DEMOGRAPHIC RESEARCH

Research Article

Divorce risk factors and their variations over time in Spain

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Abstract

The aim of this article is to analyse the risk factors for marriage dissolution in Spain and their variations over time for women married between 1949 and 2006. The data are drawn from the Survey of Fertility, Family, and Values of 2006. The article analyses the transition from first marriage to marital dissolution for couples who married in two eras: the period prior to the enactment of the Divorce Law of 1981, when there were many social and legal barriers to marriage dissolution; and the period after the law went into effect, when there were far fewer barriers to marriage dissolution. The analyses were conducted using a continuous time event history analysis. The results indicated that there are similarities between Spain and other countries, such as the positive relationship between the typical features of unconventional families and marital dissolution; but that there are also certain differences, such as an increase in the importance of premarital pregnancy and/or of not having children. It is also essential to stress the declining importance of socioeconomic variables, such as education and the labour market situation for women.

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1. Introduction

The rise in divorce rates has become a key phenomenon underlying significant changes in family structure and in patterns of social stratification in contemporary societies. A large body of evidence on determinants of divorce is now available for the United States (Rogers and DeBoer 2001; South 2001; Teachman 2002; Raley and Bumpass 2003), Northern and Eastern Europe (Jalovaara 2002; Walke 2002; Lyngstad 2004, 2006; Fischer and Liefbroer 2006; De Graaf and Kalmijn 2006), Central Europe (Diekmann and Engelhardt 1999), and Asia (Raymo, Iwasawa, and Bumpass 2004; Yi Zeng et al. 2002). Various studies have also addressed the question of whether the determinants of divorce have changed over time (Hoem 1997; Teachman 2002; Chan and Halpin 2005; De Graaf and Kalmijn 2006; Härkönen and Dronkers 2006). Much less is known, however, about divorce in Southern Europe, particularly in Spain, where union dissolution has traditionally been a rare event, but has increased considerably in recent years. In 2008, the crude divorce rate in Spain was 2.4‰, which is much higher than in other Southern European countries, and is similar to or even higher than in many countries in Central and Northern Europe.[3] However, the only available detailed studies on divorce in Spain are either based on old data or do not explicitly address the question of whether the determinants of divorce have changed over time (Härkönen and Dronkers 2006; Simó Noguera and Solsona 2007; Coppola and Di Cesare 2008). The aim of this article is, therefore, to fill this gap in the literature and to produce evidence on divorce patterns in Spain and their variations over time.[4]

Compared to other Western countries, divorce became legal in Spain at a very late date. Until 1981, divorce was forbidden by Spanish law, and could be achieved only through an annulment issued by the Catholic Church, which was granted only under special conditions, and following a very costly procedure. However, de facto separation was possible by "fleeing" the household. Not surprisingly, only a small number of unions formed before 1960 have ended in permanent dissolution (approximately 3%). For unions formed after 1980, however, the risk of dissolution has increased considerably, and one in six unions now ends in divorce (see Figure 1 below). Moreover, in the last several decades, Spanish society has also changed significantly with regard to those factors identified by previous studies in other countries as key determinants of divorce. Only 30% of women born in Spain in the first half of the

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[3] Part of the increase is due to the Divorce Law of 2005, which eliminated legal separation as a prior step to divorce. Statistics before 2006 did not include legal separations as divorces, so it is likely that the divorce statistics of Spain were underestimated before that year. See Eurostat: http://epp.eurostat.ec.europa.eu/tgm/ table.do?tab=table&init=1&language=en&pcode=tps00013&plugin=1.

[4] We use the terms divorce, marriage dissolution, disruption, and breakup interchangeably throughout the paper. Divorce also includes permanent separation.

1950s attended school beyond the primary education level, whereas the corresponding figure for those born in the second half of the 1970s was almost 70%. Between 1980 and 2008, the average age at first marriage increased from 24 to 30 for women, and from 26 to 32 for men. The total fertility rate fell from 2.2 to 1.5. Cohabitation was not common in the past, but today, 20% of Spanish women between the ages of 30 and 34 who are living with a partner are not married. Moreover, the employment rate of women between the ages of 30 and 34 has now reached 70%, which is 40 percentage points higher than the employment rate for women of that age group in the mid-1970s.[5] In sum, as a result of the rapid modernisation of Spanish society, both the rate of divorce and the distributions of those factors usually associated with the risk of divorce have changed significantly. It seems, therefore, interesting and timely to investigate whether the correlates of divorce have also changed over time.

Another important objective of our analysis is to determine whether the increase in divorce risk is driven by a cohort effect or by a period effect. As Lyngstad and Jalovaara (2010) have noted, different cohorts can be associated with different expectations and resources, which can in turn translate into higher divorce rates. For instance, most women born from the 1960s onwards have pursued careers and are more open-minded than older cohorts, and these differences may have an impact on divorce. Increased divorce risk can also be driven by a period effect. In this case, several factors related to divorce can affect all of the cohorts at the same time. The legalisation of divorce in Spain in 1981 is a good example of a significant factor that affected all of the cohorts.

The structure of the paper is as follows: we present the hypotheses that guide the research in the second section; we describe the data, methods, and variables in the third section; we comment on the findings in the fourth section; and we draw our conclusions in the final section.

2. Theories and hypotheses

In this section, we outline the hypotheses on divorce risk factors and how their effects might have varied over time in Spain. We focus on socioeconomic attributes, the presence of small children, and unconventional characteristics related to union formation or previous family experience (Wagner and Weiß 2006). We also discuss whether changes over time are more properly interpreted as cohort or period effects.

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[5] The data were drawn from the Spanish Labour Force Surveys, except for the average age for first marriages and the total fertility rate, which were provided by the National Statistics Institute (see: www.ine.es).

2.1 Socioeconomic determinants

The classic hypothesis regarding the effects of education on divorce is derived from the theory formulated by Goode (1951, 1962). When divorce is complicated—or, in extreme cases, banned by legal and administrative impediments—it becomes a privilege that is predominantly restricted to an elite group who are able to cover the costs of working around the legislation. This situation is typical of pre-industrial or early industrialised societies, where a higher incidence of divorce was found among the upper classes. However, when legal and administrative barriers to divorce are lifted, divorce becomes accessible to all social classes. It is then argued that the economic strain that might lead to marital discord, and, eventually, divorce, is more common among the lower classes (Goode 1951). Moreover, the most educated have access to more cultural, social, and cognitive resources, which enable them to choose a better mate and to use their communication skills to resolve conflicts in the relationship (Amato 1996; Hoem 1997; Ono 1998). It is then predicted that the relationship between class and divorce reverses, and that, in industrial societies, a higher incidence of divorce will be observed at the bottom of the class structure (Goode 1962, 1993; Härkönen and Dronkers 2006).

