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Gender Differences in the Transition to Adulthood in France: Is There Convergence Over the Recent Period?

Le passage vers l'âge adulte des hommes et des femmes en France : y-a-t-il convergence?

Maria Winkler-Dworak · Laurent Toulemon

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Abstract Numerous studies have shown that educational attainment and labour force status have a strong impact on the timing of family formation for both men and women. The effects of educational level, school enrolment and employment seem to be different for men and women. The aim of this article is to investigate how gender-specific differences in family formation have changed over time, and more particularly, whether these differences have disappeared in recent years. We use a large-scale survey (more than 240,000 men and women born after 1940) conducted within the French 1999 census and apply event history techniques. The sample size allows us to test our hypotheses with more sophisticated models that cover several interactions. Our data fully support the convergence hypothesis for men and women with regard to the effects of educational attainment and working status (working/not working). However, it is only partly relevant for the effects of their school enrolment status on entry into first union and parenthood. For both men and women, the impact of work experience on first union disappears over time, but remains important for first parenthood.

Keywords Transition to adulthood · Family formation · Event history techniques · France · Gender · Education

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M. Winkler-Dworak (✉)
Vienna Institute of Demography, Austrian Academy of Sciences, Wohllebengasse 12–14, 6th floor, Vienna 1040, Austria
e-mail: maria.winkler-dworak@oeaw.ac.at

L. Toulemon
Institut national d'études démographiques, 133 Boulevard Davout, Paris cedex 20 75980, France
e-mail: toulemon@ined.fr

Résumé De nombreux travaux ont étudié les fortes variations sociales de l'âge à la mise en couple et à la naissance du premier enfant. Les variations selon le niveau d'éducation, le statut face aux études et à l'emploi sont en général différentes pour les hommes et pour les femmes. Comment ces différences entre hommes et femmes ont-elles évolué au cours des quarante dernières années ? Pour répondre à cette question, nous appliquons des modèles de durée aux biographies de plus 240 000 hommes et femmes nées après 1940, recueillies en 1999 dans le cadre de l'enquête Étude de l'histoire familiale, intégrée au recensement général de la population.

La taille de l'échantillon nous permet de construire des modèles comprenant de nombreuses interactions. Pour les mises en couple et l'arrivée du premier enfant, nous mettons en évidence une convergence entre hommes et femmes dans les variations selon le niveau d'éducation et le statut professionnel ; la convergence est moins forte pour les variations selon le statut face aux études. Pour les hommes comme pour les femmes, l'expérience professionnelle perd son importance comme préalable aux unions, mais la garde pour l'arrivée du premier enfant.

Mots-clés Passage à l'âge adulte · Formation de la famille · Modèles de durée · France · Genre · Éducation

1 Introduction

Educational attainment and employment define important markers in the transition to adulthood and have been shown to be the main predictors of the timing of family formation for males and females, both theoretically (Becker 1981; Oppenheimer 1988) and empirically (e.g. Goldscheider and Waite 1986; Hoem 1986; Blossfeld and Huinink 1991; Thornton et al. 1995; Liefbroer and Corijn 1999). However, these studies show differences between men and women regarding the impact of educational level, school enrolment and employment on family formation. For instance, most studies find that educational attainment and employment speed up family formation among men (e.g. Huinink 1995), but tend to delay entry into a union and motherhood among women (e.g. Blossfeld and Huinink 1991; Liefbroer and Corijn 1999).

In most western countries, the pattern of family formation has changed significantly since the 1960s. In the past decades, family formation has been characterised by a postponement of marriage and parenthood, a decline in fertility and marriage, and an increasing prevalence of non-marital cohabitation and non-marital childbearing (e.g. Billari 2005). Changes in family formation have not occurred in isolation: in the past few decades, education levels have risen for both sexes, but much more for women. In France, women nowadays study longer and obtain higher levels of education, which was far from being the case among earlier cohorts (Estrade and Minni 1996). The educational expansion has been accompanied by a steady increase in female employment. For instance, labour force participation of French women aged 25–39 years almost doubled from 42 to 74% between 1962 and 1989 (Leridon and Toulemon 1995) and was as high as 79% in 2002 (ILO 2003).

The changes in family formation have been linked to changes in norms and attitudes; cohabitation and marriage are increasingly expected to be "a partnership of equals" (Van de Kaa 1994). Moreover, the demographic changes have challenged family policies. Since the 1970s, family policies have changed significantly in industrialised countries and the incompatibility between family roles and female labour force participation has weakened over the past decades (Gauthier 1996). Do these changes in norms and family policies have implications on family formation, and more specifically, on the observed gender differences in the effects of educational attainment and employment on the timing of family formation?

The focus of this article is on investigating the impact of educational attainment and employment on the timing of family formation among men and women in France. Have the differences in the impact of education and work between men and women changed over time; in particular, have the gender differences in these effects diminished?

We studied the question of convergence by applying event history techniques, which model the entry into first union and first parenthood, respectively. In particular, we used hazard regression to analyse the time trend in the different impact of educational attainment and employment on the timing of family formation for males and females. The large sample under study allows us to consider enrolment in education, by considering separately the last year in studies, and to take into account duration since the end of studies for those who are no longer enrolled, as well as work experience and working status.

Section 2 of the article contains a detailed discussion of family formation theories in relation to education and employment and their empirical relevance. Section 3 introduces the data and methods, and Sect. 4 presents the results. Section 5 contains the conclusions and outlines areas for further analysis.

2 Theoretical Background

2.1 Theories on Family Formation

There are several theories about why people marry and have children and what factors influence the timing of marriage and entry into parenthood. In the New Home Economics theory, marriage is seen as a rational choice made by individuals for whom the gains from marriage outweigh the benefits of remaining single. Given the complementarity of men and women in the household production of commodities, these individuals would be more productive in a joint than in a single household. If each sex specialises in its comparative advantage,[1] the gendered division of labour within households will create gains from marriage (Becker 1976,

[1] A household member is said to have a comparative advantage in the household if the ratio of the marginal product in the household to his/her wage rate on the market is higher for him/her than for the other household member(s) if all contribute the same amount of time to the household and all invest in the same human capital (Becker 1981).

1981). Gains from marriage are thus based on the assumption of complementarity between men and women and on their market opportunities.

On the one hand, higher wages imply a higher total income of the joint household, which increases the gain from marriage (*income effect*). On the other hand, higher wages reduce the comparative advantage of the household sector, which, in turn, increases the opportunity costs of household work (*price effect*). Hence, the gender-specific division of labour within the family becomes less advantageous for the individual specialising in household tasks, which reduces the gains from marriage. If a traditional division of labour prevails, the income effect is expected to dominate among men, whereas the price effect is expected to outweigh the income effect among women (Liefbroer and Corijn 1999). As higher educational attainment is associated with higher expected earning levels (which implies higher economic independence), Becker's theory predicts a positive impact of education on the entry into marriage for males and a negative one for females.

Becker's economic approach also draws conclusions on the decision to have children. He argues that the "main purpose of marriage and the family is the production and rearing of their own children" (Becker 1981, p. 93). Children are produced and reared by the use of market goods and services and parental time, especially mother's time because of the sex-specific differentiation of labour in the family. A growth in the value of women's time as a result of increased investments in education and career opportunities will therefore raise the relative costs of children and thereby reduce the demand for children. As the final number of children is linked to the timing of first births, i.e. early fertility is associated with high fertility (Kohler et al. 2001; Toulemon 2006), given Becker's argumentation, the impact of educational level and labour force participation on family formation is expected to be positive for men, but negative for women.

However, Becker's specialisation model has been severely criticised for several reasons. In particular, Oppenheimer (2000, p. 285) argues that "sex role specialisation is essentially a high risk and inflexible family strategy in an independent nuclear family system." Better-educated women might also perceive economic independence as a means to share the costs of setting up a common household which will accelerate entry into a union (Oppenheimer 1994), and they may more easily be able to afford child-related costs and to purchase care in the market (Martín-García and Baizán 2006). Hence, for better-educated women, the income effect might be stronger and more dominant than the price effect (Oppenheimer 1988; Kravdal 1994), and thus the effect of educational level on family formation for women may be rising for high levels of education, which may result in a U-shaped relation between educational level and female family formation rates.

The prolonged period of education implied by higher educational levels has itself a substantial impact on the timing of union formation and parenthood. In particular, when attending school or university, students are normally not economically independent and rely heavily on their parents' financial support, which makes family formation unlikely (Blossfeld and Huinink 1991). In addition, "there exist normative expectations in society that young people who attend school are 'not at risk' of entering marriage and parenthood" (Blossfeld and Huinink 1991, p. 147).

