The Relationship between Childbearing and Transitions from Marriage and Cohabitation in Britain*

Fiona Steele, Constantinos Kallis, Harvey Goldstein and Heather Joshi
Institute of Education, University of London

Address for Correspondence:

Bedford Group for Lifecourse and Statistical Studies
Institute of Education
University of London
20 Bedford Way
London WC1H 0AL

Email: F.Steele@ioe.ac.uk

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ABSTRACT

We describe a general framework for the analysis of correlated event histories, with an application to a study of partnership transitions and fertility among a cohort of British women. A multilevel multistate competing risks model is used to examine the relationship between prior fertility outcomes (the presence and characteristics of children and current pregnancy) and the dissolution of marital and cohabiting unions and movements from cohabitation to marriage. Using a simultaneous equations model these partnership transitions are modelled jointly with fertility, allowing for correlation between the unobserved woman-level characteristics that affect each process. The analysis is based on the partnership and birth histories collected for the 1958 birth cohort (National Child Development Study) between the ages of 16 and 42. We find that preschool children have a stabilising effect on their parents' partnership, whether married or cohabiting, but the effect is weaker for older children. There is also evidence that while pregnancy precipitates marriage among cohabitors, the odds of marriage decline to pre-pregnancy levels following a birth.

INTRODUCTION

The trend towards greater instability and informality of partnerships is associated with later and less childbearing at the national level in many developed countries. To some extent, childbearing is postponed until a stable partnership is formed, but also an increasing number of children experience the break-up of their parent's union such that the majority of lone parent families are the outcome of partnership dissolution rather than unpartnered motherhood, and a growing number of step-parent families are also being formed. The nature and dynamics of partnership stability and childbearing are of interest in understanding the changing demographic and social structure, and in helping to inform the expectations of policy makers and the public at large.

Although some adults may benefit from the greater freedom to dissolve unsatisfactory relationships, partnership instability tends to be detrimental to the well being of the children involved, and is therefore a matter of particular concern for public policy. When a family breaks up there are likely to be greater claims on collective resources for income and other support[1]. Although many lone mothers are reconstituted as stepfamilies, such families tend to face extra stresses and strains (Ferri and Smith 1998) and may also be at heightened risk of dissolution. Furthermore, a climate of opinion that children suffer when parents split up may help to reinforce couples' preferences to stay together.

This paper distinguishes partnerships which are marriages from those which are informal cohabitations not only because marriage is thought to signal a greater commitment, but also because there is more of a formal mechanism for dealing with the division of

[1] Joshi et al. (1999) analyse the educational and economic disadvantages of some of the children of the 1958 birth cohort studied here when not in intact families (those in a 1 in 3 sample aged 4-17 in 1991).

resources and parental rights on dissolution. We set out to investigate the relationship between partner transitions and childbearing in the life histories, so far, of a large set of women born in 1958. We ask under what circumstances does the presence of children stabilize a partnership, and which children are particularly at risk of parental break-up.

Correlated Processes and Multiple States

Transitions from marital and nonmarital partnerships and childbearing within those partnerships are two related dynamic processes. An individual is observed over the course of one or more partnerships, and a transition may occur at any time. A partnership may dissolve or, in the case of cohabitation, may be converted to marriage and, within a partnership, one or more children may be born. The decision to end a union, or to move from cohabitation to marriage, is likely to be jointly determined with the decision to have a child with that partner. In other words, there may be factors, both observed and unobserved, which drive both processes. If decisions about partnerships and childbearing are jointly determined, the unobserved components of the models for each process will be correlated. In that case, indicators of the presence of children will not be independent of the residuals in the model for partnership transitions, and estimates of their effects on partnership outcomes will be biased.

In this paper, we model jointly the processes of partnership stability and fertility using a multiprocess model, which allows for unobserved factors that influence both partnership and fertility outcomes. Multiprocess modelling of event history data was first proposed by Lillard (1993), with an application to an analysis of marital dissolution and marital fertility. While marriage remains the favoured setting for childbearing in North America and much of Western Europe, the increasing prevalence of unmarried cohabitation among young couples has contributed to a rise in extra-marital births (Kiernan 2001; Loomis and Landale 1994). In England and Wales, for example, 41% of live births in 2002 were outside marriage, and 64%

of those were jointly registered by parents resident at the same address (Office for National Statistics 2004). The increase in extra-marital childbearing has in recent years led to research that extends Lillard's work to include nonmarital fertility. Brien, Lillard and Waite (1999) examine the interrelationships between nonmarital fertility and the formation of marital and cohabiting unions. This framework is extended by Upchurch, Lillard and Panis (2002) to incorporate the processes of marital dissolution, marital fertility and educational transitions. In Britain, Aassve et al. (2004) use a multiprocess model to analyse union formation and dissolution, jointly with fertility and employment decisions.

While the above studies consider nonmarital childbearing, neither Brien et al. nor Upchurch et al. distinguish between conceptions within cohabiting unions and those that occur outside co-resident relationships. Similarly, in their analysis of the formation of marital unions, Brien et al. do not consider transitions from cohabitation separately from direct marriage from an unpartnered state. This approach is consistent with the idea that cohabitation is closer in nature to being single. In contrast, Aassve et al. view cohabitations as *de facto* marriages and therefore combine marriage and cohabitation into one 'union' state. In practice, however, individuals may live with their partners for different reasons and their perception of cohabitation may change over time (Kiernan 2001; Loomis and Landale 1994; Murphy 2000; Rindfuss and Vandenheuvel 1990). Cohabitors may regard cohabitation as an alternative to marriage, a precursor to marriage, or more convenient than a non co-resident sexual relationship. Furthermore, partners' views of cohabitation are likely to influence whether they have children outside marriage, and the effect of the presence of children on the outcome of their partnership. For these reasons, we treat marriage and cohabitation as two separate partnership 'states' in our analysis of the interrelationship between partnership stability and childbearing. Following previous research on the duration of cohabitation we consider two outcomes of a cohabiting union, separation and marriage to the same partner, in a competing risks framework.

The hazards of partnership transitions are modelled jointly with the hazard of a conception, again distinguishing between marital and nonmarital unions. A simultaneous equations model is used to allow for correlation between the unmeasured individual-specific determinants of partnership durations and fertility. We also build on previous research based on the National Child Development Study (NCDS), which considers only the first partnership (e.g. Berrington 2001; Kiernan and Cherlin 1999). These studies consider the predictors of divorce and the outcomes of cohabitation in Britain for ages 16-33, while we make use of event history data collected in 2000 to consider transitions up to the age of 42. By their early forties, many individuals have lived with more than one partner and have had children with those partners, making for increasing complexity in the living arrangements of families (Ferri and Smith 1998). We analyse repeated partnership transitions, using a multilevel model to allow for correlation between the durations of partnerships for the same individual.

We focus on the processes of partnership dissolution and entry into marriage from cohabitation, and their correlation with fertility during partnerships. We do not consider the formation of cohabiting unions or direct marriage from an unpartnered state. Therefore episodes spent outside a partnership are omitted from the analysis, which leads to the exclusion of a small number (7%) of respondents who delay formation of their first partnership until after age 33. This may result in a selection bias if those who delay partnering differ in unobserved ways from those who partner earlier. Further, because we do not model conceptions that occur outside co-resident partnerships, children arising from extra-union conceptions are treated as exogenous. There are two major reasons for restricting our focus to the *outcomes* of partnerships. First, our primary substantive interest is in the effects of the presence of children on union stability, that is, the risk of dissolution and the chance that cohabitation is converted to marriage. Second, incorporating equations for partnership formation and extra-union fertility would greatly increase the complexity of the multiprocess model. We return to this issue in the Discussion.

PREVIOUS RESEARCH ON THE LINK BETWEEN PARTNERSHIP OUTCOMES AND CHILDBEARING

We begin with a review of the current evidence for effects of fertility outcomes on marital dissolution. There is a growing body of research on the effect of the outcomes of previous fertility decisions on the dissolution of marriage and cohabiting unions, and on the translation of cohabitation into marriage. Most authors contend that having children together raises the costs of separation and increases the gains from marriage, leading to a negative effect of the presence of children on the risk of divorce, particularly while children are young (e.g. Koo and Janowitz 1983; Lillard and Waite 1993). This theory is largely supported by the empirical evidence. In Britain, using data from the NCDS up to age 33, Berrington and Diamond (1999) consider the effect of the timing of childbearing on dissolution of the first marriage. They find that women who experience a premarital birth or conception are less likely to separate than childless women, and that women with a marital birth are the least likely to separate. Using data from a nationally representative British survey carried out in 1976, Murphy (1985) finds evidence of a U-shaped relationship between the number of children born within marriage and marital disruption, with the lowest risk observed among women with two children.   A lower risk of partnership dissolution among women with one or two children is also found in an analysis of more recent British data (Aassve et al. 2004). Lillard and Waite (1993) report that women in the US who had one child with their current husband were at lower risk of divorce than childless women, but that subsequent children had a destabilising effect on the marriage. In an earlier study, Waite and Lillard (1991) consider also the age of children and find that the presence of a young child delays separation until the child is of school age.