A similar hypothesis of a reversal of the gradient of education on the risk of divorce over time was developed by Blossfeld et al. (1995) and Hoem (1997), although their emphasis was on social norms and social diffusion of innovative behaviours. Thus, they argued that divorce is a stronger deviation from social norms in a low-divorce context than it is in a high-divorce context. The social costs of divorce are higher in a society where divorce is a rare event. In this context, however, education can have a "liberating effect" from normative pressures (Blossfeld et al. 1995) because highly educated women are less traditional and are thus more willing to dissolve an unhappy union, even if doing so violates social norms (Levinger 1979). However, as divorce gradually becomes more common, it is argued that the stigma associated with divorce declines, and that the positive effect of education on divorce should, therefore, decrease.

Based on the arguments of Goode (1962, 1993), and Blossfeld et al. (1995), we might expect to find that the relationship between education and divorce is positive in the older marriage cohorts in Spain. In fact, prior to the enactment of the Divorce Law of 1981, the economic and social costs associated with marital disruption were very high. But with the spread of divorce and the reduction in its social and economic costs, we might anticipate that the relationship between education and marriage dissolution will have weakened after 1981.

With regard to the effect of women's employment, various studies have found that employed women are more likely to divorce than those who are inactive (South 2001). Working women are able to bear the economic costs of divorce because they receive a salary and are less economically dependent on their husbands. Moreover, the chances of

meeting a new partner are higher for employed women, given that employment broadens the range of social contacts (Udry 1981; South and Lloyd 1995). Other studies have suggested that female employment calls into question the traditional division of labour within the household. The advantages of a gendered specialisation in paid and unpaid work are lost, and the interdependence of spouses is reduced (Becker 1981; Parsons 1949). Female employment is then associated with an increase in union dissolutions.

Very few analyses have investigated the impact of a wife's unemployment on divorce, although there is some evidence of a positive relationship between the two factors in Scandinavia (Jalovaara 2003; Hansen 2005). Unemployment usually generates greater stress for a couple, which facilitates marital breakdown. Moreover, unemployment status might be an indicator of an anticipated or expected union dissolution. Inactive women, when faced with an unsatisfactory marital relationship, might decide to start looking for a job in anticipation of a breakdown of the union. In this case, the causal relationship would go from the decision to divorce to unemployment.[6] However, given the limited coverage of unemployment benefits in Spain, unemployed women are less economically independent than employed women. According to the fourth quarter of the 2009 Spanish Labour Force Survey, only 31% of unemployed women received unemployment benefits. Because their economic independence is limited, the risk of divorce for unemployed women should be lower than that of employed women.

With respect to changes in these effects over time, Goode's argument should also apply to the effect of female employment. When union separation was rare, economic resources were crucial for covering the legal costs associated with divorce, and for starting an independent life. It has also been argued that the positive effect of women's economic independence on divorce is stronger in more traditional settings, such as Spain in the 1950s and 1960s (Kalmijn, De Graaf and Poortman 2004). Moreover, Becker's and Parsons' arguments that working women had a disruptive effect on unions referred to societies in which the model of the male breadwinner was dominant. Given the higher costs of divorce and the stronger mismatch between traditional values and female employment when relatively few women were employed and the breadwinner model was largely dominant, we might expect to find that the positive effect of female employment on divorce was stronger in the past, and that it has declined over time.

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[6] The analysis undertaken by Poortman (2005) suggests, however, that the evidence for anticipatory behaviour in the case of women's work and divorce is weak.

2.2 Children

Previous studies have consistently shown that couples who have children, especially young children, are less likely to divorce than couples who do not have children (Walke 2002; Wagner and Weiß 2006). On the one hand, having children can be viewed as the specific capital of a couple. According to New Home Economics, the benefits from that kind of specific capital are greater if the parents live together (Becker, Landes, and Michael 1977). Moreover, social psychology has demonstrated that having children increases the marital commitment; thus, it is less likely that the parents will divorce (Brines and Joyner 1999). It may also be the case that spouses who are not confident of the durability of their marriage are less likely to have children (Lyngstad and Jalovaara 2010).

With regard to variations over time, Wagner and Weiß (2006) have hypothesised that modernisation has reduced the effect that children have on marital disruption. Their hypothesis is based on the notion that the economic risks of living as a single parent have decreased. In countries where a developed welfare state exists, there are mechanisms that make living as a single parent less difficult, such as the option of accessing alimony and maintenance support programs. Moreover, it has been argued that the normative climate that has traditionally encouraged parents with children to stay together has weakened (Lyngstad and Jalovaara 2010). In the case of Sweden, for example, there is evidence that the protective effect of children on the risk of divorce has actually attenuated over time (Anderson 1997).

2.3 Unconventional characteristics

Among the unconventional characteristics related to union formation or previous family experience are age heterogamy, parental divorce, previous cohabitation, premarital pregnancy, and having children by someone other than the current spouse. Our interest in these attributes mainly lies in the unobserved processes of self-selection and their consequences for divorce (Lillard, Brien, and Waite 1995). Self-selection occurs when a predictor of divorce is itself the result of an individual choice based on some unobservable characteristic that might be positively associated with the risk of divorce.[7]

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[7] Unobserved self-selection might also occur with regard to marriage. The problem of self-selection into marriage might be seen as a problem of sample selection bias. Sample selection occurs when unobserved factors that influence the process of interest (marriage dissolution in our case) help to determine whether that process (marriage) is observed in the first place (Breen 1996). In a companion paper (Bernardi and Martínez-Pastor 2010), we have examined in detail the implications of selection into marriage with regard to the effect of education on marriage dissolution. We do not consider this type of process here.

In the case of prior cohabitation, it has been argued that, when cohabitation is rare, those who cohabitate before marrying are already a select group characterised by unconventional values and a lesser degree of commitment to marriage than those who marry without cohabitating first (Hall and Zhao 1995). In other words, the decision to cohabit prior to marriage is driven by certain values that are unobservable or difficult to observe, and which also make subsequent divorce more likely. However, when cohabitation is more common, and people decide to cohabit for reasons that go beyond an adherence to unconventional values, those who cohabit before marriage are likely to become more internally heterogeneous. The effect of prior cohabitation on divorce due to self-selection should, then, diminish. Some authors have also argued that, given the increasing heterogeneity of marriage markets and the growing individualisation of societies, it is necessary to spend more time searching for a partner. As a consequence, cohabitation has been converted into a prerequisite for a stable marriage (Wagner and Weiß 2006). In this sense, cohabitation functions as a test period to ensure that a good selection has been made. If this is the case, then previous cohabitation should reduce the probability of marital disruption.

