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Changes in the Relationship between the Outcomes of Cohabiting Partnerships and Fertility among Young British Women: Evidence from the 1958 and 1970 Birth Cohort Studies

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Abstract

We investigate the effects of a range of time-varying fertility indicators, including pregnancy, and the presence and characteristics of children, on the outcomes of nonmarital unions for two cohorts of British women. We compare the effect of conceptions and births on the odds that a cohabiting partnership is dissolved or that it is converted to marriage for women born in 1958 and 1970. The analysis uses a multilevel competing risks model to allow for multiple partnerships and conceptions, and to distinguish between two outcomes of cohabiting unions (separation and marriage). We also use a multiprocess model, in which the outcomes of cohabitation are modelled simultaneously with fertility, to allow for the potential joint determination of partnership and childbearing decisions. The analysis is based on partnership and birth histories between the ages of 16 and 29, and social background, in the National Child Development Study and the 1970 British Birth Cohort Study.

Key words: Cohabitation, partnership dissolution, fertility, competing risks, multilevel modelling, simultaneous equation modelling

Introduction

In Britain, as in much of Western Europe, there has been a dramatic rise in unmarried cohabitation in recent decades (Ermisch 2005; Ermisch and Francesconi 2000a; Haskey 2001; Kiernan 2001; Murphy 2000). While cohabitation was once the domain of previously married couples, it is now the most prevalent form of first union among young people, adopted by around three-quarters of British men and women entering their first partnership in the early 1990s (Ermisch and Francesconi 2000a). Not only has incidence increased, but couples are cohabiting for longer. Murphy (2000) found that the median duration of cohabitation increased by about one year between 1987 and 1995 to almost three years, having remained almost constant for the previous decade. Although cohabitations still tend to be much shorter in duration than marriages, Murphy suggests that this recent increase in duration may signal a change in the nature of cohabitation, with a move towards cohabitation being viewed as more of an alternative to marriage. The increase in the prevalence and duration of cohabitation has been accompanied by a rise in childbearing within nonmarital unions. Cohabiting conceptions form an increasing proportion of conceptions outside marriage. In the 1970s, a large proportion of these precipitated ‘shotgun marriages’, where the bride was pregnant at the time of the wedding. By the late 1990s such marriages had become rare (Berthoud et al. 1999) and in 2002 almost two-thirds of extra-marital births in England and Wales were registered by parents living at the same address (Office for National Statistics 2004). According to one comparative study using data from the early 1990s, the marriage rate among women who had their first child during cohabitation was lower in Britain than in any other country in western Europe (Kiernan 2001).

It is not clear what has been driving these trends. Changes in the legal framework, official statistics, attitudes and practice have been running alongside each other. The premium on married couples in the income tax system was abolished (for those under 65) in 2000. There is still a reward for legal marriage in the benefits for widows and widowers in the state retirement pension, though occupational pension funds increasingly recognise cohabitees as eligible for survivor benefits, at the discretion of trustees. However, state insurance and means-tested benefits have been recognising the existence of cohabitation in a less generous vein throughout the post-war era by operating a ‘cohabitation rule’ which disqualifies people ‘living as man and wife’ from claiming benefits intended for single parents or widows living alone. The 1987 Family Law Reform Act abolished the legal status of illegitimacy for children, although unmarried fathers’ parental rights and responsibilities continued to differ from married fathers’, and to evolve under ensuing legislation such as the 1989 Children Act and the 1991 Child Support Act (Kiernan et al. 1998: pp 94-5). The use of the term ‘illegitimacy’ in commentators’ discourse has taken somewhat longer to die out, along with disapproval of extra-marital relationships. The attitudes of the parental generation should be taken into account as well as that of the young adults whose partnership behaviour is under investigation here. Their influence may be weakening, or their attitudes may mirror those of the young. Scott (1999) shows attitudes to pre-marital sex liberalizing within and across cohorts, but the direction of influence between attitude and practise may not be one-way. Ermisch (2005) argues that the changes across time reflect a diffusion of attitudes more tolerant to living together without marriage, operating through a ‘social contagion’ as the idea catches on, led by innovators among students and graduates in the 1970s, but spreading across all social groups over time, with the original social profile eliminated (Ermisch and Francesconi 2000b)[1] or even reversed (McRae 1999), as the labour market for least skilled men deteriorated. Certainly by the time of the Millenium Cohort births in 2000-1, the 25 per cent of all mothers who were cohabitants tended to be younger and less advantaged than married mothers (Kiernan 2004). It is suggested that their capacity to make the long-term commitments involved in marriage, or their attractiveness as long-term partners, is eroded particularly by the relative gains in education and employment made by women. (Smock and Manning (1997) put this argument in the US context.) Another argument for the increasing duration of cohabitations could be the falling credibility of formal marriage as a life-long contract in the face of rising divorce rates.

In this paper we investigate the outcome of cohabiting unions in two cohorts of British women between the ages of 16 and 29 and its link with childbearing. We make inter-cohort comparisons of the effects of pregnancy and the presence and characteristics of children on the probability that a cohabiting partnership, once formed, is dissolved or converted to marriage. We take age 29 as our cut-off because the 1970 cohort were last interviewed at this age. Although the most recent wave of the 1958 cohort study collected event histories up to age 42 (analysed by Steele et al. 2005), we apply the same cut-off to both cohorts for ease of comparison. We acknowledge, however, that although the cohorts are observed for the same chronological ages, these may not correspond to comparable 'social' ages. Previous comparisons have found that, relative to the 1958 cohort, the more recent cohort delayed young adult transitions, including leaving education, leaving home, and partnership and family formation (Berrington 2003; Ferri and Smith 2003; Makepeace et al. 2003; Schoon et al. 2001). The 1970 cohort entered their first partnership later, married later, and were more likely to cohabit, and to cohabit for longer (Berrington 2003; Ferri and Smith 2003). There have also been changes in the timing and context of childbearing. Berrington (2003) reports that the later cohort had their first child later and were more likely to have children outside marriage, either during cohabitation or while unpartnered. However, these trends across cohorts are not uniform. For example, similar rates of teenage residential partnership are displayed by both cohorts (Ferri and Smith 2003), and there has been little delay in the timing of first partnership among the least educated. The primary aim of our study is a detailed examination of inter-cohort differences in young women's transitions after forming their first and any subsequent cohabitations, and the extent to which these transitions depend on past and current fertility.

One methodological issue that must be considered when assessing the impact of fertility outcomes on transitions from cohabitation is the possibility that decisions about childbearing and partnerships are subject to shared influences. 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. Inevitably, some of the factors that drive these decisions will be unobserved, and there may be correlation between outcomes of the fertility and partnership processes which cannot be explained by their dependency on observed characteristics alone. For example, an individual's latent view of cohabitation, whether as a precursor or alternative to marriage, will influence not only their odds of marriage but also the probability that they start childbearing before marriage. To put the problem another way, fertility outcomes may be endogenous with respect to partnership outcomes. Failure to account for

[1] Ermisch and Francesconi (2000b) find a positive social gradient in the proportion who start their first partnership as cohabitation rather than marriage for the 1950-62 cohort; this is not present for the 1963-76 cohort, which is observed only up to age 24.

this source of endogeneity will lead to biased estimates of the effects of pregnancy and the presence and characteristics of children on the outcomes of cohabitation.

