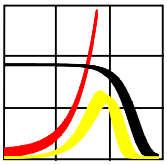

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Cohabitation, marriage, first birth: The interrelationship of family formation events in Spain

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Cohabitation, marriage, first birth: The interrelationship of family formation events in Spain

Short title: Union formation and first birth in Spain

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Abstract

In this paper we investigate (1) the mutual causal relationship between first union formation and first childbirth, and (2) the existence of constant common determinants of these two events. It is argued that (unmeasured) common factors reflect differentials among the population in value orientations and in norms about the sequencing of events. We apply event history techniques to retrospective survey data for Spain, allowing for the correlation between unobserved heterogeneity components belonging to each process. Our findings confirm the strong interrelationship between union formation and first birth. After controlling for these common factors, we find that the risk of conception increases immediately at marriage, and it continues to be high during the following four years. Entry into cohabitation produces much smaller increases in the relative risk. The effect of the conception of the first child on union formation is especially strong during pregnancy, but declines sharply after delivery.

1. Introduction

The fact that childbearing in most cases takes place within unions (traditionally within marital unions and increasingly within consensual unions) has made demographers to devote less attention to this relationship than it actually deserves. More specifically, the causal (versus possibly spurious) nature of such a relationship needs to be explicitly addressed if researchers are interested in policy-related issues. For instance, what happens to the timing of first births if first unions are experienced at later ages? And what are the differences between consensual unions and marital unions? Does the conception of a baby have a causal impact on subsequent union formation? Do value orientations and life plans make the relationship between union formation and first births a spurious relationship? Only a few studies, which we shall review later on, have addressed this issue from such a point of view.

In this paper, we would like to focus our attention on a situation in which first births are increasingly postponed, and in which fertility reaches very low levels. We concentrate on the case of Spain, which has been among the first countries to reach levels of "lowest-low" fertility (Kohler et al., 2001) and is currently among the leading countries in low fertility. Spain is also a country where cohabitation is spreading at a very slow speed. In fact, not much is known about cohabitation in contemporary Spain at all. The general questions we shall address in this paper are the following. Is postponed first birth in Spain a consequence of postponed union formation, net of the possible common factors? Can the emergence of cohabitation reverse the fertility trends or at least prevent the trends from moving towards an even lower fertility?

Entering parenthood and forming first unions are closely linked events both in terms of their timing over the life course and in terms the intentions and life plans of individuals. If a union is viewed as the appropriate setting for bearing children, individuals wanting to have a child may speed-up their union formation, considering this event as part of their family building strategy. Conversely, a pregnancy may precipitate marriage formation for couples that already had plans in that direction. The

lack of independence between the processes of first union formation and first birth is reflected in the modelling strategy that is to be applied. Intensity regression analyses usually include one of these events, for example by focusing on the duration from union formation to first birth, and possibly use the experience of the other event as a time-varying covariate (see i.e., Blossfeld and Mills, 2000).

In this paper, we use a modelling strategy that allows for the presence of constant unmeasured common factors simultaneously influencing the timing of first childbirth and first union formation. If such common factors (whose nature we shall discuss in more detail) exist, then an individual with a high risk of childbearing will also be more likely to form a union early in his or her life course. As a consequence, we expect an increasing selection effect at higher ages, where individuals with lower family formation risks become over-represented. Moreover, the time order of the events may not reflect a causal relationship. A reversed causality between the dependent event and the explanatory variables may be present, for instance, when the anticipation of having a child affects the decision to form a union. In that situation, the estimated parameters in hazard regression will be biased and unlikely to reflect the independent effect of union formation on conception leading to a first birth.

In order to overcome these potential biases, we model simultaneously as dependent events first birth and first union formation. Furthermore, we model union formation in a competing-risk framework with two destinations: cohabitation and marriage. We follow the approach developed by Lillard (1993), that estimated the correlation between the unobserved heterogeneity components belonging to different processes. In this way, we will be able to answer the questions we posed at the beginning of this paper.

In studying these issues, the model employed in the study of childbirth takes into account the impact of the duration since union formation; and when studying union formation, we use the time since the start of a pregnancy. It is substantively important to depict the "shape" of each of these time effects. Furthermore, only when the time dependencies between processes are modelled in a detailed way is it possible to control for the interrelationships between processes with confidence.

Finally, we will pay attention to the effects of several socio-economic variables on the timing of first birth, cohabitation, and marriage. The effect of some of these variables may well differ for each of the processes studied. In addition, their effects may operate through a related process. For instance, a given variable may have an effect on the marriage process, which in turn will influence the likelihood of having the first birth. We use data from the Family and Fertility Survey for Spain, which was conducted in 1995. We restrict the study to include female birth-cohorts only.

The paper is organised as follows. Section 2 provides a short description of some elements of the family formation process in Spain, paying somewhat more attention to the inter-cohort development of cohabitation. Then follows a proposal of several hypotheses concerning the interrelationships between union formation and first birth. Section 4 describes the variables and the model employed. Section 5 presents and discusses the results, and section 6 provides some concluding remarks.

2. The demographic setting

Since the mid-1970s a new pattern of family formation has developed for young adults in Southern European countries. This is characterised by a very low fertility, together with an increasing age at first birth and at first union formation. Few births take place before entering marriage, and marriage usually coincides with leaving the parental home. Post-nuclear family forms, such as living alone and cohabitation, have only had limited increases in Southern Europe. Several authors have interpreted such developments as being closely interrelated, attributing this pattern to cultural peculiarities (Reher, 1998; Micheli, 2000; Dalla Zuanna, 2001), or to the specific pathways of accession to modernity of these countries (Baizán et al. 2000; Flaquer, 2000). They have pointed to the centrality of the family for the economic well-being of individuals in such contexts.

The reduction, with respect to previous birth-cohorts, of first union formation and first birth intensities for the Spanish birth-cohorts born in the 1960s and 1970s has been well documented (Delgado Pérez, 1994; Miret-Gamundi, 1997; Delgado and Castro Martín, 1999; Baizán, 2001; Billari et al. 2001a). For instance, according to the Family and Fertility Survey of 1995, by the age of 24, 71.2 percent of the women of

the 1955-59 birth-cohort had entered a first partnership, while only 53.3 of the women of the 1965-69 birth-cohort did. The corresponding figures for first birth are 47.0 and 33.2. A parallel trend has been the postponement of marital births, which somewhat weakens the existing strong linkage between marriage and motherhood roles (Castro Martin, 1992). However, the proportion of first births conceived before marriage (born less than 7 months after marriage) has substantially increased for the marriage cohorts of the late 1970s and of the beginning of the 1980s (Muñoz Pérez, 1991). This is possibly a consequence of the liberalisation of sexual behaviour, coupled with an insufficient control of contraceptive use and with social and legal restrictions to abortion[1]. This may also reflect the existence of a strong norm that children should be born inside unions, and that women (or couples) want to avoid the penalties associated with out-of-wedlock births. For illustrative purposes, it can be pointed out that in our sample, based on FFS data, around 97 percent of first births took place inside a union (91% in a marriage and 6% in a consensual union), for the female cohorts born in the 1950s and 1960s.