The evidence of a change in the effect of previous cohabitation over time is limited. Liefbroer and Dourleijn (2006) have, however, shown that cohabitation before marriage increases the likelihood of divorce only in countries where premarital cohabitation is either very extended or very limited. The argument is as follows: when cohabitation is rare, those who cohabit before marriage are a highly select group whose non-traditional attitudes might also favour union dissolution; whereas when cohabitation is common, those who marry without a previous experience of cohabitation are also a select group whose traditional values make them less likely to divorce. Cohabitation without being married has traditionally been a rare event in Spain. Whereas 46% of French women and 74% of Swedish women born between 1960 and 1965 had experienced non-marital cohabitation by the age of 25, the proportion for Spanish women of the same era is only 7% (Kiernan and Estaugh 2002). From a cross-sectional perspective, in 1995 only 5% of couples who lived together were not married (Spanish Fertility Survey of 1995). That number doubled in the 10 years after that survey (Spanish Fertility Survey of 2006). Given the comparatively low proportion of cohabitants, women who have cohabited prior to marriage can be considered a select group based on certain unobserved characteristics that might be positively associated with the risk of union dissolution (Lillard, Brien, and Waite 1995). Thus, we expect to find a positive association between premarital cohabitation and divorce.

With respect to the intergenerational transmission of divorce, it has been shown that, due to socialisation, children of divorced parents inherit attitudes that are not supportive of marriage and family, which makes them more prone to marital breakdown (Amato 1996; Diekmann and Engelhardt 1999; Amato and DeBoer 2001). Some

research has also confirmed that children of divorced parents usually marry early (Thornton 1991), which tends to encourage divorce in at least two ways. First, those who marry early have dedicated less time to the search for a partner, which increases the likelihood that they have made a bad selection. Second, in the case of divorce, they still have a broader remarriage market available to them than those who delay marriage. It has also been shown that children of divorced parents cohabit more frequently and are less likely to attain socioeconomic stability, factors that are positively associated with divorce (Diekmann and Engelhardt 1999). With regard to changes over time, the main hypothesis is that, with the diffusion of divorce, children of divorced parents become a less select group, and the negative consequences of parental divorce should diminish. The intergenerational transmission of divorce should also decline. The evidence supporting this hypothesis is, however, mixed (Lyngstad and Jalovaara 2010).

The great majority of couples are similar in age. Wives are seldom older than their husbands, and husbands are rarely more than five years older than their wives (see Table 1 for Spain). In addition, despite the recent increase in couples who cohabit in Spain, the majority of births continue to take place within marriage. Heterogamous marriages and premarital pregnancy can also be considered uncommon behaviours. Based on the self-selection argument previously discussed, it might therefore be expected that couples with heterogamy of age, and those who have had children before marrying, have a greater probability of divorcing (Lyngstad 2006; Kalmijn and Poortman 2006).

To summarise, deviations from the most common patterns of behaviour can be interpreted as signals of non-traditional attitudes and values. We can therefore expect to find that they are also associated with a higher risk of divorce. However, parallel to the rise in divorce, the incidence of age heterogamy, parental divorce, previous cohabitation, and premarital pregnancy with the same or another partner, have also increased to the point where they have, to a certain extent, become common traits. As these attributes have become less unconventional, the unobserved self-selection of them, based on characteristics that also make a couple more prone to divorce, should also decrease. This leads to the general hypothesis that their observed effect on divorce has declined over time.

2.4 Cohort or period effects?

Changes over time in the effects of the determinants of divorce have traditionally been investigated as birth or marriage cohort effects (Diekmann and Engelhardt 1999; De Graaf and Kalmijn 2006; Lyngstad 2006; Hall and Zhao 1995; Härkönen and Dronkers 2006; Vignoli and Ferro 2009). However, in the case of Spain, the ban on divorce was

not lifted until 1981. This belated legal change might have also affected earlier marriage cohorts who finally had the formal right to end unsatisfactory relationships. We might ask whether social changes in the risk of divorce, if any, have come about as period or cohort effects. In the former case, a significant discontinuity in divorce patterns should exist between the pre- and post-1981 periods, regardless of the date of the marriages. If, on the other hand, the dynamics of change have been more gradual over time, the great divide should be observed across marriage cohorts, more so than before and after the passage of the Divorce Law. The date of marriage should, therefore, matter, indicating that those who married in the 1960s and 1970s remain less likely to divorce, even after the legalisation of divorce in 1981.

3. Data, methods, and variables

We draw on data from the Fertility, Family, and Values Survey (FFVS) collected in 2006 by the Centre for Sociological Research. The FFVS is a nationally representative survey of women aged 15 and above in Spain. It provides detailed retrospective information on marriage and fertility histories. The original sample size was 9,737. We excluded from the analyses 890 women who were older than 75 to avoid possible bias due to differential mortality. The remaining sample consisted of 8,847 cases. Next, we eliminated cases for which it was impossible to reconstruct the matrimonial or labour history due to errors in the recollection of dates. After eliminating cases with invalid information, we were left with a sample of 6,559 women (4,409 ever married). We were, however, concerned about the implications of analysing the data with such a high number of missing cases. Given the large number, we decided to analyse the data without using the labour trajectory of individuals. This strategy notably reduces the amount of missing data (n=8,759, 6,246 ever married). The results using the two sub-samples showed the same trends for the main variables (Bernardi and Martínez-Pastor 2010). We also compared the distributions of the variables for education and civil status in the original sample and the analytic sample with valid information on marital and employment history, and with data from the Spanish Labour Force Survey in 2006 (see Appendix Table A1). Moreover, we compared the incidence of divorce or separation by level of education in the original sample, the analytical sample, and the Spanish Labour Force data (Appendix Table A2). No major differences were found.