Roles of student and parent are sufficiently demanding and therefore the incompatibility of being a student and engaging in adult family activities thus delays family formation until the educational career has been completed (Rindfuss et al. 1980; Marini 1985; Blossfeld and Huinink 1991).

Moreover, the incompatibility of student and parental roles implies that having a child would lead to dropping out of school. This decision is costly because it results in truncation of accumulation of knowledge and skills necessary for an attractive job, resulting in poorer job and a lower standard of living than might have been possible: "[f]or young people who decide to combine marriage and student roles, marrying near the end of one's schooling minimises the time spent in conflicting roles" (Thornton et al. 1995: 764). However, it may also be the case that the conflict between forming a union and being a student makes many students drop out at an earlier stage of their education after their entry in a first union, which might raise union formation rates shortly before leaving school. In any case, these opportunity costs of dropping out of school have risen sharply over the last decades (Oppenheimer 1988). Although, schooling has become more important for men and women over the last decades, to the extent that couples adhere to gender-specific division of labour, the incompatibility between school enrolment and family formation may be stronger among women (Davis and Bumpass 1976; Liefbroer and Corijn 1999). Summing up, school enrolment is expected to have a negative impact on the timing of family formation for both women and men, but stronger for women than for men, and stronger for first birth than for entry into first union.

After completing education and entering employment, men and women accumulate labour-force experience at their workplaces. As earnings increase with work experience, following Becker's argumentation, the effect of work experience should be positive among men but negative among women. Later theories of neo-classical economics hypothesise that it is optimal for women to have the first child early in their work career if the acquired working skills are liable to decay strongly during their child-related absence from the labour market (Happel et al. 1984), but to shift the onset of fertility to a later stage of their work career if the potential lifetime earnings profile increases steeply (Cigno and Ermisch 1989). Moreover, there may be normative expectations, not only for men but also for women, to establish themselves in the labour market and to amass savings before embarking on parenthood (Toulemon and Leridon 1999; Santow and Bracher 2001). In addition, child allowances and maternal leave entitlement may be connected to previous work experience. Hence, work experience may have a positive impact on family formation for women as well.

With the spread of unmarried cohabitation, the postulated effects of educational attainment and labour force participation on union formation may have changed, since several authors suggest that the opportunity costs of unmarried cohabitation are lower than those of marriage, and that cohabitation requires less economic underpinning and may be less conflicting with student roles (Thornton et al. 1995; Oppenheimer 1988; Kravdal 1999). Furthermore, cohabitation is associated with less sex-differentiated roles (Van de Kaa 1994; Leridon and Toulemon 1995) and Cherlin (2000) argues that women incorporate premarital cohabitation into their search processes because it provides a better opportunity to observe men's income potential. Moreover, family policies have changed significantly since the 1970s in industrialised countries and the incompatibility between family and female labour force participation has weakened over the past decades (Gauthier 1996). Hence, the effects of educational attainment and labour force participation on union formation are expected to have weakened, while we do not anticipate a similar change for first birth. Last but not least, gender differences in the impact of education and employment on union formation and entry into parenthood are expected to have diminished in the most recent periods.

2.2 Findings of Prior Empirical Studies on Family Formation

Existing studies on family formation among men lend support to the New Home Economics. Rising earnings level have been found to accelerate men's family formation (Huinink 1995; Bracher and Santow 1998; Oppenheimer and Lewin 1999) and being in employment is a prerequisite for marriage and parenthood for men (Goldscheider and Waite 1986; Huinink 1995; Liefbroer and Corijn 1999; Oppenheimer and Lewin 1999). Moreover, educational level has been found in most studies to have a positive impact on the intensity of entry into union for men (Goldscheider and Waite 1986; Huinink 1995; Bracher and Santow 1998; Oppenheimer and Lewin 1999; Coppola 2004). Liefbroer and Corijn (1999) do not find differences in union formation and marriage by educational level for Flemish and Dutch males, but Brüderl and Diekmann (1994) identify for US and West German males a significantly negative impact of educational level on marriage. In contrast to other studies, Brüderl and Diekmann (1994) use a parametric model, i.e. a log-logistic function in order to capture the age dependency of marriage rates. When controlling for the institutional effect of education, the age-specific marriage rates are shifted to the right depending on the average number of schooling years after the minimum marriage age. This approach, also used by Brüderl and Klein (1993) in the analysis of female family formation rates, was criticised by Blossfeld et al. (1993) because it forces the rate function to increase steeply while the better-educated are still in education, which may explain the obtained negative effect of educational level on family formation.

For the entry into fatherhood, the empirical evidence is less clear. While Huinink (1995) reports a positive, albeit not significant, impact of educational level on fatherhood for West German males, this effect is statistically significantly negative for Dutch and Flemish males in Liefbroer and Corijn (1999). However, the negative effect of educational level on parenthood shrinks and loses significance for ages above 24 years. Both studies include educational attainment, measured by the average number of schooling years in order to obtain the degree, in a linear term. But if the educational level and family formation are related in a non-linear way, the assumption of a linear relation may result in non-significant estimates (Buber 2002).

The empirical evidence on the effects of educational attainment and labour force participation of family formation is mixed for women. In line with the New Home Economics, being not employed facilitated entry into motherhood (Kravdal 1994; Liefbroer and Corijn 1999; Santow and Bracher 2001; Buber 2002; Meron and

Widmer 2002). However for entry into a first union, being out of the labour force had either no effect (Liefbroer and Corijn 1999) or delayed union formation among women (Goldscheider and Waite 1986; McLaughlin et al. 1993; Oppenheimer and Lewin 1994; Bracher and Santow 1998). Moreover, a longer duration in the labour market facilitated entry into motherhood, where a levelling off of the effects after 2–4 years working was observed for Norwegian (Kravdal 1994) and Swedish women (Santow and Bracher 2001). Blossfeld and Huinink (1991) developed an indicator for career resources which incorporates accumulation of labour-force experience, job prestige score, and depreciation of working skills during absence from the labour market. However, this indicator proved to have no effect on the entry into marriage of West German women. In addition, higher earnings accelerated entry into a union among women in Sweden (Bracher and Santow 1998) and marriage in the US (McLaughlin et al. 1993, Oppenheimer and Lewin 1994). The above findings are clearly contrary to Becker's arguments.

A woman's educational level has empirically been found to be of major importance for the timing of family formation. However, previous studies on these effects give diverging results. In particular, Liefbroer and Corijn (1999) found a significantly negative impact of educational level on first union formation for Dutch and Flemish females,[2] while the studies of Hoem (1986) and Leridon and Toulemon (1995) did not reveal any specific effect of educational attainment on entry into a first union in Sweden and France, respectively. Moreover, two more recent studies on Italy and Spain (Coppola 2004) and Sweden (Bracher and Santow 1998) even found that a university degree accelerated union formation.

A similarly diverse picture is obtained for the timing of first marriage in relation to educational attainment. In particular, Oppenheimer and Lewin (1994) and Leridon and Toulemon (1995) found a significantly higher rate of marriage of low educated white US women and French women, respectively. In contrast, Oppenheimer and Lewin (1994) revealed a strong positive effect of educational level on entry into marriage for black US women, which was greatly reduced and lost significance when working status and earnings were controlled for. Other studies on US females (Goldscheider and Waite 1986; McLaughlin et al. 1993) identified higher marriage rates for higher educated women, but they do not consider an interaction between educational level and race. Finally for West German women, no significant effect of educational level on marriage could be identified (Blossfeld and Jaenichen 1992; Blossfeld and Huinink 1991).

