Cohabitation and Childbearing: Changing Compatibility between Two Cohorts of Young British Women

Fiona Steele
University of Bristol

Heather Joshi, Constantinos Kallis
Institute of Education, University of London

Harvey Goldstein
University of Bristol

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Abstract

We investigate the role of parenthood on the outcomes of nonmarital unions for two cohorts of British women between the ages of 16 and 29. We compare the effect of conceptions leading to births and the presence and characteristics of children on the odds that a cohabitation is dissolved, or that it is converted to marriage, for women born in 1958 and 1970. A multilevel multiprocess competing risks model allows for multiple cohabitations per woman and endogeneity of fertility status. We find that cohabiting couples' response to impending parenthood and the presence of children has changed over time. In particular, the proportion of cohabiting couples who marry before a birth has decreased and, in the 1970 cohort only, the risk of dissolution declines during pregnancy. There is also evidence that the presence of a child cements a cohabiting union for women from the 1970, but not the earlier, cohort.

Key words: Cohabitation, marriage, partnership dissolution, fertility, competing risks, multilevel modelling, multiprocess modelling, simultaneous equation modelling, selection effects

Introduction

In Britain, as in much of western Europe, there has been a dramatic rise in unmarried cohabitation in recent decades (Ermisch and Francesconi 2000; Murphy 2000; Haskey 2001; Kiernan 2001; Ermisch 2005). While cohabitation used to occur mainly after marital breakdown, by the early 1990s around three-quarters of British men and women cohabited in their first partnership (Ermisch and Francesconi 2000). 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 than marriages, Murphy suggests that this increase in duration may signal a change towards cohabitation being viewed more as an alternative to marriage. The increased prevalence of cohabitations has also seen a rise in childbearing within them. Cohabiting conceptions form an increasing proportion of births conceived outside marriage. In the 1970s, many of these precipitated marriages of pregnant brides. 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). The marriage rate among women who had their first child during cohabitation was lower in Britain than in any other country in western Europe in the early 1990s (Kiernan 2001).

It is not clear what has been driving the increases in cohabitation and extra-marital childbearing. Changes in the legal framework, official statistics, attitudes and practice have been running alongside each other. The premium on marriage in UK income tax was effectively abolished in 2000, but there is still a reward for legal marriage in the survivor benefits of the state retirement pension. The legal status of illegitimacy was abolished in

1987, although unmarried fathers' parental rights and responsibilities continued to differ from married fathers' (Kiernan et al. 1998). The use of the term 'illegitimacy' in common parlance has taken somewhat longer to die out, along with disapproval of extra-marital relationships. Scott (1999) shows attitudes to pre-marital sex liberalizing within and across cohorts born since 1900. 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', led by innovators among students and graduates in the 1970s, but spreading across all social groups over time. The original social gradient reversed as the labour market for least skilled men deteriorated (McRae 1999). 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. Another factor increasing cohabitation could be the falling credibility of marriage as a life-long contract as divorce rose.

The principal aim of this paper is to examine the role of parenthood in prolonging, converting to marriage, or ending non-marital partnerships for two cohorts of young British women born in 1958 and 1970. We explore how cohabiting couples' odds of marriage and dissolution depend on their changing fertility status, establishing inter-cohort differences in these relationships. Our secondary aim is to examine the influences on childbearing in cohabitation, both within and across cohorts. Our focus is on the years running up to age 30, the stage of transition into family formation which increasingly involves spells of cohabitation, disturbing the traditional sequence from single to married to parent. While more women have postponed childbearing until after 30, those who have not are increasingly likely to have children before being married.

Previous British studies have found that an impending birth is associated with increased odds of marriage (Berrington 2001), but this association has weakened over time

(Berrington 2003). We investigate whether the higher rate of marriage during pregnancies (which are not terminated) may be partly explained by a selection mechanism, whereby women who conceive while cohabiting (and carry the pregnancy to term) differ from those who do not bear children conceived in cohabitations on characteristics that are also associated with the likelihood of getting married. We also explore the extent to which differences between cohorts in this relationship may be explained by the following factors: (1) changes in the characteristics of those who cohabit, (2) changes in the nature of any selection effect, and (3) a real shift in the attitudes and behaviour of cohabitors with given characteristics. Each of these sources of cohort change could be aspects of the Second Demographic Transition (Van de Kaa 2003), characterised among other things by a loosening of the monopoly of marriage over childbearing during the last decades of the Twentieth Century.

