.355
Independence – studying	−0.099	11.329	−0.287	128.212

(*) The likelihood ratio test of $H_0 : \rho = 0$ follows a $\chi^2(1)$

Table B.4. Reduced form first stage estimates: females

	Work		Independence		Study	
	Coeff	t-ratio	Coeff	t-ratio	Coeff	t-ratio
Age	−0.135	−1.188	1.660	10.469	−1.256	−8.823
Age^2	−0.341	−5.534	−0.498	−5.768	0.845	11.087
Primary school	0.367	5.028	0.231	2.970	0.338	2.443
Secondary school	0.411	5.446	−0.154	−1.905	1.552	11.268
Higher education	0.907	11.293	−0.378	−4.410	1.681	11.860
Regional UR	−0.695	−1.516	−0.171	−0.324	−1.128	−1.965
Regional UR – by age (females)	−1.216	−5.358	−1.311	−4.716	0.942	3.675
City	0.003	0.097	0.049	1.153	0.264	5.956
Village	−0.064	−1.382	−0.241	−4.334	−0.048	−0.784
Non labor income	−0.046	−4.457	0.073	5.744	0.008	0.747
Owning costs	−0.037	−0.414	−0.384	−3.341	−0.063	−0.596
Age X owning costs	−0.178	−1.203	0.602	2.855	0.212	1.100
Rental values	−0.122	−0.652	−0.072	−0.312	0.208	0.952
Age X rental values	1.443	5.788	−0.468	−1.342	−0.773	−2.376
Rental accommodation	0.647	1.921	1.186	2.954	0.698	1.665
Illiteracy rate	−0.423	−0.535	2.420	2.600	−0.495	−0.500
Intercept	−0.177	−1.389	0.203	1.349	−2.241	−12.12
Log-likelihood	−5,956		−4,050		−3,728	
Number observations			9535			

Table B.5. Reduced form first stage estimates: males

	Work		Independence		Study	
	Coeff	*t*-ratio	Coeff	*t*-ratio	Coeff	*t*-ratio
Age	1.024	8.006	2.049	11.485	−1.296	−8.855
Age^2	−0.685	−9.957	−0.269	−2.626	0.670	8.241
Primary school	0.591	8.311	0.441	5.390	−0.087	−0.733
Secondary school	0.259	3.554	0.378	4.445	1.192	10.136
Higher education	0.203	2.489	0.231	2.522	1.571	12.646
Regional UR	−3.909	−7.710	−2.694	−5.012	−0.313	−0.516
Regional UR – by age (males)	−0.288	−1.051	−0.062	−0.164	0.448	1.431
City	−0.206	−5.378	0.025	0.581	0.318	7.025
Village	−0.107	−2.227	−0.272	−4.942	−0.122	−1.934
Non labor income	−0.077	−7.510	0.046	4.121	0.038	3.324
Owning costs	−0.163	−1.729	−0.394	−3.097	−0.080	−0.730
Age *X* owning costs	0.047	0.277	0.176	0.781	0.209	1.055
Rental values	0.430	2.126	0.358	1.430	−0.061	−0.270
Age *X* rental values	0.923	3.142	0.239	0.625	−0.627	−1.819
Rental accommodation	−0.147	−0.408	1.480	3.674	−0.017	−0.041
Illiteracy rate	2.176	2.604	6.162	6.531	−2.484	−2.525
Intercept	0.872	6.929	−0.915	−6.086	−1.714	−10.20
Log-likelihood	−5,086		−3,859		−3,571	
Number observations			9741			

Endnotes

[1] For an early exposition of this idea in Spanish literature, see Revenga (1991). For more recent analysis, see Robinson (1998) and Toharia et al. (1998).

[2] Among the young people studied in this paper who reside in a household headed by someone else, almost 7% live in a household headed by a relative, other than a parent. However, to simplify the terminology, in what follows we will refer to all such households as the "parental" household.

[3] On the other hand, Ermish and Di Salvo (1997) and Ermish (1999) develop a model of household formation in which parents are altruistic about their children, and housing is a local public good in the sense that housing services per person are not affected by household size. The theoretical model derives predictions about the impact of the price of housing, young adults' income and parental income on the probability that a young adult lives apart from his/her parents.

[4] Recent empirical work in this area is of this type. RW use the kinship-linked cohorts of 9 National Longitudinal Surveys from 1967 to 1981 in the U.S. Their sample consists of more than 5,000 young men who were ages 14–17 at the time of the 1966 interview. E uses people aged 16–29 in the first four waves of the 1991–1994 British Household Panel Survey (BHPS). Ermish and di Salvo (1997) use a sample of 10,500 British people born in 1958 and interviewed in 1991 when they were 33 years old. Ermish (1999) focus on persons aged 16–30 from the first five waves of the 1991–1995 BHPS.

[5] This is the Ashford and Sowden (1970), Amemiya (1975) and Zellner and Lee (1965) model, discussed as *Case 3* in Heckman (1978). For the panel data case, see Heckman (1981).

[6] Within the empirical literature, Haurin et al. (1993, 1994) also insist that the child's income is endogenous because labor supply is jointly decided along household formation.

[7] As far as the fourth possible state, in RW transfers are always positive whenever the two generations live together, while in E there is no data on non-housing transfers to the coresiding child.

[8] In a multinomial logit framework, data limitations would force us to ignore the human capital investment decision. At any rate, estimates of a 4 state multinomial logit arising from the labor force participation and the household formation decisions are available upon request.

[9] Similar simultaneous equation models have been used in other contexts in order to analyze the relationship between disability and labor force participation (Stern 1989), or the relationship between employment and care giving to elderly parents (Wolf and Soldo 1994).

[10] Mallar (1977) or Heckman (1978), among others, propose two-stage methods in this type of situation. For a revision of these methods, see Maddala (1983).

[11] Independent equation estimation at the first-stage will in general yield inefficient estimates because it ignores the possible dependence among the equations through the error terms, u_{ji}. Taking into account such dependence would typically require the use of simulation based estimators (see Hajivassiliou and Ruud 1994).

[12] If the three decisions in Eqs. (1) to (3) were not simultaneous and we were only interested in the study of the decision of leaving the parental household, a single equation with two endogenous dummies could also be estimated. However, the two-stage estimators we have discussed could not be implemented, and some simulation technique of estimation had to be applied. Finally, note that a multivariate probit with endogenous dummies will also complicate the estimation.

[13] The six European countries are the ones studied in Fernández Cordón (1997) using the Labor Force Survey. For another compartive study between Northern and Southern European countries, using the European Household Panel, see Iacovou (1998).

[14] Results for an upper bound of 30 years are available on request. They are not very sensitive to the choice of the upper bound, especially for males. For females, although in general the sign and size of the estimated parameters are comparable for both age groups, the overidentifying restrictions are marginally rejected. Moreover, the independence variable, that is marginally significant for women between 18–35 in the equations of working and studying losses completely its significance when considering only women between 18 and 30 years of age.

[15] Table 2 of Cantó and Mercader (1999) show the entry-level jobs in 1996 by new school leavers (aged 16 to 29) one year after leaving education in different selected European countries. Spain is an outlier: 80% of this group hold a temporary job. On the other hand, according to Güell and Petrongolo (2000), in 1990 less than 15% of all temporary contracts become indefinite. Finally, Jimeno and Toharia (1993) and De la Rica and Felgueroso (1999) find that the temporary workers earn approximately 10% less than permanent ones, after controlling for observable personal and job characteristics.

[16] This is confirmed by all demographic studies in the subject. See, for instance, Vergés (1997) or Jurado (1997).

[17] We thank an anonymous referee for this suggestion.

[18] The results are substantially different when ignoring the endogeneity of the working and studying decisions. For women, we find a positive effect of the propensity to work and to study on the probability of coresidence. For males, working provides incentives to leave the parental house but studying has no effect on the household formation decision.

[19] According to Ermish (1996), the positive effect found by Rosenzweig and Wolpin (1993) may reflect the different definition of coresidence among students: young men who are at college and away from home are classified as coresiding with their parents if they report themselves as attached to their parents' household. Nevertheless, in their fixed effects multinomial logit model, Rosenzweig and Wolpin (1993) also find a positive effect of attending college on the probability of receiving a parental transfer.

[20] Rosenzweig and Wolpin (1993) find a positive but small direct effect of the weeks spent in unemployment on coresidence, as well as on the parental transfers when living apart. Thus, in the U.S., the family acts as a cushion in bad times independent of living arrangements.

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Group living decisions as youths transition to adulthood

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Abstract. This study follows teens through young adulthood as they transition to independent living. We focus on a little studied issue: why some youths live in groups rather than alone or with parents. This choice is important because the size of the group has a substantial impact on the demand for dwelling units; the more youths per dwelling the lower is aggregate demand and the greater is population density. Our study also adds to the knowledge of which factors influence youths' choice of destination as they leave the parental home. The empirical testing uses a discrete hazard model within a multinomial logit framework to allow for more than one possible state transition. We find that economic variables have little impact on the decision of whether to exit to a large versus a small group, while socio-demographic variables matter. We also test a new push-pull hypothesis and find that the pull of economic variables on the probability of exiting the parental home increases as youths reach their mid to late twenties.

JEL classification: D1, J12, R20

Key words: Group living, household formation, home-leaving

1. Introduction

Interest in the process of youths leaving the parental home and moving to an independent living arrangement continues to be high. Recent studies about

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youths' tendency to leave home include analyses of the impact of the cost of shelter and potential earnings (Haurin et al. 1993, 1994; Bourassa et al. 1994; Whittington and Peters 1996; Ermisch and Di Salvo 1997; Johnson and DaVanzo 1998; Ermisch 1999), the impact of family structure (Aquilino 1991; Goldscheider and Goldscheider 1998), the impact of gender, religion, and ethnic differences (Goldscheider and DaVanzo 1989; Buck and Scott 1993), and the impact of parental and youths' normative expectations about home-leaving (Goldscheider and Goldscheider 1993b). A fairly comprehensive portrait of the sociological, psychological, demographic, and economic explanatory factors is emerging.

Understanding home-leaving is key to modeling household formation. Household formation impacts aggregate housing demand, population density, fertility, labor force mobility, and the demand for public services. We argue that while the substantial attention paid to the timing of youths' departure from the parental home is appropriate, careful study of all destination possibilities is equally important as is understanding whether the magnitude of an explanatory factor's influence differs by respondent age.

There is some controversy about the appropriate categorization of exit types. Early studies focused on a dichotomous choice: exit or remain in the parental home. More recently, multiple exit "destinations" have been considered where a destination is defined as a type of living arrangement outside the parental home. Examples include exiting to marry, live alone, live with a relative, or live in a group. Studies have attempted to sort out the impact of explanatory factors on the probabilities of exiting to each of the destination categories. However, the choice of destination categories has often been dictated by data limitations or been the result of an apparently arbitrary choice. When refocusing research from the origin (parental home) of a youth's transition to the destination, a more systematic approach to study alternative configurations of destinations should be employed. We recognize that the extent of disaggregation of destinations is limited by sample size; the more categories, the smaller the number of exits to a particular category. If destinations are highly disaggregated, empirical analysis becomes more difficult and inferences of relationships are harder to draw with confidence. However, when there is a compelling theoretical reason for making a distinction among destination types, exits should be disaggregated by type.[1]

We focus on the question of which social, demographic, and economic factors influence the tendency of a youth to exit to a small group living arrangement compared with a large group. Although exiting to a group has been included as a destination category, to our knowledge, no studies have distinguished the size of grouped arrangements. From a theoretical perspective, we conduct new tests of factors hypothesized to influence destination choice, focusing on a comparison of economic with socio-demographic variables. The goal of our analysis is to determine whether a finer differentiation of exits to groups of small or large size is needed in the analysis of youths' home-leaving.

Failure to disaggregate destination categories could lead to the statistical problem of aggregation bias. For example, an explanatory variable such as the cost of shelter could negatively influence the tendency to exit to a small group, but positively influence the tendency to exit to a large group. If exits to small and large groups are combined into a single category, it is possible that the resulting estimation would indicate that this explanatory variable had no impact on the tendency to exit to a group. Inappropriate aggregation of various group sizes into a single category may lead to biased statistical inferences.

Weighting our sample data allows us to describe U.S. patterns in youths' tendencies to exit to small compared with large groups where a small group is defined as the respondent and one other adult other than spouse or opposite sex partner. The percentage of youths selecting to exit to small groups is 18 and the percentage exiting to large groups is 15. Clearly, group living is an important destination when exiting the parental home. Further, the near equality of the two percentages suggests that both categories are important.

2. Literature

2.1. Types of transitions

Groups serve important social, psychological and economic functions. Differences in the expected outcome of living in a large compared with a small group form the basis for our hypotheses. The size of the group chosen should be a function of economic resources. Large groups may afford greater financial sharing opportunities and advantage can be taken of economies of scale in obtaining shelter. Financial risks are spread more widely in a larger group; that is, if one resident leaves the dwelling, the temporary increase in costs to remaining residents is lower in a large group. Socially, group living offers companionship with one or more others as an alternative to the social environment of the family. Individuals will tend to live with others who share similar personality characteristics, have common roles and interests, and similar biological traits (Hare 1982). Psychologically, group living provides independence from parental supervision and freedom from the psychological and sexual commitment required of married or partnered relationships. Also, the choice of group size depends on an individual's preference for privacy. While small groups offer greater intimacy, they provide less security and less stability for the individual if some group member should decide to withdraw (Hare 1981). The choice between living alone and living in a group likely distinguishes individuals with greater self-confidence in their skills for total versus partially independent living; e.g., the abilities to cook, clean, do laundry and handle financial matters. In this sense, large groups require the least independence.

We account for types of exits other than to groups by including two other destination categories: exiting to live alone and exiting to marry or live with a partner. The contemporary trend toward slower entry into marriage has had divergent effects. Myers (1992) notes the positive relationship of this trend with a rise in nonfamily households, specifically an increase in the number of youths living alone. Others have associated the trend with an increase in the tendency of youths to remain in their parental home (Heer et al. 1985; Buck and Scott 1993; Goldscheider and Goldscheider 1993a). We follow prior studies of the transition of young adults to independent living by separating marriage and other sexually partnered relationships from living alone (Thornton et al. 1993).

2.2. *Explaining the transitions: Economic factors*

In the literature, factors associated with leaving home fall into several major categories: economic factors, family background and family structure, and demographic and personal characteristics. Characteristics of both youths and parents have been considered relevant to the decision of when a youth leaves home. Underlying the inclusion of these measures are arguments regarding motivations for leaving home. These include normative or age-appropriate expectations about when to leave home, stress factors motivating an exit, opportunities motivating an exit, a general preference for autonomy and privacy, and intergenerational transfers (Rosenzweig and Wolpin 1993, 1994). Our discussion focuses on hypotheses related to the tendency to exit to small versus large groups.

Household formation should depend on the cost of independent living, with the cost of shelter as one component. While most demographic studies have omitted this intuitively influential factor, the empirical economic literature finds that high housing costs reduce the tendency of youths to reside outside the parental home (Borsch-Supan 1986; Haurin et al. 1993, 1994, 1997; Bourassa et al. 1994; Ermisch 1999). Ermisch and Di Salvo (1997) find that higher housing costs reduce the tendency of British women, but not men, to live outside the parental home. Johnson and DaVanzo (1998) find this effect for Malaysian sons, but not daughters.

Ermisch and Di Salvo (1997) and Ermisch (1999) show that the long run impact of the cost of shelter on a youth's living arrangement is theoretically indeterminate. They argue that variations in housing cost affect both the housing consumed by a youth living outside the parental home and the housing consumed by the parents. For example, comparing a high housing cost locality to a low cost area, the parents' quantity of housing consumed will be relatively small, as is a youth's potential quantity consumed. Thus, it is not clear that a youth is better off remaining with parents. Ermisch and Di Salvo find that if the absolute value of the price elasticity of housing demand is less than or equal to one (as in the U. S. and Great Britian), then youths should respond to higher housing costs by remaining with their parents. We follow Ermisch and Di Salvo and hypothesize that the greater the cost of shelter, the less likely is a youth to live alone rather than remain with a parent. We extend their model by hypothesizing that as housing costs increase, the greater is the likelihood of living in a group compared with living alone. Further, the probability of living in a large group should rise, the higher are housing costs.

As a youth's earnings ability rises, we expect the tendency to live in a large group to decline. Generally, the literature has measured an individual's ability to pay the cost of independent living by using personal income (Ermisch and Overton 1985; Goldscheider and DaVanzo 1985; Avery et al. 1992). However, as argued in Haurin et al. (1993), Bourassa et al. (1994), and Whittington and Peters (1996), income is the product of the wage rate and the amount of labor supplied. Participation in the paid labor force is a decision that occurs jointly with the decision on household formation. For example, a youth may not work because he or she is subsidized in the parental household. Similarly, observed wages may not accurately reflect earnings capacity if the current job is part-time (which is more likely if the youth does not reside alone). Therefore, the potential wage, or the wage that could be earned if a youth took on the responsibility of independent living, is the theoretically preferred predictor of the tendency to reside outside the parental household. We estimate a potential wage for each respondent for each year using the procedure described in the appendix.

Neither the research on the impact of housing costs nor the more substantial research on the impact of income or wages on household formation has considered whether the influence of these factors varies with a youth's age. We argue that the impact of a unit change in an economic variable increases as a youth ages. At very young ages (15–18), we expect the impact of socio-demographic variables to dominate the explanation for leaving the parental home, but for older youths (25+), we expect both economic and socio-demographic variables to play significant roles. Underlying this hypothesis is an argument that teens leave the parental home if "pushed" by social factors. The "pull" of economic variables such as a high potential wage or low housing cost is ineffective because, typically, parents act as an economic buffer. However, for older youths, the desire for independence from parents reinforces the pull of favorable economic factors, making a transition from the parental home more likely. Parental resources have been shown to be important to a youth's transition to independent living. Avery et al. (1992), Whittington and Peters (1996), Ermisch and Di Salvo (1997), and Ermisch (1999) argue that parents with sufficient resources influence their children's choice of living arrangement by altering financial transfers they make with the children. The direction of impact may depend on the youth's age. For example, parents could use their resources to keep children at home during the teenage years, but then use their resources to promote home-leaving when the youth becomes a young adult. As an alternative to parental income (not observed in our data), we include a measure of parental education. Conditional on exiting the parental home, we expect increased parental resources to increase the tendency of exiting to small groups relative to large.

Macro-level economic characteristics of a youth's environment such as the availability of public assistance and local unemployment may affect housing decisions. Haurin et al. (1993) and Whittington and Peters (1996) argue that young adult women consider the availability of public assistance benefits through the Aid to Families with Dependent Children (AFDC) program in deciding whether to leave home. Untested is whether this availability has an impact on their tendencies to group-up or live alone. Whittington and Peters argue that a youth's response to AFDC will be age-linked. If 18 or older, a higher AFDC payment rate should encourage an exit from the parents' home, if less than 18, the youth's parents should desire to retain the child in their home. In all cases, the impact of AFDC on residence should be limited to eligible youths. We create an indicator variable reflecting eligibility and interact it with a measure of AFDC payment rates.

A higher rate of local area unemployment lowers the probability of securing a job that pays the youth's potential wage and thus increases the financial risk of independent living (Ermisch and Di Salvo 1997; Haurin et al. 1997). Thus, we expect youths in localities with high unemployment to be more likely to exit to a large group than to a small group.

2.3. *Explaining the transitions: Family background and family structure*

One of the largest areas of interest with regard to the transition of youths to residential independence has been prior family structure and relationships. Of particular interest has been analysis of the impact of prior residence with a stepmother or a stepfather (White and Booth 1985; Mitchell et al. 1989; Aquilino 1991; Avery et al. 1992; Haurin et al. 1997; Goldscheider and Goldscheider 1998). Family stress deriving from these parental residential situations and exposure to unsuccessful relationships are often cited as the underlying cause for a youth's early exit from the parental home. However, this theory does not suggest a particular impact on an exiting youth's choice of group size. Single parent households are likely to have lower income; thus, they are able to provide less financial support for an exiting child. We expect that youths exiting from single parent households will be more likely to exit to a large group.

We expect that the greater the number of siblings living outside the parental home, the more likely is a transition to a large group. The justification for this hypothesis is that there is an enlarged pool of potential roommates when there are more siblings living independently. We expect that the greater the number of siblings who continue to reside in the parental home, the more likely will be the selection of a large group, assuming exiting youths continue to prefer to live with a large number of people. First born youths are more likely to have the skills needed for independent living; thus, we expect them to be more likely to exit to live alone or in small groups compared with large.

Avery et al. (1992) suggest that responsibility for minor own-children has mixed effects on leaving home. Young unwed parents have greater need for support from their own parents, but the presence of grandchildren reduces privacy for the parents of the young adults. An increased number of own-children should decrease the likelihood of leaving home for large groups.

Because Catholic youths tend to be raised in large families, their expectation may be that large families are normative. This observation suggests that exiting to a large group is more likely. However, as noted above, we control for the number of siblings. Finding a significant impact of being Catholic on the tendency to exit to a large group suggests that the norms supporting living in large groups exist for Catholic youths independently of the actual number of siblings.

2.4. Explaining the transitions: Demographic and personal characteristics

Large groups offer enhanced safety, and we expect that women select residences to enhance safety more so than do males. Offsetting this effect is the observation that young women are more likely to have the skills for independent living than young men, hence would be less likely to live in large groups. Allowing gender specific behaviors suggests that we should test the assumption that the samples of men and women can be pooled by first estimating separate models and then testing whether a combined model is acceptable. We find that the impacts of the explanatory factors differ significantly by gender; thus, we present only separate results.[2]

The youngest members of our sample are less likely to have confidence in their skills for independent living. The implication is that we expect older youths to be less likely to exit to large groups, and somewhat less likely to exit to small groups.

Discrimination in the housing market could limit the residential choices of Black and possibly of Hispanic youths, perhaps leading to a greater tendency to exit to large groups. Following Whittington and Peters (1996), we do not estimate separate models by race, the reason being our concern about the relatively small number of exits per race/gender category.

Students in college and not living in a dorm are likely to seek financial sharing of shelter costs by living in a large group. Including as an explanatory variable a measure of a youth's earning ability is not sufficient to capture this effect because college students have relatively high earnings potential, but they are likely to have relatively low current income, hence they have an incentive to share shelter costs. We include an indicator variable for a youth being in college.

Another variable not typically considered is a youth's health. We expect that poor health reduces a youth's ability to live independently. If a youth reports a health concern and exits the parental home, there are potentially offsetting factors impacting the choice of living arrangement. A youth with a health problem may seek to live in a large group to more efficiently share household responsibilities. However, finding roommates may be more difficult for someone with a health impairment.

We test for the impact of living in an urban area on transition probabilities. Urban areas tend to have relatively high crime rates; thus, we expect youths will tend to form large groups for additional security. Urban areas also are associated with relatively high shelter costs, but we control directly for this variable. We also include three dummy variables indicating the region of residence in the U.S.

3. Hazard model

Our model describes an individual's decision at any point in time to reside in one of five possible arrangements. We estimate a reduced form model using a competing risks framework (Kalbfleisch and Prentice 1980). The occurrence of a transition from the parental home to another living arrangement removes the individual from the risk of experiencing any other transition. The competing risks framework characterizes each transition by a separate transition rate and hazard function.

The type-specific hazard function is defined as the probability that an individual will move from the parental home to living arrangement type j after $t + \Delta t$ years given that they lived in their parental home at least t years (Kalbfleisch and Prentice 1980; Allison 1984). The hazard rate h is defined to be a function of time and a set of explanatory variables:

$$h_j(t, Z) = \lim_{\Delta t \to 0} [P(t \leq T < t + \Delta, J = j \mid T \geq t, Z)/\Delta t] \quad j = 1, \ldots, m \qquad (1)$$

where j is the destination living arrangement following the transition; t is the number of years living in the parental home; and Z is a vector of sociodemographic and economic factors that may change in value over time. The overall hazard function is the sum of all the type-specific hazard functions. The period of observation begins with a youth living in his or her parental home and we follow the youth until the first exit from the parental home or until the observation is right censored. We use a discrete-time framework to estimate the model because of the annual nature of the data; that is, we can identify the time of transition only by comparing responses in adjacent survey years. The model is multinomial logit because there are four exit types (Greene 1993).

4. Sample and variables

4.1. Sample characteristics

We employ annual data from the 1979 to 1992 waves of the U.S. National Longitudinal Survey of Youth (Center for Human Resource Research 1993). The NLSY79 contains a national sample of youths aged 14 to 21 in 1979. Oversamples of Blacks, Hispanics and economically disadvantaged whites permit statistical analyses of these population subgroups. Survey attrition rates are low with approximately 90 percent of the eligible sample retained as of the 1992 survey.

We limit our study sample to respondents age 14 to 17 who resided in their parental household in 1979. Residential locations are followed through the first exit or until 1992, yielding 27,472 person-year observations. Because exiting prior to age 16 is highly unlikely, we omit 919 observations when a respondent is age 14 or 15, these ages only observed in 1979 or 1980. Missing data reduce our final sample to 16,184 person-year observations (7,360 for females and 8,824 for males) for which 2,661 exits are observed.[3]

4.2. Dependent variable

The dependent variable is a categorical measure of the five possible current living arrangement of the respondent. The types of exits are coded as: '0' if continuing to reside in the parental home, '1' if exited to live alone with or without own-children, '2' if exited to live with a spouse or partner and possibly children, '3' if exited to live with a group that includes one nonspouse/nonpartner adult, and '4' if exited to live in a group with more than one other nonspouse/nonpartner adult. The distribution of these exits in our sample is 28% to living alone, 39% to living with spouse/partner, 18% to living in a small group, and 15% to living in a large group. Descriptive statistics for the dependent and independent variables are listed in Table 1. These statistics cover the 1979–1992 period and include all person-year observations.

4.3. Explanatory variables

Housing cost is a continuous variable representing the constant-quality housing cost in the area in which the respondent resides. By using a constant-quality measure, variations in the average amount of housing consumed between communities and over time are controlled.[4] Another economic factor is Potential Wage, a continuous variable that estimates the wage the respondent could obtain if he or she worked full-time. We estimate wage using a two-step framework that is described in the appendix (Heckman 1979; Greene 1995).

Table 1. Descriptive statistics for the full sample: 1979–1992

Variable	Females		Males	
	Mean	Std. Dev.	Mean	Std. Dev.
Exit from parental household (1 = yes)	0.19	0.39	0.15	0.36
Age in years	19.78	2.71	19.95	2.80
Black (1 = yes)	0.25	0.43	0.24	0.43
Hispanic (1 = yes)	0.15	0.36	0.14	0.35
In-high school (1 = yes)	0.23	0.42	0.25	0.43
In-college (1 = yes)	0.24	0.43	0.19	0.39
Out of school-LTHS (1 = yes)	0.11	0.31	0.21	0.41
Health limit (1 = yes)	0.03	0.16	0.02	0.12
Stepmother (1 = yes)	0.01	0.09	0.01	0.12
Stepfather (1 = yes)	0.06	0.24	0.05	0.22
Single parent (1 = yes)	0.33	0.47	0.33	0.47
Number of sibling-in	1.77	1.57	1.84	1.61
Number of sibling-out	1.69	2.01	1.74	2.14
First born (1 = yes)	0.25	0.43	0.25	0.43
Number of own-children	0.11	0.38	0.01	0.13
Catholic (1 = yes)	0.37	0.48	0.34	0.47
Religious attendance (1 = more than once a month)	0.61	0.49	0.51	0.50
Local housing cost	0.58	0.16	0.58	0.16
Potential wage in 1979 dollars/hour	3.88	0.89	4.96	1.20
Parental education in years	11.05	2.95	11.02	3.07
Local unemployment rate (%)	8.62	3.55	8.66	3.59
Urban (1 = yes)	0.62	0.48	0.60	0.49
AFDC eligible (%)	15.30	56.06	2.69	27.35
South (1 = yes)	0.40	0.49	0.37	0.48
Midwest (1 = yes)	0.27	0.44	0.28	0.45
West (1 = yes)	0.14	0.35	0.15	0.36
Samples size	7,360		8,824	

All variables denominated in dollars are deflated using the CPI-U with 1979 as the base year.

Two additional variables are created by interacting the housing cost and potential wage with a semicontinuous variable (Age 18) that equals the respondent's age if age is greater than or equal to 18; otherwise, if age is less than 18, the interaction variable takes the value zero. This specification allows the interaction variable to capture age-related effects once the respondent's age has passed the threshold of 18 years.[5]

Other economic variables include Parental Education, a proxy for parental resources and measured as the highest grade completed by the respondent's mother (if data on the mother are not available, then the father's value is used), and the Local Unemployment Rate, a continuous variable measuring the local unemployment rate for the labor market in which the respondent resides. We also include AFDC Eligible, the product of an indicator of whether the respondent is eligible to receive AFDC payments and the maximum AFDC benefit for a family of three in the respondent's state.[6]

Demographic and personal characteristics hypothesized to affect living arrangements include the respondent's age and its square to test for nonlinear effects. Race and ethnicity are operationalized through two indicator variables: Black non-Hispanic and Hispanic, with non-Black non-Hispanic being the omitted race/ethnicity category. We interact a dummy variable for whether the respondent is currently attending school with an indicator of whether the respondent's highest grade completed is less than high school (LTHS) or high school or beyond (GEHS). The resulting three variables are named In-High School, In-College, and Out of School-LTHS. The omitted category is Out of School-GEHS. Health Limit is an indicator variable for whether the respondent reports having a health condition that limits his or her ability to work. Urban is an indicator variable equaling unity if the respondent lives in an MSA. Three regional indicator variables are included with the eastern U.S. being the omitted category.

Family background and structure variables include indicator variables for whether the respondent is living with a stepmother, a stepfather, or a single parent. These variables are lagged one year to avoid endogeneity with the respondent's living arrangement decision. Family size variables include: Number of Siblings-In, a continuous variable for the number of the respondent's siblings that live in the parental home; Number of Siblings-Out, a similar measure for the number of siblings that live outside of the parental home; and First Born, an indicator variable for whether the respondent is the first child of his or her mother. The two measures of siblings are lagged one year. Other family variables include Own-Children, a continuous variable for the number of respondent's own-children that lives with the respondent, and Catholic, an indicator variable for whether the respondent reports his or her religious affiliation as Catholic. We include as a control variable an indicator variable for whether the respondent reports attending religious services more than once a month (Religious Attendance).

5. Results

We report the estimates from the multinomial logit model in Tables 2 (females) and 3 (males).[7] Listed are the marginal impacts of changing an explanatory variable by one unit on the probability of observing a particular living arrangement. Marginal impacts can be computed for all possible living arrangements *including* the reference category in the multinomial logit, remaining with parents (see Greene 1995, Ch. 24). An example of the interpretation of a marginal effect in Table 2 for the coefficient of a dummy variable such as Black is that a Black female respondent is 4.09 percentage points more likely to live with parents than a white respondent in any particular year. For a continuous variable such as the unemployment rate, the interpretation is that a one percentage point higher unemployment rate increases a female's probability of remaining with parents by 0.04 percentage points. The first column of data contains results for remaining with parents, followed by columns for exiting to live alone, live with a spouse or partner, live in a small group, and live in a large group. Reading across a row reveals the comparative effects of a unit change in an explanatory variable. We present a series of pair-wise tests of differences in coefficients among all types of living arrangements. Using codes of "a" to "o", we indicate whether two coefficients differ for three levels of significance (0.01, 0.05, 0.10). Of particular interest are those cases where the impact of an explanatory variable differs comparing exits to small groups with exits to large groups (categories j through o).

Table 2. Multinomial logit marginal effects: female respondents[a]

Variable	Remain with parents		Exit to live alone		Exit to live with spouse/partner		Exit to live in small group		Exit to live in large group	
	Coeff.	Std.error	Coeff.	Std.error	Coeff.	Std.error	Coeff.	Std.error	Coeff.	Std.error
Constant	1.72 d,g,j,n	(0.24)	−0.59 a	(0.13)	−0.66 a	(0.14)	−0.27 a	(0.11)	−0.21 b	(0.12)
Age	−12.11 d,g,k	(2.18)	3.90 a	(1.15)	5.11 a	(1.23)	1.71 b	(0.97)	1.39	(1.09)
Age squared	0.28 d,g,k,o	(0.05)	−0.08 a	(0.03)	−0.12 a	(0.03)	−0.04 b	(0.02)	−0.04 c	(0.03)
Black	4.09 g,n	(1.09)	0.74 g,k,m	(0.58)	−3.30 a,d,k	(0.67)	−0.05 e,h	(0.45)	−1.02 b,d	(0.48)
Hispanic	1.49	(1.33)	−0.72	(0.81)	−0.24	(0.74)	−0.13	(0.56)	−0.40	(0.51)
In-high school	19.73 e,g,j,m	(1.81)	−1.29 b,g,j,n	(0.98)	−10.27 a,d,o	(1.19)	−5.53 a,d,n	(0.83)	−2.64 a,e,i,k	(0.69)
In-college	11.40 d,g,j	(1.06)	−2.14 a,g,n	(0.57)	−8.11 a,d,j,m	(0.77)	−0.99 a,g	(0.38)	−0.16 e,g	(0.36)
Out of school-LTHS	−3.10 g,n	(1.18)	0.17 i,o	(0.71)	2.24 a,f,k	(0.60)	−0.24 h,n	(0.54)	0.94 b,f,k	(0.46)
Health limit	−6.97 e,h,n	(1.96)	2.30 b	(1.02)	2.57 b	(1.10)	0.55	(0.91)	1.56 b	(0.66)
Stepfather	−2.16	(1.53)	0.76	(0.86)	1.03	(0.90)	0.12	(0.64)	0.26	(0.56)
Single parent	1.00 n	(0.84)	0.27 o	(0.47)	−0.47 l,m	(0.50)	0.44 i	(0.35)	0.76 b,f,g	(0.33)
Number of sibling-in	0.35 f	(0.25)	−0.27 c,o	(0.15)	−0.06	(0.15)	−0.04	(0.11)	−0.03 f	(0.10)
Number of sibling-out	−0.62 g,m	(0.20)	−0.08 h,l,m	(0.12)	0.32 a,e,o	(0.11)	0.12 f,n	(0.08)	0.26 a,d,i,k	(0.07)
First born	−0.97 g,k	(0.94)	0.64 j,o	(0.52)	1.63 a,j,m	(0.53)	−1.04 b,d,g,o	(0.43)	−0.26 f,g,l	(0.40)
Number of own-children	−0.97 e,g	(1.20)	1.04 b,j,m	(0.49)	1.74 a,j,m	(0.58)	−0.89 d,g	(0.63)	−0.92 d,g	(0.71)
Catholic	2.14 h	(0.91)	−0.79 o	(0.54)	−0.99 b	(0.53)	−0.54 o	(0.39)	0.18 f,l	(0.34)
Religious attendance	0.34	(0.74)	0.35	(0.43)	−0.10	(0.42)	−0.32	(0.31)	−0.28	(0.30)
Local housing cost	−8.51 k	(6.33)	4.28	(2.99)	−1.17 l	(4.70)	5.09 b,i	(2.45)	0.30	(2.30)
Local housing cost * age	0.44 f,l	(0.29)	−0.23 c	(0.14)	0.06	(0.21)	−0.20 c	(0.12)	−0.08	(0.12)
Potential wage	−0.91	(1.33)	0.29	(0.69)	0.44	(0.90)	−0.09	(0.59)	0.27	(0.51)
Potential wage * age	−0.10	(0.06)	0.03	(0.03)	0.01	(0.04)	0.03	(0.03)	0.02	(0.02)
Parental education	−0.26	(0.15)	0.09	(0.09)	0.03	(0.09)	0.09	(0.06)	0.05	(0.06)
Local unemployment rate	0.04	(0.12)	0.01	(0.07)	0.03	(0.07)	−0.02	(0.05)	−0.06	(0.05)
Urban	0.23	(0.79)	−0.06	(0.45)	−0.73 n	(0.45)	0.20	(0.34)	0.37 h	(0.33)
AFDC eligible	0.01 d,g	(0.008)	0.01 a,g,k,n	(0.003)	−0.002 a,d,j,n	(0.005)	0.001 e,g	(0.003)	−0.001 e,h	(0.003)
South	−5.01 e,h,j	(1.16)	1.61 b	(0.69)	1.36 b,l	(0.65)	1.60 a,i,n	(0.53)	0.43 k	(0.46)
Midwest	−4.18 e,k,o	(1.19)	1.36 b	(0.72)	0.87	(0.67)	1.27 b	(0.54)	0.68 c	(0.45)
West	−8.43 d,g,j,m	(1.38)	3.25 a,i	(0.81)	1.87 a,f,l	(0.80)	1.99 a,i	(0.61)	1.32 a	(0.52)

[a] Sample size is 7,360. All coefficients and standard errors except that for the constant are $\times 10^{-2}$.
a,b,c = Significantly different from the coefficient estimate for 'remains with parents' at the 0.01, 0.05, 0.10 levels.
d,e,f = Significantly different from the coefficient estimate for 'exit to alone' at the 0.01, 0.05, 0.10 levels.
g,h,i = Significantly different from the coefficient estimate for 'exit to spouse/partner' at the 0.01, 0.05, 0.10 levels.
j,k,l = Significantly different from the coefficient estimate for 'exit to small group' at the 0.01, 0.05, 0.10 levels.
m,n,o = Significantly different from the coefficient estimate for 'exit to large group' at the 0.01, 0.05, 0.10 levels.

Table 3. Multinomial logit marginal effects: male respondents[a]

Variable	Remain with parents		Exit to live alone		Exit to live with spouse/partner		Exit to live in small group		Exit to live in large group	
	Coeff.	Std.error	Coeff.	Std.error	Coeff.	Std.error	Coeff.	Std.error	Coeff.	Std.error
Constant	1.30 d,g,k,m	(0.22)	−0.44 a,i,o	(0.14)	−0.32 a,f,k	(0.08)	−0.20 b,h,n	(0.10)	−0.33 a,f,k	(0.11)
Age	−9.55 d,g,l,m	(1.89)	2.93 a,h,n	(1.23)	2.60 a,e,k	(0.65)	1.36 c,h,n	(0.87)	2.67 a,e,k	(1.03)
Age squared	0.22 e,g,k,m	(0.04)	−0.06 b,i,m	(0.03)	−0.06 a,f,l	(0.01)	−0.03 b,i,n	(0.02)	−0.07 a,d,k	(0.02)
Black	3.50 e,g,n	(0.85)	−1.22 b,i	(0.59)	−1.08 a,f,k	(0.31)	−0.31 h,o	(0.39)	−0.89 b,l	(0.37)
Hispanic	0.37 l	(1.01)	0.34 k	(0.71)	0.40 j	(0.31)	−0.95 c,e,g	(0.52)	−0.16	(0.39)
In-high school	10.54 d,g,j,m	(1.43)	−3.41 a	(0.94)	−2.67 a	(0.75)	−2.29 a	(0.67)	−2.17 a	(0.58)
In-college	3.76 d,g,l,o	(0.88)	−1.80 a,g,m	(0.61)	−1.92 a,d,j,m	(0.45)	−0.65 c,g,m	(0.38)	0.62 c,d,g,j	(0.31)
Out of school-LTHS	0.71	(0.76)	−0.49	(0.54)	−0.10	(0.21)	−0.23	(0.36)	0.11	(0.33)
Health limit	4.48	(2.84)	0.19	(1.60)	−1.24	(1.00)	−1.55	(1.50)	−1.88	(1.69)
Stepmother	−4.63 k	(2.23)	1.54	(1.66)	0.14	(0.81)	1.87 b	(0.80)	1.08	(0.76)
Stepfather	−3.23 h,o	(1.30)	1.21	(0.94)	0.78 b	(0.41)	0.44	(0.59)	0.79 c	(0.46)
Single parent	−1.55 k,o	(0.67)	−0.08 k,n	(0.49)	−0.04 k,n	(0.20)	0.74 b,e,h	(0.31)	0.53 c,e,h	(0.28)
Number of sibling-in	0.07 n	(0.20)	−0.14 m	(0.15)	−0.08 m	(0.07)	−0.03 n	(0.10)	0.18 b,d,g,k	(0.08)
Number of sibling-out	−0.29 m	(0.15)	0.15	(0.11)	−0.02 m	(0.05)	0.01 n	(0.07)	0.16 a,g,k	(0.06)
First born	−0.37 h	(0.73)	0.39 o	(0.51)	0.44 b,l,m	(0.22)	−0.06 i	(0.34)	−0.40 f,g	(0.31)
Number of own-children	1.58 g	(2.71)	−2.69 h,k	(2.07)	1.14 a,e,n	(0.40)	0.93 e,o	(0.80)	−0.96 h,l	(1.59)
Catholic	0.84 g	(0.74)	−0.39 i	(0.54)	−0.70 a,f,j,n	(0.25)	0.16 g	(0.34)	0.09 h	(0.29)
Religious attendance	1.15	(0.60)	−0.63 i	(0.43)	0.15 f,n	(0.18)	−0.28	(0.28)	−0.39 h	(0.25)
Local housing cost	−0.64	(6.35)	3.91	(3.64)	1.17	(3.39)	−3.57	(3.56)	−0.86	(2.50)
Local housing cost * age	0.02	(0.29)	−0.12	(0.17)	−0.08	(0.15)	0.16	(0.16)	0.02	(0.12)
Potential wage	1.53 h	(0.84)	−0.54	(0.51)	−1.07 b,l,n	(0.43)	−0.11 i	(0.44)	0.20 h	(0.33)
Potential wage * age	−0.08 h	(0.04)	0.02	(0.02)	0.05 b,l,o	(0.02)	0.01 i	(0.02)	0.00 i	(0.02)
Parental education	−0.12	(0.11)	0.03	(0.08)	−0.03 k	(0.03)	0.08 h	(0.05)	0.04	(0.04)
Local unemployment rate	0.43 e,j,n	(0.11)	−0.16 b	(0.07)	−0.04	(0.03)	−0.12 a	(0.05)	−0.10 b	(0.04)
Urban	−0.82	(0.64)	0.64 l	(0.46)	0.16	(0.19)	−0.25 f,o	(0.30)	0.28 l	(0.27)
AFDC eligible	−0.04 d,g	(0.01)	0.03 a,j,m	(0.005)	0.01 a,j,m	(0.003)	0.003 d,g	(0.005)	0.000 d,g	(0.006)
South	−4.89 d,h,j,n	(0.92)	2.08 a	(0.64)	0.54 b	(0.28)	1.39 a	(0.48)	0.88 b	(0.38)
Midwest	−3.68 e,j	(0.95)	1.24 b,j	(0.68)	0.06 j	(0.28)	1.88 a,d,g,m	(0.48)	0.50 j	(0.40)
West	−5.52 e,j,m	(1.08)	1.43 b,j,n	(0.78)	0.01 j,m	(0.34)	2.60 a,d,g	(0.53)	1.48 a,e,g	(0.43)

[a] Sample size is 8,824. All coefficients and standard errors except that for the constant are $\times 10^{-2}$.
a,b,c = significantly different from the coefficient estimate for 'remains with parents' at the 0.01, 0.05, 0.10 levels.
d,e,f = significantly different from the coefficient estimate for 'exit to alone' at the 0.01, 0.05, 0.10 levels.
g,h,i = significantly different from the coefficient estimate for 'exit to spouse/partner' at the 0.01, 0.05, 0.10 levels.
j,k,l = significantly different from the coefficient estimate for 'exit to small group' at the 0.01, 0.05, 0.10 levels.
m,n,o = significantly different from the coefficient estimate for 'exit to large group' at the 0.01, 0.05, 0.10 levels.

5.1. *Economic factors*

The impact of an increased potential wage on home-leaving is generally as expected and is large. A one dollar increase in female wages reduces her annual probability of remaining with parents by 2.9 to 3.9 percentage points as age rises from 20 to 30. The results are much less strong for males. Increased wages have the effect of reducing the probability of remaining with parents once a male is age 19, but even at age 30, the impact of a one dollar increase is slightly less than one percentage point. We find no difference in the impact of a higher wage on the tendency to exit to small compared to large groups for either gender.

A higher housing cost increases the probability of a female remaining with parents once she becomes at least age 20. The probability of remaining with parents is 2.7 percentage points higher for a female age 30 residing in a locality where the cost of shelter is double the sample average compared with a female living in an area with shelter cost equal to the sample average. Her probability of living alone is 1.5 percentage points lower when shelter costs are double the sample average. For male youths, we find no statistically significant effects. Finding a response for women, but not men, agrees with the results in Ermisch and Di Salvo (1997) who studied a sample of British youths. There is no impact of housing cost on the tendency to exit to small versus large groups.

Female youths eligible for AFDC are less likely to exit to marriage/partnering. There is no difference of the impact of AFDC eligibility on the tendency to exit to large or small groups. Eligible male youths are less likely to remain with parents or exit to any group than exit to marriage or live alone.[8]

A higher local unemployment rate raises the probability of remaining with parents for male youths compared to exiting to live alone or either type of group. The effect of variations in the unemployment rate on females living arrangements is small.[9] The better educated are parents (our proxy for parental resources), the lower the chance of males exiting to marriage compared with exiting to a small group. We find no impact of parents' education on the distribution of females' exit types.[10]

5.2. *Demographic characteristics*

We find that as a youth ages, the most likely destination of the first exit from the parental home differs.[11] For men, the probability of exiting to a large group peaks at age 19.9, followed by peaks for a small group (age = 21.8), living alone (23.8), and marriage/partnering (25.4). For women, the same pattern is observed; exits peak first for a large group (age = 18.3), then a small group (20.1), followed by living alone (23.0), and marriage/partnering (23.8).

Controlling for other socioeconomic and demographic variables, Black males and females are less likely to exit to marriage or a large group and are more likely to live with parents than white youths. Black males are less likely to exit to live alone than stay with parents. These observations are consistent with the hypothesis that discrimination in the housing market reduces the tendency of Black youths to leave home because of the greater difficulty of securing shelter in any type of living arrangement. Unexpectedly, Black males are more likely to exit to small compared with large groups.

In contrast, the only significant finding for Hispanic youths is a lower tendency for males to exit to small groups than remain with their parents, live alone, or marry. This result is consistent with the lower level of housing market discrimination encountered by Hispanics compared with Blacks (Yinger 1991).

Being a high school student greatly increases the probability of remaining with parents and reduces the probability of exiting to all other living arrangements. Being a post high school student increases the probability of remaining with parents (or living in a dormitory) and reduces the probability of exiting to living alone, marriage/partnering, or a small group. Both males and females are more likely to exit to a large group compared with a small group when in college, this result expected for cost conscious college students, but the result is only statistically significant for males.

The impact of ill health differs greatly by gender. A female youth with a health problem that limits her work is less likely to remain with parents and is more likely to exit to live alone, married/partnered, or a large group. In contrast, males with a health problem are more likely to remain with parents, but the coefficient is not significant.

Male youths residing in urban areas are less likely to exit to a small group than to a large group or to live alone. Female youths in urban areas are less likely to exit to marriage/partnering and are more likely to exit to large groups. Finding that women in urban areas tend to live in large groups is expected, likely a result of seeking greater safety.

Compared to respondents living in the eastern U.S., all youths are less likely to remain with their parents. Exits to all destinations are more likely for those in the south, midwest, and west relative to those in the east, although some of the estimated effects are not statistically significant.

5.3. Family background and family structure characteristics

The impact of living with a stepmother is estimated only for males because of colinearity problems when this variable is included in the estimation for females. For males, the effect of a stepmother is to reduce the probability of remaining in the parental home by five percentage points. Living with a stepfather reduces the likelihood of remaining with parents by two percentage points for female youths and three for males. These results are as expected and are consistent with the finding of Avery et al. (1992). There are no differences in the impact on exits to small or large groups.

The number of siblings and birth order affect exit choices. First born children are more likely to exit to live with a spouse or partner than are children born second or later. They also tend to live alone rather than in a group, this expected. Increased number of siblings in the parental home reduces the likelihood of exiting to live alone (females only), and increases the probability of exiting to a large group (males only), but the impacts are small. The greater the number of siblings living outside the parental home the greater the likelihood of exiting to marriage/partnering (females), and group arrangements. As expected, the largest effect of having more siblings outside the parental home is to increase the probability of exiting to a large group compared with small groups.

An increased number of respondent's own-children reduces the tendency of a female youth to remain in the parental home or exit to live in a small or large group. Males with own-children are unlikely to exit to live alone or to large groups compared with remaining with parents or exiting to marriage or small groups.

Individuals raised as Catholics are more likely to remain with parents and are less likely to exit to live alone or with a spouse/partner compared with exiting to large groups. Catholic females are less likely to exit to a small group than a large group. We are surprised by the lower probability of exiting to the marriage destination; however, this category includes living with a partner. Thus, the negative marginal impact of being Catholic on exiting to marriage/partnering may result from the dominant effect of a much lower rate of exiting to live with a partner. Males who attend church frequently are more likely to exit the parent's home to marry than to live in a large group or live alone.

The impact of the explanatory variables is shown in greater detail in Table 4. We establish a base case by applying a set of assumptions to a 16 year old white youth.[12] We follow the youth for 13 years and cumulate the probability distribution of possible exit types. The table lists the cumulative distributions for the base case and many variations. In the base case for men, the most likely exit is to marriage/partner (39%), followed by living alone (34%), exiting to a small group (13%), and to a large group (11%). There is only a 3% chance that this male youth will continuously live at home through age 29. For women, the distribution of exit probabilities differs: 49% to marriage/partner, 21% to small groups, 19% to alone, 9% to large groups, and 2% remain in the parental home.

Variations in the base case include increasing the house price by one standard deviation, reducing the unemployment rate by 25%, increasing parents' education to 16 years, and raising the youth's mental ability score by one standard deviation. Other variations include changing the youth's race/ethnicity to Black or Hispanic, having a child at age 18, being a college graduate, and having a stepfather or stepmother.

We also report the expected duration of stay in the parental home, this value equaling 4.9 years for males and 4.1 for females in the base case. The most notable changes in the expected duration of stay with parents occur when a youth's education is increased to 16 years from 12 (stay with parents for one year longer), when a youth is Black (stay with parents for one year longer), or when a male lives with a stepparent (stay with parents is nearly one year shorter).

Reading across a row reveals the impact of a change in an explanatory variable on that exit type. In general, the effects are consistent with the previous discussion. The table also reveals the overall size of the impact on the distribution of exits. The biggest effects, arbitrarily measured as the sum of absolute values of deviations from the base case probabilities, are for a youth who has a child at age 18, being Black, completing college, having a stepfather, and living in a high house price locality.

6. Conclusions

In recent years, spurred by the availability of longitudinal data, significant progress has occurred in understanding the factors that explain when youths leave their parental homes. Our study complements those analyses by adding detail to the list of potential destinations and by adding to the list of

Table 4. Impacts of variations in selected explanatory variables on a 16 year old white youth

Variable	Base case[a]	House price + 1 s.d.[b]	Mental ability + 1 s.d.	Unemployment rate 8% ⇒ 6%	Black	Hispanic	Youth has child when 18	Stepfather	Stepmother	Youth: college grad	Parents college grads
Male youths											
At home	0.03	0.02	0.03	0.02	0.10	0.02	0.01	0.01	0.01	0.02	0.03
Alone	0.34	0.38	0.34	0.35	0.35	0.35	0.17	0.34	0.36	0.32	0.35
Marriage	0.39	0.41	0.39	0.38	0.33	0.46	0.59	0.40	0.26	0.34	0.36
Small	0.13	0.10	0.13	0.14	0.15	0.08	0.17	0.12	0.23	0.13	0.15
Large	0.11	0.09	0.11	0.11	0.07	0.09	0.06	0.13	0.15	0.18	0.12
Mean duration[c]	4.9	4.8	5.0	4.8	6.0	4.9	4.7	4.2	4.0	6.0	5.0
Female youths											
At home	0.02	0.01	0.01	0.02	0.06	0.02	0.01	0.01		0.01	0.01
Alone	0.19	0.21	0.19	0.19	0.31	0.17	0.21	0.20		0.24	0.19
Marriage	0.49	0.42	0.48	0.49	0.35	0.51	0.60	0.51		0.35	0.47
Small	0.21	0.28	0.22	0.22	0.22	0.22	0.13	0.19		0.26	0.23
Large	0.09	0.08	0.10	0.10	0.06	0.08	0.05	0.09		0.14	0.09
Mean duration[c]	4.1	3.9	3.9	4.1	5.0	4.2	4.0	3.8		5.0	3.9

[a] The base case is defined in footnote 12.
[b] 1 s.d. indicates a change of one standard deviation.
[c] The value is the expected duration in the parental home in years.

explanatory factors. Our national sample is of American youths ages 16 to 30 in the period 1979–1992. We focus on exits to small and large groups, with about one-third of all exits being to groups.

We comprehensively model economic and socio-demographic factors expected to influence a youth's decision of whether to leave the parental home and what living arrangement to select. We highlight economic factors hypothesized to impact home-leaving and test whether the influence of these factors varies with a respondent's age.The estimation technique is a multinomial logit analysis.

Our first finding is that male and female home-leaving must be modeled separately. Gender specific differences in exit tendencies include responses to variations in the wage that could be earned in full-time employment, the local unemployment rate, the local cost of shelter, race and ethnicity, health problems, and residence in an urban area.

We test for and find some evidence of an age linked impact of economic factors. Our hypothesis is that parents shelter teens from economic factors such as a high cost of shelter and low potential income; however, this sheltering is reduced as the children age. Higher wages reduce the likelihood of staying with parents for youths older than 20. Higher housing costs increase the probability of staying with parents for women older than 20. No significant effect is found for men.