The findings in the literature about the effect of educational attainment also vary considerably for the entry into motherhood. In some studies, the educational level was found to delay motherhood: but the effect was found to be significant (Brüderl and Klein 1993; Liefbroer and Corijn 1999; Martín-García and Baizán 2006), weak or insignificant (Blossfeld and Jaenichen 1992). Other studies, in contrast, even revealed an accelerating effect of educational level on motherhood (Blossfeld and Huinink 1991) or a U-shaped effect, with medium levels of education showing the

[2] However, the effect of educational level on first union formation increases with age and loses significance, i.e. there is no effect of educational attainment of the timing of union formation for 24 to 30-year old Dutch and Flemish women.

lowest intensities of entry into motherhood (Bernhardt 1990; Kravdal 1994; Santow and Bracher 2001; Buber 2002). The estimated impact of educational level on the timing of first births not only varied across countries, but also within countries in different studies. In particular, the three West German studies on the timing of motherhood derived a significant negative effect (Brüderl and Klein 1993), a weak negative and insignificant effect (Blossfeld and Jaenichen 1992), and even a significant positive effect (Blossfeld and Huinink 1991). All three studies pertained to similar birth cohorts of women, but used different model assumptions. While Brüderl and Klein (1993) applied a log-logistic model (see discussion above), Blossfeld and Jaenichen (1992) and Blossfeld and Huinink (1991) used age splines and a time-varying dummy indicating whether the individual is in education. Moreover, Blossfeld and Huinink (1991) additionally controlled for the level of career resources, which proved to have a strong negative impact on entry into motherhood. The model estimated without the career resources gave an insignificant effect of educational level on motherhood rates. Hence, the effect of educational level on motherhood is affected by women's attachment to the labour market. The degree of incompatibility of work and family life is suggested to be one of the reasons for cross-country differences in the estimated effect of educational attainment on family formation (Blossfeld 1995).

With respect to the effect of school enrolment, the findings in the literature are unambiguous: being in education strongly depresses rates of entering a union and parenthood for men and women, which supports the hypothesis of incompatibility between student and family roles. Several studies also show gender differences in the impact of educational enrolment on family formation, with stronger estimated effects for women than for men (Goldscheider and Waite 1986; Bracher and Santow 1998; Liefbroer and Corijn 1999; Coppola 2004).

Our aim is twofold. First, based on these mixed empirical results, we want to test some specific hypotheses on trends on family formation in France, using models of union formation and first birth rates by educational attainment among women and men in France, while controlling for enrolment in education as well as for working status and work experience; second we want to test the global hypothesis that gender differences in family formation diminished between the 1960s and the 1990s.

2.3 Family Formation and Family Policies in France

In France, fertility declined when the baby boom period came to its end, between 1964 and 1975, and stabilised thereafter. Effective methods of contraception became widely available during the 1970s among young adults, and the transitions to first union and first birth were at the highest in the early 1970s, before a dramatic decline which lasted up to the mid-1990s due in part to the progressive diffusion of effective methods of contraception and the decline of unwanted pregnancies (Guibert-Lantoine and Leridon 1999).

Unlike other European countries, France has maintained a relatively high-level of fertility. After declining for more than 30 years, fertility levels rose again and recently reached close to replacement level (Richet-Mastain 2007). The relatively high fertility levels have been related to the family policies in France (Letablier 2003). Indeed, France is considered as the prototype of a country practising a pro-natalist policy and offering a wide range of allowances that help families to lower the direct and indirect costs of children. The benefits range from monetary support, parental leave systems, and tax deductions to childcare facilities (for a detailed description see Aglietta et al. 2002).

Up to the 1960s, the French family policies supported the traditional male breadwinner and homemaker model: there was an allowance for single-earner families, but no child-care allowance (Ekert-Jaffé et al. 2002). Moreover, the family policies targeted large families with young children. Family allowances were not given to the first child, but only from the second child, and the tax system was revised in 1945 to take account of the number of dependent members of the family (Letablier 2003).

From the 1970s onwards, family policies were adapted to changes in family structures. In particular, the rise in the labour market participation of married women implied a change in family policies to incorporate the model of "working mother" and dual-earner families. As a consequence, the single-wage allowance was progressively reduced and finally abolished in 1978 and funding for the construction of childcare centres (*crèches*) was substantially increased at the beginning of the 1980s (Letablier 2003). Between 1984 and 1994, the number of places in *crèches* doubled and they accommodate nowadays about 16% of the infant population, while nursery schools cover all 3-year olds and 40% of 2-year-olds (Ekert-Jaffé et al. 2002). In addition, care by an approved child-minder is subsidised by the state.

Besides responding to the demand for childcare, family policies also include income support and parental leave. Family allowances (*allocations Familiales*) are aimed to partially compensate for the costs of having children. They are paid to families with at least two children (up to the age of 16/20 if schooling is continued) and increase with the number of children. In addition, there are specific means-tested allowances for children with disabilities, lone parent families, and children in poor families.

Parental leave (*congé parental d'éducation*) was established in 1977. It entitles parents to take unpaid leave or work part-time following maternity leave until the child reaches age 3.[3] Additionally, parents may apply for a parental leave allowance (*allocation parentale d'éducation*), which is a flat-rate benefit for parents with at least two children, the youngest below age 3.[4] However, eligibility for the parental leave allowance is connected to prior work experience: at least 2 years out of the five preceding the birth for second births and out of the preceding 10 years for larger families (Fagnani 1999).

[3] Since introduction, entitlement regulations have been modified several times. In the beginning, only mothers were eligible for parental leave, which was extended to fathers in 1984. Under the new legislation, parents could also work part time, and in 1986 the period of parental leave was extended from 2 to 3 years.

[4] The parental leave allowance was introduced in 1985 and only parents of three children, the youngest below age 3, were eligible. In 1994, the benefit was modified and extended to parents with two children.

The different family benefit measures have different implications on income and price effects according to the New Home Economics. Allowances designed to reduce the monetary costs of children, predominant before the 1970s, mainly affect the income effect. As the costs of children are reduced by a greater proportion for parents with a lower income, family policies reduce fertility differences linked to the income effect. In contrast, policies to facilitate the work-family balance, such as the provision of childcare facilities from 1970 onwards, reduce the opportunity costs of childbearing and thus lower fertility differences associated with the price effect (Ekert-Jaffé et al. 2002).

2.4 Hypotheses to be Tested

The theoretical considerations lead us to a set of three hypotheses regarding gender differences in family formation in France, between 1960 and 2000:

Hypothesis 1 A higher educational level is positively linked to family formation rates for men, but the link is negative or U-shaped for women. In recent periods, a positive effect has been emerging for women likewise, leading to decreasing gender differences.

Hypothesis 2 Being enrolled in education has a negative effect on family formation (first union and first birth) for both sexes. This effect used to be larger for women than for men, but it is decreasing for women, leading to decreasing gender differences.

Hypothesis 3 Having a job has a positive effect, which increases with work experience, on family formation (first union and first birth) for men, and it is now having a similar effect for women, leading to decreasing gender differences.

In addition, the spread of unmarried cohabitation and the increasing delay in first births among couples imply that the stakes in family formation have moved from entering the first union to having a first child. As enrolment in education and working status are more important for the behaviours at stake in the transition to adulthood, this leads to our last hypothesis:

Hypothesis 4 In recent periods, the negative effect of enrolment in education and the positive effect of having a job have weakened for union formation and strengthened for first birth, for both men and women.

3 Data and Methods

The data for this study come from the French *Etude de l'Histoire Familiale* (EHF) 1999, which was conducted together with the census in March 1999 (Cassan et al. 2000). About 235,000 women and 145,000 men completed an additional questionnaire on their children, partnerships, working life, social origin, place of birth and languages spoken in the family (Cassan et al. 2000). We restricted our sample to birth cohorts born after 1940, because we want to describe transitions to first union and first child which took place from the 1960s to the 1990s. Immigrants were only included if they had arrived in metropolitan France before they reached age 15, i.e. they underwent their transition to adulthood in France. Moreover, we excluded observations if the event (pregnancy or union formation) took place before the age of 15.

The date of entry into a first union was identified from a table including questions about the “main dates of partnership and marriage”. Partnership was first defined as “sharing the same household, for 6 months or longer, with or without marriage”. Two lines in the table were devoted to the “first and latest” partnerships. For each of these partnerships, a question was asked about the beginning of the union: “Approximately, when did this partnership period begin (Month–Year)”. This question was asked together with other questions: the existence of children born to the partner, and whether they came to live in the household, the dates, if applicable, of marriage, separation, divorce and death of the partner.

The occurrence of a first child was identified from the following questions. First, a general question was asked: “Have you ever had a child? (Include every child you have adopted or given birth to even if s/he has since died). For those who answered “yes” to that first question, there was a question about the number of children and a table with, for each child, nine questions including month and year of birth.[5] To avoid anticipation bias, we considered that the event “entry into parenthood” took place at the conception of the child, 9-month before the date of birth of the first child.

