Consequences of Family Disruption on Children's Educational Outcomes in Norway

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ABSTRACT

Using high quality data from Norwegian population registers, we examine the relationship between family disruption and children's educational outcomes. We distinguish between disruptions caused by parental divorce and paternal death and, using a simultaneous equation model, pay particular attention to selection bias in the effect of divorce. We also allow for the possibility that disruption may have different effects at different stages of a child's educational career. Our results suggest that selection on time-invariant maternal characteristics is important and works to overstate the effects of divorce on a child's chances of continuing in education. Nevertheless, we find that the experience of marital breakdown during childhood is associated with lower levels of education, and that the effect weakens with the child's age at disruption. The effects of divorce are most pronounced for the transitions during or just beyond the high school level. In models that do not allow for selection, children who experienced a father's death appear less disadvantaged than children whose parents divorced. After controlling for selection, however, differences in the educational qualifications of children from divorced and bereaved families narrow substantially and, at mean ages of disruption, are almost non-existent.

INTRODUCTION

In the latter half of the twentieth century in almost all industrialised countries, increasing proportions of children have experienced a family disruption, with divorce replacing parental death as the main cause. These trends have, in many countries, provoked great concern because correlational evidence suggests that the changes are associated with a heightened risk of disadvantage for the children involved. Although the majority of research on family disruption and its consequences comes from the United States, there is now evidence from a wide range of nations demonstrating that, on average, children who experience a family disruption – a parental divorce, in particular – fare poorly across a wide range of adolescent and adult outcomes, including educational attainment, economic security, and physical and psychological well-being (for reviews see Amato and Keith 1991a,b; McLanahan and Sandefur 1994; Rogers and Pryor 1998; Sigle-Rushton and McLanahan 2004). Although there are many outcomes that have been linked to family structure in childhood, the negative association between family disruption and educational attainment may be especially important. In light of the positive relationship between educational qualifications and subsequent employment and earnings in post-industrial economies, poor educational outcomes may set in place pathways that lead to other kinds of disadvantage. Put simply, low educational attainment may account for persistent differences in adult physical and psychological well-being, relationship quality, and economic well-being later in adulthood.

A better understanding of the complex processes that link family structure to subsequent outcomes is substantively interesting as well as policy relevant. A frequently cited (but difficult to refute) explanation for the association between parental divorce and child outcomes is selection. Proponents of this view argue that the relative disadvantage observed among children whose parents divorce is due to differences between the kinds of people who dissolve their relationships and those people who partner and remain together. Divorce and separation tend to be more common among the lower educated and children from these families have, on average, poorer educational outcomes. Similarly, high parental conflict is associated with both union dissolution and poor child outcomes. Finally people who have problems with substance abuse, violence, mental illness, or other forms of anti-social behaviour are more likely to make poor parents and poor partners. If the negative consequences of parental divorce or separation are mainly due to selection, interpreting the association as causal could lead to unhelpful or even harmful policy interventions. For example, a causal interpretation might lead to the development of interventions aimed at making divorce or separation more difficult when in fact the most effective interventions should be targeted at the selective factors. Indeed keeping high conflict or abusive families together may result in even greater disadvantage for the children involved. For this reason, understanding the role that selection plays is essential to developing programmes that aim to support children and their families.

In this analysis, we use extraordinarily high quality register data from Norway and a simultaneous-equation approach to assess the effects of family disruption on children's educational outcomes. We investigate the effects of both causes of disruption – a parental divorce and a parental death -- but we pay particular attention to the issue of bias due to selection into divorce and following previous research assume that parental death is a more exogenous event. Norway is an interesting country to study for several reasons. Divorce rates have increased quickly and steeply (indeed Norway now has one the highest divorce rates in the world), but in a context where economic inequality remains relatively low (Statistics Norway 2007a; United Nations 2006). The Norwegian economic context is particularly salient because there is a good deal of evidence in countries like the United States that the economic consequences of divorce are responsible for much of the observed differences between children living with two parents and children living with only one parent. If the economic vulnerability that accompanies divorce (and to a lesser extent bereavement) is an important explanatory factor in countries like the United States, we should expect to find a smaller association between divorce and educational attainment in the Norwegian context. Perhaps in line with this thinking, it is noteworthy that there is little political concern in Norway about the large fraction of children who experience a parental disruption[1]. The common perception seems to be that parents will provide sufficient care for their children even if they live apart, and children will benefit from the combined resources of the resident[2] and non-resident parent[3], the possible assistance from other adult family members, the economic support from society[4], and the wide acceptance of this living arrangement. The lack of concern may also reflect a trust in the parents' ability to understand how their children will be affected by a disruption, compared to living with

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[1] For example, the liberal-conservative government 2001-2005 wrote a Family Report to the Parliament in which they expressed some concern about the high number of disruptions, but in a very diffuse and unsubstantiated manner. The red-green government that took over in 2005 did not mention disruption in their declaration of intention (and the Department of Children and Family Affairs was renamed the Department of Children and Gender Equality).

[2] The couple must take part in mediation proceedings with an officially appointed family law expert to obtain a formal separation, which is the first step towards divorce. The decision on with whom the child shall live and visitation rights are the key issues in these proceedings, which are meant to take primarily the child's interest into account.

[3] This other parent is obliged by law to contribute economically, with the amount depending on his or her income. The more time the child spends with that parent, according to formal agreements, the lower the amount.

[4] Special benefits were already given to never-married and widowed mothers in 1965. In 1972, this was extended to divorced mothers, and later to single parents more generally (see, for example, the review by Kjeldstad and Rønsen 2002). From 1981, the most important component has been the so-called transitional allowance, which was meant as temporary support to single parents who could not support themselves. Work requirements were strengthened and the benefit period shortened in 1998. For example, parents of children older than three years must work or study at least half-time or be registered as seeking employment to be eligible for the allowance. In addition, single parents are entitled to a supplementary rate of the universal child allowance (corresponding to having an additional child). They may receive an education or child care benefit, and they have a higher chance than others of getting their child admitted to a public day care centre (provided that they work or study). In most municipalities, there is also a reduced day care price for families with low incomes.

two parents who have a poor relationship, and in their willingness to take that into account when making the decision. However, few attempts have been made to examine statistically the links between family structure and children's outcomes in Norway (for some exceptions see Størkersen et al. (2005) and Breivik and Olweus (2006) who examine a diversity of outcomes including school problems and achievements, and Bratberg and Tjøtta (2007) who study economic effects). More work on these issues has been carried out in Sweden, which has a very similar public welfare system, and results from these studies suggest that selection explains almost all of the differences between children who experience a parental divorce and those who do not (see, for example, Björklund and Sundström 2006; Björklund, Ginther and Sundström 2007; Jonsson and Gähler 1997) suggesting that, consistent with arguments about greater economic equality, selection may play a decisive role in Norway as well. Nonetheless, evidence from comparative studies (see for example Pong, Dronkers, and Hampden-Thompson 2003) shows that Norwegian children whose parents have divorced have lower mathematics and science test scores than children living with both parents, and the mathematics achievement gap is not significantly different from that found in the United States. The current study adds to these findings. The method we employ allows us to identify the association of divorce and educational attainment net of time-invariant unobserved differences between families, allowing us to compare and contrast the effects of different kinds of family disruption with greater confidence. Moreover the method provides us with substantive information on the direction and importance of selection bias that might result when models are estimated assuming that parental divorce is exogenously determined.

EXPLANATIONS FOR WHY DISRUPTION SHOULD BE ASSOCIATED WITH POORER EDUCATIONAL OUTCOMES

In this section we discuss the various reasons that children who experience a family disruption might have poorer average educational outcomes than children raised by two biological parents, highlighting where we might expect to see differences between children who have experienced a divorce and those who have experienced a parental death. Of course, the strength and importance of the various explanations is likely to vary by context, so we additionally draw attention to those explanations that may be more or less relevant in a Scandinavian context. Finally, we assess the extent to which existing empirical evidence provides support for the different explanations.

Apart from selection, there are good reasons to believe that children who experience a family disruption might have lower average educational attainment relative to children in stable, two biological parent families. These explanations stem from changes due to the direct loss of a parent's presence in the same household and from other deleterious changes that might accompany the process of family disruption. The presence of the same sex parent teaches young children appropriate gendered behaviour so that father absence, *particularly at younger ages* and *particularly for boys*, affects development and child outcomes. Because parents are important role models as well as providers of social capital for their children, the absence of a parent, especially when that parent will be likely to instil values concerning the importance of an education, can be detrimental to a child's educational career. Moreover, exposure to the interactions of two parents (in a healthy relationship) helps children to develop interpersonal skills such as communication, cooperation, and conflict resolution. Children whose exposure to parental role modelling is more limited may lack the skills they need to function as healthy adults, and as a consequence, they may be less successful in school, at work, or in their own personal relationships. Nonetheless, socialization and social capital deficits can be offset when non-resident fathers are actively involved in their children's lives. Thus, holding everything else constant, we might expect to find better educational outcomes among children who maintained a (strong) relationship with their (highly educated) non-resident fathers. Children whose fathers died might have worse outcomes than otherwise similar children who benefit from high levels of father involvement. Of course, other adults may be able to substitute for the lost resources of the non-resident father, and therefore we would expect children with alternative role models and sources of support, be it a step-parent, male relative or family friend, to have better outcomes, particularly when the other sources of support have higher levels of education and can reinforce the importance of educational attainment, than children who lack these substitutes.

In addition to the loss of a role model, disruption can lead to subsequent changes that are associated with a heightened risk of disadvantage. Three of the most important disadvantages that are discussed in the literature include: economic vulnerability (which should but does not often include time poverty), inadequate parenting and reduced resilience due to acute stress. In the United States, economic hardship appears to lower food consumption affecting children's nutrition and development (Page and Stevens 2004). Similarly, reduced economic circumstances could result in moves to more deprived areas with lower quality schools and other neighbourhood effects which may have negative consequences for children's educational attainment. Although such dramatic economic effects would be unlikely to occur in a country like Norway, less dramatic drops in income and more moderate forms of economic deprivation could lead to less investment in educational resources for children such as enrichment activities, books, and computers. Because family change can be emotionally draining, most families undergoing divorce will be under a good deal of stress, as will those coping with a long term illness preceding death or dealing with the emotional and practical changes that accompany an accidental death. High levels of stress may undermine parenting capacities. In addition to being stressful themselves, family transitions are often accompanied by other events that produce stress such as moving house, changing schools, and parental dating/re-partnering. Finally the stress of the disruption may make children less resilient in coping with additional changes. Even if the effects of stress are short lived, if the stress occurs at crucial periods in the life course, such as transitions between different levels of education, the disruptive effect could have long-term repercussions. To the extent that families are better supported, both economically and socially, following a bereavement than following a divorce, we might expect any family disruption to increase the risk of disadvantage, but for children who experience a parental death to fare somewhat better than children who experience a divorce. The gap between the two may be less severe in Norway where the welfare state works to mitigate the economic consequences of family disruption without differentiating, to a great extent, between causes of disruption as happens in other countries like the United States (Biblarz and Gottaine, 2000). Nonetheless, although stress is likely to follow both disruption events, relative to a parental divorce where the parents may continue to engage in conflict and where some social ties to wider kin networks may be strained, the level of stress may be lower and the level of support offered may be greater following a bereavement. Hence the experience of death or divorce and its associated effects may be qualitatively different.