A common approach to allow for endogeneity or selection effects is to model the endogenous explanatory variable(s) jointly with the outcome of interest using a simultaneous equation model, otherwise known as a multiprocess model. Such models allow explicitly for selection on unobservables by introducing a correlation between the residual components of each equation. Multiprocess modelling of event history data was first proposed by Lillard and Waite (1993), with an application to an analysis of marital dissolution and marital fertility. Other examples of multiprocess modelling include a study of the interrelationships between nonmarital fertility and the formation of marital and cohabiting unions in the US (Brien et al. 1999), later extended by Upchurch et al. (2002) to include the processes of marital dissolution, marital fertility and educational transitions, and a British study of the link between union formation and dissolution, and fertility and employment decisions (Aassve et al. 2006). Only one study to date has used multiprocess modelling to examine the link between childbearing and the outcomes of cohabiting unions. Steele et al. (2005) extend Lillard and Waite's framework to model jointly transitions from marital and cohabiting unions and fertility within those unions (using data on the 1958 cohort to age 42). The inclusion of cohabitation requires a generalisation of Lillard and Waite's model in order to accommodate competing risks. In this paper, we adopt the methodological approach of Steele et al. (2005), but focus on the outcomes of cohabiting unions and conceptions during cohabitation, and on the comparison of two cohorts passing through the relevant ages at different stages of the trend outlined above. We aim to contribute to the understanding of increasing cohabitation, and more childbearing within it, by studying the dynamics of transitions out of the state.

Our paper builds upon previous research on cohabitation in several ways. First, we seek to contribute to the literature on inter-cohort comparisons of the effects of fertility events on the outcomes of cohabiting partnerships. There has been little research on this issue in Britain, presumably because event history data for the 1970 cohort have only recently become available. Berrington (2003) makes cross-cohort comparisons in the proportion of women who experience marriage, cohabitation and pregnancy (followed by a birth within marriage or cohabitation, or lone parenthood) as their first family event. She then considers the first event following a woman's first entry into cohabitation, and finds a sharp decline in the proportion who marry after a cohabiting conception to have a marital birth. In our paper, we investigate the impact of a conception, and the duration of pregnancy, on both the odds that a cohabitation is converted to marriage before birth and the risk of dissolution. In addition, we consider the effects of the presence, age and parentage of children on these outcomes. While most studies of cohabitation outcomes focus on the first, usually premarital, partnership we consider *all* episodes of cohabitation that begin before a woman's 30th birthday. A multilevel model is used to allow for repeated transitions, and we control for a woman's history of marriage and cohabitation prior to the current partnership. Finally, we use a multiprocess modelling approach in which we estimate simultaneous equations for the odds of marriage or dissolution and the odds of a cohabiting conception. As outlined above, this technique adjusts for selection bias due to shared dependency of cohabitation and fertility outcomes on unobserved time-invariant characteristics of the woman.

Previous Research on the Relationship between Childbearing and Exits from Cohabitation

The findings of previous research on the relationship between fertility and transitions from cohabitation in Britain and elsewhere are to some extent inconclusive. Although it is likely that this lack of consensus reflects genuine variation between countries, and between study populations and cohorts within countries, another possible explanation is differences in the fertility indicators used. Some authors consider only the number of children, while others take into account children's characteristics, including their age, sex and whether they were born before or during the current cohabitation. In addition, some researchers examine the effects of the presence of children and being pregnant (and, in a few cases, the duration of pregnancy) on the timing of separation or marriage. Another way in which studies diverge is in their definition of a partnership. Analyses of partnership dissolution do not always distinguish between marriage and cohabitation (e.g. Aassve et al. 2006; Böheim and Ermisch 2001) and some studies of entry into marriage do not consider transitions from cohabitation separately from direct marriage from an unpartnered state (e.g. Brien et al. 1999, Upchurch et al. 2002). We focus on research that treats cohabitation as a separate partnership state. All of the studies discussed below use competing risks event history analysis to consider exits from cohabitation due to separation or marriage, and most focus on the first premarital union. Only Steele et al. (2005) allow for the potential endogeneity of fertility outcomes with respect to transitions from cohabitation.

Studies also differ in the cohort composition of their samples and their treatment of cohort effects. The studies reviewed here are based on a single cohort or a sample covering multiple cohorts. Those that use data on multiple cohorts consider cohort effects on the average odds of the different outcomes, but do not consider interactions between cohort and other explanatory variables. Further, in studies from the early 1990s the experience of cohorts born around 1970 was cut off at an earlier age than in our study. These differences must be considered when comparing our findings, based on separate analyses of two cohorts, with those from pooled analyses of multiple cohorts.

We review the evidence for effects of the presence and characteristics of children on the outcome of cohabiting partnerships. Ermisch and Francesconi (2000a), in an analysis of cohabiting unions among a sample of multiple cohorts of British men and women who cohabited during the early 1990s, find that, compared to childless women, mothers are just as likely to separate but are less likely to marry. Therefore, because cohabiting couples with children are exposed to the high dissolution risk of cohabitation for longer than their childless counterparts, their partnerships are ultimately at a greater risk of separation. When pre-union births are distinguished from children born during cohabitation, they find that births within the union are associated with a lower marriage rate while mothers who had their youngest child before the start of the union have the same odds of separation and marriage as childless women. In their analysis of the 1958 British birth cohort between the ages of 16 and 42 (i.e. between 1974 and 2000), however, Steele et al. (2005) find that having a preschool age child with a cohabiting partner is associated with decreased odds of both separation and marriage. In Canada, Wu and Balakrishnan (1995) find that the presence of children, regardless of whether they were born before or within the cohabitation, reduces the chance that the union is converted to marriage. In contrast, a US study (Manning and Smock 1995) finds that, compared to childless men and women, those with children are *more* likely to marry. These North American studies also reach varying conclusions about the effect of having children on the risk of separation. Like Ermisch and Francesconi, Manning and Smock find no significant effect of the presence of children, while Wu and Balakrishnan report a decrease in the separation rate with the number of children and, for women only, a positive effect of a pre-union birth.

Ermisch and Francesconi (2000a) and Wu and Balkrishnan (1995), among others, suggest that the negative effect of having children together on a cohabiting couple's odds of marriage may be explained by a selection mechanism. Cohabiting couples who are favourably disposed towards marriage, and who are mutually acceptable as marital partners, are likely to marry before they have children. Therefore couples who have children together while cohabiting will be a combination of two types: those with an ideological commitment to cohabitation as an alternative to marriage and an acceptable setting for childrearing (Wu and Balakrishnan 1995), and couples who do not view each other as prospective marital partners (Ermisch 2005). The selection of either type of couple into childbearing within cohabitation will lead to a negative effect of having children on the odds of marriage. Further, if the true 'causal' effect of the presence of children is to reduce the risk of dissolution, selection of the first type will lead to a weaker negative or even positive effect, while the second type of selection will lead to a stronger negative effect.

Blossfeld et al. (1993) offer an alternative interpretation for the negative effect of having children on the odds of marriage. In a study of the outcomes of cohabitations formed during the 1980s in West Germany and the Netherlands, they examine the impact of current fertility status on the rates of marriage and dissolution and, in particular, the change in rates during pregnancy and following a birth. In both countries, they find evidence of an inverse U-shaped relationship with entry into marriage: the marriage rate is low before conception, increases during pregnancy and around the birth, and then decreases to pre-pregnancy levels (or lower in the case of the Netherlands) six months after the birth. Conversely, the risk of dissolution is high before pregnancy (in West Germany), declines during pregnancy, then increases again after the birth (in both countries). Similar patterns have been observed in Britain, for cohabiting first partnerships formed among the 1958 cohort before age 33 (Berrington 2001), and in the US (Manning and Smock 1995). If current pregnancy status is not taken into account, as in the analyses of Ermisch and Francesconi (2000a) and Wu and Balakrishnan (1995), the effect of the presence of children represents a contrast between women with children and a combination of childless women and those who are currently pregnant with the couple's first child. The increased odds of marriage during pregnancy inflate the overall odds for this latter group, leading to an apparent strong negative effect of having children. However, the decline in the postnatal marriage rate to pre-pregnancy levels may be the result of selection. Blossfeld et al. (1999) argue that the conception of a couple's first child prompts a solidifying of their preferences towards marriage. Those who wish to 'legitimise' the birth, and who are favourably inclined towards marriage, will tend to marry during pregnancy. If the child is born during cohabitation, the child is already 'illegitimate' and there is no longer the time pressure to marry. Couples who do not respond to the social pressure to marry during pregnancy, and therefore have a cohabiting birth, may regard cohabitation more as an alternative to marriage or, on the other hand, they may be reluctant to marry each other.

