The Formation and Outcomes of Cohabiting and Marital Partnerships in Early Adulthood: The Role of Previous Partnership Experience

Fiona Steele†

Centre for Multilevel Modelling

Graduate School of Education

University of Bristol

Constantinos Kallis and Heather Joshi

Centre for Longitudinal Studies

Institute of Education, University of London

Summary: We estimate a joint model of the formation and dissolution of cohabiting and marital unions among British women born in 1970. The focus of the analysis is the effect of previous cohabitation and marriage on subsequent partnership transitions. We use a multilevel simultaneous equations event history model to allow for residual correlation between the hazards of moving from an unpartnered state into cohabitation or marriage, converting a cohabiting union into marriage, and dissolution of either form of union. A simultaneous modelling approach allows for the joint determination of these transitions which may otherwise bias estimates of the effects of previous partnership outcomes on later transitions.

Keywords: simultaneous equation model, repeated events, multiple states, partnership formation, partnership dissolution

† *Address for correspondence*: Fiona Steele, Centre for Multilevel Modelling, Graduate School of Education, University of Bristol, 35 Berkeley Square, Bristol BS8 1JA, Email (until 1/9/05): F.Steele@ioe.ac.uk

1. Introduction

Most studies of the formation and outcomes of coresidential partnerships in Britain consider only the first union, typically investigating the timing of entry into marriage and marital breakdown. The focus on the first marriage is consistent with the traditional route to family formation whereby an individual left the parental home to marry before having children. There is now greater diversity among young adults in their partnership trajectories and the sequencing of union formation and childbearing (Berrington 2003). A notable change is the sharp increase in premarital cohabitation (Ermisch 2005; Murphy 2000) which has led researchers to broaden their definition of a partnership to include non-marital coresidential unions, and to study the outcomes of the first cohabiting partnership (e.g. Berrington 2001). Another important shift in partnership behaviour is an increase in the dissolution rate, leading to multiple partnerships per individual (Ferri and Smith 2003, Murphy and Wang 1999). Nevertheless, studies that consider the formation and dissolution of second and higher order partnerships are rare.

In this paper, we study repartnering behaviour, and in particular how it is shaped by past partnership events. Specifically, we examine the relationship between previous cohabitation or marriage and the timing of the formation and dissolution of subsequent partnerships. We ask whether couples who lived with each other or a previous partner before marriage are at increased risk of marital dissolution. There is some evidence for first marriages in Britain between 1970 and 1989 that those which were preceded by premarital cohabitation were more likely to dissolve (Haskey 1999, using data from the 1989 General Household Survey). There was, however, no control for other characteristics observed or otherwise. In a study of first marriages among the 1958 birth cohort which adjusts for the effects of a range of background characteristics, Berrington and Diamond (1999) found that previous cohabitation, with either the marital or a previous partner, is associated with a higher risk of dissolution. Haskey (1992) also cites evidence from the USA, Canada and Sweden that show a positive association between cohabitation and the breakdown of a subsequent marriage. Following the break-up of a marriage, most divorcees (especially women) cohabit with their next partner (Haskey 1999). We explore whether the experience of divorce acts as a deterrent to converting these cohabitations into another marriage. Finally, we consider the relationship between previous marriage or cohabitation and subsequent partnership dissolution. Are later partnerships at a higher risk of dissolution because the possibility of break-up has already been experienced, or does the previous experience, on the contrary, discourage repeating it? Previous research on divorce in Britain tends to support the former hypothesis. Haskey (1983, 1996) has found that marriages in which a partner is remarrying after divorce carry an above-average risk of subsequent divorce.

A major difficulty in determining the effects of previous events on the timing of later events in the same or a related process is the potential for selection bias. There may be unobserved time-invariant characteristics that affect the probability of event occurrence across the observation period, leading to correlation in the durations between successive events in the same process. For example, the positive association observed between previous marital breakdown and the dissolution rate of remarriages may be due to differences between individuals who marry only once and those who have divorced; the latter group may contain individuals whose unobserved characteristics put them at high risk of separation from any partner. Another form of selection arises when the unobserved risk factors for events in one process are correlated with those for events in a different but related process. A study in the US found that this type of selection explained an apparent positive effect of premarital cohabitation on the risk of marital breakdown (Lillard et al. 1995). Women who chose to cohabit with their partner before marriage differed from those who married directly in unobserved ways which were related to their risk of marital dissolution. In this paper, we adjust for such selection effects by jointly modelling repeated entries into and exits from partnerships.

Using data from the 1970 British cohort, we study the formation and outcomes of all partnerships formed by women before age 30. In these years of their early adulthood nearly all women formed coresidential partnerships, but these differed from previous generations in having entered partnership a little later and for a higher proportion of their first (and current) partnerships to be cohabitations rather than marriage (Ferri and Smith 2003). These partnerships were also being formed in an era when divorce was becoming a normal occurrence which may have encouraged a more relaxed attitude to the importance of entering formal marriage, while parallel developments in education and employment meant that these young women need be less dependent on finding a husband than their predecessors had been. We focus on women for ease of comparison with our previous research on partnership transitions among the 1958 and 1970 birth cohorts (Steele et al. 2005a, 2005b), and because men's later age at first partnership results in shorter exposure to multiple partnerships. Our study builds upon previous research in several ways. First, while other studies that have jointly modelled partnership formation and dissolution (Aasve et al. 2004; Goldstein et al. 2004) grouped together marriage and cohabitation, we treat them as separate partnership states. Second, we extend the work of Steele et al. (2005a) who jointly modelled transitions *out* of marriage and cohabitation to incorporate transitions *into* both forms of partnership. Third, we extend the simultaneous equations model of Lillard et al. (1995) to examine the impact on the risk of marital dissolution of premarital cohabitation not only with the current partner but with any previous partner.

2. Assessing the Effect of Previous Partnership Events on Subsequent Transitions: The Need for a Joint Modelling Approach

In an analysis of multiple partnership transitions, and in particular estimation of the effects of previous partnership outcomes on subsequent transitions, it is important to consider the potential for observed and unobserved heterogeneity and the endogeneity of past partnership outcomes with respect to later behaviour. One covariate which may or may not be observed which links cohabitation and marital break-up is religious belief and whether the marriage was a religious ceremony (Murphy 1983). Other relevant personality traits are inherently less observable. To illustrate the implications of unobserved heterogeneity, suppose we are interested in the effect of divorce on the stability of a later marriage. There are likely to be unobserved individual characteristics, constant over time, that influence an individual's dissolution risk in any marriage they form. The presence of unobserved time-invariant risk factors leads to unobserved heterogeneity between individuals. If such heterogeneity is not taken into account, the effect of previous marriage on subsequent marital dissolution cannot be interpreted in a causal way; rather the dissolution risk among the previously married is likely to be inflated by the presence of individuals whose unobserved attributes place them at increased risk, leading to an overstated positive effect of previous marriage. Similarly, failure to account for unobserved heterogeneity is likely to lead to a biased estimate of the effect of previous cohabitation on the dissolution risk of later cohabiting partnerships. Aasve et al. (2004) and Lillard et al. (1995) highlighted the need to control for unobserved heterogeneity when considering repeated events. Using data from the British Household Panel Study (BHPS), Aasve et al. found that an apparently strong increase in the risk of separation with the number of previous partners vanished after adjusting for the effects of unobserved individual-specific factors. In the US, Lillard et al. also found that the increased dissolution rate observed among previously married women disappeared after accounting for unobserved heterogeneity.

The above examples concern the effect of a previous event on the probability that an event of the same type occurs again. More generally, we may be interested in the effect of any previous partnership event (e.g. the breakdown of a marriage or cohabitation, and premarital cohabitation) on later transitions, which include both the dissolution and formation of marital and cohabiting partnerships. To examine the effect of the break-up of a marriage (cohabitation) on the stability of a later cohabitation (marriage), it is important to allow for the possibility that individuals with a high risk of separating from a cohabiting partner (according to unobserved characteristics) might also have a high risk of marital dissolution. This may be achieved by jointly modelling the outcomes of marriage and cohabitation and allowing for residual correlation between the transitions. Using this approach, Steele et al. (2005a) found evidence of a positive correlation between the unmeasured woman-specific influences on the risk of separation from marriage and cohabitation.