The direction of the effect of the presence of children from a previous relationship on marital separation is more difficult to predict. We might expect that having children, regardless of their parentage, would reduce the risk of separation by constituting a shared interest. Furthermore, women with children who enter a new partnership have already experienced the breakdown of the relationship with the child's father. To the extent that separation is more traumatic where there are children involved, women with children may be more selective in their choice of a future partner. This form of selection would lead to a negative effect of children from a previous relationship on the risk of separation. On the other hand, if the prospect of stepchildren is an impediment in the "marriage market" or a source of conflict in a partnership, their presence could increase the risk of dissolution. To date, empirical evidence from the US and Britain supports the latter theory (Ermisch and Pevalin 2003; Lillard and Waite 1993; Murphy 1985; Waite and Lillard 1991).

The increasing prevalence of cohabitation has led to research on the determinants of the dissolution of nonmarital partnerships. While Wu (1995) suggests that sociological and economic theories about marital stability should apply to nonmarital unions as well, there is a lack of concensus in the findings from British and North American studies on the effects of fertility outcomes on separation from cohabitation. Berrington (2001) finds that neither the presence of children born within the partnership nor pregnancy has an effect on the rate of separation from cohabiting first partnerships by age 33 among women in the 1958 British birth cohort. Ermisch and Francesconi (2000) reach a similar conclusion in an analysis of a representative sample of British never-married women, although they do not distinguish between children fathered by the current partner and pre-union births; they also find no effect of having a child of preschool age. However, analyses using data from the same study, but based on women with children and not distinguishing between marriage and cohabitation, find that the risk of dissolution is lower among women with a young child but increases with the number of children (Böheim and Ermisch 2001). The findings from research in the US are consistent with those from the British studies. Manning and Smock (1995) find that the presence of children and pregnancy has no effect on the risk of separation among never-married cohabiting men and women. In contrast a Canadian study, also based on never-married men and women, finds a negative effect of the presence of children on the risk of separation, but no effects of the characteristics (i.e. the number, age and sex) of children (Wu 1995). In another study, with separate analyses by gender, pre-union births are found to increase the risk of separation among women (Wu and Balakrishnan 1995).

Most of the studies described above also consider the effects of the presence and characteristics of children on the other possible outcome of cohabitation, that is the conversion of the partnership into marriage. Couples with children, or women with children from a previous relationship, may seek a more formal marital partnership for childrearing. However, it is also possible that couples who have children while living together view cohabitation as an alternative to marriage, and are therefore less likely to marry than childless couples. A negative effect of the presence of children from a previous relationship would also be expected if men are less inclined to marry women with children fathered by another partner. Berrington (2001) finds that pregnant women are more likely to marry their partner, but there is no effect of children born within the partnership and the presence of a child born before the start of the partnership lowers the probability of marriage. Another British study (Ermisch and Francesconi 2000) concludes that mothers are less likely to marry their partner than childless women. Wu and Balakrishnan (1995) also find that the presence of children, born either during or before the partnership, lowers the probability of marriage in Canada.

The findings from US studies are inconclusive. Manning and Smock (1995) find that the presence of children increases the probability of marriage. However, Brien et al. (1999) find that while pregnant women are more likely to marry their partner, marriage is less likely after a birth. Brien et al. also find that the presence of children fathered by a previous partner reduces a woman's hazard of marriage.

While most of the studies described above treat the presence and characteristics of children as exogenous covariates in their analyses of partnership transitions, several authors have used multiprocess modelling to allow and test for the endogeneity of fertility outcomes. These studies allow for the possibility that fertility and partnership outcomes are jointly determined due to correlation between the unmeasured individual characteristics (represented by residual terms in the statistical model) that affect each process. A residual correlation that is significantly different from zero provides evidence of endogeneity. Lillard and Waite (1993) find a negative correlation between the woman-specific unmeasured components for marital dissolution and childbearing, suggesting that women who are more likely to experience marital breakdown are also less likely to have a child within marriage. Another US study that considers nonmarital fertility finds evidence of a *positive* residual correlation between the hazards of marital breakdown and of an extra-marital birth. Upchurch et al. (2002) report that women with an above average risk of marital separation tend also to have an above average probability of having a birth outside marriage. In Britain, when no distinction is made between marital and nonmarital fertility or between marriage and cohabitation, Aassve et al. (2004) report a positive correlation between the hazards of a union dissolution and of a conception. While none of these studies consider the transition from cohabitation to marriage, Brien et al. (1999) model entry into marriage (either directly from being single or via cohabitation) jointly with nonmarital childbearing (within either cohabitation or a non co-residential relationship) and find evidence of a positive cross-process correlation. Aassve et al. also report a positive residual correlation between the processes of union (marriage or cohabitation) formation and fertility.

RESEARCH QUESTIONS

In this study, we explore the following research questions:

1. What is the impact of the presence and characteristics of children on transitions from marital and cohabiting unions? Specifically, we ask the following research questions about the effects of having children on the risk of partnership dissolution: Does the presence of children promote the stability of unions? Does this differ by whether the partnership is a marriage or cohabitation? Are different effects observed for children from the current union and those from a previous relationship? Do effects differ for young (preschool) and older children? In addition, we explore whether the presence of children, from the current or a previous partnership, hastens or delays the transition to marriage by cohabiting couples.

2. Does the legal status of a partnership affect the chances of the couple having a(nother) child, and do the factors influencing the hazard of a conception depend on union status?

3. Are there unobserved factors that influence both the stability of a woman's partnerships and her fertility, which result in selection of women who are more or less prone to stable partnerships into childbearing? If so, what is the direction of any selection effect? For example, does Lillard and Waite's (1993) finding that unobserved factors lead to a negative residual association between the risk of marital dissolution and the chance of

conception during marriage extend to this British cohort, and is the nature of selection the same for cohabitation? We explore these questions by examining estimates of the pairwise correlations between woman-specific random effects in a partnership transition equation and random effects in a fertility equation, that is, the *cross*-process residual correlations. Where there is evidence for selection, we investigate how this affects estimates of the relationship between the presence of children and different partnership outcomes.

4. Is there residual correlation between the hazards of different types of partnership outcome? For example, one might expect that individuals with a high risk of marital dissolution will also have a high risk of separating from a cohabiting partner, and that this correlation will not be entirely explained by covariates included in the model. Similarly, it is of interest to examine the correlation between conception intervals within marital and nonmarital unions. As couples are likely to vary in their reasons for living together rather than marrying, cohabiting couples are expected to be a more heterogeneous group than married couples; this would lead to differences in decisions about if and when to have children across the two groups, and therefore a small correlation between marital and nonmarital conception intervals. We use estimates of the *within*-process residual correlations to investigate these questions.

METHODOLOGY

The multiprocess model is a system of simultaneous equations for partnership transitions and childbearing. Each equation defines a discrete-time hazards model. While other applications of multiprocess modelling use continuous-time models, we use a discrete-time formulation for two main reasons. First, as with many retrospectively collected event history data, the dates of events are reported in months. It is therefore natural to specify a model that assumes measurement in discrete rather than continuous time. Second, after restructuring the data, standard methods for analysing discrete response data may be used. Thus complex event history models, such as the one described below, may be fitted using existing estimation procedures. A potential disadvantage is the size of the dataset generated, which may lead to long estimation times for complex models; this issue is further discussed below.

A common feature of event history data is that events may occur more than once to an individual. In the present case, for example, women may have more than one partner and more than one conception during the observation period. Repeated events lead to a two-level hierarchical structure, with events nested within individuals. The durations between events for the same individual may be correlated due to the presence of unobserved individual-specific risk factors that affect the occurrence of each event. To allow for the dependency between event times within individuals we use multilevel models, in which unobserved characteristics are represented by random effects (Goldstein 2003).

Model for Partnership Transitions

A partnership is defined as a continuous period of at least one month spent living with the same partner. The unit of analysis is a partnership episode which is the period spent in the same partnership state, marriage or (unmarried) cohabitation, with the same partner. Thus a marriage preceded by premarital cohabitation with the same partner counts as two episodes, while direct marriage counts as a single episode. We consider three partnership transitions: marriage to separation, cohabitation to separation, and cohabitation to marriage with the same partner. In each case, we model the hazard of a partnership transition as a function of partnership duration, outcomes of previous partnership episodes, outcomes of the related childbearing process, observed background characteristics and unobserved time-invariant characteristics.

Marriage

We denote by $h_{ij}^{PM}(t)$ the hazard of a marital separation during time interval *t* of episode *i* for individual *j*. Following Steele, Diamond and Amin (1996), a multilevel discrete-time event history model for marital separations may be written (omitting subscripts) as:

$$\text{logit}\, h^{PM}(t) = \alpha_0^M D^{PM}(t) + \alpha_1^M F(t) + \alpha_2^M X^{PM}(t) + u^{PM} . \qquad (1)$$

$\alpha_0^M D^{PM}(t)$ is the baseline logit-hazard which is a function of marriage duration at time *t* or, for marriages preceded by a period of cohabitation, partnership duration. Possible choices for the baseline logit-hazard include a step function, where the duration is treated as a categorical variable, or a polynomial function. The potentially endogenous time-varying outcomes of the fertility process which may affect future partnership transitions and fertility are denoted by $F(t)$, with coefficient vector α_1^M. Other covariates which affect marital dissolution are represented by $X^{PM}(t)$. The specification in (1) is a proportional hazards model, but non-proportional hazards may be accomodated by including interactions between partnership duration and the covariates. Unobserved time-invariant individual-specific factors are represented by normally distributed random effects u^{PM}.

