Cohabitation, Marriage, and First Birth: The Interrelationship of Family Formation Events in Spain

PAU BAIZÁN[1,*], ARNSTEIN AASSVE[2] and FRANCESCO C. BILLARI[2]
[1]*ICREA and Pompeu Fabra University, Department of Political and Social Sciences, R. Trias Fargas 25, 08005 Barcelona, Spain;* [2]*Max Planck Institute for Demographic Research, Doberaner Strasse 114, 18057 Rostock, Germany (*Author for correspondence: E-mail: pau.baizan@cpis.upf.es)*

Baizán, P., Aassve, A. and Billari, F.C., 2003. Cohabitation, Marriage, and First Birth: The Interrelationship of Family Formation Events in Spain, European Journal of Population, **19**: 147–169.

Abstract. In this paper, we investigate (1) the mutual causal relationship between first union formation and first childbirth, and (2) the existence of constant unmeasured determinants shared by these two events. We argue that these determinants mainly consist of value orientations that are heterogeneous in the population. We apply event-history techniques to retrospective survey data on Spain, allowing for unobserved heterogeneity components which simultaneously affect the two processes. Our findings confirm the existence of a strong selection effect, which influences both union formation and first birth. When controlling for these shared 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, in contrast, produces a much smaller increase in the relative risk of conception. The effect of conception on union formation is particularly strong during pregnancy, but it declines sharply after delivery.

Key words: first birth, selection effect, simultaneous hazard model, Spain, union formation

Baizán, P., Aassve, A. et Billari, F.C., 2003. Cohabitation, mariage, première naissance: relations entre événements familiaux en Espagne. Revue Européenne de Démographie, **19**: 147–169.

Résumé. Cet article s'intéresse (1) aux relations réciproques entre formation du couple et première naissance et (2) aux facteurs permanents mais non observés qui jouent en même temps sur ces deux événements. Ces facteurs pourraient essentiellement consister en des systèmes de valeur vis-à-vis desquels la population est hétérogène. Nous appliquons des techniques d'analyse des biographies à des données d'enquête rétrospective espagnoles, permettant la prise en compte de facteurs d'hétérogénéité non directement observables et qui affectent simultanément les deux processus. Nos résultats confirment l'existence d'un fort effet de sélection qui joue à la fois sur la formation du couple et sur la première naissance. Si l'on contrôle ces facteurs communs, on trouve que le risque de conception s'accroît immédiatement après le mariage et reste élevé au cours des quatre années qui suivent. Au contraire, l'entrée en cohabitation n'augmente que légèrement le risque relatif de conception. L'effet d'une conception sur la formation du couple est particulièrement élevé pendant la grossesse mais il décroît fortement après l'accouchement.

1. Introduction

Childbearing generally takes place within a union, especially in (continental) Western Europe (Kiernan, 1999). Maybe for this reason, demographers devote less attention to the relationship between union formation and childbearing than it actually deserves. More specifically, the causal (in contrast to the possibly spurious) nature of such a relationship has rarely been specifically addressed. For instance, what consequences does a later age at first union have on the timing of first birth? And what are the differences between consensual and marital unions? Does the conception of a baby have a causal impact on subsequent union formation? Is the relationship between union formation and first birth a spurious one due to the presence of different value orientations and life plans in the population? Only a few studies, which we review later, have addressed this issue from this perspective.

In this paper, we focus our attention on a situation in which first births are increasingly postponed, and in which fertility has reached very low levels. We concentrate on 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. Also, the spread of cohabitation is still limited in Spain. However, we believe that our approach has a wider relevance than the particular case of Spain, since our results imply a reassessment of the relationship between union formation events and first birth, in that they question the assertion that early marriage leads to early childbearing (Rindfuss et al., 1988), and that early pregnancy leads to early marriage (Manning, 1995). We show that the association normally found between first union formation and first childbirth is partially spurious. We also suggest an interpretation of the factors that may be responsible for this interrelationship, in line with contemporary demographic theories.

Entering parenthood and forming a first union are closely linked events, both in terms of their timing over the life course and the intentions/life plans of individuals. If a union is viewed as the appropriate setting for bearing children, individuals wanting to have a child may accelerate their union formation because they consider 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 time order of the events may therefore not always reflect a causal relationship. More generally, selection effects may arise if there is an unmeasured antecedent which simultaneously influences both processes of first childbirth and first union formation. If such shared factors (whose nature we shall discuss in more detail later) 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. In this situation, the estimated parameters in a standard hazard regression analysis will be biased and will be

unlikely to reflect the independent effect of union formation on conception leading to a first birth (and vice versa).