The sharp increase in the marriage rate observed among cohabiting expectant parents is not sustained after a birth. Studies using data from Britain and northern Europe reveal a postnatal decline in the odds of marriage and increase in the dissolution risk (Blossfeld et al. 1993; Ermisch and Francesconi 2000). As noted by Ermisch and Francesconi, these effects may be due to a tendency among cohabiting couples who are favourably disposed towards marriage, and who find each other acceptable partners, to marry before having children. We compare across cohorts the relationship between the outcomes of cohabitation and the age and parentage of existing children, and consider whether any inter-cohort differences might be explained by a change in any selection effects, or changes in the characteristics of those who cohabit and have children while cohabiting.

Links between Childbearing and the Outcomes of Cohabitation: Hypotheses and Empirical Evidence

The effect of an impending birth

There is overwhelming evidence that among cohabiting couples an effective conception (i.e. one which subsequently results in a live birth) is associated with increased odds of marriage and a decreased risk of separation (e.g. Blossfeld et al. 1993; Manning and Smock 1995; Berrington 2001, 2003; Steele et al. 2005a). For some couples impending parenthood merely hastens marriage, but for others an accidental pregnancy leads to a marriage which may not otherwise have taken place. Blossfeld et al. (1993) provide a summary of theories that might explain why the prospect of a birth hastens or forces marriage among cohabitors. One theory is that a couple's actions are influenced by social norms which favour the traditional route of marriage before childbearing. However, Blossfeld et al. contend that this normative model is likely to be operating in conjunction with rational choice whereby couples act in a way that satisfies their own preferences. Couples facing parenthood may choose to marry because marriage offers a longer-term commitment and therefore a more stable setting in which to have children. There may also be legal and economic reasons behind a couple's decision to marry. Men may prefer to have children within marriage because of differences in the parental rights of married and unmarried fathers, while women may choose to marry for greater financial security in the event of break-up.

Under both the normative and rational choice model, one would expect the association between pregnancy and marriage to weaken over time. As non-marital childbearing becomes more common and more widely accepted, traditional norms should exert less influence on behaviour. At the same time, increasing divorce together with convergence of the rights of married and unmarried fathers and the increased labour participation of mothers may remove some of the advantages of marriage for cohabiting men and women. For similar reasons, we expect more potentially stable couples to remain unmarried after a conception, and hence a lowering of cohabitors' risk of separation approaching (or after) parenthood in the later cohort.

The effects of the presence and characteristics of children

Social and economic theories of marital dissolution suggest that having children together raises the costs of separation and increases the gains from marriage, leading to greater marital stability among couples with children (e.g. Koo and Janowitz 1983; Lillard and Waite 1993). Wu (1995) argues that the same theories should apply to cohabitation. The direction of the relationship between the presence of children and the probability that a cohabiting couples marries is more difficult to anticipate. On the one hand, couples with children may marry to signal a longer-term commitment. On the other hand, couples who chose not to marry before a birth may view cohabitation as an alternative to marriage, and may therefore be less likely to marry or split up than childless couples. Although, as suggested above, the presence of children may increasingly be associated with union stability, the increase in single parenthood may reduce the stabilising effect of children.

The direction of the effects of the presence of children from a previous relationship on the odds of dissolution and marriage are also difficult to predict. To the extent that both biological and step children constitute a shared interest, the presence of either should reduce the risk of separation and increase the odds of marriage. Furthermore, women with children who enter a new partnership have already experienced the breakdown of the relationship with the child's father and may prefer a more formal union, leading to increased odds of marriage. Women with children may also be more selective in their choice of future partner, resulting in a lower risk of dissolution. If, however, the prospect of stepchildren is an impediment in the "marriage market" or a source of conflict in a partnership, their presence could decrease the probability of marriage or increase the risk of dissolution. As the living arrangements of families become increasingly complex and diverse (Ferri and Smith 1998), we anticipate that an inter-cohort comparison of these relationships will show weaker effects for the 1970 cohort.