Investigation of the impact of prior cohabitation on the risk that a subsequent marriage is dissolved raises another methodological issue. There has been much debate about the effect of premarital cohabitation on the risk of marital breakdown. An increased dissolution rate is often observed among couples who lived together before marriage, but many commentators argue that this is due to selection rather than a causal effect. Lillard et al. (1995) tested the selection hypothesis using a simultaneous equations model in which the hazard of marital dissolution was modelled jointly with the decision to cohabit before marriage; they found that the apparent increased risk of separation among the cohabitors was due to self-selection of women with a high risk of dissolution into cohabitation. This selection effect was not captured by covariates and therefore led to a positive correlation between the woman-specific unobservables, constant across marriages for the same woman, that affect the hazard of dissolution and the propensity to cohabit. Put another way, the decision to cohabit before marriage is endogenous to marital dissolution. No British study of the link premarital cohabitation and marital dissolution has allowed for selection on unobservables. Of course cohabitation is not always a precursor to marriage; cohabitation may also take the form of a short-term convenient alternative to a non-coresident sexual relationship, or a longer-term alternative to marriage (Murphy 2000). The association between *any* previous cohabitation experience and marital dissolution may be subject to the same form of selection reported by Lillard et al., i.e. that individuals with an above-average propensity to cohabit also have an above-average propensity to separate should they marry.

As discussed above, failure to account for correlation between different types of partnership transition may lead to incorrect inferences. There are at least two additional benefits of adopting a joint modelling approach in the analysis of partnership histories. First, estimates of the residual correlations across transitions can provide useful insights into partnership patterns. Goldstein et al. (2004), in an analysis of unions formed by male members of the 1958 birth cohort between the ages of 16 and 33, found a weak negative residual correlation between the hazards of entry into and exit from partnerships. A negative correlation may be due to the presence of individuals with a strong latent desire to be in a partnership, who therefore have a tendency towards long partnerships and short unpartnered episodes. Aasave et al. (2004), however, using data on multiple cohorts from the BHPS, find that the correlation between the hazards of partnership formation and dissolution is positive for both men and women, although not statistically significant from zero for men. A positive correlation suggests the presence of women who make quick transitions into and out of partnerships. A second advantage of estimating one joint model for all types of partnership transition rather than independent models is that tests of the equivalence of covariate effects across transitions may be carried out. For example, one may be interested in comparing the effects of previous partnership experience on the dissolution risks of marital and cohabiting partnerships.

3. Methods

3.1 Definitions of state transitions and episodes

Over the course of their event history, an individual will usually move between different ‘states’ over time. Here we distinguish between three partnership states – single, cohabitation and marriage. In theory, there are a total of eight possible transitions between these states. From the single state, an individual may move into marriage or cohabitation. For a cohabiting couple there are potentially four types of transition: they may marry each other or they may separate, after which a cohabitor may become single or move immediately into cohabitation or marriage with a new partner. (Strictly the start of a new cohabitation is a transition within the same state, but it is natural to treat a change in partner as the start of a new episode.) Similarly, following a marital separation, an individual may move into either the single or cohabitation state. In practice, however, and in common with previous research, we do not take account of the partnership state occupied immediately after dissolution in our definition of transitions; dissolution is our event of interest. We therefore consider a total of five transitions, and distinguish between partnership formation (transitions out of the single state) and partnership outcomes (dissolution of a marriage or cohabitation, and the conversion of cohabitation into marriage). Whenever one of these five types of event occurs, a new partnership episode begins.

3.2 Discrete-time data structure

We analyse the formation and outcomes of cohabiting and marital unions using a multilevel discrete-time modelling approach. For each episode i in the partnership history of woman j, we observe the current state s_{ij} (coded 1 if single, 2 if cohabiting and 3 if married), the episode duration t_{ij}, and an event indicator δ_{ij}, the coding of which depends on the current state as follows:

$$\delta_{ij} = \begin{cases} 0 & \text{no event occurs, i.e. censored} \\ 1 & \text{partnership dissolution } (s_{ij} = 2,3) \textit{ or } \text{entry into cohabitation } (s_{ij} = 1) \\ 2 & \text{entry into marriage } (s_{ij} = 1,2) \end{cases}$$

From the observed data $(s_{ij}, t_{ij}, \delta_{ij})$ we create, for each time interval t, a binary response $y_{ij}(t)$ for $s_{ij} = 3$, and a bivariate binary response $(y_{ij}^{(1)}, y_{ij}^{(2)})$ for $s_{ij} = 1,2$. For $t < t_{ij}$, all binary responses are coded 0. For $t = t_{ij}$, the binary responses for each state are coded as follows.

If $s_{ij} = 1,2$, $y_{ij}^{(r)} = \begin{cases} 1 & \delta_{ij} = r \\ 0 & \delta_{ij} \neq r \end{cases}$ $\quad(r = 1,2)$;

if $s_{ij} = 3$, $y_{ij} = \delta_{ij}$.

The coding of the binary indicators $y_{ij}^{(r)}$ for partnership formation and the outcomes of cohabitation follows the commonly applied latent survival time approach to competing risks (Box-Steffensmeier and Jones 2004). This approach assumes that for single and cohabitation episodes, there are two latent durations $t_{ij}^{(1)*}$ and $t_{ij}^{(2)*}$, the durations to events of type 1 and 2 respectively. We observe $t_{ij} = \min(t_{ij}^{(1)*}, t_{ij}^{(2)*})$. When an event of one type occurs, an individual is removed from the risk of experiencing the other type of event. For example, if a cohabiting woman breaks up with her partner $(r=1)$ at t_{ij}, we fully observe $t_{ij}^{(1)*}$ while $t_{ij}^{(2)*}$ is censored at t_{ij}; we assume that she would have continued to be at 'risk' of marrying that partner had they not separated.

In any discrete-time analysis, the length of the time intervals t must be specified. In the present case, it is possible to use monthly intervals because the start and end dates of cohabitations and marriages were recorded to the nearest month. We decided to use broader intervals, however, due to the length of the observation period, together with the need to have two records for each time interval for unpartnered and cohabitation episodes. After exploratory analysis in which the impact of grouping on parameter estimates was investigated, six-month intervals were chosen. Thus we assume that partnership transition hazards and the values of time-varying covariates remain constant within six month intervals. Under this assumption, grouping will not lead to loss of information provided that intervals are weighted by exposure time. For each six-month interval, a weight is defined as the number of months during the interval that a woman was exposed to the risk of making a transition.

3.3 Model for partnership transitions

Partnership formation model

The two possible transitions from the single state, to cohabitation $(r=1)$ or marriage $(r=2)$, are treated as competing risks as described above. The hazard that woman j enters a partnership of type r in time interval t of episode i is given by $h_{ij}^{S(r)}(t) = \Pr(y_{ij}^{(r)} = 1 \mid y_{ij}^{(r)} = 0, y_{ij}^{(r')} = 0,\ s_{ij} = 1)$. We fit a two-level random effects logit model to the bivariate responses $(y_{ij}^{(1)}, y_{ij}^{(2)})$:

$$\text{logit}[h_{ij}^{S(r)}(t)] = \alpha^{S(r)} D_{ij}^{S}(t) + \beta^{S(r)} X_{ij}^{S}(t) + u_{j}^{S(r)}, \qquad r = 1, 2 \tag{1}$$

where $\alpha^{S(r)} D_{ij}^{S}(t)$ is a function of the cumulative duration of the current single episode by time t, the baseline logit-hazard; $X_{ij}^{S}(t)$ is a vector of explanatory variables, which may be defined at the time interval, episode or woman level, with coefficient vector $\beta^{S(r)}$, and $u_{j}^{S(r)}$ is a woman-level random effect. The inclusion of individual-specific random effects allows for correlation between the duration of episodes from the same woman. Event times will be correlated if there are unobserved characteristics, fixed over time, which affect a woman's hazard of a transition in all episodes. For example, some women may form coresidential partnerships quicker than others because of unmeasured personality traits. The random effects $u_{j}^{S(r)}$ represent unobserved factors which affect a woman's probability of forming a cohabiting or marital union during any unpartnered episode. We assume that the random effects $\mathbf{u}_j^S = [u_j^{S(1)} \quad u_j^{S(2)}]^T$ follow a bivariate normal distribution with zero mean and variance $\mathbf{\Omega}^S$.