We hypothesized that socio-demographic factors dominate explanations of why youths leave the parental home when they are teens, but economic factors increase in influence as a youth matures. Our results support this hypothesis most strongly for female youths. A possible modification of our hypothesis is to argue that parents shelter their daughters from economic factors more so than they do their sons. Similar age linked responses to variations in parental income and AFDC payments were hypothesized by Whittington and Peters (1996), but we do not find support for their hypotheses.

Our results for housing costs and wage impacts on home-leaving are not as strong as in Haurin et al. (1993). One reason is that this paper's sample includes young teens while Haurin et al. (1993) did not. Another substantial difference is that Haurin et al. studied the current living arrangement of all youths while we analyze only their first exit destination. It may be that returns to the parental home are influenced by housing costs and earnings. If true, these variables would affect the results in Haurin et al. but not this study. Ermisch (1999) finds strong support for the argument that high housing costs and low incomes increase the rate of return to the parental home.

While understanding the likelihood of exiting to living alone or marriage is of interest, we highlight the study of exits to small and large groups. Exiting to a large group tends to be the first path out of the parental home for teens, followed by exiting to a small group. Table 5 summarizes the statistically significant impacts of the explanatory variables on youths' exits from the parents' home to large compared with small groups. Some results are as expected; for example, youths with a larger number of siblings living outside the parental home tend to exit to a large group (which may include some siblings). Also, college students have a greater tendency to exit to live in large groups. Black youths are more likely to exit to small groups than large, the unexpected effect statistically significant only for males. Youths dropping out of high school, Catholics, and youths in urban areas tend to exit to large groups.

We find that the determinants of whether youths exit the parents' home

Table 5. Significant impacts on the choice of exiting to small or large groups[a]

Group Size	Females		Males	
	Small	Large	Small	Large
Black			×	
Age	×		×	
Number of siblings-out		×		×
Number of siblings-in				×
First born		×		
Out of school-LTHS		×		
In high school		×		
In college				×
Number of own-children			×	
Catholic		×		
Urban				×
Midwest			×	
South	×			

[a] An '×' indicates that the marginal effect of this variable for this category (small group or large group) is significantly greater than the marginal effect for the other category (large group or small group) at least at the 0.10 level.

to large or small groups are dominated by socio-demographic factors, with economic factors being unimportant. The lack of significance of any of the economic variables in explaining the size of the group selected by exiting youths is surprising because economic variables are important in explaining whether youths remain in the parental home or leave. We believe the reason is that exits to groups typically occur at younger ages (teens and very early twenties), and economic factors tend to become important later when a youth is in his or her early to mid twenties.

Appendix: Potential wage estimation procedure

A potential full-time wage is estimated for each respondent for each year using a two-step Heckman procedure (Greene 1995; Heckman 1979). This wage represents the resources available to the respondent if she or he chooses to work full-time; thus, it must be estimated for those respondents not observed working full-time. The first step of the procedure requires estimating a probit model of full-time work status. We estimate identical models for males and females using a sample of respondents age 16 or older: 24,408 observations for males and 21,855 for females. The dependent variable is an indicator of full-time work status defined as 1600 or more hours worked during the calendar year. Explanatory terms include: age, age squared, Black, Hispanic, not a high school graduate, attended college but did not graduate, graduated college, the local unemployment rate, health limit, in-school, age of spouse and its square, number of children age 0 to 6, number of children age 7 to 17, and urban. Results of these regressions are available from the authors. They correspond well with our expectations with outcomes correctly predicted for 76% of males and 71% of females.

The second step is to estimate a wage equation for those respondents working full-time. For male respondents, full-time work was performed in 14,404 (59.0%) of the person-years. Females worked full-time in 10,599 (48.5%) of the person-years. The dependent variable is a continuous measure of the hourly rate of pay the respondent received in his or her current job. Preliminary exploration with the dependent variable transformed into logs yielded inferior results.

Explanatory terms included in the wage equation are age, age squared, Black, Hispanic, not a high school graduate, attended college but did not graduate, graduated college, urban, health limit, the local unemployment rate, job tenure, a measure of the respondent's mental ability, a series of 13 year specific dummy variables, and the sample selection correction variable (inverse Mills ratio) generated from the full-time work status equation. Results for these regressions are available from the authors. The equations for males and females are generally similar. Significant and negative is Black (4% lower wage for males and 6% lower wage for females) and the local unemployment rate. Significant and positive are attended college but did not graduate (a 4%–8% higher wage) and graduated college (a 25%–33% higher wage). Also significant is urban (a 12%–14% higher wage) and the score on the mental ability test (a one standard deviation increase raises wage by 10%–11%). In both cases the potential wage rises with respondent's age; for example, as age rises from 20 to 30, potential wages for males and females rise by 15% and 18% respectively.

The variable included in the multinomial logit is the predicted value from the wage equation. Over time, it is updated as a respondent's age and other characteristics change.

Endnotes

[1] We do not focus on exits to institutional group arrangements, but should clarify how we categorize exits to college dormitories and military. In the literature, exits to institutions have been considered separate destinations (Goldscheider and DaVanzo 1985, 1989), or have been undifferentiated from parental living arrangements (Buck and Scott 1993). Whittington and Peters (1996) generally treat college students as residing with parents (unless self-supporting), but combine those in the military with those living alone in noninstitutional arrangements. Our study classifies college students and military members according to their current residence. We consider youths residing in dormitories, fraternities, or barracks to be in temporary quarters and classify them as not exiting the parental home. College students and military members living elsewhere are distributed among our destination categories (alone, married/partnered, or in a small or large group). Our conceptualization reflects a youth's choice to reside in a particular living arrangement rather than the choice to enter college or the military.

[2] The critical Chi-squared value at the 0.01 level with 25 degrees of freedom is 45. Twice the difference in log likelihood values comparing the pooled model to the sum of the gender specific models is 126. These results clearly reject pooling the male and female samples.

[3] In the sample of 26,553 youths, missing values occur most often for wages, housing costs, and the measure of the mother's intellect (12%, 11%, and 15% of the sample, respectively).

[4] Dwelling cost is derived from the Freddie Mac-Fannie Mae (FF) repeat sales house price index, augmented by data from the American Chamber of Commerce Research Association (ACCRA 1993). The FF index covers more than 100 MSAs and all states and is a pure time series price index. We use the 1982 ACCRA data for 88 MSAs and rural areas in 50 states to develop a baseline cross-sectional price index. The final index is developed by applying the FF index to the ACCRA data, yielding a nominal house price index with excellent spatial coverage.

[5] We experimented with age thresholds other than 18 but found 18 yields the best fit.

[6] A respondent is eligible if there is at least one own-child living in the household and the respondent is not married. The maximum benefit is the amount of assistance a family would receive if it had no income (Committee on Ways and Means 1996). We select this measure because it is exogenous; that is, it is invariant to the respondent's choices of participation in the AFDC program and supply of labor.

[7] We do not report multinomial estimation coefficients because their values are difficult to interpret; however, they are available from the authors.

[8] AFDC Eligible was interacted with Age 18, but the estimated coefficient is not significant.

[9] We found no impact of including a variable interacting Unemployment with Age 18.

[10] We tested for differing impacts depending on the age of the respondent, but failed to find supportive evidence.

[11] These calculations include the effects of changing age, age squared, and the age interaction terms with house price and wage.

[12] The assumptions include: the youth stay in school until age 18 and leaves with a high school degree. The youth lives in an urban area in the southern U.S. where the housing cost equals the sample's mean and the unemployment rate is eight percent. The youth's parents have 12 years of education. The youth has good health, no children, no stepmother or stepfather, is not Catholic but attends church at least once per month, is not first born, has two resident parents, and is not AFDC eligible. The youth has one sibling in the parental home and one out until age 24 when both siblings are out of the parent's home. Wages also are set by the base case assumptions and vary as we change the values of explanatory variables.

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Residential location and youth unemployment: The economic geography of school-to-work transitions

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Abstract. In response to increased international policy attention to youth unemployment this study investigates post-secondary school transitions of school leavers. Multinomial logit models are estimated for male and female German youth. The models control for individual, parent, and household characteristics, for those of the youth's region of residence and local labor markets. The findings suggest that immigrant youth has particularly low participation rates in continued education, and that youth unemployment is centered in high unemployment states and metropolitan areas. More generous academic benefit policies seem to be correlated with increased academic enrollment, and men's transitions to the military do reflect recent changes in defense policies.

JEL Codes: J24, J64, J68

Key words: School-to-work, youth unemployment, local labor markets

1. Introduction

As of 1998 youth unemployment in the European Union was at 19.1% of the youth labor force. The literature provides ample evidence on the lifetime scars early unemployment experiences leave on workers' labor market and criminal records (Ellwood 1982; Freeman and Rodgers 1999; for Germany Franz et al. 1997). This dramatic situation prompted government responses in several

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countries. France launched an ambitious youth unemployment program in 1997, which as of November 1998 covered 152,000 persons aiming at 350,000 by 2000 (OECD 1999a). The German government passed a "100,000 Jobs for Youth" program, with DM 2 bio allocated for the fiscal years 1999 and 2000 each.

With youth unemployment high on the policy agenda, it is important to understand its determinants and the potential role for labor market policies. So far, only a few studies investigated the labor market transitions of school leavers. The issue was discussed in the United Kingdom due to a concern about declining participation in continued education. The German case found attention in the United States following the Clinton administration's suggestion to install some features of the German apprenticeship system there (Clinton and Gore 1992). Overall, existing studies are highly specific in their topics. Only few attempt to answer the broader questions of what young people do after leaving school, how their choices are affected by regional labor markets and by policy changes, and to what degree parental and household characteristics affect these transitions.

This study addresses these important issues at the example of Germany. It applies a comprehensive modelling approach to investigate the transition choices of all youth leaving school between 1984 and 1997, as observed in the German Socioeconomic Panel (GSOEP). This dataset permits the consideration of detailed household and parent background variables, which were omitted in prior analyses. The findings are relevant to the evaluation of the German government's "100,000 Jobs for Youth" program, as one of the criticisms of this program relates to the equal spreading of program activities across labor market regions. To the degree that local factors are important for the success of youth labor market entry, regional differentiations in active labor market policies may be required. Finally, the study evaluates whether changes in training grant and military policies affected school-to-work transition patterns.

The paper proceeds with a summary of the German institutional framework for school-to-work transitions, and of policy changes, that may have affected transition decisions. It describes the German youth labor market, the main features of the "100,000 Jobs for Youth" program, and briefly surveys the school-to-work literature in Sect. 2. Section 3 discusses the econometric specification, presents the data and the estimation strategy. The results are discussed in Sect. 4. The study concludes with a summary, highlighting policy implications.

2. Institutional background and review of the literature

2.1. School-to-work in Germany

In contrast to other countries, the German school system introduces differentiated educational tracks after the first four grades of primary education. The tracks differ in their academic orientation and requirements. The basic school (*Hauptschule*) graduates individuals after six years of secondary education and is a preparation for blue collar occupations. The middle school (*Realschule*) also lasts six years and provides training for white collar jobs. Only the highest track (*Gymnasium*) provides another nine years of schooling. Graduating from the *Gymnasium* is a precondition for university studies. In addition, comprehensive schools (*Gesamtschule*) were introduced in the 1970s, which grant degrees of either track. Depending on the track, pupils typically finish school aged 16 or 19. Of the 1.1 mio school leavers in 1997 7% had not obtained a degree, 25% graduated from basic school, 38% from middle school, and 22% from the *Gymnasium* (with the rest in the "other" category).

Once they leave school, individuals can choose from a number of alternative paths. This choice is restricted only for healthy young men above age 18, who are typically drafted for military or conscientious objector service. The most common transition after school is that into apprenticeship training. Apprenticeships last between 2 and 4 years, and combine vocational on the job training with formal education in vocational schools (*Berufsschule*).

Particularly in the mid 1980s, when the German baby-boom generation left school, insufficient apprenticeship positions were available for school leavers. For them and to provide training for certain occupations without apprenticeship programs, vocational schools are available to meet excess demand for vocational training. These offer (i) fulltime general schooling for those not previously qualified for apprenticeships in a one year 'vocational preparation year' (*Berufsvorbereitungsjahr*) program. Here individuals can complete their basic school degree (*Hauptschulabschluß*). In a program (ii) called 'elementary vocational year' (*Berufsgrundbildungsjahr*) students may learn occupation-specific skills which – if successfully completed – allows them to shorten a later apprenticehip. The third type (iii) labelled 'special vocational school' (*Berufsfachschule*) offers a variety of training opportunities. More than fifty percent of the students graduating from these three vocational schools continue their education with an apprenticeship.

In addition to the military, apprenticeship, and vocational school options, graduates may choose employment without training, they may leave the labor force, become unemployed or begin an academic education. A policy affecting the decision to take up academic training is the financial support program *Bafög*. Since 1971 benefits are available to children of non-wealthy parents, who pursue an academic education.[1] The law has been changed repeatedly, with the most influential adjustments in 1983, when the program switched from grants to loans, and again in 1990 when it was stipulated that only half of the program benefits had to be repaid. Figure 1 shows that after 1983 the share of school leavers going on to university declined. However, the fraction of university students increased again in subsequent years with a steep jump in 1990, when the new grant system was passed.

A set of policies likely to affect males' transition decision relates to the military draft. Most influential here is the leniency in the requirements for physical fitness and the flexibility of postponing the draft to complete vocational and academic training. The evidence suggests that prior to the end of the cold war in 1990 the draft was rather strict. The size of the German army declined from half a million in the mid 1980s (West Germany) to about 350,000 soldiers (united Germany), as determined in the unification treaties after 1990 (Rotte 1996). At that time drafting procedures were loosened and the duration of military service was cut from 15 to 12 months in 1991. The treatment of draftees was tightened again recently (since 1995) after the military engagements in Bosnia and Kosovo had caused an increase in the share of conscienious objectors (BREG 1996). The impact of these developments on individual transitions is analysed below.

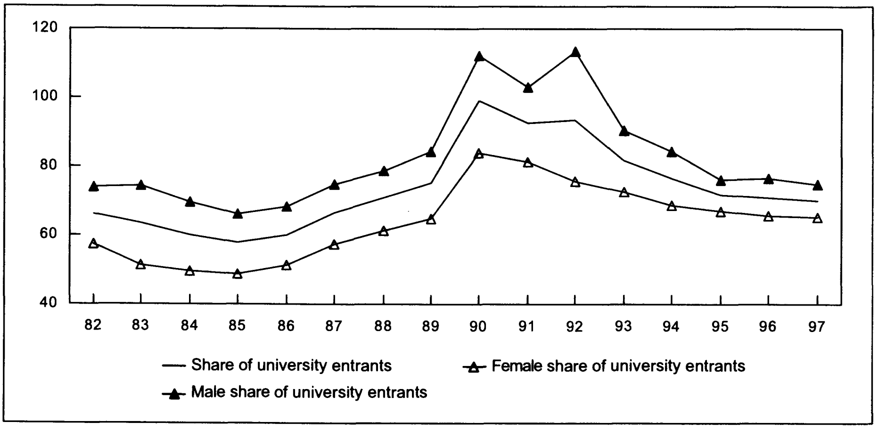

Fig. 1. New university students as share of *Gymnasium* graduates (in percent). *Source:* Statistisches Bundesamt, *Statistical Yearbook of the Federal Republic of Germany*, various years. *Note:* (1) The ratios do *not* depict the share of *Gymnasium* graduates moving on immediately to academic training. Instead the number of new university students (as of the winter semester) is divided by the number of *Gymnasium* graduates of the same year. Since military service or other vocational training might intervene between leaving school and taking up university studies in some years the ratio can take on values above one. (2) Through 1990 the figures represent West Germany only, after 1991 they represent united Germany

2.2. Youth unemployment in an international perspective

By international comparison youth unemployment in Germany is moderate. Table 1 describes the unemployment rates by agegroups across countries for 1990 and 1998. By OECD definitions Germany had very low overall and youth unemployment in 1990, and even in 1998, with higher overall unemployment, youth unemployment rates were relatively modest. Table 2 presents the ratio of youth unemployment to total unemployment across countries, and confirms the comparatively positive situation for German youth.

Figures 2 and 3 present the development of east and west German unemployment rates over the last decades. Youth unemployment rates in the 20 to 24 agegroup almost permanently exceed overall unemployment and that of the younger agegroup, with a particularly striking difference in East Germany.[2] When West German baby boomers flooded the labor market in the early eighties, their unemployment was acutely above the overall average. Since 1987 unemployment among the very young remained below average and those above age 20 slightly exceed overall unemployment rates.

In order to gauge whether regional differences play a role in youth unemployment, Fig. 4 presents the ratio of youth to overall unemployment rates, averaged across two types of west German states: The states of Berlin, Hamburg, and Bremen contain almost exclusively urban regions, whereas the others[3] ('area states') combine urban and rural areas. Averaging relative unemployment rate ratios across city and area states shows that for the last two decades relative youth unemployment has been a more pressing problem in cities than in rural areas.

The increase in youth unemployment rates since 1990 has prompted the government in 1998 to install a "100,000 Jobs for Youths" program. This

Table 1. Unemployment by agegroup, country, year, and sex

Country	1990 Unemployment rates				1998 Unemployment rates			
	All (15–64)	Youth (age 15–24)			All (15–64)	Youth (age 15–24)		
		All	Men	Women		All	Men	Women
France	9.2	19.1	15.3	23.9	11.9	25.4	21.9	30.0
Germany	4.9	4.6	4.4	4.7	8.6	9.4	10.4	8.2
Italy	9.9	28.9	23.4	35.4	12.2	32.1	28.1	37.2
Netherlands	7.4	11.1	10.0	12.3	4.3	8.2	7.8	8.7
Spain	16.1	30.1	23.2	39.7	18.8	34.1	27.1	43.4
United Kingdom	6.8	10.1	11.1	9.0	6.2	12.3	13.8	10.5
United States	5.7	11.2	11.6	10.7	4.5	10.4	11.1	9.8
European Union	8.1	15.7	13.5	18.1	9.9	19.1	17.6	20.8
Total OECD	5.9	11.5	11.1	12.1	6.8	12.8	12.5	13.1

Source: OECD Employment Outlook June 1999.

Table 2. Youth unemployment relative to total unemployment by country, year, and sex

Country	1990			1998		
	All	Men	Women	All	Men	Women
France	2.08	1.66	2.60	2.13	1.84	2.52
Germany	0.94	0.90	0.96	1.09	1.21	0.95
Italy	2.92	2.36	3.58	2.63	2.30	3.05
Netherlands	1.50	1.35	1.66	1.91	1.81	2.02
Spain	1.87	1.44	2.47	1.81	1.44	2.31
United Kingdom	1.49	1.63	1.32	1.98	2.23	1.69
United States	1.97	2.04	1.88	2.31	2.47	2.18
European Union	1.94	1.67	2.24	1.93	1.78	2.10
Total OECD	1.95	1.88	2.05	1.88	1.84	1.93

Note: Figures present youth unemployment rate (age 15–24) relative to overall unemployment rate across all agegoups, both as presented in Table A. Values smaller than 1 indicate that youth unemployment is below average unemployment, values bigger than 1 describe the reverse situation. – German figures for 1990 refer to West Germany, for 1998 to East and West Germany.
Source: Own calculations based on OECD Employment Outlook June 1999.

program (a) provides subsidies to firms who offer apprenticeship positions to unemployed youth, (b) offers training programs through regional employment offices, and (c) provides funding for various types of vocational training. It started in 1999, covers individuals up to age 25, and intends to focus on east Germany, were unemployment is particularly trenchant. While the program had reached more than 170,000 youths by August 1999, the effectiveness of the measures is difficult to judge so soon after initiation. Critics point out that employers free ride on employment subsidies, that the program is run even in areas where unemployment is low, and that it only postpones young people's unemployment. Most recently the OECD (1999b) pointed to the often discouraging results of yout labor market policies, but lists success stories, as well. Clearly, youth unemployment is high on the agenda of policy debates.

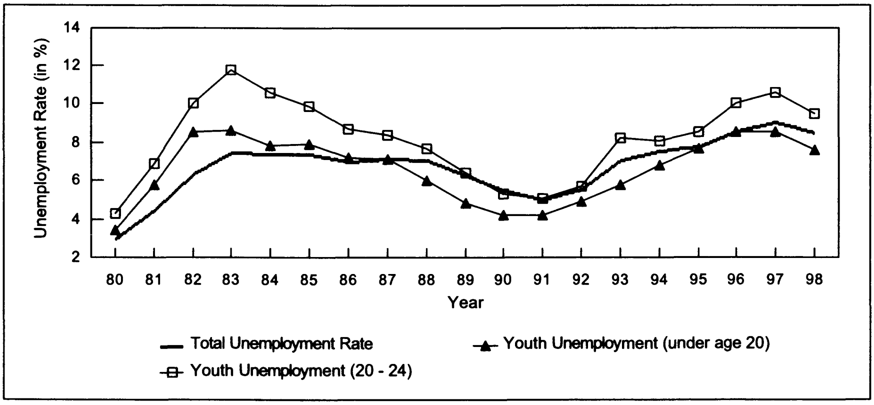

Fig. 2. West German unemployment rates by agegroup. *Source:* see Figure 3

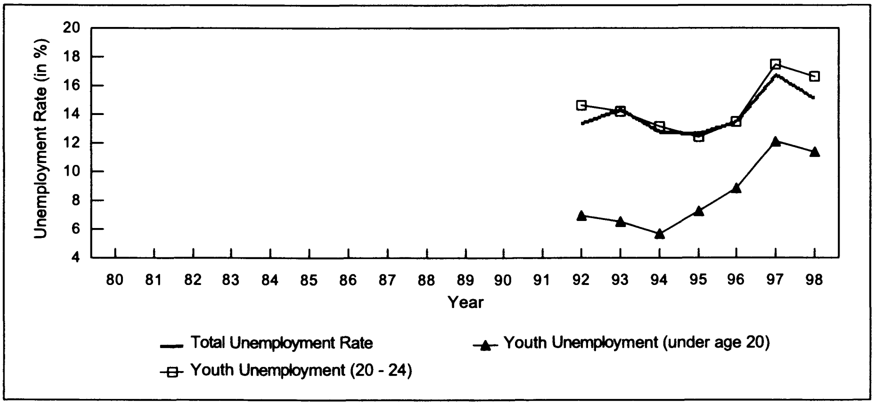

Fig. 3. East German unemployment rates by agegroup. *Source:* Own calculations based on the number of registered unemployed as of September (Bundesanstalt für Arbeit, *Strukturanalyse 1993* and Bundesanstalt für Arbiet, *Strukturanalyse 1998*), and the labor force as of April (Statistisches Bundesamt, Fachserie 1, Reihe 4.1.1)

2.3. The school-to-work literature

The literature on school-to-work transitions falls in a British, a German, and an internationally oriented tradition. In the latter category OECD (1998) points to educational attainment, labor market tightness, and institutional settings as determinants of the employment probability of school leavers. McIntosh (1998) investigates the changes in post-secondary education participation in four countries and finds that changes in initial academic attainment and expected returns to schooling are decisive.

The British literature is concerned with post-compulsory education choices of youth and its determinants. The contributions are distinguished either by rich and comprehensive datsets, or by specialized focuses: Rice (1987) looks at the effects of household income, and discusses the introduction of education subsidies. Micklewright (1989) confirms that family background remains in-

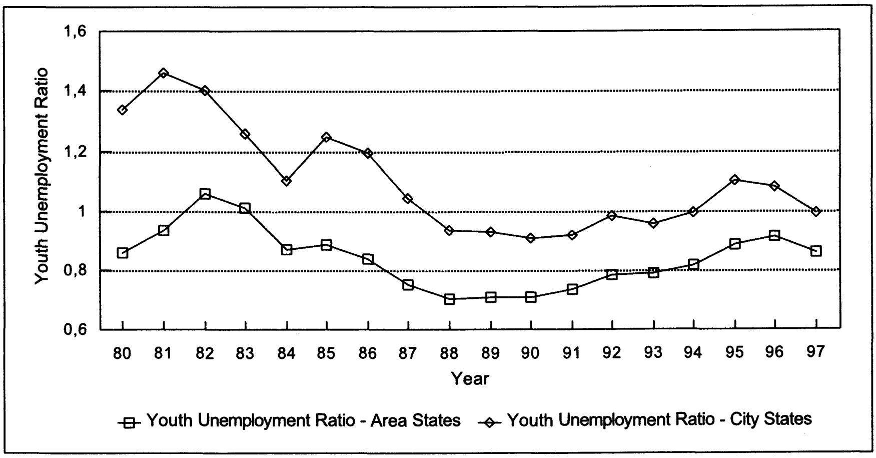

Fig. 4. Averaged youth unemployment (under age 20) relative to total unemployment by area and city states. *Source:* Own calculations based on state unemployment rates (Bundesanstalt für Arbeit, *Amtliche Nachrichten der Bundesanstalt für Arbeit: Jahreszahlen 1988* and Bundesanstalt für Arbeit, *Amtliche Nachrichten der Bundesanstalt für Arbeit: Jahreszahlen 1997*)

fluential even when detailed controls for student ability are considered. Whitfield and Wilson (1991) interprete aggregate time series evidence to support the introduction of special employment and training measures for youth. Rice (1999) confirms the relevance of individual academic attainment, as well as of family and social background. She emphasizes the relevance of labor market tightness for educational decisions of low ability males. Leslie and Drinkwater (1999) focus on ethnic minorities and find that their high participation in continued education is related to expected increases in future benefits and to fewer current opportunities. Dustmann et al. (1998) investigate the effects of school quality on the level of education achieved. The study most closely related to the approach followed below is Andrews and Bradley (1997), which models a set of transitions open to school leavers using a multinomial logit estimator. They confirm the relevance of local labor market variables, and show that the characteristics of the previous school are influential for transitions.

In contrast to the British studies, German studies of school-to-work transitions addressed broader sets of issues, including institutional design and cross-national comparisons of educational outcomes (Buechtemann et al. 1994; Gitter and Scheuer 1997; Lindner 1998; Winkelmann 1996). The most recent German empirical school-to-work studies focus on graduates of apprenticeship programs: Franz et al. (1997) (abbreviated F), Inkmann et al. (1998) (I), and Franz and Zimmermann (1999) (FZ) investigate possible problems in the transition from apprenticeship training to work. The research questions concern the duration of non-employment (F) or employment spells (FZ) following apprenticeship training, the probability of being hired by the firm which provided the apprenticeship training (FZ), as well as earnings effects of failure during the apprenticeship program or in the transition to employment (F, I). Only Merz and Schimmelpfennig (1999) (MS) look at the group of those who graduated from *Gymnasium* and who choose between academic and vocational training. The authors test whether skill-specific wage differentials and unemployment rates affect transition decisions. The main findings of this empirical literature are as follows: (i) Failures during or after vocational training cause long term harm to earnings trajectories. (ii) Early unemployment spells are more detrimental for long run earnings than failures in completing apprenticeships. (iii) The higher aggregate unemployment, the shorter the expected duration of first employment spells after vocational training. (iv) MS find that particularly male graduates' career choices respond to expected returns.

Overall, the issues addressed in this literature are specific to certain groups of graduates. No analysis has studied the role of regional effects for individual behavior, and only FZ and MS considered the effects of *aggregate* labor market tightness at all. The contribution of this study is to comprehensively investigate school-leavers' transition decisions and to fill the gap in the literature regarding the role of local labor markets and regional effects.

3. Specification, data and econometric approach

3.1. The empirical model

Most studies on the transition decision of school leavers refer to human capital theory, where investments in education follow a comparison of expected costs and benefits. Each school leaver chooses that transition option (e.g. employment or apprenticeship), which generates the highest expected lifetime utility. The literature distinguishes three factor groups that may affect utility directly and through their role in the school-to-work transition decisions: factors relating to the individual (I), to social and family background (F), and to characteristics of the local labor market (L) as determinants of the probability that individual i chooses a transition to state j over its alternatives (the set of alternatives is discussed below):

$$\mathrm{Pr}_i(\text{transition into state } j) = f(I_i, F_i, L_i)$$

Individual characteristics (I) account for the individuals' age at the time of completing secondary education, their sex, nationality, health, and educational attainment. The type of secondary school in combination with age measures individual ability, which is expectedly higher for *Gymnasium* and middle school graduates than for basic school graduates, and for a given degree should be higher for those who graduate at an earlier age. While the British literature showed that the probability of continued education is higher for females than for males, such a prediction does not seem warranted for Germany: Given the young age of school leavers, labor force exits for family reasons should be infrequent. Since ethnic background might affect educational choice – be it through cultural or human capital differences, such as language – foreign origin ought to be considered. Finally, handicapped youth might be particularly disadvantaged and face limited opportunities for vocational training.

The studies reviewed above provide ample evidence for the relevance of *family background* in educational transitions. The structure of a household might indicate the degree of financial restrictions school leavers face in their transition decision, e.g. if they live in a single parent versus two parent household. Since Micklewright (1989) and Andrews and Bradley (1997) show that the number of younger siblings affects youth educational choices, this indicator is considered as well. Parental background is likely to influence tastes and preferences of the offspring as well as their academic attainment in secondary school. Therefore we control for the impact of parental education, ethnic background, and labor force status.

Finally, *regional and labor market indicators* may be important factors in school leavers' transition decisions. These variables have been neglected in German studies, but the international literature speaks to their relevance.[4] The measures considered here describe the individuals' regions of residence by community size, the local unemployment rate, and an indicator of whether it is in East or West Germany. The underlying hypothesis is that the size and structure of the local labor market affects youth behavior. Clearly, perfect mobility would render such effects irrelevant, but particularly for young school leavers financial constraints may restrict mobility. We expect youth to be more likely to find a job or an apprenticeship the larger the local labor market and the more extensive labor demand. The same rationale applies to transitions to academic training. Whereas youth living in a major town might have the opportunity to study while living at home, academic training may be much more costly for those living in the countryside. In addition to regional indicators the consideration of unemployment rates permits more direct controls for labor market effects.

3.2. The data

The data is taken from the first 14 waves of the German Socioeconomic Panel (GSOEP, 1984–1997). The sample is restricted to individuals aged 15 to 25. Observations are censored when the individual does not respond to a survey, when measures on core variables such as age, sex, or household identifiers are missing, or when the interview was incompletely conducted.[5]

The dependent variable describes the individuals' training or employment status at the annual interview in the year after leaving secondary school. To generate this variable, one first has to determine in which year the person left school. There are two ways to determine schooling status within the GSOEP: The *first* is based on a survey question where individuals above age 16 are asked about current participation in schooling or training activities and about the type of schooling currently pursued. We consider an individual as in school if either *Grundschule*, *Sonderschule*, *Hauptschule*, *Realschule*, *Gymnasium* or *Gesamtschule* are indicated.[6] For individuals under age 16 information on current schooling is gathered from the household head. A school exit is coded if an individual was in school in one year and out of school in the next year. The *second* way to determine schooling status is based on a question, which asks individuals whether they obtained a degree in the preceding calendar year. 68% of all school exits are identified by both indicators. In cases where the coding procedure resulted in more than one exit from secondary school, only the exit indicated last was considered, assuming that the intermittent failure to indicate school attendance resulted from measurement error.[7]

A wide set of alternative transitions is available for school leavers. It ranges from continued schooling in vocational schools, polytechnical schools, and universities over immediate employment, apprenticeship, military or substitute service for men, to unemployment, or out of the labor force spells. In contrast to the recent German school-to-work literature a comprehensive set of alternatives is captured in the dependent variable here, where all school leavers are considered and transitions into the following activities are modelled:

(1) fulltime or parttime employment without training,
(2) apprenticeship,
(3) vocational training without employment,
(4) university or polytechnical schools,
(5) military or substitute service for men,
(6) unemployment or out of the labor force (OLF).

Military and substitute service are considered as endogenous outcomes, because it is possible to influence the timing of these activities. While unemployment and OLF are typically separate outcomes, they are combined, because school leavers have no claim to unemployment benefits and because unemployed seekers of apprenticeship positions are not defined as unemployed.

The distribution of the 2,702 school leavers across these states is presented in Table 3 by various characteristics. Almost half of all graduates take up an apprenticeship, and 22% seek vocational training through the alternatives offered in the vocational schooling system. The shares of individuals in non-training employment, in academic training, and military/substitute service are below ten percent each. A substantial fraction of about 12% of all school leavers is either unemployed or out of the labor force one year after exiting secondary school.

The main difference between the two sexes lies in the share of immediate transitions into academic training, which is higher among females because male *Gymnasium* leavers are typically drafted immediately. More surprisingly, the share of nonemployed females clearly exceeds the sample average. The comparison between East and West German transitions yields a substantially higher coverage of East German youth with training programs: Jointly 76% of East German school leavers are in vocational training, compared to 66% in the West. This is balanced by a higher share of West German youth in nonemployment one year after finishing school. While the latter outcome seems surprising in view of higher overall unemployment in East vs. West Germany, Figs. 2 and 3 do suggest that the difference in unemployment rates for youth under age 20 is indeed minor.

The comparison of transitions by national origin shows substantive differences, in that immigrant youth have much higher probabilities of immediate employment and nonemployment than native youth. Panel B of Table 3 describes transition distributions by community size, health, and regional unemployment. Most striking are the high nonemployment rates in large communities, and the sensitivity to state unemployment rates. Health does not appear to be strongly correlated with transition decisions. Panel C shows the distribution of transitions by type of school and agegroup. Nonemployment rates are highest for those in 'other' schools. Transition into military or substitute service is highest among *Gymnasium* graduates because they reach draft age when leaving school. This difference by age is also reflected in the last two columns of panel C. Panel D presents the distribution of destination states

Table 3. Distribution of school leavers across destination states by characteristics (in percent)

A	All	Males	Females	East Germans	West Germans	Natives	Immigrants
Employed	6.1	6.0	6.2	2.0	6.7	4.9	13.1
Apprenticeship	45.6	46.7	44.4	62.4	43.1	46.9	37.8
Vocational training	21.8	20.1	23.6	13.8	23.0	21.7	22.4
Academic training	8.2	6.3	10.1	7.6	8.3	8.7	5.0
Military	6.3	12.3	–	7.1	6.2	7.3	1.0
Not employed	12.0	8.6	15.7	7.1	12.8	10.5	20.7
Number of obs.	2,702	1,388	1,314	354	2,348	2,305	397

B	All	Community size		Health		State unemployment	
		Small	Large	Good	Poor	Low	High
Employed	6.1	6.1	6.0	5.8	7.3	6.9	5.3
Apprenticeship	45.6	46.9	44.7	46.2	42.9	44.6	46.6
Vocational training	21.8	25.4	19.3	21.8	21.7	24.9	18.9
Academic training	8.2	6.7	9.3	8.3	7.7	7.8	8.5
Military	6.3	6.4	6.3	6.2	7.1	6.5	6.1
Not employed	12.0	8.5	14.5	11.8	13.3	9.3	14.6
Number of obs.	2,702	1,112	1,590	2,236	466	1,302	1,400

C	All	Type of school				Age	
		Basic	Middle	Highest	Other	under 19	19 or above
Employed	6.1	6.9	4.7	6.8	6.0	3.1	10.1
Apprenticeship	45.6	47.5	61.7	23.9	41.2	53.4	34.7
Vocational training	21.8	32.3	21.1	9.7	24.7	29.9	10.7
Academic training	8.2	0.2	1.4	27.6	2.2	1.3	17.7
Military	6.3	–	0.2	22.9	0.6	–	15.1
Not employed	12.0	13.0	10.8	9.1	25.3	12.3	11.6
Number of obs.	2,702	854	933	733	182	1,568	1,134

D	All	Women		Men		
		1985–90	1991–97	1985–90	1991–94	1995–97
Employed	6.1	7.02	5.45	7.38	6.03	3.01
Apprenticeship	45.6	42.81	45.81	44.14	54.52	43.37
Vocational training	21.8	25.25	22.21	21.56	16.16	21.39
Academic training	8.2	9.70	10.47	7.81	6.58	3.01
Military	6.3	–	–	12.01	10.14	15.36
Not employed	12.0	15.22	16.06	7.09	6.58	13.86
Number of obs.	2,702	598	716	691	365	332

Note: Communities with at least 20,000 inhabitants are defined as "large." State unemployment is "low" if in the considered year it remains under 9%.
Source: Own calculations based on GSOEP.

Table 4. Descriptive statistics of explanatory variables

Variable	Description	Men		Women	
		Mean	St.Dev.	Mean	St.Dev.
Individual characteristics					
Age	Age in years	18.581	1.786	18.574	1.647
Foreign	Born abroad (0/1)	0.154	0.361	0.139	0.346
Handicap	Handicapped (0/1)	0.099	0.299	0.079	0.270
Basic school	Last school: basic school (0/1)	0.372	0.483	0.257	0.437
Middle school	Last school: middle school (0/1)	0.285	0.451	0.409	0.492
Gymnasium	Last school: Gymnasium (0/1)	0.275	0.447	0.267	0.443
Other school	Last school: other school (0/1)	0.068	0.253	0.066	0.249
Characteristics of parents					
F high education	Father has advanced degree (0/1)	0.281	0.450	0.264	0.441
M high education	Mother has advanced degree (0/1)	0.273	0.446	0.265	0.442
F Foreign	Father born abroad (0/1)	0.323	0.468	0.276	0.447
M Foreign	Mother born abroas (0/1)	0.325	0.469	0.296	0.456
F Employed	Father currently employed (0/1)	0.793	0.405	0.724	0.447
M Employed	Mother currently employed (0/1)	0.540	0.499	0.490	0.500
F Missing	Father information missing (0/1)	0.102	0.303	0.187	0.390
M Missing	Mother information missing (0/1)	0.048	0.213	0.097	0.296
Characteristics of household					
No. of children	No. of children < age 16 in household	1.052	1.114	0.967	1.060
Single parent	Single parent household (0/1)	0.086	0.280	0.101	0.302
Two parent	Two parent household (0/1)	0.890	0.313	0.821	0.383
Regional and labor market indicator					
East	Residence in East Germany (0/1)	0.122	0.327	0.140	0.347
Unempl. rate	State unemployment rate	9.463	3.630	9.553	3.745
Community 1	Community < 5,000 inhabitants (0/1)	0.184	0.388	0.178	0.383
Community 2	Community 5–20,000 inhabitants (0/1)	0.229	0.420	0.232	0.422
Community 3	Community 20–50,000 inhabitants (0/1)	0.180	0.384	0.181	0.385
Community 4	Community 50–100,000 inhabitants (0/1)	0.099	0.299	0.091	0.287
Community 5	Community 100–500,000 inhabitants (0/1)	0.156	0.363	0.175	0.380
Community 6	Community > 500,000 inhabitants (0/1)	0.151	0.358	0.144	0.351
Number of observations		1388		1313	

over time to evaluate possible policy effects. After 1990 the benefit program for academic training became more generous, but the share of transitions into academics increased only slightly for women and not at all for men. Based on the military policy changes we expect a decline in the share of male graduates going to the military after 1990, and an increase in most recent years. These developments are indeed borne out by the frequency distributions.

The independent variables considered in the multivariate analysis are described in Table 4 for the two subsamples. Among the *individual characteristics* age, nationality, type of school attended, and health as measured by

Table 5. Results of the Hausman test for independence of irrelevant alternatives

Omitted category	Men – original version		Men – adjusted version		Women – original version	
	Test statistic	p-Value	Test statistic	p-Value	Test statistic	p-Value
1 Employed	−20.16	(a)	−1.32	(a)	−2.57	(a)
2 Apprenticehip	4,116.50	0.00	(b)	(b)	127.63	0.13
3 Vocational train.	−191.46	(a)	−60.82	(a)	−34.28	(a)
4 Academic train.	776.38	0.00	(b)	(b)	−7.50	(a)
5 Military	−6.76	(a)	0.19	1.00	–	–
6 Not employed	−3,991.34	(a)	79.75	0.25	−4.64	(a)

Note: (a) The test statistic takes on a negative value, which can be interpreted as strong evidence against rejecting the null hypothesis that the IIA assumption holds. (Hausman and McFadden, 1984, footnote 4, or Stata 6 Manual, vol. 2, p. 12)
(b) This outcome is combined with outcome 3 in the adjusted version of the model.

whether the person suffers from a handicap, are considered. While individual characteristics are measured as of the year after the transition, when the dependent variable is observed, parent and household characteristics are gathered in the last year of school attendance, i.e. before the transition to avoid endogeneity problems. *Parent characteristics* combine the level of parent education, nationality, and employment status. Since parent information could not be matched for all school leavers, separate indicators are considered if that information is missing and the missing values of the parent variables are set to zero. The data provides information on a parent-child relationship only for the youth and the household head. We assume that the household's partner is the other parent.[8] As *household characteristics* we consider the number of siblings and whether it is a single or a two parent household.

The set of regional and labor market indicators are East versus West German location, state unemployment rate, and size of the community. The state unemployment rates follow the East/West trends as depicted in Figs. 2 and 3. The majority of school leavers covered in the data resides in communities with less than 50,000 inhabitants. The distribution of school leavers across community sizes closely matches the aggregate figures for Germany (STBA 1998).

3.3. Econometric method

A multinomial logit model is applied to investigate the determinants of school leavers' transition decision. This estimator provides a very flexible approach, as all possible transitions can be considered and no *a priori* restrictions are imposed on the parameters and the set of transition alternatives. However, two features of the model must be discussed. The first concerns the independence of irrelevant alternatives (IIA) assumption, which underlies the multinomial logit estimator. Under IIA the odds of chosing one transition over another are independent of the set of alternatives considered. If IIA does not hold, this may lead to inconsistent estimates. Hausman and McFadden (1984) introduce an IIA test, which was performed here. The results are presented in Table 5. For women the test results suggest that the null hypothesis that IIA holds cannot be rejected. For men this was not the case. When outcomes 2 (apprenticeship) or 4 (academic training) were eliminated from the set of alternatives, the null hypothesis was rejected at high significance levels. Therefore the outcomes apprenticeship, vocational, and academic training are combined to an adjusted dependent variable, for which the IIA hypothesis cannot be rejected.[9]

Second, Moulton (1990) showed that in models where aggregate information, such as state unemployment rates, is considered jointly with individual characteristics, the disturbances may be correlated within aggregation groups (here states). This can bias the standard errors downwards. To address this problem a nonparametric random effects estimator, as developed by Heckman and Singer (1984), was applied to the multinomial logit estimator, which permits tests for correlation among the unobservables of individuals in a given state (see Riphahn 1999 for a more detailed description). This yielded a significant improvement of the likelihood function only for the female sample.[10] Therefore a standard multinomial logit model was estimated for men and a model with corrected standard errors is provided for women.[11] The next section discusses the determinants of transition choices.

4. Results

The two estimated models differ in that the model for women does not contain the transitions into military or substitute service and that the transitions into further training (apprenticeship, vocational, and academic training) are considered jointly in the model for men. The interpretation of the coefficients is complicated because they describe the probability of each outcome relative to the omitted category – the transition to apprenticeship training – and because the signs of coefficients can deviate from those of marginal effects. Therefore, the interpretation first evaluates the statistical significance of the estimates, and then interprets the substantive evidence based on simulation results.

Table 6 presents the results of joint significance tests for coefficients across the full models, estimated for the male and female samples. The indicators for age and prior schooling have the most significant impact on transition outcomes for both samples.[12] Jointly they represent potent indicators of student ability, as weaker students may take additional years before they are able to graduate from a given secondary school. Of high overall statistical significance are the indicators for parent education. The effect of these variables is well established in the literature (Merz and Schimmelpfennig 1999; Rice 1999).[13] In the model for men the indicators for the number of children yield surprisingly significant influences on labor market transitions.

Among the regional effects we observe significant differences between the transition probabilities of the East and West German samples. The state unemployment rate, which is hypothesized to affect transitions through the availability of employment and training opportunities and at the same time as a determinant of the expected future payoff of additional training, is not jointly significant in either subsample. This outcome is sensitive to the consideration of year fixed effects in the model: In models, which consider only a linear time trend instead of the fixed effect control, unemployment was significant for both samples. Finally, the indicators of the size of an individual's community of residence are jointly significant.

Table 6. Joint significance tests of coefficients

	Men		Women	
	χ^2	p-Value	χ^2	p-Value
Individual characteristics				
Age	172.95**	0.000	86.69**	0.000
Foreign	2.54	0.468	9.52*	0.049
Handicap	1.75	0.626	4.60	0.330
Middle/Gymnasium/other school	26.00**	0.000	99.36**	0.000
Characteristics of parents				
High education	16.31*	0.012	20.42**	0.009
Foreign nationality	31.34**	0.000	12.13	0.145
Employed	10.27	0.114	14.71□	0.065
Information missing	11.03□	0.088	9.84	0.277
Characteristics of household				
No. of children	11.27*	0.010	3.37	0.498
Single parent	2.29	0.514	3.30	0.508
Two parents	4.02	0.259	3.75	0.440
Regional and labor market indicator				
East	8.43*	0.038	12.69*	0.013
State unemployment rate	5.02	0.170	7.02	0.135
Community size effects	23.49□	0.074	34.00*	0.026

Note: The coefficients for the transition into apprenticeship, vocational or academic training are restricted to zero in the model for men, in the model for women coefficients for transitions into apprenticeship are restricted to zero. Reference categories are basic school, and the largest community size. The estimations control for a set of year dummies (jointly significant at the 2% level). **, *, and □ indicate statistical significance at the 1, 5, and 10% level.

To aid the substantive interpretation of the estimated effects, Table 7 presents simulations of the variables' effects on the transition probabilities. The simulations are obtained in two steps. First, baseline probabilities are predicted for each observation based on the estimated coefficients. The average predicted baseline probabilities (first rows in the panels of Table 7) agree with those presented in Table 3. Next, single variables are set to fixed values for all observations and the predictions are repeated. The difference in predicted probabilities, e.g. when "foreign origin" is set to 1 minus the probability when it is set to 0, is divided by the baseline probability for each observation. The figures in Table 7 indicate effects of single variables on transition probabilities, measured in percent of the baseline probabilities.

Since the transition into nonemployment is of prominent policy relevance, we focus on the last columns in Table 7. A first finding is that the explanatory variables frequently have larger effects on the probability of nonemployment for women than for men. Being of foreign origin increases both samples' probability of nonemployment, but the effect on females is almost three times as large as that on men. The consequence of a handicap shows the same pattern. The probability of nonemployment increases by 48% for handicapped men and about doubles for handicapped women. However, here the underlying coefficients are not precisely determined.

The schooling effects are as expected: The higher the level of schooling, the less likely school leavers are to become nonemployed. Only the effect of middle versus basic school for men does not fit the pattern, but this coefficient is

Table 7. Simulation results – Determinants of transitions into: **0:** apprenticeship and vocational training and academic training **1:** employment **2:** apprenticehip **3:** vocational training **4:** academic training **5:** military or substitute service **6:** nonemployment

A: Men – transition into:	0	1	5	6
Baseline probability	0.731	0.060	0.123	0.086
Individual characteristics				
Foreign origin (1 vs. 0)	−0.050	0.522	−0.097	0.239
Handicap (1 vs. 0)	0.174	−0.456	−0.866	0.483
School: Middle vs. basic school	−0.005	−0.196	−0.005	0.187
School: Gymnasium vs. basic school	0.087	−0.926	0.087	−0.234
School: Other school vs. basic school	−0.185	0.798	−0.185	1.671
Parent and household characteristics				
Both parents of foreign origin (1 vs. 0)	0.110	1.787	−2.632	0.291
No. of children under 16 in household (1 vs. 0)	0.015	0.234	−0.436	0.171
Household type: Two parents vs. single parent	0.218	−0.067	−0.076	−1.993
Regional and labor market characteristics				
Region: East vs. West Germany	0.077	−0.705	0.642	−0.970
Unemployment: 13 vs. 6%	0.014	0.421	−0.569	0.397
Community size: <5K vs. >500K Inhabitants	−0.032	1.209	0.161	−0.798
Community size: <20K vs. >500K Inhabitants	0.022	0.342	0.441	−1.151
Community size: <50K vs. >500K Inhabitants	0.074	−0.361	0.093	−0.593
Community size: <100K vs. >500K Inhabitants	−0.039	0.561	0.577	−0.900
Community size: <500K vs. >500K Inhabitants	0.067	0.293	0.070	−0.983
Policy effects (before 1991 vs. after 1991)	−0.046	−0.495	0.243	0.621

B: Women – transition into:	1	2	3	4	6
Baseline probability	0.063	0.426	0.229	0.123	0.158
Individual characteristics					
Foreign origin (1 vs. 0)	0.410	−0.211	0.058	−0.409	0.687
Handicap (1 vs. 0)	−0.331	0.202	−0.176	−1.318	1.058
School: Middle vs. basic school	−0.904	0.335	−0.188	1.225	−0.513
School: Gymnasium vs. basic school	−1.560	−0.563	0.588	13.225	−0.734
School: Other school vs. basic school	−0.987	−0.156	−0.083	1.687	0.626
Parent and household characteristics					
Both parents of foreign origin (1 vs. 0)	0.037	−0.311	0.090	0.405	0.412
No. of children under 16 in household (1 vs. 0)	0.108	−0.068	0.042	0.069	0.019
Household type: Two parents vs. single parent	0.985	0.195	−0.818	0.401	−0.191
Regional and labor market characteristics					
Region: East vs. West Germany	0.126	0.345	−0.601	1.395	−0.724
Unemployment: 13 vs. 6%	−0.488	−0.082	0.243	−0.635	0.511
Community Size: <5K vs. >500K Inhabitants	0.024	0.066	0.111	−0.166	−0.183
Community Size: <20K vs. >500K Inhabitants	−0.261	0.142	0.228	0.789	−0.242
Community Size: <50K vs. >500K Inhabitants	−0.478	−0.145	0.183	0.215	0.135
Community Size: <100K vs. >500K Inhabitants	−0.355	−0.070	0.354	0.562	0.043
Community Size: <500K vs. >500K Inhabitants	−0.132	−0.433	0.167	1.785	−0.023
Policy effects (before 1991 vs. after 1991)	0.019	0.122	−0.054	−1.471	0.275

Note: The figures describe the deviation between the two predicted probabilities relative to the baseline probability and can be interpreted as percentage deviation from the baseline due to changes in variables.

statistically insignificant. Females with middle school or *Gymnasium* degrees have lower risks of nonemployment relative to basic schooling. Women in the "other school" category have an elevated risk of nonemployment, but also a surprisingly high transition probability into academic training.

Independent of the nationality of the graduates themselves, the nationality of their parents has sizeable effects on youth labor market outcomes. If both parents are non-natives, the probability of nonemployment increases by 29% for men and by again a more sizeable 41% for women. Also, all indicators of non-native nationality yield increased probabilities of immediate employment after school. These results raise concerns, as they show that independent of ability, non-native ethnic background seems to hinder continued training. The effects of non-native ethnicity in the British data are typically reverse: Rice (1999) and Andrews and Bradley (1997) show that non-whites have higher probabilities of participating in continued education. A possible rationale for the different ethnicity effects in the U.K. and German framework might be that the German non-native population consists mostly of blue-collar guest-workers whereas British immigrants are more favorably self-selected in terms of preferences for human capital.

The indicator of the number of young siblings in the household has a significant and sizeable effect on men's transitions into nonemployment. Having one versus no child under age 16 in the household increases the probability of nonemployment (as well as of immediate employment) by about 20%. The indicator has not been considered in the German literature so far, however, the results confirm Micklewright (1989), who finds that youth with more siblings have a lower probability of continued education. The mechanism behind this effect is likely to be related to parents' financial and time constraints. The more siblings there are to care for, the fewer financial resources are available to finance continued education, or to support job search activities.[14] The same pattern is likely to be behind the result that youth from two parent households have a smaller risk of nonemployment than those from single parent households. For men also the probability of immediate employment falls and that of continued training increases, if they come from a two as opposed to a single parent household.

Confirming the distribution discussed above, East German youth has a lower probability of nonemployment than youth in the West. East German men have a high transition probability into military service, an about equal probability of training, and a much lower probability of starting employment immediately after leaving school. For East German females the probability of academic training exceeds that of women in the West by far, a phenomenon which is likely to be related to higher female labor force participation and thus higher expected lifetime returns to education in East compared to West Germany.

As expected, state unemployment yields sizeable effects on nonemployment transitions. Again, female school leavers are more strongly affected by a given change in unemployment than their male classmates. Women's transition probabilities into immediate employment also decline drastically if state unemployment is high. The effect on male youths surprisingly points in the opposite direction. Their negative response in the probability of a transition to military service, when unemployment is high, is also difficult to rationalize. For females the probability of taking up apprenticeship training declines in times of high unemployment and that of other vocational training increases, a plausible pattern, as the availability of apprenticeship positions should be correlated with the overall labor market tightness.

The community size effects are evaluated by comparing transition probabilities in smaller communities with those for a metropolitan center. Four patterns emerge: (i) The probability of nonemployment is largest in the metropolitan center. (ii) The probability of taking up military service is lowest for men in metropolitan areas. (iii) Men are least likely to start immediate employment in metropolitan areas. (iv) The probability of a transition into vocational or academic training for women is particularly low in metropolitan areas, where a wider pool of apprenticeship positions might be available instead. It appears that school leavers outside of metropolitan centers make up for a lack of apprenticeship positions by chosing vocational school or academic training. For women in midsize towns this pattern is quite clear.[15]

Finally, it is interesting to look at policy effects. The first hypothesis is that the probability of academic training increased after the generosity of the *Bafög* benefit system was expanded in 1990. The probability distributions in Table 3D indicated no such effect for men, and only a slight increase in the probability of academic training for women. In the multivariate framework this correlation can only be examined for females, since the dependent variable for men combines academic training with other outcomes. The simulations suggest that there is a strong increase in the probability of taking up academic training after 1990 for women by about 150% of the baseline probability. This suggests a clear response in individual behaviors to the relaxation of financial constraints.