In this article, we focus on the risk of dissolution of first marriages. We define dissolution as both divorce and separation. Approximately 68% of respondents (n=4,409) entered a first marriage, and 7.7% (n=339) of respondents subsequently saw their unions end in dissolution. We studied the risk of marriage dissolution using continuous time event history models. Completed spells were measured by the duration

in months between the date of marriage and the date of divorce or separation. Right-censored spells were measured by the duration between the date of marriage and the time of the interview for those marriages that had not ended in divorce, or between the date of marriage and the date of the death of the spouse for widowed women. In the multivariate analyses, we used a piecewise constant exponential model (pcem). This model allows for more flexibility than other specifications, such as the Weibull and Gompertz model, and does not require any assumptions concerning the form of the time dependence of the process (the transition to marital disruption). We split the time axis into nine periods, from having married less than two years ago, to having married 16 or more years ago; that is, one interval every two years. PCEM assumes that transition rates are constant in each of these periods, but can change between them.

Table 1: Descriptive statistics

	Entire populat.	Married (All cohorts)	Marriage cohort 1949-1980	Marriage cohort 1981-2006
Level of education				
Primary or less	30.3	39.1	63.0	11.9
Lower secondary	26.2	24.9	20.0	30.4
Upper secondary	26.9	22.9	11.0	36.4
University	16.5	13.2	6.0	21.3
Parental divorce				
Yes	7.1	4.5	2.4	6.9
No	92.9	95.5	97.6	93.1
Age of marriage (mean)		24.1	23.2	25.2
Parity and age of youngest child (at the time of censoring)				
None		8.3	3.3	13.7
Parity 1, youngest 0-6		7.8	0.9	15.5
Parity 1, youngest 7 +		11.6	10.2	13.2
Parity 2, youngest 0-6		11.0	1.7	21.3
Parity 2, youngest 7 +		61.3	83.9	36.3
Employment (at marriage time for those married)				
Inactive	70.6	56.8	71.4	40.0
Unemployed	5.4	7.9	4.3	12.1
Employed	24.0	35.2	24.3	47.9
Age heterogamy				
Homogamy		60.8	61.6	59.9
Husband 5 years older		25.7	27.7	23.5
Wife older		13.4	10.6	16.7
Premarital cohabitation				
Yes		10.3	3.3	18.2
No		89.7	96.7	81.8

Table 1:   (Continued)

	Entire populat.	Married (All cohorts)	Marriage cohort 1949-1980	Marriage cohort 1981-2006
Premarital pregnancy				
Yes		6.2	5.2	7.4
No		93.8	94.8	92.6
Children with partner other than spouse				
Yes		1.7	1.1	2.3
No		98.3	98.9	97.7

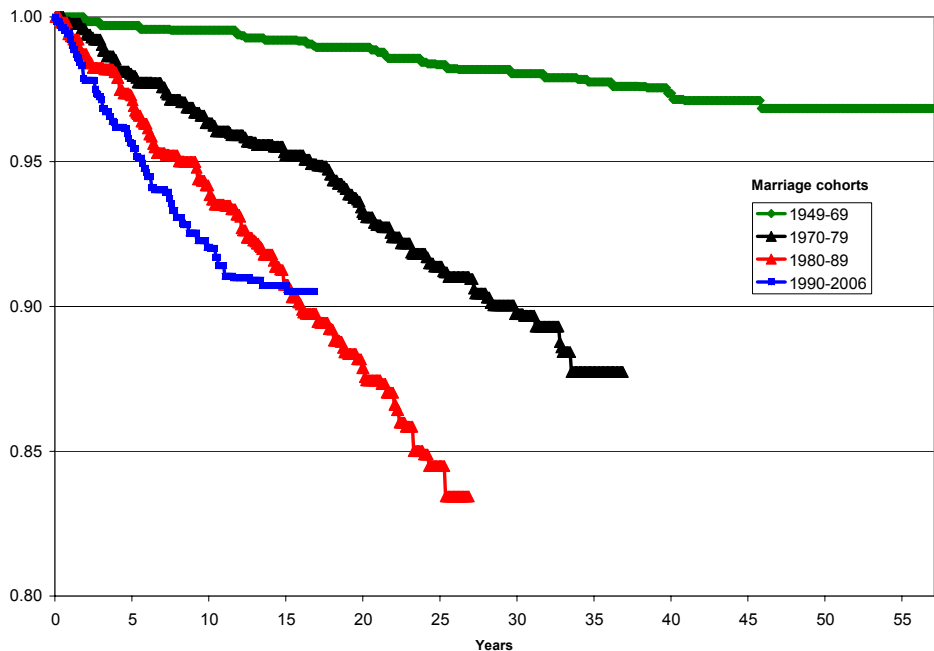*Source*: Spanish Fertility, Family and Values Survey 2006.

The other independent variables included in the analyses are education, marriage year, age heterogamy, employment status, age at marriage, and number and age of children. Education was coded into four categories: primary education or less (reference category), lower secondary education, upper secondary education, and university education. Age heterogamy with the spouse was coded into three categories: couples in which the husband is at least five years older than the wife, couples in which the wife is older than the husband, and homogamy (reference category). Employment status was coded into three categories: inactive (reference category), unemployed, and employed. Employment status was a time-varying variable. Age at marriage was a continuous variable that indicated the age (in years) at which the marriage was contracted. The number and ages of children were measured with five time-varying categories that refer to the parity and the age of the youngest child: no children (reference category); one child younger than age six; one child older than age six; two or more children, of whom one is younger than age six; and two or more children older than age six (Lyngstad 2006). The number and ages of children were also time-varying variables. The following variables were also included in the models as dummy variables: parental divorce (yes=1), pre-marital cohabitation (yes=1), pre-marital pregnancy (yes=1), and children from a different previous relationship (yes=1). Table 1 presents descriptive statistics for the independent variables.

Finally, the marriage year was rescaled so that 1949, the year of the first observed marriage in the sample, was equal to zero. We also defined four marriage cohorts: marriages that started between 1949 and 1969, those that started between 1970 and 1980, those that started between 1981 and 1990, and, finally, those that started between 1991 and 2006. The period effect was identified using a time-varying dummy that received a value of one from the year 1981.

Figure 1 shows Kaplan–Meier estimators of the survival functions by marriage duration for the different marriage cohorts. This figure shows that the probability of marital dissolution was extremely low for the marriage cohorts of the 1950s and 1960s,

and that it started to rise rapidly for those married in the 1970s. It continued to increase for those who married in the 1980s and 1990s. At the end of the observation window, 13% of women who married in the 1970s were separated or divorced. For those married in the 1980s, the cumulative proportion of marital dissolution increased to 17%.