In order to estimate (1) each marriage duration, D_{ij}^{PM}, is converted to a sequence of D_{ij}^{PM} binary responses, $y_{ij}^{PM}(t)$. For *t*=1, . . ., $D_{ij}^{PM}-1$, $y_{ij}^{PM}(t)$=0; and for *t*=D_{ij}^{PM}, $y_{ij}^{PM}(t)$=1 if separation occurs at duration D_{ij}^{PM} and $y_{ij}^{PM}(t)$=0 otherwise (right-censored durations). As start and end dates of episodes were recorded to the nearest month, it is possible to have a binary response for each month. However, using discrete time intervals of one month leads to a very large dataset, which increases model fitting times. For example, an individual who married (directly) in January 1980 (at age 21) and was still married when they were interviewed in January 2001 (at age 42) would contribute 253×2=506[2] records to the dataset. One strategy to reduce the number of records generated is to group discrete time intervals. In our application, partnership durations (and conception intervals) are grouped into six-month periods and weighted by exposure time. For each six-month interval, a weight is defined as the number of months during that interval for which the individual was exposed to the risk of an event. These weights form denominators for the binary responses, leading to binomial response data. If it is reasonable to assume that the hazard function and covariate values are constant within each six-month period, grouping will not lead to loss of information. If more than one partnership event occurs within any six-month interval, a new interval begins after each event. We also note that the time-varying indicators of conceptions and the presence of children denote fertility status *before* the occurrence of any partnership transition which it might affect. To explore whether aggregation of time intervals leads to a loss in accuracy which might affect our conclusions, three-month intervals were considered in preliminary analyses of partnership transitions. The results, including estimates of the effects of prior fertility outcomes, were very similar to those obtained using six-month intervals.

Cohabitation

We consider two transitions from the cohabitation state: separation and marriage to the same partner. Denote by $h_{ij}^{PC(r)}(t)$ the hazard of a transition of type r from cohabitation, in time interval t of episode i for individual j, where r=0 (no transition), 1 (separation), or 2

[2] There would be 253 months of observation, and two responses per month (as described below).

(marriage). Transitions from cohabitation may be modelled using a multilevel discrete-time competing risks model (Steele, Diamond and Wang 1996):

$$\log\left[\frac{h^{PC(r)}(t)}{h^{PC(0)}(t)}\right] = \alpha_0^{C(r)} D^{PC(r)}(t) + \alpha_1^{C(r)} F(t) + \alpha_2^{C(r)} X^{PC(r)}(t) + u^{PC(r)}, \quad r = 1,2 \qquad (2)$$

where $\alpha_0^{C(r)} D^{PC(r)}(t)$ is a function of cohabitation duration at time *t*, $X^{PC(r)}(t)$ are covariates that affect the hazard of a transition of type *r* from cohabitation, and $u^{PC(r)}$ are individual and transition-specific random effects.

To estimate (2) each cohabitation duration, D_{ij}^{PC}, is converted to a sequence of multinomial responses, $y_{ij}^{PC}(t)$. The response at time *t* is coded 0 if still cohabiting, 1 if separation occurs, and 2 if marriage to the same partner occurs.

Equations (1) and (2) define a multilevel, multistate model (Goldstein, Pan and Bynner 2004; Steele, Goldstein and Browne 2004), where in the present case the states are marriage and cohabitation. To allow for unobserved individual-level characteristics that affect each type of transition, the random effects may be correlated across transitions with covariance Ω_u^P. Simultaneous estimation of (1) and (2) is achieved by pooling all episodes and defining indicator variables for marriage and cohabitation. These indicators are interacted with the explanatory variables to allow for marriage- and cohabitation-specific effects of partnership duration, fertility outcomes and background characteristics. The coefficients of the indicators themselves are allowed to vary randomly across women to produce the state-specific random effects. Further details are given in Steele et al. (2004).

Model for Fertility within Partnerships

Denote by $h_{ij}^{FM}(t)$ the hazard of a conception leading to a live birth[3] within marriage during time interval t in partnership episode i for individual j. We denote by $h_{ij}^{FC}(t)$ the hazard of a conception within a cohabiting union. The model for fertility consists of separate equations for marriage and cohabitation, which are estimated simultaneously.

Marriage

A multilevel event history model for the waiting time to conception within marriage may be written (omitting subscripts):

$$\text{logit}\, h^{FM}(t) = \beta_0^M D^{FM}(t) + \beta_1^M F(t) + \beta_2^M X^{FM}(t) + u^{FM} \tag{3}$$

where $\beta_0^M D^{FM}(t)$ is a function of the partnership duration and, for second or higher order births, the duration since the previous birth, $X^{FM}(t)$ are covariates affecting the fertility process, and u^{FM} is an individual-level random effect.

[3] A conception date is calculated as the date of birth minus nine months. Still births and pregnancies that end in abortion or miscarriage are not considered for two reasons. First, these pregnancy outcomes do not lead to the presence of children which can affect partnership transitions. Second, data on abortions and miscarriage are likely to be incomplete. Berrington (2001) compared abortion rates calculated from the NCDS to age 33 with national rates and found that the NCDS figures were underreported by 50%. Because conception outcomes are unknown for women pregnant at the time of interview, the last nine months of the observation period are omitted from our analysis of fertility and partnership transitions.

Cohabitation

The model for conceptions within cohabitation is written:

$$\operatorname{logit} h^{FC}(t) = \beta_0^C D^{FC}(t) + \beta_1^C F(t) + \beta_2^C X^{FC}(t) + u^{FC} \tag{4}$$

where $X^{FC}(t)$ are covariates and u^{FC} is an individual-level random effect, which may be correlated with u^{FM} with covariance Ω_u^F.

Estimation

Equations (1), (2), (3) and (4) define a multiprocess model. These equations must be estimated simultaneously as there may be non-zero correlations between the woman-specific random effects across equations. We assume that $u = (u^{PM}, u^{PC(1)}, u^{PC(2)}, u^{FM}, u^{FC}) \sim N_5(\mathbf{0}, \Omega_u)$. Correlated random effects would arise if the unobserved characteristics that influence the timing of partnership transitions are correlated with those that affect childbearing within partnerships. Non-zero correlations between elements of $u^P = (u^{PM}, u^{PC(1)}, u^{PC(2)})$ and of $u^F = (u^{FM}, u^{FC})$ would suggest that any or all elements of $F(t)$ are endogenous with respect to partnership transitions.

The discrete-time multiprocess event history model can be framed as a multilevel bivariate discrete response model where for each time interval *t* of a partnership there are two responses: 1) a binary or multinomial response for the partnership status, and 2) a binary response indicating the occurrence of a conception. The model may be estimated using existing methods for mixtures of binary and multinomial responses (Steele et al. 2004) after defining indicators for the partnership and fertility responses and interacting these with the duration variables and covariates. The results presented in this paper were obtained using

Monte Carlo Markov chain (MCMC) estimation, as implemented in *MLwiN* (Rasbash et al. 2004)[4].

Identification

Identification of simultaneous equations models typically requires exclusion restrictions to be placed on the covariates. In the case of the model described above, this involves including in the fertility equations a set of covariates that are excluded from the partnership transition equations. In practice, however, it is difficult to find, and justify on theoretical grounds, variables that affect fertility decisions but do not have direct effects on partnership behaviour. Fortunately, the observation of repeated events for a subset of women, with some overlap in events across processes, means that identification is possible without covariate exclusions. The model is identified under the assumption that all residual dependency between processes can be accounted for by allowing cross-process correlation between individual-level residuals that are constant across replications for the same individual. After accounting for this

[4] MCMC methods are used to estimate statistical models in a Bayesian framework. In the Bayesian approach, each unknown parameter in the model is viewed as a random variable with an associated probability distribution that incorporates any prior beliefs about the value of that parameter. MCMC methods are simulation-based procedures in which a chain of random draws is taken from the current conditional probability distribution for each parameter. A point estimate of a parameter may be obtained by taking the mean, median or mode of the parameter values across the chains, while the standard deviation of parameter values corresponds to a frequentist standard error. It is usual to calculate these summary statistics after a "burn-in" period, which involves discarding initial draws before the chains have converged. See Browne (2003) for an introduction to MCMC methods for multilevel analysis.

residual correlation, the remaining variation in the fertility outcomes $F(t)$ between partnership episodes represents the effects of prior fertility on the outcomes of marriage and cohabitation, controlling for selection bias. Identification of similar multiprocess models is discussed further in Lillard et al. (1995) and Upchurch et al. (2002). In both studies, sensitivity analyses were carried out and the authors' conclusions were unaffected when covariate exclusions were imposed.