In order to overcome this potential bias, we use a modelling strategy which simultaneously estimates first birth and first union formation as dependent events, and which allows for the presence of unmeasured factors affecting the timing of both processes. The statistical specification is based on the approach developed by Lillard (1993), who extended the concept of simultaneous equations to include hazard or duration (failure time) models. For instance, Lillard et al. (1995) and Upchurch et al. (2002) demonstrate the importance of modelling inter-related demographic processes jointly. This technique provides a test for the existence of selection effects, by measuring the correlation in the heterogeneity components across processes. The control provided for these shared factors makes it possible to estimate the direct (or net) impact of the interrelated processes on each other. In addition, when studying the effect of first birth and first union formation, 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 model the time since the start of a pregnancy. It is important to depict the "shape" of each of these time effects. Furthermore, only when the time dependence between processes is modelled in detail is it possible to control for the interrelationship between processes with confidence. Finally, we will focus on 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 be influenced by the heterogeneity in the population composition (selection) discussed above. Therefore, controlling for these unmeasured factors should provide more reliable estimates of the effects of these variables. We use data from the Family and Fertility Survey on Spain, which was conducted in 1995. We restrict the study to include female birth cohorts only.

The paper is organized as follows. Section 2 provides a short description of some elements of the family formation process in Spain, looking in particular at the inter-cohort development of cohabitation. Then follow several hypotheses concerning the interrelationship 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 in Spain

The reduction (relative to previous birth cohorts) of first union formation and first birth intensities in 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., 2002[1]). According to the 1995 Family and Fertility Survey (FFS), 71.2 percent of the 1955-59 female birth cohort had entered a first partnership by the age of 24, while only 53.3 of the 1965–1969 female birth cohort did so. The corresponding figures for first

births 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 Martín, 1992). However, the proportion of marital first births conceived before marriage (born less than 7 months after marriage) has substantially increased for the marriage cohorts of the late 1970s and the early 1980s (Muñoz Pérez, 1991). In our sample, which is based on FFS data, around 97 percent of first births took place within a union for the female cohorts born in the 1950s and 1960s (91% within a marriage and 6% within a consensual union).

Qualitative studies conducted in Spain suggest that consensual unions are characterized by more egalitarian gender roles than marriage does and by an explicit criticism directed at the institution of marriage by the individuals involved (Alabart et al., 1988). The FFS survey contains the first representative data set for the whole of Spain that allows a detailed analysis of cohabitation. These data show that the proportion of women who started their union as cohabitees was 3.7 percent in the birth cohort 1945–1954, 9.5 percent in the birth cohort 1955–1964, and 22.9 percent in the birth cohort 1965–1974.[2] Thus, starting from an extremely low proportion in the older birth cohort, cohabitation has substantially increased as a means of initiating a union. Nevertheless, this increase is far from sufficient in compensating for the sharp decline in the cumulative proportion of individuals marrying in the cohorts born in the 1960s, as shown by the declining proportion of people in partnerships.

3. Theoretical Perspective and Hypotheses

Our theoretical perspective integrates several insights from sociology and economics on family formation, the framework of decision-making theory (Ajzen, 1991) and the life-course approach (Buchmann, 1989; Liefbroer, 1999; De Bruijn, 1999). According to the latter, individuals are involved in the achievement of several interrelated events during the transition to adulthood. These events are outcomes of processes, which interact dynamically with each other and with the different contexts in which the individual is placed.

3.1. SELECTION EFFECTS

The risk of first birth and first-union formation may be influenced by joint factors involving value orientations and attitudes towards family life.[3] An important research tradition has examined the role of these factors in explaining family formation. Certain types of value orientation, such as secularization, post-materialistic values, and women's emancipation, are important elements in the explanation of current and past demographic behaviour. Value orientations produce a selection effect, by which individuals choose consistent paths over the life course. Some authors have emphasized that the contribution of value orientations to the study of family formation is not redundant with respect to structural or economic

explanations (Lesthaeghe and Moors, 1995). Other authors have argued that this is also true for Spain (Holdsworth, 1998; Billari et al., 2001). Referring to values, Delgado Pérez and Livi-Bacci (1992) report differences on ideal family sizes by region and social class. Reher (1998) emphasizes the importance of historically long-standing cultural forms of family formation in differentiating between Northern and Southern Europe.

Moreover, focusing on pre-marital birth, recent research has found that marital and fertility behaviour is to a large extent interchangeable, suggesting that the notion of family building strategy would be appropriate (Brien et al., 1999; Bumpass, 1990). According to this approach, women make their fertility decisions in the light of their union status preferences. Although some authors claim that union formation and first birth have become less connected to each other (see for instance, Van de Kaa, 1997; Corijn and Klijzing, 2001; Beck and Beck-Gernsheim, 2002), it is clear that value orientations play a role in connecting these events to a higher degree for some individuals than for others. For instance, according to the notion of individualism, people with post-materialistic values will not only be less likely to enter a union, they will also behave more individualistically once they are in a union, postponing childbearing (Jansen and Kalmijn, 2002).