The previous research on the relationship between the presence of children, partnership dissolution and entry into marriage in Britain and elsewhere lacks consensus. This may reflect genuine variation between countries, but another possible explanation is differences in the fertility indicators used. Some authors consider only the number of children, while others take into account their characteristics, including their age, sex and whether they were born before or during the current cohabitation. 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. Böheim and Ermisch 2001; Aassve et al. 2004) and some studies of entry into marriage do not consider transitions from cohabitation separately from marriage from an unpartnered state (e.g. Brien et al. 1999, Upchurch et al. 2002). Studies also differ in the cohort composition of their samples and their treatment of cohort effects. Some studies reviewed here are based on a single cohort. 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.

Ermisch and Francesconi (2000), 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. 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. (2005a) 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 different 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.

The evidence to date therefore suggests some instability in the relationship between the presence of children and the odds that a cohabiting couple marries or separates, with variations across time and place. In this paper we seek to contribute to the literature by investigating temporal changes in the link between childbearing and cohabitation using data for two cohorts, controlling for an identical set of background characteristics for each.

Selection Effects

Estimates of the effects of current fertility status on cohabitation outcomes may be subject to selection bias. Selection mechanisms may at least partly explain observed within-cohort associations and, if the nature of selection bias changes over time, cross-cohort differences. A selection bias will arise if women who get pregnant during cohabitation and subsequently have a birth differ from those who do not on observed or unobserved characteristics that are also associated with their chance of marriage or separation. For example, women inclined towards solo living may be less likely to marry, more likely to end a partnership and less likely to have a child. Alternatively, women who hold more 'traditional' family values may be more likely to marry, be less inclined to have a child during cohabitation and more likely to marry before the birth in the event of a cohabiting conception. Selection on unobserved individual attributes that are fixed over the observation period, e.g. attitudes towards marriage and family, can be said to operate at the individual level. Selection may also operate at the level of the partnership if there are unobserved characteristics of couples, or an interaction between the characteristics of each partner, which affect both decisions about the future of the partnership and having children together. For instance, couples in a stable relationship may be more likely to marry and less likely to separate, and more likely to have children together.

The potential for selection bias has been recognised by several authors, most commonly as an explanation for the apparent negative effect of having children together on a cohabiting couple's odds of marriage (Wu and Balkrishnan 1995; Blossfeld et al. 1999; Ermisch and Francesconi 2000). They suggest that 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, i.e. a positive effect on the persistence of cohabitation. 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.

If selection effects are constant over time, cross-cohort comparisons of the relationship between childbearing and the outcomes of cohabitation will be unaffected by selection bias. However, as cohabitation becomes more widespread and the link between marriage and childbearing loosens, selection effects of the type described above should grow weaker over time. For this reason, it is important to allow for selection in the estimation of both within-cohort effects of current fertility status and cross-cohort differences in these effects.

Methods

Multilevel multiprocess modelling

As discussed above, an important 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 may be subject to shared, or correlated, unobserved influences. Failure to account for selection on unobserved characteristics will lead to biased estimates of the effects of pregnancy and the presence and characteristics of children on the outcomes of cohabitation. One way to allow for selection effects is to use a multiprocess model in which the endogenous explanatory variable(s), i.e. fertility outcomes, are modelled jointly with transitions from cohabitation. Such models allow explicitly for selection on unobservables by introducing a correlation between the residual components of each process in the system. 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 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. 2004). Only one study to date has used multiprocess modelling to examine the link between childbearing and the outcomes of cohabiting unions. Steele et al. (2005a) 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, and allowing for competing risks in the outcome of cohabitation). In this paper, we adopt the methodological approach of Steele et al. (2005a), but focus on the outcomes of cohabiting unions and fertility during cohabitation, and on the comparison of two cohorts.