The selection effects described above could operate at the level of the individual or the partnership, or a combination of both. If there are some women who are more inclined to marry than others due to unobserved characteristics that remain fixed over the observation period (e.g. holding more 'traditional' views towards marriage and family), this will lead to selection at the individual level. Such women will be less inclined to have a child during cohabitation with *any* partner and a cohabiting conception with *any* partner is likely to be followed by a marital birth. Selection may also operate at the partnership level, that is, there may be some couples who are more likely to marry than others due to unobserved characteristics of the couple, or an interaction between the unobserved characteristics of each partner. For example, cohabiting couples in a more stable relationship, or an individual whose partner is strongly pro-marriage, may have higher odds of marriage.

In the present study, we investigate the effects on the outcomes of cohabitation of a set of time-varying fertility indicators. We contrast the impact of the presence of children born to the current cohabiting partner, a previous co-resident (marital or cohabiting) partner and a non co-resident partner. We also examine the importance of the age of children, distinguishing between preschool ($\leq$ 5 years old) and older children. In addition, we consider the effects of being pregnant and the current duration of pregnancy on the odds of separation and marriage. Of prime interest are inter-cohort differences in the effects of these fertility indicators on transitions from cohabitation. A multiprocess model is used to account for selection at the individual level by allowing for correlation between the woman-specific residuals in the cohabitation outcomes equations and the cohabiting conception equation. The directions of these residual correlations indicate the nature of any selection effects due to unobserved individual-specific factors. For example, a tendency for women who favour marriage to wait until marriage to start childbearing would lead to a negative residual correlation between the odds of marriage and the odds of a cohabiting conception.

Methods

A discrete-time multilevel multiprocess model for the outcomes of cohabiting unions and fertility within cohabitation

The multiprocess model consists of a system of equations, each defining a discrete-time event history model. The main reason for adopting a discrete-time approach is to allow the use of existing estimation procedures for multilevel discrete response models, which are now implemented in several software packages. (See Steele et al. (2004) for further discussion of the advantages of discrete-time event history models.)

The fact that respondents may live with more than one partner between the ages of 16 and 29, and may have multiple conceptions within those cohabitations, leads to a two-level hierarchical structure with cohabitations and conceptions (at level 1) nested within women (at level 2). We use a multilevel event history model, also known as a *shared frailty* model, to allow for correlation between the durations of cohabitations, and the intervals between conceptions, contributed by the same woman. The model for the outcomes of cohabitation includes woman-specific random effects, representing time-invariant unobserved characteristics that affect the odds of marriage or dissolution for all cohabitations. The random effect variance measures the extent of unobserved heterogeneity between women due to unobserved individual characteristics that are fixed in time.

Our model is a system of three simultaneous equations, one for each cohabitation outcome and a further equation for conceptions within cohabitation. Each equation contains a woman-specific random effect. These random effects are permitted to correlate across equations to allow for correlation between the unobserved woman-level characteristics that affect each process. The model described below is a special case of the more general model proposed by Steele et al. (2005) for analysing transitions from both marriage and cohabitation jointly with fertility within either form of partnership.

Separate models are estimated for each cohort. While it is possible to fit a single model to a pooled dataset, with cohort dummies and their interactions with explanatory variables, this approach is not practically feasible at present due to the already large size of the discrete-time dataset.

Competing risks model for the outcomes of cohabitation. We consider two transitions from the cohabitation state: separation and marriage to the same partner. Denote by $h_{ij}^{C(r)}(t)$ the hazard of a transition of type *r* from cohabitation, in month *t* of episode *i* for individual *j*, where *r*=0 (no transition), 1 (separation), or 2 (marriage). Transitions from cohabitation may be modelled using a multilevel discrete-time competing risks model (Steele et al. 1996) which may be written (omitting subscripts) as:

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

where $\alpha_0^{C(r)} D^{C(r)}(t)$ is a function of cohabitation duration at month *t*, $F(t)$ is a vector of fertility indicators of current pregnancy status and the presence and age of children from the current or a previous partnership, $X^{C(r)}(t)$ are covariates that affect the hazard of a transition of type *r* from cohabitation (described below), and $u^{C(r)}$ are individual and transition-specific random effects.

To estimate (1) each cohabitation duration, D_{ij}^C, is converted to a sequence of D_{ij}^C multinomial responses, $y_{ij}^C(t)$. The response at month *t* is coded 0 if cohabitation episode *i* of woman *j* is still in progress, 1 if separation occurs, and 2 if marriage to the same partner occurs.

Model for fertility within cohabitation. Denote by $h_{ij}^F(t)$ the hazard of a conception during month *t* in cohabitation episode *i* for individual *j*. We consider only those conceptions that lead to a live birth. Conception dates are 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. In a comparison of abortion rates calculated from the NCDS to age 33 and national rates for the same cohort, Berrington (2001) found that the NCDS figures were underreported by 50 per cent.

The model for conceptions within cohabitation is written:

$$\operatorname{logit} h^F(t) = \beta_0^F D^F(t) + \beta_1^F F(t) + \beta_2^F X^F(t) + u^F \quad (2)$$

where $X^F(t)$ are covariates and u^F is an individual-level random effect.

Estimation of the multiprocess model

Equations (1) and (2) 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 the random effects follow a multivariate normal distribution, i.e. $u = (u^{C(1)}, u^{C(2)}, u^{F}) \sim N_3(\mathbf{0}, \Omega_u)$. Correlated random effects would arise if the unobserved characteristics that influence the timing of transitions from cohabitation are correlated with those that affect the hazard of a conception during a cohabiting relationship. Non-zero correlations between elements of $u^{C} = (u^{C(1)}, u^{C(2)})$ and u^{F} would suggest that $F(t)$, the number and/or age of children from the current or a previous partnership, is endogenous with respect to the outcomes of cohabitation.

The discrete-time multiprocess event history model can be framed as a multilevel bivariate discrete response model where for each month *t* of a partnership there are two responses, $y_{ij}(t) = (y_{ij}^{C}(t), y_{ij}^{F}(t))$. The model may be estimated using existing methods for mixtures of binary and multinomial responses (Steele et al. 2004) after defining indicators for the cohabitation 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. 2005). See Browne (2005) for an introduction to MCMC methods for multilevel analysis, and their implementation in *MLwiN*.

Identification of the multiprocess model

Identification of simultaneous equations models typically requires exclusion restrictions to be placed on the covariates. In the case of the model described above, this would involve including in the fertility equation a set of covariates that are excluded from the cohabitation outcome equation. In practice, however, it is difficult to find, variables that affect fertility decisions but do not have direct effects on partnership behaviour. Fortunately, the observation for a subset of women of repeated cohabitation, and more than one conception within cohabitation, 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 (Lillard et al. 1995; Upchurch et al. 2002). After accounting for this residual correlation, the remaining variation in the fertility outcomes $F(t)$ between episodes of cohabitation represents the effects of prior fertility on the outcomes of cohabitation, controlling for selection bias.