Furthermore, family formation may be embedded in a web of choices, in which young adults seek a strategic balance between their family-life goals and goals in other domains (Liefbroer, 1999). Attitudes favourable to union formation and childbearing are likely to reinforce each other (Barber et al., 2002). Conversely, attitudes towards roles that are potentially competing with family building (such as being a student or being employed) are likely to be adjusted to conform with family building attitudes (and vice versa). Attitudes toward childbearing and career have been shown to be negatively correlated (Crimmins et al. 1991). Therefore, initial heterogeneity in the population with respect to the value orientations associated with these attitudes, lead to distinct life-course patterns. In addition, parental attitudes are also important determinants of the attitudes and behaviour of young adults, via the process of social influence. Orientation towards a career or the family then results from parental views (Barber, 2000).

The life-course approach has emphasized the importance of norms in explaining the timing of demographic events (Marini, 1985), and the fact that they vary according to, for example, the regional context or the social class. There is cumulating empirical evidence about the relevance of perceived norms[4] (Ajzen, 1991) for explaining the timing of marriage, fertility, and the departure from the parental home (Billari and Liefbroer, 2001; Billari and Micheli, 1999). Norms pertaining to the age of entering parenthood, the acceptance of out-of-wedlock childbearing, and female labour-force participation are thought to be of particular relevance in lowest-low fertility countries, which are usually characterized by strong familial and social ties (Kohler et al. 2001). In addition, a growing literature has pointed out the importance of social networks in the explanation of family related behaviour, particularly in the context of fertility decline (Kohler, 2001; Montgomery

and Casterline, 1996). This approach stresses that the uncertainty associated with fertility decisions can be reduced through learning about the experience of friends, neighbours, siblings, or other network partners.

Given the lack of appropriate longitudinal (panel) data, it has not been possible to assess in a proper way the impact of these cultural factors in Spain and most European countries. For these reasons, unobserved characteristics have potentially affected and biased the results in the analyses performed so far; moreover, the role of cultural factors might have been understated in empirical analyses.[5]

Our first hypothesis derives directly from the literature cited above:

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

3.2. THE EFFECTS OF MARITAL STATUS ON FIRST BIRTH

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 shared factors.

First, economic theory predicts an increase in the risk of first birth after union formation. The union can be considered as an institution where the production of children, i.e. child-bearing and child-rearing, is 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 specialization 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 future marriage prospects.

Second, social pressure and expectations (from parents or others) related to procreation might also increase once individuals are in a union (Barber and Axinn, 1998). As a consequence of these normative expectations and the economic incentives cited above, individuals in a union may develop more favourable attitudes and intentions towards childbearing than single people do.

These arguments suggest 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 has a stronger 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 these arguments, we make the following hypothesis:

H2: *Union formation has a strong effect on the risk of first birth, which is independent of factors common to both processes. This impact is higher in the*

case of marriage than in cohabitation. Through birth cohorts, the gap between cohabitation and marriage increasingly narrows in terms of bearing children.

3.3. EFFECTS OF PREGNANCY AND THE AGE OF THE FIRST CHILD ON FIRST UNION FORMATION

Our third hypothesis concerns the effect of first childbirth on union formation. One can expect a positive and independent impact of pregnancy and first birth on first union. 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 incentive to legitimize 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 for a partner 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 birth has a strong positive effect on marriage. In a recent contribution, Brien et al. (1999) considered separately the effect of a pre-union pregnancy on cohabitation and marriage. They found that a premarital birth generally accelerates the process of entering marriage, but that this acceleration fades quickly for those who do not marry immediately after the birth. In terms of cohabitation, in contrast, the overall effect is considerably weaker.

We expect in addition the effect of pregnancy/birth to be highly time-dependent. There are several reasons for this. First, there is a time lag from the actual conception time to the moment at which women become aware of the pregnancy. 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 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 be taken around the middle of the pregnancy, i.e. when the pregnancy is less visible, than in a state of more advanced pregnancy or when the child is already born. The effect of pregnancy on first union formation should therefore be concentrated during pregnancy or shortly after delivery. 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 factors jointly affecting both processes. This impact decreases shortly after the birth of the child. A pre-union pregnancy results more often in marriage than in cohabitation.*

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 analysis are first union formation (modelled as a competing risks process) and first birth. 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 consider 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).

The timing of first birth and the timing of 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 specifically 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 shared 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, respectively.

$$\begin{cases} \ln h^B(t) = y^B(t) + \Sigma_k z_k^B(u_k + t) + \Sigma_j a_j x_j + \Sigma_i \alpha_i w_i(t) + \varepsilon \\ \ln h^C(t) = y^C(t) + \Sigma_k z_k^C(u_k + t) + \Sigma_j b_j x_j + \Sigma_i \beta_i w_i(t) + \delta \\ \ln h^M(t) = y^M(t) + \Sigma_k z_k^M(u_k + t) + \Sigma_j b_j x_j + \Sigma_i \beta_i w_i(t) + \delta \end{cases} \tag{1}$$

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 δ.[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 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 cohabitees). 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 $\{w_i(\cdot)\}$ 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 ε and δ capture unobserved heterogeneity, and are assumed to have a joint bivariate normal distribution:

$$\begin{pmatrix} \varepsilon \\ \delta \end{pmatrix} \sim N \left(\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \sigma_\varepsilon^2 & \rho_{\varepsilon\delta} \\ \rho_{\varepsilon\delta} & \sigma_\delta^2 \end{pmatrix} \right) \tag{2}$$

in which $\rho_{\varepsilon\delta}$ 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 come 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. In our analysis, we use the female sample only. Several data cleaning procedures were implemented. Twenty-one respondents were excluded from the analysis; this is 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 in which particular events occurred was missing, and therefore, these values were assigned randomly.