Successive fertility surveys show an increase in the prevalence of contraception (Instituto Nacional de Estadística, 1978 and 1987; Delgado and Castro Martín, 1999; Spinelli et al. 2000). The contraceptive prevalence rate among married women was 47 percent in 1977, and 58 percent in 1985, rates which are rather low, compared to other industrialised countries. Single women even had substantially lower levels of contraceptive use than married women did. Furthermore, an important segment of both married and single women relied on traditional methods, such as withdrawal. These patterns of fertility regulation have been affected by the legal banishment - until 1978 - on the use, sale and distribution of contraceptives, as well as on the dissemination of family planning information.

Qualitative studies suggest that consensual unions are characterised by more egalitarian gender roles than marriages and by an explicit criticism to the institution of marriage by the individuals involved (Alabart et al., 1988; Cabré et al., 1988). The FFS survey, which is the first representative data set for the whole of Spain that allows a detailed analysis of cohabitation, enables us to compute some useful

[1] Induced abortion has been legally permitted in Spain since 1985 on a restricted basis.

descriptive statistics. Tables 1 and 2 present the proportions of first unions that began outside of marriage, by birth cohort and sex. They show the extremely low proportions of first unions that started in that way by the cohorts born in the 1940s and 1950s, as well as a jump in the proportion of cohabitants for the birth-cohorts born in the 1960s. Nevertheless, this increase in cohabitation has been far from sufficient to compensate for the sharp decline in the proportion of individuals married for the cohorts born in the 1960s, as shows the declining proportions of people in partnerships. Given the still socially deviant status of cohabitation in Spain, some cohabitants may have hidden their situation, presenting themselves as married from the beginning of their union[2]. Therefore, the FFS data presented in Tables 1 and 2 should be interpreted as a rough estimate of the phenomenon, reflecting its general trend, at least until new data allows for a better perspective on non-marital cohabitation.

Finally, some references to the characteristics of the gender system in Spain can be made, as they are useful in interpreting the process of family formation. Recent decades have witnessed some emerging patterns that point in the direction of greater gender equality. They include important improvements in women's education and attachment to the labour force; the weakening of the figure of the male provider (Luxán et al, 2000); the inclusion of gender equality principles in law (Valiente, 1995) and in some elements of the expansion of the welfare state (Guillén, 1992; Sarasa and Moreno, 1995); and the relaxation of traditional family hierarchies (Valero and Lence, 1995). However, these changes have been very partial, and the traditional gender order is still to a large extent in place (González et al., 2000). Furthermore, important differences are still present within the socialisation process (Morales, 2000), as well as in the values and attitudes concerning work and family life according to gender (Centro de Investigaciones Sociológicas, 1994, 2000).

[2] In a country like Great Britain where cohabitation is much more frequent, there are important discrepancies between data sources (Murphy, 2000). Moreover, the number of individuals that report having started their first union as a cohabitation shows several inconsistencies between subsequent waves of a survey for the same individuals, suggesting changes in respondents' perceptions of their relationships over time, coupled with changes in the social desirability of reporting informal unions (Berrington, 1995).

3. Hypotheses

During the transition to adulthood, individuals are involved in the achievement of several interrelated events. These events are outcomes of processes, which interact dynamically with each other, and with the different contexts in which the individual is placed (Buchmann, 1989; Liefbroer, 1999; De Bruijn, 1999). An important feature of the life course approach is that it emphasises the role of norms on the sequence of events (Marini, 1985), and that such norms vary according to regional context, or according to social class, for example. Prevalent family models prescribe that having children is only expected within a stable (cohabiting or married) relationship (Roussel, 1989; Jurado Guerrero and Naldini, 1997; Alberdi, 1999). Consequently, forming a union becomes part of the strategy leading to procreation. Furthermore, one's desire for children will influence union formation and its timing. In addition, sequencing norms may help to explain that a pregnancy generally leads to a union formation before the birth or shortly after it. Therefore, we expect differences among members of a population in the strength and in the compliance to the norms concerning the sequencing of union formation and childbirth.

Other normative sequences of events that affect the time link between first union formation and first birth have also been proposed. For instance, union formation and childbirth should take place after school completion (Blossfeld and Huinink, 1991). And, at least for men (and increasingly for women), having a consolidated position in the labour market is often seen as a pre-requirement for family formation.

Some authors have argued that value orientations contribute significantly to explaining family formation (among others, see Lesthaeghe and Moors, 1995). For Spain, this has also been argued to explain behaviours which are not fully explained when using economic variables (Holdsworth, 1998; Billari et al., 2001b). Referring to values, Delgado Pérez and Livi-Bacci (1992) report differences on ideal family sizes by regions and social class. Reher (1998) emphasises the importance of historical cultural forms of family formation to differentiate between Northern and Southern Europe.

Parental attitudes are also important determinants of the attitudes and behaviour of young adults through the processes of social influence. Orientation towards career or family can then be a consequence of the views of parents (Barber, 2000).

The risks of first birth and first union formation may be affected by joint factors involving favourable attitudes and intentions towards family life. Many people may see childbearing and union formation as elements of a whole and same process, namely the attainment of a family. For instance, the idea of a "bourgeois"[3] family supposes this type of identification. Even if some authors claim that union formation and first birth have increasingly become disconnected to each other (see for instance Van de Kaa, 1997; Corijn and Klijzing, 2001), it is clear that value orientations play a role in making such behaviours more connected for some individuals than for other individuals.

Furthermore, family formation may be embedded in a choice process in which young adults seek strategic balancing of their family-life goals with goals in other domains (Liefbroer, 1999). Consequently, motherhood is seen by some women as being incompatible with a self-directed biography. Some authors have even opposed a "family orientation" dimension versus a "work orientation" to explain inter-individual differences in the trajectories and their consistency for each individual (Willekens, 1991).

It has not been possible so far, given the lack of appropriate panel data, to properly assess the impact of cultural factors on family formation behaviour in a context such as Spain. For these reasons, in the analyses done so far, unobserved characteristics have potentially affected and biased the results; moreover, the role of cultural factors might have been understated in empirical analyses.

Our first hypothesis derives directly from this literature:

[3] The "modern" (Roussel, 1989; Rezsohazy, 1991) or "bourgeois" (Sorokin, 1947; Mitterauer and Sieder, 1982) family model is characterised by a strong separation of roles between the members of the couple, by a limited number of children, and by an ideal of conjugal intimacy.