Partnership outcomes model

The model for the outcomes of cohabiting and marital unions consists of three equations, two for cohabitation and one for marriage.

Cohabitation

An episode of cohabitation may end in separation $(r = 1)$ or marriage to the same partner $(r = 2)$. The hazard that woman j ends an episode of cohabitation i in time interval t for reason r is given by $h_{ij}^{C(r)}(t) = \Pr(y_{ij}^{(r)} = 1 \mid y_{ij}^{(r)} = 0, y_{ij}^{(r')} = 0, s_{ij} = 2)$. The model takes the same form as the partnership formation model (1), i.e.

$$\text{logit}[h_{ij}^{C(r)}(t)] = \alpha^{C(r)} D_{ij}^C(t) + \beta^{C(r)} X_{ij}^C(t) + u_j^{C(r)}, \qquad r = 1, 2 \tag{2}$$

where the random effects $\mathbf{u}_j^C = [u_j^{C(1)} \quad u_j^{C(2)}]^T$ follow a bivariate normal distribution with zero mean and variance $\mathbf{\Omega}^C$.

Marriage

The hazard that woman j separates from a marital partner during time interval t of episode i is $h_{ij}^M(t) = \Pr(y_{ij} = 1 \mid y_{ij} = 0,\ s_{ij} = 3)$. The model for marital separation is a two-level logit model of the form:

$$\text{logit}[h_{ij}^M(t)] = \alpha^M D_{ij}^M(t) + \beta^M X_{ij}^M(t) + u_j^M \tag{3}$$

where $u_j^M \sim N(0, \Omega^M)$.

3.4 Estimation

We consider two specifications of the model defined by (1), (2) and (3). In the first, referred to as the single process model, the random effects $(\mathbf{u}_j^S, \mathbf{u}_j^C, u_j^M)$ are assumed to be mutually uncorrelated.

Note that we still allow for pairwise correlation between $u_j^{S(1)}$ and $u_j^{S(2)}$, and between $u_j^{C(1)}$ and $u_j^{C(2)}$. In the second specification, referred to as the multiprocess or joint model, we allow for non-zero correlation between any pair of the five random effects, i.e. correlation across the processes of partnership formation, partnership dissolution and the conversion of cohabitation into marriage.

The simplest way to fit the single process model is to estimate the three components of the model implied by (1), (2) and (3) independently. The two equations within each of (1) and (2) must be estimated simultaneously because of the pairwise random effect correlations across equations. The multiprocess model is fitted by simultaneous estimation of all five equations. Here, we outline estimation of the more general multiprocess model; an alternative way of fitting the single process model is to constrain the cross-process random effect correlations in the multiprocess model to equal zero. In most cases, the first step will be to convert an episode-based file to the discrete-time format described above. Next, dummy variables for each of the three states are created. In addition, for episodes in the single and cohabitation states where there are two possible end events, the responses $y_{ij}^{(1)}(t)$ and $y_{ij}^{(2)}(t)$ are stacked into a single response and a response indicator defined for each. This leads to five dummy variables, one for marriage and two each for single and cohabitation. Further details, with an example, are given in Steele et al. (2004). Equations (1), (2) and (3) may then be expressed in the form of a single model equation, where the explanatory variables are the state and event type dummies and their interactions with duration and covariates. The random effects are fitted by allowing the coefficient of each dummy variable to vary randomly across women. The model may be estimated using existing algorithms for two-level logit models. In this paper, all estimation was carried out using Markov Chain Monte Carlo (MCMC) methods in *MLwiN* v2.0 (Rasbash et al. 2004).

4. Data

4.1 Partnership histories and details of the analysis sample

We analyse data from the 1970 British Cohort Study, a prospective longitudinal study of all those living in Great Britain who were born in a single week of 1970 (Bynner et al. 1997). Since birth, there have been five attempts to contact cohort members at ages 5, 10, 16, 26 and 30. Data were collected from parents, cohort members and supplementary sources on respondents' physical, educational and social development over the life course. At age 30, respondents were asked to recall the start and end dates of all coresidential relationships which lasted for at least a month. Such relationships are referred to as partnerships. Information on both marriage and (unmarried) cohabitation was collected and, where marriage was preceded by cohabitation with the same partner, the dates of the start of cohabitation and marriage were both recorded. Periods between partnerships are classified as single (unpartnered) episodes in the analysis.

As described in Section 3.2, the dependent variable is a partnership transition indicator, or pair of indicators in the case of partnership formation and the outcomes of cohabitation. The indicator(s) are defined for each six-month interval between the ages of 16 and 30. Episodes that are in progress at the time of interview are treated as right-censored, as are partnership episodes that end due to the death of the partner. We examine the partnership transitions of female respondents. Partnership histories were collected from 5790 women. Of these, 295 were excluded from the analysis because of missing or inconsistent dates in their partnership, fertility or employment histories, resulting in a final sample of 5495. These women contribute 15032 partnership episodes, of which 48.0% are single episodes, 32.2% are cohabitations and 19.8% are marriages. The discrete-time analysis file, using six-monthly time intervals, contains 274472 records.

4.2 Choice and definition of explanatory variables

The focus of our study is the effect of previous partnership experience on the formation and dissolution of subsequent partnerships. Thus, the explanatory variables of major interest are indicators of previous marriage and cohabitation. In addition, we consider a set of background characteristics as control variables, the choice of which is guided by previous research on the determinants of partnership transitions. The variables we consider may be classified as follows: current (time-varying) duration unpartnered or in a partnership, age, current fertility status, current educational enrolment and attainment, and family background. Table 1 shows the coding and percentage distributions of all explanatory variables for single, cohabitation and marriage episodes.

By jointly modelling the formation and outcomes of partnerships we allow for the potential endogeneity of previous partnership events with respect to subsequent transitions. We treat all other variables as exogenous. This assumption may be questionable for outcomes of processes which are contemporaneous to partnership transitions, i.e. the time-varying indicators of fertility and educational status. While it is possible to extend our model to allow for the joint determination of partnership, childbearing and educational decisions, we do not pursue this approach here. The model defined by eq. (1), (2) and (3) already consists of five equations and the addition of further equations for childbearing and education would greatly increase model complexity and therefore estimation times. Although Aasve et al. (2004) simultaneously model partnership, fertility and employment histories, they simplified the partnership process by grouping together marriage and cohabitation. In previous work based on the 1958 and 1970 cohorts (Steele et al. 2005a, 2005b), we demonstrate that allowing for the endogeneity of fertility actually has little effect on estimates of the relationship between partnership outcomes and fertility status. Nevertheless, the estimated coefficients of the fertility and education variables presented in this paper should be interpreted as associations rather than causal effects.

Partnership duration

The durations of partnership and single episodes are derived directly from the partnership histories, as are the indicators of previous partnership experience. For marriage episodes, partnership duration includes any period of premarital cohabitation.

In a discrete-time event history model, the analyst must specify the form of the baseline hazard function by including as explanatory variables some function of the current duration in a given state. The functional form of the baseline hazard was chosen after inspection of hazard plots for each of the five transitions. For each type of transition, the hazard was found to have an inverse U-shaped relationship with duration which was well approximated by a quadratic function. Therefore duration and duration-squared terms were included as explanatory variables in the model. In the case of partnership formation, we fitted distinct baseline hazards for formation of the first and subsequent partnerships by including interactions between a dummy for partnership order and the time-varying duration variables.

Age

Age is treated as a time-varying covariate in the model for partnership formation. In the case of the first partnership, the age at time t is equal to the current duration single because, for all women, we model partnership transitions from an origin of age 16. Therefore age effects are estimated only for entry into the second and higher order partnerships. In common with most previous studies of the transition from cohabitation to marriage and partnership dissolution, we consider the effect of the age at the start of the partnership. Both the time-varying and partnership-level age variables are treated as categorical to allow for nonlinear age effects, with five categories each of three years width.