When studying the effect of marital status on first birth, it is useful to account for the impact of premarital pregnancies, since some marriages and consensual unions are the direct outcome of conceptions. In order to minimize the contamination of reversed causation, the dependent variable is taken as the time of birth minus eight months. In this way, the marriages or consensual unions that started 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 of conception with marriage formation or the start of cohabitation (that is, if these events take place during the same unit of time, i.e. the month), the conception can be seen as the result of a change in marital status. If both a conception and a change of marital status take place during the same month, the change can not be regarded as the direct result of pregnancy. This is because the women are most likely unaware of the pregnancy at the time. The conception may rather be seen as the result of the union. An accurate measure of the conception hazard will be obtained when these synchronized situations are counted as made by already married/cohabiting individuals.

When studying the effect of first birth/conception on first union formation, we distinguish between the impact of being pregnant (leading to first birth), of not being pregnant, and of already having a child. Union formation behaviour exhibited 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. Therefore, it is adequate that the spline for having a child starts ticking the ninth month before birth.

We used a time-varying dummy variable to control for educational enrolment, and the level of the respondent's education as an approximate measure of human capital. The latter variable is often used to test the "independence hypothesis" (Goldscheider et al., 2001; 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 tested the inclusion of other background variables in the two equations: the number of siblings of the respondent, divorce or separation of the respondent's parents, and the respondent's job experience (i.e. the cumulated duration of employment in all the jobs held up to the moment of measurement).[9] Only the variable "job experience" was kept because it complements the information on work status and educational level. That is, it better captures the opportunities individuals face in the labour market, as well as their attachment to it. Job 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

We present the main results of our analysis in Table 1 (for first birth), Table 2 (for cohabitation), and Table 3 (for marriage). For comparative purposes, we report two models: Model 1, which does not include the correlated unobserved heterogeneity components,[10] and Model 2, which does include such components. 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. SELECTION EFFECTS

In the first hypothesis of the paper, we proposed that the processes of first birth and first union formation share some unmeasured factors that influence them. These factors are 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 ($p < 0.01$). The 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 we would get a biased picture if these shared factors were not taken into account in the models.

Table 1. Hazard of first birth (conception)

Parameter	Model 1			Model 2		
	Parameter	S.E.	Relative risk	Parameter	S.E.	Relative risk
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			0.00	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=1945–1954)			1			1
1955–1964	−0.05	0.05***	0.95	−0.01	0.08	0.99
1965–1975	−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

Notes: *** = $p < 0.01$, ** = $p < 0.05$, * = $p < 0.1$. We have added in model 2 an unobserved factor influencing simultaneously first birth and first union.
§The reference category is the age group 15–18 of Model 2.
#The relative risk for each age group is calculated by exponentiating the value of the average hazard for the ages concerned.

Table 2. Hazard of first union formation (cohabitation)

Parameter	Model 1			Model 2		
	Parameter	S.E.	Relative risk	Parameter	S.E.	Relative risk
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=1945–1954)			1			1
1955–1964	1.01	0.24***	2.75	1.14	0.25***	3.13
1965–1975	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: *** = $p < 0.01$, ** = $p < 0.05$, * = $p < 0.1$. We have added in model 2 an unobserved factor simultaneously influencing first birth and first union.
§The reference category is the age group 15–18 of Model 2.
#The relative risk for each age group is calculated by exponentiating the value of the average hazard for the ages concerned.

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 which are a consequence of 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 probabili-

Table 3. Hazard of first union formation (direct marriage)

Parameter	Model 1			Model 2		
	Parameter	S.E.	Relative risk	Parameter	S.E.	Relative risk
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	9.09	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=1945–1954)			1			1
1955–1964	0.15	0.04***	1.16	0.30	0.07***	1.35
1965–1975	−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: *** = $p < 0.01$, ** = $p < 0.05$, * = $p < 0.1$. We have added in model 2 an unobserved factor simultaneously influencing first birth and first union.
§The reference category is the age group 15–18 of Model 2.
#The relative risk for each age group is calculated by exponentiating the value of the average hazard for the ages concerned.