H1: *The timing of first union and of first birth is, at least partially, determined by joint factors, and such factors make them positively interrelated.*

Once the interrelationship between first birth and first union has been identified, we are able to discuss the mutual impact of these events, net of the effects of common factors. The literature has repeatedly shown that being in a union drastically raises the risk of conception. This cannot be a surprise, especially in (continental) Western Europe, where births take place almost exclusively within marriages or consensual unions (Kiernan, 1999, 2001). Prevailing institutional models concerning nuptiality and fertility in Spain include the idea that the environment of a union is the most suitable situation for rearing children (Jurado Guerrero and Naldini, 1997; Marí-Klose and Nos Colom, 1999). For the couple, a union normally implies a certain level of commitment, which brings stability. This is especially the case for marriage, which favours a shared long-term commitment associated to having a child. Thus, a union often acts as insurance against an unfavourable position in terms of work and education. In addition, individuals may prefer to avoid the difficulties associated with having children outside a union, in particular its potential negative effect on one's educational career and work career. As a consequence of these circumstances and normative expectations, individuals in a union may develop more favourable attitudes and intentions towards childbearing than single people do. Social pressure and expectations (from parents or others) related to procreation might also increase once individuals are in a union (Barber and Axinn, 1998).

Additional impacts may be the ones affected by sexual activities: a higher sexual activity of cohabiting and married people compared to singles might raise the risks of conception if contraception is not perfect, as has been observed in the United States (Rao and DeMaris, 1995).

Economic theory also predicts an increase in the risks of first birth after union formation. The union can be considered an institution where the production of children, i.e. child bearing and rearing, is more efficient due to the division of labour (Becker 1981). Children are union-specific capital, and can be viewed as a rational investment based on the long-term prospective of the union, which allows a certain degree of role specialisation for the couple. Consequently, bearing children outside a

union is generally considered to be less beneficial. In addition to having a higher direct cost, it may also hamper the individual's attractiveness in the marriage market, lowering any future marriage prospects.

These arguments suppose the existence of differential effects for marriage and cohabitation. To sum up, marriage is viewed by many as a more permanent living arrangement than cohabitation. It is laden with a higher degree of normative component and, in case of union dissolution, it offers several legal compensations for the economically weakest partner. However, as cohabitation loses its marginal status in recent times, and as the social acceptance of cohabitation increases, there may be less pressure to marry in order to have children, while a partnership context may still be viewed as necessary (Mulder and Manting, 1993).

Based on this reasoning we state the following hypothesis:
H2: *Union formation has a strong effect on the risks of first birth, which is independent of common factors. This impact is higher in the case of marriage than in the case of cohabitation. Through birth-cohorts, the gap between cohabitation and marriage increasingly narrows in terms of bearing children.*

Our third hypothesis concerns the effect of first childbirth on union formation. A positive independent impact of pregnancy and first birth on first union can be expected. Single women who become pregnant may form a union because of a desire to offer their child the social and economic protection that normally accompanies a union. Furthermore, normative pressures are likely to increase the incentives to legitimise the birth through an acceleration of union formation. In terms of theories of marriage market search (see for instance Keeley, 1977), the cost of searching for a partner is believed to increase after a woman experiences an out-of-union birth. The theory predicts that a woman who has an out-of-union birth spends less time searching before she settles into a partnership. However, in terms of empirical research, relatively little has been done to establish the effect of pre-union childbearing on union formation. An exception is Goldscheider and Waite (1986), who find that premarital births have a strong positive effect on marriage. In a recent contribution, Brien et al. (1999) consider separately the effect of a pre-union pregnancy on cohabitation and marriage. They find that a premarital birth generally

accelerates marriage, but that this acceleration fades quickly for those who do not marry right after the birth. In terms of cohabitation, on the other hand, the overall effect is considerably weaker.

In addition, we expect the effect of pregnancy/birth to be highly time-dependent. There are several reasons for this. First, becoming aware of a pregnancy implies a time lag from the actual conception time. Consequently, being pregnant is only going to influence the decision to enter a union one or two months after conception. Second, women may want to avoid an out-of wedlock birth or an out-of-union birth in order to comply with social norms and expectations. In that case, their decision to enter a union will probably take place around the middle of the pregnancy, when the pregnancy is less visible than in a more advanced pregnancy stage or when the child is already born. Otherwise, a woman may decide to have an abortion[4]. The effect of pregnancy on first union formation should therefore be concentrated during pregnancy or shortly after delivery, under the conditions of the birth-cohorts studied.
The above discussions take into account the fact that contraception is far from being perfectly regulated for the birth-cohorts studied, as mentioned in section 2. Furthermore, the age at first sexual intercourse is considerably lower than the age at first union, and has shown a tendency to decline through the birth-cohorts[5].

Finally, we expect that the effect of pregnancy is higher on marriage formation than on entering cohabitation, given the stronger normative and contractual nature of marriage.

H3: Pregnancy and first birth have a strong effect on union formation, independent of common factors. This impact decreases shortly after the birth of the child. A pre-union pregnancy results more often in a marriage than in cohabitation.

[4] The fact of not observing abortions, as is often the case in demographic surveys, should lead to an overestimation of the pregnancy/birth effect, because pregnant women that would not have wanted to enter a union are systematically under-represented.
[5] The median age at first sexual intercourse was 23.1 for the female birth-cohort 1945-49 and 20.1 for the women born in 1970-75 (Delgado and Castro Martín, 1999, table 21).

4. Methods and data

4.1 Statistical models

We develop our modelling strategy in order to disentangle the various effects depicted in our theoretical hypotheses. The transitions included in the analyses are first birth on the one hand, and first union formation on the other, modelled as a competing risks process. We do not include life segments where individuals are separated, divorced, widowed, or in a subsequent union. In other words, any of the trajectories that are disrupted are censored at the point of disruption of their first union. We considered married individuals that got married after a period of cohabitation as having a different state than the marriage of singles (i.e. women that never cohabited prior to entering their first unions). In the competing risk process of union formation, we include a covariate for having had a first child (conception).

Separate models for men and women are desirable in this setting, since each gender may differ considerably in their behaviour. However, given a sample size of men that was too small, we restricted the estimation to include women only. In the models to be estimated, the basic time unit is the age attained by the individual. Since we want to be able to compare the risk of cohabitants with that of single and married people, the timing of the event is expressed in terms of the life of the individual, not in terms of union duration, for example.

Time to first birth and time to first union are endogenous in that survival in one state depends on the outcome of the other process. Here, the endogeneity of union formation in the hazard of first birth is explicitly addressed by allowing unobserved heterogeneity to be correlated across the two processes of marriage/cohabitation and entry into parenthood. In that way, it is possible to control for common unmeasured factors that simultaneously influence first birth and union formation.