The second policy hypothesis refers to changes in military policy. The above discussion suggests that the probability of a transition to military or substitute service should be high before 1990, decline after the cold war, and rise again in the last years of our data. Averaging the predicted probabilities of a transition into military service over the relevant periods yields a mean probability of 11.6% for 1985–1990, of 8.6 for 1991–1993, and of 17.2 for 1994–1997. These figures (not presented in Table 7) indicate a close correlation with the expected pattern.

5. Conclusion

Motivated by the increasing attention of labor market policies to the youth unemployment problem, this study investigates labor market transitions of secondary school leavers in Germany. Using a detailed dataset and a flexible modelling approach, the determinants of such transitions are evaluated. As potentially influential factors characteristics of the youth, household, and parents, as well as indicators of the region of residence, and local labor markets are considered.

The findings confirm many of the conclusions from the British school-to-work literature. Overall the most significant effects are those describing the ability level of the individual, as reflected in the indicators of age and school type. The latter expectedly yields that the probability of nonemployment is least for those youth with the highest completed degree. The results permit conclusions as to the groups most vulnerable to the risk of nonemployment, or of foregoing continued education: Youth with a handicap and immigrants suffer substantial nonemployment problems. The latter effect is obtained based both on the nationality of the individual as well as on the parents', and it is consistent in both, the male and female subsamples.

Two findings that are new to the literature are that the educational attainment of parents has a significant effect on youth labor market transitions and that youth from two-parent families are in a much better position to take up vocational training after secondary school than individuals from single parent families, an outcome which holds particularly for young men. The simulation results yield a clear correlation between the probability of taking up academic training and expansions in the student-benefit program in 1990. This agrees with Rice (1987), who found that the introduction of an educational allowance would have positive effects on the participation in continued education. – Additionally, men's participation in military service seems to respond to general trends in military policy.

With respect to regional and local labor market effects, living in a high unemployment state is correlated with higher risks of nonemployment for school leavers. Given state unemployment rates, the risk of nonemployment is highest for youth in large metropolitan areas. These results suggest that youth labor market policies should focus on disadvantaged regions with tight overall labor markets, emphasize metropolitan areas, and possibly pay particular attention to the needs of immigrant youth.

Endnotes

1 The benefit program also provides payments to high school students and participants in certain vocational training programs under restricitive and complex regulations. Since most of the program expenditures are alloted to university students, the discussion focuses on this program aspect. For detail see e.g. BMA (1995).

2 This may in part be due to the definitional exclusion of "apprenticeship seekers" from the ranks of the unemployed. Franz et al. (1997) show that their inclusion would drive up youth unemployment rates by about 20%.

3 Bavaria, Baden-Württemberg, Hesse, Lower Saxony, Northrhine-Westfalia, Rhineland-Palatinate, Saarland, Schleswig-Holstein.

4 Micklewright (1989) and Rice (1999) find highly significant differences in the educational choices of students in different areas of the country, as well as by local labor market conditions.

5 Imposing an age limit on the considered observations is important and necessary to avoid measurement errors, that might otherwise occur with the consideration of outlier observations (of school leavers above age 25) in the sample. When normal school exit age is 16 or at most 19, considering an additional 9 to 6 years covers the relevant population. Also, it is unlikely and not plausible that the omission of observations with missing values on explanatory variables or on school participation biases the estimation, since survey responses of typically other members of the household should be uncorrelated with post-secondary schooling transition decisions. Therefore the selection mechanisms should be neutral to the estimation results.

6 For a description of these institutions see Sect. 2 above. *Sonderschule* provides special education for those unable to pass basic school. *Grundschule* is elementary school.

7 About 20% of all observations were affected by multiply coded school exits.

8 In addition, parent information was used, which was explicitly gathered in the survey of 1986.

9 Initial adjustments of combining smaller subsets of transition outcomes did not satisfy the IIA assumption.

10 For the male sample the log likelihood could not be improved beyond the value of 888.20 obtained before, for women the consideration of controls for correlated errors improved the log likelihood from 1,455.55 to 1,449.56, which is statistically significant at the 5% level given that 5 additional parameters were estimated.

[11] To save space the full set of coefficient and standard error results is not presented. Preliminary estimations showed that a different health indicator did not affect the results, that a richer set of indicators for parent schooling degrees, and the consideration of parents' age, or of immigrant students' language capacity did not significantly add to the explanatory power of the model, that unemployment rates measured at a more disaggregated level had less explanatory power, than the state level ones, and that a variable controlling for the number of persons living in a household did not significantly improve the model fit.

[12] The schooling coefficients in the model for a transition into military service were restricted to zero, because the fact that *Gymnasium* graduates have reached draft age at graduation caused an unreasonably large coefficient in this model. To provide an indirect test for the potential endogeneity of the school type, estimations without these variables were performed. The results did not differ in major ways.

[13] It is a striking advantage of the GSOEP that it provides indicators for parent educational status. The studies by Inkmann et al.(1998) and Franz and Zimmermann (1999) are based on a dataset with more observations, but which apparently does not permit this type of control. Only Franz et al. (1997) are able to consider controls for the *vocational background* of the household head, which they find to be influential. Neither partner information nor the educational status of the household head are considered.

[14] For women this conclusion is somewhat counterbalanced by the higher probability of a transition into vocational or academic training.

[15] Critics may argue that the community effects are subject to the reflection problem pointed out by Manski (1993). He argues that the effect of (the characteristics of) an aggregate, e.g. a state, on the behavior of a unit of observation typically cannot be identified. For our case this implies that parents' decision to move e.g. to a large town is determined by similar factors and motives as the outcome of their child's post-school transition. One condition for such an argument to hold is that households can indeed be observed to move. A dataset generated before defining the transition variable, contained panel information on a multitude of households with and without a graduating youth, before and after a possible graduation. In this dataset, with more than 90,000 household-year observations, 0.74% of the observations moved between states. In the final sample the share of moving households dropped to 0.1%, suggesting that very little moving takes place. Overall the argument that the decision of parents to move between states is correlated with a transition decision of their children at a possibly much later date is not convincing. Therefore endogenous moves are not likely to bias the presented results.

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Transitions from employment among young Norwegian workers

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Abstract. In a large representative sample of young Norwegian workers, we estimate gross transitions to unemployment, education, and other exits in a multinomial logit. In line with received literature, we find that individuals with high education, experience, and income have significantly lower probabilities of job exits. While female education rates have increased to surpass those of males, female labour market outcomes are still more responsive to family related background characteristics as compared with the outcomes for males.

JEL classification: J64, J21

Key words: Competing risk, youth employment, youth unemployment

1. Introduction

There is a wealth of studies evaluating policies to facilitate transitions into work from unemployment. Among several notable American contributions, we mention Card and Sullivan (1988), and Meyer (1990), while Ackum (1991) and Torp (1994) have studied Scandinavian labour markets, and Winkelmann (1997) offers a recent comparison between Germany and the United States. However, even when the main focus is on unemployment, it is important to

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understand events within the work state: why do some individuals exit from work while others don't? This point has been convincingly made in a series of pioneering papers by Clark and Summers, several of which are collected in Summers (1990). It is reiterated by Layard, et al. (1991, p. 297) in an evaluation of youth unemployment where they conclude that: "*Unemployment is, of course, almost everywhere more common among young people than among adults. Very often, the difference results from higher inflow rates – and certainly not from unusual duration.*"

We follow this lead, and investigate exits from work (including both layoffs and quits) among young Norwegian workers between January 1989 and December 1990. This was a period of a cyclical downturn in Norway, such that quite a few youngsters lost their jobs due to layoffs. A recent survey of the Norwegian labour market situation in this period is given in Torp (1996). Our investigation is made possible by the availability of a new Norwegian database that contains a wealth of information on a randomly selected 10% sample of the Norwegian population. We have extracted as many as 32202 workers between ages 18–29 from this base. The available data set permits us to identify different end-states out of employment such as unemployment, further education, or other exits out of the labour force. We analyse straightforwardly the gross exits to these states by means of a reduced form multinomial logit. Thus, we capture the competing risk aspect of the inflow from employment to youth unemployment.

In the next section, we briefly discuss our methodological approach. Section 3 presents the data, and an overview of destination states. Summary statistics and the results of the multinomial logit are presented, and interpreted in Sect. 4, while Sect. 5 concludes.

2. Method

Our reduced form multinomial logit formulation answers a simple empirical question: What factors affect the probability of being observed in the end states *work, unemployment, education*, and *other* in December 1990, conditional on being in the state *work* in January 1989? If there were no round-tripping, i.e. individuals moving in and out of the different states in the interim period, our approach would also capture net flows, but not durations. The occurrence of round tripping in our material is illustrated by the following figures: 70% of those that occupied a job in the end state had held that same job during the two years, while 80% of the unemployed became unemployed within the last six months before our time of final observation. These figures indicate that round tripping is nowhere nearly as prominent in our sample as in the sample of teenagers analysed in Clark and Summers (1982), for instance.

Clark and Summers (1982), Flinn and Heckman (1982), and more recently Gönül (1992) have addressed the question whether unemployment and out of the labour force are distinct states for youngsters. They investigate whether reported job search promotes employment. The results are mixed. From our perspective, we focus on the idea that the policy, and labour market, implications for youngsters that leave the labour force to enter education are different from those that exit altogether. Further education may reasonably be interpreted as a commitment to return to the labour force at a later date, while stopping search may be a less active alternative. Therefore, we have separated those two alternatives of leaving the labour force, so we are left with exits from work to three alternative states, where the individuals are *unemployed*, under *education*, or belonging to *others*. We implement the estimation through the multinomial logit

$$Prob(Y_i = j) = \frac{e^{\mathbf{b}_j'\mathbf{x}_i}}{\sum_{j=0}^{3} e^{\mathbf{b}_j'\mathbf{x}_i}}, \tag{1}$$

where $Y_i = j$ indicates that the dependent state variable Y for individual i takes the value $j = 0, 1, 2, 3$. In Sect. 4 we focus on competing risks, and report the marginal effects found by differentiating equation (1):

$$\frac{\partial P_{ji}}{\partial \mathbf{x_i}} = P_{ji}\left[\mathbf{b_j} - \sum_{k=0}^{3} P_{ki}\mathbf{b_k}\right], \quad j = 0, 1, 2, 3, \tag{2}$$

where P_{ji} equals $Prob(Y_i = j)$ as defined in equation (1). Our explanatory variables, $\mathbf{x_i}$, comprise personal and demographic variables such as number of children, marital status, education and age, and labour market related variables such as experience, income, replacement rate, and unemployment rates. We allow these effects to have various impacts on the flow to different states for the two genders by doing the analysis separately for males and females.

3. Data

3.1. Data sources

The analysis draws on data from the KIRUT database. (KIRUT is a Norwegian acronym which roughly translates to "Clients into and through the Social Insurance System"). The base contains detailed individual information on socio-economic background, labour market participation, and social insurance payments for a random 10% sample of the Norwegian population between ages 16–67. The information is merged from several different public registers, with the consent and supervision of the Norwegian Data Protection Agency. The large sample provided in KIRUT (the total sample exceeds 300 000 individuals), combined with the fact that such wealth of information on each individual is merged from diverse administrative records, makes this database an interesting source of analysis to complement analyses based on tailor-made surveys.

Our sample includes observations of individuals aged 18–29 that occupy a permanent job on January 1, 1989. To eliminate students working part-time, we only include youngsters that are registered with taxable earnings greater or equal to NOK 60 000 (about £ 6 000) in the previous year. This corresponds to the limit set by Norwegian educational authorities, above which all rights to public support for students are forfeited. Individuals working less than 4 hours per week and individuals who are registered as job-seekers are excluded from the sample. Finally, we drop 788 observations due to missing variables. We end up with a final sample of 32 202 individuals.

All individuals in the final sample are classified into one of four groups based on their labour market status two years later, on December 15, 1990. These groups are; *work, unemployment, education* and *other*. The classification procedure is as follows. First, all individuals are classified as other. Second, an individual that is registered as a student and works less than 20 hours per week is classified as student. Individuals registered as job-seekers, and again, work less than 20 hours per week, are classified as unemployed. Finally, students or job-seekers working 20 hours or more per week, together with individuals registered working 4 hours or more and not found in any of the other registers, are classified as workers. This latter requirement is applied to ensure consistency with the criterion used to exclude non-workers in the beginning of the sample period. Note that all classifications are based on public, administrative records. Therefore, to be registered as unemployed, for instance, an individual must be registered as a job seeker with the labour market authorities.

3.2. Variables

All the explanatory variables are measured at January 1, 1989. There are four family related variables; marital status (1 if married, 0 otherwise), number of children less than 11 years old, an interaction term between unmarried and number of children less than 11 years old, and spouse income. Citizenship (1 if Norwegian, 0 otherwise), and number of years of education are also included in our model. The probabilities of transiting from one state to another may also be a function of age and four age dummies are therefore included. We use age dummies to accommodate potential non-linearities. The dummies are grouped over three-year intervals, while differences over single-year cohorts may be dominated by stochastic elements.

Our proxies of experience are based on information about pension points earned in the National Social Insurance Scheme (referred to as P.P. in the tables). Pension points earned is a piecewise linear function of income and only earned in years when income exceed NOK 37 300.[1] Our two experience proxies are as follows. Experience A; the number of years with positive pension points. Experience B; the aggregate number of pension points earned. In addition to the experience variables, information about the quality or skills of an individual may be contained by income in the previous year. This income also proxies expected future income. Therefore, it is expected to increase the probability of staying in the work state. To allow for non-linearities in the income variable, we also include income squared in the model. Our replacement rate proxy variable is also based on income. The unemployment benefits in Norway are 62.4% of income previous year up to NOK 223 800. For incomes higher than NOK 223 800 unemployment benefits are constant such that the replacement ratio declines. Therefore our replacement variable is defined as the deviation in income from the threshold value of NOK 223 800 for those individuals with income higher than this value, and zero otherwise. An increasing value of this variable indicates a decreasing replacement rate.

Therefore, we expect the probability of entering the unemployment state to decrease in this proxy. Note that income, and income squared, already enter the equation, such that the replacement rate proxy captures whether something happens around the break-point of NOK 223 800. All income variables (spouse income, own income, own income squared and replacement rate) are measured in NOK 10 000.

Finally, the transition to unemployment is likely to depend both on the industry where an individual works and the local unemployment. To control for such effects, the unemployment rate in the local municipality of each individual is included in the model together with nine industry dummies.

4. Empirical results

For well-known reasons we follow standard practice in labour market analysis and run separate analyses for males and females.

4.1. Descriptive statistics

Summary statistics by labour market status at the end of 1990 are given in Table 1. The frequency of marriage, and the number of children, are lower for men relative to women. This difference is caused by the fact that the average age of marriage is higher for males relative to females. The average age of first marriage in 1989 was 28.3 for males and 25.8 for females (NOS Population Statistics). Comparing different end-states, workers have the highest frequency of marriage, while students are at the other end of the scale. The frequency of foreigners in the sample is low, and they are more likely to end up taking more education relative to Norwegian youngsters. Finally, individuals with high initial education tend to return to further schooling. Men have more work experience, and higher incomes, relative to women. Males also have a slightly higher probability of staying employed (79.6%) as compared with females (76.4%). For incomes higher than the threshold value, where the replacement rate starts to decrease, individuals are more likely to stay in work. Clearly, further analysis is needed to disentangle income and replacement effects. The summary statistics also indicate that individuals living in municipalities with a higher unemployment rate are more likely to be unemployed, as expected.

4.2. Multinomial logit model results

In Table 2 the marginal effects of the multinomial logit model are given. The marginal effects both reflect an impact from the coefficients, and from competing risk, as shown in equation (2).[2] The reported marginal effects cannot be given a clear structural interpretation in our reduced form regression. Moreover, some of our regressors may be plagued by potential endogeneity problems due to selection processes on unobserved background characteristics. For instance, the estimated marginal effects of education and experience may reflect both the impact of the variables themselves, and potential unobserved innate characteristics. The estimated implicit returns to these variables

Table 1. Summary statistics by end status. Mean values. Standard errors are reported in parentheses; measuring units in square brackets. For more detailed variable definitions, see text

	Male				Female			
	Work	Unemployment	Education	Other	Work	Unemployment	Education	Other
Married[a] [percentage]	21.7 (41.2)	12.9 (33.6)	5.61 (23.0)	14.1 (34.8)	26.8 (44.3)	25.2 (43.4)	11.8 (3.3)	31.3 (46.4)
# children	0.23 (0.58)	0.17 (0.49)	0.10 (0.37)	0.19 (0.54)	0.32 (0.62)	0.37 (0.67)	0.15 (0.43)	0.34 (0.61)
Unmarried* # children	0.03 (0.18)	0.05 (0.24)	0.05 (0.26)	0.05 (0.27)	0.10 (0.33)	0.14 (0.40)	0.07 (0.29)	0.11 (0.35)
Spouse income [10^4 NOK]	1.71 (4.27)	0.77 (2.79)	0.39 (2.06)	1.02 (3.42)	4.65 (8.59)	4.26 (8.14)	1.77 (5.70)	5.60 (9.59)
Norwegian[a] [percentage]	98.7 (11.4)	97.9 (14.3)	97.2 (16.5)	98.4 (12.6)	99.1 (9.65)	98.9 (10.4)	97.7 (14.9)	98.9 (10.4)
Education [years]	11.3 (1.76)	10.5 (1.25)	11.6 (1.51)	10.9 (1.56)	11.7 (1.84)	10.8 (1.32)	12.0 (1.53)	11.4 (1.80)
Experience A [years with p. p.]	6.01 (2.79)	5.21 (2.90)	3.50 (2.17)	5.04 (2.82)	5.24 (2.58)	4.49 (2.54)	3.53 (2.09)	5.05 (2.59)
Experience B [p. p.]	16.3 (12.0)	12.2 (10.4)	6.56 (7.04)	12.1 (10.61)	10.9 (8.43)	7.72 (7.08)	5.80 (5.56)	9.89 (8.14)
Income [10^4 NOK]	16.1 (5.32)	13.7 (4.64)	11.5 (4.41)	14.2 (5.38)	12.6 (4.00)	10.7 (3.24)	10.4 (3.50)	11.7 (3.84)
Replacement rate proxy [10^4 NOK]	0.44 (1.88)	0.14 (1.00)	0.11 (1.10)	0.31 (1.86)	0.07 (0.70)	0.01 (0.16)	0.01 (0.13)	0.04 (0.51)
Unemployment	0.02 (0.01)	0.03 (0.01)	0.02 (0.01)	0.02 (0.01)	0.02 (0.01)	0.03 (0.01)	0.02 (0.01)	0.02 (0.01)

Number of observations	14548	1060	856	1804	10649	830	705	1750
by age [percentage]								
Age 18–20	7.0	17.7	25.4	18.6	5.5	10.7	13.9	7.5
Age 21–23	20.3	26.7	40.2	26.6	26.1	35.2	45.5	28.1
Age 24–26	33.5	29.0	21.0	30.4	33.7	30.2	25.5	30.7
Age 27–29	39.3	26.6	13.4	24.4	34.7	23.9	15.0	33.7
by *industry* [percentage]								
Primary sector	2.8	4.2	3.3	4.3	0.9	1.4	1.6	1.0
Oil and mining	1.2	0.8	0.7	0.5	1.0	0.2	0.4	0.7
Manufacturing	26.3	21.2	19.7	24.2	12.0	15.1	6.8	12.6
Electricity, construction	17.0	32.3	21.3	23.0	2.2	3.7	1.1	1.7
Trade and hotels	17.4	18.1	15.1	18.1	21.1	32.7	26.1	26.8
Transport	9.5	7.7	7.8	9.3	7.1	4.5	4.3	7.0
Finance	7.7	4.2	6.4	5.3	14.3	10.0	11.1	12.5
Public services	15.1	9.6	23.6	14.4	33.7	29.9	46.7	35.7
Unregistered	3.1	1.8	2.1	1.0	5.7	2.5	2.0	2.0

[a] These numbers are measured in absolute terms in the multinomial logit model

Table 2. Marginal effects, multinomial logit model. Standard errors are reported in parentheses. For variable definitions, see text and Table 1

	Male				Female			
	Work	Unemployment	Education	Other	Work	Unemployment	Education	Other
Married	0.003	0.014	−0.011	−0.005	−0.011	0.001	−0.010	0.020
	(0.017)	(0.009)	(0.008)	(0.014)	(0.018)	(0.009)	(0.008)	(0.014)
#children	0.004	−0.008	0.001	0.003	0.036**	−0.005	−0.005	−0.025**
	(0.009)	(0.005)	(0.004)	(0.007)	(0.009)	(0.004)	(0.005)	(0.007)
Unmarried* #children	−0.008	0.007	−0.003	0.004	−0.039**	0.010	−0.005	0.034**
	(0.015)	(0.008)	(0.006)	(0.011)	(0.013)	(0.006)	(0.007)	(0.010)
Spouse income	3.7E-3**	−2.5E-3**	−4.4E-4	−7.6E-4	−1.4E-3	3.7E-4	−6.9E-4	1.7E-3*
	(1.4E-3)	(7.9E-4)	(6.3E-4)	(1.1E-3)	(8.9E-4)	(4.7E-4)	(4.7E-4)	(6.7E-4)
Norwegian	0.041	−0.013	−0.019**	−0.009	0.025	0.003	−0.033**	0.005
	(0.021)	(0.011)	(0.006)	(0.018)	(0.033)	(0.016)	(0.009)	(0.028)
Education	0.017**	−0.015**	0.006**	−0.007**	0.015**	−0.014**	0.007**	−0.007**
	(0.002)	(0.001)	(0.001)	(0.002)	(0.002)	(0.001)	(0.001)	(0.002)
Experience A	0.001	−0.001	−0.001*	0.001	0.008**	−0.003*	−0.002	−0.003
	(0.002)	(0.001)	(0.001)	(0.002)	(0.003)	(0.001)	(0.001)	(0.002)
Experience B	2.2E-3**	−7.3E-4**	−5.8E-4**	−9.0E-4**	1.7E-3*	−1.2E-3**	−4.7E-4	−5.6E-5
	(4.5E-4)	(2.3E-4)	(2.1E-4)	(3.5E-4)	(7.7E-4)	(4.0E-4)	(3.9E-4)	(6.2E-4)
Income	0.015**	−0.004*	−0.003*	−0.007**	0.031**	−0.005	−0.011**	−0.015**
	(0.004)	(0.002)	(0.001)	(0.003)	(0.006)	(0.003)	(0.002)	(0.005)
Income^2	−1.6E-4	6.7E-5	−1.3E-5	1.1E-4	−7.5E-4**	1.0E-4	3.0E-4**	3.6E-4
	(1.3E-4)	(8.0E-5)	(5.2E-5)	(9.3E-5)	(2.2E-4)	(1.3E-4)	(9.0E-5)	(1.8E-4)
Replacement rate proxy	−0.010	0.000	0.006*	0.005	0.019	−1.40E-04	−0.013	−0.005
	(0.006)	(0.004)	(0.002)	(0.004)	(0.014)	(8.7E-03)	(0.009)	(0.009)
Unemployment	−0.110	0.389**	−0.082	−0.196	0.034	0.799**	−0.320**	−0.514*
	(0.211)	(0.107)	(0.070)	(0.172)	(0.275)	(0.123)	(0.109)	(0.235)

Age 21–23	0.041**	0.000	−0.006*	−0.035**	−0.009	0.010	−0.003	0.002
	(0.010)	(0.005)	(0.003)	(0.008)	(0.014)	(0.006)	(0.004)	(0.012)
Age 24–26	0.059**	0.004	−0.020**	−0.043**	−0.001	0.016*	−0.021**	0.006
	(0.012)	(0.006)	(0.004)	(0.010)	(0.017)	(0.007)	(0.006)	(0.014)
Age 27–29	0.067**	0.016	−0.024**	−0.059**	−0.019	0.020*	−0.030**	0.029
	(0.016)	(0.009)	(0.005)	(0.013)	(0.020)	(0.009)	(0.008)	(0.017)
χ^2_{industry} (df=8)	128.09**	121.78**	32.98**	44.80**	73.54**	35.27**	43.70**	38.47**
Log Likelihood	−12067.9				−10321.8			
Pseudo R^2	0.081				0.056			
Number of observations	14548	1060	856	1804	10649	830	705	1750

** Indicates significance at 1% level
* Indicates significance at 5% level

may therefore be biased upwards. However, with due caution in interpretation, our analysis does identify factors that affect the probability of entering the different states. Furthermore, the multinomial formulation aids a richer understanding of exits from work as compared to models with single exit routes. For instance, the factors behind a decision to stay at home with children are different from those that may explain the transition from work to unemployment.

Family related variables. Although there are some surprises, family related characteristics confirm some traditional gender differences. The presence of children has no significant effects on males, while females are affected. For married females there is a significant tendency to increase the probability of staying employed with increasing number of children, while the probability of leaving the labour force decreases. These marginal effects of children are erased for unwed mothers.[3] However, remember that we are investigating the behaviour of females that had chosen to work at the time. Labour force participation by a mother of several children is probably an indication of a high endowment of personal human resources. Non-participating mothers with many children may not be as fortunate on the average, so selection effects may play a role. Increased spouse income increases the probability for women to leave the labour force. For males, the effect of increased spouse income is opposite, increasing the probability of staying in a job and decreasing the probability of being unemployed. This pattern for the males may also be due to selection, through assortative mating.

Other individual characteristics. Non-Norwegian citizenship increases the probability of transiting from work to school, like in Table 1. Individuals with higher education have higher chances of staying employed, and correspondingly, lower probability of being unemployed. In addition, we find that individuals with a relative high education tend to take more education if leaving a job. All effects relating to education are statistically significant for both males and females. For the transition from work to *other*, the marginal effects of education are negative and significant, while the coefficients are positive and significant for both genders. This illustrates that the direct effects measured by the estimated coefficients can deviate significantly from the marginal effects due to competing risk, as shown in equation (2). However, this is the only instance where we obtain such significant differences between coefficients and marginal effects.

With few exceptions, the age dummies show a monotonic increasing or decreasing pattern, depending on the transition. For males, age increases the probability of staying employed, and decreases the probability of taking more education or ending in the unspecified group. For females, the effects of age on the transition to education are similar to the ones for males, while age increases the probability of being unemployed. While age obviously is strongly exogenous, it interacts in the regression with experience, income and other variables that may be prone to potentially endogenous selection processes as we have discussed earlier. For instance, when we run a regression for females with only age dummies as regressors, age decreases the probability of becoming unemployed. While such a regression is misspecified, it inspires caution in the interpretation of the results, both regarding age and obviously those variables that are prone to selectivity.

Labour market variables. Increased experience is associated with a higher probability of staying employed, and lower probabilities of entering all other states. For males, this effect is significant only for the aggregate pension point experience variable (Experience B). For females, both aggregate pension points and number of years affect the probability of staying employed, or transiting to unemployment, in a significant and expected manner. The marginal effects of income are as expected both for males and females. Increased income increases the probability of staying employed and decreases the probability of transiting to the other labour market states. There are some significant second order effects of income for females, while such effects are not present for males. The replacement rate proxy does not explain much. This may both result from low variance in the variable, and that income variables catch some of the potential effects.

A high local unemployment rate, not surprisingly, increases the probability of becoming unemployed. However, it decreases the probability of dropping out of the labour force. This is one of the few observed patterns in Table 2 that points at systematic differences in transitions to unemployment as compared to out of the labour force. In most instances, explanatory variables tend to influence all transitions out of employment in the same qualitative manner. In this way, inspection of Table 2 does not reveal strong indications that the recruitment processes to unemployment as compared to out of the labour forces are different. However, when we test whether all the marginal effects are equal across two and two states by a chi-square test, the null hypothesis of no differences are strongly rejected, in all states and for both genders.

The industry chi-square test indicates that there are important sectoral differences. These differences may be due to business cycles, since the demand and activity level varied substantially for different sectors and industries in the period 1989–1991. In particular the construction industry experienced a serious recession these years. Finally, the data fit the multinomial logit model (measured by Pseudo R^2) somewhat better for males as compared to females.

Gender differences. Looking at the number of observations in each state in Table 1, and the pattern of marginal effects in Table 2, we find that there are significant gender differences. To test what the driving forces behind these differences are, we run a regression of the multinomial logit model where the two sub-samples of males and females are merged.[4] We include a gender dummy (1 if female, 0 otherwise) separately, in addition to interaction terms between the explanatory variables and this gender dummy.[5] A significant interaction indicates gender differences in transitions due to the explanatory variable concerned. We wish to explore gender differences in transitions further. To accommodate this, we group the interaction terms according to our previous labelling; family related variables, other individual characteristics, and labour market variables. For each of these groups, we run chi-square tests to evaluate gender differences in transitions to the end states. The chi-square tests indicate that the family related variables are most important in explaining gender differences in the transits. However, the transition from work to education is an exception. For this latter transition, the labour market variables are jointly statistically significant, while the family related variables are not. Again, caution is in order. The summary statistics over end states, reported in Table 1, reveal that both males and females returning to more education are less likely to have a family on their own. For instance, only 11.8% of females in the end state *education* are married, while the corresponding number in all other states is above 25%. Given that few individuals taking further education have a family, it may come as no great surprise that family related variables are not significant explanatory factors in this transition. This conclusion is also borne out by the insignificant marginal effects of family related variables for the end state *education* in Table 2. However, for all states where the incidence of marriage is higher, family related variables are the most important factors explaining gender differences.

5. Conclusion

The results of our analysis can be interpreted in several directions. First, the industry dummies and the local unemployment rate are highly significant variables. This indicates that many exits may be due to insufficient demand in some sectors. Furthermore, our replacement rate proxy gives no indication of adverse incentive effects from the social insurance system. However, the available data are not particularly well suited to reveal such effects. Second, experience as well as higher income increase the probability of staying employed and decrease the probabilities of transiting to one of the three other states. One interpretation of these findings may be that it is not only important to get youths into a job to secure future labour market participation. It is also important to hold a good job with a reasonable high income to reduce the probability of future exits. This latter interpretation relies on exogenous income and experience variables. However, selectivity biases may operate here. If only the ablest of the youngsters manage to build experience in a well-paid job, the benign effect of these observable characteristics is overstated. In the most extreme case, placing less able workers in well-paid jobs with high demands on skills may even produce the opposite results, i.e. higher exit rates. Finally, the numbers of years of education increase the probability of staying employed, and decrease the probability of being unemployed or ending in the residual group. These results may be interpreted to support the clearly stated policy of the Norwegian labour market authorities to increase resources to education as a primary strategy to combat youth unemployment. However, it is not clear that the importance of education will be the same in a population of youths where everyone has a higher education level both due to selectivity bias and potential job competition.[6]

In recent years, the education pattern and labour market behaviour of young males and females in Norway have displayed a converging pattern. Notice, for instance, that the mean education levels reported in Table 1 are higher for females as compared to males in all destination groups. While the young males in this respect are lagging behind in educational attainments, traditional gender roles still matter for early labour market experiences. Family related variables, age, and experience still play a different role for males and females; and these differences are compatible with the hypothesis that females still respond more in their labour market decisions to changes within the family.

Endnotes

[1] The pension points function is defined as:

$$\text{pension points} = \begin{cases} \dfrac{i-g}{g} & \text{if } g \leq i \leq 8g \\ 7+\dfrac{1}{3}\dfrac{i-8g}{g} & \text{if } 8g < i \leq 12g' \\ 8.33 & \text{if } i > 12g \end{cases}$$

where

$g = \text{NOK } 37\,300$

$i = \text{income}.$

[2] The coefficients of the multinomial logit are available from the authors upon request.

[3] The marginal effects of an increased number of children for unmarried women are −0.003 and −0.009 (work and other, respectively).

[4] These results are not reported for sake of brevity, but are available from the authors upon request.

[5] The following model is estimated:

$$Prob(Y_i = j) = \frac{e^{\mathbf{b}_j'\mathbf{x}_i + \mathbf{b}_j^{diff'}(\mathbf{x}_i \cdot D_i^{females})}}{\sum_{j=0}^{3} e^{\mathbf{b}_j'\mathbf{x}_i + \mathbf{b}_j^{diff'}(\mathbf{x}_i \cdot D_i^{females})}},$$

where $\mathbf{b}_j^{diff}$ is a vector of parameters which give the difference in the importance of the relevant regressors, and $D_i^{females}$ is a dummy-variable taking the value of one for females and zero otherwise. The coefficient vector will be $\mathbf{b}_j$ for males and $(\mathbf{b}_j + \mathbf{b}_j^{diff})$ for females. We refer to the variables that are different for males and females, $\mathbf{x}_i \cdot D_i^{females}$, as the difference variables.

[6] Further analysis is needed to resolve this issue. A relevant recent study is van Ours and Ridder (1995).

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Parental disruption and the labour market performance of children when they reach adulthood

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Abstract. This paper uses data from the age 33 wave of the British National Child Development Survey (NCDS) to analyze the effects of a parental disruption (divorce or death of a father) on the labour market performance of children when they reach adulthood. The NCDS is a longitudinal study of all children born during the first week of March 1958 in England, Scotland, and Wales. Controlling for a rich set of pre-disruption characteristics, the results indicate that a parental disruption leads to moderately less employment among males and considerably lower wage rates among females at age 33. If pre-disruption characteristics are not controlled for, larger effects are estimated for both males and females. Parental disruption also seems to cause substantial reductions in educational attainment for both males and females.

JEL classification: J12, J22, J24

Key words: Marital disruptions, labour supply, educational attainment, wage rates

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1. Introduction

As is well known, the rise in parental marital disruptions in recent decades (i.e., a divorce or the death of a parent) has meant that children have become increasingly likely to spend at least part of their childhood living in a single-parent household.[1] There is considerable evidence suggesting that living in

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a disrupted family has short-run detrimental effects on children. However, comparatively little evidence exists concerning the longer-run effects of living arrangements on children's labour market performance when they reach adulthood. Knowledge of such longer-run effects is important because, if they exist, children who are subjected to a parental disruption will suffer a loss of income and society will lose productive output.

The purpose of this paper is to examine the effects of parental disruption on the labour market performance of children when they reach adulthood, using a database (the British National Child Development Survey, or NCDS) that follows individuals from birth to age 33. We investigate whether a parental disruption affects the amount of formal education obtained and whether it affects two labour market outcomes at age 33: hourly earnings and labour market status (i.e., whether individuals are employed, unemployed, or out of the labour force).

The remainder of this paper is organized as follows. Section 2 discusses the theoretical relationship between parental disruptions and adult labour market performance. Section 3 describes the NCDS and its advantages for examining the effects of parental disruption on labour market performance. Section 4 reviews previous empirical studies of the effect of parental disruptions on adult labour market performance. Section 5 describes the methodology used in the empirical analysis. Section 6 presents the findings. Finally, Sect. 7 offers some conclusions.

2. The relation between parental marriage disruptions and adult labour market performance

There are at least three reasons why a parental marital disruption might adversely affect the longer-term labour market performance of a child. First, income in the household in which the child grows up is likely to be lower, often much lower (Duncan and Hoffman 1985). Even if the absent parent is alive and provides child support, the economic cost of running two households, rather than one, reduces the amount of income available to the child. As implied by the household production model developed by Becker and others (see, for example, Becker 1981; Becker and Tomes 1986; and Michael and Becker 1973), a decline in available income will reduce investments in the child's human capital. Such a reduction in human capital may manifest itself as a decrease in formal education, because the child is forced to leave school earlier than normal as a result of funding constraints and the need to obtain full-time employment. It may also occur in other ways – for example, fewer private out-of-school lessons and living in less desirable neighborhoods with lower quality schools. These reduced investments in human capital may ultimately result in detrimental effects on the labour market performance of the child when he or she reaches adulthood.

A second adverse consequence is that a marital disruption tends to reduce the amount of time parents devote to their children. The absent parent will probably spend less time with his or her child, even if the cause of the disruption is divorce, rather than death, and he or she exercises full visitation rights. Furthermore, the custodial parent may spend additional time participating in the labour market and, in addition, have more responsibilities within the household. Consequently, the custodial parent may have less time to devote to the child. The household production model implies that these reduced time inputs, like the diminished income, will decrease the human capital of children – for example, there will be less opportunity for parents to serve as role models or to pass on their knowledge to their children.

Third, a divorce or a parent's premature death may subject a child to emotional stress that may negatively affect the child's labour market performance in later years. Emotional stress could be caused directly by the disruption itself or from the changes in lifestyle that result – for example, from residential moves and transfers to new schools.

The emotional stress that may result from a parental disruption has somewhat different implications for children's labour market performance as adults than do losses of income and time inputs. For example, the loss of income and time inputs may have larger effects if the disruption occurs at a younger, rather than an older, age simply because the loss of these inputs occurs over a longer time. It is not clear how the emotional stress that accompanies a marital disruption is related to the age at which the event occurs, but a longer passage of time between the event and adulthood may well allow for a fuller recovery. Furthermore, in the case of a divorce, there is evidence that non-custodial fathers maintain more contact with sons than with daughters (Hetherington et al. 1982) and provide more educational financing for sons than for daughters (Wallerstein and Corbin 1986). The emotional stress from the loss of a male role model in the home might be greater for boys, however. In addition, losses of income and time inputs are typically more severe if the disruption results from a parent's death, rather than from a divorce. Emotional stress, in contrast, could be greater as the result of a divorce, especially if the child is caught in the middle of a continuing struggle between the parents.

Being subjected to a family disruption does not necessarily have an unambiguously negative impact on the labour market performance of children as adults. As adults, individuals may well be able to overcome apparently devastating events that occurred earlier in their lives. In addition, time spent with some parents – for example, those that are active alcoholics or mentally and physically abusing – may exert negative, rather than positive, influences on children. Moreover, living in a household in which parents are in constant conflict may conceivably cause more emotional stress than living in a single-parent household. Furthermore, the impact of a parental disruption on the labour market performance of children may be mitigated to some degree by the fact that the missing parent was a consumer of household inputs (i.e., income and the remaining spouse's time), as well as a provider of inputs to his or her children. Thus, the effect of a parental disruption on the labor market performance of children when they reach adulthood becomes ultimately an empirical question.

3. Data

To investigate the effects of marital disruptions on the labour market performance of children when they reach adulthood, we use the National Child Development Survey (NCDS). The NCDS is a longitudinal study of all British children born during the first week of March 1958 in England, Scotland, and Wales. A total of 17,414 mothers, representing 98% of all births occurring that week, were interviewed (Shepherd 1985).[2] The original purpose of the survey was to identify social and obstetric factors associated with stillbirth and death in early infancy, but the survey was later expanded to measure other important economic and social phenomena associated with the interviewed parents and children.

The first three follow-up interviews were conducted when the children were ages 7 ($n = 15{,}458$), 11 ($n = 15{,}503$), and 16 ($n = 14{,}761$). These ages were chosen because they represented important transition points in the British school system for individuals born in 1958. On each occasion, wide-ranging information was collected from the parents, teachers, and school medical officers. In 1981, a professional survey research organization was used to trace and re-interview the children, who at this time were age 23. A sample of 12,537 young adults participated in the 1981 interview. All four interviews (at ages 7, 11, 16, and 23) collected detailed information on the socioeconomic and psychological characteristics of the parents and children.

The fifth and most recent follow-up interview was conducted in 1991, when sample members were age 33. Over 13,400 cohort members were traced, although only 11,407 interviews were ultimately conducted (Ferri 1993). (The 84% response rate among those traced to age 33 is reasonable given the comprehensiveness of the questionnaires and the great demands placed on the time of the respondents.) As is evident, the combined effects of adulthood and the lengthy period between interviews resulted in substantial sample attrition.[3] The fifth followup interview was the first to ask detailed questions about earnings and work experience. In addition, a question about parental disruption and the age at which a disruption occurred was asked of all respondents.

The NCDS provides researchers an unprecedented opportunity to learn about the effects of parental disruption on adult labour market performance, while being able to control for other factors that also affect labour market performance. Because of the long time span covered by the NCDS, parental disruptions at earlier stages of life and labour market performance at a later prime-age stage in life can both be observed. With the wealth of data collected by the NCDS, it is possible to control for children's psychological and intellectual functioning and family socioeconomic status prior to a parental disruption.[4] In addition, the size of the sample allows researchers to distinguish between disruptions that result from the death of the parent and those that stem from parental divorce and between impacts on sons and those on daughters. Finally, the NCDS is ideal for investigating whether the age at which the disruption occurred influences the size of effects resulting from the disruption. One disadvantage of the NCDS is that detailed information about family income is not available.

4. Previous studies

Numerous studies of the long-term effects of marital disruptions on children in both Britain and the United States have now been conducted. Most of these studies are concerned with divorce; studies of the long-term consequences of the death of a parent are comparatively rare, although sometimes disruptions that occur as a result of either divorce or death are combined into a single explanatory variable. Moreover, most previous studies focus on the effects of divorce on children's educational attainment. Until very recently, only a few studies examined effects on labour market outcomes. For example, in a meta-analysis of studies of the long-term consequences of parental divorce, Amato and Keith (1991) found 18 studies that examine educational attainment. In contrast, they did not find any studies of the effects of divorce on labour market status or earnings. However, they did uncover nine studies that examined outcomes that tend to be correlated with earnings – for example, occupational status, job autonomy, and job satisfaction. (Blau and Duncan's (1967) study of the relation between divorce, educational attainment, and occupational status was the first of these.) Moreover, they excluded studies by Mueller and Cooper (1986) that used unemployment status as an outcome, Wadsworth and Maclean (1986) that used income, and Greenberg and Wolf (1982) and Krein (1986), that used earnings.

Nonetheless, it seems evident that before the 1990s studies of the effects of divorce on children's labour market status and earnings were fairly rare. Perhaps stimulated by increases in divorce rates and improvements in available data, the number of such studies greatly expanded during the 1990s. Examples include Haveman et al. (1991), Peters (1992), Wojtkiewicz (1993), Garasky (1995), Grogger and Ronan (1995), Lillard and DeCicca (1995), Couch and Lillard (1995), Kiernan (1997), Francesconi and Ermisch (1998), Gregg and Machin (1998), and Boggess (1998).

Although they do not always obtain statistically significant findings, previous studies consistently imply that the effects of parental divorce on educational attainment, employment, unemployment, occupational quality, and earnings are adverse. However, the estimated effects are usually moderate. For example, Amato and Keith's sample of studies implied that a parental divorce caused educational attainment to fall by a bit over one-quarter of a standard deviation and occupational quality to fall by just over one-tenth of a standard deviation.[5] In addition, Amato and Keith found that studies that used control variables to take account of pre-divorce family economic status and parental characteristics obtained substantially smaller estimated effects than studies that did not use such controls.

The NCDS offers some important advantages over the data sets used in many previous studies of parental disruption on adult labour market performance. First, many (but not all) previous studies were based on limited information on family circumstances prior to the parental disruption. Therefore, it is difficult to determine whether estimates of detrimental effects on the child are linked to the disruption itself or merely to family and individual conditions that existed before the disruption.[6] For example, if earnings appear to be negatively related to parental divorce, it is important to distinguish whether this is due to the divorce event itself or due instead to having lived in an intact, but possibly, dysfunctional family prior to the divorce. The NCDS contains richer information than is available in most data sets covering individuals during their childhood, permitting one to better make such a distinction.

A second important advantage of the NCDS is that it follows children for a longer period of time than most panel studies. The data used in many (although far from all) previous studies only allow children to be observed during early adulthood, a time when their careers may still be in considerable flux. Even the education data for the sampled population is incomplete in some studies and, thus, the measure of educational attainment that is used is restricted to a simple indicator of high school graduation. More importantly, in studies that use earnings as an outcome, this variable is often measured while the individuals sampled were still in their 20s. (See, for example, Greenberg and Wolf 1982, Peters 1992, Grogger and Ronan 1995, and Lillard and DeCicca 1995.) Earnings this early in the life cycle may be only weakly related to earnings during prime-age. Moreover, the effects of a parental disruption may fade over time and, thus, be stronger in one's 20s than later in life. As previously indicated, the NCDS measures the educational attainment, earnings, and labour market status of individuals at age 33.

A third advantage of the NCDS is that it allows investigation of whether the age at which the disruption occurred influences the size of effects resulting from the disruption. We found only two non-NCDS studies that do this.[7] Lillard and DeCicca (1995), who used U.S. data, compare the effects of marital disruption on the earnings of two groups of male children: those for whom the disruption occurred between ages 14 and 21 and those for whom the disruption occurred between ages 22 and 25. They find larger effects for the younger group. Francesconi and Ermisch (1998), who used British data, distinguish among three groups of children who spent at least part of their childhood in a single parent family: those who began their spell in a single parent family between the ages of zero and 5, those who began between 6 and 10, and those who began between 11 and 16. They examine a number of outcomes, including educational attainment and unemployment, and find evidence that children who lived in a single parent family early in life experienced the largest adverse effects as adults.

Given the advantages of the NCDS for examining the effects of parental disruption on education and adult labour market outcomes, it is not surprising that two studies in addition to ours have used the NCDS for this purpose: Gregg and Machin (1998) and Kiernan (1997).[8] However, our study focuses exclusively on education and labour market outcomes, while the other two studies examine additional outcomes as well. Furthermore, our study places greater emphasis on formal tests of whether competing model specifications are statistically different than do the other two studies.

The objectives of Gregg and Machin's 1998 study differ from those of our study. They are concerned with how a broad range of factors associated with childhood disadvantage, only one of which is parental marital disruption, affect various outcomes at ages 16, 23, and 33. Thus, while we examine whether the effects of parental disruption vary with the age at which it occurs and with whether it results from divorce or death, they do not. Consequently, they estimate the effects of disruption using a single dummy variable, while we use a number of variables to distinguish among divorce, death of a parent, and the age of the disruption. To construct their disruption variable, they assigned a value of one to individuals who ever lived in a single parent family before age 16, but whose families never faced economic difficulty during this period, and a value of zero to all other persons.[9] Thus, effects that are estimated with this variable pertain only to persons who experienced parental disruption, but not economic hardship, as children. Such individuals may be atypical products of disruption.

Gregg and Machin find that these individuals had a lower level of educational attainment than other persons did. They also find that males (but not females) who lived in single parent families as children had higher unemployment at age 23 and received lower hourly wages at age 33 (but not at age 23).

Kiernan (1997) focuses exclusively on parental divorce, while we examine parental death, as well as divorce. Her objectives, however, are similar to ours. Like us, she is interested in whether disadvantages that children of divorce have as adults are attributable to the divorce itself or to factors that antecede the disruption. In addition, both her study and our study examine whether the age at which a parental divorce occurs influences the size of effects of divorce on children as adults. Unlike our study, however, she does not use formal statistical tests to examine this issue. In conducting her study, Kiernan first constructs several dichotomous measures of educational and labour market outcomes – specifically, whether at age 33 individuals have no educational qualifications, have high level qualifications, were unemployed, and have earnings in the upper quartile. She then runs logistic regressions on these measures and reports the resulting odds ratios. As will be seen, we use a somewhat different specification of educational and labor market outcomes.

Kiernan finds that children whose parents divorced before they were 16 are worse off than other adults at age 33 in terms of most (but not all) the education and labour market outcomes she examines. However, she also finds that these effects usually diminish, or even disappear, when family circumstances that existed before the divorce, especially financial circumstances, are taken into account. This suggests that the children of divorce not only fare worse because of the divorce itself, but also because they were worse off before the divorce occurred. Finally, Kiernan finds that parental divorce occurring after men were 21 is associated with reduced levels of educational attainment and increased levels of unemployment. Similar associations were not found for women, however. As will be seen, there are both important similarities and differences between our findings and Kiernan's.

5. Methodology

5.1. Outcomes analyzed

As indicated earlier, we analyze two labour market outcomes at age 33 – gross hourly earnings (or wage rates) and labour market status.[10] We also examine the effects of parental disruption on educational attainment. The means of these outcomes are presented in Table 1. Following previous studies, as well as the previously mentioned evidence that income and time input losses from divorce are greater for girls than for boys, separate analyses of the outcomes are conducted for males and females. (Previous research by Cherlin et al. (1991) found that the effect of parental divorce or separation on behavioral and educational outcomes differed for boys and girls.)

We use the multinomial logit model to estimate the effects of parental disruption on adult labour market status. A 3-choice model is estimated for males (employed, unemployed, and not in the labour force), and a 4-choice model is estimated for females (employed full-time, employed part-time, unemployed, and not in the labour force). According to Table 1, 90.6% of the males in our sample were employed at age 33, 6.6% were unemployed, and 2.8% were not in the labour force. Less than 1% of the males was employed part-time (not shown in Table 1) and that is why we do not distinguish between full-time and part-time employment for males. Females were much more likely than males to be employed part-time and they were much more likely to be out of the labour force. Specifically, 35.3% of the women were employed full-time, 31.2% were employed part-time, 2.3% were unemployed, and 31.2% were not in the labour force.

Table 1. Definitions and means of outcome variables

Variable	Definition	Males	Females
LNWAGE	Gross Log of Hourly Wages (in 1991 British pounds)	1.97	1.61
	Full-time workers	–	1.79
	Part-time workers	–	1.39
		($n = 3{,}003$)	($n = 2{,}448$)
Work status	= 1 if employed	90.6%	66.5%
	= 1 if employed full-time		35.3%
	= 1 if employed part-time		31.2%
	= 1 if unemployed	6.6%	2.3%
	= 1 if not in labor force	2.8%	31.2%
		($n = 3{,}662$)	($n = 4{,}166$)
Education[a]	= 0 if no qualification (If NVQ0 = 1)	11.1%	13.0%
	= 1 if some qualification (If NVQ1 = 1)	10.9%	13.8%
	= 2 if O level or equivalent (If NVQ2 = 1)	29.0%	37.7%
	= 3 if A level or equivalent (If NVQ3 = 1)	18.6%	10.1%
	= 4 if other higher education (If NVQ4 = 1)	15.4%	14.6%
	= 5 if higher education (If NVQ5 = 1)	15.0%	10.9%
		($n = 3{,}603$)	($n = 4{,}129$)

Source: National Child Development Survey, 1991, Wave 5.
[a] See appendix for definitions of NVQ variables.

The effects on parental disruption on adult hourly wages were estimated for only those individuals who were working and not self employed at the time of the age 33 survey. A variable was included in the OLS wage regression model to correct for the selectivity bias that results from this sample restriction. This variable was computed from the estimated multinomial logit equations described in the previous paragraph.[11]

By age 33 almost all individuals in the sample had completed their formal education. A categorical variable is used to measure educational attainment and an ordered probit model is used to estimate the effects of parental disruption on educational attainment.[12]

5.2. Modeling strategy

Our basic approach is to regress the educational and labour market outcome variables described above against indicators of whether a parental disruption occurred when individuals were children. In addition, most of the regression models include measures of childhood behavior (childhood psychological status, family socioeconomic status while growing up, school performance, and so-forth) that were obtained at the individuals' birth and when they were 7 years of age. These birth and age 7 variables are described in the following subsection. They are included in the regressions to control for children's psychological and intellectual functioning and family socioeconomic status prior to age 7, as these factors may be correlated with both parental disruption and children's labour market performance as adults.

This modeling strategy raises a serious issue that must be dealt with in estimation. Parental disruptions occur when children are at different ages. As a consequence, variables in the NCDS are measured prior to the disruption for some individuals and subsequent to the disruption for others. Compare, for example, a 33-year-old male whose parents divorced before he was age 7 with one whose parents divorced between the time he was ages 7 and 11. The childhood psychological and socioeconomic status variables at age 7 will be measured post-divorce for the first individual, but pre-divorce for the second individual. Thus, the age 7 childhood status variables for the first individual will probably be influenced by the divorce event, but the age 7 childhood status variables for the second individual should not be influenced by the divorce event (although they could be influenced by factors that caused the divorce). Consequently, if both individuals are included in the same regression, it will be extremely difficult to interpret the estimated coefficients on these variables.

We employ an estimation strategy that circumvents this endogeneity problem. In many of the regression specifications we report, the sample is restricted to only individuals whose parents were intact at age 7.[13] Because the control variables are measured at birth and at age 7, they can be included in these regressions without introducing endogeneity into the estimated coefficients.[14] Other regressions that we report, however, are estimated for a sample that includes individuals whose parents were not intact at birth or at age 7. We avoid endogeneity in these regressions by limiting the independent variables to those that are truly exogenous. For example, only disruption variables are included in regressions based on samples that include individuals whose parents were not intact at birth; other control variables are omitted. Similarly, only variables measured at birth in addition to the disruption variables are used as explanatory variables in regressions based on samples including individuals whose parents were intact at birth, but not necessarily at age 7.

The main sample used to estimate labour market status at age 33 consists of 4,166 females and 3,662 males. This sample includes all individuals interviewed at age 33 whose parents were intact at age 7 and who had valid data on parental divorce and death, hourly earnings, and work status.[15] The wage equations are estimated for a sub-group of this sample – namely, those who worked at age 33. To minimize the missing data problem for explanatory variables other than the disruption variables, each explanatory variable in our regressions is paired with a dummy variable equaling 1 whenever the explanatory variable has a missing value, and 0 otherwise. The explanatory variables are set equal to 0 whenever their value is missing.[16]

In a few specifications of the labour market outcome regressions, education is included as an explanatory variable. Because education is measured at age 33, it is likely to be endogenous with respect to parental disruptions occurring earlier. The purpose of estimating these models is to determine the extent to which the effects of parental disruptions on labour market outcomes are attributable to the effects of disruption on educational attainment. Because of the potential for endogeneity, these results should be interpreted with caution. (Estimating a simultaneous equation model that takes into account the endogeneity of education is beyond the scope of the present paper. As Francesconi and Ermisch (1998) demonstrate, finding a suitable instrumental variable for education to enable identification is difficult.)