As previous analyses of the evidence have concluded, no one theory of the association between divorce and children's education receives consistent and conclusive empirical support. There is some evidence that the loss of a resident parent has important educational consequences for children (Lang and Zagorsky 2001), the effects of which are mediated by parental-like inputs from other adults, but not stepparents (Amato and Keith 1991a; Thompson, Hanson, and McLanahan 1994; Painter and Levine 1999; DeLeire and Kalil 2002; McLanahan and Sandefur 1994). The findings for stepparent families are worthy of further scrutiny, because, on average, these families have higher incomes and more time available for parenting than do single-mother families. Theorists can invoke sociological (Cherlin 1978), psychological (Hetherington, Bridges, and Insabella 1998; Rogers and Pryor 1998) or evolutionary explanations for this apparent inconsistency, but these require further elaboration and verification (Biblarz and Raftery 1999; Case, Lin, and McLanahan 2001). More troubling, because families make the choice to divorce or remarry, the predominance of selection bias always remains a strong possibility.

Consistent with socialization deficits are findings by Krein and Beller (1988) that father absence during the preschool years had the largest negative effects on years of educational attainment for black and white individuals of both sexes. Comparing siblings, Ermisch and Fransconi (2001) found that for the UK disruptions before age five were more strongly associated with negative outcomes later in life, such as low academic achievement, economic activity, and smoking, than later disruptions. Using a similar modelling strategy and data drawn from a US study, Yeung, Duncan and Hill (1995) failed to find support for the early absence hypothesis, however. In addition, there is also little evidence of significant differences between boys and girls (Sigle-Rushton, Kiernan, and Hobcraft 2005). Overall, there is only partial support for theories related to the direct loss of the parent figure from the household.

There is fairly strong evidence that the psychological adjustment of the custodial parent is associated with better child outcomes. Children of depressed mothers have higher externalizing and internalizing problems than children whose mothers are not depressed (Downey and Coyne 1990; Covey and Tam 1990) which may be related to school performance and longer-term educational achievement (Huerta and Sigle-Rushton 2007). Insofar as mothers suffer short-term depression and anxiety before, during and subsequent to a divorce or the death of a spouse, mothers' psychological adjustment might be an important mediating factor (Heatherington et al 1982). However, few studies have found evidence of such a relationship. Empirical studies have also documented differences in the parenting practices of single-mother and two-parent families. Controlling for differences in parenting does little to attenuate the effects of disruption, however (Brown 2006). For example, Astone and McLanahan (1991) found, using data from the United States, that controlling for parental inputs did little to narrow the difference in the risk of dropping out of high school between children in intact and non-intact families. In another study, analysing data from the United States, Thomson et al. (1994) found that parenting practices accounted for practically none of the disadvantage associated with living in a single-mother family, but between 13 to 35 percent of the disadvantage associated with living in a stepfather or mother-partner family.

There is clear evidence that differences in the educational attainments of children in single-mother and two-parent families are far less pronounced once income is held constant (Amato and Keith 1991a). Indeed, previous studies have suggested that in the United States income differentials account for between 30 and 50 percent of the gaps in rates of high school graduation among children living in one- and two-parent families (McLanahan 1985; McLanahan and Bumpass 1988; McLanahan and Sandefur 1994). The effects of family structure on children's educational outcomes frequently attenuate after income is controlled, but in many cases differences by family structure remain significant. Painter and Levine (1999) present some evidence that measurement error reduces the effectiveness of income as a mediating variable in mother-only families in the United States. Even with improved measures of income, there are additional problems with testing the importance of income changes that accompany a family disruption. Most researchers hold income constant but do not hold time and parenting inputs constant. Mothers with higher incomes are likely to be devoting less time to childcare, especially in contexts where the gender gap in earnings is high. The time deficit will be particularly important in contexts where access to good quality alternative forms of childcare is limited such as the United States and Britain. Under these circumstances, an income effect estimated using an undifferentiated measure of income will underestimate the extent to which income transfers, child support, or higher wages improve children's well-being. These other sources of income do not require time inputs and might possibly have larger and more positive effects for children than income from labour market earnings. Because we only have access to a measure of gross labour market earnings and because income is an endogenous variable which is affected by parents' human capital and by the amount of time they devote to the labour market, we do not control for income in our models. However, we hope to learn something about income by comparing our findings to those obtained from other countries. If income substantially mediates the effect of family structure on educational attainment, we would expect the effects of divorce, particularly after controlling for some sources of selection, to be much less negative in a country like Norway which has far less income inequality, greater levels of support to mother-only families, and a generously subsidised public school system with no tuition fees[5], relatively generous loans to students (from the Norwegian State Educational Load Fund), and grants to students living away from their home.[6]

METHODS

Alternative Approaches to Dealing with Selection into Divorce

Without running an experiment in which children were randomly assigned to different kinds of families, it is impossible to ascertain how children *would have* fared if their parents had not divorced. However, a range of second-best strategies have been developed to try to deal with selection and to better identify causal associations.

One strategy for isolating the direct effect of parental divorce has been to exploit longitudinal study designs and control as much as possible for pre-divorce circumstances. In some instances, measures of children's well-being prior to divorce are included in the models, while in other cases, measures of parental conflict and other characteristics of the family environment (low income, poor parenting) are included. In general, when pre-divorce characteristics are included, the associations between parental divorce and child outcomes become smaller, and in some cases, they become insignificant. Sometimes important moderating factors are identified. For example, using data

[5] There are only a few private schools, largely at the college level, which charge fees.

[6] Those who study abroad get supplementary grants and loans to cover tuition fees, as well as travel and living expenses.

from the United States, Jekielek (1998) shows that in high-conflict families, children whose parents divorce fare better than those whose parents stay together, but for low-conflict families the children whose parents divorce fare worse. In some instances, associations that become insignificant at one stage of childhood, reappear when outcomes at later stages of the life course are considered (Sigle-Rushton, Hobcraft, and Kiernan 2005). Although our data contain a good deal of high quality longitudinal information for a large, representative sample of individuals, the data we use are administrative so we do not have rich measures of the characteristics of the families and children that comprise our sample. Importantly, we have no information on conflict or parenting practices, so it is unlikely that the set of variables we have available would control sufficiently for the selective factors that might bias the association between divorce and educational outcomes.

Another approach compares children living in places where divorce is more easily available to those who live in places where obtaining a divorce is more difficult. Often these have been US studies where differences across states are exploited. Employing this sort of method, Gruber (2004) finds that living in a state with unilateral divorce laws is associated with less education, more dropping out of high school, more early marriage and more divorce. In some instances the effects differ between men and women but, in all cases, the association for at least one sex is significant. Interpreting the negative associations as causal is problematic however. Changes in divorce law may have altered bargaining power within families and these changes may have had implications for the children involved. If this is the case, negative outcomes associated with different divorce laws might not be due to increases in divorce but rather to changes affecting parental obligations to children. Even if these problems could be surmounted, this approach is not suitable for our purposes because there is no within-country variation in divorce laws in Norway.

A final approach to dealing with selection is to compare children who share parents (or a parent) but whose exposure to different family structures differs. Usually this is operationalised as time spent in different family types (see for example, Ermisch and Francesconi 2001; Hao and Xie 2001; Ginther and Pollak 2004; Björklund and Sundström 2006). For example, if parents separate when one child is ten years old and the other child is five, the older sibling can be said to have experienced eight years of father absence before the age of 18 while the other can be said to have experienced 13. This approach (which can be extended to examining step-families as well) allows researchers to include family- or parent-specific fixed effects but assumes that the longer a child lives in a father-absent family the greater the negative effect. Moreover, sibling models assume that parents treat their children exactly the same and that children respond similarly to family-wide risk influences, both of which are highly unlikely (Carbonneau et al 2002; Jenkins, Rasbash, and O'Connor 2003). It is possible that some parents wait to divorce until the oldest child leaves home (or is older) precisely because they believe the child will be harmed by the divorce. If two children in such a family have similar outcomes this does not necessarily mean that divorce would not have had a negative impact on the older child (for further discussion, see Ermisch and Francesconi 2001). There are good reasons for researchers wanting to employ this sort of method to proceed cautiously. When the outcome is discrete, fixed effects models are only identified from the sub-sample of families in which there is within-family variation in the outcome being considered. Because in our application educational outcomes are measured as a series of discrete variables, we had concerns about employing this method.

A Simultaneous Equation Model of Children's Education and Marital Dissolution

In this paper we model children's educational attainment jointly with mothers' risk of marital dissolution using a multilevel multiprocess (simultaneous equation) model (for further details of multiprocess model see, for example, Lillard and Waite 1993). This approach is similar in some ways to the sibling models (and suffers from some of the same limitations) but imposes fewer theoretical assumptions about the age effects of divorce (that the longer a child has been exposed to divorce, the more disadvantage they experience) and overcomes some of the identification problems associated with fixed effects sibling models. A multilevel model is used because a mother may have more than one child and more than one marriage over the course of the observation period. We therefore allow for the presence of unobserved woman-specific factors that affect the educational outcomes of siblings with the same mother and the dissolution risk of marriages formed by the same woman. The model assumes that, conditional on covariates, any remaining selection effect of disruption on children's education is due to correlation between unobserved time-invariant characteristics of the mother that affect the probability of continuing in education and the risk of marital dissolution. In each process these unmeasured characteristics are represented by woman-specific random effects, which may be correlated across processes to allow for the possibility of shared unobserved risk factors. We do not control for unmeasured characteristics that are time-varying or child-specific. Thus our random effects model makes assumptions of within-family similarity that are similar to those made in fixed effects models.

The multiprocess model is a system of simultaneous equations for women's marriage durations and children's educational outcomes. The duration to parental divorce is modelled using an event history model, and the highest educational qualification achieved is modelled using a sequential probit model. Similar to previous research, we assume that parental death is exogenous. A parental death censors the duration to marital dissolution and is introduced as an independent variable in the education process.