The statistical specification is derived from the framework developed by Lillard (1993). It consists of three simultaneous hazard rate equations, capturing time (since age 15) to first birth and to first union formation (by destination), respectively.

$$
\begin{aligned}
\ln h^{B}(t) &= y^{B}(t) + \sum_{k} z_{k}^{B}(u_{k} + t) + \sum_{j} a_{j} x_{j} + \sum_{i} a_{i} w_{i}(t) + e \\
\ln h^{C}(t) &= y^{C}(t) + \sum_{k} z_{k}^{C}(u_{k} + t) + \sum_{j} b_{j} x_{j} + \sum_{i} b_{i} w_{i}(t) + d \\
\ln h^{M}(t) &= y^{M}(t) + \sum_{k} z_{k}^{M}(u_{k} + t) + \sum_{j} b_{j} x_{j} + \sum_{i} b_{i} w_{i}(t) + d
\end{aligned}
\tag{1}
$$

The subscript for an individual is suppressed for simplicity. The superscripts *B*, *C* and *M* denote first birth, entering first cohabitation, and first marriage formation, respectively. The union formation intensity equations are modelled as a competing risk process, with a common heterogeneity component d[6]. Each *y(t)* denotes a piecewise linear spline[7] that captures the effect of the duration on the intensity. The $\{z_k\}$ are splines that capture the effects of some covariates that are continuous functions of *t* and z_k starting from an origin u_k relevant to each individual. In the equation for the process of first birth, these splines consist of the effects of duration since marriage formation (first union), the duration since the start of cohabitation (first union), and the duration since marriage formation (for individuals having started their first unions as cohabitants). In the equation for the process of union formation, the spline represents time since the conception of the first child.

The $\{x_j\}$ denotes fixed time-invariant covariates; and $\{w_l(\cdot)\}$ are a set of time-varying covariates whose values change at discrete times in the spell, and are constant over the time span between those changes. The random variables e and d capture unobserved heterogeneity, and are assumed to have a joint bivariate normal distribution:

$$
\begin{pmatrix} e \\ d \end{pmatrix} \sim N\left(\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} s_{e}^{2} & r_{ed} \\ r_{ed} & s_{d}^{2} \end{pmatrix} \right) \tag{2}
$$

[6] Unfortunately, it is not possible to identify separate heterogeneity components in the setting of a competing risk process (Lancaster, 1990, p. 154).

[7] Piecewise linear splines are used to approximate continuous functions (such as a baseline hazard or a non-proportional relative risk), by using functions that are linear within each (possibly open-ended) interval. Those linear functions are connected at knots given *a priori*: piecewise linear splines are then also continuous functions.

in which $?_{ed}$ is the correlation between the unobserved heterogeneity terms of the process[8].

The model estimation was performed using full-information maximum likelihood, as implemented in the package aMl (Lillard and Panis, 2000).

4.2 Data and construction of the variables used

The data comes from the Spanish Fertility and Family Survey, conducted in 1995 (Delgado and Castro Martín, 1999). This is a retrospective survey on a sample of 4021 women and 1991 men, born between 1945 and 1976. Here we only use the female sample. Some data cleaning procedures were implemented. Twenty-one respondents were excluded from the analyses, because they started a first union and/or had a first child before the age of 15. In a small number of cases, the month that particular events occurred was missing, and therefore, these values were assigned randomly.

When studying the effect of marital status on first births, it is useful to account for the effect of premarital pregnancies, since some marriages and cohabitations are the direct outcome of conceptions. In order to avoid this problem of reverse causation, the dependent variable is taken as the time of birth minus eight months. In this way, those marriages or cohabitations that start eight months or less before delivery will not be taken into account when computing the first birth hazard. The reason for not subtracting nine months is the following. In the case of an exact simultaneity with marriage formation or the start of cohabitation (that is, if the two events take place the same month), the conception can be the result of the marital status change. If a conception and a change of marital status both take place the same month, the marital status change can not be the direct result of the pregnancy. This is because the women are most likely unaware of the pregnancy at the time. In contrast, the conception may be the result of the union. An accurate measure of the conception hazard will be

[8] Several experiments where conducted in order to test the sensitivity of the results to different values of the variance of the unobserved heterogeneity component. A value of 1 for the standard deviation was finally adopted. This is also in analogy with discrete-time event history probit models. Though the magnitude of effects was somewhat affected by the value of the variance, the sign and significance was not (see table 6).

obtained when these synchronised situations are counted as made by already married/cohabiting individuals.

In the first birth process, the observation is considered to be censored when (a) the individual has not had a first birth at the time of the interview, or (b) the first union is disrupted, in which case censoring occurs at the time when the partnership ends.

In the process of union formation, the observation is considered as censored when the individual has not entered a first union at the time of the interview.

When studying the effect of first birth/conception on first union formation, we want to distinguish the effect of being pregnant (leading to first birth), as opposed to not being pregnant (leading to a first birth), or to already having a child. Union formation behaviour taking place from the eighth month before birth onwards may be affected by the pregnancy or the birth of the first child, but not union formation taking place the same month as conception or before, since the women do not know at that stage if they are pregnant. Therefore, it is adequate that the spline for having a child starts ticking the ninth month before birth.

The data on cohabitation is based on the questions, "In what month and year did you first start living with your first partner in the same household?" and "Were you married to him/her when you started living together?" These questions were preceded in the interview by the question, "Have you ever lived in the same household with someone you had an intimate relationship with but did not marry?" These questions are adequate tools to measure the prevalence of cohabitation. However, as discussed in section 2, we suspect some underreporting of cohabitation in our data.

A time-varying dummy variable is used to control for educational enrolment. As an alternative specification we experimented with a duration spline that kicked in at the moment of completion of the latest educational enrolment (see Brien et al. 1999). However, this specification was not used since the shape of the duration dependence was essentially flat after an initial jump at the moment of leaving education, i.e. it is *grosso modo* equivalent to a dummy variable. We use the level of the respondent's education as an approximate measure of human capital. This variable is often used to

test the "independence" hypothesis[9] (Goldscheider et al., 2000; Becker, 1981). The Spanish FFS contains full histories of educational enrolment including the date of attainment of each particular level of education. Given the relatively high number of individuals (22 percent) in the sample that did not attain a lower secondary level of education, a category for primary education was included. The International Standard Classification of Education was used as the basis to generate this variable (Eurostat, 1996). Data from questions on the start of employment and work interruptions were used to indicate the months the respondent was employed.

We also included other background variables in the two equations. These were the number of siblings of the respondent, a dummy indicating if the parents ever disrupted their union (before the respondent was 15 years old), and job experience. The number of siblings was included only in the equation for first birth, whereas an indicator for whether parents had experienced a disruption was included in the equation for union formation. These variables tried to test identification and the robustness of the specification. However, their inclusion did not appreciably improve the fit of the model, nor did they have a great impact on the remaining coefficients, and were consequently dropped from the final specification. Moreover, the results of these variables were similar to those obtained by other studies, i.e. a positive effect of the number of children and of the parental union disruption on the risks of first birth and of cohabitation (the coefficients for marriage were close to zero and not significant). The variable "job experience" was kept because it complements the information on work status and educational level. That is, it captures better the opportunities individuals face in the labour market, as well as their attachment. Work experience is an indicator of human capital accumulation in the labour market, and thus may reflect an 'income effect' on the transitions studied, as well as the opportunity costs of having children (see for instance Cigno and Ermisch, 1989; Kravdal, 1994).