Current fertility status

Fertility status is denoted by a set of time-varying variables which indicate current pregnancy status and the presence of children. A unique feature of the pregnancy histories collected in the cohort studies is that respondents were asked to identify the other parent of each of their children, if not the partner at interview, with reference to the partners named in the partnership histories. Thus, for each time interval t, we are able to distinguish between children with the current partner, those fathered by previous cohabiting or marital partners, and children from non-coresident relationships. We count only children who are living with the partner at time t, using information on the date at which a child left home. Finally, we consider the current age of each child, classified as preschool (less than five years old) or older.

Educational enrolment and attainment

Education has been identified as an important determinant of a range of outcomes throughout the lifecourse, including partnership formation and dissolution. In this study, we consider time-varying indicators of enrolment in full-time education and the number of years of post-16 education. The indicator of educational enrolment is derived from employment histories which include full-time education as an employment state. The attainment measure is based on the cumulative number of months of post-16 education. Where employment histories were incomplete, values for the educational indicators were imputed on the basis of responses to a question on the highest qualification achieved; in such cases, we had to assume that a woman was continuously in education until her final qualification was obtained.

Family background

Although indicators of educational attainment and enrolment are treated as time-varying covariates, and therefore reflect contemporaneous life course characteristics, their values at time t are largely determined by family and school characteristics. As such, our educational measures are to some extent proxies for family background. More direct indicators of family background include parental social class, the experience of parental divorce during childhood, and region of residence.

Our measure of paternal social class is based on the occupation of the father figure at the time of the respondent's birth, which we classify as I or II (professional, managerial or technical occupations), III (manual or non-manual skilled), and IV or V (partly skilled or unskilled). Women who had no father figure are placed in a separate category. The indicator of family disruption before age 16 includes the experience of parental divorce or any other living arrangement where the father or mother figure was not one of the natural parents. Region of residence at birth was initially represented by dummies for the 12 standard regions, but grouped into five categories following previous research and our own exploratory analysis.

Missing data

Of those successfully interviewed at age 30, there is missing covariate information, either because they were not present at all earlier sweeps or because the data collected were subject to item nonresponse. There is therefore missing data for some of the explanatory variables included in our models, namely, the family background variables. Where possible, missing values were imputed using information collected at earlier or later ages. For example, if father's social class at the cohort member's birth was unavailable, we used information collected at a later stage of childhood if it was reasonable to assume that the child's father figure at that age was the same person as the father figure at birth. Even after imputation, however, there remained some missing data. Rather than drop cohort members with incomplete data from the analysis, we created an extra 'missing' category for those variables affected by nonresponse. This approach may introduce bias to estimates of their effects on partnership transitions (Greenland and Finkle 1995). However, this bias will be small if background characteristics are weakly correlated with the outcomes under study which, as we shall see, is the case here.

5. Results

5.1 Partnership trajectories and the timing of transitions

Figure 1 shows the partnership trajectories made by women in the analysis sample before age 30. Thirteen percent of women had not yet formed a partnership. Most (70%) women had one partner and 81% of those were still with that partner at the time of interview (percentages calculated from Figure 1). The remaining 17% of women had more than one partner. The trajectories reveal the high incidence of cohabitation among this cohort. Seventy-two percent of all women cohabited with their first partner, which led to marriage for half of them. Of those who married their first partner without living together first, 17% subsequently separated and an overwhelming majority of those who repartnered (97%) cohabited with their second partner.

We now turn to the timing of partnership formation, the conversion of a cohabiting union to marriage, and partnership dissolution. Table 2 shows the quartiles of the distribution of the number of years to each type of transition, calculated from life tables. In the case of partnership formation and the outcomes of cohabitation, where there are competing risks, these distributions are based on associated single decrement life tables. For partnership formation, we consider separately the timing of the first and subsequent cohabitations. The duration to the first cohabitation is calculated from the sixteenth birthday for all women. The duration to the first partnership, usually cohabitation (see Figure 1), is likely to be longer than the duration of single episodes following partnership dissolution for two reasons. First, most women delay partnering until their 20s, leading to a long median duration of the first single period and, second, the duration to first cohabitation includes those who have never partnered. In the analysis of the timing of direct marriage, i.e. without premarital cohabitation, we do not distinguish the first marriage and remarriages. Because very few women remarry without living with their partner first (see Figure 1), the distribution of the duration to marriage is based predominantly on first marriages.

Starting with partnership formation, we see that women enter cohabitation much quicker than marriage, and that the formation of second and subsequent cohabitations is particularly fast. We note, however, that direct marriage is rare for this cohort and women who choose to marry without premarital cohabitation are likely to be a select group. Women who do not live with their partner before marriage may have a negative perception of premarital cohabitation. This group of women might find it more difficult to find a partner who shares their (negative) attitude towards cohabitation and this may increase the length of time they spend unpartnered. Another possible explanation for the long duration to direct marriage is that women who prioritise career over family might delay partnership formation altogether. By the time they decide to form a coresidential relationship, they may find marriage preferable to cohabitation. This is one reason for including time-varying age in the models for partnership formation.

Turning to the outcomes of partnerships, we can see that cohabitations tend to be of much shorter duration than marriages, suggesting that marriage is still viewed by most couples as a longer term commitment than cohabitation. However, there is evidence that young women now cohabit for longer with a partner before marriage. In this cohort the median duration at which cohabitation is converted to marriage is just over four years, compared to 2.7 years for the 1958 cohort (Steele et al. 2005a).

5.2 Joint modelling of partnership formation and outcomes

As described in Section 3.4, we consider two model specifications in our analysis of partnership formation and partnership outcomes. In the single process model, the random effects associated with transitions from the three states $(\mathbf{u}_j^S, \mathbf{u}_j^C, u_j^M)$ are assumed to be uncorrelated across states, although the random effects for different exits from the same state may be correlated. In the multiprocess model, pairwise correlations are estimated between each of the five random effects.

Residual correlations across transitions

The first step of the analysis was to assess whether the multiprocess model was a better fit to the data than the simpler single process model. Table 3 shows the random effects covariance matrix from the multiprocess model. Non-zero correlation(s) across the three partnership states, i.e. pairwise correlation between any element of $(\mathbf{u}_j^S, \mathbf{u}_j^C, u_j^M)$, would suggest that there is residual cross-state correlation, in which case the multiprocess model is the preferred specification.

From an examination of the 95% interval estimates for the covariance terms, we find that two cross-state correlations are significantly different from zero at the 5% level. The significant correlations are between the random effect for marital separation (u_j^M) and the random effects for entry into cohabitation ($u_j^{S(1)}$) and entry into marriage via cohabitation ($u_j^{C(2)}$). Both correlations are positive and moderately strong with point estimates of 0.487 and 0.457 respectively. Because the most common exit from the single state is into cohabitation, the first correlation may be interpreted as a positive residual association between the hazard of partnership formation and the hazard of marital separation. Thus women who cohabit quickly also tend to have shorter marriages. We note, however, that a short time to cohabitation is not associated with a higher dissolution risk for cohabitations; in fact there is some suggestion of a negative association between the hazards of forming and dissolving a cohabiting union (the 95% interval estimate for $Cov(u_j^{S(1)}, u_j^{C(1)})$ is -0.390 to 0.012, r = -0.322). The second significant correlation suggests that a quick transition from premarital cohabitation to marriage is associated with a higher risk of marital dissolution. Taken together, these results are consistent with those of Aasve et al. (2004) who found that fast partnership formation (cohabitation and marriage combined) is associated with shorter times to separation.

Steele et al. (2005a) allowed for residual correlation between the outcomes of partnerships formed by the 1958 cohort between ages 16 and 42. They found strong evidence of a positive correlation between the hazards of dissolution for marriage and cohabitation, suggesting that women with a high risk of marital separation also tend to have a high risk of separating from a cohabiting partner. Surprisingly we find no evidence of such a correlation in the 1970 cohort for partnerships formed before age 30 (the 95% interval estimate for $Cov(u_j^{C(1)}, u_j^{M})$ is -0.526 to 0.544, r = -0.072). Possible explanations for the difference in findings across the two cohorts are the shorter observation period for the earlier cohort, and later marriage for those born in 1970 which means that they are less likely to have married by age 30.