Most studies of cohabitation outcomes focus on the first, usually premarital, partnership. In this paper we consider *all* episodes of cohabitation that begin before a woman's 30th birthday, controlling for her partnership history prior to the current cohabitation. The possibility that respondents may live with more than one partner between the ages of 16 and 29, and may have multiple conceptions (leading to births) within those cohabitations, implies 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.

Our multilevel multiprocess model is a system of three simultaneous equations, one for each outcome of cohabitation and a further equation for effective conceptions within cohabitation. Each equation defines a multilevel discrete-time event history model. The model is a special case of the more general model proposed by Steele et al. (2005a) for analysing transitions from both marriage and cohabitation jointly with fertility within either form of partnership. We estimate the model using Monte Carlo Markov chain (MCMC) methods, as implemented in *MLwiN* (Rasbash et al. 2004). Further details of estimation and model identification can be found in Steele et al. (2005a).

Separate models are estimated for each cohort. Fitting a single model to a pooled dataset, with cohort dummies and their interactions with explanatory variables, is not practically feasible at present, given the already large size of the discrete-time datasets which contain an observation for each month of cohabitation.

Competing risks model for the outcomes of cohabitation.

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:

$$\log\left[\frac{h^{C(r)}(t)}{h^{C(0)}(t)}\right] = \boldsymbol{\alpha}_0^{C(r)}\mathbf{D}^{C(r)}(t) + \boldsymbol{\alpha}_1^{C(r)}\mathbf{F}(t) + \boldsymbol{\alpha}_2^{C(r)}\mathbf{X}^{C(r)}(t) + u^{C(r)}, \quad r = 1,2 \qquad (1)$$

where $\boldsymbol{\alpha}_0^{C(r)}\mathbf{D}^{C(r)}(t)$ is a function of cohabitation duration at month *t*, $\mathbf{F}(t)$ is a vector of fertility indicators of current pregnancy status and the presence and age of children from the current or a previous partnership, $\mathbf{X}^{C(r)}(t)$ 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. The random effects represent time-invariant unobserved characteristics that affect the odds of marriage or dissolution for *all* of a woman's cohabitations. The random effect variance measures the extent of unobserved heterogeneity between women due to unobserved individual characteristics that are fixed in time.

Model for fertility within cohabitation.

Denote by $h_{ij}^{F}(t)$ the hazard of an effective conception during month *t* in cohabitation episode *i* for individual *j*. We consider only those conceptions that lead to a live birth and 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:

$$\text{logit}\, h^{F}(t) = \boldsymbol{\beta}_0^{F}\mathbf{D}^{F}(t) + \boldsymbol{\beta}_1^{F}\mathbf{F}(t) + \boldsymbol{\beta}_2^{F}\mathbf{X}^{F}(t) + u^{F} \tag{2}$$

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

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. $\mathbf{u} = (u^{C(1)}, u^{C(2)}, u^{F}) \sim N_3(\mathbf{0}, \boldsymbol{\Omega}_u)$. 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 conceiving a live birth during a cohabiting relationship. Non-zero correlations between elements of $\mathbf{u}^C = (u^{C(1)}, u^{C(2)})$ and u^F would suggest that $\mathbf{F}(t)$, the number and age of children from the current or a previous partnership, is endogenous with respect to the outcomes of cohabitation.

The above model allows for selection at the individual level, i.e. unobserved characteristics of women which are constant over time. To allow additionally for selection at the partnership level would require instrumental variables; these would be variables which affect the chances of a cohabiting conception but not the outcomes of cohabitation. Such variables are difficult to find so we do not pursue this further here. See Steele et al. (2005a) for further discussion of model identification.

In the analysis that follows, two specifications of the multilevel event history model were estimated. In the first model, the full multiprocess model, the pairwise correlations between random effects across the three equations are freely estimated. In the second model, the single process model, the residual correlations between each of the cohabitation outcome equations and the fertility equation are constrained to zero, i.e. each element of $\mathbf{u}^C$ is assumed to be uncorrelated with u^F. Note, however, that the correlation between $u^{C(1)}$ and $u^{C(2)}$ is still permitted to be non-zero in this model. Placing a zero constraint on the cross-process correlations is equivalent to fitting separate models for transitions from cohabitation and fertility. Estimates of the coefficients of the fertility indicators $\mathbf{F}(t)$ were compared across these two models to assess the impact of allowing for selection on unobservables.