ties of completing both events early in the life course leave the population at risk at a young age. Thus, there is indeed an overrepresentation of individuals with a low propensity to family formation at older ages. When this selection effect is accounted for, the shape of the baseline risks shows a more marked contrast according to age.
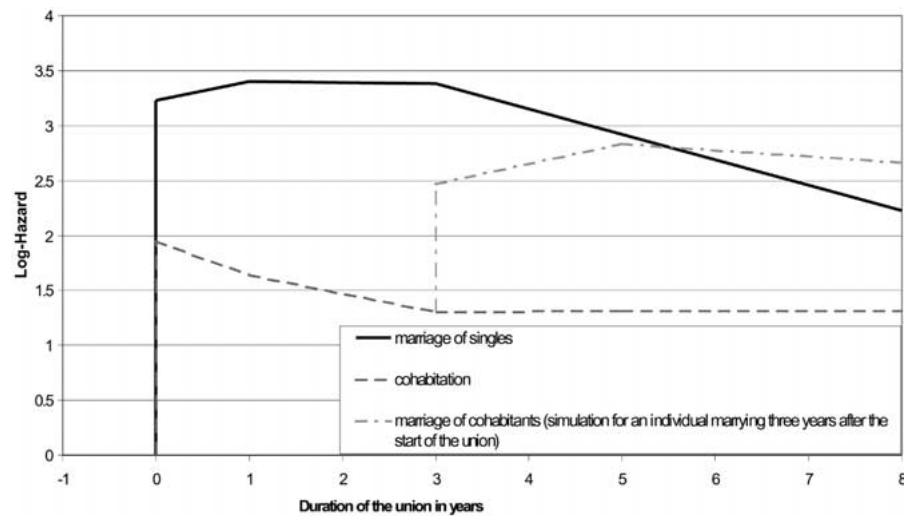
Figure 1. First birth process: effects of duration since start of first union (cohabitation or marriage) and of marriage of cohabitants (model with unobserved heterogenity).

5.2. THE EFFECTS OF MARITAL STATUS ON FIRST BIRTH

The results of Model 2, reported in Table 1 and illustrated in Figure 1, show that entering a first union sharply increases the hazard of first birth. This effect can be seen immediately after the start of the first union. The relative risk of first birth increases 25 times for direct marriage and nearly 7 times 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 the union, and it remains stable thereafter. The marriage of a cohabiting couple produces a considerable additional increase (3 times) in the relative risk of first birth.

These results provide strong support for our second hypothesis, which states that union formation has an independent effect on the risk 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 rather than in a cohabiting union, possibly due to normative influences. The first few years of the union produce the highest rates of childbearing.

The results discussed so 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 no correlated unobserved heterogeneity is taken into account, show some differences in the "shapes" of the effects of union on first birth. They indicate some over-estimation of the immediate effect of marriage formation: the coefficient for Model 1 is 3.64, while for Model 2 it is 3.23. However, after the first few months of marriage formation, the results of Model 1 show a substantial and increasingly important underestimation of the effects of marriage. A similar pattern is found for

Table 4. Effects of union status on first birth behaviour (relative risks). Birth cohort 1945–1954 as reference. Women. Model 3, with heterogeneity component

Birth cohort	Union status			
	Single	Cohabitant	Married (no previous cohabitation)	Married (premarital cohabitation)
1945–1954	1	1	1	1
1955–1964	2.07	1.38	0.80	1.14
1965–1975	1.93	1.63	0.39	0.44

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

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 under-estimated in the former model. Finally, the underestimation of the effects of the marriage of cohabitees 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 (Model 3) in which the endogenous time varying variables were not estimated as piece-wise linear splines. Instead, a time-varying covariate for marital status is used, which provides a coefficient for each cohort. The other explanatory variables remain as in Model 2, including the correlation between the heterogeneity components. The results of Model 3 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, 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 cohabitees. The birth cohort 1965–1975 presents a relative risk of 1.63 with respect to the 1945–1954 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 closer 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. In contrast to the trend for those who have

Table 5. Effects of union status on first birth behaviour (relative risks). Married with no previous cohabitation as reference. Women. Model 3, with heterogeneity component

Birth cohort	Union status			
	Single	Cohabitant	Married (no previous cohabitation)	Married (premarital cohabitation)
1945–1954	0.02	0.09	1	0.67
1955–1964	0.05	0.15	1	0.96
1965–1975	0.09	0.36	1	0.76