Data

The cohort studies and measures

We analyse data from the National Child Development Study (NCDS) and the 1970 British Cohort Study (BCS70), prospective longitudinal studies of all those living in Great Britain who were born in a single week of 1958 and 1970 respectively (Shepherd 1997; Bynner et al. 1997). Since birth, attempts have been made to contact the 1958 cohort on six further occasions (at ages 7, 11, 16, 23, 33 and 42) and the 1970 cohort on five occasions (at ages 5, 10, 16, 26, and 29). In both studies, data were collected from parents, and then cohort members, and a number of supplementary sources. The cohort studies therefore provide a rich source of information on respondents' physical, educational and social development from birth to early adulthood.

Partnership histories have been collected retrospectively at ages 23, 33 and 42 for the 1958 cohort and at age 29 for the 1970 cohort. In the NCDS, respondents were asked at age 33 to recall the start and end dates of all cohabiting relationships and marriages since age 16 which lasted for at least one month. These data were later reconciled with reports at age 23 to form a single partnership history (Di Salvo 1995a). Partnership histories from age 16 were collected from the 1970 cohort at age 29. Therefore, to aid inter-cohort comparisons, histories for the 1958 cohort were curtailed at their 30^{th} birthday. For our analysis, episodes of (nonmarital) cohabitation were extracted from the partnership histories. The dependent variable indicates, for each month of cohabitation, whether separation or marriage has occurred (at which point the episode ends) or whether the cohabiting relationship continues. The very small number of episodes which ended in a partner's death are treated as right-censored, as are cohabitations in progress at the time of interview (or, for the 1958 cohort, their 30^{th} birthday).

In this paper, the explanatory variables of major interest are time-varying indicators of pregnancy status, and the presence and characteristics of children. These variables were constructed from birth history data collected at the same time as the partnership histories (Di Salvo 1995b; Dodgeon 2002). Respondents were asked to identify the other parent of each child, and in particular whether this was the current partner at the time of interview or a previous partner named in the partnership histories. From this information, we are able to distinguish children fathered by the current partner at month t, a previous co-resident partner or a non co-resident partner. In addition to the number and parentage of children, we consider the current age of each child, classified as preschool (younger than five years) or school age (five or older). In calculating the number of children present at each month, we count only those children living with the respondent, using information on the date of leaving home. Finally, we consider an indicator of current pregnancy status and, if pregnant, the duration in trimesters. This information is also used to construct a binary conception indicator, coded 1 in the month that conception occurs and 0 otherwise, which is included as a second dependent variable in the multiprocess models.

Although the impact on cohabitation outcomes of changes in fertility are of prime interest, we adjust for the effects of a range of other factors that have previously been found to predict partnership transitions. We control for characteristics of the current cohabiting partnership, including its duration and the respondent's age at the start of the partnership, and of the partnership history, including indicators for previous marriage and cohabitation. In addition, we consider the number of years of post-compulsory education (treated as time-varying), based on employment histories collected at the same time as the partnership and birth histories, and several family background measures: region of residence at birth[2], father's social class at the respondent's birth, the experience of family disruption before age 16[3], and

[2]Region was represented by dummy variables for the 12 standard regions. Because preliminary analysis found little evidence of regional variation in cohabitation outcomes, or in fertility within cohabiting unions, region was excluded from the final models.

[3]Family disruption includes the experience of parental divorce or any other living arrangement where the father or mother figure was not one of the natural parents.

housing tenure at age 16[4]. Descriptive statistics for all explanatory variables included in the final models are shown in Table 1.

The analysis samples

In common with most other studies of the link between fertility and partnership transitions, our analysis is restricted to women. While the focus on women permits easier comparison with earlier research, there are two additional, pragmatic, reasons for this decision (see also Steele et al. 2005). First, we expect some unreliability in men's reports of children from previous, particularly nonmarital, relationships. Second, the absence of longitudinal information on step-children means that they are excluded from the time-varying counts of the number of children living with a respondent; to the extent that children stay with their mother following a partnership breakdown, this omission will have a greater effect for men than for women.

Our analysis is based on the subsample of women from each cohort who had formed at least one co-residential nonmarital relationship by their 30th birthday. Of the 5800 women from the 1958 cohort interviewed at age 33, 39 per cent had experienced cohabitation before age 30. In BCS70, 5790 women were interviewed at age 29 and 73 per cent of those had cohabited. There are a number of further exclusions: women for whom an accurate partnership history could not be constructed, childless women who had been told by a doctor that they should or could not have children, women with adopted children, and those who had lived with a same-sex partner. The final analysis samples consist of 2650 cohabitation episodes from 2140 women for the 1958 cohort, and 4836 episodes from 3964 women for the 1970 cohort. Thus the mean number of cohabitations per woman is 1.24 and 1.22 for the 1958 and 1970 cohorts respectively.

The NCDS and BCS70, like other longitudinal studies, suffer from attrition. Our analysis is based on the subsamples of original respondents who were successfully interviewed at age 33 (NCDS) and age 29 (BCS70). In each survey, the observed sample represents approximately 70 per cent of the target sample (Plewis et al., 2004), and previous research on the nature of attrition in the cohort studies suggests that respondents are a non-random subsample of the eligible sample. In a study of drop-out in the NCDS, Hawkes and Plewis (2006) report that low reading ability, unstable employment patterns and indicators of disadvantaged circumstances were positively associated with non-response at age 33, although none was a strong predictor. Berrington (2003) found that in both cohorts the socially disadvantaged were the most likely to be lost to follow-up. In addition, she reports that women who began childbearing in their teens are underrepresented among respondents at ages 29 and 33. By controlling for educational attainment and indicators of social disadvantage (paternal social class and the experience of family disruption) in our models, we minimise attrition bias due to these factors. Of course, it is almost certain that there is further nonresponse bias due to the association between attrition and the processes under study. However, it seems reasonable to assume that the nonresponse mechanisms are similar for the two cohorts. Hence estimates from each cohort are subject to the same biases, which are expected to cancel out when looking at inter-cohort differences – the focus of this study.

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[4]Housing tenure was represented by binary variable contrasting owner-occupied accommodation versus any other form of tenure. This variable was not included in the final models due to a high proportion of missing values and, after imputation, a weak association with the outcomes of cohabitation.

Of those successfully interviewed at age 29 and 33, 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 and years of education. 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[5].

Results

Changes in cohabitation

We begin with an inter-cohort comparison of cohabitation behaviour before age 30, including the number and duration of cohabitations and fertility within those partnerships. Table 2 shows the current partnership status of respondents at their 30th birthday, according to their partnership history since age 16. While the percentage of women without a partner at age 29 is only slightly lower for the 1970 cohort (29 per cent compared with 32 per cent for the 1958 cohort), the proportion of those who have *never* partnered has increased. In both cohorts, the majority have had only one co-residential partner and this has remained constant at about 56 per cent. However, among those in their first partnership at age 29, there has been a dramatic decrease in direct marriage without pre-marital cohabitation and a substantial increase in the proportion who do not marry. Overall, nearly a quarter of the 1970 cohort are cohabiting at their 30th birthday, compared to only 8 per cent of the earlier cohort.

The proportion of the cohort who has ever cohabited before age 30 has doubled from approximately one to two thirds. At the same time there have been many other changes. As shown in Table 1, 52 per cent of women in the 1958 cohort had no schooling post 16, compared with 32 per cent of the 1970 cohort. The proportion of women with six or more years of post-compulsory education roughly doubles from 7.4 to 12.5 per cent. To a lesser extent the social backgrounds of their families of origin reflected secular upskilling of the labour force, with 21.2 per cent of fathers in the least skilled class (IV and V) compared with 198.7per cent in BCS70. We do not explicitly allow for cyclical macro-economic change, but note here that the 1970 cohort and their potential partners faced lower chances of employment in their early labour market years than did the previous cohort who were already in their thirties by the time of the recession around 1990 (see Makepeace et al. 2003).