Model for Mother's Marital Dissolution

Denote by $h_{ij}(t)$ the hazard of marital dissolution in year *t* of marriage *i* of woman *j*. A multilevel continuous-time event history model allowing for unobserved heterogeneity between women may be written

$$\log h_{ij}(t) = f(t) + \boldsymbol{\alpha}^T \mathbf{w}_{ij}(t) + v_j. \qquad (1)$$

The log-hazard of dissolution is assumed to depend on the marriage duration at year t through a function $f(t)$, the baseline log-hazard rate. Common choices for $f(t)$ include linear or quadratic functions or a piecewise-constant formulation. Here, we assume that $f(t)$ is a piecewise-linear spline with nodes spaced at bi-annual intervals up to 10 years and the slopes within intervals denoted by a vector $\boldsymbol{\delta}$:

$$\begin{aligned} f(t) = \delta_0 &+ \delta_1 \min[t,2] + \delta_2 \max[0, \min[t-2,2]] + \delta_3 \max[0, \min[t-4,2]] \\ &+ \delta_4 \max[0, \min[t-6,2]] + \delta_5 \max[0, \min[t-8,2]] + \delta_6 \max[0, t-10]. \end{aligned}$$

Covariates $\mathbf{w}_{ij}(t)$ may be time-varying or characteristics of a particular marriage (e.g. the education or age of spouse *i*). Time-invariant unobservables that affect a woman's risk of separation from any marital partner are represented by the random effects v_j, which are assumed to follow a normal distribution with mean zero and variance σ_v^2. The above specification is a proportional hazards model that assumes elements of $\mathbf{w}_{ij}(t)$ have the same effect on the risk of dissolution at each duration *t*.

Model for Children's Educational Qualifications

Children's educational qualifications are modelled using a sequential probit model. Sequential probit models are appropriate for ordered outcomes that can be viewed as the result of a sequence of decisions or transitions, and are widely used to study educational progression (see, for example, Brien and Lillard (1994), and Upchurch, Lillard and Panis (2002)). Here, the educational outcome used is the highest qualification achieved by the end of the observation period, which is measured by a ordered categorical variable with five levels (described in the Data section below).

Denote by y_{ij} the level of education achieved by child i of woman j. We view the observed y_{ij} as the result of a sequence of binary transitions made from one education level to the next. For five levels, a child can make up to four sequential transitions, where the transition from level r ($r = 1,\ldots,4$) is possible only for those for whom we observe $y_{ij} \geq r$. Specifically, the sequence of transitions is as follows: (i) level 1 to 2, (ii) level 2 to 3 (conditional on having attained level 2), (iii) level 3 to 4 (conditional on level 3), and finally (iv) level 4 to 5 (conditional on level 4). The transition from level r is indicated by a binary variable $y_{ij}^{(r)}$, coded 1 if the child continues to and completes level $r+1$, and 0 if he or she stops at r. These binary transition indicators are constructed from the qualification level achieved by the end of observation, y_{ij}. Thus, for example, a child who has either attained level 3 or went on to but did not (yet) complete level 4 will contribute three binary responses $(y_{ij}^{(1)}, y_{ij}^{(2)}, y_{ij}^{(3)})$ = (1, 1, 0) and a child who achieved the highest level 5 will have response vector (1, 1, 1, 1). One advantage of breaking down y_{ij} into a set of binary transitions is that children who are still in education at the end of observation period, and whose educational career is therefore censored, can be retained in the analysis and will contribute $y_{ij}^{(r)}$ up to the level they are observed[7].

The sequential probit model is defined in terms of a set of continuous latent variables or propensities, $y_{ij}^{(r)*}$ underlying the observed binary responses $y_{ij}^{(r)}$, where $y_{ij}^{(r)}$=1 if $y_{ij}^{(r)*}$>0 and

[7] Another way of viewing educational progression is as a type of discrete-time event history process where 'time' corresponds to an educational level and the 'event' is leaving education. Individuals who have not yet experienced the event of interest are right-censored.

$y_{ij}^{(r)}$=0 otherwise. A multilevel model that allows for unobserved heterogeneity between mothers can be written:

$$y_{ij}^{(r)*} = \boldsymbol{\gamma}^{(r)^T} \mathbf{z}_{ij}^{(r)} + \boldsymbol{\beta}^{(r)^T} \mathbf{x}_{ij}^{(r)} + \lambda^{(r)} u_j + e_{ij}^{(r)},\ r = 1,...,4 \qquad (2)$$

where $\mathbf{z}_{ij}^{(r)}$ is a vector of potentially endogenous indicators of family disruption with coefficients $\boldsymbol{\gamma}^{(r)}$, and $\mathbf{x}_{ij}^{(r)}$ is a vector of background characteristics of the child and the mother with coefficients $\boldsymbol{\beta}^{(r)}$. The model also includes two random terms: a mother-specific random effect u_j representing unobserved characteristics of the mother that affect the probability of attaining level r ($r = 1,...,4$) for each of her children, and transition-specific residuals $e_{ij}^{(r)}$. The random effects have transition-specific coefficients or 'loadings' $\lambda^{(r)}$. Thus, although the same unmeasured mother characteristics are assumed to influence progression at all levels of education, the variance of the mother-specific effect, is allowed to differ across transitions. In fact, we assume that the u_j follow a normal distribution with mean zero and variance σ_u^2[8], and $e_{ij}^{(r)}$ are independently and identically distributed N(0,1) random variables. The presence of u_j in each equation means that the four equations must be estimated simultaneously.

The sequential probit model (and its logit counterpart, the continuation odds model) is extremely flexible. The r superscript on $\mathbf{z}$ and $\mathbf{x}$, and their coefficients $\boldsymbol{\gamma}$ and $\boldsymbol{\beta}$, allow both the values and effects of the disruption indicators and background characteristics to vary across transitions. Therefore, in theory, it is possible to have different sets of explanatory variables affecting progression to successive levels. Furthermore, rather than focusing on one specific educational

[8] For identification, some constraint must be placed on $\lambda^{(r)}$ or the variance σ_u^2 in order to fix the scale of the random effect. Common choices are to fix one of the loadings to 1 or to fix the variance to 1.

transition – such as graduating from high school or graduating from college – this model specification allows us to examine the individual educational career as a whole and to identify those stages of the educational system where the variables of interest have stronger and weaker associations.

Equations (1) and (2) together define a multilevel multiprocess model. The equations are linked in two ways. First, the family disruption indicators $\mathbf{z}_{ij}$ in (2) are prior outcomes of the marital dissolution process in (1). Second, we allow for the possibility of a non-zero correlation between the unmeasured woman-specific components u_j and v_j. Specifically, u_j and v_j are assumed to follow a bivariate normal distribution with correlation ρ_{uv}. A value of ρ_{uv} that is significantly different from zero would suggest that at least one element of $\mathbf{z}_{ij}$ is endogenous with respect to educational transitions.[9]

Estimation

The presence of u_j in all four educational transition equations in (2), and the correlation between u_j and v_j, mean that the equations in (1) and (2) must be estimated simultaneously. The multilevel multiprocess model can be estimated via maximum likelihood as follows. For each process the likelihood conditional on the woman-specific random effects is derived. Conditional on (u_j, v_j) and covariates ($\mathbf{z}_{ij}^{(r)}, \mathbf{x}_{ij}^{(r)}, \mathbf{w}_{ij}(t)$), the probability of continuing in education and the hazard of marital disruption are assumed to be independent. Thus the joint conditional likelihood is the product of the conditional likelihoods for the two processes. The marginal (unconditional)

[9] The loadings $\lambda^{(r)}$ have no effect on the correlation between mother specific unobservables in (1) and (2). In other words, the value of ρ_{uv} does not vary across the transitions.

likelihoods are not independent, with correlation induced by the random effects. The joint marginal likelihood (ignoring subscripts) is obtained by 'integrating out' the random effects as follows:

$$L(\boldsymbol{\delta},\boldsymbol{\alpha},\boldsymbol{\gamma}^{(r)},\boldsymbol{\beta}^{(r)},\lambda^{(r)},\sigma_v,\sigma_u,\rho_{uv}) = \int_v \int_u LL(\boldsymbol{\delta},\boldsymbol{\alpha} \mid v) L(\boldsymbol{\gamma}^{(r)},\boldsymbol{\beta}^{(r)},\lambda^{(r)} \mid u) g(v,u)\, du\, dv$$

where $g(u,v)$ is a bivariate normal density. The integration is performed numerically using Gauss-Hermite Quadrature and the joint marginal likelihood is then maximised to obtain estimates of the unknown parameters.

All estimation was carried out using the *aML* software (Lillard and Panis, 2003), which specialises in fitting multilevel multiprocess models.

Identification

In order to estimate a simultaneous equations model it is usually necessary to impose some identification conditions on the exogenous covariates in the model, represented by $\mathbf{x}_{ij}^{(r)}$ and $\mathbf{w}_{ij}(t)$ in (1) and (2) above. Here, this would involve identifying at least one variable (an instrument) that on theoretical grounds affects the risk of dissolution but not educational transitions, i.e. $\mathbf{w}_{ij}(t)$ should contain one or more elements that are not included in $\mathbf{x}_{ij}^{(r)}$. In the present case, however, exclusion restrictions are not required for model identification due to the presence in the sample of women with more than one child and more than one marriage. After accounting for the mother-level random effects and their correlation, the remaining variation in $\mathbf{z}_{ij}^{(r)}$ between siblings represents the effect of experiencing disruption on the probability of progressing to the next level of education adjusting for selection on mother-specific unobservables. Nevertheless, the dissolution equation does contain several variables that are not included in the educational transition equations.

In particular, age heterogamy is considered as a predictor of marital dissolution, but not of children's educational outcomes. Omitting this variable from the analysis does not affect the estimate of ρ_{uv} nor the estimates of the disruption effects $\gamma^{(r)}$.

DATA

Norwegian Registry Data

Our analysis is based on data up to 2003 from the Norwegian Population Register, the Population Censuses, and Statistics Norway's Educational Registration System. These data were collected for everyone living in Norway after 1960[10]. For each person, there is information about their marital status on 1 January each year from 1974 and the spouse's identification code (which allows information on spouses to be merged), the highest educational level achieved as of 1 October in 1970 and annually from 1980, municipality of residence for every month since 1964, dates of any in- and out-migration, and dates of death. Parents are identified for almost everyone born in Norway after 1953. A series of record linkages, based on these links between parents and children and between spouses, were performed to build up an analysis file from the raw data.

The analysis file includes women who were born in 1935-70, who married for the first time after 1973, whose oldest child born after 1974 had turned 16 by 2003, and who did not live outside Norway when any of their children were aged less than 16 years. For each of these women, there is information on i) their education and year of birth/death/emigration, ii) the dates of birth, education

[10] For each person, there is a unique anonymised identification code which enables linkage of marital partners and parents with their children.

in 2003, and father's education for up to five children born in 1974-1987, and iii) the start and duration of up to five of their marriages and the education of each spouse.