Source: based on the Family and Fertility Survey data on 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.
Standard deviation σ_ε (first birth)	1	–	1.44	0.18
Standard deviation σ_ε (union formation)	1	–	2.82	0.19
Correlation σ_ε	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 1955–1964	−0.01	0.08	0.04	0.09
Birth cohort 1965–1975	−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 1955–1964	1.14	0.25	1.47	0.28
Birth cohort 1965–1975	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 1955–1964	0.30	0.07	0.66	0.14
Birth cohort 1965–1975	−0.30	0.08	−0.35	0.15
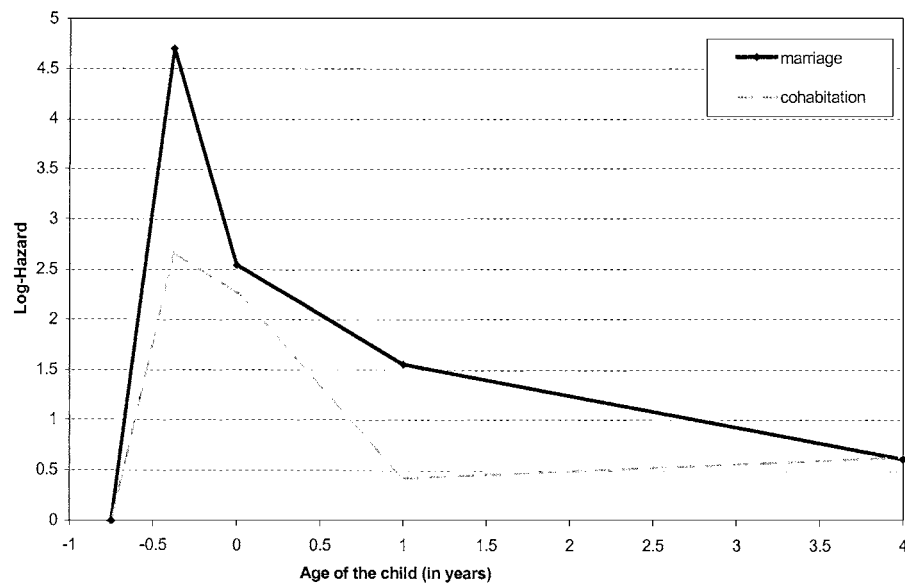
Figure 2. First union formation process: effect of pregnancy and age of first child (model with unobserved heterogeneity).

cohabited, the trend for married women with no previous experience of cohabitation (Table 4) 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 Tables 2 and 3). It can be seen that during the period of pregnancy there is a dramatic increase in the risk 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 it 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 risk of union formation. During the first year after the birth, the risk of union formation declines, implying average levels of 7.77 for marriage and 3.84 for cohabitation. After the first year of birth, the relative risk of union formation is 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-dependence 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 bear a child. This is also the case in our study. 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 have a child (Billari et al., 2001). However, being a student is much less incompatible with entering cohabitation than with marriage formation, 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 1, Model 2).

The results for educational attainment show that there is a strongly negative gradient for first birth and marriage. Higher education means higher professional career expectations, which may often be difficult to attain, given the very high unemployment rate in Spain in recent decades and the need to consolidate one's career before forming a family. It is interesting to see that this effect is stronger in Model 2, where unobserved heterogeneity is accounted for, than in Model 1, where it is not. The results for cohabitation indicate, by contrast, a *positive* gradient in the effect of the educational level, consistent with the interpretation that young people with higher education 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 behaviour.

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 female employment on family formation. On the one hand, a positive impact may be expected, since employment increases the resources necessary to form a family and because it may increase the attractiveness of individuals 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, because of the high incompatibility between working and having children in Spain. 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 or 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 studied, on the one hand, whether the timing of first childbirth and first union formation share any constant factor that influences both events, and on the other hand, the causal mutual impact of these events. The modelling strategy employed here, based on the simultaneous hazard equations approach, overcomes the bias due to the correlation between the unobserved heterogeneity affecting each process. Our findings confirm the existence of a positive and significant correlation between these heterogeneity components in Spain, indicating that 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. In this paper, we have pointed to value orientations, norms on the timing of events, and to the individual's network interactions, as possible candidates for explaining what is now unobserved. Furthermore, we would like to emphasize the importance of measuring subjective dimensions in a dynamic way, as well as the socio-economic status of individuals, in standard demographic surveys.

Our findings also show that, if the strong interrelationship between the events studied is not controlled for, 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 an 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. It can be noted that in most low-fertility countries, especially in Southern Europe, it is at these late ages where most first births take place. Therefore, a proper account of the determinants of fertility postponement should consider these stronger effects of union formation.

The results obtained also show that, net of the common heterogeneity factors, union formation and first birth have an independent impact on each other. This has a practical methodological consequence. 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 (completely) 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 formalization of the consensual union. This effect is considerably more important for marriage than for cohabitation, but it may also depend on the meaning of cohabitation and marriage in each context. In this respect, we have found indications that cohabitation plays an increasingly important role in first childbearing through birth cohorts, parallel to the (still limited) spread of cohabitation in Spain. This finding suggests that normative or other influences hampering the development of cohabitation, may have contributed to the rapid decline in fertility in countries where "new" demographic behaviours have encountered difficulties to become established. Finally, we have shown that the impact of conception leading to a first birth on union formation is mainly restricted to the period of pregnancy until shortly after delivery, while afterwards the impact is very low.

Acknowledgements

Notes

[1] The study of Billari et al. (2002) provides survival functions for the events of first union formation and first birth, involving the same groups of birth cohorts as in the present paper, and using the same survey (Family and Fertility Survey, 1995).
[2] Note, however, that these proportions are not strictly comparable, due to differences in exposure time.
[3] We follow the definition of Rokeach, who sees values as "enduring beliefs that a specific mode of conduct is personally or socially preferable to an opposite or converse mode of conduct or en-state of existence" (1973, cited in Moors, 2001). Attitudes refer to domain-specific opinions.
[4] Norms can be defined as situation-specific rules of conduct that prescribe or prohibit, that may (or may not) be followed by sanctions.
[5] Other factors that we do not discuss in detail may be part of the explanation of the unobserved heterogeneity affecting both processes, such as health status, the partner's characteristics, or psychological traits.