The changes in the social composition of the whole cohort are also reflected in the composition of the samples of cohabitants which we analyse in this paper, but not proportionately. In NCDS, women ending up with the highest educational record were over-represented among cohabitants, while the social class distribution was similar for cohabitants and all women. In the later cohort, all social groups participated in the incoming tide of cohabitation, but it was the less educated and those from less auspicious backgrounds (often the same people, see Bynner and Joshi 2002) who rode its crest. Thirty-eight per cent of the woman-months of cohabitation in the 1970 cohort were contributed by

[5] We used social class information collected at an older age only if was reasonable to assume that the child's father figure at that age was the same person as the father figure at birth. Therefore, imputation was carried out only if the father figure was identified as the natural or adoptive father on both occasions.

the 32 per cent of women who had no post-16 schooling. Women whose families of origin had been disrupted before they were 16 were more likely than other contemporaries to cohabit, in both cohorts, at similar levels of over-representation.

Table 1 presents the social composition of the cohabiting samples in terms of woman months, rather that the number of women who ever experience at least one spell, to reflect the longer cohabiting experience of those whose partnerships started earlier and lasted longer. The less qualified tended to start earlier. When we looked at these distributions among men who have ever cohabited, there is a similar shifting pattern in the social profile, but the fact that the more advantaged eventually 'catch up' means that the differences in incidence are smaller[6]. These data are compatible with the findings of Ermisch and Francsconi (2000b) and McRae (1999) that cohabitation is becoming more common among less privileged women, when comparing the last two decades of the Twentieth Century. However, since the propensity to cohabit and educational composition are moving in opposite directions, the social composition of those who are cohabiting does not change much.

Among those who have cohabited before the end of the observation period, the distribution of the number of cohabitations per woman is very similar for the two cohorts; four-fifths have cohabited only once, and only 2-3 per cent have lived with more than two partners (Table 3). While there has been little change in the frequency of cohabiting unions, however, their duration has increased. The median number of months spent in cohabitation was 25 (SD=31) for the 1958 birth cohort, compared to 34 (SD=33) for the 1970 cohort. Further information on the duration of cohabiting partnerships is provided in Table 4, which shows the cumulative percentage of cohabitations ending in separation or marriage within two and four years. The main determinant of the increase in the duration of cohabitation is the lower rate of marriage among the more recent cohort. This trend is particularly pronounced among cohabitors in their first partnership; for the 1958 cohort, 58 per cent of first cohabiting partnerships were converted to marriage within four years, compared to 41 per cent for the 1970 cohort. There is also a marked cohort effect on the marriage rate for women in their second partnership, but not for the small group who have already experienced at least one marriage and cohabitation. Inter-cohort differences in the rate of dissolution are less dramatic. The only notable difference between cohorts is observed among previously married women in their first cohabitation, a relatively small group in the 1970 cohort for whom premarital cohabitation was the norm.

Next we investigate whether the increase in the duration of cohabiting relationships, and the overall time spent in cohabitation, led to a commensurate rise in the number of conceptions during cohabitation. Table 5 shows, for each cohort, the distribution of the number of cohabiting conceptions (leading to live births) per woman. While the proportion of women who did not conceive with a cohabiting partner remained constant at just under one-fifth, a slightly higher proportion of the 1970 cohort experienced multiple conceptions (8.3 per cent compared to 5.9 per cent of the 1958 cohort). The major change in the fertility behaviour of cohabiting couples, however, is in the relative frequency of conceptions that led to births within cohabitation. For the 1958 cohort, 26.5 per cent of women experienced at least one conception while cohabiting, and 17.5 per cent gave birth during cohabitation. The same

[6] The results reported below suggest that there is little effect of socio-economic factors on transitions *out* of cohabitation. We do not analyse entry *into* cohabitation in this paper. Current research (Kallis et al. 2005) considers the determinants of transitions both in and out of cohabitation and marriage in the 1970 cohort.

proportions for the 1970 cohort were 28 per cent and 24.9 per cent respectively. The main reason for this difference is a greater tendency among the earlier cohort for a cohabiting conception to be followed by marriage before the birth.

Multilevel event history analysis of outcomes of cohabitation and fertility

Two specifications of the multilevel event history model were considered, with separate models fitted for each cohort. In the first model, referred to as the single process model, the random effects in the cohabitation equation, $u^C = (u^{C(1)}, u^{C(2)})$, are assumed to be uncorrelated with the random effect in the fertility equation, u^F. This approach is equivalent to fitting separate event history models for each process, and treats conceptions during cohabitation as exogenous. In the second specification, the multiprocess model, correlations between each element of u^C and u^F are freely estimated. Non-zero cross-process correlations would suggest that cohabiting conceptions are endogenous with respect to the outcomes of cohabitation, in which case the single process model will yield biased estimates of the effects of children conceived during cohabitation on the odds of either or both separation and marriage.

Separate models are estimated for each cohort. While it is possible to fit a single model to a pooled dataset, with cohort dummies and their interactions with explanatory variables, this approach is not practically feasible at present due to the already large size of the discrete-time dataset.

Random effect correlations. Table 6 shows the estimated random effects covariance matrices from the multiprocess models fitted to each cohort. Of particular interest are estimates of the pairwise covariances between u^F ('conception') and each element of $(u^{C(1)}, u^{C(2)})$ ('separation' and 'marriage'), shown in the 'conception' rows of the table. Although, for both cohorts, neither covariance is significantly different from zero at the 5 per cent level, we observe a weak negative residual correlation between the odds of marriage and the chance of a cohabiting conception in the more recent cohort. This negative correlation implies that a high (low) probability of conception during cohabitation is associated with a low (high) marriage rate, a result which is consistent with the idea that women who are less inclined to marry are more likely to initiate childbearing during cohabitation.

It is interesting to note that, for both cohorts, estimates of the residual correlation between the odds of marriage and a cohabiting conception change dramatically after adjusting for the effects of current pregnancy status. When pregnancy status is omitted from the equations for cohabitation outcomes, the estimated correlation between u^F and $u^{C(2)}$ is 0.70 (with 95 per cent interval estimate 0.54 to 0.82) for the 1958 cohort, and 0.17 (-0.16 to 0.45) for the 1970 cohort. The strong positive correlation for the earlier cohort implies that women with above average odds of marrying a cohabiting partner tend also to have an above average propensity to conceive during cohabitation. It transpires, however, that this correlation is explained by the increased odds of marriage during the first six months of pregnancy, as described below. The finding that controlling for pregnancy status leads to a substantial change in the woman-level residual correlation may suggest that the positive association between the occurrence of a cohabiting conception and the likelihood of subsequent marriage is constant across repeated partnerships for the same woman.

Also shown in Table 6 is the correlation between the random effects for the two outcomes of cohabitation, $u^{C(1)}$ and $u^{C(2)}$, estimated as 0.32 and 0.36 for the 1958 and 1970 cohorts respectively. There is weak evidence (at the 10 per cent level of significance) that both correlations are non-zero. A positive correlation implies that women with a high (low) hazard of dissolution tend also to have a high (low) hazard of marrying a cohabiting partner. Thus we might tentatively distinguish between short-term cohabitors (with high odds of both separation and marriage) and longer-term cohabitors (with low odds of exiting cohabitation for either reason).