5.3. Explanatory variables

Summary statistics for the explanatory variables are given in Table 2. A description of the construction of selected explanatory variables is given in the Appendix (for further details, see Cherlin et al. 1995). The explanatory variables are of three types: parental disruption variables, control variables measured at birth, and control variables measured at age 7.

The parental disruption variables, which are the major variables of interest in this study, consist of a set of dummy variables denoting whether a parental divorce or father's death[17] occurred at various ages (DV0, DV06, DV710, DV1115, DV1622, DV2333, DTH015, and DTH1622).[18] In addition, we also estimate partially constrained models in which a single divorce dummy (whether the individual's parents ever divorced, DV) and a single death dummy (whether the individual's father ever died, DTH) are used in place of the age-specific dummies. Finally, we also estimate fully constrained models that replace the DV and DTH variables with a single disruption dummy (DISRUPT), which indicates whether a disruption occurred, regardless of the age at which it occurred or whether it was due to a parental divorce or a father's death.

The control variables measured at birth consist of the age of the mother (MOMAGE), the age difference between the mother and the father (AGEDIFF), the birth weight of the child (BIRTHWGT),[19] the time between the parent's marriage and the mother's first birth (INTERVAL1-INTERVAL4), whether the mother was single at first pregnancy (INTERVAL0), the maternal grandfather's social class (MGCLASS1, MGCLASS2, MGMINER), whether the maternal grandfather was dead or away (MGDEAD), the number of persons in the household per room (PERSPROOM), and the occupational social class of the father (FSOC-CLASS1 and FSOC-CLASS2). In addition, there is a set of variables denoting whether the mother worked during pregnancy and, if she did, the number of weeks she worked (PREGNOWORK, PREGWORK1, PREGWORK2, and PREGWORK3). The age 7 variables consist of measures of behavioral problems (BEHPROB7), reading achievement scores (READTEST7), arithmetic achievement scores (MATHTEST7), and school performance (SCHPER7). The age 7 variables also include measures that denote whether the mother's or father's education exceeds the minimum (MOMEDUC and DADEDUC), the age the father left school (AGELEFT), and the birth order of the child (YOUNG2, OLDEST2, YOUNG3, MIDDLE3, OLDEST3).

6. Results

Seven different models are estimated to determine the effects of parental disruption on education and labour market outcomes. One of the models uses the full sample of families at birth, including both intact and non-intact families,

Table 2. Definitions and means of explanatory variables[a]

Variable	Definition	Sample of individuals in intact families at age 7		Sample of employed individuals in intact families at age 7	
		Males	Females	Males	Females
		($n = 3{,}662$)	($n = 4{,}166$)	($n = 3{,}003$)	($n = 2{,}448$)
Disruption variables					
DV710	= 1 if parents divorced between ages 7 & 10	0.02	0.02	0.02	0.02
DV1115	= 1 if parents divorced between ages 11 & 15	0.04	0.04	0.04	0.04
DV1622	= 1 if parents divorced between ages 16 & 22	0.03	0.04	0.03	0.04
DV2333	= 1 if parents divorced between ages 23 & 33	0.02	0.02	0.02	0.02
DTH015	= 1 if father died before age 16	0.03	0.03	0.03	0.03
DTH1622	= 1 if father died between ages 16 & 22	0.05	0.05	0.05	0.05
DV	= 1 if parents divorced	0.11	0.12	0.11	0.12
DTH	= 1 if father died	0.08	0.08	0.08	0.08
DISRUPT	= 1 if parents divorced of father died	0.19	0.21	0.19	0.20
Variables measured at birth					
MOMAGE	age of mother at birth	27.78	27.80	27.78	27.72
AGEDIFF	difference between mothers and fathers age at birth	2.95	2.86	2.90	2.89
MISSDIFF	= 1 if AGEDIFF is missing	0.01	0.01	0.01	0.01
BRTHWGT	birth weight (in ounces)	118.80	113.93	118.81	114.58
MISSBRWT	= 1 if BRTHWGT is missing	0.00	0.00	0.00	0.00
INTERVAL1	= 1 if 1–2 year interval between marriage and first birth	0.25	0.27	0.25	0.27
INTERVAL2	= 1 if 2–3 year interval between marriage and first birth	0.13	0.11	0.13	0.11
INTERVAL3	= 1 if 3–4 year interval between marriage and first birth	0.07	0.07	0.08	0.07
INTERVAL4	= 1 if 4–8 year interval between marriage and first birth	0.14	0.13	0.14	0.13
INTERVAL0	= 1 if single at first pregnancy	0.06	0.06	0.06	0.06
INTMISS	= 1 if interval between marriage and first birth is missing	0.02	0.02	0.02	0.02
MGCLASS1	= 1 if maternal grandfather is in Social Class II or III	[b]	[b]	0.55	0.55
MGCLASS2	= 1 if maternal grandfather is in Social Class IV or V	[b]	[b]	0.22	0.23
MGDEAD	= 1 if maternal grandfather is dead or away	[b]	[b]	0.08	0.08
MGMINER	= 1 if maternal grandfather is a miner	[b]	[b]	0.07	0.06

Table 2 (continued)

Variable	Definition	Sample of individuals in intact families at age 7		Sample of employed individuals in intact families at age 7	
		Males	Females	Males	Females
		($n = 3{,}662$)	($n = 4{,}166$)	($n = 3{,}003$)	($n = 2{,}448$)
MGMISS	= 1 if maternal grandfather social class is missing	[b]	[b]	0.04	0.04
PERSPROOM	number of persons in household per room at birth	1.42	1.47	1.39	1.45
ROOMMISS	= 1 if AOPHH is missing	0.02	0.02	0.02	0.02
FSOC-CLASS1	= 1 if father's social class is professional or intermediate	[b]	[b]	0.19	0.18
FSOC-CLASS2	= 1 if father's social class is skilled or partly skilled	[b]	[b]	0.72	0.73
FSOC-MISS	= 1 if father's social class is missing	[b]	[b]	0.02	0.02
PREGWORK1	= 1 if mother worked 13–24 weeks during pregnancy	0.09	0.08	[b]	[b]
PREGWORK2	= 1 if mother worked 25–36 weeks during pregnancy	0.18	0.17	[b]	[b]
PREGWORK3	= 1 if mother worked 37 or more weeks during pregnancy	0.01	0.01	[b]	[b]
PREGNOWORK	= 1 if mother did not work during pregnancy or missing	0.68	0.70	[b]	[b]
Variables measured at age 7					
BEHPROB7	factor score for behavioral problems at age 7	0.28	0.28	0.28	0.29
BPMISS	= 1 if BEHPROB7 is missing	0.01	0.01	0.01	0.01
READTEST7	reading test score at age 7	[b]	[b]	22.93	24.70
READMISS	= 1 if READTEST7 is missing	[b]	[b]	0.03	0.03
MATHTEST7	arithmetic test score at age 7	[b]	[b]	5.30	5.15
MATHMISS	= 1 if MATHTEST7 is missing	[b]	[b]	0.04	0.03
SCHPER7	school performance scale	[b]	[b]	14.74	15.28
MISSSCH1	= 1 if SCHPER7 is missing	[b]	[b]	0.03	0.02
MOMEDUC	= 1 if mother stayed past minimum school age	0.25	0.25	0.26	0.26
MOMMISS	= 1 if MOMEDUC is missing	0.01	0.01	0.01	0.01
DADEDUC	= 1 if father stayed past minimum school age	0.23	0.24	0.24	0.24
DADMISS	= 1 if DADEDUC is missing	0.04	0.04	0.03	0.04
AGELEFT	Age father left school	3.92	4.06	4.09	4.13
MISSLEFT	= 1 if AGELEFT is missing	0.78	0.77	0.77	0.77

Table 2 (continued)

Variable	Definition	Sample of individuals in intact families at age 7		Sample of employed individuals in intact families at age 7	
		Males	Females	Males	Females
		($n = 3{,}662$)	($n = 4{,}166$)	($n = 3{,}003$)	($n = 2{,}448$)
YOUNG2	Child is youngest of 2 children	0.18	0.18	0.18	0.18
OLDEST2	Child is oldest of 2 children	0.19	0.18	0.20	0.19
YOUNG3	Child is youngest of 3 or more children	0.14	0.15	0.14	0.14
MIDDLE3	Child is between youngest and oldest of 3 or more children	0.29	0.29	0.27	0.28
OLDEST3	Child is oldest of 3 or more children	0.12	0.11	0.12	0.11
Variables measured at age 33					
NVQ0	= 1 if no qualifications (omitted category)	0.11	0.13	0.08	0.10
NVQ1	= 1 if some qualifications	0.11	0.14	0.10	0.13
NVQ2	= 1 if O level qualifications or equivalent	0.28	0.37	0.29	0.37
NVQ3	= 1 if A level qualifications or equivalent	0.18	0.10	0.19	0.10
NVQ4	= 1 if other higher education	0.15	0.14	0.16	0.16
NVQ5	= 1 if higher education	0.15	0.11	0.16	0.13
MISSNVQ	= 1 if education is missing	0.02	0.01	0.02	0.01

Source: National Child Development Survey, 1991, Wave 5

[a] Sample of intact families at age 7 for whom valid labor market data are available at age 33. See appendix for detailed description of selected variables.

[b] Variable excluded from equation explaining this outcome.

two of the models use the sample of intact families at birth, and four of the models use the sample of intact families at age 7. (Non-intact families at birth are those in which the parents were never married or the parents separated or divorced prior to the birth of the child.) The samples and model specifications are as follows:

1. All families at birth, disruption variables are the only explanatory variables;
2. Intact families at birth, disruption variables are the only explanatory variables;
3. Intact families at birth, disruption variables plus birth variables are the explanatory variables;
4. Intact families at age 7, disruption variables only;
5. Intact families at age 7, disruption variables plus birth variables are the explanatory variables;
6. Intact families at age 7, disruption variables plus birth variables plus age 7 variables are the explanatory variables;

7. Intact families at age 7, disruption variables plus birth variables plus age 7 variables plus age 33 education variables are the explanatory variables.

A comparison of Models 1 and 2 provides an indication of whether excluding non-intact families at birth alters the effects of a parental disruption. A comparison of Models 2 and 3 provides an indication of whether controlling for pre-disruption characteristics at the time of the birth of the child alters the independent effects of disruption. A comparison of Models 2 and 4 provides an indication of whether the effects of parental disruption on labour market outcomes are altered by restricting the sample to families that were intact at age 7. A comparison of Models 4 and 5 provides an indication of whether controlling for pre-disruption characteristics at the time of the birth of the child alters the independent effects of disruption on the age 7 intact sample. A comparison of Models 5 and 6 provides an indication of whether controlling for pre-disruption characteristics between the time of birth and age 7 alters the independent effects of disruption on the age 7 intact sample. Finally, a comparison of Models 6 and 7 provides an indication of whether the effects of parental disruption on labour market outcomes are mitigated by effects on education. Of course, when education is the outcome, Model 7 is not estimated.

As indicated above, the effects of disruption are captured by a series of dummy variables representing the age range during which a disruption occurred. Specifically, in keeping with the survey design, we distinguish among divorces occurring between the ages of zero and 6, 7 and 10, 11 and 15, 16 and 22, and 23 and 33. Because there are few deaths of a father prior to age 15, we only distinguish between deaths occurring from ages 0 to 15 and from ages 16 to 22 (recall that the NCDS does not contain information about the death of a father after the child reaches age 22). As indicated earlier, much of the sociological and psychological literature suggests that the size of the effects of parental disruption depends on whether the disruption occurs in early or late childhood, the type of disruption (divorce or death), and the sex of the child (Cherlin et al. 1991). We explicitly test for differences in the effects of disruption by age and type of disruption, but we maintain separate analyses for males and females.

6.1. Effects on educational attainment

Table 3 presents the estimated effects of disruption on educational attainment at age 33. Both the fully unconstrained and the constrained disruption coefficients are reported. Chi-square statistics for the test of equality of the disruption coefficients across the various age and disruption categories are also reported. The full set of coefficient estimates for the unconstrained version of Model 6 is presented in Appendix Tables A.1 (for males) and A.2 (for females).[20]

As Table 3 indicates, males and females that grew up in a family that experienced a marital disruption have significantly lower levels of educational attainment. Even after controlling for the effects of individual characteristics measured at birth and at age 7 (Model 6), both males and females who experienced a marital disruption have lower levels of education. In the case of males, this effect is reduced by about three fifths after controlling for the birth and age 7 variables (Model 4 versus Model 6). Although in Models 4–6 we cannot reject the hypothesis of a constant disruption effect across disruption types, most of the effect for males appears to be due to a parental divorce that occurred between the ages of 7 and 15.[21] A statistically significant effect of a father's death (almost equal in magnitude to the effect of a parental divorce) becomes statistically insignificant after controlling for family circumstances at age 7 (e.g., comparing Model 6 with Model 5). The results for the effects of disruptions subsequent to age 7 do not appear to be sensitive to restricting the estimation sample to families intact at age 7. Moreover, it appears that divorce at younger ages (but subsequent to birth) has an effect similar to that between ages 7 and 10.

In the case of females, the effect is reduced by about two fifths when we control for pre-disruption characteristics (Models 4 and 6). Unlike males, we can reject the hypothesis of a constant disruption effect across age and disruption type for females for all models except Model 6. As in the case of males, most of the effect for females is due to a divorce that occurred before age 16. In addition, the effect of the death of a father prior to age 16 is also statistically significant. Like males, the results for females do not appear to be very sensitive to restricting the sample to intact families at age 7. However, the effects of divorce between the ages 0 and 6 appear to be slightly larger than the effects of a divorce between the ages of 7 and 15. Unlike males, the effect of a father's death remains for females even after controlling for family circumstances at age 7.

Table 4 uses the ordered probit coefficients in the unconstrained specification of Models 1–6 to determine which levels of education are affected by a parental divorce occurring between the ages of 7 and 10 and 11 and 15 (the coefficients are statistically significant for both males and females). As all six model specifications reported in the table indicate, parental divorce causes an increase in the probability of not going beyond the lower education levels (that is, levels 0–2; see the Appendix for definitions) for both males and females and a decrease in the probability of reaching the higher education level (levels 3–5). For example, in the case of a divorce between the ages of 7 and 10 for Model 6, there is about a 5-percentage point lower probability of attaining education level 5 for males and about a 4 percentage point lower probability of attaining the same education level for females (first degree, postgraduate diploma, masters, Ph.D.). As indicated by Table 2, this is about a one-third reduction for both males and females. Clearly, a parental divorce occurring before age 16 for boys and girls has substantial adverse consequences on ultimate educational attainment.

Our finding that a parental disruption adversely affects educational attainment is consistent with the results of previous studies. The fact that the effect for both males and females is substantially reduced when we control for pre-disruption family circumstances illustrates the importance of accounting for such family circumstances when drawing inferences about the effects of a parental disruption on educational attainment. In this sense, our results agree with those of Kiernan (1997), who finds that families that experience a divorce are generally worse off before the disruption occurred. Unlike most previous studies, we have examined the effect of a father's death on educational attainment. We find that females, but not males, appear to be adversely affected by a father's death, particularly when the death occurs before the child's 16th birthday.

Table 3. Ordered probit estimates on education at age 33. Various specifications for males and females. (Standard errors in parentheses)

	1		2		3		4		5		6	
	b	se(b)	b	se(b)	b	se(b)	b	se(b)	b	se(b)	b	se(b)
Males												
1												
DV0	−0.30***	(0.10)										
DV06	−0.52***	(0.11)	−0.52***	(0.11)	−0.43***	(0.11)						
DV710	−0.53***	(0.16)	−0.53***	(0.16)	−0.43***	(0.15)	−0.53***	(0.16)	−0.43***	(0.15)	−0.32**	(0.15)
DV1115	−0.25***	(0.09)	−0.25***	(0.09)	−0.20**	(0.09)	−0.25***	(0.09)	−0.20**	(0.09)	−0.23***	(0.09)
DV1622	−0.08	(0.09)	−0.08	(0.09)	−0.04	(0.10)	−0.08	(0.09)	−0.04	(0.10)	0.03	(0.10)
DV2333	−0.31***	(0.11)	−0.31***	(0.11)	−0.19*	(0.10)	−0.31***	(0.11)	−0.19*	(0.11)	−0.10	(0.11)
DTH015	−0.30***	(0.08)	−0.30***	(0.08)	−0.14*	(0.09)	−0.30***	(0.09)	−0.12	(0.09)	−0.07	(0.10)
DTH1622	−0.27***	(0.08)	−0.27***	(0.08)	−0.17**	(0.08)	−0.27***	(0.08)	−0.17**	(0.08)	−0.06	(0.08)
2												
DV	−0.30***	(0.05)	−0.30***	(0.05)	−0.23***	(0.05)	−0.25***	(0.05)	−0.18***	(0.06)	−0.13**	(0.06)
DTH	−0.28***	(0.06)	−0.28***	(0.06)	−0.16***	(0.06)	−0.28***	(0.06)	−0.15**	(0.06)	−0.06	(0.07)
3												
Disrupt	−0.29***	(0.04)	−0.29***	(0.04)	−0.20***	(0.04)	−0.26***	(0.04)	−0.17***	(0.04)	−0.10**	(0.05)
Chi-square (1 & 2)	12.6*		12.6*		9.9		7.5		5.6		6.0	
Chi-square (1 & 3)	12.6**		12.6**		10.8*		7.6		5.7		6.6	
Chi-square (2 & 3)	0.1		0.1		0.9		0.1		0.2		0.6	
n	3,853		3,743		3,743		3,603		3,603		3,603	
Females												
1												
DV0	−0.27***	(0.09)										
DV06	−0.50***	(0.09)	−0.50***	(0.09)	−0.43***	(0.09)						
DV710	−0.39***	(0.10)	−0.39***	(0.10)	−0.32***	(0.10)	−0.40***	(0.10)	−0.32***	(0.10)	−0.24**	(0.10)
DV1115	−0.43***	(0.09)	−0.43***	(0.09)	−0.36***	(0.09)	−0.43***	(0.09)	−0.36***	(0.09)	−0.31***	(0.09)
DV1622	−0.13	(0.09)	−0.13	(0.09)	−0.03	(0.09)	−0.13	(0.09)	−0.03	(0.09)	−0.05	(0.09)
DV2333	−0.11	(0.10)	−0.11	(0.10)	−0.05	(0.11)	−0.11	(0.10)	−0.05	(0.11)	0.04	(0.11)
DTH015	−0.23***	(0.08)	−0.23***	(0.08)	−0.17**	(0.08)	−0.20**	(0.09)	−0.18*	(0.10)	−0.19**	(0.10)
DTH1622	−0.09	(0.07)	−0.09	(0.07)	−0.07	(0.07)	−0.09	(0.07)	−0.07	(0.07)	−0.08	(0.07)

2												
	−0.31***	(0.04)	−0.32***	(0.04)	−0.24***	(0.05)	−0.27***	(0.05)	−0.19***	(0.05)	−0.15***	(0.05)
DV												
DTH	−0.16***	(0.05)	−0.16***	(0.05)	−0.11**	(0.06)	−0.13**	(0.06)	−0.11*	(0.06)	−0.12**	(0.06)
3												
Disrupt	−0.26***	(0.03)	−0.25***	(0.04)	−0.19***	(0.04)	−0.21***	(0.04)	−0.16***	(0.04)	−0.14***	(0.04)
Chi-square (1 & 2)	16.8**		16.5**		16.7**		10.6		11.4*		9.2	
Chi-square (1 & 3)	22.3***		22.2***		20.2***		14.1**		12.6*		9.3	
Chi-square (2 & 3)	5.5		5.7		3.5		3.4		1.2		0.1	
n	4,481		4,325		4,325		4,129		4,129		4,129	

* significant at 0.10 level; **significant at 0.05 level; ***significant at 0.01 level.
Samples and specifications are as follows:
(*1*) All families at birth, disruption variables only.
(*2*) Intact families at birth, disruption variables only.
(*3*) Intact families at birth, disruption variables plus birth variables.
(*4*) Intact families at age 7, disruption variables only.
(*5*) Intact families at age 7, disruption variables plus birth variables.
(*6*) Intact families at age 7, disruption variables plus birth and age 7 variables.

Table 4. Marginal effects of parental disruption on probability of achieving particular education levels. Unconstrained model, various specifications for males and females. For a parental divorce occuring between ages 7 and 10, and between ages 11 and 15

	1	2	3	4	5	6
Males						
Effect **DV710** on Probability of						
EDUC = 0	0.100	0.100	0.071	0.098	0.069	0.035
EDUC = 1	0.057	0.057	0.050	0.057	0.050	0.042
EDUC = 2	0.052	0.053	0.050	0.055	0.051	0.050
EDUC = 3	−0.028	−0.027	−0.026	−0.026	−0.025	−0.025
EDUC = 4	−0.063	−0.063	−0.057	−0.062	−0.056	−0.050
EDUC = 5	−0.119	−0.119	−0.088	−0.121	−0.089	−0.052
Effect **DV1115** on Probability of						
EDUC = 0	0.048	0.047	0.033	0.047	0.032	0.026
EDUC = 1	0.027	0.027	0.023	0.027	0.023	0.031
EDUC = 2	0.025	0.025	0.023	0.026	0.024	0.036
EDUC = 3	−0.013	−0.013	−0.012	−0.012	−0.012	−0.018
EDUC = 4	−0.030	−0.030	−0.027	−0.029	−0.026	−0.037
EDUC = 5	−0.056	−0.057	−0.041	−0.058	−0.042	−0.038
n	3,853	3,743	3,743	3,603	3,603	3,603
Females						
Effect **DV710** on Probability of						
EDUC = 0	0.086	0.085	0.061	0.083	0.059	0.032
EDUC = 1	0.046	0.046	0.042	0.047	0.042	0.037
EDUC = 2	0.014	0.015	0.014	0.017	0.016	0.015
EDUC = 3	−0.021	−0.021	−0.020	−0.021	−0.019	−0.018
EDUC = 4	−0.053	−0.053	−0.047	−0.053	−0.047	−0.039
EDUC = 5	−0.071	−0.072	−0.050	−0.073	−0.051	−0.026
Effect **DV1115** on						
EDUC = 0	0.092	0.092	0.069	0.089	0.067	0.042
EDUC = 1	0.050	0.050	0.047	0.051	0.048	0.048
EDUC = 2	0.015	0.016	0.016	0.018	0.018	0.020
EDUC = 3	−0.022	−0.023	−0.022	−0.022	−0.022	−0.024
EDUC = 4	−0.057	−0.057	−0.053	−0.057	−0.053	−0.051
EDUC = 5	−0.077	−0.077	−0.057	−0.079	−0.057	−0.035
n	4,481	4,325	4,325	4,129	4,129	4,129

Derived from ordered probit results. See the appendix for definitions of the education levels. Samples and specifications are as follows:
(*1*) All families at birth, disruption variables only.
(*2*) Intact families at birth, disruption variables only.
(*3*) Intact families at birth, disruption variables plus birth variables.
(*4*) Intact families at age 7, disruption variables only.
(*5*) Intact families at age 7, disruption variables plus birth variables.
(*6*) Intact families at age 7, disruption variables plus birth and age 7 variables.

6.2. Effects on labour market outcomes

Tables 5 and 6 present the effects of a parental disruption on the labour force status of males and females, respectively. The tables present the multinomial logit coefficients for the disruption variables, their standard errors, an indication of their statistical significance, and their marginal effects (because the multinomial logit model is nonlinear, the coefficients do not represent marginal effects). The marginal effects are the change in the outcome associated with a parental disruption. The full set of coefficient estimates for Model 6 in the unconstrained version of the model is presented in Appendix Tables A.1 (for males) and A.2 (for females).[22]

For males, the results indicate a statistically significant effect of divorce and death on unemployment at age 33, but the effect becomes statistically insignificant after adjusting for the effects of the age 7 control variables (Model 6). In the fully unconstrained specification, the largest effects in Model 6 on unemployment are for a father's death occurring between the ages of 16 and 22 and for a parental divorce occurring between the ages of 23 and 33 (although neither effect is statistically significant in Model 6). These adverse effects arising from later disruptions are consistent with the results of Kiernan (1997), who finds greater unemployment among men whose parents divorced after the child reaches the age of 21 (Kiernan doesn't examine the effects of a father's death).

There is also an effect for males of divorce (but not of death) on being out of the labour force. Unlike the effect on unemployment, the effect of divorce on being out of the labour force persists through each model specification. However, the effect declines by about one-quarter from Model 1 to Model 6. It doesn't appear that further controlling for education level alters the disruption effects for either males or females (Model 7 versus Model 6). Thus, the endogeneity of education that has plagued researchers using other data sets does not appear to be as serious a problem for researchers using the NCDS, perhaps because key economic and demographic characteristics during childhood can be controlled for in the estimated model.

The chi-square tests for Model 6 (which is our preferred model) indicate that the fully constrained specification (a single disruption dummy variable) fits best. Overall, the results suggest that a parental divorce prior to age 15 increases the probability of a male being out of the labour force by about one percentage point, or an increase of about one-third.

For females, there are few significant effects of a parental disruption on labour force status, about what would be expected by chance alone. In Model 6, there is only one statistically significant coefficient across the three labour force statuses. In Model 7, which controls for education, there is a significant *negative* effect of a parental disruption on the probability of part-time employment. The chi-square statistics indicate that the fully unconstrained model fits best for females. Although none of the disruption measures have a statistically significant effect on unemployment in Model 6, Models 1, 2, 4, and 5 indicate that a divorce occurring between the ages of 7 and 10 significantly increases the probability of being unemployed at age 33. However, the effect becomes statistically insignificant for Model 6. Overall, there is only very weak evidence that a parental disruption adversely affects the labour force status of females at age 33.

Table 7 presents the selectivity-corrected estimates of the effects of parental disruption on the hourly wage rates of males and females. For males, effects are presented for all workers, while for females, separate effects are presented for full-time and part-time workers. The tables also report the estimated coefficients of the selectivity-correction terms and results of tests for the constraints on the disruption coefficients across age at which the disruption occurred and disruption type.

Table 5. Multinomial logit estimates on labour force status at age 33. Various specifications of three-choice model for males (Standard errors in parentheses, marginal effects in brackets)

	1			2			3		
	b	se(b)	m.e.	b	se(b)	m.e.	b	se(b)	m.e.
LFP = Unemployed									
1									
DV0	0.59*	(0.33)	[0.03]						
DV06	0.96***	(0.31)	[0.06]	0.96***	(0.31)	[0.06]	0.86***	(0.32)	[0.04]
DV710	0.80*	(0.41)	[0.05]	0.80*	(0.41)	[0.05]	0.59	(0.42)	[0.03]
DV1115	0.42	(0.31)	[0.02]	0.42	(0.31)	[0.02]	0.33	(0.32)	[0.02]
DV1622	−0.07	(0.40)	−[0.01]	−0.07	(0.40)	−[0.01]	−0.14	(0.40)	−[0.01]
DV2333	0.78**	(0.35)	[0.05]	0.78**	(0.35)	[0.05]	0.64*	(0.35)	[0.03]
DTH015	0.46	(0.28)	[0.03]	0.46	(0.28)	[0.03]	0.22	(0.29)	[0.01]
DTH1622	0.68***	(0.25)	[0.04]	0.68***	(0.25)	[0.04]	0.51**	(0.26)	[0.03]
2									
DV	0.56***	(0.16)	[0.03]	0.55***	(0.17)	[0.03]	0.45**	(0.18)	[0.02]
DTH	0.58***	(0.19)	[0.04]	0.58***	(0.19)	[0.04]	0.38*	(0.21)	[0.02]
3									
Disrupt	0.57***	(0.13)	[0.03]	0.56***	(0.14)	[0.03]	0.42***	(0.14)	[0.02]
LFP = Not in Labor Force									
1									
DV0	1.17***	(0.39)	[0.03]						
DV06	1.09**	(0.44)	[0.03]	1.09**	(0.44)	[0.03]	0.96**	(0.45)	[0.02]
DV710	1.14**	(0.53)	[0.03]	1.14**	(0.53)	[0.03]	0.94*	(0.56)	[0.02]
DV1115	0.78*	(0.41)	[0.02]	0.78*	(0.41)	[0.02]	0.81*	(0.42)	[0.02]
DV1622	0.68	(0.44)	[0.02]	0.68	(0.44)	[0.02]	0.74*	(0.45)	[0.02]
DV2333	0.76	(0.53)	[0.02]	0.76	(0.53)	[0.02]	0.81	(0.54)	[0.02]
DTH015	0.45	(0.43)	[0.01]	0.45	(0.43)	[0.01]	0.23	(0.45)	[0.00]
DTH1622	0.38	(0.43)	[0.01]	0.38	(0.43)	[0.01]	0.17	(0.45)	[0.00]
2									
DV	0.93***	(0.21)	[0.02]	0.86***	(0.23)	[0.02]	0.84***	(0.24)	[0.02]
DTH	0.41	(0.32)	[0.01]	0.41	(0.32)	[0.01]	0.20	(0.33)	[0.00]
3									
Disrupt	0.77***	(0.19)	[0.02]	0.70***	(0.20)	[0.02]	0.60***	(0.21)	[0.01]
Chi-square (1 and 2)	9.16			6.64			5.60		
Chi-square (1 and 3)	9.60			8.36			8.60		
Chi-square (2 and 3)	0.44			1.72			3.00		
n	3,916			3,804			3,804		

*significant at 0.10 level; **significant at 0.05 level; ***significant at 0.01 level.
Samples and specifications are as follows:
(*1*) All families at birth, disruption variables only.
(*2*) Intact families at birth, disruption variables only.
(*3*) Intact families at birth, disruption variables plus birth variables.
(*4*) Intact families at age 7, disruption variables only.
(*5*) Intact families at age 7, disruption variables plus birth variables.
(*6*) Intact families at age 7, disruption variables plus birth and age 7 variables.
(*7*) Intact families at age 7, disruption variables plus birth, age 7, and education variables.

Table 5 (continued)

4			5			6			7		
b	se(b)	m.e.	b	se(b)	m.e.	b	se(b)	m.e.	b	se(b)	m.e.
0.80*	(0.41)	[0.05]	0.58	(0.42)	[0.03]	0.27	(0.45)	[0.01]	0.21	(0.44)	[0.01]
0.42	(0.31)	[0.02]	0.32	(0.32)	[0.02]	0.26	(0.33)	[0.01]	0.23	(0.33)	[0.01]
−0.07	(0.40)	−[0.01]	−0.15	(0.40)	−[0.01]	−0.28	(0.41)	−[0.01]	−0.30	(0.41)	−[0.01]
0.78**	(0.35)	[0.05]	0.63*	(0.35)	[0.03]	0.53	(0.36)	[0.02]	0.58	(0.37)	[0.02]
0.26	(0.36)	[0.02]	0.02	(0.37)	[0.00]	−0.01	(0.37)	[0.00]	−0.03	(0.38)	[0.00]
0.68***	(0.25)	[0.04]	0.52**	(0.26)	[0.03]	0.43	(0.27)	[0.02]	0.44	(0.27)	[0.02]
0.44**	(0.19)	[0.03]	0.32	(0.20)	[0.02]	0.20	(0.20)	[0.01]	0.17	(0.21)	[0.01]
0.53**	(0.21)	[0.03]	0.34	(0.22)	[0.02]	0.27	(0.23)	[0.01]	0.27	(0.23)	[0.01]
0.48***	(0.15)	[0.03]	0.33**	(0.16)	[0.02]	0.23	(0.16)	[0.01]	0.22	(0.16)	[0.01]
1.14**	(0.53)	[0.03]	0.97*	(0.55)	[0.02]	0.71	(0.57)	[0.01]	0.61	(0.58)	[0.01]
0.78*	(0.41)	[0.02]	0.81*	(0.42)	[0.02]	0.76*	(0.43)	[0.01]	0.70	(0.43)	[0.01]
0.68	(0.44)	[0.02]	0.74*	(0.45)	[0.02]	0.63	(0.45)	[0.01]	0.63	(0.46)	[0.01]
0.76	(0.53)	[0.02]	0.81	(0.54)	[0.02]	0.67	(0.54)	[0.01]	0.65	(0.55)	[0.01]
0.36	(0.52)	[0.01]	0.16	(0.54)	[0.00]	0.09	(0.55)	[0.00]	0.05	(0.55)	[0.00]
0.38	(0.43)	[0.01]	0.18	(0.45)	[0.00]	0.09	(0.45)	[0.00]	0.06	(0.46)	[0.00]
0.81***	(0.25)	[0.02]	0.82***	(0.26)	[0.02]	0.70**	(0.27)	[0.01]	0.65**	(0.27)	[0.01]
0.37	(0.34)	[0.01]	0.17	(0.36)	[0.00]	0.09	(0.37)	[0.00]	0.05	(0.37)	[0.00]
0.65***	(0.22)	[0.02]	0.57**	(0.23)	[0.01]	0.46**	(0.23)	[0.01]	0.42*	(0.23)	[0.01]
5.00			4.20			3.80			4.00		
6.46			6.60			6.00			6.20		
1.46			2.40			2.20			2.20		
3,662			3,662			3,662			3,662		

For males, the coefficient of DISRUPT is never larger than 0.02 and is never statistically significant in any model, implying that there is little overall effect of disruption on wage rates. The hypothesis that the disruption coefficients are the same across age and disruption type cannot be rejected for any model. However, in Model 6 there is a statistically significant reduction in the wage rate of about 13% resulting from a divorce occurring when the child was between the ages of 7 and 10. Thus, for males, the adverse effects of a divorce seem to work primarily through a reduction the probability of

Table 6. Multinomial logit estimates on labour force status at age 33. Various specifications of four-choice model for females. (Standard errors in parentheses, marginal effects in brackets)

	1			2			3		
	b	se(b)	m.e.	b	se(b)	m.e.	b	se(b)	m.e.
LFP = Part-time									
1									
DV0	−0.14	(0.21)	−[0.04]						
DV06	0.49**	(0.22)	[0.07]	0.49**	(0.22)	[0.07]	0.45**	(0.22)	[0.06]
DV710	0.31	(0.25)	[0.06]	0.31	(0.25)	[0.06]	0.23	(0.25)	[0.06]
DV1115	0.15	(0.20)	[0.01]	0.15	(0.20)	[0.01]	0.08	(0.21)	[0.01]
DV1622	−0.25	(0.20)	−[0.02]	−0.25	(0.20)	−[0.02]	−0.27	(0.20)	−[0.02]
DV2333	−0.05	(0.25)	−[0.03]	−0.05	(0.25)	−[0.03]	−0.11	(0.26)	−[0.04]
DTH015	−0.29	(0.19)	−[0.06]	−0.29	(0.19)	−[0.06]	−0.21	(0.20)	−[0.05]
DTH1622	−0.20	(0.17)	−[0.02]	−0.20	(0.17)	−[0.02]	−0.11	(0.17)	−[0.01]
2									
DV	0.07	(0.10)	[0.01]	0.12	(0.11)	[0.02]	0.07	(0.11)	[0.01]
DTH	−0.24*	(0.13)	−[0.04]	−0.24*	(0.13)	−[0.04]	−0.15	(0.14)	−[0.03]
3									
Disrupt	−0.03	(0.08)	−[0.01]	−0.02	(0.09)	[0.00]	−0.02	(0.09)	[0.00]
LFP = Unemployed									
1									
DV0	0.41	(0.48)	[0.01]						
DV06	0.61	(0.54)	[0.01]	0.61	(0.54)	[0.01]	0.67	(0.55)	[0.00]
DV710	0.98**	(0.50)	[0.02]	0.98**	(0.50)	[0.02]	0.84	(0.51)	[0.01]
DV1115	0.52	(0.49)	[0.01]	0.52	(0.49)	[0.01]	0.49	(0.49)	[0.00]
DV1622	−0.26	(0.60)	[0.00]	−0.26	(0.60)	[0.00]	−0.30	(0.61)	[0.00]
DV2333	−0.06	(0.74)	[0.00]	−0.06	(0.74)	[0.00]	−0.06	(0.75)	[0.00]
DTH015	−0.75	(0.73)	−[0.01]	−0.75	(0.73)	−[0.01]	−0.98	(0.74)	−[0.01]
DTH1622	0.52	(0.37)	[0.01]	0.52	(0.37)	[0.01]	0.53	(0.39)	[0.01]
2									
DV	0.38	(0.25)	[0.01]	0.38	(0.27)	[0.01]	0.35	(0.28)	−[0.01]
DTH	0.14	(0.34)	[0.01]	0.14	(0.34)	[0.01]	0.04	(0.35)	[0.00]
3									
Disrupt	0.30	(0.21)	[0.01]	0.28	(0.23)	[0.01]	0.23	(0.23)	[0.00]
LFP = NILF									
1									
DV0	0.11	(0.20)	[0.03]						
DV06	0.35	(0.23)	[0.02]	0.35	(0.23)	[0.02]	0.33	(0.23)	[0.02]
DV710	−0.03	(0.27)	−[0.04]	−0.03	(0.27)	−[0.04]	−0.10	(0.27)	−[0.05]
DV1115	0.15	(0.20)	[0.01]	0.15	(0.20)	[0.01]	0.09	(0.21)	[0.01]
DV1622	−0.36*	(0.21)	−[0.05]	−0.36*	(0.21)	−[0.05]	−0.39*	(0.21)	−[0.06]
DV2333	0.17	(0.24)	[0.04]	0.17	(0.24)	[0.04]	0.13	(0.24)	[0.04]
DTH015	0.04	(0.18)	[0.04]	0.04	(0.18)	[0.04]	0.08	(0.18)	[0.04]
DTH1622	−0.24	(0.17)	−[0.04]	−0.24	(0.17)	−[0.04]	−0.18	(0.18)	−[0.03]
2									
DV	0.06	(0.10)	[0.00]	0.04	(0.11)	−[0.01]	0.00	(0.11)	−[0.01]
DTH	−0.11	(0.13)	[0.00]	−0.11	(0.13)	[0.00]	−0.06	(0.13)	[0.00]
3									
Disrupt	0.00	(0.08)	[0.00]	−0.02	(0.09)	[0.00]	−0.02	(0.09)	[0.00]
Chi-square (1 & 2)	22.02***			19.72***			20.00***		

Table 6 (continued)

4			5			6			7		
b	se(b)	m.e.	b	se(b)	m.e.	b	se(b)	m.e.	b	se(b)	m.e.
0.31	(0.25)	[0.06]	0.23	(0.25)	[0.06]	0.17	(0.25)	[0.06]	0.08	(0.25)	[0.05]
0.15	(0.20)	[0.01]	0.08	(0.21)	[0.01]	0.02	(0.21)	[0.00]	−0.06	(0.21)	−[0.02]
−0.25	(0.20)	−[0.02]	−0.27	(0.20)	−[0.02]	−0.30	(0.20)	−[0.02]	−0.35*	(0.21)	−[0.03]
−0.05	(0.25)	−[0.03]	−0.11	(0.26)	−[0.04]	−0.18	(0.26)	−[0.04]	−0.15	(0.26)	−[0.04]
−0.34	(0.24)	−[0.08]	−0.22	(0.24)	−[0.06]	−0.29	(0.24)	−[0.07]	−0.33	(0.25)	−[0.08]
−0.20	(0.17)	−[0.02]	−0.10	(0.17)	−[0.01]	−0.10	(0.18)	−[0.01]	−0.13	(0.18)	−[0.01]
0.02	(0.12)	[0.01]	−0.04	(0.12)	[0.00]	−0.09	(0.12)	−[0.01]	−0.14	(0.12)	−[0.02]
−0.24*	(0.14)	−[0.04]	−0.14	(0.15)	−[0.03]	−0.16	(0.15)	−[0.03]	−0.19	(0.15)	−[0.05]
−0.09	(0.09)	−[0.01]	−0.08	(0.10)	−[0.01]	−0.12	(0.10)	−[0.02]	−0.16	(0.10)	−[0.03]
0.98**	(0.50)	[0.02]	0.87*	(0.51)	[0.01]	0.75	(0.53)	[0.01]	0.62	(0.53)	[0.01]
0.52	(0.49)	[0.01]	0.50	(0.49)	[0.00]	0.50	(0.50)	[0.01]	0.42	(0.50)	[0.01]
−0.26	(0.60)	[0.00]	−0.26	(0.61)	[0.00]	−0.33	(0.62)	[0.00]	−0.32	(0.62)	[0.00]
−0.06	(0.74)	[0.00]	−0.05	(0.75)	[0.00]	−0.16	(0.75)	[0.00]	−0.15	(0.75)	[0.00]
−0.35	(0.73)	−[0.01]	−0.56	(0.75)	[0.00]	−0.70	(0.75)	−[0.01]	−0.75	(0.75)	−[0.01]
0.52	(0.37)	[0.01]	0.51	(0.39)	[0.01]	0.46	(0.39)	[0.01]	0.41	(0.39)	[0.01]
0.33	(0.30)	[0.01]	0.30	(0.31)	[0.00]	0.22	(0.31)	[0.00]	0.18	(0.31)	[0.01]
0.30	(0.34)	[0.01]	0.21	(0.35)	[0.00]	0.13	(0.36)	[0.00]	0.09	(0.36)	[0.00]
0.31	(0.24)	[0.01]	0.26	(0.24)	[0.00]	0.18	(0.25)	[0.00]	0.14	(0.25)	[0.00]
−0.03	(0.27)	−[0.04]	−0.11	(0.27)	−[0.05]	−0.26	(0.28)	−[0.08]	−0.35	(0.28)	−[0.09]
0.15	(0.20)	[0.01]	0.09	(0.21)	[0.01]	0.01	(0.21)	[0.00]	−0.06	(0.21)	−[0.01]
−0.36*	(0.21)	−[0.05]	−0.39*	(0.21)	−[0.06]	−0.43**	(0.21)	−[0.06]	−0.47**	(0.21)	−[0.07]
0.17	(0.24)	[0.04]	0.13	(0.24)	[0.04]	0.03	(0.25)	[0.02]	0.05	(0.25)	[0.02]
0.09	(0.21)	[0.05]	0.16	(0.22)	[0.06]	0.12	(0.22)	[0.05]	0.07	(0.22)	[0.05]
−0.24	(0.17)	−[0.04]	−0.17	(0.18)	−[0.03]	−0.20	(0.18)	−[0.04]	−0.22	(0.18)	−[0.04]
−0.03	(0.12)	−[0.01]	−0.09	(0.12)	−[0.02]	−0.17	(0.12)	−[0.03]	−0.22*	(0.12)	−[0.06]
−0.11	(0.14)	[0.00]	−0.04	(0.14)	[0.01]	−0.07	(0.14)	[0.00]	−0.10	(0.15)	−[0.01]
−0.07	(0.09)	−[0.01]	−0.07	(0.10)	−[0.01]	−0.13	(0.10)	−[0.02]	−0.17*	(0.10)	−[0.04]
14.04**			13.60**			14.00**			13.00**		

Table 6 (continued)

Chi-square (1 & 3)	26.27***	24.89***	22.20***
Chi-square (2 & 3)	4.25	5.17	2.20
n	4,522	4,364	4,364

*significant at 0.10 level; **significant at 0.05 level; ***significant at 0.01 level.
Samples and specifications are as follows:
(*1*) All families at birth, disruption variables only.
(*2*) Intact families at birth, disruption variables only.
(*3*) Intact families at birth, disruption variables plus birth variables.
(*4*) Intact families at age 7, disruption variables only.
(*5*) Intact families at age 7, disruption variables plus birth variables.
(*6*) Intact families at age 7, disruption variables plus birth and age 7 variables.
(*7*) Intact families at age 7, disruption variables plus birth, age 7, and education variables.

employment at age 33 and only very slightly through a lower wage rate at age 33. There is some evidence that selection bias is important, even after we control for birth and age 7 characteristics.

For females, in contrast, there is considerable evidence that wage rates are adversely affected by a parental disruption. In Model 6, females who work full-time and experience a parental divorce have 12% lower wages at age 33 and females who experience the death of a father have 2% lower wages, although the latter effect is not statistically significant. As in the case of males, we cannot reject the hypothesis that the disruption coefficients are the same across age and disruption type for any of the models.

The effects of disruption on wages of women do not diminish when we control for pre-disruption characteristics. As in the case of males, selection bias (as measured by the coefficient on LAMBDA) remains important after including the pre-disruption variables. Model 6 also suggests that women who work part-time and experience a parental divorce between the ages of 11 and 22 receive lower wages. However, Model 6 also indicates that women who work part-time and experience a parental divorce between ages 7 and 10 receive higher wages. We have no explanation for this unexpected finding.

Thus, we find just the opposite pattern for females as for males. Whereas a disruption appears to manifest itself for males primarily through an adverse effect on employment at age 33, it appears to manifest itself for females primarily through a lower wage rate at age 33.

It is not clear why these differential effects of disruption on labour market status and wage rates have occurred for males and females. With regard to the absence of wage effects for males, there is some evidence from previous studies (Hetherington et al. 1982 and Wallerstein and Corbin 1986) that losses of income for education and time input losses from parental divorce are greater for girls than for boys. Thus, effects on human capital – and hence wage rates – could be greater for girls than for boys. (Amato 1991 finds stronger effects of disruption on education for girls, but we do not. However, the effects could be manifesting themselves in informally obtained human

Table 6 (continued)

14.40**	15.00**	14.00**	
16.50**			
2.46	2.20	1.00	1.00
4,166	4,166	4,166	4,166

capital, which is not captured by our education variables.) Regarding the absence of effects on labour force status for females, it is possible that boys suffer more emotional effects from not having a male role model in the house than do girls. We find that males who experienced a parental divorce had a one percentage point greater probability of not being employed at age 33 than males who did not experience a parental divorce. It is possible that long-term emotional problems resulting from parental divorce could be explaining this very small effect.

Another possibility for the differing effects by gender (more specifically, the lack of an effect of parental disruption on employment at age 33 for females) may be related to the timing of childbearing. Francesconi and Ermisch (1998) and Kiernan (1997) find that women who experience a parental divorce are more likely to start childbearing early. Consequently, such women may be more likely to be employed at age 33, all else constant. But, as we have seen, these women also have lower educational attainment and perhaps less human capital than women who do not experience a parental divorce. This lower human capital would tend to reduce their probability of employment at age 33. The lack of any significant effects of a parental disruption on employment for women at age 33 may be reflecting these two offsetting forces, which are not adequately captured by the other explanatory variables in our empirical model.

7. Conclusions

Using a longitudinal database of British individuals born during the first week in March 1958, we have investigated the effects of a parental disruption (divorce or death of a parent) on educational attainment and several labour market outcomes at age 33 (in 1991). Unlike many previous studies, we have been able to control for pre-disruption characteristics of the individuals, including a rich set of variables measured at birth and a set of socioeconomic characteristics measured at age 7. We have also examined the behavior of individuals from disrupted homes at a later stage of their life cycles than most previous studies.

We have estimated a variety of models to determine the effects of controlling for pre-disruption characteristics and the effects of limiting our sample to families that were intact at age 7. Our results indicate that a parental disruption has a strong effect on educational attainment. In particular, parental divorce between the ages of 11 and 15 seems to reduce significantly

Table 7. Selectivity corrected log of wage equations at age 33. Various specifications for males and females (Standard errors in parentheses)

	3		5		6		7	
	b	se(b)	b	se(b)	b	se(b)	b	se(b)
Male all workers								
1								
DV06	0.11**	(0.05)						
DV710	−0.06	(0.06)	−0.06	(0.06)	−0.12*	(0.06)	−0.09	(0.06)
DV1115	0.09**	(0.04)	0.09**	(0.04)	0.03	(0.04)	0.05	(0.04)
DV1622	−0.03	(0.04)	−0.03	(0.04)	−0.03	(0.04)	−0.03	(0.04)
DV2333	0.07	(0.05)	0.08	(0.05)	0.03	(0.05)	0.02	(0.05)
DTH015	−0.03	(0.04)	−0.06	(0.04)	−0.05	(0.04)	−0.05	(0.04)
DTH1622	0.00	(0.04)	0.01	(0.04)	−0.02	(0.04)	−0.03	(0.04)
LAMBDA	−0.87***	(0.12)	−0.92***	(0.13)	−0.23	(0.16)	−0.19	(0.16)
2								
DV	0.04*	(0.02)	0.03	(0.03)	−0.01	(0.03)	0.00	(0.02)
DTH	−0.01	(0.03)	−0.02	(0.03)	−0.03	(0.03)	−0.04	(0.03)
LAMBDA	−0.90***	(0.12)	−0.95***	(0.13)	−0.28*	(0.16)	−0.16	(0.16)
3								
Disrupt	0.02	(0.02)	0.01	(0.02)	−0.01	(0.02)	−0.01	(0.02)
LAMBDA	−0.92***	(0.12)	−0.96***	(0.13)	−0.29*	(0.16)	−0.17	(0.16)
Chi-square (1 & 2)	4.20		4.80		3.60		3.40	
Chi-square (1 & 3)	4.20		5.00		3.60		4.00	
Chi-square (2 & 3)	0.00		0.20		0.00		0.60	
n	3,100		3,003		3,003		3,003	
Female full-time workers								
1								
DV06	0.02	(0.07)						
DV710	−0.12	(0.08)	−0.12	(0.08)	−0.13*	(0.07)	−0.12*	(0.07)
DV1115	−0.07	(0.06)	−0.07	(0.06)	−0.09*	(0.06)	−0.05	(0.05)
DV1622	−0.19***	(0.06)	−0.19***	(0.06)	−0.13**	(0.05)	−0.10**	(0.05)
DV2333	−0.17**	(0.07)	−0.17**	(0.07)	−0.13*	(0.07)	−0.13**	(0.06)
DTH015	−0.06	(0.05)	−0.05	(0.06)	−0.02	(0.06)	−0.04	(0.05)
DTH1622	−0.03	(0.05)	−0.03	(0.05)	−0.03	(0.04)	−0.03	(0.04)
LAMBDA	−0.65***	(0.16)	−0.63***	(0.16)	−0.36***	(0.14)	−0.20	(0.13)
2								
DV	−0.11***	(0.03)	−0.14***	(0.03)	−0.12***	(0.03)	−0.10***	(0.03)
DTH	−0.04	(0.04)	−0.04	(0.04)	−0.02	(0.04)	−0.04	(0.03)
LAMBDA	−0.64***	(0.15)	−0.62***	(0.15)	−0.35***	(0.14)	−0.23*	(0.12)
3								
Disrupt	−0.08***	(0.03)	−0.10***	(0.03)	−0.08***	(0.03)	−0.07***	(0.03)
LAMBDA	−0.68***	(0.15)	−0.67***	(0.15)	−0.39***	(0.14)	−0.22*	(0.12)
Chi-square (1 & 2)	2.00		2.00		0.80		1.00	
Chi-square (1 & 3)	6.40		6.40		4.60		3.00	
Chi-square (2 & 3)	4.40		4.40		3.80		2.00	
n	1,406		1,350		1,350		1,350	
Female part-time workers								
1								
DV06	−0.11*	(0.06)						
DV710	0.17**	(0.07)	0.17**	(0.07)	0.18**	(0.07)	0.22***	(0.06)
DV1115	−0.15**	(0.06)	−0.15**	(0.06)	−0.13**	(0.06)	−0.07	(0.05)
DV1622	−0.14**	(0.07)	−0.14**	(0.07)	−0.14**	(0.06)	−0.05	(0.06)

Table 7 (continued)

	3		5		6		7	
	b	se(b)	b	se(b)	b	se(b)	b	se(b)
DV2333	−0.12	(0.08)	−0.12	(0.08)	−0.11	(0.08)	0.03	(0.07)
DTH015	−0.10	(0.07)	−0.11	(0.08)	−0.10	(0.08)	0.01	(0.08)
DTH1622	0.01	(0.06)	0.01	(0.06)	0.04	(0.05)	0.06	(0.05)
LAMBDA	0.14	(0.15)	0.11	(0.16)	0.13	(0.16)	−0.08	(0.15)
2								
DV	−0.08**	(0.03)	−0.08**	(0.04)	−0.06*	(0.04)	0.02	(0.03)
DTH	−0.03	(0.04)	−0.03	(0.05)	0.00	(0.05)	0.05	(0.04)
LAMBDA	0.12	(0.15)	0.09	(0.16)	0.11	(0.16)	−0.11	(0.15)
3								
Disrupt	−0.07**	(0.03)	−0.06*	(0.03)	−0.04	(0.03)	0.03	(0.03)
LAMBDA	0.14	(0.15)	0.11	(0.15)	0.13	(0.15)	−0.10	(0.14)
Chi-square (1 & 2)	10.60		10.00		10.40		10.80	
Chi-square (1 & 3)	11.40*		10.60		11.40*		11.00	
Chi-square (2 & 3)	0.80		0.60		1.00		0.20	
n	1,154		1,098		1,098		1,098	

*significant at 0.10 level; **significant at 0.05 level; ***significant at 0.01 level.
Samples and specifications are as follows:
(3) Intact families at birth, disruption variables plus birth variables.
(5) Intact families at age 7, disruption variables plus birth variables.
(6) Intact families at age 7, disruption variables plus birth and age 7 variables.
(7) Intact families at age 7, disruption variables plus birth, age 7, and education variables.

the probability of attaining higher levels of education for both males and females, although the effect diminishes considerably once the effects of individual characteristics at birth and age 7 are taken into account. For males, we find that after controlling for pre-disruption family circumstances, parental divorce has a greater adverse effect on educational attainment than the death of a father. For females, however, both parental divorce and death of a father adversely affects educational attainment, even after controlling for pre-disruption family circumstances.