Figure 1: **Kaplan-Meier survival function. Transition from being married to marital dissolution (Marriage cohorts between 1949 and 2006)**

Source: Spanish Fertility, Family and Values Survey of 2006.

In some of our analyses, we have reduced the four cohorts to two, distinguishing those married before 1981 from those who married in 1981 or thereafter, in order to test whether the risk factors for marriage dissolution have varied over time. Although it might be preferable to define more marriage cohorts and keep the cohort 1970-1980 as distinct, the small total number of divorces in the sample (N=339) prevents us from doing so. Nonetheless, this twofold distinction, which places the marriage cohorts between 1970 and 1980 and the older cohorts in a single category, allows us to make a comparison that is theoretically interesting and empirically sound. First, although the Divorce Law of 1981 certainly lowered the social and economic costs of divorce for

women who married between 1970 and 1980, they did not know that divorce would eventually become legal when they married. Still, they can be considered the "frontrunners" in Spain in their experience of marital dissolution because the rate of union dissolution in the previous cohorts had been almost zero. Moreover, we performed additional analyses that focused only on the comparison between the 1970-80 cohort and the subsequent cohorts. These analyses showed that divorce determinants differed substantially between the marriage cohorts that occurred between 1970 and 1980, and those that occurred after 1980 (see Appendix Table A3). These findings support the decision to consider marriages that took place in the 1970s as distinct from those that took place after 1980 in the analysis of trends over time.

4. Results

Table 2 illustrates an attempt to test whether the change in union disruption over time in Spain was mainly due to period effects as a result of the liberating effect of the Divorce Law in 1981, or whether it was due to differences across cohorts. The variable "1981 onward" in model 1 measured the period effect. The cohort effect was captured by a continuous variable "year of marriage: 1949=0" (model 2), and, alternatively, through a set of dummy variables: "marriage cohort 1949-69, 1970-80, 1981-90, and 1991-2006" (model 3). In the last two models, we jointly specified the period effect, the effect of the year of marriage (model 4), and the four cohort dummies (model 5). We assessed the fit of these different specifications using the Akaike Information Criterion (AIC), with the model with the lowest AIC value being the preferred one (Burnham and Anderson 2002). The period effect was shown to have a clear impact on the likelihood of divorce. Since 1981, the risk of union dissolution has been three times higher [exp (1.11)=3.03] than that for marriages occurring before 1981 (see model 1). Strong differences were also found across marriage cohorts. Model 3 showed that the risk of divorce for women married between 1949 and 1969 was approximately six times lower [exp=-1.80=0.16] than for women married between 1981 and 1990 (reference category), whereas for women born between 1970 and 1980, the risk of divorce was 0.70 that of women born between 1981 and 1990. However, after the period and cohort variables were introduced simultaneously in models 4 and 5, the effect of the period variable was greatly reduced, and was no longer statistically significant. Moreover, a comparison of the AICs suggested that the data were most consistent with model 3, which posited that the variation over time has come about as differences across cohorts. Model 1, which depicted change over time as a period effect, had the worst fit to the data among all of the models considered. In the following section of the paper, we therefore focus on

cohort differences in order to investigate whether the determinants of divorce have changed over time.

Table 3 presents the estimates of interactions between the marriage cohort dummy variable (married before or after 1981) and the key independent variables.[8] The main effect in Table 3 referred to women married before 1981, whereas for those married after 1981, the interaction quantified the variation in the effect of interest. In model 2, we also added the variables that refer to the employment situation. First, distinct patterns were found with respect to changes in the risk of divorce with the duration of marriage. Although the estimates were not found to be statistically significant due to the low number of events, the risk of divorce was shown to decline as the duration of the marriage increased for women married before 1981 (main effect), while the risk was found to increase for those married after 1981. We also used different time intervals to truncate the maximum duration at 15 years for both cohorts, and the same pattern of results was found (see Appendix Table A5).

The coefficients of age at the time of marriage also changed over time. For women married before 1981, marrying at an older age was associated with a lower risk of divorce. This result is consistent with analyses carried out in other countries (Hoem 1997; Lyngstad 2006; Yi Zeng et al. 2002). However, this effect disappeared for those married after 1981.

With respect to education, among women married before 1981, those with a university degree or secondary education were found to be more likely to divorce than women with only a primary education. The risk of union dissolution was 3.2 times [exp (1.15)] higher for women with a university education than for women with a primary education or less. The risk for women with upper secondary education was even higher, at approximately four times that of women with a primary education or less. The effect of education was reduced when employment was taken into consideration, but it remained statistically significant. For the women married after 1981, however, the findings indicated that the positive coefficient of the constitutive term for education was largely compensated for by the negative coefficient of the interaction term (column 2). As a result, the positive relationship between education and divorce vanished for the women married after 1981. Indeed, an analysis restricted to women married after 1981 found no significant differences between educational levels (see Appendix Table A4). These results refine the findings of Härkönen and Dronkers (2006), which were based on data collected between 1994 and 1995, and placed Spain in the same category as countries such as France, Greece, Italy, and Poland, where the educational gradients for divorce were positive. The present analysis, undertaken for more recent marriage cohorts, partially corroborates the hypothesis of a reversal in the educational gradient on divorce originally formulated by Goode (1962). It appears, then, that Spain is

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[8] The results for the two marriage cohorts are also reported separately in appendix Table A4.

situated halfway between those countries in which the positive relationship between divorce and education still exists, such as Italy (de Rose 1992; Vignoly and Ferro 2009), and those where it has completely inverted, such as Holland (De Graaf and Kalmijn 2006), Sweden (Hoem 1997), Norway (Lyngstad 2004), Japan (Raymo, Iwasawa, and Bumpass 2004), Austria, and the United States (Härkönen and Dronkers 2006).