The assumption that the woman-specific residual terms account for all sources of correlation between processes implies that any selection effect is due to unmeasured characteristics of the women that remain constant between the ages of 16 and 42. This does not imply that a woman's attitudes and aspirations regarding partnerships and having children are fixed over time; rather, we assume that any such change that leads to a shift in the hazard of a partnership transition or the conception risk is due to observed circumstances, including changes in the age and number of children or in education level. Our model does not control for changes in unobservables which affect both the partnership and fertility processes (see also Upchurch et al. 2002).

DATA AND VARIABLES

The National Child Development Study

We analyse data from the National Child Development Study (NCDS), a prospective longitudinal study of all those living in Great Britain who were born in a single week of March 1958 (Shepherd 1997). Following the initial birth survey, there have so far been attempts to contact members of the birth cohort on six further occasions, corresponding to ages 7, 11, 16, 23, 33 and 42. Information was collected from parents and then cohort members, and also from teachers, school health services and other supplementary sources.

The NCDS therefore provides a rich source of data on physical, educational and social development over the life course.

Retrospective partnership and birth histories were collected in 1981, 1991 and 2000, when the respondents were age 23, 33 and 42. At age 33, respondents were asked for the start and end dates of all co-residential relationships between the ages of 16 and 33 which lasted for at least one month. Information was collected using a self-completion questionnaire and face-to-face interview. These data were later reconciled to form a single partnership history which also incorporated data collected at age 23 (Di Salvo 1995). At age 42, information was collected on partnerships since age 33 (or since age 16 if the cohort member was not interviewed in 1991). One task of the current study was to link the age 16-33 and age 33-42 data to form a continuous partnership history from age 16 to 42 (Kallis, forthcoming). Similarly, a continuous birth history was derived from retrospective data collected at age 33 (Di Salvo 1995) and at age 42 (Dodgeon 2002).

Variable Definitions

As described above, the dependent variable for partnership transitions is an indicator for each six-month interval of whether a transition has occurred and, in the case of cohabitation, the type of transition. The durations of partnerships ending due to the death of a partner are treated as right-censored, as are partnerships that are still intact at the time of interview. The dependent variable for childbearing within partnerships is a binary indicator, coded 1 for months during which conceptions occurred and 0 otherwise.

The explanatory variables of major interest are outcomes of the fertility process. Respondents were asked to identify the father of each child and for the date that each child left home. Thus it was possible to create time-varying counts of the number of children living with a woman, distinguishing between preschool and older children, and between children born to the current partner at time *t* and those fathered by a previous partner or a non-coresident partner. In addition to variables relating to previous births, the indicators of prior fertility include a binary indicator for current pregnancy status (for pregnancies that result in live births).

Also included as explanatory variables were characteristics of the current and previous partnerships. These include an indicator of the legal status of the current partnership (marriage or cohabitation), the woman's age at the start of the partnership, and indicators of previous marriage and cohabitation. We considered a range of background characteristics found to be important in earlier studies of partnership transitions and childbearing, based on the NCDS (e.g. Berrington and Diamond 1999; Kiernan and Cherlin 1999) and other British data (e.g. Ermisch and Francesconi 2000). These included the number of years of post-compulsory education (treated as time-varying) and measures of the respondent's family background: the father's social class at the respondent's birth, region of residence[5], the experience of family disruption by age 16, and housing tenure at age 16[6]. The cumulative years of education was created from longitudinal data on educational enrollment and employment collected retrospectively at ages 33 and 42. The measure of social class is based on the occupation of the father, or of the father figure if the natural father was not present. After exploratory analysis, social class was categorised as: 1) professional, managerial and technical, 2) skilled (manual or non-manual), 3) partly skilled and unskilled, and 4) no father

[5] Following preliminary analysis which found no evidence of regional variation in any partnership or fertility outcome, region was excluded from further analysis.

[6] Housing tenure was classified as owner-occupied versus any other type of accommodation (mainly rented). This variable was removed from the final models due to a large amount of missing data, even after imputation, and weak or nonsignificant associations with partnership transitions and fertility.

figure[7]. The indicator of family disruption is derived from data collected at age 33 on whether the respondent's parents had ever separated or divorced and, if so, the respondent's age when her parents last lived together. This information was supplemented by data collected at birth and at ages 7, 11 and 16 on the relationship between the respondent and those household members who were identified as the mother and father figures. If, at any of these ages, a respondent was living with a mother or father figure who was not their natural or adoptive parent, they were classified as having experienced family disruption. Descriptive information on the partnership and background characteristics considered in the models is given in Table 1.

The Analysis Sample

In common with most other studies of fertility, our analysis is restricted to women. There are two main reasons for this decision. First, we expect some unreliability in men's reports of children from previous relationships, particularly nonmarital births. In an evaluation of men's retrospective fertility histories from the British Household Panel Study, Rendall et al. (1999) found that only 60 of men's births from a previous marriage were reported per 100 of women's births, with women's fertility reports matching birth registration statistics. If men's reports are unreliable, this could lead to substantial measurement error in the duration

[7] Where father's social class at birth was missing, and it was known that a father figure was present in the household, imputation was carried out subject to the following constraint. Social class collected at a later age was imputed only if it was reasonable to assume that the father figure at that age was the same person as the father figure at birth. Thus imputation was restricted to cases where the father figure was the natural or adoptive father on both occasions. Missing values on social class, and other background characteristics, that could not be imputed are represented in the analysis by 'missing' dummy variables.

between conceptions and time-dependent indicators of the number of children from the current and previous partnerships. Second, children who were not the biological offspring of the respondent were omitted from the time-varying count of co-resident children; this was necessary because longitudinal information on the presence of stepchildren was not collected. Cross-sectional data on household composition show that the omission of stepchildren would be more severe for men than for women, which is as expected given that children usually remain with their mother after partnership dissolution.

A total of 6489 women were interviewed at age 33 and/or age 42. Of these, women who were interviewed only at age 42 (10.6%) had to be excluded because, due to time limitations of the interview, they were not asked about their fertility before age 33. The analysis is based on the 93% of women who had partnered at least once by age 33[8]. Of those, 1.3% of cases for whom no accurate event history could be constructed were dropped (see Kallis, forthcoming, for details). The analysis sample is further restricted to women for whom all types of partnership transition were possible and who were able to have biological children. Based on these criteria, women with a female partner and childless women who had been told by a doctor that they could or should not have children were excluded. Women with adopted children were also omitted because it is likely that their conception intervals would have been affected by the adoption and it is possible that they or one of their partners was infertile. A further three cases were dropped because of missing covariate information after imputation. The final analysis sample contains 5142 women, who contribute 7032

[8] Although a small proportion of women were observed to have their first partnership between ages 33 and 42, they are excluded from the analysis. Because of some non-response at age 42, we do not have histories for all women between these ages. We therefore use exclusion criteria that treat respondents and nonrespondents equally to avoid selection bias.

partnerships and 9137 partnership episodes. These women conceived 9313 children while in a partnership.

The NCDS, in common with any longitudinal study, suffers from attrition. At age 33, for example, 29% of the target sample had dropped out of the study. Hawkes and Plewis (2004) investigated the impact of attrition bias on the NCDS by considering the relationship between attrition and a limited range of demographic and socio-economic factors; they concluded that cohort members “with lower educational attainments, less stable employment patterns and living in more disadvantaged circumstances” were more likely to be lost from the study, but found that none of these factors are strong predictors of attrition. By including education and parental social class in our models, we adjust for much of the non-response bias due to these factors. Of course, this does not preclude the possibility that non-response is related to the processes under study, namely, partnership outcomes and childbearing. Our analysis is based on women interviewed at age 33, of whom 12% were not interviewed at age 42. Thus the partnership and birth histories of the women who were not present at age 42 are treated as right-censored at age 33. In treating the processes of age 42 non-respondents as censored we assume that, conditional on partnership transitions and childbearing before age 33 (and other characteristics included in the models), partnership and fertility behaviour after drop-out would have been similar to the experiences of women who remained in the sample. This assumption may be questionable if, for example, among women with similar partnership patterns up to age 33, those who dropped out at age 42 went on to have multiple partners after age 33, while those who did not drop out had few partners. While we cannot discount such situations, we believe that they introduce minimal attrition bias to our analysis because partnership histories up to age 33 are generally good predictors of behaviour thereafter.

RESULTS

Descriptive Analysis of Partnership Transitions and Fertility

In this section we present preliminary analysis of partnership transitions and childbearing, starting with an examination of the timing of separation and moves from cohabitation to marriage. From a lifetable analysis of the distribution of the number of years to each type of transition, we find that about three quarters of marriages were still intact after 12 years. Transitions from cohabitation tend to occur much earlier suggesting that, at least for this cohort, marriage is seen as a longer-term commitment than cohabitation. A quarter of cohabitations end within 3.5 years and a further quarter are converted to marriage after just over a year. Both the hazard of separation and the hazard of a move from cohabitation to marriage increase in the early stages of a partnership, and then decrease rapidly. For marriage the hazard of separation peaks at around four years, while for cohabitation the risk of separation is highest at about 3.5 years and the hazard of marriage peaks at three years. The forms of these hazard functions suggest that the effects of partnership duration can be represented by cubic polynomials in the models for each transition.