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 (Bynner et al. 1997; Shepherd 1997). Since birth, contacts been made with 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) and used here up to the 30th birthday. Partnership histories from age 16 were collected from the 1970 cohort at age 29. For our analysis, episodes of (nonmarital) cohabitation were extracted from these histories. One 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 30th 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 if leading to a live birth and its 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 two family background measures: father's social class at the respondent's birth, and the experience of family disruption before age 16. Region of residence at birth and housing tenure at age 16 were also considered, but excluded from the final model due to weak association with the outcomes and, in the latter case, missing data. Descriptive statistics for all explanatory variables included in the final models are shown in Table 1.

[TABLE 1 ABOUT HERE]

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. 2005a). 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. As children usually 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 (see Table 1). 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 those eligible. 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, which would mean biases should cancel out when looking at inter-cohort differences. There are also missing data for the family background variables and years of education. Where possible, missing values were imputed using information collected at earlier or later ages.

Results

Descriptive analysis of inter-cohort differences

Changes in the incidence and duration of cohabitation. The proportion of cohort members who ever cohabited before age 30 almost doubled from 37 to 68 per cent (Table 1). Much of this increase can be attributed to a rise in pre-marital cohabitation. In the 1970 cohort, 73 percent of those in their first marriage at age 30 cohabited with their partner before marriage, compared to only 26 per cent in the 1958 cohort. Among those who cohabited before age 30 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. 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. The main determinant of the increase in the duration of cohabitation is the lower rate of marriage 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. Inter-cohort differences in the rate of dissolution are less dramatic.

Changes in childbearing within cohabitation. 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 effective conceptions during cohabitation. The major change in the fertility behaviour of cohabiting couples is in the relative frequency of effective conceptions that led to births within cohabitation. For the 1958 cohort, 26 per cent of women experienced at least one conception while cohabiting, and 17 per cent gave birth during cohabitation. The same proportions for the 1970 cohort were 28 per cent and 25 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.

Changes in the characteristics of those who cohabit. 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 nearly 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) in NCDS compared with 18.7 per cent in BCS70. We do not explicitly allow for macro-economic change in our models, 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 analysed 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 BCS70, all social groups participated in the incoming wave of cohabitation, but it was the less educated and those from less auspicious backgrounds who rode its crest. Thirty-eight per cent of the woman-months of cohabitation in the 1970 cohort were contributed by 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. The less qualified tended to start earlier. 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.

Multilevel event history analysis

As noted above, we considered two specifications of the multilevel event history model – the multiprocess and single process models – which differ according to assumptions made about the random effects correlations. In the multiprocess model, estimates of the pairwise covariances between u^F and each element of $(u^{C(1)}, u^{C(2)})$ are of particular interest since they provide a test of endogeneity of fertility status with respect to cohabitation outcomes. We find, for both cohorts, that neither covariance is significantly different from zero at the 5 per cent level (see Table A1 in the Appendix for the estimated covariance matrices). This implies that, conditional on the covariates included in the model, there is no selection on woman-specific unobservables. Because the estimated coefficients of the fertility indicators $\mathbf{F}(t)$ are very similar for the two models, we present estimates only from the multiprocess model. (A discussion of the covariance estimates from this model is given in Steele et al. (2005b).)

Effects of fertility indicators. We begin by discussing the effects of prior outcomes of the fertility process on transitions from cohabitation, shown in the upper panels of Tables 2 and 3. As expected, current pregnancy status has a strong positive effect on the odds of marriage, particularly among the 1958 cohort, in the first two trimesters. Cohabitors born in 1970 experience a lower risk of dissolution in the first and third trimester of pregnancy. In contrast, expectant mothers in the 1958 cohort are no more or less likely to separate than non-expectant 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 (Wu and Balakrishnan 1995; Ermisch and Francesconi 2000).