Table 2: **Event history analysis: Transition from marriage to marital dissolution. Coefficients.**

	Model 1	Model 2	Model 3	Model 4	Model 5
Marriage duration (years)					
<2	-0.41	-0.58*	-0.51*	-0.56*	-0.49
2 and 3	-0.06	-0.16	-0.12	-0.15	-0.11
4-5 (ref.)					
6-7	-0.22	-0.13	-0.16	-0.14	-0.17
8-9	-0.31	-0.14	-0.21	-0.16	-0.23
10-11	-0.34	-0.12	-0.20	-0.14	-0.23
12-13	-0.36	-0.07	-0.18	-0.10	-0.21
14-15	-0.25	0.10	-0.02	0.06	-0.06
16 or more	-0.36	0.38	0.27	0.31	0.20
Educational levels					
Primary or less (ref.)					
Lower sec.	0.53***	0.25	0.17	0.25	0.17
Upper sec.	0.75***	0.44**	0.40*	0.44**	0.40*
University	0.68***	0.42	0.35	0.41	0.35
Age at marriage	-0.06***	-0.08***	-0.07***	-0.08***	-0.07***
Parity and age of youngest child					
None (ref.)					
1 child <7	-0.36	-0.35	-0.39*	-0.34	-0.39*
1 child >=7	-0.14	-0.13	-0.20	-0.13	-0.20
2 children <7	-0.67***	-0.68**	-0.72***	-0.67**	-0.71***
2 children>=7	-0.72**	-0.61**	-0.68**	-0.62**	-0.68**
Age heterogamy					
Homogamy					
Husband 5 years older	0.30**	0.33**	0.36**	0.33**	0.36**
Wife older	0.52***	0.52***	0.51***	0.52***	0.51***
Parental divorce	0.93***	0.86***	0.91***	0.86***	0.90***
Premarital cohabitation	0.41**	0.26	0.32	0.26	0.32
Premarital Pregnancy	0.70***	0.65***	0.69***	0.65***	0.69***
Children with other than spouse	0.74**	0.75**	0.69**	0.75**	0.69**

Table 2: **(Continued)**

	Model 1	Model 2	Model 3	Model 4	Model 5
Employment					
Inactive (ref.)					
Unemployed	1.23***	1.17***	1.10***	1.17***	1.10***
Employed	0.87***	0.76***	0.73***	0.76***	0.73***
Year of marriage				0.05***	
(1949=0)		0.05***			
Year 1981 onwards	1.11***			0.22	0.24
Marriage cohort					
1949-69			-1.80***		-1.69***
1970-80			-0.36**		-0.31*
1981-90 (ref.)					
1991-2006			0.20		0.17
Constant	-8.12***	-8.47***	-6.58***	-8.52***	-6.79***
AIC	2856.2062	2816.701	2807.6966	2819.984	2810.7348
Number of subjects	4,409	4,409	4,409	4,409	4,409
Number of failures	339	339	339	339	339

* p<.1; ** p<.05; *** p<.01

Table 3: **Event history analysis: Transition from marriage to marital dissolution. Summary of interactions: Independent variables*married from 1981. Coefficients.**

	Model 1		Model 2	
		*interaction 1981		*interaction 1981
Marriage duration (years)				
<2	-1.06	0.62	-1.04	0.61
2 and 3	0.11	-0.35	0.11	-0.34
4-5 (ref.)				
6-7	-0.49	0.47	-0.49	0.47
8-9	-0.50	0.43	-0.50	0.44
10-11	-0.79	0.88	-0.78	0.88
12-13	-0.90	1.10	-0.90	1.12
14-15	-1.39*	2.03**	-1.39*	2.06**
16 or more	-0.55	1.03*	-0.44	0.97

Table 3:   (Continued)

	Model 1		Model 2	
		*interaction 1981		*interaction 1981
Educational levels				
Primary or less (ref.)				
Lower sec.	0.88***	-0.95**	0.75***	-0.90**
Upper sec.	1.40***	-1.28***	1.14***	-1.15***
University	1.15***	-0.83*	0.74**	-0.62
Age at marriage	-0.13***	0.08*	-0.11***	0.06
Married from 1981	0.30		0.72	
Parity and age of youngest child				
None (ref.)				
1 child <7	0.21	-0.86	0.23	-0.83
1 child >=7	0.61	-1.20	0.66	-1.20
2 children <7	0.00	-1.18*	0.08	-1.19*
2 children>=7	0.45	-2.03***	0.51	-2.05***
Age heterogamy				
Homogamy				
Husband 5 years older	-0.13	0.63**	-0.10	0.62**
Wife older	0.52	-0.01	0.50	0.05
Parental divorce	1.33***	-0.50	1.20***	-0.40
Premarital cohabitation	-0.19	0.65	-0.36	0.79
Premarital Pregnancy	0.10	1.06*	0.07	1.06*
Children with other than spouse	1.12**	-0.52	1.16**	-0.48
Employment				
Inactive (ref.)				
Unemployed			1.78***	-0.88**
Employed			1.01***	-0.39
Number of subjects	4,409		4,409	
Number of failures	339		339	

* p<.1; ** p<.05; *** p<.01

Consistent with the standard expectations of New Home Economics, our analysis showed that employed and unemployed women were at greater risk of divorcing than inactive women, although the coefficient was found to be smaller for women married after 1981, especially for the unemployed (see Table 3 model 2). Thus, the hazard rate of divorce for employed women married before 1981 was almost three times higher [exp (1.01)=2.7] than that of inactive women, and approximately two times higher for

marriages after 1981 [exp (0.62)=1.9]. This result confirms for Spain a solid finding from the majority of Western countries, where the relationship between the employment of women and the probability of divorce has been documented (South 2001; Lyngstad 2004; Poortman and Kalmijn 2002; Kalmijn and Poortman 2006). Notably, the likelihood of divorce was shown to be higher among unemployed women than those who are employed.[9] This might be due to the fact that unemployment generates stress within a marriage, and thus fosters union breakup. It might also reflect reverse causation, with inactive women starting to look for a job once the decision to divorce has been made, even though they are still unemployed when the divorce finally occurs. Still, this result confirms at an individual level the findings of an analysis by Fischer and Liefbroer (2006) for Holland, which showed a positive relationship between particularly unfavourable economic conditions at a macro level and dissolution rates.

The estimates of the interactions suggest that the positive effect of being economically active on the risk of divorce has been smaller among women married after 1980, especially in the case of the unemployed. The effect of being employed has also diminished, although the difference is not statistically significant. The results partly corroborate the hypotheses previously mentioned. First, it might be the case that, in older cohorts, economic barriers to divorce or separation were much higher than for couples married after 1981, which make being employed a crucial prerequisite to exiting an unsatisfactory union. Second, over time, the dual-earner couple model has spread: it may have become a more efficient and stable arrangement than the traditional breadwinner model, especially in times of high labour market insecurity for male workers (Oppenheimer 1997). Finally, an additional explanation is related to the self-selection argument. The high proportion of active women in Spain who were born in the 1960s has made the group of women who are employed a more heterogeneous— and, probably, a less self-selected—group with respect to the unconventional values that usually increase the risk of divorce. Therefore, while being economically active still increases the likelihood of divorce, the differences between marriage cohorts are lower.