[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.

[8] Several experiments have been 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. Although the magnitude of effects was somewhat affected by the value of the variance, the sign and significance was not (see Table 6).

[9] The number of siblings was included in the equation for first birth only, whereas an indicator for whether parents of the interviewee had separated or divorced during the respondent's childhood was included in the equation for union formation. These variables were used 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 they were consequently dropped from the final specification. Moreover, the results of these variables were similar to those obtained in 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).

[10] Note that this model, which is a standard specification in event-history studies, also assumes a zero correlation between the endogenous "independent" variables and the heterogeneity terms (residuals).

References

Ajzen, I., 1991. 'The theory of planned behavior', *Organizational Behavior and Human Decision Process* 50: 179–211.

Alabart, A., Cabré, A. Domingo, A. Fabré, A. and Stolcke, V., 1988. *La cohabitación en España. Un estudio en Madrid y Barcelona*. Centro de Investigaciones Sociológicas, Serie Estudios y Encuestas, Madrid, 8.

Baizán P., 2001. 'Transitions to adulthood in Spain', in M. Corijn and E. Klijzing (eds), *Transitions to adulthood in Europe*. Series European Studies of Population 10, Dordrecht: Kluwer Academic Publishers, 279–312.

Barber, J. S., 2000. 'Intergenerational influences on the entry into parenthood: Mothers' preferences for family and nonfamily behavior', *Social Forces* 79(1): 319–348.

Barber, J. S. and Axinn, W. G., 1998. 'The impact of parental pressure for grandchildren on young people's entry into cohabitation and marriage', *Population Studies* 52: 129–144.

Barber, J. S., Axinn, W. G. and Thornton, A., 2002. 'The influence of attitudes on family formation processes', in R. Lesthaeghe (ed), *Meaning and Choice: Value Orientations and Life Course Decisions*. The Hague/Brussels: NIDI/CBGS Publications, 45–96.

Beck, U. and Beck-Gernsheim, E., 2002. *Individualisation: Institutionalized individualism and Its Social and Political Consequences*. Sage, London.

Becker, G., 1981. *Treatise on the Family*. Harvard University Press, Cambridge, MA.

Billari, F. C., Castiglioni, M., Castro Martín, T., Michielin, F. and Ongaro F., 2002. 'Household and union formation in a Mediterranean fashion: Italy and Spain', in E. Klijzing and M. Corijn (eds), *Comparative Research on Fertility and the Family in Contemporary Europe: Findings and Lessons*. United Nations, New York/Geneva.

Billari, F. C. and Liefbroer, A. C., 2001. "Should I stay or should I go? The impact of age norms on leaving home". Paper presented at the Annual Meeting of the Population Association of America, Washington D.C.

Billari, F. C. and Micheli, G. A., 1999. 'Le scelte demografiche. Le percezione dei costi e delle norme sociali', in L. Mauri and F. C. Billari (eds), *Generazioni di donne a confronto. Indagine sociodemografica*. Franco Angeli, Milano, 163–190.

Billari, F. C., Philipov D. and Baizán P., 2001. 'Leaving home in Europe. The experience of cohorts born around 1960', *International Journal of Population Geography* 7(5): 339–356.

Brien, M. J., Lillard, L. A. and Waite, L. J., 1999. "Interrelated family-building behaviors: Cohabitation, marriage, and nonmarital conception", *Demography* 36(4): 535–551.

Buchmann, M., 1989. *The Script of Life in Modern Society. Entry into Adulthood in a Changing World*. The University of Chicago Press, Chicago/London.

Bumpass, L. L., 1990. 'What's happening to the family? Interactions between demographic and institutional change', *Demography* 27(4): 483–490.

Castro Martín, T., 1992. 'Delayed childbearing in contemporary Spain: Trends and differentials', *European Journal of Population* 8: 217–246.

Cigno, A. and Ermisch, J., 1989. 'A microeconomic analysis of the timing of births', *European Economic Review* 33: 737–760.

Corijn, M. and Klijzing, E. (eds), 2001. *Transitions to Adulthood in Europe*. European Studies of Population, 10. Kluwer Academic Publishers, Dordrecht.

Crimmins, E. M., Easterlin, R. A. and Saito, Y., 1991. 'Preference changes among American youth: Family, work and goods aspirations, 1976–1986', *Population and Development Review* 17(1): 115–133.

De Bruijn, B. J., 1999, *Foundations of Demographic Theory. Choice, Process, Context*. NethurD Publications, Thela Thesis, Groningen.

Delgado Pérez, M., 1994. 'Cambios recientes en el proceso de formación de la familia', *Revista Española de Investigaciones Sociológicas* 64: 123–153.

Delgado Pérez, M. and Livi-Bacci, M., 1992. 'Fertility in Italy and Spain: The lowest of the world', *Family Planning Perspectives* 24(4): 162–171.