Finally, about 145,000 women and about 95,000 men remained in our sample, of whom about 82% of all women and about 74% of all men had entered into a first union before the time of the survey. Furthermore, about 70% of all women and 60% of all men had entered into first parenthood.[6]

In our study, we followed individuals from their 15th birthday until the time of the event, i.e. union formation or conception of first child. Having a first birth and entering a first union are not treated as competing events: when studying the occurrence of a first union, a pregnancy or a child are treated as explanatory variables among others; similarly, a union or a marriage are explanatory covariates for studying the occurrence of a first child. Furthermore, we censored by 1 January 1999 or by their 40th birthday, as events occurring after age 40 may not be considered as part of transition to adulthood.[7] Sample weights were used to correct for the higher non-response rates of certain population groups. In order to get

[5] The other questions were the following: First name, Sex, Date of arrival in the household for adopted children, Place of birth of the child, Date of leaving and place of residence if the child left the parental home, Live or still birth, and Age at death if the child is dead.

[6] These percentages were derived by using sample weights in order to correct for the bias related to higher non-response rates of certain population groups. The raw percentages are 82 and 78 for all women and men, respectively, entering a first union; and 72 and 60% for all women and men, respectively, who had a first child before the time of the survey.

[7] Moreover, events such as first birth and first union are very rare after the age of 40. Less than 1.3% of men and 0.9% of women born in the 1940s have a first birth after the age of 40 (they are aged 49–58 at the time of the survey); first unions are slightly more common: 2.3% of men and 1.4% of women of the same cohorts experienced a first union after the age of 40.

consistent estimates of parameter variances and likelihood ratios, we normalised the weights to an overall mean of 1 for each of the female and male samples. In any case, due to the large sample size, almost all tests led to “statistically significant” results, even for very small contrasts.

We modelled the intensity of forming a union or conceiving the first child by using a piecewise constant exponential model (Blossfeld and Rohwer 2002). We assumed the effect of age to be constant over single years of age in order to achieve maximum flexibility. Moreover, we controlled for social origin, educational level, school enrolment status, employment status and experience, and calendar period. For the entry into first union we additionally controlled for pregnancy or the presence of a child, and for parenthood we included union status in the control variables.

3.1 Explanatory Variables

In order to test our hypotheses, we focused on the effects of educational and employment variables, while controlling for individual characteristics and characteristics of the family of origin. Since we were particularly interested in the differences of these effects by gender, we ran separate model regressions for males and females.

With regard to individual characteristics, we first controlled for the effect of age. The process of family formation is highly dependent on age, and union formation rates and the rate of conception of a first child usually show a bell-shaped pattern with increasing age (Blossfeld and Huinink 1991; Buber 2002; Coppola 2004). By assuming a piecewise constant exponential model as outlined above, we incorporated age in single-year steps, as we know that age effects are very large, which was possible because of the large sample size.

For the entry into a first union, we controlled for the presence of pregnancy or a child (time varying covariate). Conception outside a union may accelerate first union formation because the father-to-be is a potential partner for pregnant women and people want to offer the child the social and economic protection of a union (Brien et al. 1999). Moreover, normative pressures may increase the incentive to ‘legitimise’ the birth (Baizán et al. 2004). Indeed, several studies show a strong positive effect of pregnancy on union formation (Goldscheider and Waite 1986; Blossfeld and Huinink 1991; Brien et al. 1999; Baizán et al. 2004; Toulemon 1997). Since there is a time lag of about 1 month between conception and the detection of pregnancy, we followed pregnancies from 1 month after conception.

However, once the child is born, its presence may impose constraints on resources and time, which may hamper union formation (Baizán et al. 2004). In fact, Brien et al. (1999) found that the risk of entering into marriage or cohabitation drops to approximately the level before conception or even below that level for US women. Baizán et al. (2004) obtained similar findings for German and Swedish women, as well as Toulemon (1997) for French women.

When investigating the entry into first parenthood, we also considered the union status, because being in a union is viewed as the appropriate setting for having children, and individuals tend to avoid having a birth out of union, in France as in Germany and Sweden (Baizán et al. 2004; Toulemon 1995). Moreover, there exists in France a preference for having the first child in a marriage rather than in a cohabiting union. Hence, when controlling for union status in the hazard regression of first conception rates, we distinguished between out of union, consensual union, and marriage.[8]

Apart from individual characteristics, we also controlled for characteristics of the family of origin, i.e. the parents' socioeconomic status and the number of siblings (both time–constant variables). Indeed, several studies show that family formation is strongly influenced by some characteristics of the family of origin. In particular the parents' socioeconomic status influences the timing of family formation. The effect is not limited to income levels or family strategies relating to property ownership or modes of consumption, which create social opportunities for children, but also comprises their social orientations, values, and beliefs, which influence family, educational and career decisions (Blossfeld and Huinink 1991). Therefore, controlling for the socio-occupational status of the father is common in analyses of family formation (Goldscheider and Waite 1986; Blossfeld and Huinink 1991; Leridon and Toulemon 1995; Thornton et al. 1995). In particular, we distinguished between farmers, self-employed persons, unskilled workers, skilled workers, low-level white-collar, medium-level white-collar, and high-level white-collar employees.

We also tried to incorporate the mothers' socio-occupational status. However when controlling for the other characteristics, the effects turned out to be minor and mostly insignificant and were therefore not included for reasons of parsimony.[9]

Moreover, we took into account the number of siblings, because there is empirical evidence that individuals in large households tend to enter a first union and first parenthood earlier (Blossfeld and Huinink 1991; Billari and Philipov 2004). We considered zero, one, two, three, four or five and more siblings.

Our main explanatory variables were educational level, school enrolment and working status (all time-varying covariates). Concerning the educational level, we distinguished between primary (including no qualifications), secondary and tertiary education. Within the secondary level of education, we distinguished further between short secondary studies *(brevet d'études du premier cycle*, BEPC, including *brevet élémentaire* and *brevet des collèges*), long vocational apprenticeship (*certificat d'aptitude professionnel*, CAP), long vocational studies (*brevet d'études professionnelles*, BEP), and completed secondary studies (*baccalauréat*).

In France, schooling between ages 6 and 16 has been compulsory since 1967, and was compulsory from 6 to 14 before 1967. Primary schooling lasts from ages 6 to 11, and lower secondary education from ages 11 to 15. At the end of lower secondary education, pupils can take the *brevet* exam (BEPC and its predecessors, the *brevet*

[8] Several authors found that family-building behaviours such as union formation, particularly marriage and conception, are interrelated and estimation procedures not taking into account the possible endogeneity may lead to a biased interpretation of the effects they have on each other (see, e.g. Brien et al. 1999 for modelling the interrelations between cohabitation, marriage and non-marital conception). We tested for such endogeneity by comparing models on first births including or excluding conjugal status.

[9] The Bayesian Information Criteria improved when we left out the mothers' socio-occupational status.

élémentaire and *brevet des collèges*). Before compulsory schooling was prolonged, the vast majority of children attended extended primary education comprising eight grades, which led to the *certificat d'études primaires* (Grenet 2004). In the upper secondary level, children either enter a *lycée*, which ends with the *baccalauréat*, or start vocational training, which leads either to a BEP or a CAP (Eurydice 2005).

We do not have the complete educational history, but only the highest level attained and the respondents' age at the end of studies. When constructing the time-varying educational level we therefore assumed that students reach the BEPC by age 15, the *baccalauréat* by age 18, and the first university level by age 21.

Concerning the school enrolment status, we used the answers to the question about the age at which the respondent left school or university "for the first time". As adult education is rare in France, this information is considered as sufficient to distinguish periods of enrolment in education and periods following the end of initial studies. We extended the commonly used dichotomous variable by taking into account the number of years after leaving school, since this has been shown to have a significant effect, at least on entry into motherhood (Buber 2002). As the end of studies is a foreseeable event, at least in the final year of enrolment, we distinguished between more than 1 year before finishing school and the last year of schooling. Therefore, we distinguished between enrolled and more than 1 year before finishing education, being in the final year of education, and first, second, third and more than 3 years after leaving school.

Our third main explanatory variable was employment status, i.e. working versus not working. The form included a question on the year of first employment (for a duration of at least 3 months, in order to exclude summer jobs), as well as the occurrence of work interruptions of at least 2 years. In order to incorporate the speed and difficulties of moving into a stable work career, we took into account the time elapsed since the person started his/her first job in the employment status. Hence, we distinguished between not working, and having worked for one, two, three and more than 3 years.

As enrolment in education and working status are strongly correlated, the covariance of the estimates is likely to be very large; furthermore, many interactions were present in our preliminary models, especially between gender, duration since the end of the studies and duration since first job. Furthermore, as enrolment in education and employment are strongly correlated, describing the changes by, for example, working status and work experience, "controlling for enrolment in education", seemed meaningless. Thus we combined covariates on enrolment in education, working status, duration to or since the end of studies, and working experience. However, if we had incorporated all interactions between the different categories of enrolment and employment status, we would have had to estimate an additional 20 coefficients. Therefore, we combined interaction categories in a meaningful way, which is also confirmed by the fact that the Bayesian Information Criteria (BIC) improved relative to a model with two covariates on enrolment in education and working status, as well as relative to the model including all the interactions (result not shown).