5. *Results*

The main results of the analyses are presented in Table 3, where panel A concerns the parameters for the process of first childbirth, panel B reports the coefficients for

[9] This hypothesis states that the rise in female employment provided women the independence to avoid family roles, either by not entering them or by leaving them.

cohabitation, and panel C the coefficients for marriage. For comparative purposes, two models are reported: Model 1, which does not include the unobserved heterogeneity components, and Model 2, which does. Apart from the heterogeneity terms, the specifications are the same as in equation (1). Unless explicitly stated, the estimates and the comments in the text refer to Model 2.

5.1 Interrelationship between events

In the first hypothesis of the paper, we proposed that the processes of first birth and first union formation share some common unmeasured factors. These common factors are essentially captured by the correlation between the heterogeneity components of each process. As expected, there is a positive and significant correlation between these heterogeneity components, with a value of 0.65. Those young women who are most likely to have a first birth (for reasons we do not measure), are also most likely to form a union. This suggests that both events are part of the same process, which is essentially a couples' family formation, and that if these common factors are not taken into account in the models, we would get a biased picture. Although it is not possible to elucidate the exact nature of the variables underlying the heterogeneity components, in the section on hypotheses we proposed some possible common factors underlying the timing of first birth and first union formation. These factors could include values and intentions towards family life, driven by normative and behaviour guiding elements of the family institution.

By comparing the baseline hazards of the models with and without heterogeneity components (Models 1 and 2 respectively), it is possible to explore the selection effects generated by the unobserved heterogeneity. From age 15 to around age 23, the baseline risks of Model 1 are considerably higher than the risks of Model 2, while from that last age onwards the situation is reversed. In the model without the heterogeneity component, individuals with high probabilities of completing both events early in the life course leave the population at young ages. Thus, there is indeed an overrepresentation of individuals with low family formation proneness at higher ages. When this selection effect is accounted for, the shape of the baseline risks shows a more marked contrast according to age.

Furthermore, the introduction of correlated heterogeneity has important consequences for the estimated effects of the covariates. In particular, the effects of marital status (or parenthood status), modelled as piece-wise linear splines, generate a lower baseline level at younger ages and a higher level at more advanced ages. A comparison of the two models shows that the effects of these covariates are very different. Thus, failing to control for the shared heterogeneity factor produces an overestimation at younger ages and an underestimation at older ages.

5.2 The effects of marital status on first birth

The results of Model 2, reported in Table 3 (panel A) and illustrated in Figure 1, show that entering a first union sharply increases the hazard of first birth. This effect takes place immediately after the start of the first union. The relative risk of first birth increases 25.22 times for direct marriage and nearly by 7 for cohabitation. It further increases during the three initial years of marriage, then it declines, while in the case of cohabitation, it continuously declines until the third anniversary of union, and remains stable afterwards. A marriage of a cohabiting couple produces a considerable additional increase (3.22) in the relative risk of first birth.

These results provide strong support for our second hypothesis, which stated that union formation has an independent effect on the risks of first birth. They suggest that union formation is indeed viewed by couples as the most appropriate setting for having a child. In addition, they clearly show a preference for having the first child in a marital union rather than in a cohabiting union, possibly due to normative influences. The first few years of the union produce the highest levels of childbearing.

The results discussed thus far are based on estimates that account for the endogeneity of the marriage of singles in the first birth process. The results of Model 1, where zero correlation between the processes is assumed, show some differences in the shapes of the effects of union on first birth. They indicate some overstatement of the immediate effect of marriage formation: the coefficients for Model 1 is 3.64, while for Model 2 it is 3.23. However, after the first few months of marriage formation have passed, it would have resulted in a substantial and increasingly important underestimation of the effects of marriage. A similar pattern is found for cohabitation. The shift in the hazard

of a first birth when entering cohabitation is higher for Model 1 than for Model 2, while afterwards the hazard is slightly underestimated in the former model. Finally, the underestimation of the effects of the marriage of cohabitants is especially important when endogeneity is not taken into account. These results are consistent with the estimated positive and statistically significant correlation between the processes of union formation and first birth.

5.3 Inter-cohort changes in the effects of union status on first-birth patterns

In order to facilitate an investigation into the changes taking place over birth-cohorts, we have estimated a separate model. Here the endogenous time varying variables were not estimated as piece-wise linear splines. Instead, a time-varying covariate for marital status is used, thus providing a coefficient for each cohort and its statistical significance. The other explanatory variables remain as in Model 2, including the correlation between the heterogeneity components.

The results are presented in Tables 4 and 5. They provide a good picture of the evolution in the relative risk over birth-cohorts. The two tables are based on the same coefficients (for which the reference category corresponds to the cohabitants of the 1945-54 birth-cohort), but they focus on different comparisons, intending to facilitate the interpretation of coefficients.

Table 4 shows the expected pattern of an increase in the relative risk of first childbirth for cohabitants. The birth-cohort 1965-75 presents a relative risk of 1.63 with respect to the 1945-54 birth-cohort. If the comparison is made with married women without a previous cohabitation (Table 5), the results show that the effect of cohabitation becomes close to that of marriage in terms of bearing children, although there are still considerable differences in the younger cohort. It should be noted that the differences across birth-cohorts are not statistically significant, a feature which is likely to be connected to the small numbers involved. Nevertheless, the results indicate that cohabitation is increasingly viewed as a suitable situation for bearing children.

The trend for married women with no previous cohabitation (Table 4), in contrast to the trend for those who have cohabited, shows a sharp decline in the relative risks, which is connected with the increasing postponement of marital fertility.

5.4 Effects of pregnancy and age of first child on first union formation

Figure 2 presents the risks of entering marriage and cohabitation, according to the age of the child (see also Table 3, panels B and C). It can be seen that during the period of pregnancy there is a dramatic increase in the risks of entering a union, with respect to childless individuals. The average relative risk of marriage for the first half of the pregnancy is 10.50 and for the second half is 37.48, while the corresponding figures for cohabitation are 3.79 and 11.79. These results show that couples generally want to avoid an out-of-union childbirth, and in particular an out-of-wedlock birth. As expected, the discovery of a pregnancy leads to an increase in the risks of union formation. During the first year after the birth, the risks of union formation decline, implying average levels of 7.77 for marriage and 3.84 for cohabitation. After the first year of birth, the relative risks of union formation are low, but still higher than that of single individuals. Overall, a conception and a birth are more often followed by marriage than cohabitation, indicating that marriage is still seen by many as a more appropriate setting for rearing children than cohabitation is.