An analysis of the distribution of the total number of marriages and cohabitations for each woman reveals that almost half of the women in the sample have been married only once with no periods of cohabitation. A further quarter have had one marital partner, but have also cohabited with the same or a different partner. Almost a quarter of women have experienced two or more marriages and/or cohabitations, highlighting the need for a multilevel approach to allow for correlation between durations of partnerships contributed by the same woman.

Table 2 contains additional information on the sequence of transitions for the most common combinations of marriages and cohabitations among women with at least two partnership episodes. For the women who have exactly one cohabitation and one marriage in total, we see that four out of five moved from cohabitation to marriage with the same partner. Cohabitation followed marital separation for 17% of this group of women. For women who experienced two marriages and one cohabitation, the vast majority cohabited after marital separation and subsequently married their second partner. Of those women who had cohabited twice, almost three-fifths began with a cohabiting partnership which later dissolved, and then lived with their second partner before marriage; a further 29% married their first partner after premarital cohabitation, and cohabited with their next partner following the breakdown of their marriage.

Turning to childbearing, we find that the majority of cohabiting unions were childless (Table 3). Further analysis reveals that a conception occurs within five years of the start of an episode for 36% of cohabitations and 70% of marriages. Of those partnerships that produced children, there was more than one conception in 53% of marriages and 5% of cohabitations; the multilevel model allows for correlation between the duration of conception intervals within the same partnership episode that is due to woman-specific unobservables.

Table 4 shows the distribution of fertility outcomes according to partnership status at ages 30 and 40. By age 30, most married women have had at least one child with their current (for most, the only) partner, and a small proportion has children fathered by a previous partner. A markedly different pattern is observed among women who were cohabiting at their 30th birthday. As noted above, relatively few children are conceived in nonmarital unions, but cohabiting women are considerably more likely than married women to have children from a previous partner. The main change in fertility outcomes between ages 30 and 40 is, as expected, an increase in the proportion with school age or older children.

Multilevel Event History Analysis

We consider two specifications of a multilevel event history model. The first model is a single process model, where partnership transitions and fertility within partnerships are modelled as separate processes. The three partnership transition equations are still estimated jointly, as are the equations for fertility within marriages and cohabitations, but the random effects across processes are assumed to be uncorrelated, i.e. the pairwise correlations between each element of u^P and an element of u^F are constrained to zero. This model assumes that prior fertility outcomes are exogenous with respect to partnership transitions. The second model considered is a multiprocess model in which the correlations between u^P and u^F are estimated freely. A correlation that is significantly different from zero provides evidence that prior fertility outcomes are endogenous, in which case the estimated effects from the single process model will be biased. The nature of this bias will depend on the direction of the correlation and on the direction of the effect of the endogenous prior fertility outcome on the risk of the partnership transition.

<u>Correlations between random effects</u>

The estimated random effects covariance matrix obtained from the multiprocess model is shown in Table 5. There is substantial unobserved heterogeneity in the hazards of all partnership transitions and in the hazards of conceptions within partnerships. Of most interest, however, are the covariance terms, several of which differ significantly from zero.

Among partnership transitions, the random effect for marital separation is positively correlated with the random effect for separation from cohabitation (r = 0.52); this suggests that women with above average propensities of marital separation ($u^{PM}>0$) will tend also to have above average propensities to separate from a nonmarital union ($u^{PC(1)}>0$). Although not significant at the 5% level, there is also a positive correlation between the random effects for marital separation and for the transition from cohabitation to marriage. Allowing for this particular correlation has a substantial impact on the estimate of the effect of premarital cohabitation on the risk of marital separation. In a model where this correlation is assumed to equal zero, i.e. when the equations for marital separation and the outcomes of cohabitation are estimated independently (results not shown), we observe a strong positive association between premarital cohabitation with the current partner and the risk of dissolution. This apparently positive effect of premarital cohabitation has been found in previous studies that have analysed marital dissolution separately from the outcomes of cohabitation (e.g. Berrington and Diamond 1999), an effect which a US study has shown to be due to selection of individuals with a high propensity to divorce into cohabitation (Lillard, Brien and Waite 1995). The positive, albeit nonsignificant, residual correlation that we observe between the hazards of marital separation and the transition from cohabitation to marriage could be picking up this form of selection: the estimate of this correlation is based on women who have experienced at least one marriage and cohabitation, the majority of whom have had a marriage preceded by a period of premarital cohabitation with the same partner (Table 2).

The correlation between the random effects for marital fertility and fertility within cohabiting unions is small and nonsignificant, implying that the unobserved time-invariant characteristics that influence a woman's chance of conceiving during marriage are uncorrelated with those that influence her chance of conceiving during cohabitation. We therefore conclude that the unobserved factors driving fertility decisions differ according to the type of union.

Across the partnership and fertility processes, we find significant residual correlations between the hazard of a conception within a cohabiting union and the hazards of all three partnership transitions. There are positive correlations between the chance of conceiving during cohabitation and the risk of dissolution from both marital and cohabiting unions (r=0.425 and r=0.316 respectively). The corresponding correlations with the chance of a marital conception are not significantly different from zero. Taken together, these findings suggest that women prone to unstable partnerships tend to have a high chance of conceiving during cohabitation, but are no more or less likely to conceive during marriage. Although we cannot infer the direction of causality from these correlations, we note that the majority of cohabitations in this dataset, and cohabiting conceptions, precede marriage; therefore a plausible interpretation of the positive correlation with marital separation is that the type of woman who gets pregnant during cohabitation (where 'types' are defined by unobserved characteristics) tends to experience a higher risk of dissolution should she later marry.

There is also a significant correlation between the unobserved woman-specific factors affecting the hazard of converting a cohabitating union into marriage and those affecting the hazard of a conception within cohabitation (r=0.299). This positive correlation persists after controlling for the increased odds of marriage experienced during pregnancy (see Table 8), although the inclusion of current pregnancy status did reduce the correlation estimate from 0.425. A possible interpretation of this correlation is that women who view cohabitation as a precursor to a more formal marital partnership (and therefore have a high probability of marriage) have an increased chance of initiating childbearing during cohabitation, in anticipation of marriage.

Including current pregnancy status as an explanatory variable in the partnership transition equations also affects the estimate of the residual correlation between the risk of marital dissolution and conception intervals within marriage. Before adjusting for the substantially reduced dissolution risk during pregnancy (Table 6), we find a significant, negative residual correlation between the risk of marital dissolution and conception intervals within marriage (r=-0.279). A negative correlation implies that women with a below average risk of separation, i.e. long marriages, have an above average risk of conceiving during marriage. After accounting for the negative effect of current pregnancy status on the risk of marital separation, this correlation decreases in magnitude to –0.071 and is no longer significant at the 5% level. This result may suggest that the negative residual correlation observed before controlling for pregnancy status is due to the stabilizing effect of pregnancy on marriage, rather than a selection of women with a tendency towards stable marriages into childbearing.

Effects of prior fertility outcomes

Tables 6 to 8 show estimates of the effects of prior fertility outcomes and other covariates on the hazards of partnership transitions. The single process estimates are from a model which assumes that, conditional on covariates, the processes of partnership transitions and childbearing are independent. The multiprocess estimates are from a model which allows for the dependence between processes via correlated random effects. Of prime interest in this study are the effects of prior outcomes of the fertility process on the odds of separation and, for cohabiting women, of marriage. A comparison of estimates from the single and multiprocess models demonstrates the effect of allowing for the endogeneity of the presence of children with respect to partnership transitions.

The estimated coefficients from the equation for marital separation are very similar for the single and multiprocess models (Table 6). In the case of fertility outcomes from marriage, including current pregnancy status and children born to the current partner, this is expected due to the non-significant residual correlation between the hazards of a marital separation and a marital conception (Table 5). While the positive residual correlation between the hazards of marital separation and a conception during cohabitation could have led to bias in the estimated effects of having children from a previous partnership, there is in fact little change in the estimates between models. The results from both models show a decreased risk of marital dissolution during pregnancy, and following the birth of the child. The presence of preschool age children fathered by the current husband reduces the risk of marital separation, and this effect strengthens with the number of young children. Older children born to the current partner also have a negative, though weaker, effect on separation. The presence of children from a previous partnership does not significantly affect the risk of marital dissolution. However, there is a strong positive association between having a child with a non co-resident partner and the risk of separation.

Allowing for the endogeneity of fertility within partnerships has some impact on estimates of the effects of pregnancy and the presence of children on the risk that a cohabiting union dissolves (Table 7). While we would conclude from either model that expecting a child or having young children together reduces the risk of separation for a cohabiting couple, the effects obtained from the multiprocess model are slightly stronger, due to the positive residual correlation between the dissolution risk and the odds of a conception during cohabitation (Table 5). In the single process model, the negative effects of pregnancy and the presence of children are subject to selection bias. Women with a high risk of separation are more likely to conceive during cohabitation; these women increase the risk of separation for women who are pregnant or who have children with a cohabiting partner, leading to an understatement of the 'true' negative effect of these particular fertility outcomes. The results from both models show that the presence of school age or older children fathered by the current partner, or children of any age from a previous partnership or non-coresidential relationship, does not significantly affect the risk of dissolution.