Let us explain these restrictions. First, it is obviously meaningful to consider only children born 1974-1987. This is because the full marital history of the parents is not known for those born earlier, and because those born later have not had a chance to take any education beyond the compulsory level. The mothers of these children are born between about 1930 and 1970, but we nevertheless restrict the analysis to mothers born 1935-1970 (leaving out 0.02% of the mothers of children born 1974-87), because those born before 1935 may have incomplete birth histories (births in the 1940s or early 1950s that are not registered). The restriction to five children is purely a matter of convenience, but it leads to the omission of only 0.7% of births in 1974-1987. The reasons for the marriage restrictions are that: i) we do not know the time of marriage for those who were already married in 1974, ii) we do not know the number and length of earlier marriages for those who married as widows or divorcees after 1974, and iii) children of never-married women cannot experience any marital disruption. Finally, mothers who lived outside Norway when any of their children were aged less than 16 years were excluded because information on changes in marital status taking place abroad may be incomplete, and there may be missing information on men who fathered a child abroad but did not accompany the mother and child back to Norway. Fathers were identified for 97.7% of the children, but for simplicity we did not collect information for the 0.01% of the fathers born before 1915.

The analysis is based on a 50% random sample of the linked files. This sample contains 200498 children from 113980 mothers who had 129189 marriages.

Outcome Variables

Marriage durations. The marital history data contain the years in which marriages start and end, but not the exact dates. If the marital status is 'never-married' on 1 January of year t and 'married' on 1 January of year t+1, we can infer only that the couple married at some point during year t. Similarly, if a status 'separated', 'divorced', or 'widowed' is seen for the first time on 1 January of year t^*, we know that a disruption took place some time during year t^*-1. In that case, the duration is set to t^*-t-1. If no such disruption is indicated, the duration is of course 2003-t. The duration of marriages ending in the death of the husband are treated as censored at the time of death, under the assumption that such marriages would have continued had the spouse survived.

Children's education. Educational level was coded according to the 2000 standard (Statistics Norway 2001). The levels mentioned above are as follows. 1: compulsory education only (10 years of schooling), 2: lower secondary education (11-12 years), 3: higher secondary education (13 years), 4: some college or university education, up to and including a Bachelor's degree (14-17 years)[11], and 5: all college education taking 5 or more years, for example a Master's degree (18 or more years). The small proportion of children with unknown education or who had not yet completed compulsory school were omitted from the analysis.

The registration system does not give a full picture of the upward movement through the educational levels 1-5. Until the mid-1990s, those who set out to take a theoretical higher-secondary education were recorded as having no more than compulsory education until they had earned their degree. If they dropped out, they would remain registered with the compulsory level. Students on a vocational track, however, were registered as passing through a lower-secondary level. This means that some of those registered with only compulsory level actually have taken some secondary education, which may have added to their real qualifications. Similarly, some students are registered

[11] This category is a very broad one: passing a one-semester course at a college may be all that is needed to be placed here, depending on the nature of the course. The category also includes a small subgroup of people who have taken post-secondary courses that are not part of the requirements for a college degree.

as passing directly from a higher secondary-level to the equivalent of a Master's degree. This is most common for educations that do not consist of one- or two-semester modules.

The following age constraints were placed on the analysis samples for estimation of the four equations in the sequential probit model: i) the probability of progressing from compulsory to lower secondary school is estimated for children aged 16+ years, ii) lower to higher secondary is estimated for age 18+ years, iii) higher secondary to Bachelors degree for age 19+ years, and iv) Bachelors to Masters or a higher degree for age 23+ years. The age restrictions are set taking into account the age that someone would be typically expected to complete each educational level. A rather young lower age restriction is used for the estimation of the equation for the third transition because level 4 captures a wide range of university courses, from a short course to a full Bachelor's degree. However, it is important to note that the educational system in Norway is unique relative to many other countries. The educational system is extremely flexible in the sense that students generally have good opportunities to change from one track to another, exit and return several times, and study part-time. As a result, many spend more years as formally enrolled in school to reach these levels than indicated by the numbers in parentheses, which indicate the number of years that would be needed according to the stipulated 'normal progress'. For this reason, the age constraints we employ mean that the effects of disruption will reflect differential educational attainment as well as differences in pathways to the same level of attainment.

Indicators of Family Disruption

From the marital and birth history data, it is possible to create a range of variables indicating a child's experience of family disruption up to the year of their sixteenth birthday. Marital dissolution is represented by a dummy variable indicating whether the biological parents separated and the child's age at the time of the separation. Children whose parents did not marry during this period (but whose mother married someone other than the biological father between 1974 and 2003) are indicated by a dummy variable and coded zero on the disruption indicator. Similarly, children whose parents did marry but who did not separate are coded zero on the age at disruption variable. Thus, the coefficients of the divorce indicator and age at dissolution are respectively interpreted as the effect of divorce among children whose parents married, and the effect of the age at disruption among children who experienced a parental divorce before age 16. A corresponding set of variables were created to indicate disruption following the father's death (whether disruption was due to death and the child's age at that time), again defined only for children whose parents married.

For ten percent of children who experienced marital breakdown, their parents subsequently reunited meaning that the mother was coded as formally separated from and later married again to the same person. These annulled separations are treated as disruptions in the analysis, but are indicated by a dummy variable to explore whether a reunion reduces any negative impact of the separation. Finally two variables are defined to indicate whether a child had a stepfather and, if so, whether that marriage broke up. For children whose parents married before age 16 the dissolution of a marriage to a stepfather represents a repeated experience of disruption, while it could be the first disruption for children whose biological parents did not marry.

Other Covariates

Model for marital disruption. The event history model for marital disruption includes a range of explanatory variables specific to marriage i: an indicator of previous marriage, the presence of children from a previous relationship (for both the woman and spouse i), an indicator of whether the woman had a pre-marital birth with spouse i, a time-varying count of the number of children fathered by spouse i, the woman's age at marriage and the age difference between the woman and her spouse, and the education level of each partner. Table 1 shows descriptive statistics for variables included in the final disruption equation.

Model for educational transitions. The covariates of prime interest in the analysis of children's educational transitions are the various indicators of family disruption. We also control for characteristics of the child and the family, including the child's sex, their age in 2003, the number of older and younger siblings and the highest level of education achieved by each of the parents. (See Table 2 for descriptive statistics for the selected variables.)

Although each of the four sequential probit equations is estimated for the subsample of children who are eligible to make a given transition, based on their age in 2003 and whether they have achieved the previous level in the sequence, an additional control for age is included to allow for between-child variation in the timing of transitions and the fact that the probability of making a transition increases with age.

The number of siblings present at the time a child is completing compulsory education is included as an indicator of competition for household resources, although Black, Devereux and Salvanes (2005) showed that family size may have little impact on educational outcomes in Norway. The number of younger siblings is defined at age 17 rather than 16 to allow for pregnancies.

Based on evidence from previous studies (d'Addio, 2006), we expect parental education to be a strong predictor of children's education. Because mother's and father's education are likely to be highly correlated, we represent parent's education by a composite categorical variable (following Lyngstad 2004). For each parent, the five categories considered for children's education are grouped into three (see note to Table 1). Each cell in the 3×3 cross tabulation of mother's and father's education forms a category in the composite variable, leading to nine categories. Such a parameterisation allows us to explore whether, on average, one parent's education makes a dominant contribution to the child's educational choices as well as possible interactive effects of mother's and father's education. For example, does an increase in father's education have the same effect for different levels of maternal education?

RESULTS

Model Selection

The final model specifications were reached after preliminary analysis in which random effects models were fitted separately for marital dissolution and children's education, i.e. the single process models represented by (1) and (2) with zero correlation assumed between the mother-level random effects u_j and v_j. Initial analysis revealed large differences in the effects of covariates on the probability of continuing in education depending on the type of a transition. Because these differences included the effects of some of the family disruption indicators, all subsequent models included transition-specific effects.

In addition to the covariates listed in Tables 1 and 2, the main region of residence and the population size of the municipality of residence were considered as predictors of both dissolution and educational transitions. Region and municipality of residence were defined at the birth of the oldest child in the dissolution model and at age seven in the model for children's education. Their inclusion did not affect the other parameter estimates in either model and were therefore omitted from the subsequent analysis.

In preliminary analyses, age at disruption was treated as a continuous variable with linear effects on each educational transition and different age effects according to whether the disruption was the result of divorce or the father's death. Polynomial functions of age were also considered, but the coefficients of higher order terms were not significantly different from zero. The nature of the age dependence was further explored by treating age as categorical, with dummy variables created for each year. We found that age at disruption effects are fairly monotonic for separation, while there are no significant effects of age among children whose father died. Furthermore, the effects of other covariates in the model were unaffected by the specification of age at disruption effects. We therefore fitted linear age effects in all subsequent analyses. The child's age at the time of death was not found to be significant for any transition or any of the different age specification and is therefore not included in the models presented below.

The child's age in 2003 was included in the model for educational transitions to allow for differences in children's exposure to the probability of making a transition and individual variation in the timing of transitions. In each equation of the sequential probit model, age effects are modelled as a step function, where the age categories used differ across transitions (see Table 6).

As described earlier, parental education in the model for children's education, and husband and wife's education in the model for marital dissolution, is measured by a single categorical variable which combines the educational level of both parents or partners. We therefore allow for an interaction effect of mother's and father's education, and of husband's and wife's education. In both cases, the combined variable was found to be a significantly better fit to the data than two separate variables allowing only for main effects[12].

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[12] Two sequential probit models were fitted and compared using a likelihood ratio test: a model with main effects of maternal and paternal education, and the extended model with an interaction effect (using the combined variable described in Table 1). There is strong evidence of an interaction between mother's and father's education (LR=61, d.f.=16, p<0.001). The same approach was used to compare a main effects model and an interaction model for the effects of a woman's and her husband's education on their risk of separation. The interaction effect was found to be strong and statistically significant (LR=44.52, d.f.=4, p<0.001).

Because previous research suggests that girls and boys may respond to family disruptions in different ways (see, for example, Jekielek, 1998), interaction effects between gender and selected indicators of family disruption were considered. Specifically we tested whether the experience of a parental separation or death and the age at either type of disruption affect boys and girls differently. Interactions between gender and the child's age at separation were also considered but none were significant so they are not included in the models presented below.

Predictors of Marital Dissolution

We begin with a discussion of the results for marital dissolution. The parameter estimates from this model are given in Table 3. All coefficients are in the expected directions and are broadly in line with the findings of previous research suggesting that the risk of divorce is higher amongst more socio-economically disadvantaged couples (Kiernan and Mueller, 1999; Clarke and Berrington, 1999; Lyngstad 2004). A woman's age at the start of the marriage and the age difference between her and her spouse are both important predictors of the risk of divorce: a young age at marriage and age heterogamy, especially when the woman is the older partner, are associated with shorter marriages. There is also evidence of an effect of educational heterogamy. Regardless of the woman's education level, having an educated husband is linked to a reduced risk of disruption. However, marriages of highly educated women to men with a low level of education are the most likely to be dissolved. Because men tend to work longer hours, to earn higher hourly wages, and to earn higher wages as they get older, the findings suggest that economically advantaged families are over-represented among the population of families that do not experience a divorce.