Effects of fertility indicators. Our finding that the cross-process residual covariances do not differ significantly from zero implies that, conditional on the covariates included in the model, the individual-level unobservables affecting the hazard of a transition from cohabitation may be assumed uncorrelated with the unobservables affecting the hazard of a cohabiting conception. As a result, the estimates from the single and multiprocess models are very similar. We therefore present only the results for one model specification, the multiprocess models. The estimated effects of the fertility indicators and other covariates on the odds of dissolution and marriage are shown in Tables 7 and 8 respectively. Estimates from the models of the hazard of a cohabiting conception are shown in Table 9. We begin by discussing the effects of prior outcomes of the fertility process on transitions from cohabitation, shown in the upper panels of Tables 7 and 8.

Current pregnancy status has a strong effect on the odds of marriage, particularly among the 1958 cohort; relative to women who are not pregnant, those in the first six months of pregnancy have an increased probability of marriage. Further, cohabitors born in 1970 experience a lower risk of dissolution in the first and third trimester of pregnancy. In contrast, pregnant members of the 1958 cohort are no more or less likely to separate than non-pregnant women. While the lack of significance of the correlation between the cohabitation and fertility random effects in the multiprocess model implies that we can rule out selection due to individual unobservables that are constant across partnerships as an explanation for the strong association between pregnancy and marriage, it is possible that selection is acting at the level of the individual partnership. Rather than a causal effect of pregnancy on marriage, the observed positive association may be due to women in stable cohabiting partnerships (presumably with a high chance of being converted to marriage) choosing to conceive with their partner before marriage (Ermisch and Francesconi 2000a; Wu and Balakrishnan 1995).

In common with previous research (Berrington 2001; Blossfeld et al. 1993; Manning and Smock 1995) we find that the marriage rate, while high during pregnancy, declines once the child is born. For both cohorts, having at least one preschool age child with their current cohabiting partner is negatively associated with the odds of marriage, particularly among the more recent cohort. We also find that, for the 1970 cohort only, having a child with their partner reduces their risk of separation, particularly when the child is of school age. Taken together, these results suggest two distinct patterns of childbearing behaviour among women of the more recent cohort who conceive during cohabitation: those who take the more traditional route, favoured by the earlier cohort, of marriage before the birth, and those who give birth during cohabitation and continue to cohabit with a lower risk of separation than childless couples. The second group of women may view cohabitation more as an alternative to marriage and a suitable setting in which to raise children. Among the 1958 cohort, the lack of any significant effect on the risk of separation of the presence of children fathered by the current partner may be due to selection of those in more stable partnerships, with a low separation risk, into marriage before the birth.

Turning to the effects of the presence of children from a previous relationship, we find that relative to women who do not have school age children from a previous co-residential partnership, members of the 1958 cohort with older children by another partner have reduced odds of marriage but are no more or less likely to separate. In contrast, among the 1970 cohort, the presence of school age children from a previous partnership has no effect on the likelihood of marriage, but reduces the chances of separation. For the more recent cohort, having children from a non co-resident relationship is also associated with a lower chance of marriage.

Effects of partnership history and background characteristics. Finally, we discuss the effects on the hazards of dissolution and marriage of the age at the start of the current cohabitation, partnership history, and background characteristics. For both cohorts, there is evidence that women who formed a partnership in their twenties have a lower risk of dissolution than those who started to cohabit in their teens. Further, for the 1970 cohort only, cohabitations formed at an older age are more likely to be converted to marriage. Cohabiting women who have been previously married have reduced odds of remarriage, with the strongest effect observed among the 1958 cohort; this effect may reflect a reluctance among the previously married to marry again, or a delay in getting released from the previous legal marriage. For the 1958 cohort only, previous cohabitation is associated with a lower marriage rate. There is weak evidence that previous cohabitors in the 1970 cohort are less likely to separate than those in their first cohabitation.

The relationship between years of education and the hazard of dissolution is non-monotonic, and somewhat difficult to interpret. There is some suggestion, particularly for the more recent cohort, that those with some post-16 education have an increased risk of separation, possibly because these women are more likely to be financially independent due to better employment prospects. There is no effect of education on the marriage rate, for either cohort. Paternal social class has no effect on the marriage rate for either cohort, but there is evidence that cohabitors from more socio-economically advantaged family backgrounds experience a higher rate of dissolution. Among the 1970 cohort, those who experienced family disruption are less likely to marry and more likely to separate; therefore the cohabitations of these women tend to be of shorter duration than those of women who lived with both biological parents throughout childhood.

Predictors of conceptions within cohabitation. The effects on the chance of having a conception during cohabitation of the presence and characteristics of children and other covariates are shown in Table 9. Surprisingly, women in the 1958 cohort who already have children, either from their current or a previous partnership, are no more or less likely to have a cohabiting conception than are childless women. In contrast, having a young child increases the odds of having another in the 1970 cohort. Women who began a cohabiting relationship while in their teens are more likely to get pregnant during that partnership than those who started to live with their partner at a later age, and this effect has strengthened. In the 1970 cohort, the previously married have an increased chance of conceiving during cohabitation. There is also evidence in both cohorts of a positive effect of previous cohabitation. These partnership history effects may be due to unobserved differences between women who have had more than one co-residential partnership and those who are still with their first partner. The former group may have more liberal views which may lead them to be more relaxed about getting pregnant before marriage.

There is a strong monotonic, negative effect of education on the odds of a cohabiting conception in both cohorts, but the magnitude and gradient of this effect has changed. It is stronger for the earlier cohort, in which the effect of increasing years of education is almost linear. The relationship in the 1970 cohort takes the form of a step function with the 'step' occurring at six or more years. In both cohorts, women whose father was from social classes I or II are less likely to conceive during cohabitation than those from less advantaged backgrounds, although this effect is weaker for the later cohort. This may be indirect evidence of the earning power or marriageability of the partners of these women, which one would expect to be greater for the women from more favourable backgrounds and with more education themselves. Based on this interpretation, couples intending to marry wait until they have done so to conceive, but are somewhat less inclined to wait in the second cohort.

As in the unadjusted data shown in Table 5, there is little difference between the cohorts in the average propensity to conceive within cohabiting partnerships. The rise in births to unmarried couples is largely accounted for by cohort differences in the propensity to marry, which is not completely accounted for by the variables in Table 8. Presumably social attitudes and economic changes beyond the rise in women's education helped to delay, if not reduce, marriage for the later cohort.

Discussion

This study extends existing research on the link between fertility and the odds that a cohabiting partnership is dissolved or converted to marriage in several ways. First, we provide a detailed comparison of this relationship for two cohorts of British women as they pass through early adulthood. Second, rather than focusing only on the outcome of the first cohabiting partnership, we analyse all cohabitations experienced before age 30 using multilevel models with controls for partnership history. Finally, we model jointly transitions from cohabitation and conceptions within cohabitation. Thus we allow for the possibility that current pregnancy status and the presence of children fathered by a cohabiting partner may be endogenous.

Although we conclude that the effects of being pregnant and having children on the outcomes of cohabiting unions cannot be explained by selection on unobserved, time-invariant characteristics of women, the multiprocess models nevertheless provide some interesting insights into the relationship between fertility and partnership decisions among cohabiting couples. In particular, there is weak evidence that women in the earlier cohort who have an above average propensity to marry a cohabiting partner tend also to have an above average propensity to conceive during cohabitation, presumably in anticipation of marriage. When cohabitations up to age 42 are considered, this positive residual correlation attains significance at the 5 per cent level (Steele et al. 2005). In contrast, there is some suggestion that women in the 1970 cohort who are favourably inclined towards marriage tend to wait until marriage before having children. It will be interesting to see whether this effect becomes more pronounced as the cohort are followed into their thirties. We also find, for both cohorts, some evidence of a positive residual correlation between the odds of dissolution and marriage, with a stronger correlation observed in the recent cohort. This result may reflect a contrast between short and long-term cohabitation. We might expect this effect to strengthen as couples who view cohabitation as a long-term commitment make up an increasing proportion of cohabitees.