Turning to the effects of prior fertility on the probability that a cohabiting woman marries her partner (Table 8), we again find some small differences between the results of the two model specifications. From either model, we would conclude that cohabiting couples who are expecting a child together have an increased odds of marriage, but that the odds decrease after the birth. However, the positive effect of current pregnancy status is weaker and the negative effect of having children is stronger in the multiprocess model. This change in the estimates is due to the positive correlation between the random effects for the transition from cohabitation to marriage and conceptions within cohabitation (Table 5). On average, a high propensity to marry a cohabiting partner is associated with a high propensity to have a child during cohabitation. If this form of selection is ignored, the estimated odds of marriage for women who have had children with their current (cohabiting) partner will be inflated. The presence of children fathered by a previous co-resident partner does not affect the transition to marriage, but having children from non-coresident relationships is significantly associated with a decreased probability of marriage.

Table 9 shows estimates from the fertility component of the multiprocess model[9]. For both married and cohabiting women, those who have one preschool age child with their current partner have an increased hazard of having further children, while married couples with two or more delay or stop childbearing. Married couples with older children together have a markedly lower conception hazard than childless couples. Taken together, these results suggest a preference towards having two children with a partner, with the second child conceived while the first is less than five years old; these effects are particularly strong for married women. For married women, the effects of children from a previous partnership depend on their age; having one or more young children increases the probability of having another, while having an older child reduces the hazard of conception. A prior birth from a non-coresident relationship reduces the odds of conception within marriage.

[9] As in the analysis of partnership transitions, single and multiprocess models of fertility were considered. As the results obtained from these alternative specifications were very similar, we present only estimates from the multiprocess model.

Effects of other covariates

The effects of most background characteristics on the risk of separation from either marriage or cohabitation are similar for the single and multiprocess models (Tables 6 and 7). The woman's age at the start of a partnership has the strongest effect, with older women experiencing lower risks of separation from either form of partnership. Being more educated is negatively associated with the risk of separation from marriage, but not from cohabitation. While paternal social class does not have a significant effect on marital dissolution, having a father in a professional, managerial or technical occupation increases the odds of separation for cohabiting women. There is also evidence that cohabitations are more likely to dissolve if the respondent experienced family disruption during childhood. Turning to the effects of the outcomes of previous partnership transitions, we find that both marriages and cohabitations of women who have been previously married are longer in duration than partnerships for women who have not yet experienced a marital breakdown. Compared to married women who have never cohabited, those for whom the only experience of cohabitation is with their current partner are no more likely to separate. Women who have cohabited with a previous partner, however, have an increased risk of separation. Previous cohabitation has no effect on the risk of dissolution for cohabiting partnerships.

Cohabiting women who have previously been in another co-residential union, whether marriage or cohabitation, are less likely to marry than those in their first partnership (Table 8). A lower marriage rate is also observed among women who started to cohabit in their early thirties, rather than their teens or twenties. None of the other characteristics considered has a significant effect on the odds of marriage.

The effects of observed background characteristics on conception intervals are rather different for married and cohabiting women (Table 9)[10]. Further, the lack of correlation between the residuals for conceptions within cohabitation and marriage noted earlier suggests that there is little overlap in the unobserved determinants of these processes. The hazard of having a(nother) child declines as the age at the start of both marital and nonmarital partnerships increases, but the effect is considerably stronger for cohabiting women. The odds of a conception within cohabitation decline with increasing education, while education has little effect on the duration between marital conceptions. Although paternal social class has an effect on the hazard of a conception for cohabiting women, with longer intervals for the professional, managerial and technical classes, the effect for married women is not significant.

DISCUSSION

Regarding the research questions posed earlier, our findings are as follows.

1. *What is the impact of the presence and characteristics of children on transitions from marital and cohabiting unions?* We find some support in this cohort for the notion that children cement a partnership. The effect on union dissolution of having preschool children is similar for married and cohabiting couples, though the effect is somewhat weaker for children born to cohabitors. This stabilizing effect of children weakens as they get older, and

[10] It is usual in discrete-time event history analyses of birth intervals to include as a covariate the current duration since the last birth (or some suitable origin in the case of the first birth). This is equivalent to including the age of the youngest child (which is represented in our models by the number of preschool and older children) and the duration of the current partnership. After controlling for these variables, we find no additional effect of the number of months since the last birth.

becomes nonsignificant for cohabiting couples. There is strong evidence that pregnancy precipitates marriage among cohabitors, but following a birth the marriage rate declines to below the pre-pregnancy level.

Children most at risk of parental separation are older, with a mother who had cohabited with previous partners, has little post-compulsory education and has started her first partnership in her teens or early twenties. Children living with a stepfather are at particular risk of experiencing the breakdown of their mother's marriage if they were born outside any co-residential relationship. The partnerships of married or cohabiting mothers who have been previously married are less precarious. Reconstituted families are a minority of all families, but they require more complicated support and contact arrangements if they then split up. The risks of early family formation may not have been well appreciated by those who embarked upon it. These results suggest they might have been averted with more education in general and about these risks in particular. For those families at particular risk of dissolution, policy should be sensitive to this fragility.

2. *Does the legal status of a partnership affect the chances of a couple having a(nother) child, and do factors influencing the hazard of a conception depend on union status?* Overall, cohabitors, in the cohort we observe, were less likely than married couples to bear children, and women who get pregnant during cohabitation are likely to marry before the birth. We also find that childbearing within cohabitation is more sensitive to the woman's age at the start of the partnership and her education level than childbearing within marriage. Conceptions during cohabitation occur mostly among women with little post-compulsory education and those who started to live with their partner at a young age.

3. *Are there unobserved factors that influence both the stability of a woman's partnerships and her fertility, which result in selection of women who are more or less prone to stable partnerships into childbearing? If so, what is the direction of any selection effect*

and what is the bias in estimates of the effect of having children on partnership transitions? After controlling for the effects of current fertility status, partnership history and background characteristics, we find positive residual correlation between the risk of partnership dissolution, for both marriage and cohabitation, and the odds of a cohabiting conception. This implies that women who are prone to unstable partnerships have an above average risk of conceiving during cohabitation. A positive residual correlation between the processes of marital dissolution and fertility during cohabitation was also found by Upchurch et al. (2002), although they include births with non co-resident partners in their definition of nonmarital fertility. We suggest that this correlation may be picking up a negative effect of getting pregnant during cohabitation on the dissolution risk of a subsequent marriage, i.e. the instability of 'shotgun marriages'. We find that a moderate negative residual correlation between the hazards of marital dissolution and of a marital conception disappears and becomes nonsignificant, after adjusting for the low rate of marital dissolution during pregnancy. One possible explanation for this result is that the negative correlation observed before controlling for pregnancy status is due to a stabilizing effect of pregnancy on marriage, rather than a selection of women prone to stable marriages into childbearing. The findings of Lillard and Waite's (1993) US study provide some support for this interpretation. Based on an analysis which does not adjust for the effects of current pregnancy status, they find a very strong negative correlation between the woman-specific unobservables affecting dissolution and fertility rates which persists after explicitly controlling for a negative effect of the hazard of marital dissolution on the chance of conception.

The positive residual correlation that we find between the hazard of converting a cohabiting partnership into marriage and the hazard of a conception during cohabitation is consistent with the findings of Brien et al. (1999), who consider entry into marriage from cohabitation or an unpartnered state and use the same definition of nonmarital fertility as

Upchurch et al. (2002). Again, our estimate of the residual correlation is sensitive to the inclusion of an indicator for current pregnancy status; the positive residual correlation is partly due to the increased odds of marriage during pregnancy. Nevertheless, we still find that high odds of marriage are associated with high odds of a cohabiting conception. In a recent study, which compares the outcomes of cohabitation for the 1958 and 1970 cohorts up to age 30, we find weak evidence of a *negative* correlation among the more recent cohort (Steele et al. 2005). This suggests that there has been a shift towards women with a high propensity to marry a cohabiting partner having a low propensity to conceive during cohabitation, i.e. women in the later cohort who are favourably inclined towards marriage tend to wait until marriage before having children. This further illustrates that patterns of partnership and childbearing change across time, and presumably country.

Although the multiprocess model provides some interesting insights into the nature of the relationship between partnership outcomes and childbearing in the British 1958 cohort, allowing for these cross-process correlations has little impact on our conclusions about the effect of having children on partnership stability. It is more important to allow for changes in the rate of dissolution and marriage before, during and after pregnancy (as discussed by Blossfeld et al. 1993). In particular, if we fit a single process model and omit current pregnancy status, we would conclude that having a young child has no effect on the odds that cohabitation is converted to marriage. A negative effect is revealed either by controlling for the increased odds of marriage during pregnancy *or* by fitting a multiprocess model. Thus, the multiprocess model offers some protection against this form of model misspecification.