Consistent with previous studies, second marriages are at higher risk of dissolution, and the risk is further increased if either the woman or her husband has children from a previous relationship. The presence and number of children fathered by the current spouse, however, are associated with longer marriage durations. Taken together the results from the model of marriage disruption suggest that there may be important differences between families that experience a divorce and those who do not. To the extent that some of the differences are not measured and controlled for in models of educational attainment, parameter estimates for parental divorce will be biased.

Effects of Parental Divorce on Children's Educational Transitions: Evidence for Selection

Estimates of the parameters associated with the mother-specific random effects in the simultaneous equation model are given in Table 4. These results suggest that there is significant unobserved heterogeneity both between mothers in the chance that their children continue in education and between women in their risk of marital dissolution. Furthermore, there is residual correlation at the mother level between a child's propensity to continue in education and a mother's risk of divorce which is reflected in the estimate of the correlation between the random effects for the education and dissolution equations, ρ_{uv}. A correlation that is significantly different from zero provides evidence that dissolution is endogenous, in which case the estimated effects from the single process model will be biased. We find that the correlation is significantly different from zero (LR = 336.1, d.f. = 1, p < 0.001), and therefore conclude that there is a moderate, negative association between the unmeasured mother-level determinants of children's educational transitions and the mother's risk of divorce. The children of women with an above-average risk of dissolution (v_j>0) tend to have below-average chances of continuing in education (u_j<0). In other words, unobserved factors cause the association between parental divorce and educational attainment to be biased upward in models that treat parental divorce as an exogenous variable.

The estimates of the random effect loadings $\lambda^{(r)}$ suggest that these unmeasured mother characteristics are less important for the transitions to university and to postgraduate study than for transitions through high school; unobserved factors have stronger effects at early stages in the academic careers of children. This is perhaps not surprising because unobserved heterogeneity is measured at the level of the mother and most children are still living in the parental home when they make these early transitions. The same factors that are associated with poorer educational attainment at early stages may become less salient after the child leaves the parental home. A better understanding of the processes leading to educational failure and withdrawal at the secondary level may provide valuable information about increasing educational attainment and reducing social exclusion. These results suggest that simply expanding opportunities at higher levels without targeting the observed and unobserved factors that create differential attainment early on may reinforce patterns of inequality and intergenerational immobility.

Table 5 shows the estimated effects of parental disruption on educational transitions from the simultaneous equation model (Model 2) and a single process model (Model 1). In Model 2 the correlation between u_j and v_j is estimated freely. In Model 1 children's education and mother's marital dissolution are modelled as separate processes, which is equivalent to constraining the correlation between the unmeasured woman-specific components in each process to equal zero. A zero correlation implies that indicators of divorce (prior outcomes of the marital dissolution process) are exogenous with respect to children's educational outcomes, in which case the parameter estimate relating parental divorce to educational attainment will be free from selection bias. Both sets of estimates are net of the effects of the covariates given in Tables 1 and 2.

From either model, we would conclude that the experience of marital breakdown during childhood is associated with lower levels of education, and that the effect weakens with the child's age at disruption. However, if the negative random effect correlation is ignored, the effects of disruption are substantially overstated. The negative effects of divorce estimated from Model 1 are partly explained by a selection into lower levels of education of children whose mothers have a high dissolution risk. Similar to findings from the US, the effects of divorce seem strongest for transitions during or just beyond the high school level – the first two transitions in our sequential probit model (Mare 1995: Biblarz and Gottainer, 2000). The effects of marital breakdown are most pronounced for the transitions from compulsory to lower secondary, and from lower to higher secondary, and non-significant for the transition from undergraduate to postgraduate education. The only significant interaction between gender and divorce is on the transition from secondary school to university; the negative effect of parental divorce is substantially stronger for girls than for boys. That said, the main gender effect (0.418) is larger so although divorce reduces the chances of going to university for females more than it does for males, the predicted probability that a female goes on to university is higher, regardless of her family background experience, than an otherwise similar male.

Comparing the Effects on Educational Attainment of Different Types of Disruption

For both types of disruption, the effects appear to be strongest at earlier stages in the educational process, achieving qualifications beyond lower secondary in particular, but there are some noticeable differences between the two sources of family disruption. Although parental divorce is associated with a reduction in a child's probability of going to university, the effect of a father's death is not significantly associated with progression at that level. In addition, neither type of disruption is associated with achieving a postgraduate qualification amongst those who successfully complete a Bachelor's degree. Importantly, age at divorce is significantly associated with educational attainment but age at death is not. These differences make comparing the effects of the two types of disruption less straightforward.

To facilitate the comparison of the results for different types of disruption and to demonstrate the way in which controlling for selection changes our results, we have used the parameter estimates from Models 1 and 2 to construct the predicted probabilities of specific educational transitions, taking into account different family background experiences.[13] Figure 1 presents the (unconditional) predicted probabilities for the transition from lower secondary to upper secondary with solid bars representing estimates based on Model 1 and striped bars representing estimates based on Model 2.[14] When we assume that there is no unobserved heterogeneity bias in the estimated effect of parental divorce on education (Model 1), it appears that, consistent with studies based on data from the U.S., although any family disruption reduces the likelihood of making the transition from the lower secondary level, children who experience a parental divorce are slightly more disadvantaged than children who experience the death of their father. Although children who experience divorce at older ages are more similar to those who experience a death, we can see from Table 2 that the average age at which a child experiences any disruption is about eight, so there is around a five percentage point gap in progression at the average age. The results change to some extent when we control for selection, however. Controlling for selection on unmeasured time-invariant mother characteristics narrows attainment gaps between single-parent and two-parent families to some extent. Moreover, differences by type of disruption become almost indistinct. Children who experience a parental divorce at younger ages are slightly less likely than those who experience a death to make the transition from lower secondary school, but those who experience a parental divorce at older ages are somewhat more likely to make the transition. At the average age of disruption, the two effects are nearly equal. The results suggest that although family disruptions are associated with poorer educational outcomes, for otherwise similar families, the type of disruption has little consequence. This finding is consistent with our hypothesis that the economic

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[13] To calculate the predicted probabilities, we draw 100 random effects from a normal distribution with mean zero and standard deviation σ_u. We then use the parameter estimates to calculate predicted probabilities for each of the 100 random effect values and take the average. The predicted probabilities are calculated for a male aged 25 years in 2003 with 1 older and 1 younger sibling and parents who both obtained a medium level of education.

[14] The overall patterns described with reference to Figure 1, for the transition for lower secondary to higher secondary, are similar to those obtained for other transitions although the gaps by family background are more narrow for the first transition which almost all children make.

consequences and the levels of welfare support provided to different kinds of single-mother families should be more similar in Norway than in countries like the US. Nonetheless, it is striking that even in a country like Norway where there is far less economic inequality, there are large and significant differences in the educational attainment of children living with both biological parents and those living only with one.

Evidence from previous studies on the importance of the economic consequences of divorce led us to expect that the relationship between family disruption and educational attainment might be smaller than what is found in contexts where the economic consequences are graver and society more unequal. We have not estimated similar models for other countries, so we cannot make definitive conclusions about the relative size of the effects of divorce. Nonetheless, we can compare our results with those of similar models reported in the literature. Differences across models makes the comparison crude[15] and our conclusions tentative, but it is noteworthy that the parameters we estimate for parental divorce and disruption are similar in size if not slightly larger than those obtained from models estimated using US data (Biblarz and Gottainer 2000, Table 1). Comparisons with other studies (see, for example Lehrer (2006) for high school graduation) and results from meta-analyses (Amato and Keith 1991a; Amato 2001) led to similar conclusions. Although far from decisive, but consistent with work focusing on academic test scores, our findings indicate more similarity to than difference from countries with more economic inequality and more poverty amongst single-mother families (Pong, Dronkers, and Hampden-Thompson 2003).

[15] For example, Biblarz and Gottainer (2000) estimate logit models of educational attainment, similar to our sequential probit model but without random effects. Their models control for family structure, gender, race, mother's education, year and child's age. In models of high school completion and college completion, the parameters for living in a single-mother family are -0.74 and -0.41 respectively. If we divide those parameters by 1.6 to approximate probit parameters, we obtain parameters of -0.46 and -0.26. Compared to the parameters for parental separation for Model 1 in the two middle columns of Table 5 (lower to higher secondary and higher secondary to Bachelor's degree) they are both similar in size to those that we obtain for Norway. (Taking into account age at divorce, the parameters we obtain for Norway are more negative for those who experience a divorce at younger ages and less negative for those who experience a divorce at older ages.)

If income changes were the primary explanation for lower educational attainment amongst those children who experienced a parental divorce, we might expect to find a compensatory effect when separated parents later reconcile, as well as negative effects of subsequent parental separations. Drawing on income and sociological, psychological and evolutionary perspectives on stepfamilies, we might also expect to find better outcomes amongst children whose parents reconcile than for those who live in step-families. We find little evidence for any of these. There is no evidence of a compensatory effect in any of the models. In Model 1, we find that having a stepfather is weakly associated with a higher probability of continuing to higher secondary level, but these effects are no longer significant at the 5% level once selection on unobserved mother characteristics is controlled (Model 2). Similarly, apparently negative effects of separation from a stepfather vanish for all but the first transition after adjusting for selection.

Effects of Other Covariates on Children's Educational Transitions

Estimates for variables other than disruption indicators are very similar for Models 1 and 2, so in this section we present and discuss only those from Model 2 (see Table 6). As expected, there are strong, positive effects of both the mother's and the father's education on all transitions, especially on progression beyond lower secondary school and completion of an undergraduate degree. In contrast to the results we obtained for marital dissolution, the education level of the father does not seem to be more important. In fact, the effects of different combinations of parents' education seem symmetrical (high/low is similar to low/high). Consistent with U.S. research showing greater academic achievement amongst girls (Carlson and Corcoran 2001; Manning and Lamb 2003) our findings show that girls are more likely than boys to progress through secondary school and to go to university for undergraduate study (Table 5). Among those who complete a Bachelor's degree, however, men are more likely than women to achieve postgraduate qualifications.

The child's age at the end of the observation period is included to allow for individual variation in the age at which the different levels of education are reached. As expected, the probability of making a transition increases sharply with age (Table 6). The age effect may also be picking up cohort changes in educational progression.

Children whose parents did not marry have a lower chance of progressing through school and on to university. Selection is the most plausible explanation for this result. Parents who choose to have an extra-marital birth and remain unmarried are likely to differ from parents who marry on unobserved factors that are also associated with their children's educational outcomes. One way to account for this form of selection would be to extend the multiprocess model by jointly modelling a woman's decision to marry her child(ren)'s father(s) with educational transitions.