Having children, especially two or more, decreases the likelihood of marital dissolution, but only for women married after 1981. It was expected that having children would be more decisive for the older marriage cohorts. These hypotheses were based on two arguments. The first one cited the lower economic risks of living as a single parent for the more recent marriage cohorts. The second cited the more permissive normative climate that allows parents with children to divorce. The results have shown the opposite effect: having children was found to affect only the more recent marriage cohorts.

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[9] Various models have been tested by changing the category of reference relative to employment, and the difference between working and being unemployed has been found to be statistically significant ($p<0.05$).

Contrasting patterns were found for unconventional characteristics. Our expectation was that the positive effect of these attributes would decline parallel to their diffusion in the Spanish society. The more common these characteristics have become, the weaker their association with divorce. However, a reduction between marriage cohorts regarding the positive association with divorce was only observed for parental divorce and having a child from a previous relationship. In neither case was the reduction statistically significant. On the other hand, the effects of premarital cohabitation, premarital pregnancy, and age heterogamy were found to be stronger for marriages after 1981 (although the increase in the coefficient for premarital cohabitation was not statistically significant). As a result, for recent marriage cohorts, what we termed "unconventional characteristics" were generally shown to have positive associations with divorce.

5. Conclusions

In this article, we investigated how divorce risk factors have changed over time in Spain. Our main finding was that the impact of socioeconomic resources has declined in the more recent marriage cohorts, but that the effects of having children and the effects of unconventional attributes related to union formation or previous family experience have become more important. More precisely, among women married before 1981, being highly educated and active in the labour market were the key factors associated with the risk of divorce. The small number of individuals who had divorced parents or children from a previous relationship were also at greater risk of union dissolution. Among women married after 1981, the educational level became irrelevant, and the effect of being unemployed or employed, versus being inactive, decreased; although in the case of being employed, the reduction was not statistically significant. Consistent with previous findings from other countries, our results showed that, for women married after 1981, having children largely decreased the risk of divorce (Wagner and Weiß 2006). The findings for other unconventional characteristics related to union formation or previous family experience were also in line with evidence from other countries (Härkönen and Dronkers 2006; Lyngstad 2006; Wagner and Weiß 2006). However, the expected relationship was found only for women married after 1981. Thus, having experienced a premarital pregnancy or having a husband at least five years older was shown to be positively associated with union disruption in the more recent marriage cohorts. According to Appendix Table A4, premarital cohabitation was also associated with an increase in the likelihood of marital disruption for women married from 1981 onwards, although the difference of the coefficients between the older and the younger cohort was not statistically significant (see Table 3 in the main text).

What are the theoretical implications of these results? First, the finding that the association between higher education and the risk of divorce was positive and strong in the older marriage cohort, but has waned in the younger cohort, lends support to the prediction made by Goode (1962) and Blossfeld et al. (1995) that there will be a reversal of the educational gradient over time. Other studies from the Netherlands and Sweden that used the same specification as this study for female education, without considering the husband's education or the wife's income, have shown a negative association between education and the risk of divorce among the more recent marriage cohorts (Hoem 1997; De Graaf and Kalmijn 2006). For present-day Spain, we did not find such a negative association. The important question with substantive implications for the structuring of social inequality is whether the reversal in the educational gradient will continue into the future, to the point where a negative association between education and divorce will also be observed in Spain.

Second, unconventional traits seem to matter more now than they did in the past. It should be noted that, although most of these traits are more common today than they were in the past, they are still far from being widespread in the Spanish population. For example, while a willingness to enter into pre-marital cohabitation has become more common, only 18% of the women married after 1981 cohabitated. In other words, while these traits have become more prevalent, they are still rather unconventional. Teachman (2002) has argued that, when divorce is a rare event, the effects of any determinants that are positively associated with the risk of divorce will be suppressed. However, when legal and administrative barriers are lifted and divorce becomes more common, the effects of those determinants begin to be expressed as higher risks of union disruption. This interpretation seems to fit well with the evidence from Spain.

Our findings place Spain at a middle stage, in which the educational gradient of divorce has disappeared, but has not completely reversed, and in which the influence of some unconventional features of divorce have emerged. Given the sharp rise in divorce in the most recent years, it will be necessary to replicate the analysis for more recent marriage cohorts in order to gain a full understanding of the unfolding of the patterns highlighted in this article. We might hypothesise that, in the next several years, the educational gradient will completely reverse, with the less educated being at a greater risk of divorce, and that the association between unconventional traits and divorce will decrease.

6. Acknowledgments

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Appendix

Table A1: **Women younger than 75 years old married from 1949. Distribution of educational levels (%). Comparison between original file, used file and Spanish Labour Force Survey**

	Original sample	Analytical sample	LFS*
Primary or less	39	39	38
Lower secondary	25	25	23
Upper secondary	23	23	24
University	13	13	15
Total	100	100	100
n	5,995	4,409	81,538

*LFS: Spanish Labour Force Survey, second quarter of 2006.

Table A2: **Women younger than 75 years old married from 1949. Distribution of marital disruption by educational levels (%). Comparison between original file, used file and Spanish Labour Force Survey**

	Original sample	Analytical sample	LFS
Primary or less	3.8	3.6	4.7
Lower secondary	9.0	8.6	7.4
Upper secondary	11.8	10.3	8.1
University	9.9	8.4	6.7
n	441	339	4,870

*LFS: Spanish Labour Force Survey, second quarter of 2006.

Table A3: **Event history analysis: transition from marriage to marital dissolution. Only women married from 1970.**
Col 1 and 3: Women married from 1970-80.
Other columns: Interaction having married from 1981. Coefficients.