After allowing for the effects of the fertility and background variables the models show, in their intercept terms, little difference between cohorts in their propensity to separate from cohabiting unions, but a marked decrease in their propensity to marry. This fall in the proportion 'legalising' the relationship is especially strong in the first two trimesters of pregnancy. There is also a significant difference in the chances of a cohabitation dissolving during the first trimester of pregnancy. In the later cohort, but not the first, pregnancy consolidates the union rather than having no effect on the chances of it splitting up. This illustrates that the growing propensity for childbearing in cohabiting unions is due to both a drop in the chances of cohabiting parents splitting up and to their proceeding to marriage. This means more children growing up with parents who are not legally married. On past evidence, these children are at higher risk of experiencing parental break-up, although our evidence suggests to some extent the fragility of cohabiting partnerships is increasingly cemented by the presence of children. To this extent the assumption that married and cohabiting partnerships do not need to be distinguished (e.g. Aasave et al 2004) is increasingly warranted. Even if the chances of parental break-up are converging between married and cohabiting families, there is still a policy concern about the rise of childbearing in non-marital unions in that the legal arrangements for protecting children and their co-resident parent are still better defined for marriages than broken cohabitations, especially given the practical difficulties of implementing the Child Support Agency.

We have looked at some socio-economic influences on the outcomes of cohabitation, and found remarkably little relationship within cohorts of the woman's education or social background on the chances of a union being dissolved or converted into a marriage. There are stronger social variations in the chances of conception in a union, favouring the less educated. Although we have not incorporated a model of the influences on entry to cohabitation in this study, we have confirmed other research that the social composition of people entering cohabitation up to age 29 has changed for these two cohorts in the direction predicted by Ermisch's (2005) 'social contagion' model, from the elite to the masses. However, we do find evidence compatible with some couples treating permanent cohabitation as a first best alternative to marriage. These data can only suggest what may be expected of later cohorts, or of these cohorts themselves in new conditions of the twenty-first century. The less advantaged may find the labour market more conducive to settling down, or the more advantaged may find the increasing climate of uncertainty and the liberalisation of values leads to a new model of long-term commitment. Children may find it a blessing or a curse to have cohabiting parents, or it may not matter….

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Table 1. Descriptive statistics for the explanatory variables included in the event history models

Variables	**1958 cohort**		**1970 cohort**	
Prior fertility outcomes [a]	Percentage who have *ever* been in a given fertility state during cohabitation (base = all cohabiting women)			
Pregnant	29.0		31.9	
Preschool child(ren) with current partner	19.2		28.0	
Older child(ren) with current partner	4.0		7.5	
Preschool child(ren) with previous partner	12.5		6.8	
Older child(ren) with previous partner	14.9		7.5	
Child(ren) with non co-resident partner	3.7		4.1	
Characteristics of current/previous partnerships	Percentage of cohabitation episodes			
Age at start of partnership				
16-19	20.7		22.0	
20-24	45.9		48.0	
25+	33.4		30.1	
Previously married	23.6		5.8	
Previously cohabited	19.3		18.0	
Background characteristics	Percentage of women			
	Cohabiting sample[c]	All women	Cohabiting sample[c]	All women
Post-16 years of education[b]				
0	50.6	52.2	38.3	32.3
1	16.6	16.4	18.5	18.7
2	11.0	12.4	16.9	17.2
3-5	12.2	11.6	16.8	19.3
6+	9.6	7.4	9.5	12.5
Paternal social class				
I-II	19.0	18.0	14.5	18.1
III	55.9	58.6	59.9	57.7
IV-V	21.7	21.2	19.9	18.7
Unknown	3.4	2.2	5.7	5.5
Family disruption before age 16	15.0	10.0	24.6	19.2
No. cohabitation episodes	2650	-	4833	-
No. women	2140	5800	3962	5790
No. woman months of cohabitation	74485	-	171052	-

Notes:

[a]Fertility indicators are treated as time-varying in the analysis.
[b]Number of post-16 years of education is treated as time-varying in the models. Here, the distribution is of educational status at age 29.
[c]Weighted by number of months of cohabitation by age 29.

Table 2. Partnership status of female cohort members at age 29 (per cent)

	1958 cohort	1970 cohort
Unpartnered		
Previously partnered	23	16
Never partnered	9	13
Married		
First partnership	40	12
1st marriage, previously cohabited with current partner only	13	29
1st marriage, previously cohabited with previous partner only	1	1
1st marriage, previously cohabited with both current and previous partners	1	4
2nd + marriage, previously cohabited	4	2
2nd + marriage, never cohabited	1	0
Cohabiting		
First partnership	4	15
Previously cohabited, never married	1	5
Previously married, never cohabited	2	1
Previously cohabited and married	1	2
Total	100	100
n	5720	5613

Table 3. Number of cohabitation episodes per woman between ages 16 and 30, among those who have cohabited before age 30

	1958 cohort		1970 cohort	
No. cohabitations	No. women	%	No. women	%
1	1708	79.8	3188	80.5
2	364	17.0	687	17.3
3+	68	3.2	87	2.2
Total	2140	100.0	3962	100.0

Table 4. Cumulative percentage of cohabitation episodes ending in separation or marriage by partnership duration and partnership history

	Within 2 years		Within 4 years			
	Separate	Marry	Separate	Marry	No. episodes	% episodes
1958 cohort						
First partnership	12.9	42.5	20.4	58.1	1739	65.6
Previously cohabited, never married	17.4	36.9	28.3	49.6	287	10.8
Previously married, never cohabited	6.4	34.1	14.1	51.8	401	15.1
Previously cohabited and married	11.0	25.7	19.3	46.1	223	8.4
1970 cohort						
First partnership	14.9	24.5	24.7	40.8	3876	80.2
Previously cohabited, never married	15.0	23.0	28.3	37.9	677	14.0
Previously married, never cohabited	18.8	21.9	24.6	44.6	86	1.8
Previously cohabited and married	14.5	26.1	21.2	39.1	194	4.0

Note: The above calculations are based on multiple decrement lifetables

Table 5. Number of conceptions (leading to live births) and births during cohabitation per woman between the ages of 16 and 30, among those who have cohabited before age 30

	Conceptions				Births			
	1958		1970		1958		1970	
No. conceptions/births	n*	%	n	%	n	%	n	%
0	1573	73.5	2854	72.0	1766	82.5	2975	75.1
1	441	20.6	779	19.7	273	12.8	652	16.5
2	102	4.8	249	6.3	79	3.7	256	6.5
3+	24	1.1	80	2.0	22	1.0	79	2.0
Total	2140	100.0	3964	100.0	2140	100.0	3962	100.0

*Number of women.

Table 6. Estimated random effect covariance matrices from the multiprocess models

	Separation	Marriage	Conception
1958 cohort			
Separation	0.991 (0.499, 1.569)		
Marriage	0.194 (-0.045, 0.437) **0.323**	0.366 (0.205, 0.604)	
Conception	0.042 (-0.100, 0.202) **0.100**	0.041 (-0.040, 0.141) **0.162**	0.162 (0.099, 0.244)
1970 cohort			
Separation	0.624 (0.348, 0.971)		
Marriage	0.233 (-0.012, 0.508) **0.364**	0.623 (0.361, 0.984)	
Conception	-0.067 (-0.159, 0.020) **-0.241**	-0.065 (-0.158, 0.011) **-0.232**	0.121 (0.080, 0.170)

Note: The values in each cell are the point estimate (the mean of the MCMC samples) and the corresponding 95 per cent interval estimate (the 2.5 per cent and 97.5 per cent 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 50000 MCMC samples, with a burn-in of 5000.