The presence of two or more younger siblings is associated with a decreased probability of continuing through school and to university, and the presence of any older sibling is associated with all transitions. The negative effect of younger siblings may be due to their parents wishing to save for those children's futures and therefore having limited financial resources with which to maintain support of an older child, particularly during university study. Similarly, older siblings may have depleted the family's resources to the extent that they cannot afford to support younger children for an extended period of full-time education. However, selection effects cannot be ruled out. The apparent negative effects of having older or younger siblings may be due to a negative association between the decision to have a larger family and unobserved family characteristics that influence children's educational outcomes. Accounting for this form of selection would mean extending the multiprocess model by jointly modelling partnerships, fertility and educational transitions which would increase the complexity of the model substantially.

CONCLUSION

In this paper we set out to examine the relationship between family disruption and children's educational attainment. A key focus of our study was to determine the extent to which selection biased the estimated association of educational outcomes and parental divorce. We posited that family disruptions in Norway, with its low levels of economic inequality and highly subsidised educational system, might not be very strongly associated with negative outcomes for children. This hypothesis was strengthened by results from studies in Sweden which suggest that selection explains almost all of the differences in educational outcomes by family type. In contrast to these expectations, however, our results suggest that although selection is important and works to overstate the effect of divorce on child outcomes – particularly when divorce happens at a young age – substantial differences between one and two parent families remain, even after time-invariant unobserved factors are controlled. Our findings suggest that once selection on mother-level unobservables is included in the models, children who experience a parental disruption are still 6 – 13 percentage points less likely to successfully make the transition from lower secondary and to complete upper secondary education. This gap is about half as wide as the one that was estimated using a model that did not control for selection but, nonetheless, differences in child outcomes across different family types remain substantial. Moreover disruption has the strongest effect early rather than later in the educational career. As a consequence, children who experience a family disruption could face a heightened risk of unemployment and social exclusion later in the life course. If most of the effects were contained at higher educational levels, the life course consequences might be less severe.

Secondly, we were interested in exploring how selection bias might affect comparisons across children who have experienced different types of family transitions. In models that did not control for selection, children who experienced a father's death appeared less disadvantaged than (younger) children who experienced a parental divorce. However, controlling for selection made differences in the educational attainment between children from divorced and bereaved families narrow substantially, and at mean ages of disruption, almost non-existent. These findings are contrary to a good deal of work in the U.S. (work that frequently does not control for selection as we have) that suggests children in bereaved families usually fare better (Biblarz and Gottainer 2000).

Although our data are not rich enough to distinguish between the different explanations for why there is a significant effect of parental disruption and divorce, the findings raise questions about the theoretical perspectives that emphasise the pathways that are put in place when families undergo a disruption (the changes in economic security, stress and parenting) as well as those that focus on the direct socialization deficit that results from the loss of a parent figure in the same household. The latter perspective would predict that children who experience a divorce might still benefit from parental inputs from a non-resident father, that age at disruption would be negatively and significantly associated with educational outcomes for both types of disruption, and that boys would be more disadvantaged than girls because they would lose the role modelling of their same sex parent. Here the large sample size is particularly beneficial. Because few children growing up in wealthy countries experience a parental death, with smaller survey samples, we might have concluded that age effects were not found to be significant because our sample size was too small. The extremely large sample we use to estimate the models means that we have around 1,900 children who experienced a death. Hence we can be more confident about our finding that age at disruption is insignificant for parental death but not for divorce. Additional questions are raised by our finding that the reunion of separated parents confers no significant benefits suggesting that it is the process of separation and not simply the loss of the parent from a household that leads to lower educational success. Our findings are not clearly consistent with the pathways perspective either. The size of the divorce effects and the fact that parental reunions (as well as our findings for the formation and dissolution of step-families) confer few benefits suggests that while economic security is important, the pathways through which disruption leads to poor educational outcomes are not fully explained by income changes that accompany family transitions.

Although our models provide some interesting results and highlight important areas for future research, several limitations of our study must be considered when we interpret the results. First, it should be noted that, because of data limitations, our study is restricted to children who were born within marriage or to parents who later married. In the period under study, a large proportion of the children were born in consensual unions (Statistics Norway 2007b), and while many of these parents later married, there were also many who dissolved their relationship without ever having married. In fact, the chance of disruption has been 2-3 times higher for cohabiting parents than for married parents (Jensen and Clausen 2003), and the disruptions have often involved particularly young children. Our analysis does not provide any information about the consequences of dissolved consensual unions, which may well be different from those of dissolved marriages, for example because the disruption process is not legally regulated and the fathers have fewer rights.

Moreover, like sibling fixed effects models, we control for time-invariant unobserved factors at the household (in our case, the mother level) rather than the individual child level . Our models control for stable characteristics of the mother that are associated with divorce or educational success but not time-varying environmental conditions that may also have a strong influence, particularly at key stages of transition. Thus exposure to differentially high levels of short-term stress, inadequate parenting or parental conflict will not be well captured because they are acute experiences. In addition, our models assume that the mother-specific effect is the same for all of her children. To the extent that children have different temperaments and/or parents treat children differently, failure to allow for unmeasured child-specific variables might lead to a biased estimate of the effect parental divorce. Given the importance of parenting and the stability of the home environment to child outcomes, we would argue, however, that we have controlled for a particularly important source of bias. Nonetheless, we cannot conclude that our results are entirely free from all confounding factors that might bias the relationship between parental divorce and educational attainment. We can, however, argue that our results are less biased than those that are obtained from models that make no effort to control for selection at all.

Taking into account our findings and the above caveats, the results from this study should be of interest to researchers, policymakers and parents. Perhaps most noteworthy for researchers is our finding that selection bias is substantial. Models that fail to control for unobserved family background characteristics will tend to overstate the effects of divorce on children's outcomes. Care should be taken in the interpretation of parameters as causal, not just in models that make no effort to control for selection but in those that control as we do for only some sources. Moreover, the presence of selection bias may have important implications for those who are, for theoretical reasons, interested in identifying and testing for differences between children who experience different types of family disruptions. Our findings also highlight the importance of looking at different stages of the educational process. If we were to estimate a simple probit of university graduation (our third transition) using the whole sample, we would have found significant effects of divorce on that outcome. But our results suggest differences at that level are due to some extent to children falling out of the system earlier on. Taking into account the different transitions may be important for understanding the pathways that lead to poor educational outcomes. Qualitative research to understand the family processes that lead to disadvantage and to identify things we need to measure is clearly warranted. Comparative research to get a better understanding of institutional factors that enhance or mitigate family risk factors would also provide additional interpretative information.

Although selection is important and accounts for a large portion of the observed differences in educational outcomes by family experiences, it remains true that children who experience a family disruption are disadvantaged. Predicted probabilities show large gaps in educational transitions by family experience, but this does not mean all children who experience a parental divorce are doomed to educational failure. The effects of family disruption are actually small relative to differences between children with high and low educated parents. Our findings concerning the importance of selection – as we have modelled it – should not be overlooked by policymakers either. When they consider how to design policies to equalise the opportunities of different children, it is important to take into account that much of the association is not causal. The findings concerning mother-specific effects suggest that what happens in the home and what happens early on have important implications for children. Although further research is needed to identify the most important pathways, additional parenting support for those who are going through a disruption may benefit children. Moreover, policymakers wanting to create a fair and more just society may want to target those children at greater risk of low educational attainment for additional educational support. The limitations of the study – particularly the fact that we have not controlled for all possible sources of unobserved heterogeneity – mean that we cannot conclude from our negative and significant parameter estimates that parents in an unhappy relationship should not divorce or that policymakers should make it harder for them to do so. We have already mentioned that an unobserved and time-varying factor that is not controlled in our models is parental conflict. Other time-varying aspects of the parents' relationship quality may matter as well. Those parents who are in low-quality, high-conflict relationships may cause more problems for their children if they stay together than if they divorce.

In advanced welfare states where a good deal of resources and support are now provided publicly, it is nonetheless true that what happens in the private sphere remains crucial to the well-being of children. Learning more about this, including the implications of family disruptions, would obviously be important both to parents themselves and to those involved in the development of (family) policies that can make life better for our children. Although our findings do not settle the dispute about the "true" relationship between divorce and education and cannot shed much light on the underlying mechanisms, they do provide important theoretical and practical information about the relationship between family background and educational success. In particular, our findings suggest the need for future research on the processes that accompany family disruption, particularly in a comparative context, in order to identify ways that children experiencing family disruption can be better supported and encouraged to successfully complete their education.

REFERENCES

d'Addio, A.C. 2006. *Intergenerational Transmission of Disadvantage: Mobility or Immobility across Generations? A Review of the Evidence for OECD Countries* OECD Social, Employment and Migration Working Paper No. 52

Amato, P.R. 2001. Children of divorce in the 1990s: An update of the Amato and Keith (1991) meta-analysis. *Journal of Family Psychology* 15: 355-370.

Amato, P.R., and B. Keith. 1991a. Parental divorce and the well-being of children: A meta-analysis. *Psychological Bulletin* 110: 26-46.

Amato, P.R. and B. Keith 1991b. Parental divorce and adult well-being: A meta-Analysis. *Journal of Marriage and the Family* 53: 43-58.

Astone, N.M. and S,S. McLanahan. 1991. Family structure, parental practices and high school completion. *American Sociological Review* 6: 309-320.

Biblarz, T.J. and G. Gottainer. 2000. Family structure and children's success: A comparison of widowed and divorced single mother families. *Journal of Marriage and the Family* 62(2): 533-548.

Biblarz. T.J. and A. Raftery. 1999. Family structure, educational attainment, and socioeconomic success: Rethinking the 'pathology of the matriarchy'. *American Journal of Sociology* 105: 321-365.

Björklund, A. and M. Sundström 2006. Parental Separation and Children's Educational Attainment: A Siblings Analysis on Swedish Register Data. *Economica* 73: 605–624

Björklund, A. D.K. Ginther and M. Sundström. 2007. Family structure and child outcomes in the USA and Sweden. *Journal of Population Economics* 20: 183-201.

Black S.E., P.J. Devereux, K.G. Salvanes. 2005. The more the merrier? The effect of family size and birth order on children's education. *The Quarterly Journal of Economics* 120: 669-700.

Bratberg, E. and S. Tjøtta. 2007. Income effects of divorce in families with dependent children. Forthcoming in *Journal of Population Economics.*

Breivik K. and D. Olweus. 2006. Children of divorce in a Scandinavian welfare state: are they less affected than US children? *Scandinavian Journal of Psychology* 47: 61-74.

Brien, M.J. and L.A. Lillard. 1994. Education, marriage and first conception in Malaysia. *The Journal of Human Resources* 29: 1167-1204.

Brown, S.L. 2006. Family structure transitions and adolescent well-being. *Demography* 43(3): 447-463.