Effects of previous partnership experience on subsequent transitions

The main focus of our analysis is the effects of prior partnership experience on the formation and outcomes of later partnerships. The estimated coefficients for these and other explanatory variables are shown in Tables 4-6 for the single process and multiprocess models. As expected, the effects of the indicators of previous cohabitation and marriage are the most sensitive to model specification. We also note that the additional model complexity that results from allowing for cross-state residual correlation leads to loss of precision in the estimated coefficients for the partnership history variables.

Starting with partnership formation (Table 4) we see that previously partnered women are more likely to cohabit with a new partner, but are no more likely to marry (directly) than those embarking on their first partnership. Our conclusions are the same for each model, although the positive effect of previous partnership experience is slightly understated in the single process model. Turning to entry into marriage via premarital cohabitation (the last two columns of Table 5) we find that previous experience of marital dissolution is associated with decreased odds of marriage to a subsequent cohabiting partner. The negative effect of previous marriage is much stronger, and attains significance at the 5% level, in the multiprocess model. The change in the magnitude of this coefficient is due to the positive residual correlation between the hazard of marital separation and the hazard of marrying a cohabiting partner (Table 3). On average, women with a high hazard of marriage also have a high risk of marital dissolution, so that women with high odds of making the transition from cohabitation to marriage are disproportionately represented in the previously married category. After accounting for this form of selection in the multiprocess model, the marriage rate among the previously married is no longer artificially inflated, and a stronger negative effect emerges.

In their analysis of the outcomes of cohabitation among the 1958 cohort, for the same age range, Steele et al. (2005b) found that compared to women in their first partnership the marriage rate is lower not only among the previously married but also among those who have cohabited with a previous partner. However, their estimate of the effect of prior cohabitation should be interpreted with some caution as they did not model the decision to cohabit jointly with transitions from cohabitation.

Next we consider the effects of partnership history on the dissolution risk of later cohabitation (Table 5). Based on the single process model, we would conclude that there is some evidence of a negative effect of previous cohabitation, but not previous marriage, on the stability of later cohabiting partnerships. The effect is considerably weaker and no longer significant at the 5% level in the multiprocess model. Again, the explanation for the change in estimate lies in the cross-state residual correlations, and specifically the negative correlation between the durations to the formation and dissolution of cohabitations (Table 3). Women with a high hazard of entering cohabitation tend also to have a low hazard of separating from a cohabiting partner, leading to a selection of women with a low dissolution risk into cohabitation, thus inflating the negative effect of previous cohabitation. Steele et al. (2005b) also found no effect of previous cohabitation or marriage on the dissolution risk of cohabiting women born in 1958.

The most striking differences between the single and multiprocess models are in the estimated effects of previous marriage and cohabitation on the risk of marital separation (Table 6). In particular, the single process model shows a strong positive effect of previous cohabitation with both the current and an earlier partner. The apparent positive association between prior cohabitation and subsequent marital dissolution is due to the positive residual correlations between the hazard of marital dissolution and the hazards of entry into cohabitation and moving from cohabitation to marriage (Table 3). Women with an above-average risk of marital breakdown are selected into cohabitation and, once cohabiting, have an above-average propensity to marry; the presence of these women inflates the dissolution rate among previous cohabitors. For similar reasons, the positive estimated effects of previous marriage and previous cohabitation with *either* the current or a previous partner change in the same direction when we move to the multiprocess model, but are significant in neither model. Our findings are consistent with those of Lillard et al. (1995) who concluded that a positive association between premarital cohabitation and marital separation was due to selection of women with a high dissolution risk into cohabitation. Lillard et al. also found no impact of previous marriage on the separation rate of later marriages. In contrast Steele et al. (2005a) found that marriages formed by the 1958 cohort before age 42 were at decreased risk of dissolution if the woman had been married before.

Effects of other characteristics on partnership transitions

Next, we discuss briefly the effects of the other explanatory variables on partnership transitions, and contrast our findings with those of previous studies. We interpret estimates from the multiprocess model, although we note the coefficients of these variables, and their standard errors, are very similar for both specifications of the model.

Duration effects

In the equations for partnership formation, we allow the effect of the duration unpartnered to differ for first and subsequent partnerships. In Table 4, the estimated main effects for duration and duration-squared indicate a sharp increase in the hazards of first entry into both cohabitation and marriage; both hazards peak in the early twenties before starting to decrease[1]. Aasve et al. (2004) found that the hazard of union formation increases until age 21 before decreasing. As they did not consider separate duration or age effects for the first and higher order unions, their estimated hazard is likely to reflect mainly entry into the first partnership. A very similar pattern was found by Berrington and Diamond (1999) in a study of first partnership formation among the 1958 birth cohort. We find a weakly negative effect of duration on the hazard of repartnering. Thus the longer a woman remains single, the lower her probability of cohabiting with or marrying a new partner.

Among cohabitors, the hazard of marriage initially increases with duration (Table 5) reaching a peak at about three years before starting to decline. The hazard of dissolution increases with partnership duration for both cohabitation and marriage (Tables 5 and 6). Steele et al. (2005a) also found weakly positive duration effects on the dissolution risks of married and cohabiting women born in 1958.

Age

As noted in Section 4.2, the effects of age and duration on the formation of the first partnership are confounded because we model transitions from age 16 for all women. The hazard of forming a second or higher order cohabitation increases up to age 21, then remains constant (Table 4). There is no effect of age on the hazard of remarriage, but we note that few cohort members married more than once before age 30 (Figure 1). Among cohabitors, the odds of marriage increases until age 27 (Table 5). Although Ermisch and Francesconi (2000) found no relationship between age and the odds of marriage among cohabiting women of the BHPS, most other authors report some effect of age. In a study based on members of the 1958 cohort in their first partnership, Berrington (2001) reports a decreased marriage rate among those who started to cohabit during their teens. However, when all cohabitations to age 42 are considered for the same cohort, it is cohabitors in their early thirties who are the least likely to marry (Steele et al. 2005a).

In common with most other studies of partnership dissolution (e.g. Berrington 2001; Berrington and Diamond 1999; Ermisch and Francesconi 2000; Kiernan and Cherlin 1999; Steele et al.

[1] The maximum of the baseline hazard for a given transition is calculated by taking the first derivative of the estimated quadratic function in duration.

2005a, 2005b), we find that women who partner at a young age are at increased risk of separation from both cohabitation and marriage (Tables 5 and 6).

Fertility status

There is strong evidence that pregnancy hastens entry into cohabitation and marriage among unpartnered women (Table 4), and the transition from cohabitation to marriage (Table 5). These findings are consistent with previous studies of partnership formation (Berrington 2001; Blossfeld et al. 1999; Steele et al. 2005a, 2005b). Like most earlier studies, we also find that the marriage rate is lower among cohabitors who have a young child together (Table 5); this effect has been attributed to selection, whereby couples with a favourable attitude towards marriage marry before the birth leaving behind those who, for whatever reason, do not wish to marry (e.g. Ermsich and Francesconi 1999). Further, we find that the presence of a young (preschool) child fathered by a previous coresident partner delays cohabitation (Table 4), while the presence of an older child from a previous partnership inhibits marriage among both single and cohabiting women (Tables 4 and 5).

We find stabilising effects of pregnancy and the presence of children fathered by the current partner on both marriage and cohabitation (Tables 5 and 6), effects which are broadly consistent with those reported by Steele et al. (2005a) for the 1958 cohort. However, there is some suggestion that the presence of a school age child from a previous partnership increases the risk that a cohabitation dissolves, and evidence that married women with a child from a non-coresidential relationship are at increased risk of separation.

Education

The hazard of partnership formation decreases with cumulative years of education (Table 4). In an analysis of first partnership formation in the 1958 cohort, Berrington (2003) also found that women with little or no post-compulsory education form partnerships earlier than those with higher qualifications, which might reflect better employment prospects for more educated women. A negative effect of education on repartnering behaviour would also be expected if educated women enjoy greater financial independence and are therefore better able to afford to live alone than their less educated counterparts. Turning to our other educational indicator, current enrolment in full-time education, we find negative effects of enrolment on the hazard of partnership entry (Table 4) and on the odds of marriage among cohabitors (Table 5). Our findings agree with those from previous research (Berrington 2003; Ermisch and Francesconi 2000).