Table 7. Estimated coefficients (and standard errors) from multilevel competing risks model for outcomes of cohabitation: effects on log-hazard of **dissolution versus continuing cohabitation**

	1958 cohort		1970 cohort	
Variables	**Coeff.**	**(SE)**	**Coeff.**	**(SE)**
Prior fertility outcomes[a]				
Current pregnancy status (ref=not pregnant)				
1-3 months pregnant	-0.065	(0.284)	-0.598	(0.240)
4-6 months pregnant	-0.427	(0.374)	-0.041	(0.191)
7-9 months pregnant	-0.350	(0.377)	-0.672	(0.258)
Preschool child(ren) with current partner	-0.084	(0.156)	-0.217	(0.084)
Older child(ren) with current partner	-0.091	(0.312)	-0.424	(0.179)
Preschool child(ren) with previous partner	-0.450	(0.302)	-0.127	(0.252)
Older child(ren) with previous partner	-0.531	(0.252)	0.304	(0.182)
Child(ren) with non co-resident partner	-0.002	(0.313)	-0.013	(0.174)
Characteristics of current/previous partnerships				
Age at start of partnership (ref=16-19)				
20-24	-0.124	(0.128)	-0.219	(0.078)
25+	-0.451	(0.174)	-0.107	(0.105)
Current partnership duration[b]	0.036	(0.007)	0.021	(0.005)
Previously married	-0.138	(0.170)	0.175	(0.178)
Previously cohabited	-0.064	(0.166)	-0.305	(0.127)
Background characteristics				
Post-16 years of education[a] (ref=0)				
1	0.292	(0.147)	0.311	(0.089)
2	0.175	(0.178)	0.162	(0.098)
3-5	0.215	(0.169)	0.398	(0.094)
6+	0.169	(0.207)	0.072	(0.131)
Paternal social class[c] (ref=III)				
I-II	0.325	(0.138)	0.096	(0.092)
IV-V	-0.189	(0.144)	-0.343	(0.092)
Unknown	-0.115	(0.304)	0.111	(0.134)
Family disruption before age 16	0.090	(0.150)	0.159	(0.075)
Constant	-5.598	(0.202)	-5.315	(0.129)

[a]Time-varying covariate.
[b]Duration is measured in one-month intervals.
[c]Social class refers to the current or most recent occupation of the father (or mother's husband) at the respondent's birth. The codes are I: Professional, II: Managerial and Technical Occupations, III: Skilled occupations (manual or non-manual), IV: Partly skilled occupations, V: Unskilled occupations. Unknown parental social class includes cases with no resident father at birth.

Table 8. Estimated coefficients (and standard errors) from multilevel competing risks model for outcomes of cohabitation: effects on log-hazard of **marriage versus continuing cohabitation**

	1958 cohort		1970 cohort	
Variables	**Coeff.**	**(SE)**	**Coeff.**	**(SE)**
Prior fertility outcomes[a]				
Current pregnancy status (ref=not pregnant)				
1-3 months pregnant	1.297	(0.106)	0.822	(0.109)
4-6 months pregnant	1.344	(0.115)	0.911	(0.108)
7-9 months pregnant	0.067	(0.211)	-0.225	(0.182)
Preschool child(ren) with current partner	-0.185	(0.098)	-0.314	(0.072)
Older child(ren) with current partner	-0.361	(0.264)	-0.199	(0.150)
Preschool child(ren) with previous partner	-0.081	(0.152)	0.001	(0.213)
Older child(ren) with previous partner	0.100	(0.130)	-0.568	(0.198)
Child(ren) with non co-resident partner	-0.087	(0.178)	-0.281	(0.151)
Characteristics of current/previous partnerships				
Age at start of partnership (ref=16-19)				
20-24	0.015	(0.077)	0.237	(0.066)
25+	0.048	(0.097)	0.397	(0.087)
Current partnership duration[b]	0.021	(0.005)	0.031	(0.004)
Previously married	-0.491	(0.105)	-0.347	(0.165)
Previously cohabited	-0.225	(0.100)	-0.107	(0.113)
Background characteristics				
Post-16 years of education[a] (ref=0)				
1	0.066	(0.086)	0.072	(0.075)
2	0.103	(0.103)	0.033	(0.077)
3-5	-0.010	(0.104)	-0.105	(0.081)
6+	-0.013	(0.127)	-0.110	(0.103)
Paternal social class[c] (ref=III)				
I-II	-0.057	(0.089)	0.104	(0.077)
IV-V	-0.078	(0.080)	-0.018	(0.071)
Unknown	-0.390	(0.198)	-0.097	(0.128)
Family disruption before age 16	0.032	(0.090)	-0.310	(0.070)
Constant	-4.036	(0.102)	-5.057	(0.121)

[a]Time-varying covariate.
[b]Duration is measured in one-month intervals.
[c]Social class refers to the current or most recent occupation of the father (or mother's husband) at the respondent's birth. The codes are I: Professional, II: Managerial and Technical Occupations, III: Skilled occupations (manual or non-manual), IV: Partly skilled occupations, V: Unskilled occupations. Unknown parental social class includes cases with no resident father at birth.

Table 9. Estimated coefficients (and standard errors) from multilevel competing risks model for outcomes of cohabitation: effects on log-hazard of **conception during cohabitation (among cohabitors)**

	1958 cohort		1970 cohort	
Variables	**Coeff.**	**(SE)**	**Coeff.**	**(SE)**
Prior fertility outcomes[a]				
Preschool child(ren) with current partner	0.139	(0.106)	0.250	(0.064)
Older child(ren) with current partner	-0.093	(0.253)	-0.145	(0.141)
Preschool child(ren) with previous partner	0.136	(0.148)	0.594	(0.148)
Older child(ren) with previous partner	-0.081	(0.138)	0.109	(0.145)
Child(ren) with non co-resident partner	0.133	(0.196)	0.299	(0.122)
Characteristics of current/previous partnerships				
Age at start of partnership (ref=16-19)				
20-24	-0.223	(0.098)	-0.393	(0.064)
25+	-0.214	(0.125)	-0.365	(0.091)
Current partnership duration[b]	-0.010	(0.004)	-0.005	(0.003)
Previously married	0.170	(0.112)	0.394	(0.121)
Previously cohabited	0.232	(0.111)	0.170	(0.095)
Background characteristics				
Post-16 years of education[a] (ref=0)				
1	-0.299	(0.111)	-0.183	(0.073)
2	-0.703	(0.159)	-0.223	(0.076)
3-5	-1.233	(0.193)	-0.457	(0.091)
6+	-1.478	(0.285)	-1.050	(0.167)
Paternal social class[c] (ref=III)				
I-II	-0.344	(0.145)	-0.175	(0.095)
IV-V	0.089	(0.093)	0.191	(0.065)
Unknown	-0.079	(0.225)	0.112	(0.117)
Family disruption before age 16	0.228	(0.104)	-0.045	(0.064)
Constant	-4.193	(0.104)	-4.357	(0.078)

[a]Time-varying covariate.
[b]Duration is measured in one-month intervals.
[c]Social class refers to the current or most recent occupation of the father (or mother's husband) at the respondent's birth. The codes are I: Professional, II: Managerial and Technical Occupations, III: Skilled occupations (manual or non-manual), IV: Partly skilled occupations, V: Unskilled occupations. Unknown parental social class includes cases with no resident father at birth.

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