Our results also indicate that a parental disruption adversely affects labour market outcomes of males and females. However, the way in which these effects occur differs for males and females. For males, the effect occurs primarily through decreased employment, although the effect is diminished by the addition of pre-disruption family circumstances. For females, the effect occurs primarily through decreased wage rates and does not diminish when pre-disruption family circumstances are taken into account. Including educational attainment as a control variable does not materially affect the disruption effects for either males or females, despite the fact that education itself is adversely affected by family disruptions.

Because divorce and parental death are relatively rare events, we are not able to identify precisely separate effects of disruption by age and disruption type. However, our results seem to imply that disruptions occurring prior to the middle teenage years have somewhat greater adverse effects on educational attainment, while disruptions occurring into young adulthood have

Table A.1. Full set of results for males. Unconstrained model, specification 6 (Standard errors in parentheses)

	Education (Ordered probit)		Unemployed (Multinomial logit)		Not in labor force (Multinomial logit)		Log of hourly wages (Selectivity corrected)	
Constant	−2.098***	(0.466)	−0.19	(0.851)	−1.05	(1.225)	1.43***	(0.110)
DV710	−0.318**	(0.145)	0.27	(0.447)	0.71	(0.574)	−0.12*	(0.064)
DV1115	−0.233**	(0.092)	0.26	(0.327)	0.76*	(0.427)	0.03	(0.041)
DV1622	0.035	(0.104)	−0.28	(0.410)	0.63	(0.453)	−0.03	(0.041)
DV2333	−0.098	(0.110)	0.53	(0.361)	0.67	(0.544)	0.03	(0.054)
DTH015	−0.073	(0.096)	−0.01	(0.374)	0.09	(0.546)	−0.05	(0.043)
DTH1622	−0.060	(0.084)	0.43	(0.269)	0.09	(0.455)	−0.02	(0.037)
MOMAGE	0.005	(0.004)	−0.01	(0.015)	0.02	(0.021)		
AGEDIFF	−0.001	(0.005)	−0.02	(0.017)	0.01	(0.026)		
MISSDIFF	0.098	(0.257)						
BRTHWGT	0.002**	(0.001)	0.00	(0.003)	0.00	(0.005)	0.00**	(0.000)
MISSBRWT	0.573*	(0.311)						
INTERVAL1	0.080*	(0.047)	−0.16	(0.179)	0.01	(0.264)		
INTERVAL2	0.144**	(0.062)	−0.16	(0.245)	0.05	(0.333)		
INTERVAL3	0.254***	(0.073)	−0.09	(0.308)	−0.23	(0.465)		
INTERVAL4	0.159**	(0.066)	−0.30	(0.289)	−0.61	(0.426)		
INTERVAL0	−0.035	(0.077)	−0.21	(0.272)	−0.51	(0.464)		
INTMISS	0.003	(0.156)	−0.59	(0.558)	−0.33	(0.773)		
MGCLASS2	−0.093	(0.098)	−0.26	(0.363)	−0.25	(0.504)	−0.02	(0.039)
MGCLASS3	−0.263**	(0.103)	−0.19	(0.378)	−0.33	(0.531)	−0.07*	(0.041)
MGDEAD	−0.182	(0.114)	0.01	(0.411)	0.04	(0.568)	−0.03	(0.046)
MGMINER	−0.336***	(0.117)	0.21	(0.416)	−0.93	(0.707)	−0.13***	(0.047)
MGMISS	−0.387***	(0.137)	−0.85	(0.556)	−1.20	(0.879)	−0.09	(0.056)
PERSPROOM	−0.042**	(0.021)	0.11	(0.069)	0.17*	(0.100)		
ROOMMISS	−0.239*	(0.135)	−0.14	(0.546)	0.23	(0.764)		
FSOC-CLASS1	0.322***	(0.087)	−0.09	(0.298)	−0.37	(0.518)	0.00	(0.036)
FSOC-CLASS2	0.183***	(0.070)	−0.36*	(0.205)	0.18	(0.359)	0.00	(0.030)
FSOC-MISS	0.189	(0.144)	−0.31	(0.521)	0.37	(0.708)	0.05	(0.060)
BEHPROB7	−1.875***	(0.287)	2.12**	(0.969)	2.26	(1.440)	−0.34***	(0.128)
BPMISS	0.035	(0.543)						

READTEST7	0.021***	(0.004)	0.01	(0.013)	−0.01	(0.019)	0.00	(0.002)
READMISS	0.505**	(0.252)	0.41	(1.077)	−0.45	(1.980)	0.16	(0.101)
MATHTEST7	0.042***	(0.010)	−0.03	(0.038)	−0.08	(0.058)	0.02***	(0.004)
MATHMISS	0.211	(0.229)	−2.61*	(1.404)	−3.28	(2.222)	0.02	(0.105)
SCHPER7	0.067***	(0.008)	−0.10***	(0.033)	−0.07	(0.049)	0.02***	(0.004)
MISSSCH1	1.139***	(0.306)	0.61	(1.146)	1.61	(1.379)	0.36***	(0.124)
MOMEDUC	0.233***	(0.048)	−0.03	(0.199)	0.28	(0.277)	0.06***	(0.019)
MOMMISS	0.206	(0.480)	1.08	(0.755)	0.65	(1.105)	−0.08	(0.089)
DADEDUC	0.215	(0.271)	−0.21	(0.222)	0.15	(0.310)	0.06***	(0.020)
DADMISS	−0.074	(0.113)	−0.29	(0.478)	0.36	(0.527)	−0.03	(0.046)
AGELEFT	0.064***	(0.015)						
MISSLEFT	1.002***	(0.380)						
YOUNG2	−0.115	(0.075)	0.47	(0.347)	−0.39	(0.402)	0.01	(0.030)
OLDEST2	−0.016	(0.075)	−0.12	(0.360)	−0.47	(0.427)	−0.02	(0.030)
YOUNG3	−0.205**	(0.081)	0.56	(0.358)	−0.80*	(0.438)	−0.03	(0.032)
MIDDLE3	−0.232***	(0.072)	0.71**	(0.329)	−0.55	(0.385)	−0.06**	(0.030)
OLDEST3	−0.050	(0.085)	0.30	(0.361)	−0.83*	(0.502)	0.02	(0.033)
PREGWORK1			−0.52	(0.399)	−0.42	(0.612)		
PREGWORK2			−0.15	(0.342)	−0.67	(0.564)		
PREGWORK3			−1.03	(1.082)	0.44	(0.895)		
PREGNOWORK			−0.50	(0.317)	−0.18	(0.493)		
Mu(1)	0.611***	(0.029)						
Mu(2)	1.652***	(0.038)						
Mu(3)	2.270***	(0.042)						
Mu(4)	2.950***	(0.049)						
LAMBDA							−0.19	(0.160)

*significant at 0.10 level; **significant at 0.05 level; ***significant at 0.01 level.

Note: Variables are excluded from the labour force status equations because of convergence problems due to multicollinearity.

Table A.2. Full set of results for females. Unconstrained model, specification 6 (Standard errors in parentheses)

	Education (Ordered probit)		Employed part-time (Multinomial logit)		Unemployed (Multinomial logit)		Not in labor force (Multinomial logit)		Log of hourly wages for full-time (Selectivity corrected)		Log of hourly wages for part-time (Selectivity corrected)	
Constant	−1.985***	(0.424)	3.53***	(0.973)	−3.55	(2.471)	4.07***	(1.005)	2.29***	(0.365)	0.96***	(0.287)
DV710	−0.236**	(0.098)	0.17	(0.253)	0.75	(0.527)	−0.26	(0.276)	−0.13*	(0.070)	0.18**	(0.092)
DV1115	−0.311***	(0.090)	0.02	(0.210)	0.50	(0.502)	0.01	(0.210)	−0.09*	(0.056)	−0.13*	(0.072)
DV1622	−0.051	(0.093)	−0.30	(0.205)	−0.33	(0.620)	−0.43**	(0.212)	−0.13**	(0.055)	−0.14*	(0.078)
DV2333	0.037	(0.113)	−0.18	(0.259)	−0.16	(0.750)	0.03	(0.247)	−0.13*	(0.066)	−0.11	(0.096)
DTH015	−0.195**	(0.095)	−0.29	(0.245)	−0.70	(0.754)	0.12	(0.220)	−0.02	(0.058)	−0.10	(0.101)
DTH1622	−0.084	(0.072)	−0.10	(0.176)	0.46	(0.392)	−0.20	(0.179)	−0.03	(0.045)	0.04	(0.066)
MOMAGE	0.013***	(0.004)	−0.02**	(0.009)	0.03	(0.022)	−0.02**	(0.009)				
AGEDIFF	−0.011**	(0.005)	0.01	(0.011)	0.02	(0.030)	0.01	(0.011)				
MISSDIFF	−0.378	(0.254)										
BRTHWGT	0.001*	(0.001)	0.00	(0.002)	−0.01	(0.005)	0.00	(0.002)	0.00	(0.000)	0.00	(0.001)
MISSBRWT	0.485*	(0.291)										
INTERVAL1	0.111**	(0.044)	0.05	(0.101)	−0.05	(0.265)	0.04	(0.103)				
INTERVAL2	0.149**	(0.059)	−0.28**	(0.141)	−0.41	(0.412)	0.02	(0.136)				
INTERVAL3	0.178**	(0.076)	0.15	(0.169)	0.13	(0.475)	0.33**	(0.168)				
INTERVAL4	0.086	(0.062)	−0.08	(0.142)	−0.84*	(0.485)	0.15	(0.140)				
INTERVAL0	−0.199**	(0.078)	−0.27	(0.180)	−0.54	(0.474)	−0.15	(0.177)				
INTMISS	0.100	(0.125)	−0.14	(0.298)	0.13	(0.671)	0.23	(0.279)				
MGCLASS2	0.011	(0.089)	0.48**	(0.226)	0.79	(0.750)	0.33	(0.208)	0.11**	(0.053)	−0.05	(0.112)
MGCLASS3	−0.054	(0.095)	0.70***	(0.238)	0.90	(0.776)	0.44**	(0.222)	0.15**	(0.061)	−0.09	(0.119)
MGDEAD	−0.021	(0.105)	0.81***	(0.259)	0.46	(0.872)	0.36	(0.248)	0.13**	(0.067)	−0.09	(0.131)
MGMINER	−0.145	(0.107)	0.49*	(0.273)	1.59**	(0.804)	0.43*	(0.255)	0.12*	(0.069)	−0.21*	(0.124)
MGMISS	−0.064	(0.119)	0.86***	(0.295)	0.74	(0.953)	0.53*	(0.285)	0.22***	(0.081)	−0.01	(0.137)
PERSPROOM	−0.006	(0.019)	−0.11**	(0.047)	0.06	(0.109)	−0.06	(0.046)				
ROOMMISS	−0.243**	(0.120)	−0.11	(0.287)	0.73	(0.659)	−0.25	(0.301)				
FSOC-CLASS1	0.267***	(0.081)	0.02	(0.192)	−0.76	(0.478)	−0.30	(0.187)	−0.04	(0.051)	0.10	(0.069)
FSOC-CLASS2	0.168***	(0.063)	−0.03	(0.152)	−0.76**	(0.315)	−0.29**	(0.146)	0.01	(0.042)	0.05	(0.054)
FSOC-MISS	0.165	(0.144)										

BEHPROB7	−2.658***	(0.332)	0.65	(0.766)	2.49	(1.873)	1.10	(0.759)	−0.76***	(0.201)	−0.59**	(0.274)
BPMISS	−0.082	(0.593)										
READTEST7	0.022***	(0.004)	−0.01	(0.009)	0.00	(0.024)	−0.01	(0.009)	0.00	(0.003)	0.00	(0.003)
READMISS	0.407*	(0.216)	−0.72	(0.534)	−0.45	(1.437)	0.02	(0.494)	0.12	(0.133)	0.01	(0.195)
MATHTEST7	0.034***	(0.009)	0.01	(0.020)	0.03	(0.058)	−0.02	(0.020)	0.00	(0.005)	0.01	(0.007)
MATHMISS	0.100	(0.195)	0.20	(0.509)	−0.54	(1.601)	−0.03	(0.500)	−0.06	(0.142)	−0.18	(0.157)
SCHPER7	0.060***	(0.008)	−0.03*	(0.019)	−0.07	(0.054)	−0.04**	(0.019)	0.01*	(0.005)	0.01*	(0.007)
MISSSCH1	1.217***	(0.224)	−0.19	(0.588)	0.23	(1.598)	−0.61	(0.572)	0.31*	(0.171)	0.45**	(0.202)
MOMEDUC	0.308***	(0.044)	−0.10	(0.102)	0.20	(0.298)	0.01	(0.101)	0.07***	(0.026)	0.10***	(0.040)
MOMMISS	0.553	(0.552)	0.44	(0.531)	1.18	(1.282)	0.88*	(0.490)	0.22	(0.145)	0.04	(0.178)
DADEDUC	−0.221	(0.284)	−0.51	(0.586)	0.60	(1.096)	−0.99	(0.690)	0.44***	(0.146)	0.04	(0.046)
DADMISS	0.073	(0.095)	−0.54**	(0.233)	−1.13	(0.888)	−0.46**	(0.234)	−0.00	(0.063)	0.08	(0.097)
AGELEFT	0.047***	(0.013)	−0.09***	(0.033)	0.04	(0.078)	−0.06*	(0.029)				
MISSLEFT	0.279	(0.367)	−1.95**	(0.814)	1.12	(1.775)	−2.00**	(0.858)				
YOUNG2	−0.130*	(0.072)	0.02	(0.169)	−0.70	(0.460)	0.10	(0.167)	0.01	(0.040)	−0.04	(0.062)
OLDEST2	0.018	(0.071)	0.08	(0.162)	−0.40	(0.484)	0.12	(0.163)	0.00	(0.040)	0.00	(0.060)
YOUNG3	−0.321***	(0.079)	0.42**	(0.183)	−0.09	(0.448)	0.35*	(0.183)	−0.05	(0.046)	−0.04	(0.063)
MIDDLE3	−0.205***	(0.069)	0.19	(0.165)	−0.46	(0.422)	0.17	(0.165)	−0.05	(0.039)	0.00	(0.057)
OLDEST3	0.008	(0.080)	0.28	(0.183)	−0.16	(0.531)	0.23	(0.187)	−0.05	(0.049)	0.02	(0.066)
PREGWORK1			−0.31	(0.233)	0.50	(1.174)	−0.29	(0.238)				
PREGWORK2			−0.38*	(0.210)	0.92	(1.067)	−0.54**	(0.218)				
PREGWORK3			−0.68	(0.439)	1.22	(1.461)	−0.44	(0.424)				
PREGNOWORK			−0.36*	(0.200)	1.57	(1.029)	−0.18	(0.205)				
Mu(1)	0.677***	(0.026)										
Mu(2)	1.976***	(0.035)										
Mu(3)	2.339***	(0.036)										
Mu(4)	3.075***	(0.045)										
Lambda									−0.36***	(0.138)	0.13	(0.190)

*significant at 0.10 level; **significant at 0.05 level; ***significant at 0.01 level.
Note: Variables are excluded from the labour force status equations because of convergence problems due to multicollinearity.

adverse effects on certain labour market outcomes. In all our specifications, if we do not control for pre-disruption characteristics, we estimate much larger effects on educational attainment for both males and females and on labour market outcomes for males. This suggests that the pre-disruption characteristics often control for unobserved, shared incapacities or difficulties of the parents and children (e.g., family dysfunctioning). For example, if there is a shared family history of depression (which, in part, might be genetically derived), this could have affected both the child's age 7 behavioral problems and the parents' likelihood of divorcing.

Although we find strong effects of a parental disruption on educational attainment, we also find independent effects of disruption on labour market outcomes. However, to fully investigate the avenues through which family disruption affects labour market outcomes requires a more comprehensive structural model of educational attainment and labour market behavior, something that has not been attempted here. Future research on the economic consequences of parental disruption using a structural framework might be better able to unravel the routes through which family disruption affect economic well being in the adult years.

Appendix: Construction of selected explanatory variables

FSOC-CLASS2 = 1 if father's occupational class is professional (i.e. doctor, lawyer) or intermediate (i.e. manager, teacher)

FSOC-CLASS3 = 1 if father's occupational class is skilled non-manual (i.e. clerk, shop assistant), skilled manual (i.e. miner, bricklayer), or partly skilled (i.e. mail carrier, bus conductor) (base group for FSCL is unskilled (i.e. cleaner, labourer), unemployed, student, sick, or retired)

BEHPROB7 factor score for behavioral problems at age 7 (from Cherlin, Kiernan, and Chase-Lansdale 1995); includes logarithm of parent-rated behavior problems, logarithm of teacher-rated behavior problems, use of social services, and family reported difficulties.

READTEST7 Southgate Group Reading Test at age 7-standardized test of how well children have learned to read in school (involving both word recognition and sentence completion). As such, it probably confounds children's cognitive ability with their mastery of reading.

MATHTEST7 Mathematics Achievement Test at age 7-score represents the number of correctly answered questions in a 10 question problem arithmetic test. Developed specifically for NCDS.

SCHPER7 School performance scale at age 7-summed scale of the teacher's responses to 5 school performance items, each of which was rated 1 to 5. Includes oral ability, the child's awareness of the world around him or her, reading, creativity, and number work.

NVQ1	= 1 if some qualifications (RSA Stage 1, CSE Grades 2 to 5, other technical and business qualifications)
NVQ2	= 1 if O levels or equivalent (Joint Industry Board Technicians Certificate CGLI Operative, Insignia Award and other qualifications, RSA Stages 2 and 3, Scottish Standard Grade, O Grade, O Level, CSE Grade 1, GCSE)
NVQ3	= 1 if A Level Equivalent (TEC, BEC, BTEC National Certificate or Diploma, Joint Industry Board ONC/OND, CGLI Advanced Part 2, Scottish Certificate of 6th Form Studies, GCE and Scottish Highers, Scott Beck Awards, GCE A level)
NVQ4	= 1 if Other Higher Education (Training College Certificate, non-validated diploma or certificate, nursing qualification, professional qualification, TEC/BET/BETEC and Scottish equivalent, Higher National Diploma, Joint Industry Board, HNC/HND, CGLI full technical qualification)
NVQ5	= 1 if Higher Education (First Degree, Postgraduate Diploma, Masters, Ph.D.)

Endnotes

[1] For example, the divorce rate for England and Wales in 1961 was 2.1 per 1,000 married persons, by 1971 it was 6.0, and in 1980 it was 12.0 and has remained more or less steady since then (Kiernan 1988). Many of these divorcing parents – 55% in 1985, for example – have children under age 16 (Kiernan 1988). At present, divorce rates in Britain are the highest in Western Europe.

[2] In addition, between 1958 and 1965, a supplementary sample was added consisting of 1,142 children from recent immigrant families, who were also born during the first week of March 1958. Hence, the total initial sample was 18,556. We do not use the supplementary sample in our analysis.

[3] While only 65 percent of the original cohort was interviewed at age 33, Shepherd (1993) argues that the age 33 sample is representative of the original cohort and is similar to other nationally representative samples of a comparable age group.

[4] Because the NCDS survey contains comprehensive measures of childhood experiences over a number of years, it is not likely to suffer from the "window problem" that occurs when children's attainments are linked causally to events or circumstances occurring at a single point in time during childhood (e.g., see Wolfe et al., 1996).

[5] Interestingly, at least two U.S. studies have obtained rather large negative estimates of the impact of parental divorce on the earnings of adult white males. Findings from Greenberg and Wolf (1982) imply a 28% earnings reduction and those from Grogger and Ronan (1995) imply a 12% earnings reduction. However, both studies also found that parental divorce had a positive impact on the adult earnings of black males, although this finding was not statistically significant in the Greenberg and Wolf study.

[6] In some studies (for example, Couch and Lillard, 1995), few or no pre-disruption control variables are available for use. Other studies (for example, Krein 1986; Boggess 1998 and Kiernan 1998) use independent variables, such as measures of family financial status, that were measured after the marriage disruption occurred and, hence, are potentially endogenous to the disruption. Sometimes the researchers seem unaware of this problem. More often, however, the researchers are aware, but want to determine whether the effects of disruption on children are more attributable to the disruption per se or to changes in family circumstances that result from the disruption. Nonetheless, because of the endogeneity problem, the findings in these studies are somewhat difficult to interpret.

[7] While not directly examining this issue, however, Haveman and Wolfe (1994) and Boggess (1998) used U.S. data to find that the length of time that children live with one parent adversely affects the educational attainment of children in most demographic groups. Boggess attributes most of this duration effect to the economic deprivation that often occurs in single parent families.

[8] In addition, Wadsworth and Maclean (1986) used longitudinal data for 5,362 children born in Great Britain in March 1946 to examine the effects of parental disruptions on children's education attainment and adult income. The data they used are similar in many respects to the NCDS. They found that educational attainment and income were both lower for children who experienced a parental disruption before age 14.

[9] This somewhat restrictive measure of disruption was apparently adopted to eliminate collinearity with another variable of interest, whether individuals lived in families that faced financial difficulties before they were 16 years old.

[10] Gross hourly earnings are derived from a survey question about usual gross pay before deductions. The respondent was asked the time period covered by the earnings response: 37% said one week, 1% said a fortnight, 4% said four weeks, 25% said one month, and 32% said one year. The responses were all converted to hourly terms using usual hours of work. Self-employed individuals, who comprised 14% of the sample, were asked to report earnings in weekly terms. Self-employed individuals were not asked about usual hours of work so they are not included in our sample.

[11] Lee (1983) describes the computation of this estimator and explains the rationale for using it (the procedure is summarized in Greene 1998). To identify the wage regressions, the explanatory variables in the wage regressions differ from those in the multinomial logit equations (see Tables A.1 and A.2).

[12] In England, an individual's formal education is represented by the qualifications they receive. Qualification levels are a reflection of course work and level of education in years. In addition, some qualification levels can be obtained only if certain tests are passed. England's National Council for Vocational Qualifications have grouped qualification levels into six National Vocational Qualification (NVQ) levels in order to have a standardized measure of education levels. These six levels are represented as follows: no qualifications (NVQ0), some qualification (NVQ1), O levels or equivalent (NVQ2), A level or equivalent (NVQ3), other higher education (NVQ4), and higher education (NVQ5). For further details, see the Appendix.

[13] However, restricting the sample to individuals living in intact households at age 7 introduces further potential sample selectivity problems. We tested for the possibility of sample selection bias from using the age 7 sample. These tests indicated that using the age 7 sample does not significantly alter our results. As discussed in the text, we also estimate several models using more inclusive samples.

[14] Because we use this approach, in estimating wage regressions, we exclude variables that are typically included in a wage equation – for example, firm size, education, and work experience. Such variables, which are measured in the NCDS at age 33, may be endogenous with respect to parental disruption. In one specification, however, we do include education as an explanatory variable.

[15] Out of the original total sample of 17,414, 9,723 were interviewed at both age 7 and age 33. Of these, 70 did not have valid data on parental disruptions, 753 did not have valid data on work status at age 33, and 677 had experienced a parental disruption by age 7. This leaves a sample of 7,828 (3,662 males and 4,166 females). A slightly smaller sample (3,603 males and 4,129 females) is used in the education regressions because of missing data on education.

[16] In our sample, 13.2% of the observations have at least one missing value for an explanatory variable. Nearly 70% of these observations (9.1% of the total sample) have only one missing value. The missing data problem is thus minimal with regard to individual observations, as most have missing data for only one variable. Moreover, for most individual variables, fewer than 3% of the observations have missing values (see Table 2). A comparison of the means of the samples with and without missing data indicates that the two samples are quite similar.

[17] Information about the death of a mother is not included in the NCDS and information about the death of a father is not included subsequent to the child reaching age 22. It is not clear how the complete lack of information about the death of a mother affects our findings. On the one hand, for this age cohort, the death of the father was more likely to be associated with a reduction in family income than the death of a mother. On the other hand, the loss of time inputs and emotional effects resulting from the death of a mother could be significant. The lack

of data about the death of a father subsequent to age 22 probably is not an important shortcoming since the child had probably already left home by that time.

[18] The variables we use to measure a parental divorce were carefully constructed to be consistent across waves of the NCDS. Information on parental divorce is available in NCDS1-3 and retrospectively in NCDS5 (no information on divorce was obtained in NCDS4). Briefly, if a divorce was indicated in either NCDS1-3 or NCDS5, the divorce was considered to have occurred. If the information was inconsistent across surveys, the information in NCDS5 was considered correct. A detailed memorandum on how the divorce variables were constructed is available from the authors.

[19] To test for possible threshold effects for birth weight, we estimated an alternative model with a series of four dummy variables for birth weight (< 5 lbs., between 5 and 6 lbs., between 6 and 7 lbs., and between 7 and 8 lbs.). We could not reject the hypothesis that birth weight effects were the same across categories, so a linear effect was specified.

[20] Although not the focus of this study, there are some interesting results for the nondisruption variables. The results indicate, for example, that education is positively related to the child's birth weight, the length of time between the parents' marriage and the birth of their first child, the social class of the maternal grandfather, the social class of the father, reading and arithmetic test scores at age 7, school performance at age 7, the mother's education, and the age the father left school, and negatively related to the degree of behavioral problems at age 7 and the number of persons per room in the household at the time of the birth.

[21] This result that most of the effect is for divorces that occur prior to age 15 differs somewhat from Kiernan (1997), who finds that parental divorce after age 21 also adversely affects the educational attainment of males. In Models 1 through 5, we find an effect for males for divorces occurring after age 23, but the effect becomes statistically insignificant in Model 6 when we control for age 7 family characteristics.

[22] Again, although not the focus of this study, there are a few interesting results for the nondisruption variables. The most important is that for both males and females, the probability of being unemployed or out of the labour force at age 33 is positively related to the extent of behavioral problems at age 7.

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Optimal age at motherhood. Theoretical and empirical considerations on postponement of maternity in Europe

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Abstract. The age at which women become mothers has increased to an all time high in most European countries in the past decennia. This increase of age at first birth is the main explanatory variable for the rapid decrease in fertility in European countries which has occurred at different points of time earlier in North and West Europe than in South Europe. To understand the development of the period fertility rate it is therefore crucial to understand the determinants of optimal age at maternity. This paper reviews empirical and theoretical literature and tries to give suggestions on future research directions. The econometric so called timing and spacing literature has used current female wages and male incomes as the main explanatory variables. However, theoretical research identifies on the one hand consumption smoothing, and on the other hand career planning of the woman as the main explanations to the postponement of maternity.

JEL classification: D1, J1

Key words: Postponement of maternity, economic theories of timing of fertility

1. Introduction

Increasingly young women in all European countries educate themselves for a lifelong labour market career. Many women find that there is no room for children, and sometimes they are given promotion only on the condition that

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they are not pregnant or do not plan on a child in the near future, which results in postponing or refraining from children altogether. For example, in 1990 are 61% of Dutch women of the age group 25–29 still childless, a number, which surpasses with a broad margin the second runner up, which is Germany with 57 percent. In Sweden, the corresponding number is 49% (Bosveld 1996). Until recently the research agenda on the lack of equal opportunities between women and men focussed on labour force participation of women with young children, whereas the consequences of working life for fertility and timing and spacing of births has received less attention. On the other hand, research on the determinants of fertility has recognized the importance of timing in explaining period total fertility rates. (Ermisch 1990; Cigno 1991; Heckman and Walker 1990; Hotz et al. 1997). The purpose of this paper is to review related literature in demography, economic theory and econometrics in order to review the extent of postponement in different European countries, to formulate hypotheses on the determinants of optimal age at maternity and to present empirical evidence if available.

The postponement of births that we have observed in Europe creates also a biological strain. Biological and medical literature reviewed by Wetzels (1999 Chapt. 7) stress for example the medical costs associated with the realization of a desire to give birth by older couples who are less fecund. One estimate cited by Wetzels is (Fauser 1998) for the Netherlands in 1995 close to 2,000 births out of 190500 live births, were the result of In Vitro Fertilization (IVF). Costs for IVF are estimated by another study Bonsel and Van der Maas (1994) for the Netherlands in 1991 at 33 million Dutch guilders (1.8 guilders to a US dollar). Many of these costs could probably have been avoided, if the couple had chosen for parenthood at a younger age.

This paper starts with a presentation of the extent of postponement of maternity in some European countries in Sect. 2. The third section reviews the main results of the econometric timing and spacing literature which uses current incomes and wages as main determinants. Section 4 explains the theoretical shadow price at giving birth. Section 5 analyses consumption smoothing and husbands earnings as determinants of mother's age of first birth. Section 6 develops the career-planning motive which is subdivided into five subsections of variables that affect the woman's life time earnings by different decisions on timing of maternity. In Sect. 7 some results from biological, medical and health research are reviewed. Section 8 concludes.

2. The ageing of fertility in Europe

Bosveld (1996) is a careful demographic analysis of the ageing of fertility in Europe. This section presents and evaluates her main findings. Total fertility rates declined sharply in Europe during the first demographic transition, which for most European countries occurred between 1880 and 1940. In some countries fertility fell below replacement level by 1940. This was the case in Belgium, Switzerland, France, England and Wales and Sweden. Demographers at that time expected fertility to stabilize at a low level but between 1940 and 1960 total fertility rates increased in many countries creating what has been termed the 'post second world war baby boom.' However, since 1965 a second decreasing fertility trend started, resulting in fertility levels far below replacement in almost all European countries. Van de Kaa (1987) termed this

Table 1. Contribution of women aged 30 and over to the period TFRs, 1960–1992

	1960	1965	1970	1975	1980	1985	1990	1992	1996
Western Europe									
Austria	32.0	35.2	28.3	24.4	23.0	24.9	28.1	28.5[1991]	35.4
Belgium	35.6	34.2	27.9	22.2	24.7	27.7	–	–	32.4[a]
France	34.0	31.7	30.9	27.4	27.2	31.1	37.7	–	42.6[b]
Germany*	35.7	33.3	31.9	28.2	30.1	34.8	38.5	37.6	46.8
The Netherlands	47.6	42.1	33.4	25.6	27.5	34.3	48.1	51.8	53.5
Switzerland	40.5	36.6	31.8	29.3	34.2	37.9	43.6	44.3[1991]	50.0
Northern Europe									
Denmark	27.3	26.3	24.6	21.7	23.5	29.5	35.4	42.2	42.3[c]
England-Wales	30.2	28.9	25.8	23.1	26.0	29.6	33.5	–	41.2
Finland	35.5	33.6	28.8	26.8	31.2	35.6	39.2	39.3	47.2
Iceland	36.8	35.4	32.5	29.6	30.9	31.9	36.6	38.9[1991]	42.1
Ireland	57.1	53.5	49.5	44.3	45.7	47.4	47.7	53.9	52.8
Norway	32.4	29.7	27.9	23.6	25.8	29.1	37.8	40.5	40.6
Sweden	31.6	32.5	26.8	24.9	29.8	35.2	41.0	43.1	44.3
Southern Europe									
Greece	37.4	34.5	31.4	27.1	22.5	23.4	27.7	–	33.7[c]
Italy	41.0	38.4	35.6	31.5	29.9	34.0	44.2	45.5[1991]	47.3[c]
Portugal	43.3	43.8	39.8	36.5	29.2	29.2	33.5	34.6	35.3
Spain	45.8	45.7	41.5	38.0	34.6	36.3	43.9	45.5	49.0[c]
Eastern Europe									
Bulgaria	17.2	16.3	14.6	13.1	10.7	11.5	11.6	11.0[1991]	12.2[c]
Czech Republic*	21.4	21.1	18.0	17.3	14.8	13.9	16.1	–	17.9
Hungary	21.4	19.6	18.9	18.2	15.0	16.2	18.0	18.8[1991]	20.8
Poland	31.4	28.6	27.4	25.6	23.3	23.1	22.0	22.3[1991]	26.7
Rumania	26.3	23.1	27.5	22.7	18.4	18.2	19.6	16.8[1991]	12.8
Yugoslavia*	–	–	26.0	24.1	22.7	21.5	–	–	25.2

Source: Bosveld (1996, Table 1.2).
* Countries with border changes. Numbers for 1996 are from the 1997 Demographic Yearbook, United Nations, New York 1999. The following are for a different year, [a] = 1992, [b] = 1994 and [c] = 1995.

decline the 'second demographic transition' pointing at individualization and emancipation of women as the driving forces. Period total fertility rates can change as an effect of 'quantum' of fertility i.e. how many children a couple have and as an effect of tempo of fertility i.e. at which age of the mother births are realized. One source of decreasing period total fertility rates is if one cohort of women postpone their childbearing in comparison to the previous cohort. Similarly, period total fertility rates can increase again if the postponing cohort later catches up its fertility behavior so that cohort fertility rates are equalized between the cohort with the early births and the cohort, that postponed fertility. If the fertility postponing cohort fails to catch up with the previous cohort the quantum of fertility or the cohort total fertility rate shows a decline. Some of the decline is a result of fewer children per family with children, and some will be caused by increasing ultimate childlessness.

In Table 1, replicated from Bosveld (1996) the contribution to period total fertility rates of women aged 30 and over is presented. Starting from 1975 or 1980 most countries in Western, Northern and Southern Europe have experienced substantial increases in the contribution to the period total fertility rates of women aged 30 and over. The most dramatic increase has occurred in the

Table 2. Proportions of childless women by age group

Age and year	Be	Fr	Nl	Wg	No	Sw	It	Po	Hu	Cz
Age 25–29										
1980	31.7	29.5	44.0	40.9	30.1	38.2	31.8	26.4	19.9	16.4
1985	36.3	34.9	54.6	48.6	38.5	45.6	38.7	26.3	18.3	16.5
1990	41.0[1988]	10.6[1989]	61.4	56.9	44.7	48.7	51.1	35.1	20.3[1989]	17.3
Age 30–34										
1980	14.5	13.8	18.3	18.9	13.4	18.4	15.5	13.6	12.1	10.3
1985	16.8	15.3	25.4	25.9	17.1	22.0	17.9	14.3	11.8	9.8
1990	17.7[1988]	17.1[1989]	30.3	31.2	22.0	23.0	23.0	13.3	10.6[1989]	9.9
Age 35–39										
1980	9.6	9.1	11.8	13.1	10.1	13.7	12.0	7.4	9.9	9.1
1985	10.8	9.5	14.6	14.7	10.2	13.9	11.6	9.8	9.9	8.6
1990	11.5[1988]	9.3[1989]	18.6	21.0	12.7	15.4	13.0	10.7	9.8[1989]	8.1

Be: Belgium; Fr: France; Nl: Netherlands; Wg: West Germany; No: Norway; Sw: Sweden; Po: Portugal; Hu: Hungary; Cz: Czechoslovakia.
Source: Bosveld (1996, Table 8.1, p. 216).

Netherlands from 25.6% in 1975 to 51.8% in 1992 occurring to women aged 30 and over. However, it is evident from Table 1 that there is a *U*-shaped pattern in the contribution of women aged 30, and over to the period total fertility rates.

Spain for example in 1960 had 45.6% of all births to women aged 30 and over and in 1992 the figure was almost the same 45.5% having first declined to 34.6% in 1980 and then risen again. The similarity of the figures for 1960 and 1992 of the contribution of women over 30 conceals a decrease in quantum and a decrease in tempo. The total fertility rate in Spain decreased monotonically over the same period from 2.81 to 1.3 and the mean age of the mother at first birth increased from being below age 25 in 1975 to above 27 in 1990, (Bosveld 1996, Fig. 1.1).

The large proportions for Ireland of women over 30 in Table 1 is due to the fact that fertility remained high during the whole period in Ireland in comparison to other countries, whereas the mean age of the mother at first birth remained between 25 and 26 years of age for most of the period covered.

In Table 2 the proportions of childless women per age group in some countries is given. Again, in 1990 the Netherlands stands out with 61.4% of women aged 25–29 childless and West Germany with 56.9%. However, the catching up is larger in the Netherlands, than in Germany, also reflected in the large contribution of women aged above 30 in Table 1 above. Therefore, the ranking of the proportion childless women aged 35–39 is reversed placing Germany in the first place followed by the Netherlands. None of the other countries included arrives at a figure above 13% except Sweden at 15.4%.

Bosveld presents also estimates of ultimate childlessness of women per birth cohort, here replicated as Table 3. By the estimates of Table 3 West Germany, Finland and the Netherlands stand out for the cohort of women born in 1960. But Bosveld also remarks that ultimate childlessness in Italy is high at 14% for the cohort of women born in 1955, and remarks in her conclusions (p. 254) that fertility behaviour in West Germany and possibly in Italy might result in increasing levels of childlessness over the entire lifespan,

Table 3. Estimated proportions of childless women per birth cohort

Generations		1940	1945	1950	1955	1960
Western Europe	Austria	14.3	15.1	20.6	–	–
	Belgium	13.1	12.8	13.4	19.7	–
	France	8.3	8.1	8.3	8.3	10.2
	West Germany*	10.6	12.7	18.8	20.3	22.9^{1958}
	Netherlands	11.9	11.7	14.7	17.8	19.5^{1958}
Northern Europe	Denmark		8.9	10.8	13.7	15.0^{1958}
	England-Wales	11.1	10.2	14.0	16.0	18.0
	Finland	15.2	16.5	17.4	19.1	21.2^{1958}
	Ireland	19.8	17.3	12.2	13.1	14.5^{1958}
	Norway	9.5	9.2	10.0	13.5	–
	Sweden	–	–	10.8	12.6	12.9
Southern Europe	Italy	13.7	11.9	12.2	14.0	–
	Portugal	–	–	11.0	9.7	9.5^{1958}
	Spain	12.0^{1938}	11.0^{1943}	10.0	9.5^{1954}	–
Eastern Europe	Bulgaria	–	7.3	6.9	6.6	5.2
	Czechoslovakia*	7.9	9.2	7.7	7.7	8.3
	Hungary	9.3	10.0	9.6	8.7	8.7
	Poland	–	10.9	9.5	11.4	9.9
	Yugoslavia*	8.9	8.5	8.1	8.7	8.2
	East Germany*	11.3	8.5	7.3	7.5	8.0

Source: Bosveld (1996 Table 1.3 p. 20).
Primary Source Prioux (1993).

because catch-up is relatively small there compared with the delay that has built up.

In Table 4 mean age of the mother at first birth for selected countries is presented. The purpose of this paper is to analyse what economic explanations can contribute in explaining the increasing age of the mother at first birth.

The figures of Table 4 represent the dependent variable of this paper. The remainder of the paper concentrates on explaining the pattern which is visible in Table 4. We observe in Table 4 that there is a *U*-shaped pattern over time with the bottom in 1970 or 1975, i.e. the lowest age at giving birth occurs in all these countries around 1970 or 1975. Age of the mother at first birth first decreases from those births that occurred in 1950 to the lowest level around 1970 and then it increases again to the highest level observed in the data in our latest year of observation. For example, in 1950 in the Netherlands mothers' age at their first birth averaged 26.5 years, in 1970 it had decreased to 24.7 years, in 1991 it had increased to 27.7 years of age and in 1997 the mean age of the mother at first birth was as high as 29 years. There are also clear differences between countries with the East European countries having the youngest mothers. Beets (1997) presents age of the mother at first birth according to birth cohort of the mother and in addition to median age reports figures for the first and third quartiles. The age of the mother at first birth at the third quartile has increased spectacularly comparing the cohort of women born in 1945 to that of women born in 1955. For 15 European countries analysed by Beets (1997), the third quartile is older than age 30 for seven countries namely Ireland, the Netherlands, Sweden, Denmark, England and Wales, Finland and West Germany. For West Germany, the third quartile is as high as 34 years,

Table 4. Mean age of the mother at first birth, selected countries, 1940–1990

	1950	1955	1960	1965	1970	1975	1980	1985	1990	1991	1997
Belgium		25.2	24.9	24.6	24.4	24.4	24.8	25.6			27.0[a]
France	24.7	24.3	24.4	24.1	24.0	24.1	24.6	25.5			28.1[b]
Netherlands	26.5	26.1	25.7	25.1	24.7	25.2	25.7	26.6	27.6	27.7	29.0
West Germany*			24.9	24.2	23.8	24.4	25.0	25.9	26.3	25.9	28.4[c]
Norway							24.3	24.9	25.6	25.8	27.0
Sweden						24.4	25.3	26.1	26.3	26.5	27.3[b]
England-Wales	24.6	24.3	24.0	23.6	23.2	23.6	24.2	24.6	25.0	25.1	26.7[c]
Denmark					23.8	23.9	24.6	25.7	26.4	26.8	27.7[c]
Finland								25.4	26.5	26.6	27.7
Iceland						21.8	21.7	22.8	23.9	24.3	25.0
Ireland						25.0	25.0	25.6	26.2	26.3	27.0
Italy		25.3	25.3	24.9	24.6	24.2	24.6	25.4	26.4		27.5[a]
Portugal	25.6	25.6	25.5	25.3	25.0	24.4	24.0	24.2	24.9	25.1	25.8[c]
Spain						25.1	25.0	25.8	26.8	27.1	27.7[b]
Hungary		23.4	22.9	22.9	22.8	22.5	22.5	22.9			23.4
Czech Republic*		23.2	22.8	22.7	22.5	22.6	22.5	22.5			24.1
East Germany*		23.6	23.0	22.7	22.5	22.5	22.3	22.3			27.3[c]

* Former (countries with border changes around 1990).
Source: Bosveld kindly supplied the figures until 1992, for 1997 or latest year available the source is Council of Europe (1998), Recent Demographic Developments in Europe, Council of Europe Publishing. The following are for a different year than 1997, [a] = 1993, [b] = 1995 and [c] = 1996.

which means that 25% of women older than 34 years have not yet become mothers. Many of these women will be ultimately childless.

3. Current incomes and wages as determinants of timing of maternity

Theoretical research on fertility originally modelled completed family size (Becker 1981; Willis 1974; Hotz et al. 1997). In these models, the expected effect on fertility from husband's income is positive, since his time use is assumed not to be affected by the child bearing and child caring. The expected effect from female wages on the other hand is negative since her time use is extracted away from market earnings and spent on bearing and caring for children. Many studies have found empirical evidence of this negative effect of female wages and positive effect of male incomes on total fertility rates.

One of the most interesting is by Schultz (1985), where he analyses the first fertility transition in Sweden 1860–1910. The most interesting aspect is that Schultz manages to find exogenous changes in the female to male price of time, by studying the ratio of butter to rye prices, thereby using butter price as a proxy to female price of time, since women were dominant in pre industrialist dairy production. The reason these price changes are truly exogenous is, that Sweden lost its competitive position as a grain exporting country during the period under study. In 30 years, from 1860–1890, the ratio of butter to rye prices in Sweden increased by 43%.

For some time Richard Easterlin (e.g. Birth and Fortune 1980) was a rival theory to the mainstream theory. His theory also has a hypothesis for the explanation of timing of maternity. The Easterlin hypothesis says that young people adjust their fertility downwards, if they experience lower standards of living than their parents. This also implies an effect for ageing of fertility as pointed out by Macunovich (1998, p. 100). This is because younger people may adjust their fertility downwards for this reason, and later catch up. The catching up would then not be caused by the Easterlin hypothesis, but the postponement of fertility would. Also, Macunovich points out, that the Easterlin hypothesis says that young people make their decisions based on level of living standards in the family of origin, but when they grow older this influence is very likely weakened.

The econometric timing and spacing literature takes over the hypothesis of a negative effect of female wages and a positive effect of male wages (Butz and Ward 1979; Heckman and Walker 1990; Tasiran 1995; Merigan and St Pierre 1998). In fact, it also takes over the research interest in aiming at explaining development over time of the 'period total fertility rate' or period TFR. The dependent variable in the timing and spacing literature collapses all the different components of the development of the TFR into one measure, the hazard rate. In the words of Heckman and Walker (1990, p. 235): 'Our model explains parity choices, sterility, childlessness, interbirth intervals and initiation of pregnancy within a unified framework'. In this paper it is important to distinguish between tempo of fertility and quantum of fertility because the focus is on the age at first birth, i.e. tempo of fertility, whereas the number of children born, 'the quantum of fertility' is left out. Therefore, it is also not an advantage to discuss the different aspects together for the purpose of this paper.

Another contribution of the Heckman and Walker is the estimation of birth transitions of lower order jointly with birth transitions of higher order instead of, what Heckman and Walker call 'a piecemeal approach of estimating one birth transition at a time'.[1] The main reason for this is that there can be unobserved heterogeneity between individual women's fertility. Heckman and Walker interpret this unobserved heterogeneity as a measure of individual differences in fecundity, but they observe that: "Unlike for societies like the Hutterites where serially correlated fecundity differences play a central role, in accounting for fertility in modern Sweden serially correlated unobservables play a negligible role." This result is in my view not surprising because the Hutterites are a very special sample. This group of people did not practice any family planning and the women averaged 11 births. For these people the timing of the first birth was very decisive on the total number of births. Fecundity certainly in such a setting explains differences in timing and spacing of births. However, in modern European societies we would expect that economic variables would play a more decisive role.

Merrigan and St Pierre (1998) replicate the Heckman and Walker model on Canadian data and find that in 'Canada, a country with two languages and a population with diverse ethnic and cultural backgrounds, they observe non-parametric individual heterogeneity' (p. 40). They motivate the tests for non-parametric individual heterogeneity by a desire to see if 'economic variables would swamp out biological variation' (p. 40).

Heckman and Walker use current wages of males and females to explain fertility transitions. They motivate the use of only current wages, rather than including past and or future wages by the fact, that 'the correlation between past, current and future wages is very large, which makes current wages a good prediction for future wages'. Both in the study of Swedish fertility by Heckman and Walker (1990) and in the study of Canadian fertility by Merrigan and St Pierre (1998) there are significant positive effects of male wages and significant negative effects of female wages. Increases in the average age at first births similar to the figures for some European countries given above have occurred also in the United States (Hotz et al. 1997, p. 281) and in Canada (Merrigan and St Pierre 1998, p. 33).

Tasiran (1995) analysing timing and spacing of births in Sweden using basically the same data set as Heckman and Walker, the Swedish Fertility Survey (SFS) and in addition the Swedish household panel data set (HUS) gets results that contrast with Heckman and Walker (1990). He gets much weaker effects of current male and female wages on birth transitions. Tasiran ascribes this discrepancy to the fact, that in one case he uses individual observations on wages and he also uses a larger time series of aggregate male and female wages, than was available to Heckman and Walker (1990). In my view, the differences in results can also be caused by the fact that Tasiran adds other explanatory variables like parental benefits and childcare into the hazard models. Tasiran also estimates similar hazard models on the American PSID 1985–1988 'Birth History File'. The effect of the female wage rate is statistically significant and positive. He also finds that male income is negative but not always statistically significant. He concludes taking both the results on Sweden and the results on the USA into account (p. 232), 'that the common belief in a negative (female) wage rate effect and a positive (male wage) income effect might not hold generally'.

Do these results of the timing and spacing literature mean that current wages can explain the ageing of fertility in Europe discussed in Sect. 2 above? In a simulation exercise Table 13, p. 268 Heckman and Walker analyse the effect of wage increases on the components of life cycle fertility and find that 'the effect of the female wage on the time to the first birth is especially strong. These results indicate that the strongest effect of wages is on the postponement of first birth'.

There are, however, several reasons for believing that the 'timing and spacing' econometric literature, using current female wage and male income as the main explanatory variables have not given the ultimate explanation. First, theoretical work on fertility decisions emphasize that having children is a lifelong undertaking, which requires lifetime perspective in the economic variables that have an influence. It would be naive to think that a couple will be influenced only by the current wage rather than by lifetime earnings. Second, the theory should explain differences between countries in addition to development over time in one country. Therefore, the effect of public policies should be integrated into the analysis. There are many public policies varying between countries, that have an effect on the economics of the timing decision.

4. The period shadow price of giving birth

There are now more than 10 years of research in life cycle models of fertility. In his presidential address to the ESPE on 8 June, 1989, John Ermisch (1990) emphasizes the Cigno and Ermisch model (1989), also presented in Cigno (1991). Generally in dynamic models, the utility of having a child at time $t+1$ rather than at time t must equal the ratio of the shadow price of having a birth at time $t+1$ rather than at time t. In other words, to have a child earlier in life increases the utility of the parents because they have a longer life together with

A. Opportunity cost

$$\eta_t = \sum_{j=0}^{T-t}((1-\tau_{t+j})w_{t+j}\phi^j - \vartheta_{T+j})\delta_{t+j}$$

+B. Net direct expenditures

$$+\sum_{j=0}^{T-t}(m^j_{t+j} - a_{t+j} + (1-\phi^j)C^j_{t+j})\delta_{t+j}$$

+C. Forgone return to forgone HC investment

$$+\mu_t \sum_{j=1}^{T-1}\left(\sum_{l=0}^{j-1}\phi^l\right)(1-\tau_{t+j})h_{t+j}w_{t+j}\delta_{t+j}$$

η = period shadow price of giving birth
$j = 0$ child's birth year
τ = income tax rate
T = end of planning horizon
w = wage rate
ϑ = parental leave benefit
ϕ = fraction of year spent for parental care
δ = discount rate
m = direct expenditure for child
a = child allowance
c = expenditure on child care
μ = return on human capital
h = fraction of year spent working

Fig. 1. The period shadow price of giving birth (Walker 1995)

the children. Early children also means early grandchildren to enjoy. However, if costs of having a child later in life decreases in comparison to having it earlier, the timing of births results from the tension between having children early in life, in order to enjoy them longer, and the desire to have them when their price is low. It will then be important to analyse what it costs to have a child at any point in life. Hotz et al. (1997:309) state that: 'life cycle models of fertility blend features of static models of fertility with those from at least four different strands of dynamic models of behavior: (1) models of optimal life cycle consumption; (2) models of life cycle labor force participation; (3) models of human capital investment and accumulation; and (4) stochastic models of human reproduction. However, human reproduction has become less stochastic over time as contraceptive technology has improved. Therefore, I think that the loss of insights by leaving uncertainty of conception out of the picture is not serious. A few models have been developed to analyze lifetime earnings as an effect of choice of timing of maternity.

Walker (1995) arrives at a shadow price of giving birth at time t which consists of three terms: (A) the opportunity cost for the time actually spent caring at home away from paid work; (B) the net direct expenditures; and (C) forgone return to human capital investments forgone. The period shadow price of giving birth according to Walker (1995) is reproduced as Fig. 1. The first term consists of the net wage after income tax during the time periods that the woman spends at home caring for her child after deduction for any parental leave benefits. The second term the net direct expenditures consist of direct outlays for housing, clothes, toys, private outlays for schooling, sports, music lessons, etc., minus any child benefits or allowances plus expenditures on children during periods when the mother is working. The third term takes into account investments in human capital forgone during periods not spent in the labor force under the assumption of a proportional rate of return to human capital. Walker then proceeds to put numbers on all the component parts of the period shadow price of giving birth for Sweden for a period from about 1955 to 1990.

One of the theoretical predictions from the intertemporal arbitrage is that holding wealth constant but tilting the profile so that it becomes steeper implies that it is cheaper to have the child soon. In fact, Walker shows using wages of female shop assistants from work of Tasiran and Gustafsson (1990) that later born cohorts of women have faced flatter wage profiles with larger initial wages than earlier born cohorts of women. He uses these data to compute the relative price of fertility at age 35 to age 24 (Walker 1995:245) and finds that it has become cheaper to have children at a later age. The model of Cigno and Ermisch (1989; Cigno 1991) gives the prediction that higher pay per unit of human capital lowers tempo and it suggests that women with steeper profiles will have a slower tempo of fertility (Ermisch 1990:12), i.e. have their children later. Empirical analysis by Cigno and Ermisch (1989) shows that women in occupations characterized by steeper earnings profiles tend to have their children later in life. We have two results here. Walker concludes that *flatter* earnings profile makes later births relatively less costly while Cigno and Ermisch (1989) conclude that a *steeper* earnings profile make later births relatively less costly. These two seemingly opposing results have to do with the woman's career planning problem and emphasize different aspects of the lifetime earnings loss. Before developing on this point a different motive for delaying births will be analysed, namely the desire to have enough income before having children.