The direction of the effect of attainment on the odds of dissolution is more difficult to predict. To the extent that more educated women have better employment opportunities, one might expect that such women would be less likely to remain in an unsatisfactory partnership than those with little education, leading to a positive effect of education on dissolution. On the other hand, having a low level of education is associated with social disadvantage, which is usually found to have a destabilising effect on partnerships. While the empirical evidence from other studies (Berrington and Diamond 1999; Steele et al. 2005a) supports the latter hypothesis, we find a non-monotone relationship between education and the risk of dissolution for cohabitation (Table 5), and no effect of education on the risk of marital separation (Table 6). However, there is evidence that cohabitors in full-time education experience a higher risk of separation; a similar finding is reported by Berrington and Diamond (1999) and Ermisch and Francesconi (2000). Taken together with the negative effect of educational enrolment on partnership formation, the findings from this and previous studies suggest that full-time education is incompatible with a coresidential relationship.

Family background

In common with most previous research (Aasve et al. 2004; Berrington and Diamond 2000; Steele et al. 2005a, 2005b) we find weak effects of social class on partnership formation and dissolution. This is perhaps unsurprising given that the effects of family background are likely to be mediated by individuals' own characteristics and experiences, particularly childbearing and education. We find that having a father in a professional, managerial or skilled occupation is associated with delayed entry into cohabitation. There is also evidence that women from less advantaged backgrounds (social classes IV and V) are at decreased risk of experiencing partnership dissolution, whether they are married or cohabiting.

We find that the experience of parental separation during childhood is associated with earlier cohabitation and later marriage, lower odds of moving from cohabitation to marriage, and higher risk of dissolution (Tables 4-6). Similar effects have been found for the 1958 cohort (Berrington and Diamond 1999; Kiernan and Cherlin 1999; Steele et al. 2005a, 2005b). Thus there is evidence that women who experience marital breakdown during childhood favour cohabitation over marriage when partnering themselves, and have more fragile partnerships than women who lived with both parents throughout childhood. The effects for the 'missing' category are also worthy of mention. Compared to women whose parental status is known, this small group of women stayed single for longer, were less likely to marry a cohabiting partner, and experienced a higher risk of marital dissolution. Clearly missingness on this particular indicator cannot be assumed random. As such, the estimated effects of parental separation should be interpreted with caution.

There is some evidence of regional variation in partnership transition rates. We find that single women born in the north of England, Scotland or Northern Ireland[2] are the least likely to cohabit and the most likely to marry (Table 4). In their study of first partnership formation in the 1958 cohort, Berrington and Diamond (2000) considered the effect of region at age 16 and found that men and women from London or South East England at age 16 cohabited sooner and married later than their counterparts from other parts of Britain. We also find that, compared to women born in other parts of Britain, cohabitations formed by women from London and the South East are shorter in duration and less likely to be converted to marriage (Table 5). There is no regional variation in the risk of marital dissolution in this cohort (Table 6), nor in the 1958 cohort (Berrington and Diamond 1999; Steele et al. 2005a).

6. Discussion

After allowing for selection on woman-specific unobservables, we find that previously married cohabitors are less likely to marry than never-married cohabiting women. There is, however, no effect of previous partnership breakdown on the odds that a later cohabitation or marriage dissolves. We also find that couples who lived together before marriage are no more or less likely to separate than couples who married directly. These findings are broadly consistent with those of Steele et al. (2005a) for the 1958 birth cohort between ages 16 and 42, although that study did not model partnership outcomes jointly with their formation. Taken together, the results for both cohorts suggest that previous experience of marital dissolution makes women more cautious when considering remarriage, but after repartnering they are at no greater risk of separation than women in their first partnership. Although the presence of unobserved heterogeneity in the dissolution risks of cohabiting and married women implies that there are women who are more prone to separate than others, it is not the experience of dissolution that places them at greater risk in their later partnerships. In other words, partnership dissolution does not appear to have an internally generated dynamic.

Our results should be interpreted in light of the following limitations. The analysis is restricted to women and the conclusions about the effects of previous partnership experience may not generalise to men. For example, Haskey (1999) reports that among divorcees men are more likely to remarry than women. The analysis is also limited to the early partnership transitions of a single birth cohort although, as noted above, our findings are broadly consistent with those for the 1958 cohort to age 42 (Steele et al. 2005a). Finally, we assume that the residual correlation between the hazards of partnership formation and dissolution is due to their shared dependency on unobserved characteristics that are fixed between ages 16 and 30. We do not allow for selection on transient unobservables, which would include time-varying attributes of women (other than age, fertility status, and educational enrolment and attainment) and partner characteristics that vary between different partners of the same woman.

A major difference between this study and others that have used a joint modelling approach (Aassve et al. 2004; Goldstein et al. 2004) is our treatment of cohabitation and marriage as separate partnership states. This decision is supported by our finding that the characteristics associated with partnership formation and dissolution vary according to the legal status of the partnership. Starting

[2] Note that our sample contains only five respondents born in Northern Ireland (as those who who were born and

with formation we find that, among unpartnered women, those who are embarking on their second partnership are more likely to cohabit and less likely to marry than women entering their first partnership. We also find that the presence of a young child from a previous partnership is associated with delayed cohabitation but not marriage. Shotgun marriages are much more likely among unpartnered women than among women who are already cohabiting with the father. The effects on the risk of dissolution of previous partnership experience and children are similar for cohabitation and marriage. For both forms of partnership, there is no evidence of an association between a previous partnership breakdown and the risk of subsequent dissolution, while the presence of a young child appears to have a cementing effect. Nevertheless, overall, marriages are more stable than cohabitations and there are differences in the effects of other predictors by type of partnership. For example, pregnancy has a stronger stabilising effect for married couples, and the positive association between parental separation and the respondent's risk of dissolution is stronger for marriage than for cohabitation.

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Table 1. Distribution of explanatory variables included in the final models

	Single (%)	Cohabiting (%)	Married (%)
Current partnership and partnership history[a]			
Age in years[b]			
16-18	34.2	13.3	12.6
19-21	27.9	28.7	34.6
22-24	19.5	27.9	33.0
25-27	13.5	22.9	18.1
28-30	4.9	7.2	1.7
Previously partnered	24.2	23.8	74.5
Previously married	-[c]	5.8	2.8
Previous cohabitation			
Any	-	18.0	71.7
With current partner only	-	-	61.7
With previous partner only		-	0.9
With current and previous partner(s)	-	-	9.1
Current fertility status[d]			
No children	90.2	62.9	36.4
Currently pregnant	2.6	11.7	21.6
Preschool age child(ren) with current partner	-	23.5	47.5
School age child(ren) with current partner	-	5.1	13.0
Preschool age child(ren) with previous partner	2.0	1.9	0.4
School age child(ren) with previous partner	1.2	2.8	1.1
Child(ren) with non-coresident partner	6.6	3.8	1.6
Education[d]			
Currently enrolled in full-time education	28.3	4.3	1.2
Number of post-16 years of education			
0	33.7	38.9	38.1
1	22.5	19.0	22.2
2	18.1	17.2	18.2
3-5	19.3	16.7	14.5
6+	6.4	8.2	7.0
Family background[e]			
Paternal social class at birth			
I and II	17.9	16.7	17.0
III	56.6	57.8	57.8
IV and V	18.2	18.3	18.6
No father figure	3.5	3.6	3.0
Missing	3.8	3.6	3.6
Family disruption before age 16			
No	20.5	23.0	18.1
Yes	77.2	76.0	81.1
Missing	2.3	1.0	0.8
Region of residence at birth			
London and South East	28.3	29.3	26.5
Scotland, Northern England and Northern Ireland	37.5	36.1	37.1
Wales and Midlands	21.1	21.5	21.8
South and East	10.6	11.1	12.0
Overseas or missing	2.5	2.0	2.6
Number of six-month intervals	92483	30603	28300
Number of episodes	7230	4833	2969
Number of women	5477	3962	2886

[a]Defined at the episode level; [b]Defined as current (time-varying) age in equations for partnership formation, and age at the start of the partnership in equations for transitions out of cohabitation and marriage;
[c]'-' indicates that a variable is not included in the equation for transitions out of that partnership state;
[d]Treated as time-varying in the analysis. Here, percentage distributions for all time-varying variables are based on the total number of six-month intervals spent in each partnership state; [e]Defined at the woman level.