	Model 1		Model 2	
	col 1 (1970-80)	*interaction 1981	col3 (1970-80)	*interaction 1981
Marriage duration (years)				
<2	-0.83	0.38	-0.77	0.34
2 and 3	-0.06	-0.17	-0.05	-0.18
4-5 (ref.)				
6-7	-0.33	0.32	-0.35	0.33
8-9	-0.26	0.19	-0.28	0.22
10-11	-0.98	1.07	-1.01	1.11
12-13	-1.20	1.40*	-1.24*	1.46*
14-15	-1.24	1.88**	-1.27	1.95**
16 or more	-0.29	0.77	-0.27	0.81
Educational levels				
Primary or less (ref.)				
Lower sec.	0.25	-0.32	0.16	-0.31
Upper sec.	0.84***	-0.71*	0.67**	-0.69*
University	0.25	0.07	-0.06	0.18
Age at marriage	-0.10***	0.04	-0.08**	0.03
Married from 1981	0.43		0.86	
Parity and age of youngest child				
None (ref.)				
1 child <7	0.40	-1.05	0.49	-1.09*
1 child >=7	0.43	-1.02	0.58	-1.12
2 children <7	0.14	-1.31	0.29	-1.40*
2 children>=7	0.53	-2.11**	0.68	-2.23**
Age heterogamy				
Homogamy				
Husband 5 years older	0.29	0.21	0.34	0.18
Wife older	0.46	0.04	0.48	0.06

Table A3: (Continued)

	Model 1		Model 2	
	col 1 (1970-80)	*interaction 1981	col3 (1970-80)	*interaction 1981
Parental divorce	1.04**	-0.21	0.87*	-0.07
Premarital cohabitation	-0.17	0.63	-0.33	0.76
Premarital Pregnancy	-0.89	2.04***	-1.12*	2.25***
Children with other than spouse	0.92	-0.31	0.96	-0.29
Employment				
Inactive (ref.)				
Unemployed			1.63***	-0.73
Employed			0.86***	-0.24
Constant	-6.38***		-7.29***	
Number of subjects	3,101		3,101	
Number of failures	303		303	

* p<.1; ** p<.05; *** p<.01

Table A4: Event history analysis: Transition from marriage to marital dissolution. Two separate analyses: Married before 1981 and married from 1981. Coefficients.

	Married before 1981	Married from 1981
Marriage duration (years)		
<2	-1.06	-0.45
2 and 3	0.11	-0.24
4-5 (ref.)		
6-7	-0.49	-0.02
8-9	-0.50	-0.07
10-11	-0.79	0.09
12-13	-0.90	0.20
14-15	-1.39*	0.64
16 or more	-0.55	0.48
Educational levels		
Primary or less (ref.)		
Lower sec.	0.88***	-0.07
Upper sec.	1.40***	0.12
University	1.15***	0.32
Age at marriage	-0.13***	-0.06**

Table A4:   (Continued)

	Married before 1981	Married from 1981
Parity and age of youngest child		
None (ref.)		
1 child <7	0.21	-0.64**
1 child >=7	0.61	-0.59
2 children <7	0.00	-1.17***
2 children>=7	0.45	-1.59***
Age heterogamy		
Homogamy		
Husband 5 years older	-0.13	0.50**
Wife older	0.52	0.50**
Parental divorce	1.33***	0.83***
Premarital cohabitation	-0.19	0.46**
Premarital Pregnancy	0.10	1.16***
Children with other than spouse	1.12***	0.61
Constant	-6.25***	-5.95***
Number of subjects	2,317	2,092
Number of failures	145	194

* p<.1; ** p<.05; *** p<.01

Table A5:   Event history analysis: transition from marriage to marital dissolution. Duration truncated at 15 years (models 3 and 4). Coefficients.

	Without truncation				Duration truncated at 15 years			
	Model 1		Model 2		Model 3		Model 4	
		*Interaction 1981		*Interaction 1981		*Interaction 1981		*Interaction 1981
Marriage duration (years)								
<2	-1.06	0.62	-1.04	0.61	-0.95	0.45	-0.90	0.42
2 and 3	0.11	-0.35	0.11	-0.34	0.13	-0.40	0.14	-0.40
4-5 (ref.)								
6-7	-0.49	0.47	-0.49	0.47	-0.51	0.50	-0.50	0.49
8-9	-0.50	0.43	-0.50	0.44	-0.52	0.47	-0.51	0.47
10-11	-0.79	0.88	-0.78	0.88	-0.75	0.88	-0.73	0.87
12-13	-0.90	1.10	-0.90	1.12	-0.80	1.07	-0.79	1.07
14-15	-1.39*	2.03**	-1.39*	2.06**	-1.25*	2.01**	-1.23*	2.02**
16 or more	-0.55	1.03*	-0.44	0.97				

Table A5: (Continued)

	Without truncation				Duration truncated at 15 years			
	Model 1		Model 2		Model 3		Model 4	
		*Inter action 1981		*Inter action 1981		*Inter action 1981		*Inter action 1981
Educational levels								
Primary or less (ref.)								
Lower sec.	0.88***	-0.95**	0.75***	-0.90**	0.82**	-1.07**	0.71*	-1.05**
Upper sec.	1.40***	-1.28***	1.14***	-1.15***	1.72***	-1.71***	1.53***	-1.65***
University	1.15***	-0.83*	0.74**	-0.62	1.75***	-1.41**	1.38**	-1.23*
Age at marriage	-0.13***	0.08	-0.11***	0.06	-0.19***	0.14**	-0.18***	0.13**
Married from 1981	0.30		0.72		-0.32		-0.06	
Parity and age of youngest child								
None (ref.)								
1 child <7	0.21	-0.86	0.23	-0.83	0.41	-1.13	0.47	-1.17
1 child >=7	0.61	-1.20	0.66	-1.20	0.99	-1.49	1.09	-1.55
2 children <7	0.00	-1.18*	0.08	-1.19*	0.26	-1.58**	0.38	-1.65**
2 children>=7	0.45	-2.03***	0.51	-2.05***	0.38	-2.53**	0.48	-2.59**
Age heterogamy								
Homogamy								
Husband 5 years older	-0.13	0.63**	-0.10	0.62**	0.20	0.36	0.24	0.35
Wife older	0.52	-0.01	0.50	0.05	1.15***	-0.70	1.07***	-0.58
Parental divorce	1.33***	-0.50	1.20***	-0.40	1.17*	-0.40	1.04*	-0.29
Premarital cohabitation	-0.19	0.65	-0.36	0.79	-1.12	1.51*	-1.36	1.73*
Premarital Pregnancy	0.10	1.06*	0.07	1.06*	0.45	0.89	0.31	1.00
Children with other than spouse	1.12***	-0.52	1.16**	-0.48	0.38	0.25	0.52	0.18
Employment								
Inactive (ref.)								
Unemployed			1.78***	-0.88**			1.53***	-0.63
Employed			1.01***	-0.39			0.75**	-0.20
Constant	-6.25***		-7.15***		-5.48***		-6.18***	
Number of subjects	4,409		4,409		4,409		4,409	
Number of failures	339		339		339		339	

* p<.1; ** p<.05; *** p<.01