It should also be mentioned that the results of Model 1, where the heterogeneity components were not included, show that the effects on both cohabitation and marriage are severely underestimated during pregnancy, and overestimated afterwards. Our results clearly show the high time-dependency of the effects of pregnancy/birth on entering a union, and the importance of modelling duration effects.

5.5 The effects of education, activity status, and work experience

As several studies have shown, educational enrolment diminishes the propensity to form a union and childbearing. This is also the case here. Spanish students usually live in the parental home and are unable to live independently due to the lack of own resources, and this affects their capacity to form a union and to have a child (Billari et al. 2001b). However, being a student is much less incompatible with entering cohabitation than with entering marriage, since their relative risks are 0.55 and 0.26 respectively, with respect to not enrolled women, while the relative risk for first birth is 0.34 (Table 3, Model 2).

The results for educational attainment show a strongly negative gradient for first birth and marriage. Higher education means higher professional career expectations (that may often be difficult to attain, given the extremely high unemployment rate in Spain in recent decades), and a need to consolidate one's career before forming a family. It is interesting to see that this effect becomes sharpened in Model 2, where unobserved heterogeneity is accounted for. The results for cohabitation indicate, on the contrary, a positive gradient in the effect of the educational level, consistent with the interpretation that the educated value independence and autonomy more than young adults with low levels of educational attainment. In addition, they are probably more able to practice less conservative behaviours.

Being employed reduces the intensity of each of the three events studied. The reduction is, however, smaller for entry into cohabitation than into marriage, which is consistent with the alleged higher gender role equality among cohabitants. Theoretical arguments foresee potentially opposing impacts of women employment in family formation. On the one hand, a statistically positive impact may be expected, since employment increases the resources necessary to form a family and because it may increase individuals' attractiveness in the marriage market. On the other hand, it may have negative effects, through a self-reliance or independence effect. However, current employment does not necessarily reflect the labour force attachment of a woman, since many women may stop working in order to have a child, in a context of high incompatibility between both activities. Moreover, unemployment may provide an opportunity to have a child (or even to enter a union), especially if unemployment benefits are associated with it.

The work experience variable may better capture the attachment to the labour market and, in combination with the educational level, the individual's income potential. The relative risks in Model 2 show a strong positive impact of work experience on both union formation events, and a much less important effect on first birth. This impact is especially important in the case of marriage, since it increases the relative risks of women with four of more years of work experience by nearly three times, with respect to women without job experience. The results for the work experience variable obtained here contradict the idea that human capital accumulation for women

tends to deter union formation and childbearing. On the contrary, they point to the importance of the accumulation of resources by women, in facilitating both union formation and childbearing.

6. Conclusions

In this paper, we have simultaneously studied the presence and effect of constant common factors, unmeasured in standard retrospective surveys, influencing the timing of first childbirth and first union formation, and the mutual impact of such events. The modelling strategy employed here, based on the simultaneous hazard equations approach developed by Lillard (1993), overcomes the bias due to the correlation between the unobserved heterogeneity components belonging to each process. Our findings confirm the existence of a positive and significant correlation between these heterogeneity components for the Spanish case, indicating that those women who are most likely to have a first birth at each stage in life are also most likely to form a union. This suggests that first birth and first union formation are part of the same process of family formation and that low fertility and postponement in first unions are partially determined by joint factors.

Our findings also show that, if the strong interrelationship between the events studied is not controlled, this leads to important distortions in the estimates of the mutual effects of first birth and first union formation. These include an overestimation of the effects at younger ages and underestimation of these effects after the approximate age of 24. Moreover, the time-shape of these effects is also affected by the presence of the unmeasured heterogeneity factors between the individuals of a population.

The results obtained show that, net of the common heterogeneity factors, union formation and first birth have an independent impact on each other. This also has a practical methodological consequence. That is, in order to obtain reliable estimates, studies that focus on the process of first birth should include the effects of union formation and a heterogeneity component that accounts for their mutual correlation. The influence of the union status is not spurious, and therefore, if not included, an important determinant of first birth would not be taken into account. Similar

comments can be made concerning the influence of pregnancy/first birth on the timing of union formation.

The analysis concerning the effects of union formation on first childbirth shows a strong effect, essentially concentrated during the first few years after the start of the union, or shortly after the formalisation of the consensual union. This effect is considerably more important for marriage than for cohabitation, but it may also depend of the meaning of cohabitation and marriage in each context. In this respect, we have found indications of an increasing role of cohabitation on first childbearing through birth-cohorts.

Finally, we have shown that the impact on union formation of conception leading to a first birth is mainly circumscribed to the period of pregnancy until shortly after delivery, while afterwards the impact is very low.

The findings reported in this paper have important policy relevance. They suggest that any policy that encourages union formation (such as facilitating access to housing or tax relieves) will have a positive impact on fertility. This fostering impact of union formation on first births also includes consensual unions, albeit to a lesser extent than marital unions.

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Table 1. Proportion of first unions begun outside marriage, by birth-cohort and sex

Birth-cohort	*1945-49*	*1950-54*	*1955-59*	*1960-64*	*1965-69*	*1970-77*
Men	6.3	11.6	15.1	27.2	30.5	37.2
Women	2.6	4.5	7.4	11.5	19.0	31.3
Total	4.0	6.6	9.7	16.5	21.8	32.1
Number of cohabitations	25	50	78	141	149	91

Source: based on the Family and Fertility Survey data for Spain

Table 2. Proportion of first unions (before birthday 30) begun outside marriage, by birth-cohort and sex

Birth-cohort	*1945-49*	*1950-54*	*1955-59*	*1960-64*	*1965-69*
Men	4.3	9.5	15.5	28.7	30.5
Women	1.9	3.6	6.9	11.4	19.0
Total	2.7	5.3	9.3	16.6	21.8
Number of cohabitations	15	37	69	134	149

Source: based on the Family and Fertility Survey data for Spain

Table 3. Results of estimation (panel A).