Table 2. Distribution of years to partnership transitions

Transition	Lower quartile	Median	Upper quartile	No. episodes
Partnership formation				
Cohabitation (1st partnership)	4.8	7.8	12.0	5574
Cohabitation (2nd+ partnership)	1.4	3.3	7.9	1998
Marriage	11.4	-*	-	7572
Outcomes of cohabitation				
Marriage	1.9	4.3	10.9	5096
Separation	2.8	7.0	-	5096
Marital separation	8.5	-	-	3215

* The median and upper quartile for marital separations cannot be estimated due to women being followed only to age 30.

Table 3. Estimated random effects covariance matrix from the multiprocess model

	Partnership formation		**Outcomes of cohabitation**		**Marital separation**
	Cohabitation $u^{S(1)}$	Marriage $u^{S(2)}$	Separation $u^{C(1)}$	Marriage $u^{C(2)}$	u^{M}
Partnership formation					
Cohabitation $u^{S(1)}$	0.352 (0.252, 0.465)				
Marriage $u^{S(2)}$	0.029 (-0.167, 0.192) **0.066**	0.610 (0.210, 1.317)			
Outcomes of cohabitation					
Separation $u^{C(1)}$	-0.148 (-0.390, 0.012) **-0.322**	-0.085 (-0.420, 0.255) **-0.148**	0.585 (0.258, 1.281)		
Marriage $u^{C(2)}$	0.065 (-0.065, 0.197) **0.141**	0.107 (-0.201, 0.388) **0.173**	0.002 (-0.326, 0.253) **0.029**	0.586 (0.272, 1.104)	
Marital separation u^{M}	0.301 (0.127, 0.538) **0.487**	0.079 (-0.390, 0.420) **0.111**	-0.042 (-0.526, 0.544) **-0.072**	0.381 (0.007, 0.832) **0.457**	1.116 (0.436, 2.234)

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 50 000 MCMC samples, with a burn-in of 5 000.

Table 4. Estimated coefficients (and standard errors) from multilevel event history models of partnership formation

	Cohabitation		Marriage	
	Single process	Multiprocess	Single process	Multiprocess
Constant	-4.664 (0.049)	-4.689 (0.053)	-6.663 (0.190)	-6.643 (0.169)
Characteristics of current partnership and partnership history				
Previously partnered	0.685 (0.090)	0.734 (0.107)	-0.220 (0.405)	-0.147 (0.410)
Duration unpartnered[a,b]	0.136 (0.007)	0.141 (0.007)	0.239 (0.021)	0.233 (0.018)
(Duration unpartnered)^2	-0.009 (0.001)	-0.009 (0.001)	-0.017 (0.003)	-0.017 (0.003)
Duration*(previously partnered)	-0.172 (0.021)	-0.170 (0.021)	-0.252 (0.114)	-0.242 (0.118)
Duration^2*(previously partnered)	0.006 (0.002)	0.006 (0.002)	-0.044 (0.019)	-0.043 (0.020)
Current age (years)[a] (ref=22-24)				
16-18	-0.458 (0.347)	-0.472 (0.350)	2.018 (1.370)	1.991 (1.418)
19-21	-0.276 (0.118)	-0.276 (0.117)	0.203 (0.650)	0.208 (0.658)
25-27	-0.041 (0.083)	-0.037 (0.084)	-0.077 (0.469)	-0.107 (0.467)
28-30	-0.118 (0.114)	-0.115 (0.114)	0.354 (0.564)	0.313 (0.558)
Current fertility status[a]				
Currently pregnant	0.931 (0.062)	0.936 (0.063)	1.083 (0.150)	1.081 (0.151)
Preschool child with previous partner	-0.392 (0.101)	-0.487 (0.101)	0.524 (0.440)	0.520 (0.445)
School age child with previous partner	-0.097 (0.124)	-0.190 (0.126)	-2.364 (1.303)	-2.364 (1.297)
Child with non-coresident partner	-0.119 (0.061)	-0.092 (0.061)	-1.067 (0.199)	-1.084 (0.198)
Education[a]				
Currently in full-time education	-0.564 (0.057)	-0.557 (0.058)	-0.777 (0.147)	-0.778 (0.147)
Post-16 years of education (ref=0)				
1	-0.107 (0.047)	-0.116 (0.048)	0.172 (0.104)	0.180 (0.103)
2	-0.160 (0.051)	-0.174 (0.051)	-0.159 (0.116)	-0.145 (0.114)
3-5	-0.256 (0.054)	-0.271 (0.053)	-0.437 (0.132)	-0.412 (0.121)
6+	-0.160 (0.072)	-0.177 (0.072)	-0.214 (0.169)	-0.179 (0.159)
Family background				
Paternal social class at birth (ref=III)				
I and II	-0.190 (0.047)	-0.203 (0.050)	-0.062 (0.105)	-0.058 (0.104)
IV and V	-0.013 (0.046)	-0.016 (0.048)	0.129 (0.103)	0.124 (0.101)
No father figure	-0.212 (0.099)	-0.226 (0.100)	0.319 (0.275)	0.331 (0.267)
Missing	-0.018 (0.100)	-0.019 (0.105)	-0.013 (0.234)	-0.018 (0.233)
Family disruption before 16 (ref=no)				
Yes	0.306 (0.045)	0.338 (0.046)	-0.578 (0.132)	-0.594 (0.129)
Missing	-1.505 (0.165)	-1.544 (0.169)	-1.190 (0.306)	-1.146 (0.295)
Region at birth (ref=London , SE)				
Scotland, North, NI	-0.166 (0.041)	-0.174 (0.044)	0.207 (0.096)	0.211 (0.096)
Wales, Midlands	-0.040 (0.046)	-0.039 (0.049)	0.211 (0.109)	0.213 (0.111)
South, East	0.119 (0.057)	0.117 (0.060)	0.139 (0.144)	0.133 (0.142)
Overseas, missing	-0.223 (0.132)	-0.252 (0.140)	0.987 (0.240)	0.989 (0.239)

Note: The estimated coefficients and their standard errors are the means and standard deviations of parameter values across 50 000 MCMC chains, after a burn-in of 5 000.

[a]Time-varying covariate.
[b]Duration measured in six-month intervals.

Table 5. Estimated coefficients (and standard errors) from multilevel event history models of the outcomes of cohabitation