We built a 10-category variable, which distinguishes four categories among non-working individuals: (1) enrolled in education and more than 1 year to graduation, (2) final year in school, (3) left school less than 3 years ago and (4) left school 3 or more years ago. For those working, we differentiated between two categories of individuals also enrolled in education: (5) still enrolled in education and more than 1 year until graduation, (6) final year in school. Finally, we distinguished four categories of working individuals, depending on their work experience: (7) left school and working 1st year, (8) left school and working 2nd year, (9) left school and working 3rd year, and (10) left school and working 3 or more years.

Finally, our model included the calendar period effects (time-varying covariate) by splitting calendar time into 5-year groups, because we are interested in time trends over 40 years and not in year-to-year changes. As we only considered birth cohorts born after 1940, there were few events in the years 1955–1960. Therefore, we combined the second half of the 1950s with the first half of the 1960s. Hence, we distinguished between the calendar periods 1955–1964, 1970–1974, 1975–1979, 1980–1984, 1985–1989, 1990–1994, and 1995–1998.

In order to detect changes in the assumed effects over time, we modelled interactions with the period variable and the other explanatory variables. We employed linear splines with a node at 1975–1979 to lower the number of coefficients that had to be estimated. Taking the calendar period 1975–1979 as a reference category, we estimated the average change of the covariates' effect from 1955–1964 to 1975–1979 and the average change from 1975–1979 to 1995–1998.[10]

The models were estimated by maximum likelihood, using the statistical software package STATA (StataCorp 2006). Model selection was based on the likelihood ratio tests and Bayesian Information Criteria (BIC), since this latter test also enables us to compare non-nested models, and is more accurate than the likelihood ratios test, which led to systematically significant results, because of the large sample size.

Using the normal approximation for the parameters, we could test the differences between similar parameters for men and women. Due to the very large sample size (see Tables 1 and 4), all differences presented between trends for men and women appear to be significant at the 5% level.

4 Results

4.1 Union Formation

Table 1 shows the parameter estimates for the entry into a first union for men and women, as well as the proportions of exposure (person-years) and events. We controlled for age, number of siblings, father's socio-occupational status, pregnancy, level of education, and calendar period. Enrolment in education and working status are combined into a single variable which also includes duration to or since the end of studies, as well as duration since first job. The figures in Table 1 are the estimated exponentials of the coefficients of the regression on the log hazard. Thus

[10] We could have applied linear splines to single years in order to estimate the interaction between our covariates of interest and time, but we considered that seven 5-year periods were sufficient.

Table 1 Proportions of total exposure and proportions of events in particular states before first union formation and model estimates for hazards of entry into first union

Covariate	Women			Men		
	Exposure	Events	Hazard	Exposure	Events	Hazard
Age						
15	12.2	0.5	0.01	9.9	0.1	0.00
16	12.1	1.7	0.03	9.9	0.2	0.00
17	11.9	4.7	0.08	9.8	0.9	0.01
18	11.2	9.7	0.13	9.7	2.5	0.03
19	9.9	13.3	0.17	9.3	4.9	0.05
20	8.2	14.4	0.20	8.6	9.0	0.08
21	6.6	13.3	0.20	7.6	13.3	0.13
22	5.2	10.9	0.19	6.4	14.0	0.15
23	4.0	8.3	0.17	5.2	13.0	0.16
24	3.2	6.2	0.15	4.2	10.6	0.16
25	2.5	4.4	0.13	3.4	8.1	0.15
26	2.1	3.1	0.11	2.7	6.0	0.13
27	1.7	2.3	0.10	2.2	4.5	0.12
28	1.4	1.7	0.09	1.8	3.2	0.10
29	1.2	1.3	0.08	1.5	2.4	0.09
30	1.0	0.9	0.06	1.3	2.0	0.09
31	0.9	0.7	0.06	1.1	1.3	0.07
32	0.8	0.6	0.05	1.0	1.0	0.06
33	0.7	0.5	0.05	0.9	0.8	0.05
34	0.6	0.3	0.04	0.8	0.6	0.05
35	0.6	0.3	0.04	0.7	0.5	0.05
36	0.5	0.2	0.03	0.6	0.4	0.04

Table 1 continued

Covariate	Women			Men		
	Exposure	Events	Hazard	Exposure	Events	Hazard
37	0.5	0.2	0.03	0.5	0.3	0.04
38	0.4	0.1	0.03	0.5	0.3	0.04
39	0.4	0.1	0.02	0.4	0.2	0.03
Siblings						
0	10.4	8.7	0.87***	10.4	8.8	0.81***
1 (reference)	24.6	22.3	1	24.2	23.0	1
2	23.2	22.9	1.03**	23.4	23.2	1.02
3	14.7	15.5	1.02*	14.7	15.4	1.02
4	9.2	10.1	1.01	9.2	10.0	1.04*
5 and more	17.9	20.6	1.01	18.0	19.6	1.01
Socio-occupational status of father						
Inactive	0.5	0.5	1.00	0.5	0.4	0.99
Farmer	11.0	11.5	0.98	12.1	11.9	0.88***
Self-employed	12.7	12.4	1.01	12.1	12.6	1.03
Unskilled worker	15.0	16.8	1.02	15.5	16.1	0.99
Skilled worker (reference)	21.4	22.7	1	21.8	22.0	1
Low-level white-collar	15.8	15.7	0.98	15.9	16.2	1.02
Medium-level white-collar	12.6	11.3	1.01	11.8	11.3	1.04*
High-level white-collar	10.3	8.5	0.98	9.7	8.9	1.04*
Pregnancy						
Not pregnant/No child	94.3	82.7	1	97.4	85.2	1
Pregnant (reference)	0.9	10.7	8.02***	0.5	9.6	11.80***
Child	4.7	6.6	1.17***	2.0	5.2	1.76***

Table 1 continued

Covariate		Women			Men		
		Exposure	Events	Hazard	Exposure	Events	Hazard
Level of education attained							
Primary		23.1	26.2	0.98	24.8	24.2	0.84***
Secondary	Short (BEPC) (reference)	36.4	20.6	1	36.4	24.0	1
	Long apprenticeship (CAP)	5.5	9.1	1.05**	8.2	12.5	1.04*
	Long vocational (BEP)	5.5	8.7	1.09***	5.7	7.6	1.02
	Completed (Baccalauréat)	17.5	18.9	1.14***	14.0	14.3	1.11***
Tertiary		12.0	16.5	1.56***	10.8	17.5	1.49***
Calendar period							
1955–1964		9.5	6.5	0.79***	8.3	3.3	0.79***
1965–1969		10.8	11.1	0.88***	10.4	10.4	1.01
1970–1974		12.3	15.0	0.96**	12.1	14.8	1.03*
1975–1979 (reference)		13.2	15.0	1	13.3	15.4	1
1980–1984		14.1	14.4	0.94***	14.3	15.4	0.94***
1985–1989		15.2	13.8	0.87***	15.4	14.5	0.81***
1990–1994		15.1	14.0	0.92***	15.6	14.8	0.81***
1995–1998		9.9	10.1	0.88***	10.7	11.3	0.77***
School enrolment and employment status interaction							
Not working	More than 1 year to graduation	38.4	8.2	0.24***	30.0	5.1	0.35***
	In final year of school	8.2	6.6	0.65***	6.6	2.7	0.58***
	Left school less than 3 years ago	2.0	1.9	0.94***	2.3	1.8	0.70***
	Left school more than 3 years ago	0.7	1.0	0.79***	0.8	0.7	0.38***

Table 1 continued

Covariate		Women			Men		
		Exposure	Events	Hazard	Exposure	Events	Hazard
Working	More than 1 year until graduation	5.3	6.6	0.51***	4.3	3.0	0.65***
	In final year of school	2.2	2.5	0.82***	1.1	0.5	0.76***
	Left school; working 1st year (reference)	7.1	10.7	1	6.6	5.7	1
	Left school; working 2nd year	6.4	11.1	1.08***	6.6	6.7	1.05
	Left school; working 3rd year	5.4	11.0	1.16***	6.2	7.9	1.11***
	Left school; working 3 and more years	24.2	40.5	1.22***	35.5	65.9	1.25***
Person-years of exposure		1,186,797			964,890		
Events			118,244			69,912	
Log-likelihood				−107,013			−55,950
Bayesian information criteria				214,921			112,782

*$p < 0.05$, **$p < 0.01$, ***$p < 0.001$

they estimate the relative risk for any category, compared to a reference category for which the risk is set to 1. Figure 1 shows the estimates for men and women according to our enrolment in education and working status variable, as presented in Table 1.