Delgado, M. and Castro Martín, T., 1999. *Fertility and Family Surveys in Countries of the ECE Region. Standard Country Report. Spain*. United Nations, New York and Geneva.

Eurostat, 1996. *Eurostat Yearbook '96. A Statistical View of Europe 1985–1995*. Eurostat, Luxembourg.

Goldscheider, F. and Waite, L., 1986. 'Nest-leaving patterns and the transition to marriage for young men and women', *Journal of Marriage and the Family* 49: 507–516.

Goldscheider, F., Turcotte, P. and Kopp, A., 2001. 'The changing determinants of women's first union formation in industrialized countries: The United States, Canada, Italy and Sweden', *Genus* 57(3–4): 107–134.

Holdsworth, C., 1998. 'Leaving home in Spain: A regional analysis', *International Journal of Population Geography* 4: 341–360.

Jansen, M. and Kalmijn, M., 2002. 'Investments in family life: The impact of value orientations on patterns of consumption, production and reproduction in married and cohabiting couples', in R. Lesthaeghe (ed), *Meaning and Choice: Value Orientations and Life Course Decisions*. NIDI/CBGS Publications, The Hague/Brussels, 129–159.

Keeley, M. C., 1977. 'An analysis of the age pattern of first marriage', *International Economic Review* 20: 527–545.

Kiernan, K., 1999. 'Childbearing outside marriage in Western Europe', *Population Trends* 98: 11–20.

Kohler, H.-P., 2001. *Fertility and Social Interactions: An Economic Perspective*. Oxford University Press, Oxford.

Kohler, H.-P., Billari, F. C. and Ortega, J. A., 2001. 'Towards a theory of lowest-low fertility', Working Paper WP-2001-032. Max Planck Institute for Demographic Research, Rostock.

Kravdal, Ø., 1994. 'The importance of economic activity, economic potential and economic resources for the timing of first birth in Norway', *Population Studies* 48: 249–267.

Lancaster, T., 1990. *The Econometric Analysis of Transition Data*, Econometric Society Monographs 17. Cambridge University Press, Cambridge.

Lesthaeghe, R. and Moors, G., 1995. 'Living arrangements and parenthood: Do values matter?', in R. De Moor (ed), *Values in Western Societies*. Tilburg University Press, Tilburg.

Liefbroer, A. C., 1999. 'From youth to adulthood: Understanding changing patterns of family formation from a life course perspective', in L. J. G. Van Wissen and P. A. Dykstra (eds), *Population Issues. An Interdisciplinary Focus*. Kluwer Academic/Plenum Publishers, New York.

Lillard, L. A., 1993. 'Simultaneous equations for hazards. Marriage duration and fertility timing', *Journal of Econometrics* 56: 189–217.

Lillard, L. A., Brien, M. J. and Waite, L. J., 1995. 'Premarital cohabitation and subsequent marital dissolution: A matter of self-selection?', *Demography* 32(3): 437–457.

Lillard, L. A. and Panis, C. W. A., 2000. *aML Multilevel Multiprocess Statistical Software, Release 1.0*. EconWare, Los Angeles, CA.

Manning, W. D., 1995. 'Cohabitation, marriage, and entry into motherhood', *Journal of Marriage and the Family* 57(1): 191–200.

Marini, M. M., 1985. "Determinants of the timing of adult role entry", *Social Science Research* 14: 309–350.

Miret-Gamundi, P., 1997. 'Nuptiality patterns in Spain in the eighties', *Genus* 53(3/4): 183–198.

Montgomery, M. R. and Casterline, J. B., 1996. 'Social influence, social learning, and new models of fertility', *Population and Development Review* 22(Suppl.): 151–175.

Moors, G., 2001. 'Family theory: The role of changing values', in N. J. Smelser and P. B. Baltes (eds), *International Encyclopedia of the Social and Behavioral Sciences*. Pergamon, Oxford.

Mulder, C. H. and Manting, D., 1993. 'Strategies of nest-leavers: "Settling down" versus flexibility', *European Sociological Review* 10(2): 135–154.

Muñoz Pérez, F., 1991. 'Les naissances hors marriage et les conceptions prénuptiales en Espagne depuis 1975: Une période de profonds changements', *Population* (4): 881–912.

Reher, D. S., 1998. 'Family ties in Western Europe: Persistent contrasts', *Population and Development Review* 24: 203–234.

Rindfuss, R. R., Morgan, S. P. and Swicegood, G., 1988. *First Births in America: Changes in the Timing of Parenthood*. University of California Press, Berkeley.

Rokeach, M., 1973. *The Nature of Human Values*. Free Press, New York.

Upchurch, D. M., Lillard, L. A. and Panis, C. W. A., 2002. "Influences on nonmarital childbearing", *Demography* 39(2): 311–329.

Van de Kaa, D. J., 1997. *Options and Sequences. Europe's Demographic Patterns*. Nethur-Demography Paper 39, The Hague.