	Model 1			*Model 2*		
Parameter	*Estimate*	*S.E.*	*Relative Risk*	*Estimate*	*S.E.*	*Relative Risk*
Process: First birth (conception)						
Baseline						
constant	**-4.79**	***0.22*		**-5.44**	***0.22*	
Age 15-18 (slope)	**0.35**	***0.06*	§ 1.59	**0.44**	***0.06*	# 1
Age 19-22 (slope)	**-0.07**	***0.02*	§ 2.78	**0.03**	*0.03*	# 2.58
Age 23-25 (slope)	**-0.01**	*0.03*	§ 2.38	**0.09**	**0.04*	# 3.15
Age 26-28 (slope)	**-0.00**	*0.04*	§ 2.34	**0.10**	**0.05*	# 4.17
Age 29-31 (slope)	**-0.01**	*0.06*	§ 2.28	**0.04**	*0.06*	# 5.19
Age 32+ (slope)	**-0.10**	**0.03*		**-0.07**	**0.04*	
Marriage of single						
(reference=singles)			1			1
Enter marriage shift	**3.64**	***0.08*	37.97	**3.23**	***0.13*	25.22
0-1 year (slope)	**-0.30**	***0.09*	# 32.71	**0.18**	*0.12*	# 27.54
1-3 years (slope)	**-0.18**	***0.05*	# 23.64	**-0.01**	*0.12*	# 29.80
more than 3 years (slope)	**-0.24**	***0.03*		**-0.23**	***0.04*	
Cohabitation						
(reference=singles)			1			1
Enter cohabitation shift	**2.37**	***0.30*	10.69	**1.95**	***0.33*	7.02
0-1 year (slope)	**-0.37**	*0.39*	# 8.87	**-0.31**	*0.40*	# 6.00
1-3 years (slope)	**-0.12**	*0.17*	# 6.52	**-0.17**	*0.18*	# 4.34
more than 3 years (slope)	**0.02**	*0.06*		**0.00**	*0.08*	
Marriage of cohabitant						
(reference=cohabitants)			1			1
Enter marriage shift	**0.91**	***0.23*	2.49	**1.17**	***0.26*	3.22
0-2 years (slope)	**0.07**	*0.19*	# 2.66	**0.18**	*0.20*	# 3.85
more than 2 years (slope)	**-0.14**	*0.09*		**-0.06**	*0.10*	
Educational level						
Primary	**0.21**	***0.05*	1.23	**0.38**	***0.07*	1.46
(reference=lower secondary)						1
Upper secondary	**-0.26**	***0.08*	0.77	**-0.60**	***0.10*	0.55
University	**-0.25**	***0.09*	0.78	**-0.73**	***0.13*	0.48
Educational enrolment	**-0.95**	***0.13*	0.39	**-1.08**	***0.14*	0.34
(reference=not enrolled)			1			1
Employment status	**-0.43**	***0.05*	0.65	**-0.62**	***0.06*	0.54
(reference=not employed)			1			1
Birth cohort						
(reference=1945-54)			1			1
1955-64	**-0.05**	***0.05*	0.95	**-0.01**	*0.08*	0.99
1965-75	**-0.31**	***0.06*	0.73	**-0.37**	***0.09*	0.69
Work experience						
(reference=less than 1 year)			1			1
1 to 4 years	**0.11**	**0.06*	1.12	**0.15**	**0.08*	1.16
more than 4 years	**0.12**	**0.06*	1.13	**0.09**	*0.08*	1.09
Correlation between first union and first birth				**0.65**	***0.17*	

Notes: ***=p<0.01, **=p<0.05, *=p<0.1. In model 2 also an unobserved factor influencing simultaneously first birth and the first union is included.

§ The reference category is the age group 15-18 of Model 2.

The relative risks for each age group are calculated by exponentiating the value of the average hazard for the ages concerned.

Table 3. Results of estimation (panel B).

Parameter	*Model 1* *Estimate*	*S.E.*	*Relative Risk*	*Model 2* *Estimate*	*S.E.*	*Relative Risk*
Process: First union formation (cohabitation)						
Baseline						
Intercept	**-7.66**	****0.45*		**-8.33**	****0.45*	
Age 15-18 (slope)	**0.44**	****0.11*	§ 1.66	**0.51**	****0.11*	# 1
Age 19-22 (slope)	**0.07**	*0.06*	§ 4.58	**0.20**	****0.06*	# 4.11
Age 23-25 (slope)	**0.17**	**0.17*	§ 6.83	**0.31**	****0.09*	# 9.65
Age 26-28 (slope)	**0.01**	*0.13*	§ 8.98	**0.13**	*0.13*	# 18.60
Age 29-31 (slope)	**-0.20**	*0.19*	§ 6.73	**-0.11**	*0.19*	# 19.12
Age 32+ (slope)	**0.03**	*0.08*		**0.06**	*0.08*	
First birth (conception)						
(reference=no child)			1			1
start pregnancy to half preg. (slope)	**7.43**	****1.16*	# 4.03	**7.11**	****1.24*	# 3.79
half pregnancy to birth (slope)	**-2.94**	*2.12*	# 9.35	**-1.06**	*2.15*	# 11.79
birth to 1 year (slope)	**-1.82**	*1.21*	# 2.17	**-1.84**	*1.20*	# 3.84
more than 1 year (slope)	**0.10**	*0.12*		**0.07**	*0.12*	
Educational level						
Primary	**-0.41**	***0.21*	0.66	**-0.31**	*0.21*	0.73
(reference=lower secondary)			1			1
Upper secondary	**0.29**	**0.15*	1.34	**0.06**	*0.16*	1.06
University	**0.58**	****0.20*	1.79	**0.22**	*0.21*	1.25
Educational enrolment	**-0.53**	****0.16*	0.59	**-0.60**	****0.17*	0.55
(reference=not enrolled)			1			1
Employment status	**-0.45**	****0.14*	0.64	**-0.67**	****0.14*	0.51
(reference=not employed)			1			1
Birth cohort						
(reference=1945-54)			1			1
1955-64	**1.01**	****0.24*	2.75	**1.14**	****0.25*	3.13
1965-75	**1.53**	****0.24*	4.62	**1.57**	****0.25*	4.81
Work experience						
(reference=less than 1 year)			1			1
1 to 4 years	**0.66**	****0.16*	1.93	**0.75**	****0.17*	2.12
more than 4 years	**0.70**	****0.19*	2.01	**0.83**	****0.20*	2.29

Notes: ***=p<0.01, **=p<0.05, *=p<0.1. In model 2 also an unobserved factor influencing simultaneously first birth and the first union is included.

§ The reference category is the age group 15-18 of Model 2.

The relative risks for each age group are calculated by exponentiating the value of the average hazard for the ages concerned.

Table 3. Results of estimation (panel C).