4. *Is there residual correlation between the hazards of different types of partnership outcome?* There is strong evidence in this study that those women with a high chance of experiencing marital dissolution also have a high chance of separating from a cohabiting partner. Therefore there are some unobserved, time-invariant, attributes which make some

women more prone to unstable partnerships, regardless of the legal status of the unions they form. We also find weak evidence of a positive residual correlation between the odds of marital separation and the chance that cohabitation leads to marriage. This correlation accounts for a positive effect of premarital cohabitation on the risk of marital breakdown; we therefore conclude that the high rate of marital dissolution experienced by couples who lived together before marriage is due to selection of those with a high risk of dissolution into cohabitation.

Turning to methodological issues, the multiprocess modelling approach we adopt allows for the endogeneity of children conceived in marital and nonmarital cohabiting partnerships, but we have not controlled for the potential endogeneity of children conceived outside co-residential partnerships. One way of allowing for the endogeneity of these children would be to additionally model the duration of episodes outside marriage and cohabitation (i.e. the process of partnership formation) and conceptions outside partnerships simultaneously with partnership durations and conceptions within partnerships. However, the correlations between the residual component of the additional fertility equation and the residuals in the other equations are likely to be poorly identified because childbearing outside marriage or cohabitation was rare for this cohort of women (see Table 4). Further research is underway to explore the determinants of both the formation and the dissolution of partnerships of a 1970 birth cohort of British women, for whom conceptions outside unions were a more common occurrence (Kallis et al. 2005).

Finally, while we present estimates of the effects of prior fertility on the hazards of partnership transitions, accounting for the joint determination of partnership outcomes and fertility, we do not estimate the reverse causal path; that is, the impact of partnership stability on the hazard of a conception. Following Lillard and Waite (1993), we considered an extended model for fertility that allows for structural effects of the hazards of partnership transitions on the hazard of a conception to explore whether partnership instability might inhibit or precipitate a birth (results not shown). Identification of this model requires covariate exclusions, specifically variables that affect partnership transitions but do not have direct effects on fertility decisions. Because partnership and childbearing decisions are so closely interlinked, however, it is difficult to find suitable instruments. Lillard and Waite (1993) successfully applied a full structural model in their analysis of marital dissolution and fertility in the US by exploiting state-level variation in divorce laws and the cost of divorce to identify the effects of marital stability on the hazard of a conception. In Britain, there is no such geographical variation that can be used for identification and, in any case, one would expect measures of divorce to have only a weak relationship with the outcomes of cohabitation. Alternative candidates for instruments were found to be only weakly correlated with the hazards of partnership transitions and/or had significant direct effects on fertility. As a result, estimates of the structural parameters based on these variables had extremely large standard errors and are likely to be biased. There is a need for further research on this issue.

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Table 1. Descriptive statistics for partnership and background characteristics included in the final models

Characteristics of current/previous partnerships	Percentage of partnership episodes	
	Marriage	Cohabitation
Age at start of partnership		
<20 years	15.2	9.5
20-24	52.3	34.6
25-29	21.0	26.4
30-34	8.4	18.4
35+	3.1	11.3
Previously married	13.8	35.7
Previously cohabited	†	27.9
Previously cohabited with . . .		
No one	60.9	†
Current partner	29.1	†
Previous partner(s)	1.5	†
Both current and previous partner(s)	8.6	†
No. episodes	5595	3542
Background characteristics	Percentage of women ever partnered by age 33	
Post-16 years of education at age 33[a]		
0	55.8	
1	13.6	
2	12.3	
3-5	11.3	
6+	7.0	
Paternal social class at birth[b]		
Professional, managerial or technical	17.0	
Skilled (manual or non-manual)	58.3	
Partly skilled or unskilled	20.9	
No father figure	2.5	
Unknown	1.3	
Family disruption during childhood	10.2	
No. women	5142	

Notes:
[a]Number of post-16 years of education is treated as a time-varying covariate in the event history models.
[b] Social class refers to the current or most recent occupation of the father (or mother's husband) at the respondent's birth.
†denotes a variable excluded from the equation(s) for transitions from this partnership state, and the corresponding fertility equation.

Table 2. Sequencing of partnership events

Order of events for each type of history	**Percentage within type of history**
1 marriage, 1 cohabitation (n=1218 women)	
Cohabitation to marriage (same partner)	79
Marriage then Cohabitation (2 partners)	17
Cohabitation then Marriage (2 partners)	4
Total	100
2 marriages, 1 cohabitation (n=375)	
Marriage then (Cohabitation to marriage)	95
3 partners	5
(Cohabitation to marriage) then Marriage	0
Total	100
2 cohabitations, 1 marriage (n=289)	
Cohabitation then (Cohabitation to marriage)	56
(Cohabitation to marriage) then Cohabitation	29
3 partners	15
Total	100

Table 3. Number of conceptions leading to a live birth per partnership episode

No. conceptions	**Cohabitation (%)**	**Marriage (%)**
0	75.9	24.8
1	18.9	21.7
2	4.0	38.3
3+	1.2	15.2
Total	100	100
No. episodes	3542	5595

Table 4. Prior fertility outcomes at age 30 and 40 by partnership status, among women in partnership at each age

	Age 30		Age 40	
	Married	Cohabiting	Married	Cohabiting
No. preschool children with current partner				
0	38.4	80.2	83.0	84.2
1	42.2	15.3	14.3	13.2
2+	19.4	4.5	2.7	2.6
No. older children with current partner				
0	57.7	91.6	17.8	76.9
1	23.3	5.9	18.3	14.6
2+	19.0	2.5	63.9	8.5
Preschool child(ren) with previous partner	0.3	5.6	0.0	0.2
Older child(ren) with previous partner	3.8	25.2	7.9	40.1
Child(ren) with non co-resident partner	1.6	5.4	1.3	8.7
n	3903	595	3390	424

Table 5. Estimated random effects covariance matrix from the multiprocess model

	Conception within cohabitation	Conception within marriage	Marital separation	Cohabitation separation	Cohabitation to marriage
Conception within cohabitation	0.224* (0.156, 0.305)				
Conception within marriage	-0.007 (-0.023, 0.010) **-0.064**	0.047* (0.039, 0.057)			
Marital separation	0.217* (0.074, 0.357) **0.425**	-0.017 (-0.062, 0.027) **-0.071**	1.162* (0.679, 1.775)		
Cohabitation separation	0.131* (0.027, 0.243) **0.316**	-0.009 (-0.045, 0.025) **-0.048**	0.486* (0.261, 0.762) **0.521**	0.766* (0.532, 1.092)	
Cohabitation to marriage	0.081* (0.013, 0.166) **0.299**	-0.008 (-0.030, 0.013) **-0.063**	0.129 (-0.027, 0.319) **0.211**	0.112 (-0.015, 0.245) **0.225**	0.324* (0.212, 0.477)

Note: The values in each cell are the point estimate (the mean of the MCMC samples) and the corresponding 95% interval estimate (the 2.5% and 97.5% point of the distribution). In off-diagonal cells a point estimate of the correlation between a pair of random effects (the mean of the correlation estimates across samples) is shown in bold. The results are based on 30000 MCMC samples, with a burn-in of 5000.

*denotes significance at the 5% level

Table 6. Estimated coefficients (and standard errors) from models of marital separation

	Single process model		Multiprocess model	
Variables	**Coeff.**	**(SE)**	**Coeff.**	**(SE)**
Prior fertility outcomes[a] (ref=none)				
Currently pregnant	-1.420**	(0.119)	-1.412**	(0.120)
No. preschool with current partner				
1	-0.514**	(0.066)	-0.510**	(0.069)
2+	-1.002**	(0.101)	-0.992**	(0.109)
No. older with current partner				
1	-0.328**	(0.093)	-0.340**	(0.093)
2+	-0.669**	(0.120)	-0.695**	(0.127)
Preschool child(ren) with previous partner	0.239	(0.393)	0.211	(0.400)
Older child(ren) with previous partner	0.062	(0.174)	0.048	(0.174)
Child(ren) with non co-resident partner	0.607**	(0.200)	0.582**	(0.204)
Characteristics of current/previous partnerships				
Age at start of partnership (ref=20-24)				
<20 years	0.790**	(0.085)	0.792**	(0.085)
25-29	-0.487**	(0.096)	-0.489**	(0.099)
30-34	-0.916**	(0.171)	-0.905**	(0.175)
35+	-0.878**	(0.349)	-0.869**	(0.352)
Current partnership duration[b]				
t	0.204**	(0.019)	0.191**	(0.019)
t^2	-0.010**	(0.001)	-0.009**	(0.001)
t^3	0.000**	(0.000)	0.000**	(0.000)
Previously married	-0.569**	(0.197)	-0.626**	(0.195)
Previously cohabited with . . . (ref=no one)				
Current partner	0.106	(0.082)	0.089	(0.085)
Previous partner(s)	0.853**	(0.258)	0.927**	(0.265)
Both current and previous partner(s)	0.349*	(0.152)	0.370*	(0.150)
Background characteristics				
Post-16 years of education[a] (ref=0)				
1	-0.034	(0.094)	-0.034	(0.095)
2	-0.280**	(0.107)	-0.286**	(0.109)
3-5	-0.351**	(0.125)	-0.369**	(0.125)
6+	-0.539**	(0.176)	-0.552**	(0.178)
Paternal social class (ref=skilled occupation)				
Professional, managerial or technical	-0.037	(0.098)	-0.029	(0.097)
Partly skilled or unskilled	-0.081	(0.080)	-0.075	(0.080)
No father figure	0.016	(0.195)	0.012	(0.197)
Unknown	0.107	(0.304)	0.129	(0.311)
Family disruption during childhood	0.146	(0.099)	0.156	(0.101)
Constant	-7.037	(0.161)	-7.021	(0.180)

[a]Time-varying covariate.
[b]Duration is measured in 6-month intervals.
Statistical significance: * 5%, ** 1%. A parameter differs significantly from zero at a given significance level if the interval estimate, calculated from the MCMC samples, does not include zero.