5. Consumption smoothing and husband's earnings

In their summary of the literature on timing of births Hotz et al. (1997) give very much reference to Happel et al. (1984). This article is very nice in that it directly addresses the timing of first births and gets predictions out of the model. As pointed out by Hotz, Klerman and Willis this is at the cost of rather strong assumptions. Often one has to choose between a model with strong assumptions, which gives predictions about economic behavior and a model with weaker assumptions and no predictions. This article (Happel et al. 1984) is also the only article reviewed in the "Handbook of Population Economics" on 'The Economics of Fertility in Developed Countries' (Hotz et al. 1997), which addresses the issue of consumption smoothing as a determinant for timing of births. Most of the literature, including my own contributions Gustafsson and Wetzels (2000), address the wife's career planning as the central motive for postponing first birth. Before turning to the career-planning motive in the next section, this section will address the consumption motive. In the consumption smoothing problem the husband's earnings profile matters, and not only the level of the present value of this lifetime earnings. His rising earnings profile matters if there is no opportunity to borrow against

(1) $\text{Max } W(T) = \int_0^T U(x_t + y)\, dt$	prematernity period
$+ \int_T^{T+\tau} U(x_t - e_{T-\tau})\, dt$	parental period with labor force interruption and expenditures
$+ \int_{\tau+T}^{T+\mu} U(x_t + y - e_{t-\tau})\, dt$	parental period II: only expenditures
$+ \int_{\tau-\mu}^{L} U(x_t + y)\, dt$	postparental period

T = time at first birth
τ = periods out of work for mother
μ = periods with child expenditures
L = end of planning horizon
y = wife's earnings
x_t = husband's earnings
e_t = child related expenditures
If U quadratic:

$$(2)\ \frac{dw}{dT} = -U'' \frac{dx}{dt} [y\tau + e\mu]$$

Fig. 2. Consumption smoothing by choice of optimal time at first birth (T) (Happel et al. 1984)

future incomes. In reality, it is often rare that one can borrow against future incomes. However, state loans to finance higher education exist in Sweden and can be obtained by the single condition that an individual is a university student and passes exams at a required pace of time. No account is being taken of parents' earnings or spouse's earnings. Another Swedish example is childcare subsidies, which are available to parents with little or no income at a higher subsidy rate than to parents with higher incomes. This means that University students, if parents, have access to cheap childcare, which they later repay by paying higher taxes because of income tax progressivity once they reap their returns to education later in life.

The assumption of perfect capital markets (PCM), used in that part of the timing and spacing literature, which focuses on the career planning motive, is in the model focussing on consumption smoothing substituted by an assumption of perfectly imperfect capital market (PICM). This rules out saving for the future, which in real life is possible. The distinction between the PCM assumption and the PICM assumption are spelled out in Hotz et al. (1997). Under the PICM assumption, the consumption smoothing motive for postponing birth of first child can be illustrated in Fig. 2. I have chosen to keep the symbols used in the original texts, which help readers who want to consult the original texts. However it means that across the Figs. 1, 2 and 4 some variables who have basically the same meaning though not exactly, are noted by different symbols. We are maximizing utility W in Fig. 2 by choosing optimal time at first birth T. The planning horizon can then be broken down into 4 time periods: first the preparental period in which both husband and wife work in the market, second the parental I period during which the mother interrupts her working life, third the parental II period during which the mother returns to work but there are child related direct expenses as in the parental I period and fourth and finally the postparental period in which there are no longer any child related expenses. Then if the utility of consumption money is a quadratic function as in expression (2) of Fig. 2, we know that when the parenthesis increases, the optimal time till first birth will increase. One example given by Happel et al. (1984) is that policies which decrease the time spent out of work by the mother, such as more daycare for children will lower the tempo of first birth. Note that this is because the second period becomes shorter in Fig. 2 and it applies irrespective of whether the mother has any prospects for human capital investments. In fact, in Fig. 2 mother's earnings are assumed constant over the whole lifetime. In expression (2) also if child expenditures (e) are lowered by a child benefit, if the time until adulthood is shorter as in the case for children with shorter education, and if mother's earnings (y) are smaller the tempo of fertility decreases and women are younger at motherhood. Also in this formulation the optimal time to have the first birth is when husband's income (x_t) is the highest. In the formulation of Fig. 2, mean lifetime consumption is independent of T (time of first birth), therefore, if husbands earnings increase over time, life-cycle utility is maximized when births are delayed to the biological limit. The household smoothes its consumption profile and therefore raises its economic welfare, by delaying the τ periods of the woman's nonemployment and the μ periods of child expenses to a time when the man's earnings are relatively high (Happel et al. 1984:305)

6. Career planning

The potential detrimental effects on the mother's job career is the part of costs of children that has received most attention in economic research. See, for example, Joshi (1990, 1994, 1998) mainly for Great Britain, Dankmeyer (1996) and Meertens (1998) for the Netherlands. These costs consist of two main parts: the direct forgone wages by time spent out of the labor force and the human capital loss. In models analysing the optimal time of giving birth from the point of view of its effects on mother's lifetime earnings we assume that capital markets are perfect allowing borrowing and saving across periods. In what follows, the analysis is carried out for a given time preference for children. Another way of stating this is that we do not consider the benefits of children, only the cost of children. The fact that births are not always delayed to the fecundity limit, is an indication that benefits of children most often work in the direction of earlier births. Also, husband's earnings are left out of the analysis, assuming that his labor market career is not affected by birth timing. In the previous section the consumption smoothing motive of birth timing was seen to give a role to the earnings profile of the husband and also a higher husband's lifetime earnings generally will act as an income effect which enables a couple to fulfil their wish for a child earlier in life. However, we will disregard these influences for now and concentrate on the mother's potential earnings for a given present value of husband's earnings. The determinants of optimal time at maternity then depend on:

1. The woman's accumulated human capital at the beginning of the planning period;

2. the rate at which the woman's job skills decay;
3. the rate of return to human capital investments;
4. the profile of human capital investments;
5. the length of time spent out of the labor force;
6. the size of child quality expenditures.

Some variants of the effect of birth timing on lifetime earnings of the mother are shown in Fig. 3, with panel A showing a relatively steep linear earnings profile, panel B showing a relatively less steep linear earnings profile and panel C showing a non-linear earnings profile. In the following I will discuss how each of the six determinants mentioned above enter into the determination of optimal age at giving birth drawing on the three models Happel et al. (1984) Cigno and Ermisch (1989) as presented in Cigno 1991, Chapt. 8 and Walker (1995). I will also try to evaluate their importance drawing from empirical literature that has come to my attention.

6.1. Preparental human capital

Does a larger amount of preparental human capital have an effect on timing of maternity? A larger amount of preparental human capital will be shown as a larger intercept in Fig. 3. In the Happel, Hill and Low model the larger the preparental work experience the more likely the first birth is delayed. This is because the probability of total skill loss during home time is then less and this was the only case in which an early birth was preferred to a late birth in their model. Cigno (1991:124) concludes by his model that: 'women who enter marriage better endowed with market specific human capital will have fewer children and sooner'. They will have them sooner because of the income effect, because parents have a positive time preference. This result is at odds with the Happel, Hill and Low result that women with larger preparental human capital will tend to delay births. However, the reason the Happel, Hill and Low higher educated women would have delayed births was that they were less at risk of losing all their job skills and not positive time preference for births. Going back to Walker (1995) Fig. 1 above there is no mechanism by which preparental human capital enters the determination of optimal birth timing. However, Walker concludes using wage data for shop assistants that wage profiles have become flatter and at a higher level. This in general considering Fig. 3 would make the direct opportunity cost relatively larger for early births while the capital cost will decline. Therefore, Walker concludes in an interesting simulation Walker (1995, Fig. 13, p. 245) that the relative shadow price of having a birth at age 35 has decreased in comparison to having a birth at age 24. However, empirical work by Gustafsson and Wetzels (2000) shows that higher educated women have their children later than less educated women and they are also the ones that have postponed first birth the most comparing the 1990s to the 1980s.

6.2. Depreciation of human capital due to non-use

In Fig. 3 a situation in which there is depreciation of human capital due to a home time period to give birth and care for a child is represented by a point *p*

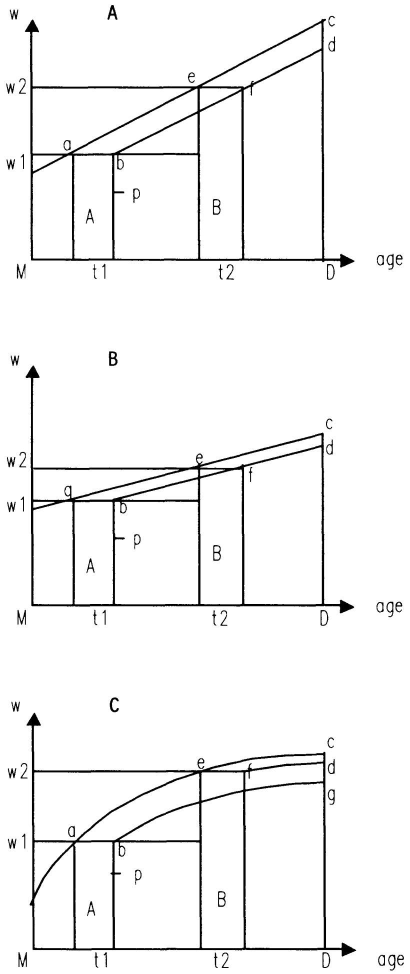

Fig. 3A–C. Timing of first birth and lifetime earnings of the mother

which means that a mother upon return to the labor market will receive a lower wage than she received the year before the labor force interruption. Accordingly cross sectional analysis of women's wages controlling for education and years of labor force experience would find a negative coefficient on years of home time. This is also what Mincer and Polachek (1974) found for data on white American women aged 30–44. However, the size of the depreciation was challenged by Corcoran and Duncan (1979) who claimed that a figure of −0.005 rather than −0.015 found by Mincer and Polachek was more correct. Gustafsson (1981) analysing Swedish cross-sectional data for 1974, matched with panel information on work histories, finds a positive wage coefficient on home time rather than a net depreciation. This positive wage coefficient on home time is however considerably smaller than the positive coefficient on years employed so that periods of home time always decreases future wages. In Fig. 3 this means that the wage at return to work after the labor force interruption is above b. Depreciation of human capital plays an important role in the Happel, Hill and Low model, since in the absence of depreciation timing of birth will have no effect on mother's lifetime earnings. The wife has an amount of human capital accumulated called w at the start of marriage and there is a linear increase in human capital for each year of labor force participation t. Then the earnings loss upon return to the labor market is $\gamma\tau$, where τ is the home time and γ the rate of depreciation. The result is that couples will begin childbearing either very early or very late in marriage. The postponement occurs when only a fraction of the wife's job skills are lost during the labor fore interruption. If most of the job skills are lost then it is better to have the child as early as possible. In fact, the model predicts that the solution is in one of the two corner solutions, either directly after marriage or at the limit of the fecund period. The models by Cigno and Ermisch (Cigno 1991) and by Walker (1995) assume a zero rate of depreciation of human capital.

6.3. The rate of return to human capital investments

In Fig. 3 the slope of the age earnings curve is the product of the rate of return to human capital investments and the investment profile. In the formulation of Cigno (1991) represented in Fig. 4 here the investment profile increases linearly whereas the rate of return to human capital is a constant. (See Fig. 4 Eq. 4.) According to Ermisch (1990:12): 'The model also predicts that higher pay per unit of human capital lowers tempo and it suggests that women with steeper earnings profiles will have a slower tempo of fertility'. The second statement about the steeper earnings profiles according to the Cigno/Ermisch formulation must be due to a higher rate of human capital investments. The reason an increase in the rate of return per unit of human capital lowers tempo, i.e. causes couples to have the first birth early in marriage, is spelled out Cigno (1991:125) in the following way: 'By contrast a rise in ω would cause the opportunity-cost component (wage plus capital loss) of P_t to grow faster if the timing of births were not modified. Since the growth rate of I_t does not depend on ω (see Eq. 6, Fig. 4) the growth rate of the ratio of P_t/V_t would then rise'. Note that P_t is the period shadow price of a birth and V_t is the utility to parents of having a birth at time period t. See Fig. 4, Eq. (3) and (6). But the ratio of the period shadow price of birth to the period utility of birth (P_t/V_t) must according to the equilibrium condition grow at the same rate as the interest rate and if the interest rate is not changed, a smaller P_t must be picked by having a birth earlier. Therefore, if the return per unit of human capital is increased, other things equal couples will have their children earlier in married life. In Walkers' formulation the rate of return to human capital (μ in Fig. 1) is predetermined by calendar time, setting it to 4.1% before 1966, 2.7% 1966–1975 and 1.3% 1976–1989, but he only uses this in his calculation of capital loss, and not in the determination of his wage, which is different from the Cigno (1991) formulation. He uses age specific earnings for female shop assistants, directly observed and he probably does not see the rate of return as a return to 'on the job investments'. Returns to on the job investment in human capital frequently is in the form of job mobility between occupations, which is ruled out by studying wages of only one occupation. Happel et al. (1984) do not discuss the capital cost part of the opportunity cost of child

Parent's utility function:

$$(1)\ U = U(C, B)$$

$B \equiv nq$ i.e. number of children times child quality

For the timing model:

$$(2)\ B = \sum_{t=M}^{D} V_t(I_t)B_t$$

$$(3)\ C = \sum_{t=M}^{D} U_t(C_t)$$

$$(4)\ w_t = \omega k_t \quad k_t = k_M + \beta \sum_{\tau-M}^{t-1} L_t$$

$$(5)\ \sum_{t=M}^{D}[C_t + (I_t + e^D - \Phi)B_t]r^{M-t} \leq A + \sum_{t=M}^{D} L_t W_t r^{M-t}$$

period shadow price of birth

$$(6)\ P_t = e^0 - \Phi + I_t + w_t + \beta\omega \sum_{\tau=t+1}^{D} L_\tau r^{t-\tau}$$

$L_t = m - B_t$
L_t = labor supply of wife
m = work capacity
B_t = the rate of birth at t
I_t = the amount invested in a child born at t
e^0 = minimum costs of procuring a birth of a child
C_t = amount of adult consumption
r = rate of discount
M = date at marriage
w_t = wage at time t
ω = rate of return to human capital
K_t = amount of human capital accumulated at t
K_M = premarital human capital
β = rate of growth of capital

Fig. 4. Model of birth timing (Cigno 1991)

timing, nor do they discuss effects of the size of rate of return to human capital.

6.4. *The profile of human capital investment*

In the Cigno and Ermisch (1989) study the investment profile is central. Women who are employed in occupations with a steeper rising earnings profile will tend to lower their tempo of fertility. They find that women in semi-skilled or manual occupations have earlier births than women in the more skilled clerical occupations. Also, Happel et al. (1984:309) find that women in high skilled occupations have their first child later than women in low skill occupations.

Gustafsson and Wetzels (2000) find that higher educated women delayed first birth more than less educated women comparing births of the 1990s to births of the 1980s for all the four countries studied (Germany, Great Britain, the Netherlands and Sweden). It is easy to see for a zero discount rate in Fig. 3 comparing panels A and B that the capital loss is higher for a steeper earnings profile panel A than for a flatter earnings profile. This works in the direction of delaying births. On the other hand the direct wage cost is bigger for later births. If the earnings profile is convex to the origin rather than linear as in panel C the capital cost of a later birth will be very much smaller than the capital cost of an early birth. Panel C is in accordance with the original Mincer (1974) formulation and is also the shape of the earnings function usually estimated in empirical work. The theoretical foundation for a quadratic earnings function is that life is finite and the investment motive becomes weaker the closer retirement a worker gets, therefore his or her incentive to further improve skills declines since fewer periods are left to reap the returns to investment in human capital. If the rate of human capital investment decreases with age instead of being linear in labor force experience as in all the models discussed in this paper, the incentive to postpone first birth will be stronger since the difference between direct wage loss of a late birth to an early birth is smaller comparing panel C to panel A and the capital loss of a late birth area delineated by efcd is smaller in panel C than in panel A. Recruitment policies of firms into career tracks reinforce the postponement incentives, since they often concentrate on recruiting young talents and it is less common to recruit a person in the postparental phase into a position involving a large amount of training.

6.5. The length of time spent out of the labor force

Does the length of time spent out of the labor force have any effect on the timing of first birth? Happel et al. (1984) answer yes in their model represented as Fig. 2 above. If there are imperfect capital markets and husband's income profile increases with time then the longer the planned time spent out of the labor force the stronger the motive to postpone childbirth. Also by the hypothetical age earnings curves of Fig. 3 it is clear that, if mothering is seen as totally incompatible with paid work it is better to delay childbearing as long as possible because that will decrease lifetime earnings loss.

Dutch mothers until recently, particularly if they have little education have followed such a pattern. Dankmeyer (1996) using a cross section from 1990 finds that low educated women with 2 children born when the mother is 25 and 28 years old will keep only 13% of the lifetime income of a corresponding woman without children. The computation is based on 1) a wage regression, 2) a participation logit, 3) an hours regression and simulations for representative women differing in education. Consequently, the labor force participation rate is predicted for each year for a woman without children and for a two child woman's lifetime earnings using the wage regression and the participation logit. Further, if the probability of labor force participation during one particular year of age exceeds 0.5, the earnings for that year are computed using the hours regression to predict hours of work and then multiply by predicted wages to get income.

Meertens (1998) using a cross section of Dutch women from 1992 estimates a tobit on annual income, where years of schooling is a crucial explanatory variable. The results of Meertens' analysis are more dismal since higher educated women with children by her method only have a marginally larger income than lower educated women with children. This means that women can always increase their lifetime income by having their child later. However, a cross sectional analysis of labor force participation and earnings are a bad predictor for the decision of young women, who now are in the situation of determining when to have their first birth, particularly as now is the case in the Netherlands, there are large cohort effects, where younger cohorts are increasing their labor force participation very much in comparison to older cohorts.

In Walker's work (1995:240) the time absent from the labor market following a birth is the single most important assumption used in the construction of the period shadow price of fertility. He varies his absence assumption from a) short absence i.e. – one year full time absence plus one year part-time employment and b) long absence, which equals six years fulltime withdrawal and one year part-time work. All his cost items shown in Fig. 2 except the direct expenditure for the child (m) depend on this assumption. The direct expenditure for the child (m) are computed in two ways. First based on an estimate from 1969 of household expenditure for a household without children in comparison to a household with children and second just assuming that a child costs 20% of household expenditures. However, Walker's period shadow price of giving birth of Fig. 2 also takes account of child care subsidies, and parental leave benefits which all have lowered the cost of children in Sweden during the period he studies. These factors are not discussed in the Cigno and Ermisch (1989) work.

Walker (1995) takes the approach of using the pattern of absence which are implied by Swedish policy makers namely, one year of fulltime home caring, followed by five years of 3/4 of fulltime work followed by fulltime work for the rest of the lifetime. This is of course a very unjust assumption for the earlier part of the time period he studies, which starts in 1955 and runs until 1990, but it is a reasonable assumption for the more recent mothers, since those policy measures have been introduced gradually starting in 1974 and becoming in full effect by 1989 (Gustafsson 1984, 1994; Gustafsson and Stafford 1992; Sundström and Stafford 1992).

A similar approach is used by Gustafsson and Wetzels (2000). Their result is that postponement of births will always be beneficial in the sense that the lifetime earnings loss is diminished. This is true also for stepping out of work only one year and then return to fulltime work, although in this case the costs are a very small proportion a few per cent of a life time income.

6.6. *The size of child quality expenses*

The more money parents spend on a child including educational expenses and the longer the period parents keep paying for their child, the higher will be the shadow price of a child. In the Happel et al. (1984) model with imperfect capital markets the larger the expenses are the later in life the parents want to have the child. This is if the consumption smoothing motive is important and if wages increase over lifetime. High educated people who have steeply rising earnings want to have high educated children, who therefore cost much. This motive will therefore strenghten the incentive for high educated people to delay parenthood.

7. Biological – medical aspects

In this paper it has been shown that age of the mother at first birth has been increased in most European countries with the exception of Eastern Europe until 1990. We have seen that there are many mechanisms which tend to push maternity towards the biological limit. The economic theoretical models discussed above all assume that the biological fecundity limit is known to the decision making couple. The closer a woman gets to her fecundity limit the more important it becomes for her to have knowledge about possible consequences of late motherhood.

One study Gilbert et al. (1999), addresses the question of childbearing beyond age 40. The study is based on data from birth certificates and hospital discharge records of all births that occurred in Californian hospitals during the period January 1, 1992 including December 31, 1993. There were 24,032 deliveries of women aged 40 and over which corresponds to 2% of all births in this two year period, of these 20% 4,777 were nulliparous i.e. had their first birth. The data of the older mothers were compared to a control group of women who were 20–29 at giving birth. The results are, that women who had their first child above age 40, have a higher risk of operative delivery (61% of which Caesarean 47%) than do younger nulliparous (35% of which Caesarean 22,5%). The study also presents medical problems of mother and child around delivery such as prematurity, malpresentation and obstructed labor and 8 other disorders. Older first time mothers are over represented as compared to younger first time mothers on all reported disorders. Older first time mothers also have children with lower birth weight and have a shorter gestational period than the control group. The authors hope that: "These data will allow us better to counsel patients about their pregnancy expectations and possible outcomes" (Gilbert et al. 1999:9). This article is a warning against increased complications around childbirth but does not say anything about any lasting consequences of being an older mother.

In order to say something about fecundity limits one would need data on probability of successful pregnancies among women who intend to become pregnant. Another study Waldron, Weiss and Hughes (1998:216) find that: "as predicted by the Age-Related Parental Role Strain Hypothesis: younger age at first birth, particularly a teenage birth, appeared to result in more harmful health effects". The data of this study is from the United States National Longitudinal Study and the study of mother's age at first birth does not include women who had their first birth after age 30. This study includes long-run effects on the health of mothers since in some cases the mothers are followed until their children are teenagers. However, there are no long run data on the health of children. What one can conclude from the two studies mentioned that it is not good to be a very young mother, and there are some immediate negative effects from being a first time mother aged 40 and over.

By a study (Treloar 1981) cited by Dorland, Kooij and te Velde it is concluded that 12.5% of women were in the menopause at age 46 and therefore a similar percentage would have fecundity problems already at age 36. At age 31 half of all women have experienced decrease in fecundity, at age 41 half of all women have experienced infecundity and at age 51 half of all women have reached the menopause.

Medical research also concludes, that it is very likely, that the age at menopause is genetically determined and also the age determined decrease in fecundity. Generally, the quality of the egg cell deteriorates which results in smaller likelihood that the egg cell gets planted and a larger risk of chromo somatic changes as Down's syndrome. The conclusion of Dorland et al. (1997:39) is that about 50% of women aged between 30 and 40 will have mild or serious fecundity problems. These women will make use of medical assistance for these problems. Therefore, medical fecundity treatment therefore will become a rather normal procedure in family planning.

Den Ouden et al. (1997) summarize medical research about health consequences for mother and child. They point out that the probability of successful pregnancy already at age 30 of the woman decreases and that the fecundity of the man from age 45 decreases. They summarize the problems of late births as the following consequences for the mother: longer waiting time to conception, more miscarriages, more multiple births, more pregnancy complications, more caesareans and more breast cancer. The consequences for the child of being born to an older mother are: more still births, more infant deaths, more premature births, more chromo somatic problems and more often learning problems.

8. Conclusions

In this paper it has been shown that European women have delayed their age at first birth substantially since the 1970s and that the postponement is large also comparing births during the 1990s to those of the 1980s. Economic theorizing on the optimal age at giving birth in general weighs the pleasure of early births against the lower cost of later births. One model reviewed (Happel et al. 1984) finds that with perfectly imperfect capital markets the man's income profile matters and its effect is to delay births until a moment when costs of the child can be offset by his higher earnings. The most important factor which works for later births is the woman's career costs. These costs are basically of two kinds her direct wage loss during labor force withdrawals and her loss of human capital investments and returns to these investments. Parameters which have effects on the career costs outcome are of five different kinds:

1. The amount of prematernity human capital;
2. the rate of depreciation of human capital due to non-use;
3. the rate of return to human capital investments;
4. the profile of human capital investments;
5. the length of time spent out of the labor force.

Increases in all these parameters tend to make postponement of birth more favorable except in one formulation by Cigno and Ermisch (1989) where a large initial human capital by its income effect will bring forward births. The analysis shows that any public policies that will lead to decreases in time spent out of work, which is the most important part of the period shadow price of giving birth will have an effect of decreasing age at maternity. Biological medical research shows that the risks around pregnancy and delivery increase for first time mothers aged 40 and above, that the probability of conception decreases for some women already after age 30 and that about half of a population will have mild or serious fecundity problems already from age 36. Research about long run effects on health of mothers and children born to older mothers are also not very optimistic. Therefore, there are good arguments for governments to consider political measures like paid parental leaves, subsidized childcare and consider the school and work organizations so as to facilitate the combination of either being a student and a mother or being a worker and a mother. Such measures will again have a tendency to decrease age of maternity, which will lead to healthier and happier mothers and children and more generally happier fathers and grandparents.

Endnotes

[1] In the mid-1980s I was a visiting scholar to the University of Chicago, and I was told by Jim Walker at that time, that the computer program took 6 hours to estimate. I thought, that there were never going to be any replications of their model but technology develops and we now have replications by other authors and other data sets) e.g. Tasiran (1994) and Merrigan and St Pierre (1998).

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The effects of female employment status on the presence and number of children

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Abstract. The main concern of this paper is to analyze the effects of female employment status on the presence and number of children in households in the Netherlands. For this purpose a hurdle count data model is formulated and estimated by the generalized method of moments. The hurdle takes explicitly into account the interrelationship between female employment status and timing of first birth. The number of children, once children are present in the household, is modeled conditional on female employment status. The empirical results show that female employment status is a major determinant of the presence and number of children in households: employed women schedule children later in life and have fewer children compared to non-employed women, holding educational attainment constant. After controlling for female employment status, the educational attainment of both the woman and the man in the households are found to have relatively small effects on the presence and number of children.

JEL classification: C35, J13, J20

Key words: Hurdle count data model, fertility, female employment

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1. Introduction

The main concern of this paper is to analyze the effects of female employment status on the presence and number of children in households in the Netherlands. For this purpose a hurdle count data model is formulated and estimated. The hurdle takes into account the interrelationship between female employment status and timing of first birth. Once children are present in the household, the number of children, i.e. the count variable, is modeled conditional on female employment status. This approach takes the endogeneity of female employment status explicitly into account and makes it possible to disentangle the effects of female employment status and educational attainment of both the man and woman in the household on the presence and the number of children. In the empirical analysis, only the conditional expectation of the number of children is specified and a generalized method of moments estimator is employed. This approach relaxes the distribution assumptions that are conventionally made when estimating count data models using fertility data.

The outline of the paper is as follows: Section 2 discusses the relevant literature and the main contribution of this paper to the literature. Section 3 discusses the data. Section 4 formulates the hurdle count data model, specifies the moment conditions and describes the estimation procedure. Section 5 discusses the estimation results and Sect. 6 concludes.

2. Previous empirical studies

Becker (1960) argues that socioeconomic variables affect the fertility decisions of households and that fertility decisions can be analyzed within an economic framework. He emphasizes the relationship between household income and the number of children and concludes that there is a positive correlation between household income and the number of children, after controlling for contraceptive knowledge. Most research on the analysis of the number of children builds on this pioneering study of Becker and is concerned with the determinants of the number of children in a household (see for instance, Willis 1973, and more recently Siegers 1985; O'Malley Borg 1989; Cigno 1991). During the last two decades the emphasis has been on life-cycle fertility behavior and research has shifted towards investigating the timing of births rather than completed fertility (see for instance, Newman and McCulloch 1984; Heckman and Walker 1990; Groot and Pott-Buter 1992). These studies on fertility dynamics employ hazard rate models to analyze the timing and spacing of births. One of the empirical findings in the literature on fertility dynamics is that women with high opportunity costs of having children (e.g. high wage women) schedule births later in life and have fewer children compared to women with low opportunity costs. An important implication of the empirical findings in these studies on household fertility decisions is that the presence and number of children are not exogenous constraints imposed on the household decision making but are outcomes of household decisions and are affected by socioeconomic variables. Furthermore, the female labor supply literature provides more than sufficient evidence that the presence of children has a significant negative effect on the female employment probability (see for instance, Heckman and Macurdy 1980; Mroz 1987). As a consequence, one may consider female employment and fertility decisions to be closely interrelated. In order to get a better understanding of life-cycle fertility behavior and its determinants, one cannot ignore this interrelationship with life-cycle female employment, and vice versa.

Conventionally, empirical studies investigating the interrelationship between female employment and fertility decisions at the household level employ a simultaneous equations model. Studies such as Willis (1973) and Siegers (1985) use a static framework and investigate jointly the female labor supply decisions and completed fertility, in line with the pioneering study of Becker (1960). Later studies of Blau and Robins (1989); Hotz and Miller (1988); Moffitt (1984); Walker (1995) and Bloemen and Kalwij (1996) model birth decisions rather than completed fertility jointly with the female employment decisions. The empirical analysis of Hotz and Miller is restricted to couples who have at least one child. This may be rationalized in their approach, but especially around the birth of the first child the interrelationship between fertility and female labor supply is observed to be strongest, hence it would be desirable to model this. Moffitt (1984) skillfully demonstrates the importance of taking the interrelationship into account but his results also suggest that timing issues cannot be investigated properly using a static econometric framework. Blau and Robins (1989) have taken a dynamic approach. Although they acknowledge the importance of the interrelationship between fertility and labor supply decisions, the econometric framework utilized (a competing risks model) does not allow for this. Basically they implicitly assume independence between female labor market transitions and the timing of births. Walker (1995) and Bloemen and Kalwij (1996) utilize a dynamic econometric model, a multiple state transition model, which explicitly takes the interrelationship between the female employment and fertility decisions into account. Such an approach makes it possible to analyze the effects of socioeconomic characteristics of the household on the timing of births and lifecycle female employment simultaneously and, consequently, the number of children at the end of a woman's fertile period.

Given the main concern of this paper, this latter approach of employing a multiple state transition model seems most appropriate[1]. However, the data requirements for estimating such a dynamic model are high: panel data with a large time dimension or retrospective data on the complete female employment and fertility histories. Usually a researcher has available only cross-section data or panel data with a short time dimension. Fertility history may be reconstructed on the basis of the age of the children in the household but the complete labor market history of the woman in the household will be more difficult or even impossible to reconstruct. Typically one observes for each household the number of children present and the employment status of the woman in the household at the time of interview. The observed values of these two variables are the outcomes of a sequential decision-making process of the household up to the time of interview. Therefore this paper adopts a count data model to analyze the effects of female employment status on the number of children. Conceptually such a model takes the underlying dynamic nature of the stochastic process into account and can be estimated on a single cross-section.

Count data models have already been applied in the demographic literature. Typically, the numbers of children observed are assumed to be realizations of a (generalized) Poisson process (e.g. Winkelmann 1995 and Wang and

Famoye, 1997) or of a more complex process taking hurdles into account (Santos Silva and Covas 1998). (An excellent discussion on hurdle count data models can be found in Mullahy (1986, 1998) and Cameron and Trivedi (1998).) These studies either do not model female employment status or include it as an exogenous explanatory variable. If female employment status is included as an explanatory variable then it is almost always found to be a dominating variable, in both the relative impact on the number of children and the level of significance. However, a discussion regarding the possible endogeneity of this variable is usually absent. If one acknowledges a possible interrelationship between female employment status and the presence and number of children, then one important reason for not explicitly modeling female employment status is that one cannot estimate a simultaneous equations model for female employment and the number of children using a Poisson based count data model. See, for instance, Windmeijer and Santos Silva (1997) for a discussion on this. Basically, for internal consistency reasons, one needs to assume that the presence and number of children does not affect the female employment probability. This makes it extremely hard, if not impossible, to come up with an instrument to identify the effect of female employment status on the number of children.

Santos Silva and Covas (1998) demonstrate the importance of taking hurdles into account when modeling completed fertility and convincingly argue that hurdles may be the reason for the observed underdispersion characterizing completed fertility data. Given the discussion above, a woman (or household) presumably faces the largest hurdle at the time of first birth, largely because of the interrelationship with female employment status and the timing of first birth. A hurdle count data model is not only considered to be a more appropriate way of modeling household fertility, relatively to a standard count data model, it also makes it possible to take into account the simultaneity between the presence of children and female employment status. This approach partly solves the limitation of modeling simultaneously female employment status and the number of children using a generalized Poisson regression model. This is the route followed in this paper and is considered to be the main contribution to the literature. As a consequence of using such an approach, the effects of female employment status and the educational attainment on the number of children can be disentangled.

3. Data: the Dutch SocioEconomic Panel

The empirical analysis is based on micro-data from the SocioEconomic Panel (SEP) of the Netherlands. At the time of starting this research all waves from 1986 up to and including 1994 were available. About 5000 households respond to the survey in each wave. There can be more than one respondent per household and each respondent is asked questions about his socioeconomic and demographic situation. (A respondent is a person at least 16 years old. In principle each person in the household over 15 should complete the questionnaire.) Up to 1990 the survey has been conducted twice a year, a wave in April and a wave in October. The relevant questions for this paper are asked in October. From 1990 onwards the survey has been conducted only once a year and all information is collected in May. Although the empirical analysis of this paper is based on panel data, the econometric model proposed in Sect. 4

can be estimated on a single cross-section. Under the assumption that there are no calendar time effects, panel data is required to identify both birth-cohort and age effects (see Sect. 5).

3.1. Sample selection

The sample is restricted to married and cohabiting women who are at most 40 years of age. The total number of children born to a woman is not observed directly and has to be inferred from the observed number of children in the household at the time of the interview. Households in which the woman is over, approximately, 40 years of age are observed to reduce in size because of children leaving the parental home. For this reason the age of 40 is chosen as the upper bound in order to reduce the potential problem of underestimating the number of children of a household. Household formation is not modeled and for this reason the sample is restricted to married and cohabiting women. The resulting sample has information on 2416 households over the period 1986–1994 (11391 observations in total). (About 900 observations were deleted from the sample because of missing values on the educational attainment variables of the man or woman in the household.) All results in this paper are conditional on this selection. Addressing possible selection problems is beyond the scope of this paper.

3.2. Descriptive statistics

Table 1 reports the number of years households are observed. This shows that about 20% of the households remain in the panel for 8 or 9 years. These households attribute to almost 50% of the observations. Table 2 reports the sample statistics per year of the variables used in the empirical analysis. A woman is on average about 33 years of age and the man in the household is on average 2 to 3 years older. There is slight increase over time in the percentage of higher educated women. The educational attainment of men is more or less stable over time. Employment is defined as having a paid job (full or part-time). A woman who has a job but works zero hours because of maternity leave is registered as being employed.[2] The employment rate is about 38%, including women on maternity leave. About 80% of the households have children and the average number of children for households who have children is just over 2. Table 3 reports the employment behavior before and after birth of the first child for each level of education. This table is based on a subsample of 339 households in which the woman has given birth to her first child during the observation period. Table 3 shows that the average age at which a woman gives birth to her first child increases with the level of education: from 27 years for women with education level 1 to 31 years for women

Table 1. The number of years households are observed in the panel

Number of years	1	2	3	4	5	6	7	8	9	Total
Number of households	441	288	301	163	247	195	172	271	338	2416

Table 2. Number of observations (NOB) and sample means of all relevant variables per year

Year	1986	1987	1988	1989	1990	1991	1992	1993	1994
NOB	1567	1462	1283	1357	1372	949	1221	1129	1051
Variable	Sample means								
Age of the woman	31	32	32	32	32	33	33	34	34
Year of birth	'55	'56	'56	'57	'58	'59	'59	'60	'60
Educational attainment of the woman[a]									
level 1	0.33	0.33	0.32	0.29	0.31	0.34	0.29	0.28	0.27
level 2	0.48	0.48	0.48	0.49	0.47	0.45	0.50	0.50	0.50
level 3	0.19	0.19	0.20	0.22	0.22	0.21	0.21	0.22	0.23
Age of the man	34	34	35	35	35	36	36	37	37
Educational attainment of the man[a]									
level 1	0.19	0.18	0.18	0.15	0.17	0.20	0.18	0.19	0.18
level 2	0.52	0.52	0.53	0.54	0.52	0.50	0.52	0.52	0.51
level 3	0.29	0.29	0.29	0.31	0.31	0.30	0.30	0.30	0.31
Employment status[b]	0.38	0.39	0.38	0.40	0.36	0.34	0.38	0.39	0.33
Presence of children	0.76	0.75	0.77	0.77	0.79	0.83	0.81	0.81	0.81
Number of children[c]	2.00	2.02	2.06	2.06	2.08	2.09	2.12	2.18	2.15

[a] Level 1 is at most primary or secondary education, level 2 is intermediate vocational education (MBO) and level 3 is higher vocational education (HBO), a university degree or higher.
[b] Equal to 1 if the woman is employed, 0 otherwise.
[c] Only for households with children.

Table 3. The average age of the women (Age), the average number of children (Kids) and the female employment rate (ER) before and after the birth of the first child for each level of education. YB is the year from first birth (for instance, YB = −5 is defined as 5 years before the birth of the first child) and n denotes the number of observations

	Education level 1				Education level 2				Education level 3			
YB	n	Age	Kids	ER	n	Age	Kids	ER	n	Age	Kids	ER
−8	1	26	0	1.00	7	23	0	1.00	2	22	0	1.00
−7	0	–	–	–	17	23	0	0.88	2	25	0	1.00
−6	1	18	0	0.00	26	24	0	0.92	14	27	0	0.86
−5	7	25	0	0.86	39	25	0	0.95	15	27	0.07	0.87
−4	11	25	0	0.82	58	26	0.02	0.95	33	28	0.06	0.91
−3	18	25	0	0.83	82	26	0.01	0.93	45	29	0.09	0.87
−2	34	25	0	0.71	114	26	0.04	0.88	62	29	0.06	0.91
−1	43	27	0	0.63	149	27	0	0.91	67	29	0	0.87
0	76	27	1.01	0.17	175	28	1.05	0.39	88	31	1.07	0.66
1	67	27	1.07	0.09	161	29	1.10	0.23	87	32	1.13	0.62
2	63	28	1.33	0.06	145	29	1.52	0.21	74	32	1.45	0.42
3	56	29	1.55	0.02	134	30	1.83	0.23	56	33	1.80	0.48
4	46	29	1.83	0.07	120	31	1.95	0.18	49	33	2.00	0.43
5	39	30	1.95	0.03	93	32	2.06	0.18	38	33	2.08	0.42
6	34	31	2.06	0.03	66	32	2.18	0.14	29	34	2.31	0.41
7	25	32	2.12	0.08	44	33	2.34	0.16	15	35	2.40	0.53
8	17	33	2.29	0.12	19	34	2.53	0.16	5	36	2.40	0.60

with education level 3. The female employment rate drops for lower educated women from 63% in the year before to 9% in the year after the birth of the first child and for higher educated women from 87% in the year before to 62% in the year after the birth of the first child. The differences in the employment rate before the birth of the first child are relatively small compared to the difference after the birth of the first child. More importantly, relatively to lower educated women, higher education women are less likely to leave employment after the birth of their first child. The average number of children is over 2 for all women 8 years after the birth of the first child. There is some evidence that higher educated women have their children in a relatively smaller time span, compared to lower educated women.

One assumption made in the empirical analysis is that a woman's employment status remains unchanged after the birth of the first child. Table 3 shows that the percentage of women who continue in employment after the first birth remains roughly constant. Or at least, there is very little evidence that women who stop working to give birth return to work shortly after the birth of their first child. Women appear to determine their employment status for the period after the birth of the first child around the birth of the first child. Bloemen and Kalwij (1996) using different data from the Netherlands make a similar observation.

Figure 1 shows that higher educated women are more likely to be employed than lower educated women, at all ages. Figures 2 and 3 show that the higher educated women schedule children later in life and, perhaps less convincingly, have fewer children than the lower educated women. These figures show that observed employment and fertility behavior already confirms most findings of the studies discussed in the introduction. Figure 4 shows that female employment status has a large impact on both the timing of children and the number of children. Figure 4 also shows that, after controlling for female employment status, educational attainment of the woman has relatively little

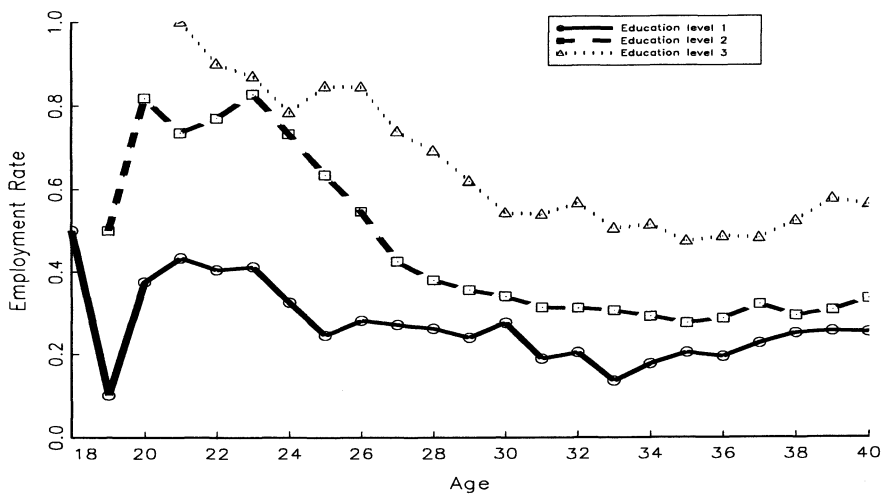

Fig. 1. Female employment rate by age and education level

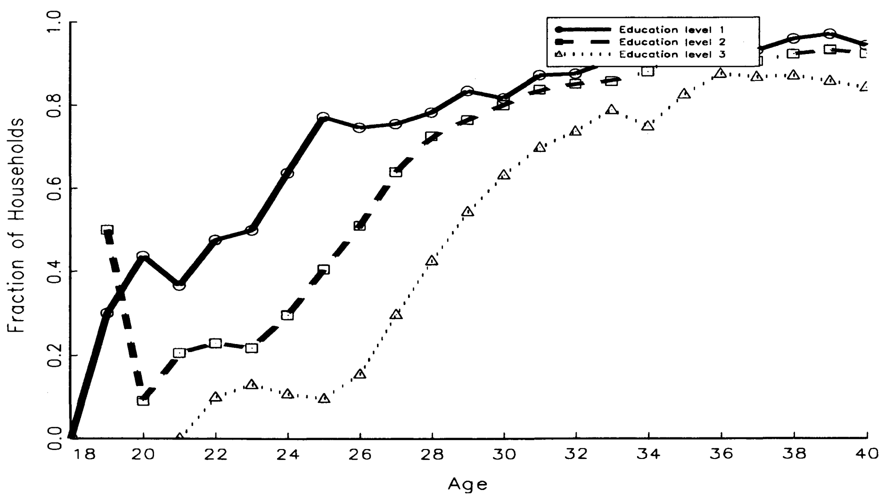

Fig. 2. The fraction of households with children by age and education level

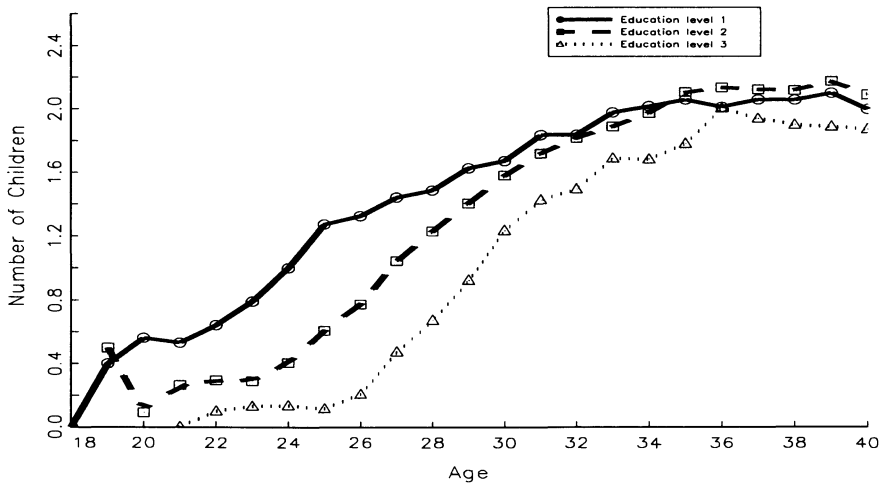

Fig. 3. The average number of children by age and education level

impact on the number of children of 40 year old women. Of course, educational attainment of the woman is highly correlated with educational attainment of the man. This and possible birth-cohort effects make any inferences based on these figures ambiguous. Therefore the main purpose of the econometric analysis in the next sections is to disentangling the effects of female employment status and educational attainment of the woman on the presence and number of children, controlling for educational attainment of the man in the household and the birth-cohort.

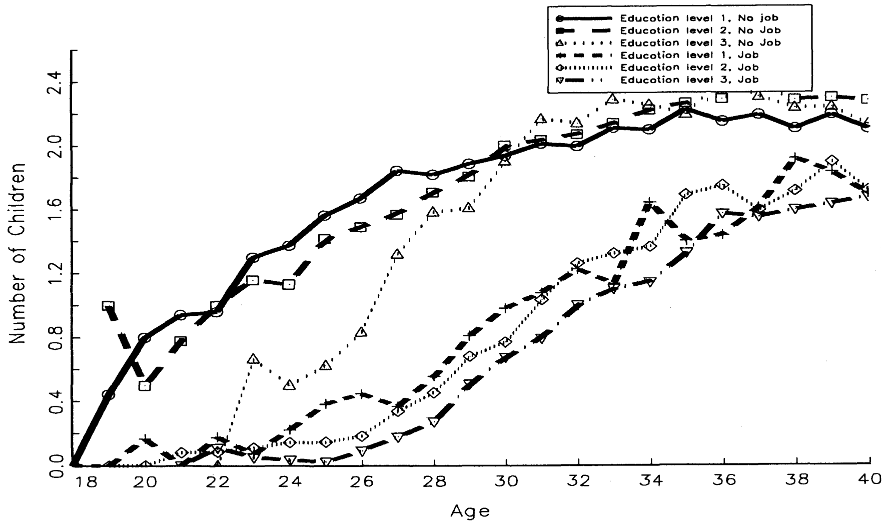

Fig. 4. The average number of children by age, employment status and education level

4. Empirical framework

4.1. Model outline: a hurdle count data model

A woman is assumed to decide simultaneously on her employment status and the timing of births. After the birth of the first child she is assumed to no longer change her employment status. Hence, she is assumed to decide simultaneously whether to combine having children with work or not. The observed number of children and female employment status at any given period in time are assumed to be the outcome of this decision making process. To model these outcomes a hurdle count data model is employed in which the interrelationship between the presence of children and female employment status is taken into account in the hurdle. The number of children, once children are present in the household, i.e. the count variable, is modeled conditional on female employment status. Hence, female employment status after the birth of the first child is assumed to be a predetermined explanatory variable for the number of children, conditional on the presence of children. Furthermore, conditional independence between the hurdle and the count process is assumed. The proposed reduced form model can be used to analyze the observed behavior of all women in the sample under 40 and is not restricted to women who completed the childbearing period.

4.2. The moment conditions and estimation procedure

The dependent variable is the number of children in the household and is denoted by Y_{th}, where t is the time index and h is the household index. The expectation of Y_{th} conditional on some exogenous household characteristics (z_{th})

is denoted by $E[Y_{th}|z_{th};\theta]$, where θ denotes the parameters of interest. Typically one is interested in the marginal effects of some exogenous characteristic on the expected number of children: $\partial E(Y_{th}|z_{th};\theta)/\partial z_{th}$. If one is prepared to specify the distribution of Y_{th} the maximum likelihood estimator can be used and, under the assumption of a correctly specified distribution, leads to the most efficient consistent estimates of θ (see for instance, Mullahy 1986; Pohlmeier and Ulrich 1995). If one assumes Y_{th} is Poisson distributed one can employ a pseudo-maximum likelihood estimator which results in consistent estimates even if the true distribution is not Poisson (see for instance, Cameron and Trivedi 1986). Unfortunately, for a hurdle type of distribution no pseudo-likelihood type result is known. Hence, misspecification of the distribution results in inconsistent estimates. For this reason a generalized methods of moments estimator (e.g. Hansen 1982), is employed as proposed by Santos Silva and Windmeijer (1998) and Mullahy (1998). Only the first moment of the distribution is specified, hence the marginal effects $\partial E(Y_{th}|z_{th};\theta)/\partial z_{th}$ are identified.

Given the model outlined above and using the law of iterative expectation, the conditional expectation of Y_{th} can be written as follows:

$$E[Y_{th}|z_{th};\theta] = E_{I_{(Y_{th}>0)},W_{th}}\lfloor E\lfloor Y_{th}|I_{(Y_{th}>0)}, W_{th}, z_{th};\theta\rfloor \,|\, z_{th};\theta\rfloor, \tag{1}$$

where W_{th} denotes female employment status ($W_{th} = 1$ if the woman is employed and 0 otherwise). $I_{(Y_{th}>0)}$ is an indicator function for the presence of children. Equation (1) shows that the interrelationship between children and female employment status is only allowed for in the first step, i.e. the hurdle. In the second step, once children are present in the household, female employment status is assumed to be predetermined. Given the binary nature of female employment status and the indicator function, Eq. (1) can be written as follows:

$$\begin{aligned}
&E[Y_{th}|z_{th},\theta] \\
&\quad = P(I_{(Y_{th}>0)} = 1|z_{th},\alpha) \\
&\qquad \times \{P(W_{th} = 0|I_{(Y_{th}>0)} = 1; z_{th},\alpha)E[Y_{th}|W_{th} = 0, I_{(Y_{th}>0)} = 1; z_{th},\beta_1] \\
&\qquad + P(W_{th} = 1|I_{(Y_{th}>0)} = 1; z_{th},\alpha) \\
&\qquad \times E[Y_{th}|W_{th} = 1, I_{(Y_{th}>0)} = 1; z_{th},\beta_2]\}.
\end{aligned} \tag{2}$$

where $\theta^T = (\alpha^T, \beta_1^T, \beta_2^T)$. Note that in the case where there are no children in the household, i.e. $I_{(Y_{th}>0)} = 0$, the conditional expectation of Y_{th} is equal to 0, irrespective of female employment status.

Once the functional forms of the choice probabilities and the conditional expectation of the number of children are specified, one can obtain estimates of all parameters of interest using the empirical moment condition implied by (2). However, in practice this is extremely difficult and for this reason the moment conditions implied by the conditional moment condition (2) are used to estimate all parameters of interest stepwise. In total, six moment conditions are formulated:

$$E\lfloor w_{th}(1 - I_{(y_{th}>0)}) - \Pr(W_{th} = 1, I_{(Y_{th}>0)} = 0|z_{th}, \alpha)\,|\,z_{th}\rfloor = 0; \tag{3}$$

$$E\lfloor(1 - w_{th})I_{(y_{th}>0)} - \Pr(W_{th} = 0, I_{(Y_{th}>0)} = 1|z_{th}, \alpha)\,|\,z_{th}\rfloor = 0; \tag{4}$$

$$E\lfloor w_{th}I_{(y_{th}>0)} - \Pr(W_{th} = 1, I_{(Y_{th}>0)} = 1|z_{th}, \alpha)\,|\,z_{th}\rfloor = 0; \tag{5}$$

$$E[(y_{th} - E[Y_{th}|W_{th} = 0, I_{(Y_{th}>0)} = 1; z_{th}, \beta_1]\,|\,z_{th}] = 0,$$
$$\text{if } w_{th} = 0 \quad \text{and} \quad y_{th} > 0; \tag{6}$$

$$E[(y_{th} - E[Y_{th}|W_{th} = 1, I_{(Y_{th}>0)} = 1; z_{th}, \beta_2]\,|\,z_{th}] = 0,$$
$$\text{if } w_{th} = 1 \quad \text{and} \quad y_{th} > 0; \tag{7}$$

$$\begin{aligned} E[y_{th} &- P(W_{th} = 0, I_{(Y_{th}>0)} = 1|z_{th}, \alpha)E[Y_{th}|W_{th} = 0, I_{(Y_{th}>0)} = 1; z_{th}, \beta_1] \\ &- P(W_{th} = 1, I_{(Y_{th}>0)} = 1|z_{th}, \alpha) \\ &\quad \times E[Y_{th}|W_{th} = 1, I_{(Y_{th}>0)} = 1; z_{th}, \beta_2]\,|\,z_{th}] = 0. \end{aligned} \tag{8}$$

The observed number of children is denoted by y_{th} and the observed female employment status is denoted by w_{th}. The estimation of the parameters of interest can be done in two stages. In the first stage GMM estimates of α are obtained by exploiting the moment conditions (3) to (5). Given the binary nature of the dependent variables entering these three moment conditions, specifying the probability distribution function and employing a maximum likelihood estimator yields identical estimates. In the second stage the sample is restricted to those couples who are observed to have children ($y_{th} > 0$) and the resulting sample is split on the basis of the observed female employment status (w_{th}). Estimates of β_1 and β_2 are obtained by exploiting the moment conditions (6) and (7). The additional moment condition (8) yields overidentification and is used for a specification test. For this purpose, a conditional moment test statistic is constructed based on the estimates obtained from using the first five moment conditions (see e.g., Davidson and MacKinnon, Chapt. 16).

For the empirical analysis one needs to specify the probability distribution function of female employment status and the presence of children and the conditional expectation of the number of children. The joint distribution of female employment status and the presence of children is specified as follows:

$$P(W_{th} = i, I_{(Y_{th}>0)} = j|z_{th}, \alpha) = \frac{\exp(z'_{th}\alpha_{(i,j)})}{\sum\limits_{(i,j)\in S} \exp(z'_{th}\alpha_{(i,j)})}, \tag{9}$$

where S is the set of feasible alternatives, $S = \{(0,0), (1,0), (0,1), (1,1)\}$. This probability distribution function is the well-known multinomial distribution and the normalization chosen is $\alpha_{(0,0)} = 0$, hence $\alpha^T = (\alpha^T_{(1,0)}, \alpha^T_{(0,1)}, \alpha^T_{(1,1)})$. Alternative one can choose a bivariate probit model. However, in a bivariate probit model the interrelationship between female employment status and the presence of children is only allowed for through the error terms. The multinomial logit model allows for an interrelationship through the observable variables as well. For instance, the multinomial model allows educational attainment to affect the presence of children differently when the woman is employed than when she is not employed. For this reason a multinomial logit model is favored in this paper. The multinomial logit model yields a more flexible empirical specification with respect to the observable characteristics but at the costs of imposing a restriction on the relationship through the error terms. Basically, the case that the unobservables determining female employment status and the presence of children are not correlated is not nested within this model. In contrast to the univariate case (probit versus logit), a comparison between the bivariate probit and multinomial logit model is complex and there is no straightforward link between the parameter estimates obtained from these two models.