Carbonneau, R., L.J. Eaves, J.L. Silberg, E. Simonoff, and M. Rutter. 2002 Assessment of the within- family environment in twins: absolute versus differential ratings and relationship with conduct problems. *Journal of Child Psychology and Psychiatry* 43: 1064-1074.

Carlson, M. and M. Corcoran. 2001. Family structure and children's behavioral and cognitive outcomes. *Journal of Marriage and the Family* 63: 779-792.

Case, A., I. Lin and S.S. McLanahan. 2001. Educational attainment of siblings in step-families. *Evolution and Human Behavior* 22: 269-289.

Cherlin, A. 1978. Remarriage as an incomplete institution. *American Journal of Sociology* 84(3): 634-650.

Clarke, L. and A. Berrington. 1999. Socio-demographic predictors of divorce. *High Divorce Rates: The State of the Evidence on Reasons and Remedies - Volume 1 (Papers 1-3) - Research Series No. 2/99.*

Covey, L.S. and D. Tam. 1990. Depressive mood, the single-parent home, and adolescent cigarette smoking. *American Journal of Public Health* 80: 1330-1333.

DeLeire, T. and A. Kalil. 2002. Good things come in threes: Single-parent multigenerational family structure and adolescent adjustment. *Demography* 39: 393-413.

Downey, G. and J.C. Coyne. 1990. Children of depressed parents: An integrative review. *Psychological Bulletin* 108: 50-76.

Ermisch, J.F. and M. Fransconi. 2001. Family structure and children's achievements. *Journal of Population Economics* 14: 249-270.

Ginther, D.K. and R.A. Pollak. 2004. Does family structure affect children's educational outcomes? *Demography* 41: 671–96.

Gruber, J. 2004. Is making divorce easier bad for children? The long run implications of unilateral divorce. *Journal of Labor Economics* 22: 799–833.

Hao, L. and G. Xie. 2001. The complexity and endogeneity of family structure in explaining children's misbehavior. *Social Science Research* 31: 1-28.

Hetherington, E.M., M. Bridges, and G.M. Insabella. 1998. What matters? What does not? Five perspectives on the association between marital transitions and children's adjustment. *American Psychologist* 53: 167-184.

Hetherington, E.M., M. Cox, and R. Cox. 1982. Effects of divorce on parents and children." Pp. 233-285 in *Nontraditional Families: Parenting and Child Development,* edited by Michael E. Lamb. Hillsdale, NJ: Lawrence Erlbaum.

Huerta M.C. and W. Sigle-Rushton. 2007. "Pathways to early motherhood in the United Kingdom". Mimeo, London School of Economics.

Jekielek, S.M. 1998. Parental conflict, marital disruption, and children's emotional well-being. *Social Forces* 76: 905-935.

Jenkins, J.M., J. Rabash, and T.G. O'Connor. 2003. The role of shared family context in differential parenting. *Developmental Psychology* 39: 99-113.

Jensen, A.-M. and S.-E. Clausen. 2003. Children and family dissolution in Norway: The impact of consensual unions. *Childhood* 10, 65 - 81.

Jonnson J.O and M. Gähler. 1997. Family dissolution, family reconstitution, and children's educational careers: recent evidence for Sweden. *Demography* 34: 277-293.

Kiernan, K.E. and G. Mueller. 1999. Who divorces? in McRae, S. (ed.), *Changing Britain: Families and Households in the 1990s*. Oxford: Oxford University Press

Krein, S.F. and A.H. Beller. 1988. Educational Attainment of Children from Single-parent Families: Differences by Exposure, Gender, and Race. *Demography* 25(2): 221-234.

Kjeldstad R and M. Rønsen. 2002. Welfare rules, business cycles and the employment of single parents. Documents 2002/7. Statistics Norway. Oslo-Kongsvinger.

Lang, K. and J.L. Zagorsky. 2001. Does growing up with a parent absent really hurt? *Journal of Human Resources* 36: 253-273.

Lehrer, E.L. 2006. Religion and high-school graduation: A comparative analysis of patterns for white and black young women. *Review of Economics of the Household* 4: 277-293.

Lillard, L.A. and C.W.A. Panis. 2003. *aML Multilevel Multiprocess Statistical Software, Version 2.0*. EconWare, Los Angeles, California.

Lillard, L.A. and L. Waite. 1993. A joint model of marital childbearing and marital disruption. *Demography* 30: 653-682.

Lyngstad, T.H. 2004. The impact of parents' and spouses' education on divorce rates in Norway. *Demographic Research* 10(5): 121-142.

Manning, W.D. and K.A. Lamb. 2003. "Adolescent well-being in cohabiting married, and single-parent families." *Journal of Marriage and the Family* 65:876-893.

Mare, R.D. 1995 Changes in educational attainment and school enrollment. In *State of the Union: America in the 1990s,* edited by R. Farley. New York: Russell Sage Foundation.

McLanahan, S.S. 1985. Family structure and the reproduction of poverty. *American Journal of Sociology* 9: 873-901.

McLanahan, S.S. and L.L. Bumpass, 1988. Intergenerational consequences of family disruption. *American Journal of Sociology* 94: 130-152.

McLanahan, S. S., and G. Sandefur. 1994. *Growing Up with a Single Parent: What Hurts, What Helps?* Cambridge, MA: Harvard University Press.

Page, M and A.H. Stevens 2004. The economic consequences of absent parents. *The Journal of Human Resources* 39(1): 80-107.

Painter, G. and D.I. Levine. 2000. Family structure and youths' outcomes: Which correlations are causal? *Journal of Human Resources* 35: 524-549.

Pong, S.L., J. Dronkers and E. Hampden-Thompson. 2003. Family policies and children's school achievement in single- versus two-parent families. *Journal of Marriage and the Family* 65: 681-699.

Rogers, B. and J. Pryor. 1998. *Divorce and Separation: The Outcomes for Children.* York, UK: Joseph Rowntree Foundation.

Sigle-Rushton, W. and S.S. McLanahan. 2004. Father absence and child well-being: A critical review. In *The Future of the Family*, edited by D.P. Moynihan, T. Smeeding, and L. Rainwater. New York: Russell Sage Foundation.

Sigle-Rushton, W., J. Hobcraft and K. Kiernan. 2005. Parental disruption and adult well-being: A cross cohort comparison. *Demography* 42(3): 427-446.

Statistics Norway 2001. Norwegian Standard Classification of Education. Revised 2000 *Official Statistics of Norway* C 617. Oslo-Kongsvvinger

__. 2007a. Fewer marriages and divorces in 2006.

Online at http://www.ssb.no/english/subjects/02/02/30/ekteskap_en/

__. 2007b. Highest fertility since 1991. Online at http://www.ssb.no/english/subjects/02/02/10/fodte_en/

Størkersen, I., E. Røysamb, T. Moum and K. Tambs. 2005. Adolescents with a childhood experience of parental divorce: a longitudinal study of mental health and adjustment. *Journal of Adolescence* 28: 725-739.

Thompson, E., T.L. Hanson, and S.S. McLanahan. 1994. Family structure and child well-being: Economic resources vs. parental behaviors. *Social Forces* 73:221-242.

Upchurch, D.M., L.A. Lillard and C.W.A. Panis. 2002. Nonmarital childbearing: influences of education, marriage, and fertility. *Demography* 39: 311-329.

United Nations. (2006). Inequality in income or consumption. In *Human Development Reports*. Online at http://hdr.undp.org/statistics/data/indicators.cfm?x=148&y=2&z=1

Yeung, J.W., G.J. Duncan, and M. Hill. 1995. A Sibling-Based Analysis of the Effects of Family Structure on Children's Achievement. Paper presented at the 1995 Meeting of the Population Association of America.

Figure 1: Probability of Continuing Beyond Lower Secondary (Before and After Allowing for Selection)

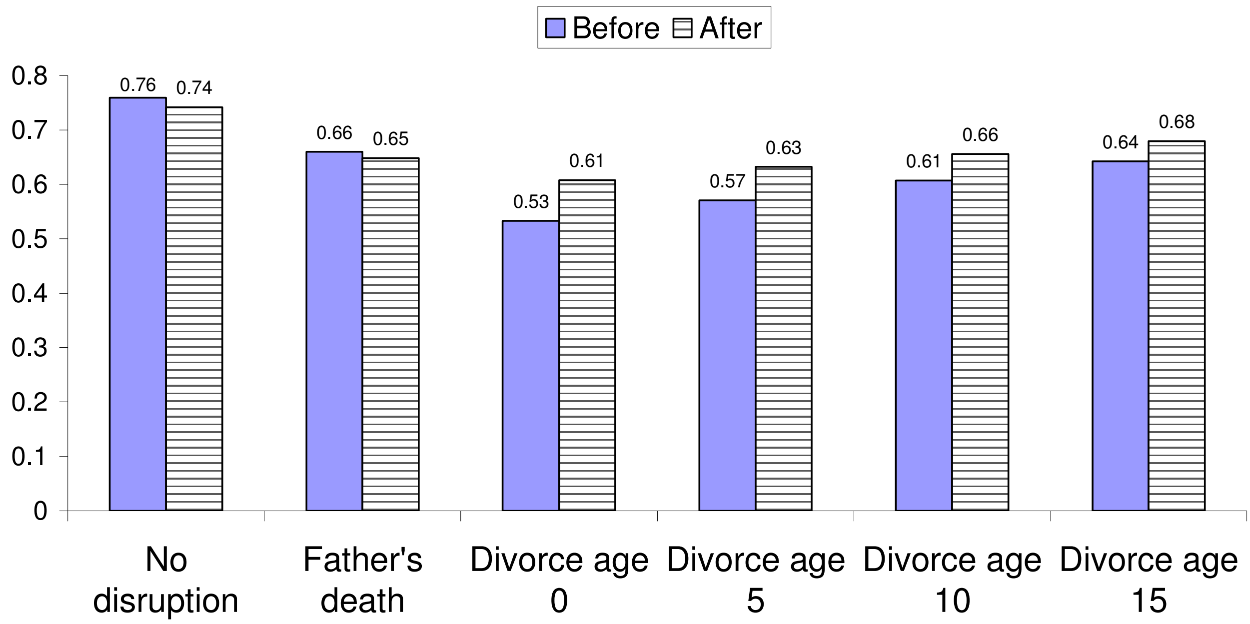

Table 1. Descriptive statistics for variables included in the model for marital dissolution (n=129189 marriages)

	Percent	**Mean**	**Std Dev**
Woman previously married	11.8		
Woman has children from previous partner	13.8	-	-
Spouse has children from previous partner	10.1		
Woman and spouse have children together before marriage	38.1		
No. children with spouse at duration t [a]			
None	5.6		
1	17.8		
2	45.6		
3+	31.0		
Woman's age at marriage	-	25.96	5.87
Age difference between woman and spouse			
Woman 5+ years older	2.3	-	-
Woman 1-4 years older	12.2	-	-
Spouse 0-4 years older (reference)	60.3	-	-
Spouse 5-9 years older	20.5	-	-
Spouse 10+ years older	4.6	-	-
Education level of woman/spouse in 2003 [b]			
Mother low/Father low	28.1	-	-
Low/medium	15.4	-	-
Low/high	9.0	-	-
Medium/low	7.7	-	-
Medium/medium	6.0	-	-
Medium/high	4.4	-	-
High/low	6.5	-	-
High/medium	5.9	-	-
High/high	17.2	-	-

Notes:

[a] Treated as time-varying in the models, but descriptive statistics refer to the number of children born by the last observed year of marriage (corresponding to dissolution or censoring).