	Separation		Marriage	
	Single process	Multiprocess	Single process	Multiprocess
Constant	-5.017 (0.114)	-5.033 (0.129)	-4.227 (0.083)	-4.255 (0.089)
Characteristics of current partnership and partnership history				
Partnership duration[a,b]	0.093 (0.020)	0.094 (0.020)	0.131 (0.021)	0.145 (0.022)
(Partnership duration)^2	-0.012 (0.003)	-0.012 (0.003)	-0.025 (0.003)	-0.025 (0.003)
Previously married	0.254 (0.186)	0.388 (0.352)	-0.258 (0.170)	-0.674 (0.308)
Previously cohabited	-0.266 (0.138)	-0.173 (0.138)	-0.020 (0.111)	-0.010 (0.135)
Current age (years)[a] (ref=22-24)				
16-18	0.291 (0.101)	0.405 (0.132)	-0.300 (0.087)	-0.384 (0.114)
19-21	0.135 (0.082)	0.203 (0.094)	0.012 (0.064)	-0.034 (0.078)
25-27	0.144 (0.095)	0.089 (0.102)	0.144 (0.073)	0.183 (0.083)
28-30	0.328 (0.225)	0.250 (0.230)	-0.191 (0.228)	-0.129 (0.239)
Current fertility status[a]				
Currently pregnant	-0.688 (0.117)	-0.691 (0.117)	0.224 (0.069)	0.225 (0.070)
Preschool child with current partner	-0.166 (0.081)	-0.177 (0.082)	-0.331 (0.066)	-0.327 (0.067)
School age child with current partner	-0.459 (0.195)	-0.458 (0.193)	-0.177 (0.160)	-0.186 (0.159)
Preschool child with previous partner	-0.100 (0.245)	-0.078 (0.245)	-0.076 (0.201)	-0.074 (0.205)
School age child with previous partner	0.321 (0.181)	0.351 (0.185)	-0.595 (0.197)	-0.612 (0.202)
Child with non-coresident partner	0.016 (0.174)	-0.002 (0.173)	-0.238 (0.144)	-0.232 (0.149)
Education[a]				
Currently in full-time education	0.270 (0.131)	0.267 (0.133)	-0.783 (0.169)	-0.772 (0.168)
Post-16 years of education (ref=0)				
1	0.292 (0.088)	0.304 (0.092)	0.062 (0.072)	0.062 (0.074)
2	0.142 (0.094)	0.163 (0.100)	0.039 (0.074)	0.025 (0.077)
3-5	0.354 (0.094)	0.391 (0.106)	-0.071 (0.080)	-0.098 (0.086)
6+	0.021 (0.134)	0.059 (0.143)	-0.020 (0.103)	-0.062 (0.111)
Family background				
Paternal social class at birth (ref=III)				
I and II	0.069 (0.091)	0.082 (0.092)	0.087 (0.074)	0.085 (0.079)
IV and V	-0.312 (0.091)	-0.318 (0.092)	-0.008 (0.067)	-0.008 (0.071)
No father figure	0.041 (0.164)	0.063 (0.170)	0.188 (0.150)	0.187 (0.158)
Missing	0.184 (0.181)	0.190 (0.184)	-0.012 (0.164)	-0.008 (0.170)
Family disruption before 16 (ref=no)				
Yes	0.160 (0.075)	0.138 (0.079)	-0.300 (0.069)	-0.302 (0.073)
Missing	0.902 (0.253)	0.991 (0.266)	-0.866 (0.351)	-0.932 (0.369)
Region at birth (ref=London, SE)				
Scotland, North, NI	-0.188 (0.077)	-0.185 (0.079)	0.104 (0.064)	0.109 (0.067)
Wales, Midlands	-0.209 (0.091)	-0.219 (0.094)	0.187 (0.074)	0.206 (0.077)
South, East	-0.144 (0.110)	-0.152 (0.116)	0.291 (0.086)	0.319 (0.091)
Overseas, missing	-0.185 (0.257)	-0.187 (0.263)	0.040 (0.219)	0.036 (0.233)

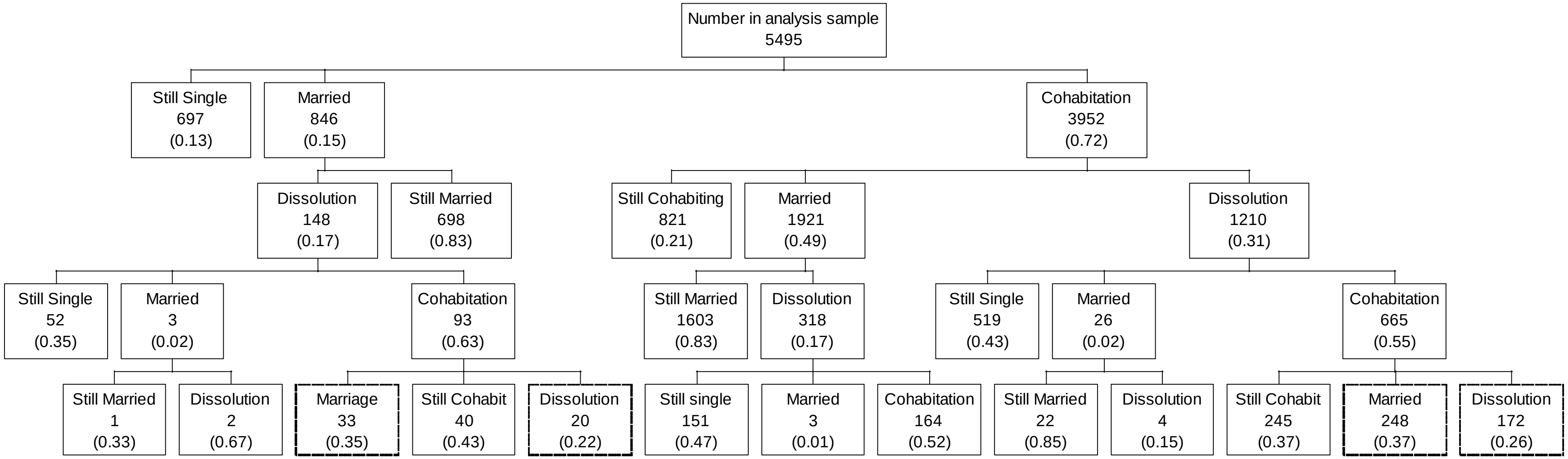
[a]Time-varying covariate.
[b]Duration measured in six-month intervals.

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

	Single process	Multiprocess
Constant	-5.979 (0.182)	-6.111 (0.271)
Characteristics of current partnership and partnership history		
Partnership duration[a,b]	0.012 (0.017)	0.044 (0.023)
(Partnership duration)^2	-0.001 (0.002)	-0.001 (0.003)
Previously married	0.462 (0.327)	-0.343 (0.441)
Previously cohabited with … (ref=no partner)		
Current partner only	0.164 (0.115)	-0.326 (0.264)
Previous partner(s) only	0.472 (0.564)	0.283 (0.677)
Current and previous partner(s)	0.611 (0.216)	-0.046 (0.433)
Current age (years)[a] (ref=22-24)		
16-18	0.974 (0.155)	0.841 (0.193)
19-21	0.389 (0.127)	0.299 (0.152)
25-27	-0.132 (0.221)	0.038 (0.246)
28-30	0.235 (0.832)	0.404 (0.880)
Current fertility status[a]		
Currently pregnant	-1.229 (0.172)	-1.248 (0.173)
Preschool child with current partner	-0.292 (0.106)	-0.368 (0.119)
School age child with current partner	-0.213 (0.180)	-0.271 (0.191)
Preschool child with previous partner	-0.644 (0.836)	-0.916 (0.909)
School age child with previous partner	0.157 (0.419)	0.088 (0.465)
Child with non-coresident partner	1.155 (0.234)	1.317 (0.287)
Education[a]		
Currently in full-time education	-0.331 (0.486)	-0.361 (0.506)
Post-16 years of education (ref=0)		
1	0.057 (0.123)	0.063 (0.146)
2	0.100 (0.132)	0.077 (0.153)
3-5	0.039 (0.157)	-0.019 (0.181)
6+	-0.172 (0.243)	-0.288 (0.271)
Family background		
Paternal social class at birth (ref=III)		
I and II	0.027 (0.140)	-0.009 (0.160)
IV and V	-0.251 (0.126)	-0.267 (0.149)
No father figure	-0.209 (0.255)	-0.216 (0.299)
Missing	-0.469 (0.309)	-0.583 (0.356)
Family disruption before 16 (ref=no)		
Yes	0.275 (0.122)	0.303 (0.147)
Missing	0.810 (0.443)	0.843 (0.520)
Region at birth (ref=London, SE)		
Scotland, North, NI	-0.189 (0.121)	-0.234 (0.143)
Wales, Midlands	-0.244 (0.135)	-0.258 (0.160)
South, East	0.045 (0.150)	0.138 (0.179)
Overseas, missing	-0.270 (0.372)	-0.270 (0.416)

[a]Time-varying covariate.
[b]Duration measured in six-month intervals.

Figure 1. Partnership transitions of women in the analysis sample before age 30



Notes:

(i) Each box shows the number of women and, in brackets, the conditional probabilities of making each possible transition given the most recent transition. For example, of the 3952 women who cohabited with their first partner 21% were still cohabiting with him at interview, 49% had married and 31% had separated.

(ii) For simplification, not all trajectories are shown. The boxes in the final row with dashed borders contain women who may have had further partnership transitions which are not shown.