The conditional expected value of the number of children, once children are in the household, is specified as follows:

$$E(Y_{th}|W_{th}=0, I_{(Y_{th}>0)}=1; z_{th}, \beta_1) = 1+\exp(z'_{th}\beta_1); \tag{10}$$

$$E(Y_{th}|W_{th}=1, I_{(Y_{th}>0)}=1; z_{th}, \beta_2) = 1+\exp(z'_{th}\beta_2). \tag{11}$$

This specification ensures that the expected number of children conditional on the presence of children is always greater than one.[3] Conditional on the correct specification of the first moments, the moment conditions as specified in (3) to (7) yield consistent GMM estimates of α, β_1 and β_2.

The assumptions underlying this model as outlined in Sect. 4.1 may considered to be strong but the resulting model is conceptually less restrictive than the count data models employed in the empirical studies mentioned in Sect. 2. In the same way that a hurdle specification and left-censored data are closely related issues, the fact that not all women have reached the end of their fertile period is closely related to the issue of right-censored data. The approach taken above implicitly takes this into account, for instance by conditioning the first moments on age.

4.3. Generalized method of moments estimator (GMM)

A standard GMM estimator is employed. For completeness, however, the estimation procedure is briefly described. The moment conditions can be written as follows:

$$E\{J(z_{th})\rho(y_{th}, w_{th}, z_{th}, \theta)\} = 0$$

where $J(\cdot)$ is the matrix with instruments. The sample analog of these moment condition is:

$$\sum_{h=1}^{H}\sum_{t=T_h}^{T^h} J(z_{th})\rho(y_{th}, w_{th}, z_{th}; \theta)/N = 0.$$

N is the total number of observations, T_h is the first observation of household h and T^h the last. Given panel data $\{y_{th}, w_{th}, z_{th}\}_{t=T_h,\dots,T^h}^{h=1,\dots,H}$, GMM estimates are obtained by solving:

$$\hat{\theta}_{GMM} = \arg\min_{\theta} \sum_{h=1}^{H} \sum_{t=T_h}^{T^h} \rho(y_{th}, w_{th}, z_{th}; \theta)' J(z_{th})' \left[\sum_{h=1}^{H} \sum_{t=T_h}^{T^h} J(z_{th}) J(z_{th})' \right]^{-1}$$

$$\times \sum_{h=1}^{H} \sum_{t=T_h}^{T^h} J(z_{th}) \rho(y_{th}, w_{th}, z_{th}; \theta).$$

The standard errors reported are asymptotically valid under heteroscedasticity and the goodness of fit measure (the R^2) reported is defined as the square of the correlation between the number of children observed and its estimated conditional expectation.

5. Empirical results

5.1. Empirical specification

In the empirical analysis two different models are estimated. The first model is a hurdle model where female employment status is not explicitly modeled (model I). In model I, the first step is whether or not children are present in the household and the distribution function for this is taken to be of a logit type. In the second stage, the conditional expected value of the number of children, once children are present in the household, is specified as $(1 + \exp(z_{th}'\beta))$. The second model is as specified in Sect. 4.2 (Eqs. (9), (10) and (11)) and takes the interrelationship between female employment status and the presence of children into account (model II). Although conceptually these two models seem to be nested, they are not nested from a statistical point of view. Also a standard count data model is not nested in the hurdle count data model as formulated in Sect. 4. The exogenous variables in both models are educational attainment of the both the man and the woman in the household, age and age squared of the woman and year of birth of the woman. Educational attainment variables are included as a proxy for the lifetime earnings of the both the man and woman. A priori one may expect higher educated women to have fewer children because of higher opportunity costs, compared to lower educated women. The educational attainment of the man may have a positive effect on the likelihood of having children and the number of children (an income effect). This would be in line with the results of Becker (1960). The year of birth is included to control for possible birth-cohort effects. One may argue that the empirical specification allows for little heterogeneity in household earnings and one should, for instance, include household income as an explanatory variable. This, however, is not possible since household income is the result of previous female employment and fertility outcomes. In other words, household income is bound to be endogenous and the only way to model this properly is to model the income process jointly with the female employment and fertility process. This is clearly beyond the scope of this paper, and for this reason time-constant regressors such as the educational attainment are included to proxy lifetime earnings. These variables are assumed to be exogenous.

Table 4. Estimation results, model I

	Presence of children	Number of children
Explanatory variables	p.e. (s.e.)**	p.e. (s.e.)**
Constant	−11.5 (1.98)*	−8.55 (0.79)*
Age/10	8.25 (1.05)*	4.54 (0.43)*
$(\text{Age}/10)^2$	−1.02 (0.17)*	−0.61 (0.06)*
Education level 2, woman	−0.55 (0.10)*	0.01 (0.03)
Education level 3, woman	−1.37 (0.12)*	−0.05 (0.04)
Education level 2, man	0.24 (0.12)*	−0.03 (0.03)
Education level 3, man	−0.26 (0.13)*	−0.01 (0.04)
(Year of birth)/10	−0.33 (0.15)*	0.07 (0.04)
R^2	0.47	
Conditional moment test	2.43*	

* Significant at the 5% level.
** p.e. = parameter estimate, s.e. = standard error

5.2. Estimation results

Estimation results for model I are reported in Table 4. Table 4 shows the by now well-known result that higher educated women are less likely to have children compared to lower educated women, after controlling for age. The relationship between educational attainment of the man and the presence of children appears to have an inverse U-shape. The effect of year of birth implies that the more recent the woman's birth year the less likely she is to have children for a given age. An interesting finding is that all of the explanatory variables, except for age, have no significant effect on the number of children. The conditional moment test statistic is significant, which may indicate some kind of misspecification.

The estimation results for model II are reported in Table 5. The conditional moment test statistic is insignificant which supports the hypothesis of no misspecification. In contrast to the results of model I, educational attainment of the man has no significant effect on the presence of children, and appears to have a negative effect on the number of children in a household where the woman is employed. Year of birth has a negative effect on the probability of having children. Although this effect is relatively small and only significant at the 10% level for employed women, this does suggest that compared to couples of earlier birth-cohorts, couples of later birth-cohorts are more likely to stay childless. To clarify the estimation results the conditional expectations, after controlling for educational attainment of the man and year of birth of the woman, are calculated and plotted in Figs. 5, 6, 7 and 8. Year of birth is set equal to 1960 and educational attainment of the man is set equal to educational attainment of the woman. Figure 5 shows the well-known finding in the labor supply literature that highly educated women are more likely to be employed at all ages, relatively to low educated women. Figure 6 shows that highly educated women schedule children later in life compared to low educated women. Figure 7 shows that highly educated women have fewer children compared to low educated women. These results are in line with the empirical findings in the fertility literature, as discussed in Sect. 2. In the search for an explanation for these findings, female employment status has

Table 5. Estimation results, model II

	Female employment and the presence of children		
	$W_{th} = 1,\ Y_{th} = 0$	$W_{th} = 0,\ Y_{th} > 0$	$W_{th} = 1,\ Y_{th} > 0$
Explanatory variables	p.e. (s.e.)**	p.e. (s.e.)**	p.e. (s.e.)**
Constant	0.51 (3.89)	−11.0 (3.85)*	−11.9 (4.22)*
Age/10	2.69 (1.87)	10.8 (1.86)*	9.73 (2.08)*
$(\text{Age}/10)^2$	−0.61 (0.30)*	−1.60 (0.30)*	−1.35 (0.33)*
Education level 2, woman	1.07 (0.19)*	0.22 (0.18)	0.75 (0.20)*
Education level 3, woman	1.11 (0.24)*	−0.86 (0.22)*	0.61 (0.24)*
Education level 2, man	0.20 (0.22)	−0.08 (0.21)	−0.06 (0.22)
Education level 3, man	0.24 (0.26)	−0.03 (0.24)	−0.14 (0.26)
(Year of birth)/10	−0.36 (0.34)	−0.64 (0.32)*	−0.64 (0.34)
	Number of Children		
	$W_{th} = 0$	$W_{th} = 1$	
Explanatory variables	p.e. (s.e.)**	p.e. (s.e.)**	
Constant	−7.83 (0.62)*	−14.2 (1.91)*	
Age/10	4.10 (0.34)*	7.64 (1.08)*	
$(\text{Age}/10)^2$	−0.55 (0.05)*	−1.03 (0.15)*	
Education level 2, woman	0.02 (0.02)	0.06 (0.04)	
Education level 3, woman	0.07 (0.03)*	−0.05 (0.06)	
Education level 2, man	−0.01 (0.02)	−0.12 (0.05)*	
Education level 3, man	0.04 (0.03)	−0.25 (0.06)*	
(Year of birth)/10	0.08 (0.03)*	0.03 (0.03)	
R^2	0.60		
Conditional moment test	1.94		

* Significant at the 5% level.
** p.e. = parameter estimate, s.e. = standard error

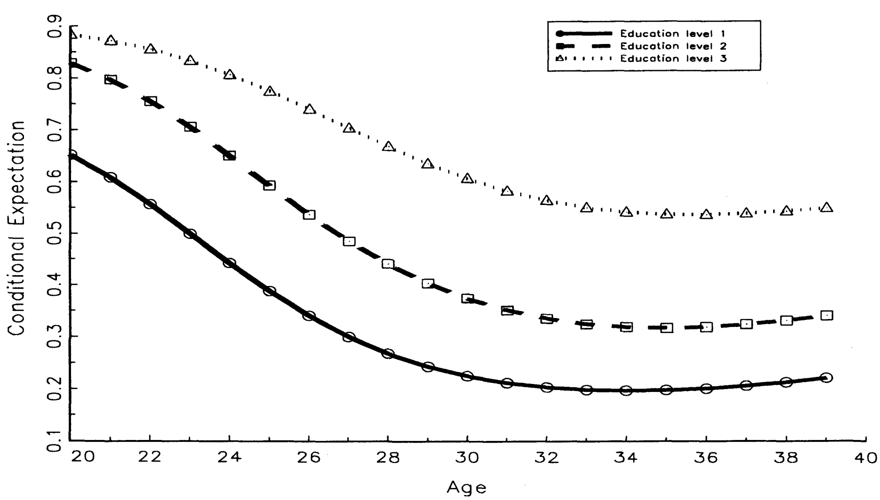

Fig. 5. The conditional expectation of female employment status, $E[E_{th}|z_{th};\theta]$, by age and education level

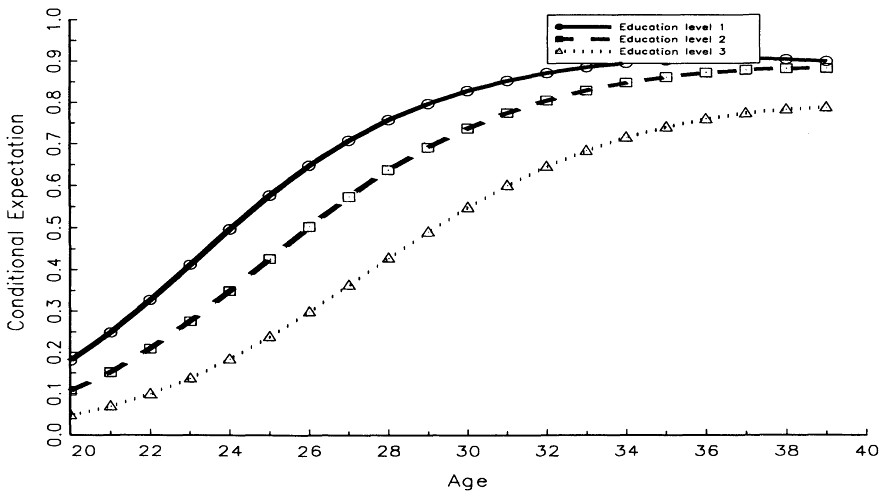

Fig. 6. The conditional expectation of the presence of children, $E[Y_{th} > 0|z_{th}; \theta]$, by age and education level

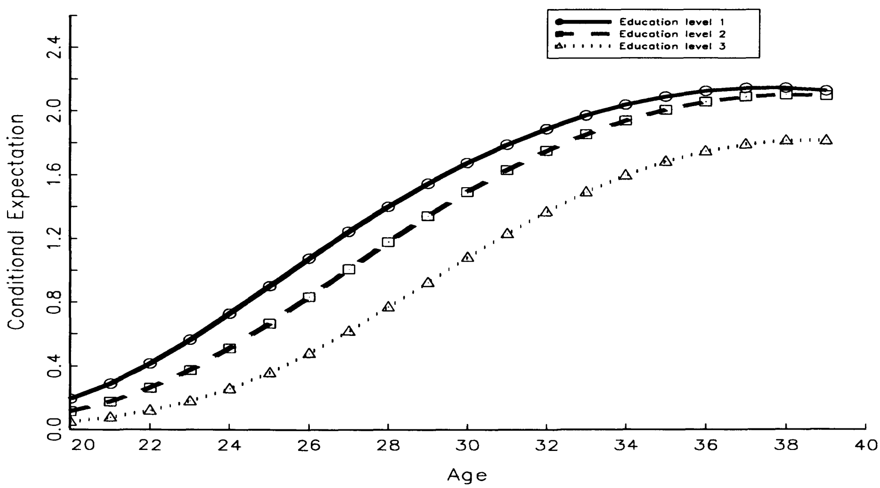

Fig. 7. The conditional expectation of the number of children, $E[Y_{th}|z_{th}; \theta]$, by age and education level

been modeled jointly with the presence of children. This makes it possible to analyze the effects of female employment status on the presence and number of children. Based on the estimation results the expected number of children conditional on female employment status and educational attainment can be calculated. These conditional expectations are plotted in Fig. 8 and show that a large proportion of the difference in the number of children between the different levels of education in Fig. 7 is explained by a difference in female employment status. Conditional on female employment status the difference

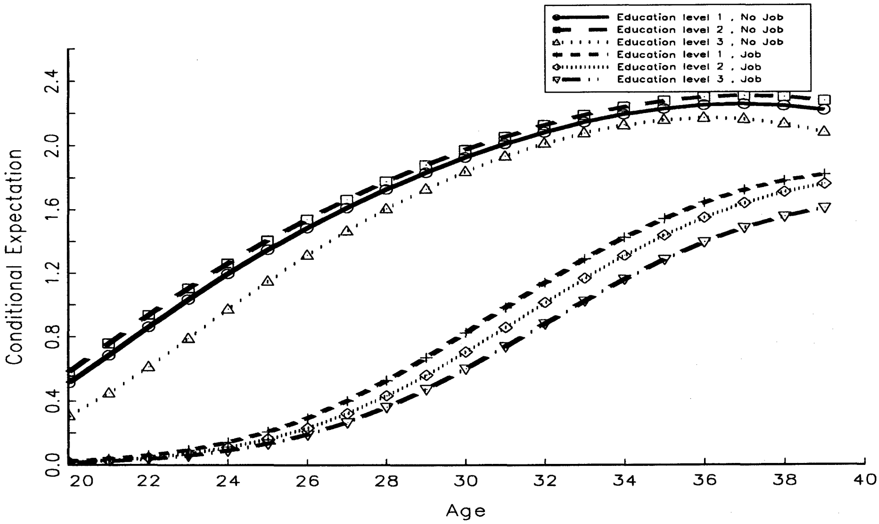

Fig. 8. The conditional expectation of the number of children given the employment status of the woman, $E[Y_{th}|w_{th}, z_{th}; \theta]$, by age and education level

in the number of children between the different levels of education is relatively small and insignificant, given the parameter estimates in Table 5.

6. Conclusions

This paper has analyzed the effects of female employment status on the presence and number of children in households in the Netherlands. For this purpose a hurdle count data model has been formulated and estimated by the generalized methods of moments. The hurdle takes into account the interrelationship between female employment status and the timing of first birth. Once children are present in the household, the number of children, has been modeled conditional on female employment status.

The main results can be summarized as follows. Relatively to lower educated women, highly educated women schedule children later in life, are less likely to have children and have fewer children. These results are in line with the empirical findings of earlier studies. The empirical results furthermore show that female employment status has a dominant effect on the presence and number of children: being employed significantly reduces both the likelihood of having children and the number of children. The direct effect of educational attainment on the presence and number of children is found to be relatively small and insignificant. In other words, the effects of educational attainment on the observed fertility pattern runs via the effects of educational attainment on female employment status, which in its turn significantly affects the fertility behavior of households. The observed delay in having children by employed women, shown in Fig. 8, is in line with the empirical findings of Bloemen and Kalwij (1996) who stress the importance of state-dependence and show that female employment status dominates the effects of educational attainment on the timing of first birth. Here, in addition, it is shown that female employment status also has a dominant effect on the number of children at all ages. These empirical results emphasize the importance of explicitly modeling female employment status when analyzing and trying to understand the observed fertility behavior of households, hence life-cycle female employment behavior.

The econometric model as outlined in this paper only partly solves the endogeneity problems of simultaneously analyzing the outcomes of female employment and fertility decisions. Intuitively this model may be considered appropriate in this particular empirical analysis. However, tackling the problems surrounding simultaneity issues in (hurdle) count data models is considered to be a necessary and fruitful route to follow for future research. Furthermore, the model employed here is a reduced form model used to gain a better understanding of observed life-cycle female employment and fertility behavior. To address more fundamental issues related to, for instance, policy schemes to increase the female employment rate after the birth of the first child, a more structural approach needs to be taken. First steps in this direction have been taken in Francesconi (1996) and Kalwij (1999, Chapt. 4).

Endnotes

[1] A perhaps more appealing approach is to formulate and estimate a structural model based on a life-cycle model of household behavior. Francesconi (1996) and Kalwij (1999) take such an approach. Because of the computational difficulties involved in estimating such a model it is fair to say that these models are not very practical. Estimating such a model is clearly beyond the scope of this paper.

[2] The maternity period is at most up to 3 months after the birth of the child. In 1991 a papental leave scheme was introduced. This makes it possible for a further 6 months leave. The maternity and parental leave schemes are a bit more complex than sketched here, but they roughly amount to a maximum period of leave, and registered as being employed, of one year. After this leave the woman has to return to work.

[3] Alternatively one may favor a truncated Poisson distribution for the underlying DGP and specify the conditional expectation as: $\exp(z'_{th}\beta)/(1-\exp(-\exp(z'_{th}\beta)))$. The main empirical findings of this paper remain unchanged when this alternative specification is used.

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Job bust, baby bust?: Evidence from Spain

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Abstract. The unemployment rate in Spain has been exceptionally high for more than two decades by now. During the same period the fertility rate dropped dramatically reaching the lowest level in the world. In this study we look for evidence of a link between the 'unemployment crisis' and the 'fertility crisis' in Spain. We examine the factors that affect individuals' ages at marriage and childbirth, focusing on the effects of *male* employment status. Our results show that spells of non-employment have a strong negative effect on the hazard of marriage. We also find negative (but smaller) effects of part-time or temporal employment on the hazard of marriage. The estimated direct effects of joblessness and part-time work on birth hazards conditional on marriage are smaller and/or not significant for most birth intervals and sample groups. Simulations based on the estimated models confirm the potential for large 'delaying' effects of joblessness on marriage. However, the delaying effect is not so large in simulations which control for the actual incidence of non-employment in the sample.

JEL classification: 13

Key words: Unemployment, age at marriage, birth intervals

1. Introduction

During the last several decades, one of the most prominent sociodemographic trends taking place across the developed world has been the decline of fer-

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tility. At the same time there was a delay in the age at marriage and a decrease in marriage rates. Traditionally, studies of family formation have focused on females, looking at female education, female wages and female labor market participation as the main determinants of marriage and fertility. In these studies men played at most a secondary role as an exogenous factor. However, historical as well as more recent observations suggest the importance of male employment in explaining fluctuations in fertility and marriage.[1] In this paper we explore the relationship between fertility, marriage and *male* employment status using micro data from Spain. We look for evidence of a link between high and persistent unemployment rates and the dramatic fertility decline experienced by some European countries such as Spain and Italy.

Figures 1a–1d illustrate the trends in aggregate fertility and selected labor market variables in Spain over the last 25 years. (Also, see Bover and Arellano 1995 and Ahn 1997.) Between the mid-1970's and the mid-1990's the total fertility rate dropped from a level around 3 which was among the highest in Western Europe to 1.15 children per woman which is lowest in the world. At the same time the labor force participation rate increased gradually from under 30% to more than 45% for all working age females, and much more rapidly among younger women (e.g., from 30 to 70% among those aged 25–34).

However, the most striking development in the Spanish labor market was the evolution of the unemployment rate which increased from a level below 5% through the mid-1970s to around 20% since the mid-1980s (Fig. 1c). This change is specially relevant for marriage and fertility because the burden of unemployment has fallen disproportionately on young workers. Among those aged 16–29, the unemployment rate has been around 40% during the last decade.[2] One may point out that unemployment declined during the second half of 1980s while fertility continued its downward trend. However, even during this expansion the unemployment rate was never below 16%. Therefore, over time it was becoming clear that high unemployment was likely to persist for a long time. These changes in expectations may have had an impact on family formation decisions among young people.

Another factor that may account for the continued decline in fertility in Spain in spite of the decreasing or stable unemployment rate was the rapid increase in the proportion of workers holding a temporary contract following a change in the labor market regulations in 1984. During the late 1980s and early 1990s most job openings were under temporary contracts, which has greatly increased the proportion of temporary contract holders among young workers. The proportion of males aged 25–39 with a permanent work contract declined from 55% during the mid-1980s to less than 40% during the 1990s (Fig. 1d).[3] High youth unemployment together with a rising proportion of temporary contract holders have brought enormous uncertainty regarding future careers and income as well as lower current income for many individuals and households. Our conjecture is that the two have combined to inhibit marriage and childbearing, both of which involve long-term commitments.[4]

A popular theory of marriage by Becker (1974) suggests that there are gains to marriage due to specialization in the production of household goods and the joint production of children and other marriage-specific capital. More recently, Becker's insights have been embedded in dynamic search-theoretic models of the 'marriage market' (see the survey by Montgomery and Trussell, 1986). The timing of marriage will be influenced by the costs of finding a suitable mate.

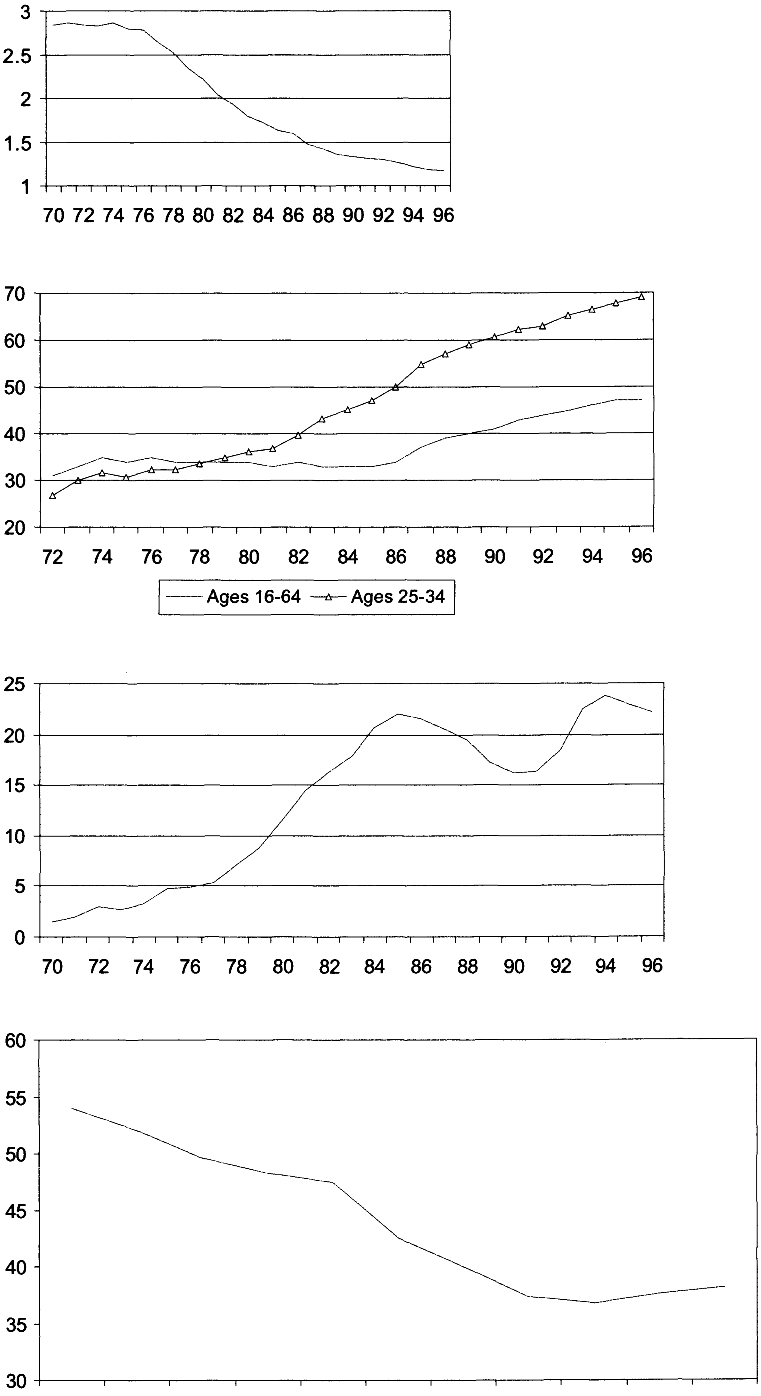

Fig. 1a–d. **a** Fertility rate in Spain; **b** Female participation rates in Spain; **c** Unemployment rate in Spain; **d** Proportion among males 25–39 with permanent contracts

Since a marriage occurs when the decisions of two individuals agree, attractiveness of one's own characteristics is important as well as those of potential spouses. Therefore, an individuals' labor market situation is likely to be one of the determinants of the availability and the quality of potential spouses, and in turn, of the age at marriage. Furthermore, marriage usually entails fixed costs in terms of housing and basic household equipment. Therefore the timing of marriage is likely to be affected by one's savings and past employment history as well as by the current situation.

Most empirical studies of the age at marriage use hazard models to estimate reduced form equations. Anderson et al. (1987), using data from Malaysia, found that skilled employment of both husbands and wives delays marriage relative to unskilled employment or non-employment. Keeley (1977) found, using US data, that high wage males marry earlier than low wage males, which is what a theory based on specialization in the household would predict. However, Bergstrom and Schoeni (1996) found, also using US data, first a positive association between male income and age at first marriage under age 30, then a negative association for those who married after age 30. A negative effect of difficulties in men's career transitions on marriage probability has been shown in recent US data (Oppenheimer et al. 1997).

The growth of the 'New Home Economics' has also lead to the development of dynamic, sequential theories of life-cycle fertility (see for a survey Hotz et al. 1997; Arroyo and Zhang 1997). However, existing dynamic fertility models tend to abstract from the marriage decision and even so their empirical implementation is difficult.[5] Because of this, most empirical studies based on dynamic fertility models have employed a strategy of reduced form estimation (See Hotz and Miller 1988 and Heckman and Walker 1990 for state-of-the-art examples.). Most studies focus on the effects of female and male wages on fertility. In a wide range of models a negative effect of female wage and a positive effect of male wage are predicted and empirical results generally have conformed to the prediction. But there are few studies which examine the impact of male employment status on childbearing.

In this paper we estimate discrete time proportional hazard models in order to learn about the relationship between men's labor market experience and their family formation behavior in Spain. In theory, the effect of male unemployment on fertility should be similar to that of a drop in current period household earnings, with an additional impact through expected lifetime income if unemployment is expected to last. As long as children are a normal good, this drop in current and expected household income should decrease the hazard of childbearing. Additional negative effects might arise if a housewife decides to enter the labor market or if a working wife delays her exit from the labor market in order to maintain household income.

We use individual data from the Encuesta Sociodemografica (Sociodemographic Survey) of 1991 which contains information on current and past economic and family situations of the members of Spanish households. First, we examine the determinants of the age at marriage. Second, we examine duration of the first three birth intervals conditional on marriage. Our emphasis is on the effects of individuals' labor market situation but we also consider the impact of family characteristics (parents' education, father's line of employment) and other relevant factors (cohort, region, and characteristics of previous children in the analyses of inter-birth intervals). Our results suggest that unemployment was indeed a factor contributing to the delay of marriages in

Spain during the last two decades, but that its direct effect on birth hazards conditional on marriage was not large.

The rest of the paper is organized as follows. In Sect. 2 we review the proportional hazards duration models used in our analysis and we discuss the selection of the sample and the choice of covariates. In Sect. 3 we present the estimation results and we carry out simulations to illustrate the impact of employment status on the timing of marriages. We conclude in Sect. 4.

2. Econometric specification and data

2.1. Model specification

In this paper we analyze decisions on the timing of marriage and childbearing within the framework of duration models. Because our data only provides yearly information on all the events of interest, we specify a continuous time hazard model which is integrated to obtain a likelihood function for grouped (discrete time) data as in Meyer (1990). The instantaneous hazard function at time t is assumed to take a proportional hazard form:

$$\lambda_i(t) = \lambda_0(i)\varepsilon_i \exp(X_i(t)\beta) = \lambda_0(t) \exp[X(t)\beta_{it} + \log(\varepsilon_i)] \tag{1}$$

where $\lambda_0(t)$ is an unknown baseline hazard at t, $X_i(t)$ is a vector of possibly time-varying explanatory variables, β is the corresponding parameter vector and ε is a random variable which describes unobserved heterogeneity across spells. It is assumed that $X_i(t)$ be a step-function, i.e., it is constant within periods (years) but it may vary between periods. Conditional on ε, the discrete interval hazard is then

$$h_j(X_{ij}) = 1 - \exp\{-\exp[X_{ij}'\beta + \gamma_j + \log(\varepsilon_i)]\} \tag{2}$$

where $\gamma_j \equiv \log \int_{a_{j-1}}^{a_j} \lambda_0(\tau)\, d\tau$. If ε is gamma-distributed with unit mean and variance v, the (unconditional) survivor function and the likelihood take convenient closed forms. The log-likelihood is

$$\log L = \sum_{i=1}^{n} \log\{(1 - c_i)A_i + c_i B_i\} \tag{3}$$

where

$$A_i = \left[1 + \nu \sum_{j=1}^{t_i} \exp[X_{ij}'\beta + \gamma_j]\right]^{-(1/\nu)}$$

$$B_i = \begin{cases} \left[1 + \nu \sum_{j=1}^{t_i - 1} \exp[X_{ij}'\beta + \gamma_j]\right]^{-(1/\nu)} - A_i, & \textit{if } t_i > 1 \\ 1 - A_i, & \textit{if } t_i = 1 \end{cases}$$

where t_i is the observed duration of spell i, $c_i = 1$ indicates a complete spell and, $c_i = 0$ a censored spell. As the variance of the gamma distribution goes to zero, the likelihood converges to that of proportional hazards model with no unobserved heterogeneity. The parameters of interest are β, v and the sequence describing the non-parametric baseline hazard.[6]

For the analysis of the age at marriage, we construct person-year data for each year since the completion of schooling until the time of the event occurrence (completed duration) or until the survey time (censored duration). For the analyses of inter-birth durations we construct similar person-year data starting at the time at the birth interval of interest. Model (3) is estimated separately for each of the events (marriage, first birth, second birth, third birth).[7] Heckman and Walker (1990) warned that this "piece-meal" approach may yield biased estimates in the presence of individual-specific unobservables correlated over spells. However, in their empirical study using Swedish data they found that this form of unobservable heterogeneity was empirically unimportant. It is worth noting that in Spain fertility outside marriage during the sample period was negligible, and cohabitation was quite rare.[8] Furthermore, very few people marry before completing school.[9]

2.2. Sample selection

The data are drawn from the Spanish Socio-demographic Survey (Encuesta Sociodemográfica) carried out by the Spanish Statistical Institute (INE) during the third quarter of 1991. The principal objective of the Survey was to gather information about individuals' history of family situation, residence and housing, economic activities and occupation, and education. The Survey contains information on 159,154 principal interviewees (a representative sample of the Spanish population of ages 10 and over) and their households.

We limit our analysis to *prime-aged males*. We do this because our survey (as most other surveys) records labor market histories only for the principal respondents (one person for each household surveyed). Given our interest in the effect of male employment status on marriage and fertility, this limitation is overcome by taking the male respondents as our working sample. An important advantage of using male samples and focusing on the effect of male employment status is that, unlike female employment status, male employment status can be treated more safely as exogenous with respect to the decisions of marriage and childbearing. For example, recent work by Angrist and Evans (1998) shows strong evidence of endogeneity of female labor supply and for exogeneity of male labor supply in childbearing decisions among couples in the United States.

In most societies the ages at marriage of male and female spouses are highly correlated. The correlation coefficient for Spain in our data is 0.97 and highly significant. Given that most childbearing occurs within stable unions (births from non-stable unions accounted for less than 5% in our data), the father's and mother's ages at birth are also highly correlated. Therefore we think the results of the analysis of the age at marriage and childbearing for males can be interpreted as applying to both sexes once we adjust for the age gaps between husbands and wives. Appendix 1 gives some descriptive proportions of ever married and the number of children by sex, age and cohort.

Our working sample consists of all principal male respondents aged 26–40 at the time of survey. The reason why we do not include people younger than 26 is that the majority (60%) of Spanish males are still unmarried by this age. By excluding those over 40 we reduce recall error about life histories arising from the retrospective nature of the Survey. Since most marriages and births occur after age 20, almost all the decisions recorded in our sample correspond to the 1970's and 80's, including the period of rapid fertility decline which started in the mid-1970s.

2.3. Choice of covariates

Using retrospective information about individuals' work histories, we construct individuals' yearly employment status. Unfortunately, within non-employment periods we cannot further distinguish between unemployment and out of labor force states. However, considering that our working samples contain only prime-aged males, we think it is reasonable to interpret non-employment periods as periods of unemployment.

One of the main factors examined in numerous studies of age at marriage and childbearing is completed education (Montgomery and Trussell 1986; Schultz 1997; Hotz et al. 1997). Most of these studies focus on women, and they generally find a strong delaying effect of education on marriage and fertility. However, a conceptual problem is that education is very likely to be an endogenous variable in marriage and childbearing decisions. In order to deal with this problem instrumental variable or simultaneous equation estimation techniques have been applied. Nevertheless the problem persists in most cases due to the difficulties in finding adequate instruments or due to tenuous identification. In any case, education is not one of our main variables of interest. Our goal is rather to establish the effects of *male* employment status on the hazards of family formation events from the age of school completion, conditional on education. However, we would like to allow for interactions between education and some of the other covariates. In particular, since men do not marry before they leave school and the ages at school completion are so different by education, it would seem too restrictive to impose the same pattern of duration dependence for all levels of education. It is for this reason that we estimate the models of marriage and the first birth separately for each educational category.[10] We break our sample of men into three groups with primary education, secondary education, and college education. For the sake of homogeneity, we further restricted the sample to include only individuals who completed schooling by a given age: age 14 for the primary education group (60% of this group), ages between 17 and 19 for the secondary education group (59% of this group), and ages 20 through 25 for the college education group (62% of this group).[11]

The duration variables are the waiting time (in years) until marriage and the duration in years of the first, second and third birth intervals conditional on marriage. The covariates included are employment status, age at the survey date (as a control of cohort and trend effects), and regional dummies. In the analyses of birth intervals we also include the wife's education and variables related to previous births, such as survival status of previous children, duration of previous birth intervals, and the gender of existing children. In the model of duration to marriage we also included the duration of the (unemployment)

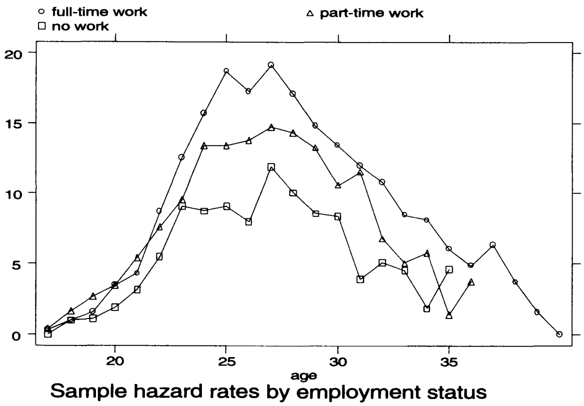

Fig. 2. Kaplan-Meier marriage hazards (%) by employment status

spell in the transition from school to work, father's and mother's education and father's labor market situation when the respondents were 16.[12]

Before we turn to the estimation results, a more descriptive analysis of the relationship between family formation and employment status will be useful. We computed Kaplan-Meier hazard rates for each event by employment status. As shown in Fig. 2, the conditional probability of marrying is clearly higher for those employed than for those without a job, and higher for the full-time workers than for the part-time or temporal workers. The differences are more pronounced after age 20 and persist well beyond age 30. The Kaplan-Meier hazard rates for birth intervals are not clearly distinguishable across different employment states. In particular, the number of individuals who do not work is substantially reduced making the comparison less precise. This suggests that few non-employed males progress to marriage leaving few non-employed males in the analyses of birth intervals.

3. Empirical results

Estimates of the parameters of the models of duration to marriage and durations of the first, second and third birth intervals are presented in Tables 1–3. In order to assess the importance of unobserved heterogeneity we look at the following statistics: a) The ratio of the estimated variance of the gamma mixture distribution and its estimated standard error; and b) Likelihood ratio tests of a model with no unobserved heterogeneity against the gamma mixture model. Both sets of statistics suggest that unobserved heterogeneity is empirically significant only for the models of the interval to the first birth.[13] Because of this the parameter estimates in Tables 1 and 3 are those of the models with no unobserved heterogeneity.

3.1. Employment status

Every year, an individual's employment status is described by a categorical variable with five possible values: full-time continuous work, part-time or tem-

Table 1. Relative risk ratios, duration to marriage

	Males with primary education (N = 56791)		Males with secondary education (N = 25133)		Males with college education (N = 13388)	
	Risk ratio (*t* statistics)	Sample mean	Risk ratio (*t*-statistics)	Sample mean	Risk ratio (*t* statistics)	Sample mean
Labor market situation one year ago (re: full-time work)						
Part-time	0.84 (3.98)	0.16	0.80 (3.03)	0.10	0.98 (0.28)	0.09
No work	0.57 (6.12)	0.07	0.42 (7.36)	0.21	0.46 (8.32)	0.28
Military	0.82 (3.89)	0.17	0.85 (2.16)	0.19	0.80 (2.77)	0.17
Search duration to first job (re: 0–6 months)						
7+	0.75 (4.84)	0.10	0.83 (2.84)	0.18	0.81 (3.16)	0.28
Cohort (re.: 1951–55)						
1956–60	1.09 (2.45)	0.39	1.01 (0.14)	0.39	0.99 (0.06)	0.42
1961–65	1.07 (1.70)	0.30	0.87 (2.26)	0.37	0.75 (3.59)	0.25
Father's education level (re.: no formal education)						
<Prim.	1.13 (2.18)	0.38	0.90 (1.00)	0.35	0.94 (0.35)	0.23
Prim. low	1.04 (0.52)	0.20	0.97 (0.31)	0.29	0.81 (1.28)	0.26
Prim. high	1.10 (0.78)	0.03	1.03 (0.24)	0.08	0.82 (1.16)	0.11
Second.+	1.06 (0.38)	0.01	0.90 (0.89)	0.10	0.97 (0.22)	0.30
Mother's education level (re.: no formal education)						
<Prim.	0.87 (2.41)	0.37	1.08 (0.81)	0.38	1.03 (0.20)	0.29
Prim. low	0.92 (0.98)	0.19	0.87 (1.33)	0.30	1.04 (0.30)	0.34
Prim. high	0.63 (2.97)	0.02	0.72 (2.37)	0.08	0.91 (0.59)	0.12
Second.+	0.88 (0.45)	0.004	0.92 (0.49)	0.03	0.94 (0.37)	0.13
Father's work status (re.: employee)						
Employer	1.14 (1.29)	0.02	1.10 (1.06)	0.07	1.21 (2.38)	0.12
Self-emp.	0.79 (6.64)	0.31	0.90 (1.94)	0.25	0.98 (0.23)	0.22

Note: Risk ratio coefficients are ratios of hazards to the baseline hazard. Baseline hazard rates are estimated non-parametrically by including a dummy variable for each year of age. We have also included regional dummy variables which are not presented in the table.

Risk ratios in the table are $\exp(\beta)$, where β is the estimated coefficient. The t-statistics in parenthesis are the unsigned ratio of the estimate of β and its estimated standard error. Therefore, a large t-statistic can be interpreted as evidence against the null that the risk ratio is 1.

poral work, no work and military duty. This variable refers to the situation the previous year to allow for the gap between decision to have a child and the occurrence of a birth. It is time-varying covariate. In general, the estimation results confirm what we learnt from the Kaplan-Meier hazard rates shown in Fig. 2.

Age at marriage. For all education samples, individuals who work full-time (the omitted or 'reference' category in estimation) are substantially more likely to marry than those who do not. Part-time or temporal work reduces the conditional probability of marriage in a given year by about 20% relative to full-time work, except for the college educated. However, the largest reduction occurs during no-work periods. Those without work are less than half as likely to become married as those with a full-time work. This result suggests that the lack of stable employment among the young has contributed to the substantial delay in the age at marriage and the increased incidence of singlehood in Spain during the last two decades. As expected, the marriage hazard is substantially lower while one is doing military service.

Table 2. Relative risk ratio from marriage to first birth

	Males with primary education (N = 8528)		Males with secondary education (N = 4790)		Males with college education (N = 3383)	
	Risk ratio (*t* statistics)	Sample mean	Risk ratio (*t*-statistics)	Sample mean	Risk ratio (*t* statistics)	Sample mean
Labor market situation one year ago (re: full-time work)						
Part-time	1.01 (0.11)	0.13	1.26 (1.38)	0.09	0.71 (2.14)	0.11
No work	1.06 (0.25)	0.02	1.13 (0.39)	0.01	0.79 (0.89)	0.02
Wife's education level (re: primary)						
<Prim.	1.18 (1.39)	0.15	1.26 (1.03)	0.05	1.46 (0.91)	0.01
Secondary	0.72 (3.50)	0.51	0.71 (2.54)	0.62	0.93 (0.36)	0.40
University	0.34 (4.34)	0.04	0.56 (3.16)	0.15	0.74 (1.42)	0.53
Cohort (re.: 1951–55)						
1956–60	0.83 (1.88)	0.43	0.62 (3.56)	0.47	0.68 (2.84)	0.49
1961–65	0.61 (4.25)	0.26	0.49 (4.42)	0.26	0.53 (3.65)	0.14
Variance of unobserved heterog.	1.51 (3.77)		1.36 (3.05)		0.78 (2.10)	

Note: Same as in Table 1.

Table 3. Relative risk ratios, second and third birth intervals

	Second birth interval (N = 29349)		Third birth interval (N = 17833)	
	Risk ratio (*t* statistics)	Sample mean	Risk ratio (*t* statistics)	Sample mean
Labor market situation one year ago (re.: full-time work)				
Part-time	1.03 (0.48)	0.13	1.09 (0.72)	0.11
No work	0.58 (3.21)	0.02	0.61 (1.10)	0.01
Own education level (re.: primary)				
Secondary	0.98 (0.57)	0.21	0.82 (1.96)	0.22
University	1.14 (2.28)	0.13	1.00 (0.03)	0.14
Wife's education level (re.: primary)				
<Primary	1.05 (0.84)	0.16	1.19 (1.56)	0.16
Secondary	0.87 (3.19)	0.47	0.85 (1.70)	0.44
University	1.04 (0.56)	0.10	0.93 (0.48)	0.10
Cohort (re.: 1951–55)				
1956–60	0.86 (4.13)	0.44	1.00 (0.01)	0.37
1961–65	0.69 (6.54)	0.16	1.09 (0.65)	0.09
Previous children				
Birth interval	0.91 (8.14)	1.89	0.82 (8.44)	4.43
Dead	4.67 (6.24)	0.001	4.22 (6.12)	0.007
Same sex	–	–	1.31 (3.52)	0.51

Note: Same as in Table 1 except that some characteristics of the previous children, survival status, previous birth interval and sex, are included.

Birth intervals. As discussed earlier, the samples are very different from the sample for age at marriage. In these samples the number of periods without work or in military duty is much smaller, less than two percent compared to more than 10% in the marriage sample. This will make it more difficult to obtain precise estimates of the effect of not having a job. It also suggests indirectly that employment status is indeed important for marriage. Many of those without work do not marry, and therefore they are not eligible for the birth interval samples. Regardless of sample selection, we would still expect negative effects of part-time jobs or no work on the hazards of births. Our results conform to that prior but are rather weak. The hazard of a first birth conditional on marriage is 39% smaller for individuals with College education working only part-time, but this is the only significant effect for that interval. The second birth hazard falls by more than 40% during periods of joblessness. The third birth hazard is also reduced by almost 40% for those who do not work, but this effect is not statistically significant. No significant effects of working part-time were found for the second and third birth intervals.

Initial unemployment duration. In the model of duration to marriage, we also included the duration of the unemployment spell that individuals experienced prior to their first job usually during the transition from school to work. This variable is not time-varying and the coefficient should not be interpreted as a direct contemporaneous effect of unemployment. Having included time-varying employment status, we interpret this variable as representing some permanent unobserved individual differences, such as abilities in the labor market and preferences for work, which influence their family formation behavior. The results are as expected. Initial unemployment duration has a negative effect on marriage hazards: the longer the transition from school to work the later one marries. Somewhat surprising is its large magnitude. Those whose initial unemployment spell was longer than 6 months are 25% less likely to get married each year than those with shorter unemployment duration prior to their first job.

3.2. Other covariates: Duration to marriage equation

Parental education and employment status. As shown in many previous studies (see Ahn and Ugidos 1996 for the Spanish case), parents' education influences children's labor market and demographic behavior mostly through children's educational achievement. Given that our samples are homogenized with respect to the education level and the age at completion of schooling, parents' education level is likely to affect children's age at marriage through family income and other relevant factors that are correlated with parents' educational level. One plausible hypothesis would be that higher education of parents makes their children more attractive in the marriage market therefore leading to earlier marriage other things equal. Given the parents' education level, father's occupation may be a proxy for the economic situation of the family. We distinguish three employment categories: employer, paid worker (reference category in our estimation models) and self-employed without employees. According to previous studies, the average income is highest among employers and lowest among the self-employed.

Our results are not clear-cut. In general, the coefficients are not significant for father's education in spite of large sample sizes. However, in several cases we observe a negative and significant effect of mother's education on their children's marriage hazard. This might reflect the existence of reverse selection. That is, given that all individuals have the same completed education level in our samples, the mother's education might be negatively correlated with the child's unobserved ability. Children with more educated mothers are supposed to have higher education, either for genetic reasons or because they receive more human capital at home.

With respect to father's employment status our results indicate that the children of 'employer' fathers have higher marriage probabilities than those of 'employee' fathers, while the marriage hazard of individuals with self-employed fathers is the lowest. This seems to be in agreement with our 'family income' conjecture – that the higher the family income the higher the marriage probability; however, the effects are not significant for all education categories.

Birth cohort. There are some differences between cohorts. Among the secondary and college education samples the marriage hazard was significantly lower for the youngest cohort (1961–1965). This decrease was greater for the college graduates, which may reflect the rapid increase in college enrollment rate among their potential spouses. Among the primary education sample, the youngest cohort (61–65) had lower marriage hazards than the intermediate cohort (56–60), but marginally higher than the oldest (51–55).

3.3. Other covariates: Birth interval equations

Birth cohort. The more recent cohorts had lower birth hazards. A strong and significant downward trend is observed for the hazards in the first and second birth intervals across all three education groups. The effect was strongest among the secondary education group; for instance, the hazard rates in the first birth interval decreased by as much as 50%. No significant trend is detected for birth hazards in the third birth interval.[14]

Wife's education. The birth hazards decrease with the wife's education. These coefficients are generally large and significant, especially so for the first birth interval.

Previous children, interval to second birth. We have included some variables regarding the first child. First, the variable indicating the death of the first child is included as a time-varying dummy equal to one for all periods after the death of the child. Since infant and child mortality rates are low, there is a very small number of person-years for which this indicator is one (less than 0.3 percent of the total). It is well known that subsequent births tend to occur sooner due to both biological and behavioral reasons when a child dies. This same result is observed in our estimates. The death of the first child increases the hazard of a second birth by a factor of more than four.

We included the duration of the interval from marriage to the first birth. This variable is likely to work as a proxy for the couples' fecundity and preferences for children, which we do not observe. Under this hypothesis, shorter previous intervals would lead to shorter subsequent intervals, which is confirmed in our results. Every year added to the previous interval reduces the second birth hazard by about 10%. However, this should not be interpreted as a causal effect since both are likely to be determined by common unobserved factors.

Previous children, Interval to third birth. Again, the death of the first or second child is very rare (less than 1 percent of person-year observations) but we observe a similar effect as in the progression to a second birth. The death of the first or second child increases the third birth hazard by a factor of four.

We also included the interval from the first to the second birth. As for the second birth interval, the longer the previous interval the lower the hazard in the current interval. The third birth hazard falls by about 20% for every year that is added to the second interval.

To control for gender preferences we introduced a dummy indicating whether the first two children were of the same gender. The results show that there are indeed preferences for a balanced gender composition. Parents with two boys or two girls have a 30% higher birth hazard in the third birth interval.[15]

3.4. *Goodness of fit*

Before we use the estimated model to perform counterfactual simulations in the next subsection, we need to check that the model can at least deliver reasonably good predictions of behavior within sample. We compute the hazard of marriage predicted by the model for each observation in the sample, and we average the predicted hazards across observations at each duration. We then compare the duration profiles of actual and predicted hazards in order to provide an informal assessment of fit. At most durations, actual and predicted hazards coincide up to the third decimal place, making them virtually indistinguishable.

3.5. *Simulations*

In order to measure the impact of non-employment spells on the process of family formation, we carry out the following simulations. In the first one, for each individual in the sample we consider two employment histories: a) his actual employment history up to the last period he is observed. b) a hypothetical history consisting of continuous full employment. For each of the two histories, we compute the sequence of predicted marriage hazards and we 'integrate' these to obtain the sequence of probabilities of marriage at each duration. That is, if $h(t)$ is the marriage hazard the probability of marrying at age a is $[1-h(0)][1-h(1)]\dots[1-h(a-1)][h(a)]$. We then sum this sequence to get the sequence of probabilities of marriage by each age, i.e., the complement of the survivor function. Finally, we compute the average of these sequences over individuals. Table 4a summarizes the results by education and cohort.

For instance, the actual incidence of joblessness or part-time work reduced the probability of marriage by age 30 from 72% to 67% for individuals with university education. Although the estimated effect of non-employment on the marriage hazard is quite large, the differences in Table 4a are not very large because the differences between the average 'actual' employment history and the hypothetical 'continuous employment' history was not important enough. However, notice that the simulated impact of non-employment is larger for individuals of the most recent cohort, for which lack of continuous employment was a more common event.

The simulations in Table 4b are meant to illustrate how the estimated coefficients could result in large reductions in the probability of marriage for individuals who suffer long spells of joblessness. We compute the predicted

Table 4. The impact of employment status on the probability of marriage. **a** Simulated impact of sample non-employment on the probability of marriage by education and cohort

	Employment history	By Age 25			By Age 30		By Age 35
	Cohort	61–65	56–60	51–55	56–60	51–55	51–55
Primary educ.	Continuous full-time	0.55	0.55	0.50	0.82	0.77	0.84
	Actual	0.51	0.52	0.48	0.79	0.76	0.83
Secondary educ.	Continuous full-time	0.40	0.46	0.46	0.79	0.78	0.87
	Actual	0.37	0.43	0.44	0.76	0.76	0.86
University educ.	Continuous full-time	0.18	0.28	0.28	0.72	0.72	0.86
	Actual	0.15	0.24	0.25	0.67	0.68	0.83

Table 4b. Simulated impact of the duration of post-school joblessness on the probability of marriage. Evaluated at sample mean characteristics

Years of non-work	Survival probability from marriage until 5th year			Survival probability from marriage until 10th year		
	Primary	Secondary	University	Primary	Secondary	University
0	0.888	0.834	0.522	0.393	0.345	0.198
1	0.915	0.862	0.614	0.495	0.413	0.279
2	0.919	0.869	0.652	0.497	0.416	0.296
3	0.925	0.880	0.703	0.500	0.421	0.319
4	0.936	0.903	0.789	0.506	0.432	0.358
5	0.950	0.938	0.892	0.514	0.449	0.405
6	–	–	–	0.528	0.476	0.462
7	–	–	–	0.554	0.514	0.534
8	–	–	–	0.587	0.560	0.606
9	–	–	–	0.628	0.621	0.677
10	–	–	–	0.668	0.687	0.753

probability that an individual with sample mean characteristics will remain unmarried 5 and 10 years after leaving school as a function of the number of years spent without a job. For instance, the probability that the average individual with university education will remain unmarried 10 years after leaving school is only 20% if he was continuously employed during the interval, but it jumps to 75% if he was continuously unemployed.