Most of the covariates display the expected effects. With increasing age, the union formation rates exhibit a bell-shaped pattern for both men and women, with female union intensities reaching their highest values about 2 years earlier than those for males (age 21 vs. age 23). Moreover, while young women display higher union intensities than young men, from their mid-twenties onwards, men and women show similar values for the risk of entering into a first union.

Growing up with at least one sibling also has a positive impact on union formation. This holds true for men and women, though being the only child, has a slightly more negative effect on union formation for men than for women. However, the presence of more than one sibling has only a minor and barely significant effect.

The father's occupation has almost no significant impact on men's and women's union formation, except for men whose fathers are farmers. Being a farmer's son reduces union formation intensities by 12%. This may be due to the fact that such men have a higher probability of becoming farmers themselves compared to children from other social backgrounds. Moreover, the risk of marrying is lower for farmers than for men in other occupations (Courgeau and Lelièvre 1986). Furthermore, men whose fathers are medium or high-level white-collar workers have slightly higher union formation rates (about 4%), while this is not the case for women. In contrast, being the daughter of a low-level white-collar worker has a significant, slightly negative effect on the girl's union formation (about 3%).

The presence of a pregnancy strongly accelerates union formation. In this case, men and women show respectively a 12-fold and 8-fold higher union formation rates than before conception. After the birth of the child, union formation intensities steeply decrease. In fact, the union formation rates for men and women are 76 and 17%, respectively, higher than before conception. The gender difference may be due to the fact that some men who did not form a union with the mother of their first child failed to report this child. If this were the case, the estimates would show an upward bias.

Educational attainment has a significant, positive effect on union formation intensities for men and women. However, low education (primary education only or no qualifications) has a significant, negative impact on union formation for men, while there is no significant effect for women. Specifically, low-educated men have about 16% lower union formation rates than those with a BEPC. In contrast, men and women with a CAP diploma show about 4% higher union formation intensities than those with a BEPC diploma. Furthermore, holding a BEP diploma increases the union formation rates for women by about 9%, while there is no significant effect for men. Having the *baccalauréat* raises the risk of entering into a first union by about 13%, while a university degree increases union formation rates by 54% for both men and women.

In line with our hypothesis, we found that the effect of enrolment in education is stronger for women than for men. In particular, enrolment decreases union formation rates by 71 and 43% for those women and men, respectively, who still have more than 1 year until graduation. In the final year of education, union formation rates are about 27 and 11% lower for female and male students, respectively.

Being in education has a significant negative impact on union formation for both men and women, with a lower effect in the final year: the risk of entering into a first union decreases by about 76% for non-working women and by 65% for non-working men as compared to those who just finished school and entered employment (reference category for our variable combining enrolment in education and working status, see Table 1 and Fig. 1). In contrast, working students have about 49 and 35% lower union intensities for women and men, respectively. The negative impact is much lower in the final year of education: non-working students exhibit union formation intensities which are only about 35 and 42% lower for women and men, respectively, while female and male students who work enter a first union about 18 and 24% less often in their final year of school than those who just finished school and started their first job (Fig. 1). By distinguishing between working status for those enrolled, we could show that those who already combine student and working roles enter a union more often than non-working students, but less frequently than those who are not enrolled and work. However, due to data limitations, we neither distinguished between full-time and part-time education nor between full-time and part-time employment, and those who combine school and work probably might be involved part-time in at least one of these activities.

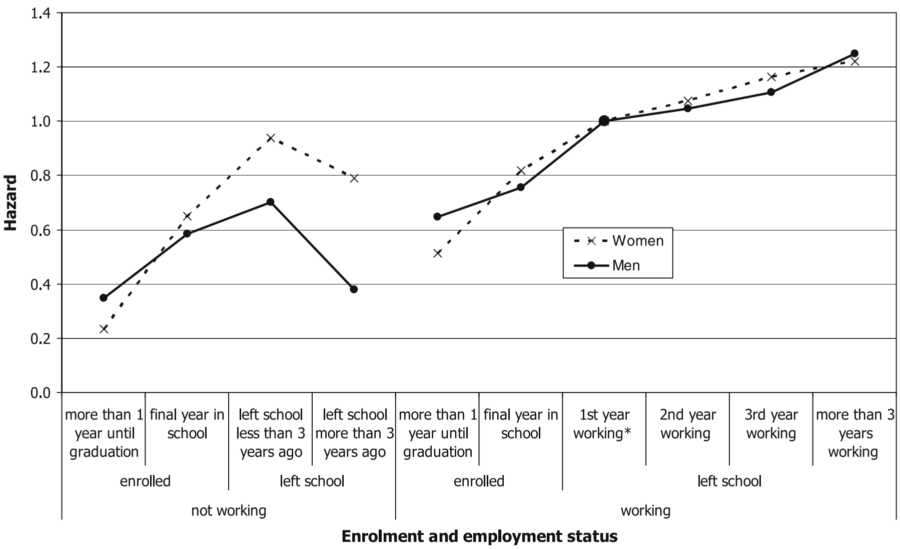

Fig. 1 Relative risks of union intensities by enrolment in education and working status interaction, as estimated by the model presented in Table 1. The asterik denotes the baseline category

Concerning the difficulties in the transition to employment, our results demonstrate that women who did not take up work during the first 3 years after leaving school show only slightly lower union formation rates than those who began working (−6%). However, if the period between leaving school and entering employment is longer than 3 years, the negative effect increases to 21%. In contrast, the status of "left school but not yet working" has a much stronger negative impact on union formation intensities for men. Indeed, union formation rates for non-working males who left school less than 3 years ago are about 30% lower, and union formation intensities are about 62% lower for those who left education more than 3 years ago. The latter effect is of a similar order as the one for non-working students.

The gender difference confirms our hypothesis, which is in line with the arguments of Oppenheimer and colleagues who claim that men's career and career maturity play a more important role for the timing of marriage than for women. Similarly, Goldscheider and Waite (1986) found that employment has a stronger impact on the marriage intensities of men than on those of women.

Among working non-student men and women, we found that the more work experience they have, the higher their union formation rates. The increase is almost linear with duration since first job, those having worked for three and more years experiencing union formation rates higher than those in the first year of employment by about 22 and 26% for women and men, respectively. Conversely to our hypothesis, we found almost no difference between male and female non-students with regard to the effect of work experience (Fig. 1).

Finally, we found a hump-shaped pattern for calendar period effects on union formation for men and women, with rather stable estimated effects since 1985.

In order to investigate trends in the gender differences regarding the effects of level of education, enrolment and employment status, we considered how these explanatory variables interacted with the calendar period.

Table 2 shows the estimated coefficients for the interaction between the effect of educational attainment and calendar period. In order to better evaluate whether the social differences among men and women become more similar, we plotted in Fig. 2 the interaction effects relative to the baseline level (short secondary studies BEPC) within selected periods, i.e. 1955–1964, 1975–1979, and 1995–1998. Men and women having a BEPC diploma entered a first union less frequently in 1955–64 and 1995–98 as compared to the mid-1970s. Moreover, in the most recent period, men with a BEPC formed a first union slightly less frequently than their female peers.