Parameter	*Model 1* *Estimate*	*S.E.*	*Relative Risk*	*Model 2* *Estimate*	*S.E.*	*Relative Risk*
Process: First union formation (direct marriage)						
Baseline						
Intercept	**-4.99**	****0.21*		**-5.81**	****0.21*	
Age 15-18 (slope)	**0.57**	****0.06*	§ 1.97	**0.64**	****0.06*	# 1
Age 19-22 (slope)	**0.24**	****0.02*	§ 9.90	**0.40**	****0.02*	# 8.12
Age 23-25 (slope)	**0.05**	**0.03*	§ 17.06	**0.21**	****0.03*	# 24.84
Age 26-28 (slope)	**-0.18**	****0.04*	§ 14.01	**-0.08**	**0.04*	# 30.32
Age 29-31 (slope)	**-0.11**	**0.06*	§ 9.11	**-0.05**	*0.07*	# 25.21
Age 32+ (slope)	**-0.20**	****0.05*		**-0.18**	****0.05*	
First birth (conception)						
(reference=no child)			1			1
start pregnancy to half preg. (slope)	**11.83**	****0.17*	# 9.19	**12.54**	****0.45*	# 10.50
half pregnancy to birth (slope)	**-7.44**	****0.59*	# 20.93	**-5.75**	****0.64*	# 37.48
birth to 1 year (slope)	**-0.94**	****0.34*	# 3.24	**-0.99**	****0.34*	# 7.77
more than 1 year (slope)	**-0.24**	****0.08*		**-0.31**	****0.08*	
Educational level						
Primary	**0.09**	***0.04*	1.09	**0.24**	****0.07*	1.27
(reference=lower secondary)			1			1
Upper secondary	**-0.23**	****0.06*	0.79	**-0.50**	****0.08*	0.61
University	**0.01**	*0.08*	1.01	**-0.38**	****0.11*	0.68
Educational enrolment	**-1.30**	****0.11*	0.27	**-1.35**	****0.12*	0.26
(reference=not enrolled)			1			1
Employment status	**-0.88**	****0.04*	0.41	**-1.17**	****0.05*	0.31
(reference=not employed)			1			1
Birth cohort						
(reference=1945-54)			1			1
1955-64	**0.15**	****0.04*	1.16	**0.30**	****0.07*	1.35
1965-75	**-0.30**	****0.06*	0.74	**-0.30**	****0.08*	0.74
Work experience						
(reference=less than 1 year)			1			1
1 to 4 years	**0.51**	****0.06*	1.66	**0.69**	****0.07*	1.99
more than 4 years	**0.87**	****0.05*	2.39	**1.08**	****0.07*	2.94
Loglikelihood	**-26840**			**-26678**		

Notes: ***=p<0.01, **=p<0.05, *=p<0.1. In model 2 also an unobserved factor influencing simultaneously first birth and the first union is included.

§ The reference category is the age group 15-18 of Model 2.

The relative risks for each age group are calculated by exponentiating the value of the average hazard for the ages concerned.

Table 4. Effects of union status on first birth behaviour (relative risks). Birth-cohort 1945-54 as reference. Women. Model with heterogeneity component

Union status *Birth-cohort*	*Single*	*Cohabitant*	*Married (no previous cohabitation)*	*Married (premarital cohabitation)*
1945-54	1	1	1	1
1955-64	2.07	1.38	0.80	1.14
1965-75	1.93	1.63	0.39	0.44

Source: based on the Family and Fertility Survey data for Spain

Table 5. Effects of union status on first birth behaviour (relative risks). Married with no previous cohabitation as reference. Women. Model with heterogeneity component

Union status *Birth-cohort*	*Single*	*Cohabitant*	*Married (no previous cohabitation)*	*Married (premarital cohabitation)*
1945-54	0.02	0.09	1	0.67
1955-64	0.05	0.15	1	0.96
1965-75	0.09	0.36	1	0.76

Source: based on the Family and Fertility Survey data for Spain

Table 6. Sensitivity of results to different values of the variance. Some examples

	FIXED VARIANCE (Model 2)		FREE VARIANCE	
	Parameter	S.E.	Parameter	S.E.
s (first birth)	1		1.44	0.18
s (union formation)	1		2.82	0.19
Correlation	0.65	0.17	0.36	0.04
Process: First birth				
Enter marriage of singles shift	3.23	0.13	3.28	0.10
Enter cohabitation shift	1.95	0.33	1.93	0.33
Enter marriage of cohabitant shift	1.17	0.26	1.25	0.22
Birth-cohort 1955-64	-0.01	0.08	0.04	0.09
Birth-cohort 1965-75	-0.37	0.09	-0.40	0.11
Process: Cohabitation				
Start of pregnancy to half of preg. (slope)	7.11	1.24	10.86	1.33
Birth-cohort 1955-64	1.14	0.25	1.47	0.28
Birth-cohort 1965-75	1.57	0.25	1.54	0.29
Process: Direct Marriage				
Start of pregnancy to half of preg. (slope)	12.54	0.45	17.44	0.67
Birth-cohort 1955-64	0.30	0.07	0.66	0.14
Birth-cohort 1965-75	-0.30	0.08	-0.35	0.15

Figure 1. First birth process: effects of duration since start of first union (cohabitation or marriage) and of marriage of cohabitants (model with unobserved heterogenity)

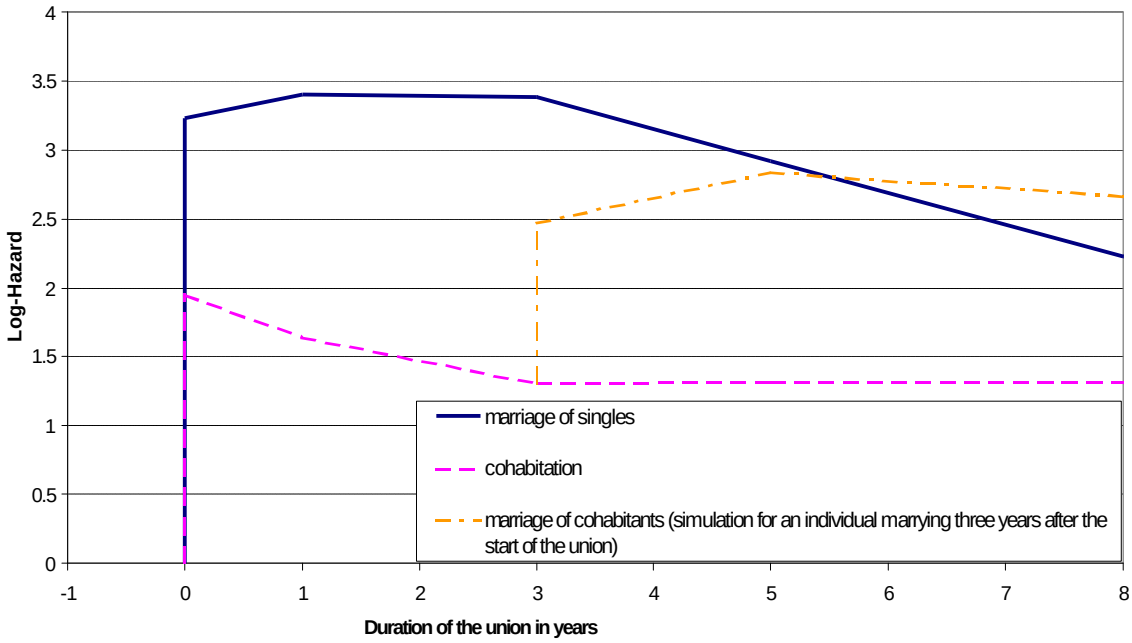

Figure 2. First union formation process: effect of pregnancy and age of first child (model with unobserved heterogeneity)