4. Discussion and conclusions

In this study we have looked for evidence of a link between the 'unemployment crisis' and the 'fertility crisis' in Spain. We have examined the factors that affect individuals' ages at marriage and childbirth, focusing on the effects of male employment status. Our results suggest that spells of non-employment have a strong negative effect on the hazard of marriage. We also find negative (but smaller) effects of part-time or temporal employment on the hazard of marriage. Simulations based on the estimated models confirm the potential for large 'delaying' effects of joblessness on marriage. However, the delaying effect is not so large in simulations which control for the actual incidence of non-employment in the sample.

We did not perform similar simulations for birth intervals because the incidence of non-employment in the sample is much smaller for married individuals. Furthermore, the estimated direct effects of joblessness and part-time work on birth hazards conditional on marriage are smaller and/or not significant for most birth intervals and sample groups. It should be noted that non-employment has an indirect effect on births through the delay of marriages. If men who delay their marriage compensate with higher birth hazards once they are married, this indirect effect might vanish. However, we found some evidence against this canceling out. First, age-at-marriage variables were not significant in our preliminary birth interval equations. Second, we estimated a model of duration to first birth unconditional on marriage and the estimates of coefficients on employment status variables were almost the same as in the duration to marriage model.

Our sample covered the relatively low-unemployment 1970's and the deep recession of the early and mid-1980's. An extended sample covering the recession of the 90's, when unemployment was extremely high again, might provide more precise estimates of the effect of employment status on birth intervals. Presumably, the simulated impact of non-employment on the delay of marriages would also be larger in an extended sample. In spite of this, we have to conclude that joblessness does not seem to be the major factor behind the downward trend in fertility. The large and significant residual 'cohort effects' in the birth hazard equations are in line with this conclusion.

Endnotes

[1] The lack of jobs during the Great Depression and full employment during the 1950s and 1960s were closely matched by corresponding fluctuations in fertility. Southall and Gilbert (1996) find a strong negative correlation between the marriage rate and the unemployment rate during the second half of the 19th century and the early 20th century in England and Wales. See other references in their article for more historical evidence. Several studies using time series data of the last several decades for developed countries also indicate a significant negative correlation of unemployment and fertility (Ahn and Mira 1999; Gauthier and Hatzius 1997; Macunovich 1996). Most recently, the experience of East Germany, Russia and other Eastern European countries also suggest the existence of strong negative correlation between unemployment and family formation (Eberstadt 1994; Witte and Wagner 1995).

[2] Unemployment rates measured from the Encuesta de Población Activa, the Spanish Labor Force Survey; definitions of unemployment are standard.

[3] The decline of permanent contract holders was even greater among younger workers. For example, among males aged 20–24 this proportion fell from 28% to 12% between 1987 and 1997.

[4] It has been suggested that current low fertility may be due to timing effects, i.e., to the delay of childbearing. We believe delayed births are not a major factor in explaining the recent fertility decline in Spain. As an example, suppose the decline in the total fertility rate below its 1982 level of 2.0 were due to the delay of births from ages 20–29 to ages 30–39. During the 10 years between 1982 and 1991 the fertility rate for women in their 20s dropped by 0.44 children. To compensate for this drop, the fertility rate for women in their 30s would have to increase approximately from 0.52 children in 1991 to 0.96 children by the year 2001. During the first 4 years, between 1991 and 1995, the fertility rate for women in their 30s increased from 0.52 to 0.58. It seems extremely unlikely that the remainder will be recovered in the following 6 years.

[5] Some studies have implemented empirical structural models. The models incorporate many simplifying assumptions in order to make the models empirically tractable. See the surveys in Hotz et al. (1997) and Arroyo and Zhang (1997).

[6] See Jenkins (1995) for an easy-to-use implementation of this model in STATA.

[7] That is, our model is one of individual and spell – specific unobserved heterogeneity, rather than permanent individual-specific heterogeneity.

8 In our sample, only 2.79% of first marriages were preceded by cohabitation.
9 Individuals filtered out of the sample for this reason were 3.8% of those with a college degree, 0.3% of those with secondary education and none of those with primary education.
10 Since these differences in the person's age at the beginning of the interval seem less of an issue for subsequent birth intervals and since sample sizes are smaller, we estimated the models for the second and third birth intervals jointly for all education categories.
11 Due to sample selection, the empirical results cannot be generalized to the whole population.
12 The beginning of the n-th birth interval is of course the date of marriage (for $n = 1$) or the date of the $(n - 1)$-th birth (for $n > 1$). It would be hard to interpret the duration of the first unemployment spell as a covariate in the birth interval equations. The parental variables were included initially, but they were not significant: their main effect is exerted through the duration to marriage. Furthermore, reduced sample sizes in the analyses of higher birth intervals is likely to lead to less accurately estimated coefficients.
13 Strictly speaking, this use of the test statistics is not completely correct because the null in a) is in the boundary of the parameter space and the null in b) is a limit of nested models.
14 Strictly speaking, cohort and calendar time effects are not separately identified. Furthermore, the cohort/time coefficient could also capture age effects in the birth hazard equations, since the mean age at marriage is not the same across cohorts. However, the coefficients on age variables were not significant in earlier specificactions which included them. Therefore, we believe it is the cohort/time effects that dominate and this is reflected in our interpretation of the results.
15 We also included a dummy for the gender of the first child in the second birth equation, but its coefficient was not significant

Appendix 1: Cumulative proportion ever married and number of children

Birth cohort	By age 25	By age 30	By age 35	By age 40	By age 45	Sample size
Proportion ever married						
Men						
1941–1945	34	72	81	84	85	4161
1946–1950	39	73	81	84		4837
1951–1955	45	76	85			5678
1956–1960	46	77				7497
1961–1965	41					6928
Women						
1941–1945	64	82	87	88	88	4044
1946–1950	65	83	87	88		4427
1951–1955	67	84	88			5231
1956–1960	67	85				7267
1961–1965	64					7749
Number of children						
Men						
1941–1945	0.23	1.05	1.65	1.92	2.05	4161
1946–1950	0.26	1.04	1.54	1.79		4837
1951–1955	0.30	0.97	1.47			5678
1956–1960	0.30	0.90				7497
1961–1965	0.28					6928
Women						
1941–1945	0.74	1.58	2.03	2.20	2.25	4044
1946–1950	0.74	1.49	1.87	2.02		4427
1951–1955	0.68	1.33	1.70			5231
1956–1960	0.64	1.24				7267
1961–1965	0.58					6928

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Effect of childbearing on Filipino women's work hours and earnings

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Abstract. The effects of childbearing and work sector on women's hours and earnings in the 8 years following an index pregnancy were examined in a cohort of more than 2,000 women in the Cebu Longitudinal Health and Nutrition Survey. Change in cash earnings and hours worked were each modeled jointly with sector of labor force participation using an estimation strategy that deals with endogeneity of childbearing decisions and selectivity into sector of work. Two or more additional children born in the 8 year interval significantly reduced women's earnings, while having an additional child under 2 years of age in 1991 reduced hours worked.

JEL classification: J13, J21

Key words: Labor force participation, childbearing, wages

1. Introduction

It is generally assumed by those who promote family planning that having fewer children improves child health and survival, and also contributes in pos-

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itive ways to the quality of women's lives. While the former assumption is well founded and documented in an extensive literature (see, for example, Rosenzweig and Schultz 1987), there remain questions about how family planning influences various aspects of women's lives (Rinehart and Kols 1984). There are significant health costs for women in developing countries of high levels of reproductive stress, particularly when coupled with chronic undernutrition and physically demanding work. Repeated cycles of pregnancy and lactation can deplete maternal energy and nutrient stores (Merchant and Martorell 1991; Adair and Popkin 1992). There are also hypothesized effects of childbearing on labor force participation, earnings, social-psychological well-being, and overall quality of life. For example, Rosenzweig and Wolpin (1980) show a negative effect of fertility on labor force participation and that this effect is negatively biased in models that assume that fertility is exogenous.

Women in developing countries are faced with competing demands of reproductive and productive roles. It has been argued that by limiting family size, constraints to productive economic activities are reduced (Podhista et al. 1990). With fewer children, women can theoretically earn more income and thereby improve their own well-being as well as that of other family members. While this hypothesis is appealing, there is a lack of empirical evidence from developing countries supporting a strong effect of childbearing on women's earnings.

Researchers uniformly acknowledge that the relationship of work for pay and childbearing is very complex. Angist and Evans (1998) point out that the economics and demography literature includes studies of how labor supply affects fertility, and how fertility affects labor supply, precluding any causal interpretation of the relationships. Since decisions about childbearing and work are likely to be made jointly, it is challenging to disentangle cause and effect (Schultz 1981). Working women may choose to limit childbearing so they can continue to work and earn income. Alternately, childbearing may decrease women's opportunities for work or affect type of work, place of work, or hours. Either sequence would result in a negative association of work and childbearing. In contrast, increased economic demands of larger family size may push women into the labor force, resulting in a positive association of work and childbearing. In either case, the nature of the work-childbearing relationship may depend on the type of work women do.

A closely related literature has examined the relationship between women's work and child care arrangements. In this research, the models are usually conditioned upon births of children and then the effects of work on child care arrangements are examined (for example, see: Leibowitz et al. 1992; Michalopoulos et al. 1992; Connelly et al. 1996). Typically, this line of research has attempted to estimate structural equations models where some measure of the woman's employment status is included as an endogenous explanatory variable in a choice of child care equation. This methodology, of course, requires exclusion restrictions in the child care equation. In contrast, in our empirical work, the number of children born to a mother is endogenous and we use longitudinal data and the fixed effects method rather than structural equation methods to estimate the effects of children on earnings.

Much of the income-producing work done by women in developing countries is carried out in the informal sector, that is, under conditions lacking routine government regulation, formal contractual arrangements or benefits such as health insurance (Portes and Walton 1981). Informal sector work often involves market sales, food production, and piece work (Bunster 1983; Tinker

1981; Cohen and House 1996). Self-employed workers in the informal sector often own their own means of production. Compared to work in the formal sector, informal sector work offers women flexibility in hours, but less security, status, and non-wage benefits such as social security, health insurance, or paid sick leave. Given its greater flexibility, women may regard informal sector work as more compatible with childbearing. In contrast, work in the formal sector offers greater stability, some non-wage benefits, and greater structure, but generally longer hours and a work site typically away from home (Anker and Hein 1985).

Becker (1981) hypothesizes that formal sector work away from home presents higher opportunity costs, and thus acts as a greater deterrent to childbearing. Under this premise, limiting family size should promote participation in the formal work sector, but have relatively little impact on the informal work sector where opportunity costs are lower and productive and reproductive roles are more compatible.

The type of work selected by women is affected by a wide range of factors including education and training, prior work experience, earning potential, compatibility with other roles (childbearing, elder care, other household responsibilities etc.), household characteristics including income, partner's education, and presence of childcare substitutes, and community level characteristics such as infrastructure, labor opportunities and wage structure. For many women in the Philippines, work is regarded solely as an economic necessity. For some, work is a choice that carries other benefits, including status and a sense of personal achievement. Thus, the choice to work or not, and the type of work selected may differ at opposite ends of the income distribution. Women from higher income households may elect not to work unless they are also highly educated while women from poor households may be driven to work by economic necessity.

To understand the effect of childbearing on women's earnings, it is important to consider not only the type of work women do but also the hours they work. Women may increase earnings by working more hours, but this may create more competition between their productive and reproductive roles, and thus result in no net gain in other quality of life outcomes. One strategy to overcome the time demand dilemma is for women with children to select jobs with more flexible hours but lower pay (Becker 1981). It is therefore important to explore the effects of childbearing on earnings both as a function of hours and of earnings per unit of time worked.

Childbearing is likely to affect women's earnings in 2 major ways. Childbearing may influence whether or not women work for pay at various points in their reproductive years, and having small children may influence the hours and earnings of women who are working. These effects will likely be mediated by the type of work women do and by the availability of childcare substitutes.

Mothers working in the formal wage sector where opportunity costs are high and flexibility for childbearing is more limited, may withdraw from the labor force for a period of time or may shift to a lower-paying but more flexible job. Women from lower income households may be less able to withdraw from the labor force and may continue working at the same or reduced hours. For women who shift to a lower-paying job without increasing hours, childbearing would result in reduced income. Self-employed women with work that is relatively compatible with child care may be able to continue their work and may see little decrease in income. In their study of piece workers in Mexico

City, Beneria and Roldan (1987) found that compatibility with childcare is a reason often given by women for why they are involved in piece work. Women engaged in piece work or whose self-employed work is fairly demanding may work similar hours, but find that they are less productive because of the demands of child-care. These women may earn less income, may increase their work hours to maintain their profits, or may switch to work that is less demanding and usually less profitable.

Finally, the effects of childbearing are likely to depend on its different phases. Late in pregnancy, women may find it difficult to continue working because of physical constraints. Mothers of infants and young children face the time demands of more intensive childcare and breast-feeding. As children get older, these demands may decline or child care may be more easily assumed by alternate care givers. In studies of US women showing a relationship of childbearing and earnings, the age of the youngest child at the time of measurement is a stronger predictor of wages than the total number of children (Lehrer 1992). In a summary of several studies that used mainly US data, Korenman and Neumark (1992) concluded that "children appear to reduce wages primarily 'indirectly' by reducing labor force participation and accumulation of human capital rather then 'directly' by lowering the productivity of otherwise similar work." In their own study, Korenman and Neumark control for the endogeneity of children through the use of longitudinal data and a fixed effects estimator. After this endogeneity is controlled, the effect of children on wages is found to be insignificant (Korenman and Neumark 1992).

Given the vast differences in labor markets in developed versus developing countries, it may not be possible to generalize results from the US. Further research in developing countries is needed to explore the effects of number and ages of children on women's earnings. Such research requires detailed longitudinal data on women at different stages of their reproductive lives. Of particular importance is the ability to document the sequence of work and childbearing events.

We use data from the Cebu Longitudinal Health and Nutrition Survey (CLHNS) to explore the determinants of change in women's hours worked and earnings over an 8 year period, while controlling for the selectivity into working for pay and the endogeneity of childbearing and work sector. The CLHNS is particularly well suited to this purpose, since data are available on work patterns, earnings, and reproductive events in a large cohort of women over an 8 year period. Prospective data allow us to examine changes in women's labor force participation and earnings subsequent to a well documented pregnancy, but also conditioned on prior reproductive experiences and a wide range of other maternal characteristics. Detailed data on type of work allow differentiation by wage, self-employment, and piece work sectors.

2. Data

CLHNS data were collected from 1983 to 1991 in Metro Cebu, the second largest metro area in the Philippines, and a major shipping center undergoing rapid urban development and population growth. Like the rest of the Philippines, the Central Visayas region where Cebu is located, has a moderately high fertility rate. In 1993, the total fertility rate (TFR) for this region was 4.38, compared to 4.09 for the entire country. These levels represent only modest declines from the previous decade, when TFR was 4.7 for the entire country (DHS 1994).

The study area encompasses the densely populated urban areas of Cebu City and other contiguous smaller cities, peri-urban areas around these cities, and more isolated rural areas in the mountains and on smaller islands. The CLHNS is a community based study of a cohort of childbearing women, ages 14 to 47 at entry into the study. Following a census of all urban and rural barangays (administrative units) of the Metro Cebu area, 33 barangays were randomly selected for the study. All pregnant women in these barangays who subsequently had a birth or pregnancy termination in a one year period from 1983 to 1984 were recruited into the study ($n = 3{,}327$). The study sample is representative of childbearing women in communities of Metro Cebu, with a preponderance of low income families, a wide range of variability in income, education, and level of modernization.

Highly trained interviewers from the Office of Population Studies at the University of San Carlos visited each household to collect sociodemographic, environment, health, nutrition, and reproductive history data. Data collection took place during the 6th to 7th month of pregnancy, then at bimonthly intervals for 2 years. The first follow-up survey was conducted in 1991, and 2,395 women (72% of the baseline sample) were located and re-interviewed. We focus this analysis on the women present at the 1991 follow-up.

In the majority of cases, loss to follow-up occurred because families moved away from the Metro Cebu area. Most migration took place during the first several years of the study. Relatively little migration occurred from 1986 to 1991 resulting in low rates of loss to follow-up during this period. Compared to the women in the 1991 survey, those lost to follow-up ($n = 932$) were on average, 1 year younger, completed about 1 additional year of schooling, of slightly lower parity, and more likely to be from urban areas. Despite measurable differences in characteristics of migrants, in previous studies of health-related outcomes, we found no significant biases in multi variate results attributable to loss to follow-up (Guilkey et al. 1989).

2.1. Definition of study variables

1. *Labor force participation* was defined for purposes of this study as "working for pay". During the baseline survey, women were asked if they were currently doing any work for pay, or whether they had worked in the previous 4 months. We define as working, those women with affirmative answers to either question, since women may have recently stopped working because of advancing pregnancy. Only 10% of the women we classified as working were not currently working but were working 4 months ago. Excluded from this classification of work are those women who are unpaid workers in a family business or farm. This type of work was excluded because we wanted to focus on women's cash earnings. The same criteria were used to identify work status of women in 1991.

Work done by Cebu women was categorized into 3 work sectors based on the mode of pay. Data were not collected on contractual arrangements, so we do not use the standard formal-informal sector dichotomy. *Wage* workers are paid on a time basis (most often daily or hourly), and may or may not have formal contracts and benefits. Most wage workers are in service-related jobs (clerical, domestic help, etc.). *Piece* workers are paid by the amount of goods they produce. Most piece work in Cebu involves manufacture of shell, wood or fabric handicrafts. *Self-employed* women run their own small businesses. Many of these are small "sari-sari" stores which have a limited inventory of sundries and food items, and are typically located in the woman's own home or a neighboring building. Other small businesses may involve selling prepared foods.

2. *Total household income* was determined based on a series of questions about earnings in primary and secondary jobs, and in family businesses (including raising of livestock, sale of produce). Income was measured during the baseline survey, at 12 and 24 months postpartum, and in 1991 using the same set of questions to insure comparability across the surveys. Using Philippines consumer price indices, income values were deflated to a common time point in 1983 so that real changes in income could be determined.

Change in women's total weekly cash earnings and change in weekly hours worked in the interval from the baseline survey to 1991 are the main dependent variables in our analysis. We restrict our analysis to cash income for several reasons. First, while the tradition in the Philippines is for men to turn over their earnings to their wives, Filipino women, like women elsewhere, have more autonomy in spending their own cash income, and tend to spend their earnings to benefit their children and themselves (Bruce et al. 1995; Kennedy and Peters 1992; Bisgrove and Popkin 1996; Haddad et al. 1997). Second, the non-cash component of income, typically derived from work on a farm or in a family business, is a measure of household activity. We have no precise measures of women's contributions, and their incomes could only be estimated with considerable error.

3. *Hours* worked. Women were asked to report the usual number of hours per day, and the number of days per week they worked in the past 4 months. Hours were summed for main and secondary jobs.

4. *Childbearing patterns* were determined from detailed reproductive histories. The multi variate model includes a variable indicating change in the number of surviving children under two years of age and a set of dummy variables indicating the number of children born after the index child, and still alive and in the household at the time of the 1991 survey.

5. *Other covariates.* Mother's age, education, place of residence (urban or rural community), and family structure (presence of spouse, presence of individuals who may serve as alternate care givers, including grandmothers, and other adult females) were also entered into analysis. Finally, we have community level information about the prices of major commodities that may indirectly affect the outcome variables. The prevailing community-specific wage rate for yayas (child-care providers) may serve as a demand side control for labor market conditions.

The mean and standard deviations of all variables used in the multi variate analysis are presented in Table 1. The earnings equations are estimated for the sub-sample of women who were working for pay both in 1983 and 1991, so the descriptive statistics are presented for this sample. In addition, this equation is estimated in differenced form and we present sample statistics for these differenced variables. The bottom half of the table presents statistics for the variables used to explain work sector. We estimate separate equations for choice of sector at baseline and 1991 and means and standard deviations for both time varying and time invariant explanatory variables are presented. Frequencies for the sector choice variable are presented in Table 2 and discussed in the next section.

Table 1. Means and standard deviations of variables in the multivariate analysis

Variables in the change in weekly income and change in hourly earnings equations ($n = 769$)		
Variable	Mean	S.D.
Change in weekly income	49.38	123.75
Change in hours per week	4.95	32.18
Change in hourly earnings	0.77	0.42
1 surviving child in interval	0.28	0.45
2 surviving children in interval	0.21	0.41
3 surviving children in interval	0.16	0.36
4 surviving children in interval	0.05	0.22
Change in number of children < 2 yrs	−0.04	0.64
Spouse absent then present	0.02	0.14
Spouse present then absent	0.06	0.24
Spouse present both times	0.89	0.31
Other adult female absent then present	0.06	0.24
Other adult female present then absent	0.13	0.34
Other adult female present both times	0.04	0.19
Stay in wage sector	0.21	0.41
Stay in piece sector	0.09	0.29
Move from wage to piece	0.04	0.20
Move from wage to self employed	0.07	0.25
Move from piece to wage	0.05	0.22
Move from piece to self employed	0.07	0.25
Move from self employed to wage	0.09	0.28
Move from self employed to piece	0.05	0.21
Change in the price of formula (pesos/100 g)	55.7	19.0
Change in yaya wage (pesos/week)	69.0	33.5
Variables in the sector choice equations ($n = 2311$)		
Mother's age	26.3	6.1
Mother's highest attained school grade	6.9	3.3
Spouse's age	27.6	8.9
Spouse's grade	6.8	3.7
Spouse present baseline	.95	.21
Spouse present 1991	.92	.27
Formula price baseline (deflated)	91.5	17.8
Formula price 1991 (pesos/100 g, deflated)	146.9	8.6
Yaya wage baseline (deflated)	115.1	22.5
Yaya wage 1991 (deflated)	404.5	116.1

Table 2. Sample sizes in the various cells for sector of work at baseline and 1991

	Not working 91	Wage sector 91	Piece sector 91	Self emp 91
Not working 83	496	295	136	409
Wage sector 83	57	162	33	74
Piece sector 83	62	41	72	51
Self emp 83	87	67	37	232

3. Methods

A critical methodological consideration concerns sources of potential bias. First, unobserved characteristics of women are likely to affect both their work and childbearing decisions, leading to biased estimates of the effects of childbearing on work. Second, women are likely to make simultaneous decisions about work and having children. They may choose not to work while children are young, or they may choose specific jobs that offer greater flexibility to care for young children, but shorter hours or lower pay. A woman with a secure, well paying job may elect to delay childbearing to maintain her work status. Thus it may be virtually impossible to identify the effect of childbearing on work with cross sectional data. One possible approach involves the use of instrumental variables for childbearing. However, this presupposes that instruments can be identified. Given the joint nature of work and childbearing decisions, it is nearly impossible to find exogenous variables that affect one decision but not the other. The difficulty of this problem has led several groups of researchers to make the simplifying assumption of treating the number and age of children as exogenous variables in models of labor force participation or earnings (Lehrer 1992; Blank 1988, 1989; Hill and Stafford 1980; Leibowitz et al. 1988).

Our estimation strategy is an extension of the method used by Korenman and Neumark (1992) to study the effects of marriage and fertility on wages for women in the US. They used the 1980 and 1982 panels from the National Longitudinal Survey of Young Women to estimate a first-differenced model for wages. The model eliminates possible bias related to unobserved fixed variables that affect both wages, fertility and marriage. Because substantial numbers of women are not employed in both 1980 and 1982, they jointly estimate the change in wage equation with a bivariate probit model that explains labor force participation at each point in time since a woman must work both in 1989 and 1982 to construct the change in wage dependent variable. In our analysis, we were interested in how childbearing affects women's earnings and work hours. As discussed above, women earn income in either the wage, piece or self-employed sectors and we hypothesize that these sectors have differing effects on wages and income. Thus rather than simply estimating a bivariate probit to control for selectivity into the change in earnings and hours equations, we estimate multinomial logit models for sector choice in 1983 and 1991. Joint estimation of the sector choice equations with the change in earnings and hours equations also allows us to control for the endogeneity of change in sector which we include as a set of variables in the change in income and wage equations as well as the selectivity of labor force participation.

Table 2 shows sample sizes in the various cells for sector of work at baseline and 1991. The lower 3×3 submatrix is the sample of 769 individuals for whom we have income and work hours for both 1983 and 1991. There was considerable shifting among the sectors through time. To see if these shifts affect the outcome variables, we define 9 dummy variables, each of which represents a cell in the table. The diagonal represents individuals who remain in the same sector.

Our complete estimation strategy deals with both the endogeneity of childbearing decisions and selectivity into work sector among women who were working at both points. The statistical specification is as follows. Consider the following reduced form equations for sector choice:

$$\ln\left[\frac{P(S_{tij}=k)}{P(S_{tij}=1)}\right] = Z_{tij}\beta_{kt} + \mu_{kjt} + \eta_{kijt} \tag{1}$$

where the dependent variable is the log odds that woman i from community

j worked in sector k (k = wage sector, piece sector, self employed sector) relative to sector 1 (not working) at time t. The Z's represent exogenous characteristics of the woman and her community that may affect sector choice. The individual level variables are the woman's age, education, whether she lives in an urban area, whether or not her spouse lives in the household and her spouse's age and education. The community level variables are the price of infant formula and the barangay-specific yaya wage rate. We also tried additional price variables but they were too collinear with the price of formula. These price variables may help determine the woman's labor supply as well as indicate general economic conditions in the community. The wage for yayas is an indicator of the demand for labor in the woman's community.

The μ's represent unobserved community level variables that affect sector choice while the η's represent unobserved, time varying individual level errors. Implicit in the logistic specification is an additional error that is independent across the $k - 1$ equations. If it is assumed that the μ's and η's are allowed to be correlated across equations then this specification does not suffer from the independence of irrelevant alternatives problem for the standard multinomial logit model.

The change in income (or hours) equation takes on the following form:

$$\Delta E_{ij} = \Delta X_{ij}\alpha + \Delta CB_{ij}\gamma + \Delta S^*_{ij}\delta + \mu_{Ej} + \eta_{Eij} + \varepsilon_{Eij} \tag{2}$$

where the dependent variable is the change in income or hours between 1983 and 1991 for woman i from community j and is only observed if the woman is working at both points in time ($S_{tij} \neq 1$ for both $t = 1$ and 2). The ΔX's represent changes in exogenous household and community variables such as the change in the price of formula and the wage rate for yayas, change in other adult females in the household, and change in the presence of the woman's spouse in the household.[1] The ΔCB's represent changes in the childbearing variables (children under 2, and number of surviving children born in the interval). The ΔS^*'s represent a set of 8 change in sector variables with staying in the self employment sector the omitted category.

The μ's and η's represent unobserved community and individual level errors that are allowed to be correlated with their counterparts in the sector equations while the ε's are assumed to be independent, identically normally distributed errors with mean zero and variance σ^2.

Since the error terms across the two equations are assumed to be correlated, the use of ordinary least squares in Eq. (2) will result in biased and inconsistent parameter estimates even though the data is already differenced. This is due to two reasons. First, Eq. (2) contains a self selected sample of women who were working at both points in time. Second, we allow sector choice to affect the level of income and wages. Our solution to these problems is joint estimation of Eqs. (1) and (2) by maximum likelihood methods using a discrete factor approximation to the distribution of the unobservables. Specifically, we use an extension of Heckman and Singer's (1984) semiparametric method that does not impose any specific distributional assumptions on the unobserved heterogeneity, but instead assumes that it can be approximated by a discrete probability distribution where both the mass points and the probabilities are estimated. This approach has been used successfully by Mroz and Weir (1990) to estimate discrete time hazard models for child spacing and by Blau (1994) to estimate the hazard of retirement. Guilkey and Stewart (2000) and Guilkey and Riphahn (1998) have applied the method to two other projects using the Cebu data: a structural analysis of infant mortality and an analysis of the effects of food industry marketing practices on child morbidity. A comparison of the assumption of normal errors with that of nonparametric error term distributions in structural equations models was done by Mroz and Guilkey (1992) and Mroz (1999). They found that when the true distribution of the errors was approximately normal, the parametric and nonparametric estimators gave very similar results. When the true distribution was far from normal, the nonparametric estimator generated much more accurate parameter estimates.

To build the likelihood function, we start with the work sector equations. Define the indicator variable I_{kijt} to be equal to one if woman i from community j worked in sector k at time t. The contribution to the likelihood function conditional on μ and η for woman i from community j at time t for the choice of sector dependent variable can then be written as:

$$L_{Sijt}(\mu_{jt}, \eta_{ijt}) = \prod_{k=1}^{4} P(I_{kijt} = 1 \mid \mu_{kjt}, \eta_{kijt})^{I_{kijt}} \tag{3}$$

The contribution to the likelihood function for Eq. (2), again conditional on the two sets of unobservables, is:

$$L_{Eij}(\mu_{Ej}, \eta_{Eij}) = \frac{1}{\sigma}\phi\left(\frac{\Delta E_{ij} - \Delta X_{ij}\alpha - \Delta CB_{ij}\gamma - \Delta S_{ij}^{*}\delta - \mu_{Ej} - \eta_{Eij}}{\sigma}\right) \tag{4}$$

where ϕ is the standard normal density function and σ is the standard deviation of ε_{ij} and all other terms are as defined previously. Using Eqs. (3), and (4), the contribution to the likelihood function for woman i from community j conditional on μ and η is:

$$L_{Sij1}(\mu_{j1}, \eta_{ij1}) L_{Sij2}(\mu_{j2}, \eta_{ij2}) L_{Eij}(\mu_{Ej}, \eta_{Ej}). \tag{5}$$

The contribution to the likelihood function unconditional on η is obtained by integrating over the range of η. Suppose we approximate the distribution of η with Q discrete points, each with probability P_q $(q = 1, 2, \ldots, Q)$, then:

$$L_{ij}(\mu_{j1}, \mu_{j2}, \mu_{Ej}) = \sum_{q=1}^{Q} P_q L_{Sij1}(\mu_{j1}, \eta_{q1}) L_{Sij2}(\mu_{j2}, \eta_{q2}) L_{Eij}(\mu_{Ej}, \eta_{Eq}). \tag{6}$$

The estimation procedure that we use does not restrict the mass points for the discrete distributions to be the same across equations. Instead, using a generalization due to Mroz (1997), each mass point is estimated separately for each equation. This more general specification allows more flexibility in the pattern of correlations across the error terms. For more details, see Mroz (1997) who refers to the more general specification as nonlinear heterogeneity.

The product of (6) over the N_j women in community j yields the contribution to the likelihood function for community j conditional on μ:

$$L_j(\mu_{j1}, \mu_{j2}, \mu_{Ej}) = \prod_{i=1}^{N_j} L_{ij}(\mu_{j1}, \mu_{j2}, \mu_{Ej}) \tag{7}$$

The contribution to the likelihood function for community j unconditional on μ is obtained by integrating over the range of μ. Suppose we approximate the distribution of μ with R discrete points, each with probability P_r $(r = 1, 2, \ldots, R)$, then:

$$L_j = \sum_{r=1}^{R} P_r L_j(\mu_{1r}, \mu_{2r}, \mu_{Er}) \tag{8}$$

where we again allow the mass points to differ across equations. The likelihood function is simply the product of (8) over the J communities.

The discrete factor FIML method is identified without exclusion restrictions. However, we do not need to rely on nonlinearities to obtain identification. The exclusion restrictions used to identify the model are discussed in the empirical results section of the paper.

4. Results

Results presented here are for the 2,309 women who were participants in the 1991 follow-up survey, and who had complete data on the variables of interest. As discussed below, two formulations of the dependent variable were used: change in the woman's weekly income and change in her work hours. For each dependent variable, Eq. (2) was estimated jointly with the two equations specified in Eq. (1) for time periods 1983 and 1991.

Before turning to a discussion of the substantive results, we first discuss estimates of the heterogeneity parameters. In the change in income model, 2 mass points were optimal for the individual level heterogeneity (η's) and 6 mass points were optimal for community level heterogeneity (μ's). The estimated mass points are at the bottom of Table 3 for choice of sector, Table 4 for change in income. Note that the first mass point is normalized to zero since it cannot be separately identified from the constant term. The estimated probability weights for the individual level heterogeneity are 0.32 and 0.68 respectively (see Eq. (6)). The probability weights at the community level are 0.22, 0.12, 0.06, 0.13, 0.34, and 0.13 (see Eq. (8)). As can be seen from the tables, many of the heterogeneity parameters are highly significant. A joint test of the null hypothesis that the heterogeneity parameters are all zero yields a chi square statistic of 344.2. Since the critical value for a 1% test with 48 degrees of freedom is 73.7, there is strong evidence of the importance of controlling for unobserved heterogeneity.

Because the parameters of the heterogeneity distributions are difficult to interpret, we also calculated the correlations in the community level errors implied by the estimated discrete distributions. Rather than presenting the entire

Table 3. Multinomial logit estimates for the determinants of work sector at baseline and in 1991

Variable	Baseline			1991		
	Wage vs Not work	Piece vs Not work	Self-Emp vs Not work	Wage vs Not work	Piece vs Not work	Self-Emp vs Not work
Mother's age	1.59 (0.57)[1]	0.03 (0.20)	1.91 (0.60)	−0.01 (0.15)	−0.29 (0.19)	0.09 (0.14)
Mother's highest attained grade	−0.72 (0.26)	0.07 (0.11)	−0.45 (0.29)	−0.18 (0.10)	0.22 (0.10)	0.03 (0.07)
Grade squared	6.97 (1.86)	−1.58 (0.79)	4.38 (1.93)	1.75 (0.42)	−1.78 (0.72)	−0.08 (0.43)
Urban	−1.28 (1.28)	−0.98 (0.25)	−1.03 (1.30)	−0.10 (0.17)	0.36 (0.23)	−0.05 (0.16)
Spouse's age	−1.33 (0.62)	−0.32 (0.20)	1.55 (0.63)	0.12 (0.14)	0.36 (0.18)	0.30 (0.13)
Spouse's grade	−0.05 (0.12)	−0.05 (0.03)	−0.03 (0.12)	−0.02 (0.02)	−0.08 (0.03)	−0.03 (0.02)
Spouse present	−2.63 (1.52)	0.65 (0.68)	−3.35 (1.64)	−0.08 (0.47)	0.08 (0.64)	−0.03 (0.45)
Formula price	−0.39 (0.59)	0.01 (0.11)	−0.23 (0.61)	−0.42 (0.70)	−0.29 (0.96)	−0.45 (0.65)
Yaya wage	2.28 (1.64)	−2.95 (0.42)	2.48 (1.76)	−0.03 (0.06)	−0.12 (0.08)	0.05 (0.05)
Constant	−2.18 (3.44)	0.94 (2.38)	−3.64 (3.61)	1.20 (1.20)	−0.99 (1.64)	0.24 (1.14)
Community heterogeneity (6 points of support)						
Parameter 1 normalized to zero						
Parameter 2	4.89 (1.91)	1.48 (0.40)	4.61 (1.90)	0.46 (0.20)	−0.58 (0.39)	0.01 (0.20)
Parameter 3	−0.35 (2.06)	2.00 (0.39)	0.09 (2.06)	0.79 (0.43)	2.67 (0.42)	−0.12 (0.46)
Parameter 4	5.30 (1.54)	1.17 (0.38)	5.99 (1.53)	0.10 (0.26)	1.17 (0.28)	0.64 (0.21)
Parameter 5	1.21 (0.77)	1.32 (0.29)	0.43 (0.79)	0.02 (0.16)	0.08 (0.23)	−0.30 (0.14)
Parameter 6	3.25 (1.72)	3.07 (0.35)	2.36 (1.74)	0.31 (0.22)	1.26 (0.27)	0.01 (0.20)
Individual heterogeneity (2 points of support)						
Parameter 1 normalized to zero						
Parameter 2	−11.57 (3.29)	0.58 (2.53)	−13.55 (3.10)	−0.83 (0.14)	−0.20 (0.19)	−0.94 (0.14)

[1] Numbers in parentheses represent standard errors.

7×7 correlation matrix, we only present correlations of the six multinomial logit equations with the change in income equation for baseline and 1991, respectively:

	Baseline	1991
Wage vs not work	0.354	0.562
Piece vs not work	0.200	−0.258
Self employed vs not work	0.313	0.068

The effects of these across-equation error correlations on the substantive results are discussed below.

The bottom of Table 5 presents the results for the heterogeneity parameters in the change in hours equation. In this equation, the distribution of the unobserved heterogeneity was best approximated with 6 mass points for the community level error and 3 mass points for individual level heterogeneity. The probability weights were 0.12, 0.29, 0.06, 0.18, 0.12, and 0.23 and 0.13, 0.46, and 0.41 for community and individual heterogeneity respectively. A chi squared test of the null hypothesis that all the heterogeneity parameters are zero resulted in a test statistic of 486. Since the critical value for a 1% test with 56 degrees of freedom is 83.5, we again see significant improvement in the likelihood function with these controls.

We now present the substantive results. We first discuss the results for the two multinomial logit estimations for the choice of sector of employment at baseline and in 1991. As stated above, we did pair wise estimation of the two set of sector choice equations (see Eq. (1)) with change in weekly income and change in hour, see Eq. (2), so two sets of results are available for the choice of sector at baseline and 1991. Since the choice of sector equations are reduced form equations, we did not expect to see much difference in the results from estimating them with the change in income or change in hourly earnings equations and this is exactly what we find. Therefore, Table 3 only presents results for the determinants of sector choice estimated jointly with the change in income equation.

4.1. Women's choice of work sector in 1983 and 1991

The results of the multinomial logit estimations for choice of sector in 1983 and 1991 are presented in Table 3. We present the log odds (and standard errors) of working in the wage, piece and self employed sectors relative to not working. This set of equations are estimated to control for the self selection of women into working at both points in time and are not the equations of primary interest. Since they are reduced form equations, they are difficult to interpret and so we only present some fairly general observations.

Not surprisingly the strongest predictors for work sector are woman's education and education squared. At both baseline and 1991, grade squared is strongly positively associated with working for wage and strongly negatively associated with doing piece work. The other education results are somewhat mixed with some sign flips between 1983 and 1991. However, the sign flips are not associated with strongly significant estimated coefficients. Mother's age is a strong predictor of wage and self employed categories versus not working in 1983 but age seems to have little effect in 1991. Interestingly, spouse's age also effects these two categories in 1983. In 1991, spouse's age has a positive effect for all three work categories relative to not working. Formula price appears to have little effect on work sector while the wage rate for yayas is a positive predictor of the wage and self employed categories relative to not working in 1983 but has little effect in 1991.

In summary, there are several significant predictors of work sector with stronger results for 1983 than for 1991. This may reflect the different composition of the reference group at each point in time, or differences in the reproductive status of working and non-working women. In 1983, all women were pregnant and very few had no other children, whereas in 1991, 8.3% were pregnant, and about 20% had no surviving child. Alternately, with the trend toward higher labor force participation rates, fewer factors differentiate workers and non-workers. These reduced form equations are not of much substantive interest for this paper. The important result is that there are significant predictors of work sector that do not have direct effects on change in earnings, the results for which we discuss below.

4.2. Change in earnings and hours

We calculated the change in weekly earnings from the baseline (1983) to 1991 for the 769 women working at both points in time. Earnings are highly variable and skewed. On average, women earned 46.6 pesos/week more in 1991, but the median change was 29.6 pesos/wk. The mean change represents a 49% increase over women's mean weekly earnings of 94 pesos per week in 1983–1984. The mean change in earnings for women with no subsequent surviving child in the 8.5 year period was 2.3 times higher than that of women with at least one additional child (69.7 vs. 26.4 pesos/wk).

The overall change in income is in part, a reflection of an increase in the mean number of hours worked per week from 41.6 in 1983 to 46.1 in 1991. More than half of sample women increased their work hours. This reflects a tendency to move from part-time to full time work. Changes in income varied substantially across sector of employment at baseline, with piece workers having the lowest mean gains (18.6 pesos/wk) and wage workers having the highest gain (62.9 pesos/wk). Among women working at both points in time, earnings represented 34.9% and 36.7% of total household earnings in 1983–1984 and 1991, respectively.

The model specification includes dummy variables representing the number of children born in the interval from 1983 to 1991 (1, 2, 3, 4+) and surviving at the time of the 1991 survey, change in the number of children less than 2 years of age , change in the presence of spouse and other adult females in the household, and a set of eight dummy variables for sector change with staying self employed the omitted category. We also tested but found no significant effect of interactions of sector of employment with childbearing and spouse present variables.

Table 4 presents results from the selectivity corrected change in weekly income regression, and Table 5 presents results from the change in hours regression. Children born during the 8 year interval decrease weekly income. The effect strengthens with each additional child going from a non-significant level

Table 4. Results from selectivity corrected linear regression models of change in weekly income (mean change in income = 46.56 pesos/week)

Variable	Coefficient	S.E.	*T*-value
1 surviving child in interval	−6.35	11.01	−0.58
2 surviving children in interval	−19.79	12.20	−1.62
3 surviving children in interval	−27.96	13.69	−2.04
4 surviving children in interval	−38.32	19.61	−1.95
Change in number of children < 2 yrs	−8.88	6.91	−1.29
Spouse absent then present	1.51	32.11	0.05
Spouse present then absent	−2.46	40.59	−0.06
Spouse present both times	19.20	28.28	0.68
Other adult female absent then present	14.08	18.11	0.78
Other adult female present then absent	16.36	12.33	1.33
Other adult female present both times	34.84	21.96	1.59
Stay in wage sector	24.17	12.21	1.98
Stay in piece sector	−3.92	22.02	−0.18
Move from wage to piece	−40.98	21.86	−1.88
Move from wage to self employed	−6.65	15.66	−0.42
Move from piece to wage	42.27	24.76	1.71
Move from piece to self employed	22.00	22.79	0.97
Move from self employed to wage	−41.06	16.37	−2.51
Move from self employed to piece	−43.78	20.99	−2.09
Change in yaya wage (pesos)	−2.38	3.51	−0.68
Change in price of formula (pesos)	2.31	6.06	0.38
Constant	56.41	34.94	1.61
Parameter 1 normalized to zero			
Parameter 2	31.85	15.51	2.05
Parameter 3	−6.24	21.58	−0.29
Parameter 4	−17.07	15.54	−1.10
Parameter 5	−16.63	12.54	−1.33
Parameter 6	10.38	14.45	0.72
Individual heterogeneity (two points of support)			
Parameter 1 normalized to zero			
Parameter 2	−58.16	18.42	−3.16

of about 6 pesos/week for women with one additional children in the interval (true for about 30% of women) to 20 pesos/week with 2 additional children (true for 29% of women), to more than 38 pesos/week for women with 4 or more additional children (true for 6% of sample women). The change in the number of children less than 2 years of age had no significant effect on weekly earnings.

Change in the presence of a spouse or other adult females in the household had no significant effect on earnings. The results for change of sector, on the other hand, were strong. Women who remained in the wage sector had increased earnings of 24 pesos/week while movement out of the wage sector to the piece sector decreased earnings by almost 41 pesos. By contrast, movement from piece to wage increased earnings by over 42 pesos/week. We also see a strong negative effect of moving out of the self employed sector to either the wage or piece sectors. Such a move results in a loss of over 40 pesos/week for either alternative.

The number of additional children born in the 8 year interval had no significant effect on hours worked, but an increase in the number of children

Table 5. Results from selectivity corrected linear regression models of change in hours worked (Mean change in hours worked = 4.9)

Variable	Coefficient	S.D.	*T*-value
1 child in interval	0.83	2.76	0.30
2 children in interval	−0.62	3.08	−0.20
3 children in interval	−1.24	3.48	−0.35
4 children in interval	2.10	5.13	0.41
Change in number of children < 2 yrs	−5.97	1.67	3.57
Spouse absent then present	−1.45	10.73	−0.13
Spouse present then absent	9.40	8.37	1.12
Spouse present both times	4.96	7.37	0.67
Other adult female absent then present	10.78	4.43	2.43
Other adult female present then absent	−0.13	3.05	−0.04
Other adult female present both times	2.95	5.33	0.56
Stay in wage sector	−21.40	5.85	−3.66
Stay in piece sector	−11.94	7.17	−1.66
Move from wage to piece	−24.04	7.10	−3.39
Move from wage to self employed	−24.48	6.45	−3.80
Move from piece to wage	−16.41	6.56	−2.50
Move from piece to self employed	13.81	5.93	2.33
Move from self employed to wage	−18.00	6.57	−2.74
Move from self employed to piece	4.82	7.04	0.68
Change in yaya wage (pesos)	1.00	0.87	1.15
Change in price of formula (pesos)	−0.24	1.52	−0.15
Constant	26.60	11.25	2.36
Standard deviation of error	26.78		
Community heterogeneity (six points of support)			
Parameter 1 normalized to zero			
Parameter 2	1.62	3.64	0.44
Parameter 3	0.46	4.77	0.10
Parameter 4	3.62	3.79	0.95
Parameter 5	−14.01	5.13	−2.73
Parameter 6	6.88	4.35	1.58
Individual heterogeneity (three points of support)			
Parameter 1 normalized to zero			
Parameter 2	−46.34	5.27	−8.80
Parameter 3	−15.54	5.82	−2.67

under the age of 2 had a strong negative effect on hours. Most often, a positive change in the number of children under the age of 2 would represent the case where a woman had a child under age 2 in 1991, but not at baseline. The addition of a child under 2 resulted in nearly 6 fewer hours worked per week. The addition of an adult female household member increased hours worked, but there were no significant effects of spouse's presence. Sector changes were strongly related to changes in hours: Relative to remaining self employed, work hours increased when women moved from piece work to self-employment, but decreased with nearly all other sector changes. This is consistent with the fact that, on average, self-employed women had the longest work hours in 1991.

5. Discussion

Our multi variate analyses of earnings and hours take selectivity into sector of work and endogeneity of childbearing into account through the use of longitudinal data and statistical methods designed to account for these likely sources of bias. For women in the labor force both at baseline and 1991, it is clear that additional children decreased deflated earnings over an 8 year time period. The overall time trend in this sample, which began with a representative cohort of childbearing women in urban and rural communities of Cebu, is toward increasing labor force participation, increased work hours and increased earnings. However, a wide range of variability exists, with about 2/3 of women working at both points in time increasing their weekly earnings between 1983 and 1991. Results confirm the notion of a "child tax": Childbearing acts as a barrier to improvements in earnings. The results are consistent with the recent findings of Angrist and Evans (1998) showing that children decrease female labor supply with each additional birth. However, our results differ from what was reported for the U.S. by Korenmen and Neumark (1992). They found that once endogeneity was controlled in a longitudinal fixed effects model, the effects of childbearing on women's earnings were not significant. We found that the effect of number of additional live births was greater than the effect of having a child under the age of 2. However, the form of the childbearing variables may have influenced this result. The likelihood of increasing the number of children under 2 years of age is increased as the number of children born in the interval increases. The majority of women with 4 or more additional births also had an increase in the number of children under 2 years of age. Thus, the net effect of 4 or more children would be better represented by adding the effect of having an additional child under the age of 2. It is important to note that all children born in the interval were less than 8 years of age in 1991.

In contrast to the earnings results, having a child under the age of 2 years of age significantly decreased women's work hours. This is consistent with the idea that the highest demands for childcare, and greatest difficulties finding alternate childcare givers are associated with having young children, especially those who are still being breast-fed. Thus women may need to adjust their work hours to meet the extra demands of caring for a young child.

We were unable to clearly demonstrate a differential effect of childbearing related to sector of employment: none of the sector by childbearing interactions terms were statistically significant, although the cell sizes for some of the interaction terms became quite small. However, we found strong effects of remaining in the wage sector or moving into this sector on weekly income. We found similarly positive effects for remaining self employed while the piece sector workers tended to be much worse off. The declines in earnings of women who moved from self employment to the wage sector and piece sectors are also worthy of note. The movement from self-employment into the wage sector is associated with a large reduction in the number of hours worked. Women with more young children may make this move in order to reduce their hours, but obtain job stability and benefits which more often accompany wage work. The movement into the piece sector is not significantly associated with a change in work hours. However, piece work, which is most often related to production of crafts, is most often done in the home, and offers flexibility in hours that enhances compatibility with childcare. In addition, having more children may facilitate piece work, since children are often recruited to assist the mother in tasks such as basket-making and shellcraft.

Our methods clearly demonstrated a high level of endogeneity in models of the effects of childbearing and sector of work on women's earnings. To see the

Table 6. Alternate models of women's earnings, 1983 and 1991

Variable	Selectivity corrected[1]		Random effects model[2]		First differences model[3]	
	Coefficient	Standard deviation	Coefficient	Standard deviation	Coefficient	Standard deviation
1 child	2.73	5.92	0.12	6.52	−6.98	10.88
2 children	−7.10	6.42	−5.63	6.84	−21.57	12.12
3 children	−12.54	7.54	−16.20	8.11	−28.80	13.68
4 children	−15.58	12.29	−30.08	14.17	−36.80	19.88
5 or more children	33.37	25.79	10.83	27.54		
Wage sector	6.64	11.61	22.95	7.41		
Self employed	21.20	11.24	32.73	6.90		
Stay in wage sector					30.15	11.81
Stay in piece sector					−47.33	15.85

[1] Regression model for weekly earnings, selectivity-corrected for working for pay.
[2] Random effects model for weekly earnings, ignoring selectivity for working for pay.
[3] First differences model for change in weekly income for women working for pay in both 1983 and 1991. Note that child variables represent children born during the 1983 to 1991 interval.

substantive impact of our methods on the empirical results, we compare the effects of young children on weekly income for 4 estimations:

1. We estimated a model in levels for weekly income as a function of dummy variables for number of children under 7 and sector of employment. The regressions also contained a set of control variables that included the corresponding variables in levels for all variables in Table 4, plus the respondent's age and education as well as the age and level of education of her spouse, plus a dummy to indicate whether the observations was from 1983 or 1991. We controlled for selection into working for pay through the use of a standard Heckman selectivity correction assuming multi variate normality for the error term distribution.[2] In this exercise, we simply pooled data for 1983 and 1991 and so a woman could have reported weekly income at either point in time. The results for the variables of interest are presented in Table 6.

The children dummies specify the number of children younger than 7 years of age. The omitted categories were no children under 7 and working for piece (income is missing for women who are not employed). The results provide weak evidence of a negative effect of young children on weekly earnings except at the extremes which are imprecisely estimated. The results for sector are also weak but we see that women who are self employed have the highest level of earnings.

2. The next estimator is also in levels and simply ignores the selectivity into working for a wage both in 1983 and 1991 and estimates a model in levels with weekly income as the dependent variable for women who where working at both points in time (i.e, the same sample of women who were used in the differenced estimations). Note that each woman contributes 2 observations to the data set.

The results for the children dummies are similar across the 2 sets of runs for 2 and 3 children under 7 years of age, while the effect of 4 children under 7 years of age is more pronounced in the random effects estimations. The sector effects are much stronger in the random effects estimations with working for wage and self employed sectors resulting in substantially higher weekly earnings than the piece sector. This should not be very surprising since we have not controlled for the self selectivity into working at both points in time in these random effects estimations.

3. The third estimation is for change in income between 1983 and 1991 without controls for selectivity into working for a wage at both points in time. The model specification is exactly that of Table 4 except that no heterogeneity parameters are estimated. The model specification is different from 1 and 2 above since all variables are specified as differences. In the results in Table 6, we use dummies for number of children born in the 8 year interval as is done in our main analysis.

It is clear that the effects of small children become much more pronounced when we control for the endogeneity of young children through the use of first differences. The bias towards zero in the effect of young children in the levels models is consistent with positive error correlation in unobservable fixed variables across equations that explain positive weekly earnings and number of small children. A comparison of the effects of the children born in the interval dummies between the table above and Table 4 where we control for the selectivity of sector of work yield very little differences across all dummy categories. This means that the simple difference model appears to have almost fully controlled for the endogeneity of children born in the interval. However, we do see that the coefficients of the sector dummies are strongly affected by controls for self selection of sector of work. This may not be surprising and seems to indicate that unobserved, time varying variables affect sector choice and earnings which means that differencing the data is not sufficient to control for this selectivity.

Finally we note that in another set of studies, we show that childbearing affects other aspects of women's lives (Borja and Adair 1997). After controlling for changes in income, an increased number of births in the 1983–1991 study interval was associated with declines in material well-being measured by the value of household assets, ownership of labor-saving devices such as refrigerators, maternal health, and well being of the index child. By reducing the number of children through use of family planning methods, our models predict that women's earnings would increase. However, it appears that women increase their earnings largely as a function of increasing hours at work. The increase in work hours may exacerbate the conflict between women's productive and reproductive roles. Work hours increased to more than 46 hours/week in 1991. Given that women also report, on average, more than 23 hours per week doing household chores and related activities, their combined market and domestic work burden is substantial. Thus improvements in income may come only at the expense of leisure time or other non-material aspects of their lives. In conclusion, these results show strong negative effects of additional live births on women's weekly earnings, and a strong negative effect of having a child under 2 years of age on work hours, suggesting a strong potential for family planning, if it limits births, to affect women's economic well-being.

Endnotes

[1] Both the presence of the spouse and other adults in the household are potentially endogenous. However, if it is assumed that any correlation is through unobserved, fixed household characteristics, then differencing the data removes this potential source of bias.

[2] In this selectivity model in levels, the same set of variables were used in the working for pay selectivity equation as were used in the earnings equation for those who were working. Identification is only achieved through the non-linearity introduced through the assumption of multivariate normality. As discussed in the introduction, it is very difficult to justify exclusion restrictions in this context and many authors have simply chosen to ignore the endogeneity of the choice to work for pay. Our estimation results were robust to minor differences in the specifications of the two equations and so we feel comfortable presenting these results for comparison to the main results of the paper.

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