Table 7. Estimated coefficients (and standard errors) from models of separation from cohabitation

	Single process model		Multiprocess model	
Variables	**Coeff.**	**(SE)**	**Coeff.**	**(SE)**
Prior fertility outcomes[a] (ref=none)				
Currently pregnant	-0.639**	(0.150)	-0.701**	(0.156)
No. preschool with current partner				
1	-0.236*	(0.120)	-0.290*	(0.120)
2+	-0.753**	(0.261)	-0.877**	(0.270)
No. older with current partner				
1	-0.032	(0.202)	-0.058	(0.208)
2+	0.239	(0.333)	0.136	(0.341)
Preschool child(ren) with previous partner	-0.330	(0.218)	-0.335	(0.224)
Older child(ren) with previous partner	-0.012	(0.128)	-0.022	(0.130)
Child(ren) with non co-resident partner	-0.019	(0.191)	-0.018	(0.194)
Characteristics of current/previous partnerships				
Age at start of partnership (ref=20-24)				
<20 years	0.161	(0.138)	0.157	(0.135)
25-29	-0.435**	(0.104)	-0.438**	(0.105)
30-34	-0.423**	(0.126)	-0.432**	(0.127)
35+	-0.714**	(0.179)	-0.725**	(0.190)
Current partnership duration[b]				
t	0.176**	(0.031)	0.198**	(0.030)
t^2	-0.011**	(0.002)	-0.013**	(0.002)
t^3	0.000**	(0.000)	0.000**	(0.000)
Previously married	-0.395**	(0.139)	-0.396**	(0.135)
Previously cohabited	-0.052	(0.103)	-0.054	(0.106)
Background characteristics				
Post-16 years of education [a] (ref=0)				
1	0.178	(0.126)	0.169	(0.129)
2	0.087	(0.134)	0.071	(0.140)
3-5	0.103	(0.130)	0.071	(0.135)
6+	0.127	(0.154)	0.096	(0.159)
Paternal social class (ref=skilled occupation)				
Professional, managerial or technical	0.230*	(0.112)	0.233*	(0.112)
Partly skilled or unskilled	-0.170	(0.112)	-0.174	(0.114)
No father figure	0.173	(0.224)	0.187	(0.223)
Unknown	0.453	(0.281)	0.457	(0.287)
Family disruption during childhood	0.256*	(0.119)	0.258*	(0.118)
Constant	-5.441	(0.138)	-5.475	(0.136)

[a]Time-varying covariate.
[b]Duration is measured in 6-month intervals.
Statistical significance: * 5%, ** 1%. A parameter differs significantly from zero at a given significance level if the interval estimate, calculated from the MCMC samples, does not include zero.

Table 8. Estimated coefficients (and standard errors) from models of the transition from cohabitation to marriage

	Single process model		Multiprocess model	
Variables	**Coeff.**	**(SE)**	**Coeff.**	**(SE)**
Prior fertility outcomes[a] (ref=none)				
Currently pregnant	0.694**	(0.062)	0.646**	(0.067)
No. preschool with current partner				
1	-0.194*	(0.080)	-0.220**	(0.082)
2+	-0.076	(0.156)	-0.136	(0.162)
No. older with current partner				
1	-0.352*	(0.183)	-0.372*	(0.187)
2+	-0.276	(0.360)	-0.350	(0.371)
Preschool child(ren) with previous partner	-0.065	(0.120)	-0.068	(0.122)
Older child(ren) with previous partner	-0.028	(0.083)	-0.032	(0.084)
Child(ren) with non co-resident partner	-0.410**	(0.138)	-0.422**	(0.139)
Characteristics of current/previous partnerships				
Age at start of partnership (ref=20-24)				
<20 years	-0.060	(0.090)	-0.057	(0.091)
25-29	-0.015	(0.065)	-0.009	(0.065)
30-34	-0.187*	(0.081)	-0.177*	(0.084)
35+	-0.190	(0.117)	-0.182	(0.119)
Current partnership duration[b]				
t	0.083**	(0.020)	0.097**	(0.022)
t^2	-0.010**	(0.002)	-0.011**	(0.002)
t^3	0.000**	(0.000)	0.000**	(0.000)
Previously married	-0.411**	(0.090)	-0.471**	(0.105)
Previously cohabited	-0.173*	(0.069)	-0.173**	(0.067)
Background characteristics				
Post-16 years of education [a] (ref=0)				
1	0.027	(0.079)	0.024	(0.081)
2	0.155	(0.084)	0.147	(0.085)
3-5	-0.004	(0.084)	-0.020	(0.087)
6+	-0.067	(0.101)	-0.082	(0.105)
Paternal social class (ref=skilled occupation)				
Professional, managerial or technical	-0.050	(0.072)	-0.056	(0.073)
Partly skilled or unskilled	-0.113	(0.067)	-0.118	(0.069)
No father figure	-0.080	(0.147)	-0.079	(0.151)
Unknown	-0.440**	(0.228)	-0.448*	(0.233)
Family disruption during childhood	0.044	(0.078)	0.051	(0.078)
Constant	-3.897	(0.075)	-3.920	(0.079)

[a]Time-varying covariate.
[b]Duration is measured in 6-month intervals.
Statistical significance: * 5%, ** 1%. A parameter differs significantly from zero at a given significance level if the interval estimate, calculated from the MCMC samples, does not include zero.

Table 9. Estimated coefficients (and standard errors) from the multiprocess model for the hazard of a conception within a partnership

	Within cohabitation		Within marriage	
Variables	**Coeff.**	**(SE)**	**Coeff.**	**(SE)**
Prior fertility outcomes [a] (ref=none)				
No. preschool with current partner				
1	0.344**	(0.086)	0.241**	(0.026)
2+	-0.221	(0.198)	-0.710**	(0.044)
No. older with current partner				
1	-0.059	(0.182)	-1.070**	(0.050)
2+	-0.604	(0.361)	-1.428**	(0.073)
Preschool child(ren) with previous partner	0.120	(0.135)	0.381*	(0.152)
Older child(ren) with previous partner	-0.132	(0.103)	-0.671**	(0.088)
Child(ren) with non co-resident partner	0.204	(0.144)	-0.320**	(0.104)
Characteristics of current/previous partnerships				
Age at start of partnership (ref=20-24)				
<20 years	0.336**	(0.108)	0.129**	(0.034)
25-29	-0.191*	(0.089)	-0.055	(0.032)
30-34	-0.420**	(0.108)	-0.251**	(0.060)
35+	-0.687**	(0.163)	-0.203	(0.138)
Current partnership duration [b]				
t	-0.049**	(0.015)	0.041	(0.006)
t^2	0.000	(0.000)	-0.003	(0.000)
Previously married	-0.004	(0.101)	-0.106	(0.064)
Previously cohabited	0.165*	(0.084)	–	–
Previously cohabited with . . . (ref=no one)				
Current partner	–	–	0.035	(0.029)
Previous partner(s)	–	–	-0.040	(0.117)
Both current and previous partner(s)	–	–	0.039	(0.059)
Background characteristics				
Post-16 years of education [a] (ref=0)				
1	-0.403**	(0.115)	0.002	(0.035)
2	-0.588**	(0.135)	-0.034	(0.037)
3-5	-0.598**	(0.134)	0.068	(0.040)
6+	-0.361*	(0.153)	0.106	(0.053)
Paternal social class (ref=skilled occupation)				
Professional, managerial or technical	-0.215*	(0.111)	-0.008	(0.034)
Partly skilled or unskilled	-0.016	(0.084)	-0.006	(0.030)
No father figure	0.068	(0.173)	0.074	(0.075)
Unknown	-0.072	(0.265)	0.126	(0.106)
Family disruption during childhood	0.172	(0.096)	-0.034	(0.040)
Constant	-4.440	(0.094)	-4.177	(0.036)

[a]Time-varying covariate.
[b]Duration is measured in 6-month intervals.
Statistical significance: * 5%, ** 1%. A parameter differs significantly from zero at a given significance level if the interval estimate, calculated from the MCMC samples, does not include zero.