As can be seen in Fig. 2, the effect of low and medium education levels clearly changed over time for women, while the relative effect of a university degree is pretty stable for them: the interaction is significant for women and men, but according to the BIC, the interaction does not improve the model for men. In the earliest period, having a primary school diploma or no qualifications at all even had a positive effect on female union formation as compared to having a BEPC. Hence, the effect of the educational level was slightly U-shaped during 1955–1964. However, between 1955–1964 and 1975–1979, the effect of only having a primary school diploma or no qualifications at all turned negative as compared to having a

Table 2 Estimated effects of the interaction between educational level and time period on entry into first union (relative risks compared to the same reference category for each 5-year period)

Educational level		1955–1964	1965–1969	1970–1974	1975–1979	1980–1984	1985–1989	1990–1994	1995–1998
Women									
Primary		1.16	1.07	0.98	0.90	0.91	0.93	0.94	0.95
Secondary	Short (BEPC) (reference)	1	1	1	1	1	1	1	1
	Long apprenticeship (CAP)	1.13	1.08	1.04	1.00	1.02	1.04	1.06	1.08
	Long voc. studies (BEP)	1.06	1.06	1.06	1.06	1.07	1.09	1.11	1.13
	Completed (Baccalauréat)	1.09	1.08	1.07	1.06	1.11	1.15	1.20	1.25
Tertiary		1.71	1.60	1.49	1.39	1.46	1.53	1.61	1.69
Men									
Primary		0.99	0.91	0.84	0.78	0.80	0.83	0.85	0.88
Secondary	Short (BEPC) (reference)	1	1	1	1	1	1	1	1
	Long apprenticeship (CAP)	1.10	1.07	1.04	1.02	1.03	1.03	1.04	1.04
	Long voc. studies (BEP)	1.10	1.06	1.03	1.00	1.00	1.01	1.01	1.02
	Completed (Baccalauréat)	1.22	1.17	1.12	1.08	1.09	1.10	1.10	1.11
Tertiary		1.62	1.56	1.51	1.46	1.46	1.47	1.47	1.48
Time trend for reference category									
Women with short secondary studies (BEPC)		0.69	0.81	0.93	1	0.92	0.83	0.85	0.78
Men with short secondary studies (BEPC)		0.70	0.94	1.00	1	0.93	0.79	0.79	0.74

Note: The BIC of the models including the interaction between educational level and time period amounts to 214,889 for women and 112,891 for men. The improvement is significant for women, but not for men

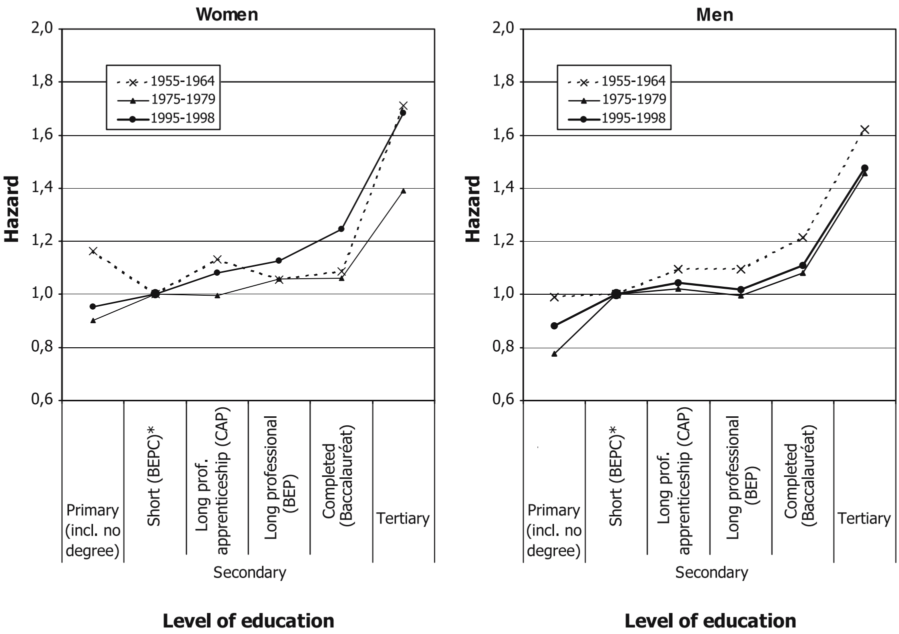

Fig. 2 Estimated effects of level of education attained on entry into first union for selected years (relative risks without main period effect). The asterisk denotes the baseline category

BEPC, and since that time, we observe that the effect of education on entry into a first union is also increasing for women.

Union formation rates increase with level of education, from primary to secondary and tertiary level for all periods. Similar to the female union formation rates, the relative risk of entering a union for men with the lowest level of education strongly decreased between 1955–1964 and 1975–1979.

Summing up, we may thus say that our first hypothesis was partly confirmed: as expected, the U-shaped relationship between level of education and union formation rates for women was replaced by a monotonous positive relationship in 1975–79 and later. Nevertheless, this trend was only due to a relative decline in union formation rates of women with a primary education, and we noted an increasing impact of the educational level on union formation rates for both men and women, and therefore, no gender differences, except for the lowest educational level.

Concerning the time trends in the effects of enrolment in education and working status on union formation rates, we found more pronounced gender differences. Since enrolment in education and labour force participation have undergone tremendous change in the past decades, we also expected more pronounced changes in the effect of school enrolment and working status on union formation rates.

Table 3 summarises the estimated effects of the interaction between calendar period, enrolment in education and employment status. Male school graduates

Table 3 Estimated effects of the interaction between school enrolment × working status and time period on entry into first union (relative risks compared to the same reference category for each 5-year period)

School enrolment × Working status		1955–1964	1965–1969	1970–1974	1975–1979	1980–1984	1985–1989	1990–1994	1995–1998
Women									
Not working	More than 1 year to graduation	0.20	0.22	0.23	0.25	0.24	0.23	0.22	0.21
	In final year of school	0.58	0.62	0.66	0.70	0.67	0.64	0.61	0.58
	Left school less than 3 years ago	1.18	1.14	1.10	1.06	0.95	0.85	0.76	0.68
	Left school more than 3 years ago	1.88	1.34	0.95	0.68	0.63	0.59	0.54	0.51
Working	More than 1 year to graduation	0.65	0.61	0.57	0.53	0.50	0.48	0.45	0.43
	In final year of school	1.11	1.01	0.91	0.83	0.79	0.75	0.72	0.68
	Left school; working 1st year (ref.)	1	1	1	1	1	1	1	1
	Left school; working 2nd year	1.31	1.24	1.18	1.12	1.06	1.01	0.96	0.92
	Left school; working 3rd year	1.73	1.51	1.32	1.15	1.09	1.03	0.98	0.92
	Left school; working 3+ years	2.16	1.72	1.37	1.09	1.06	1.03	1.00	0.97
Men									
Not working	More than 1 year to graduation	0.35	0.36	0.36	0.36	0.36	0.35	0.35	0.34
	In final year of school	0.75	0.69	0.63	0.58	0.57	0.56	0.55	0.54
	Left school less than 3 years ago	0.98	0.87	0.77	0.68	0.68	0.67	0.67	0.66
	Left school more than 3 years ago	0.76	0.59	0.47	0.37	0.36	0.35	0.35	0.34
Working	More than 1 year to graduation	0.78	0.73	0.69	0.64	0.64	0.64	0.63	0.63
	In final year of school	1.04	0.98	0.92	0.86	0.79	0.72	0.66	0.61
	Left school; working 1st year (ref.)	1	1	1	1	1	1	1	1
	Left school; working 2nd year	1.09	1.06	1.03	0.99	1.02	1.05	1.08	1.11
	Left school; working 3rd year	1.20	1.16	1.12	1.08	1.09	1.09	1.10	1.11
	Left school; working 3+ years	1.86	1.55	1.29	1.08	1.13	1.19	1.25	1.31

Table 3 continued

School enrolment × Working status	1955–1964	1965–1969	1970–1974	1975–1979	1980–1984	1985–1989	1990–1994	1995–1998
Time trend for reference category								
Women; left school; working 1st year	0.54	0.68	0.86	1	0.97	0.93	1.03	1.03
Men; left school; working 1st year	0.53	0.77	0.90	1	0.90	0.75	0.73	0.68

Note: The BIC of the models including the interaction between school enrolment × working status and time period amounts to 213,912 for women and 112,843 for men. The improvement is significant for women, but not for men entered a first union less often in 1955–1964 and in 1995–98 as compared to 1975–79, while the union formation pattern of their female counterparts has remained unchanged since the mid-1970s.

Figure 3 graphically visualises these interaction effects relative to the baseline level (left school and first year working) within the calendar periods 1955–1964, 1975–1979, and 1995–1998. Here again, the likelihood ratio tests are significant, but the BIC indicates no improvement for men. The relative union rates of non-working students are almost stable during all the time periods, for men as well as for women. During the final year of studies, the rates significantly decreased over time for men, while they remained constant for women. This may reflect the growing importance of working status for men. Moreover, Robert-Bobée and Mazuy (2003) argue that this norm is particularly binding for men, since women now more often form a union before finishing their studies. This phenomenon is particularly pronounced if they enter a union with an older man, who has already established himself in his job. Hence, we find evidence for slightly increasing divergence in union formation for non-working female and male students.

Working students are in an intermediate position between non-working students and working non-students; compared to these two groups, they are increasingly reluctant to enter a union: in the late 1990s, the relative risks were around 0.6 for working students, compared to the reference value of 1 for young men and women in their first year of work.