[b] For both the woman and her husband, a low level of education is defined as no more than lower secondary (12 years), medium is higher secondary (13 years), and high is post-secondary (14+ years).

Table 2. Descriptive statistics for variables included in the model for children's education

	Percent	**Mean**	**Std Dev**	**Base n (No. children)**
Experience of family disruption before16				
Parents did not marry	3.5	-	-	197638
Parents separated[b]	23.0	-	-	190758
Father died[b]	1.0	-	-	190758
Age at disruption (separation or death of father)[c]	-	7.73	4.28	45714
Separation later annulled[d]	10.5	-	-	43844
Child had a stepfather[e]	16.6	-	-	52594
Mother separated from stepfather	31.4	-	-	8735
Background characteristics of the child				
Female	48.8	-	-	197638
Age in 2003	-	21.24	3.62	197638
No. younger siblings in year child turned 17				197638
None	32.6	-	-	
1	41.3	-	-	
2+	26.1	-	-	
No. older siblings in year child turned 16				197638
None	53.8	-	-	
1	35.8	-	-	
2+	10.4	-	-	
Education level of mother/father in 2003[f]				197638
Mother low/Father low	28.6	-	-	
Low/medium	15.0	-	-	
Low/high	9.0	-	-	
Medium/low	7.6	-	-	
Medium/medium	5.6	-	-	
Medium/high	4.2	-	-	
High/low	6.6	-	-	
High/medium	5.7	-	-	
High/high	17.6	-	-	

Notes:

[a]Approximate year equivalents for educational qualifications are: compulsory (10 years), lower secondary (11-12 years), higher secondary (13 years), Bachelors degree (14-17 years), Masters degree and above (18 or more years).

[b]Defined for children whose parents married before age 16.

[c]Defined for children who experienced disruption before age 16.

[d]Defined for children whose parents separated before age 16.

[e]Defined for children whose parents did not marry or who married and separated before age 16.

[f]For both parents, a *low* level of education is defined as no more than lower secondary (10-12 years), *medium* is higher secondary (13 years), and *high* is post-secondary (14+ years).

Table 3. Estimated coefficients from the random effects hazards model of marital disruption (Model 2, $\rho_{uv} \neq 0$)

	Coefficient	SE
Constant	-3.051**	0.055
Duration spline (years)		
0-2	0.370**	0.021
2-4	0.073**	0.016
4-6	0.023	0.016
6-8	0.005	0.017
8-10	0.067**	0.013
10+	-0.002	0.002
Woman previously married	0.777**	0.030
Children from previous partner (woman)	0.306**	0.022
Children from previous partner (spouse)	0.399**	0.026
Children with spouse born before marriage	0.347**	0.013
No. children with spouse at duration t		
None (reference)	0	-
1	-0.328**	0.024
2	-0.683**	0.027
3+	-0.945**	0.030
Woman's age at start of marriage	-0.064**	0.002
Age difference between woman and spouse		
Woman 5+ years older	0.542**	0.043
Woman 1-4 years older	0.117**	0.020
Spouse 0-4 years older (reference)	0	-
Spouse 5-9 years older	0.090**	0.015
Spouse 10+ years older	0.213**	0.028
Education of woman/spouse		
Low/low (reference)	0	-
Low/medium	-0.256**	0.019
Low/high	-0.353**	0.023
Medium/low	0.021	0.023
Medium/medium	-0.279**	0.027
Medium/high	-0.317**	0.032
High/low	0.141**	0.024
High/medium	-0.109**	0.027
High/high	-0.394**	0.020

** Significant at 1% level; * significant at 5% level.

Table 4. Mother-level random effect parameters from the simultaneous equation model

Parameter	Estimate	SE
Education equation		
Random effect loadings		
$\lambda^{(1)}$	1[†]	-
$\lambda^{(2)}$	1.078**	0.041
$\lambda^{(3)}$	0.864**	0.041
$\lambda^{(4)}$	0.584**	0.077
Standard deviation σ_u	0.498**	0.014
Disruption equation		
Standard deviation σ_v	0.960**	0.023
Cross-equation		
Correlation ρ_{uv}	-0.431**	0.023

[†]Constrained to equal 1.

Table 5. Estimated effects of family disruption indicators (γ) from random effects sequential probit models of children's education transitions

	Compulsory to lower secondary		Lower to higher secondary		Higher secondary to Bachelor's degree		Bachelor's degree to Master's or higher	
	Model 1 ($\rho_{uv}=0$)	Model 2 ($\rho_{uv}\neq 0$)	Model 1 ($\rho_{uv}=0$)	Model 2 ($\rho_{uv}\neq 0$)	Model 1 ($\rho_{uv}=0$)	Model 2 ($\rho_{uv}\neq 0$)	Model 1 ($\rho_{uv}=0$)	Model 2 ($\rho_{uv}\neq 0$)
Parents separated	-0.580** (0.025)	-0.349** (0.028)	-0.631** (0.024)	-0.386** (0.028)	-0.350** (0.035)	-0.155** (0.037)	-0.056 (0.095)	0.072 (0.096)
Age at separation	0.019** (0.002)	0.013** (0.002)	0.018** (0.002)	0.013** (0.002)	0.012** (0.003)	0.008** (0.003)	-0.012 (0.009)	-0.014 (0.009)
Father died	-0.201** (0.066)	-0.178** (0.066)	-0.318** (0.059)	-0.295** (0.059)	-0.124 (0.083)	-0.108 (0.082)	-0.206 (0.251)	-0.198 (0.252)
Female	0.217** (0.012)	0.217** (0.012)	0.318** (0.010)	0.320** (0.010)	0.418** (0.011)	0.418** (0.011)	-0.180** (0.027)	-0.179** (0.027)
Female*separation	0.034 (0.022)	0.034 (0.022)	0.005 (0.020)	0.003 (0.020)	-0.128** (0.027)	-0.130** (0.027)	0.004 (0.074)	0.004 (0.074)
Female*father died	-0.126 (0.095)	-0.125 (0.095)	0.107 (0.084)	0.110 (0.085)	0.137 (0.109)	0.139 (0.109)	0.468 (0.311)	0.473 (0.312)
Separation annulled	-0.022 (0.031)	0.033 (0.031)	-0.070* (0.029)	-0.013 (0.030)	0.034 (0.042)	0.080 (0.042)	-0.125 (0.124)	-0.097 (0.124)
Child had a stepfather	0.112** (0.027)	0.067 (0.027)	0.081** (0.025)	0.026 (0.026)	0.028 (0.036)	-0.018 (0.036)	-0.086 (0.100)	-0.121 (0.100)
Separation from stepfather	-0.296** (0.036)	-0.140** (0.037)	-0.235** (0.036)	-0.058 (0.038)	-0.169** (0.055)	-0.022 (0.055)	-0.180 (0.170)	-0.080 (0.170)

** Significant at 1% level; * significant at 5% level.

Notes:

(i) All family disruption indicators refer to experiences up to the year the child turned 16.

(ii) All effects are net of the effects of the other explanatory variables shown in Table 6.

(iii) Model 1 assumes no correlation between the unobserved mother-specific components (random effects) in the children's education and mother's marital disruption equations. In Model 2, the random effects correlation is freely estimated.

(iv) Standard errors given in parentheses.

Table 6. Estimated effects of other explanatory variables from the random effects sequential probit model of children's educational transitions (Model 2, $\rho_{uv} \neq 0$)

	Compulsory to lower secondary		Lower to higher secondary		Higher secondary to Bachelor's degree		Bachelor's degree to Master's or higher	
	Coeff.	SE	Coeff.	SE	Coeff.	SE	Coeff.	SE
Parents still unmarried when child turns 16	-0.413**	0.026	-0.520**	0.025	-0.335**	0.036	-0.211*	0.097
Age of child in 2003[a]								
16	-3.286**	0.023	-		-	-	-	-
17	-0.856**	0.015	-		-	-	-	-
18	-0.358**	0.017	-2.644**	0.020	-	-	-	-
19	-	-	-0.835**	0.013	-2.834**	0.046	-	-
20	-	-	-0.389**	0.013	-1.912**	0.025	-	-
21	-	-	-	-	-1.359**	0.020	-	-
22	-	-	-	-	-0.940**	0.017	-	-
23	-	-	-	-	-0.646**	0.017	-1.559**	0.083
24	-	-	-	-	-0.424**	0.016	-0.881**	0.046
25	-	-	-	-	-0.224**	0.017	-0.487**	0.035
No. younger siblings at 17								
None (reference)	0	-	0	-	0	-	0	-
1	-0.012	0.013	-0.017	0.011	-0.029*	0.013	0.041	0.039
2+	-0.054**	0.015	-0.092**	0.013	-0.049**	0.016	-0.007	0.042
No. older siblings at 16								
None (reference)	0	-	0	-	0	-	0	-
1	-0.128**	0.012	-0.167**	0.010	-0.141**	0.012	-0.213**	0.039
2+	-0.226**	0.018	-0.298**	0.017	-0.230**	0.026	-0.328*	0.131
Education of mother/father								
Low/low (reference)	0	-	0	-	0	-	0	-
Low/medium	0.200**	0.016	0.180**	0.014	0.184**	0.017	0.056	0.050
Low/high	0.394**	0.201	0.471**	0.017	0.539**	0.019	0.394**	0.047
Medium/low	0.217**	0.020	0.183**	0.018	0.210**	0.022	0.185**	0.062
Medium/medium	0.357**	0.024	0.374**	0.021	0.344**	0.026	0.104	0.074
Medium/high	0.506**	0.029	0.689**	0.025	0.680**	0.026	0.597**	0.061
High/low	0.351**	0.022	0.458**	0.019	0.539**	0.022	0.290**	0.056
High/medium	0.456**	0.024	0.600**	0.021	0.648**	0.024	0.459**	0.059
High/high	0.678**	0.018	0.888**	0.015	0.920**	0.017	0.689**	0.042

** Significant at 1% level; * significant at 5% level.

[a]The reference category for age in 2003 is 19+ years for the transition from compulsory to lower secondary transition, 21+ years for lower to higher secondary, and 26+ years for both higher secondary to Bachelor's degree and for Bachelor's to postgraduate degree.