In common with previous research (Blossfeld et al. 1993; Manning and Smock 1995; Berrington 2001) 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, trodden 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 associated with a lower chance of marriage.

The estimated effects of the partnership history indicators and background characteristics, also shown in Tables 2 and 3, are discussed in Steele et al. (2005b).

[TABLES 2 AND 3 ABOUT HERE]

Predictors of conceptions within cohabitation. Table 4 shows the effects on the chance of an effective conception while cohabiting of existing children and other covariates. Surprisingly, women in the 1958 cohort who already have children, either from their current or a previous partnership, have similar chances of conceiving as the childless. In contrast, having a young child increases the odds of having another in the 1970 cohort. Women who started cohabiting in their teens are more likely to get pregnant during that partnership than those who began living 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.

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 shows the biggest contrast between six or more years and other levels of post-compulsory schooling. 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 noted earlier, 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 3. Presumably social attitudes and economic changes beyond the rise in women's education helped to delay, if not reduce, marriage for the later cohort.

[TABLE 4 ABOUT HERE]

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 allowing for the possibility that current pregnancy status and the presence of children fathered by a cohabiting partner may be endogenous.

We find evidence of important changes in the role of parenthood in cohabitors' chances of separation or marriage. The within-cohort relationships between fertility status and the outcomes of cohabitation cannot be explained by selection on women's observed or unobserved characteristics, nor can temporal changes in these relationships be explained by changes in the nature of selection. The most striking change is a fall in the proportion ‘legalising’ the relationship 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 an effective pregnancy. In the later cohort, but not the first, pregnancy (taken to term) consolidates the union rather than having no effect on the chances of it splitting up. In common with previous studies, we find that the odds of marriage decrease after a birth, and the effect is stronger in the more recent cohort. However, there is evidence that having children together reduces couples’ risk of dissolution in the 1970, but not the 1958, cohort. The effect of the presence of step-children has also changed over time. Having a school-age child from a previous partnership is associated with a reduction in the chance of marriage in the 1970 cohort, and a reduction in the risk of dissolution in the earlier cohort. These findings fit the prediction that the effects of existing children would be complex, but do not display the systematic trend towards a weakening of relationships between parenthood and partnership transitions that we hypothesised.

Taken together, our findings illustrate that the growing propensity for childbearing in cohabiting unions is due to a drop in both the chances of cohabiting parents splitting up and of 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 that the fragility of cohabiting partnerships may be increasingly cemented by the presence of children. Even so, an issue remains that the legal arrangements for protecting children and their co-resident parent are still better for marriages than broken cohabitations. The trend may, however, simplify future research. The assumption that married and cohabiting partnerships do not need to be distinguished (e.g. Aasave et al. 2004) may be increasingly warranted. We find evidence compatible with some couples treating permanent cohabitation as a first best alternative to marriage. This tale of two cohorts can only suggest what may be expected of later cohorts, or of these cohorts themselves as they approach the age to collect their pensions, but the change looks set to continue.

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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[e]	All women	Cohabiting sample[e]	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[c]				
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[d]	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]In case of missing social class, information collected at an older age was imputed if the father figure was identified as the natural or adoptive father on both occasions.
[d]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.
[e]Weighted by number of months of cohabitation by age 29.

Table 2. Estimated coefficients (and standard errors) from multilevel competing risks models 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.

Note: The estimated coefficients and standard errors are respectively the means and standard deviations of the MCMC samples. The results are based on 50000 MCMC samples, with a burn-in of 5000. See Steele et al. (2005a) and Browne (2003) for further details.

Table 3. Estimated coefficients (and standard errors) from multilevel competing risks models 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				
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	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 (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 (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)

Table 4. Estimated coefficients (and standard errors) from multilevel competing risks model for outcomes of cohabitation: effects on log-hazard of **conception during cohabitation (leading to a live birth)**

	1958 cohort		**1970 cohort**	
Variables	**Coeff.**	**(SE)**	**Coeff.**	**(SE)**
Prior fertility outcomes				
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	-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 (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 (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)

Appendix

Table A1. Estimated random effect covariance matrices from the multiprocess models

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

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.