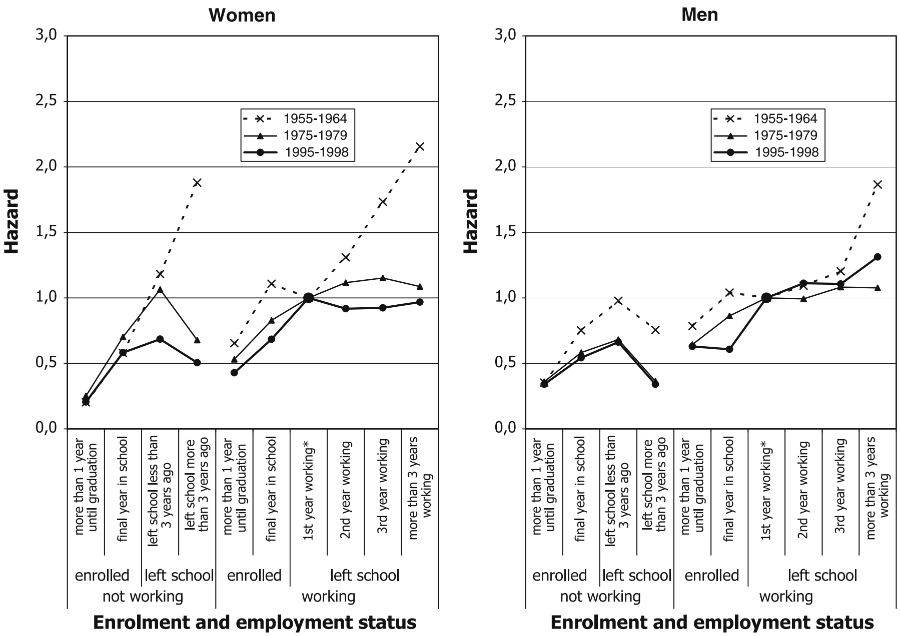

Fig. 3 Estimated effects of interaction between enrolment in education and employment status on entry into first union for selected years (relative risks without main period effect). The asterisk denotes the baseline level

Work experience is less and less associated with an increasing propensity to enter a union. Before 1965, work experience had a positive effect on union formation rates for both sexes. The less steep increase in the initial years of employment for men than for women may reflect the difficulties encountered in obtaining a stable job, and the higher importance of work for men as outlined by Oppenheimer and Lewin (1999). However, the impact of longer work experience diminished over time, and even disappeared for women.

The main change with time that we see in Fig. 3 is the dramatic decline of union rates of young adults who have already finished their studies but who are not working, compared to those who are already working. This decline is larger for women, who had a higher risk of entering a union when they did not work before 1965, while the risk became lower in the late 1990s, compared to working women. In the last period the union rates decline with duration since the end of studies for men and women who are not working.

Thus, all in all, the gender differences are much smaller in the last period than in the first. The convergence hypothesis is again confirmed, but this is not the case for our second and third hypotheses, when age and the other covariates are controlled for: the contrast in union rates among students and working adults is stable, and not increasing with time (hypothesis 2), and job experience has a decreasing effect on union rates (hypothesis 3). Most of the exposures and events concern men and women enrolled in education and not working, or working after having left school (see Table 1), and the main conclusion for the union rates of men and women in these two situations is that the contrasts are declining for women and increasing for men, in line with the convergence hypothesis.

4.2 First Parenthood

Table 4 shows the parameter estimates for the entry into parenthood for men and women. We controlled for age, number of siblings, father's socio-occupational status, union status, level of education, the interaction of enrolment in education and working status, and calendar period.

Most of the covariates had the expected effects. With growing age, the first conception rates exhibit a bell-shaped pattern for both men and women. Gender differences are particularly pronounced before age 25, with women showing higher rates of having a first child. Above age 30, men's chances of having a first child are somewhat higher.

The size of the family of origin has a significant impact on the timing of first births for both men and women. The larger the family of origin, the higher the risk of having a first child. In contrast, the father's socio-occupational status has almost no effect if we control for the respondent's socioeconomic characteristics.

Union status has a strong effect on the risk of having a first child. Individuals who do not live in a union tend to avoid a pregnancy. Compared to a consensual union, the hazard of conceiving a first child is 77% lower for women, and 82% lower for men who do not live in such a union. The risk of having a first child is two times higher for married men and women than for cohabiting partners.

Table 4 Proportions of total exposure spent and proportions of events in particular states before conception leading to first birth and model estimates for hazards of entry into first parenthood

Covariate	Women			Men		
	Exposure	Events	Hazard	Exposure	Events	Hazard
Age						
15	9.9	0.6	0.04	8.2	0.1	0.01
16	9.9	1.7	0.08	8.2	0.2	0.01
17	9.7	3.7	0.13	8.2	0.7	0.03
18	9.4	6.3	0.15	8.1	1.7	0.05
19	8.7	8.3	0.14	7.9	2.9	0.08
20	7.9	9.3	0.12	7.5	4.5	0.09
21	7.0	9.9	0.11	7.0	7.6	0.11
22	6.1	10.0	0.10	6.5	9.4	0.10
23	5.2	9.3	0.09	5.8	10.2	0.10
24	4.4	8.7	0.09	5.1	10.0	0.09
25	3.7	7.4	0.09	4.4	10.3	0.09
26	3.0	6.0	0.08	3.8	8.6	0.09
27	2.5	4.7	0.08	3.2	7.8	0.09
28	2.1	3.7	0.07	2.7	6.4	0.08
29	1.7	2.8	0.07	2.2	5.0	0.08
30	1.5	2.1	0.06	1.9	3.8	0.07
31	1.3	1.5	0.05	1.6	2.9	0.07
32	1.1	1.1	0.04	1.4	2.3	0.06
33	1.0	0.9	0.04	1.2	1.6	0.05
34	0.9	0.6	0.03	1.1	1.3	0.04
35	0.8	0.4	0.02	1.0	1.0	0.04

Table 4 continued

Covariate	Women			Men		
	Exposure	Events	Hazard	Exposure	Events	Hazard
36	0.7	0.4	0.02	0.9	0.6	0.03
37	0.6	0.3	0.02	0.8	0.5	0.02
38	0.6	0.2	0.01	0.7	0.4	0.02
39	0.5	0.1	0.01	0.6	0.3	0.02
Siblings						
0	10.2	8.4	0.97*	10.2	8.5	0.91***
1 (reference)	25.3	20.2	1	24.5	21.2	1
2	23.8	22.0	1.09***	23.7	22.4	1.05**
3	14.8	15.9	1.16***	14.8	15.9	1.14***
4	9.0	10.8	1.25***	9.2	10.7	1.18***
5 and more	16.9	22.7	1.31***	17.6	21.3	1.24***
Socio-occupational status of father						
Inactive	0.5	0.5	1.06	0.4	0.5	1.36**
Farmer	10.8	12.3	1.00	11.7	12.9	1.03
Self-employed	12.9	12.3	1.02	12.3	12.5	1.01
Unskilled worker	14.6	17.7	1.02	15.3	16.8	1.02
Skilled worker (reference)	21.2	23.1	1	21.7	22.1	1
Low-level white-collar	15.9	15.8	0.98	16.1	16.1	0.97
Medium-level white-collar	12.9	10.3	0.97*	12.0	10.3	0.97
High-level white-collar	10.6	7.5	0.99	9.9	8.3	1.00
Union status						
Not in union	79.0	28.1	0.23***	82.8	24.4	0.17***
Cohabitating (reference)	10.6	20.6	1	9.0	21.6	1

Table 4 continued

Covariate		Women			Men		
		Exposure	Events	Hazard	Exposure	Events	Hazard
Married		10.4	51.3	2.19***	8.2	53.9	2.32***
Level of education attained							
Primary		21.0	29.1	1.06***	23.9	25.9	0.98
Secondary	Short (BEPC) (reference)	32.6	21.5	1	34.0	25.4	1
	Long apprenticeship (CAP)	6.2	9.8	1.04*	8.9	12.7	1.01
	Long vocational (BEP)	6.4	8.9	1.05***	6.1	7.2	1.04
	Completed (Baccalauréat)	18.7	16.4	0.97**	14.3	12.9	0.97
Tertiary		15.1	14.4	1.14***	12.7	15.9	1.23***
Calendar period							
1955–1964		7.9	6.4	1.29***	7.0	2.9	1.42***
1965–1969		9.5	11.3	1.25***	9.2	10.6	1.48***
1970–1974		11.4	15.3	1.26***	11.2	15.3	1.26***
1975–1979 (reference)		13.0	14.8	1	13.0	16.0	1
1980–1984		14.5	14.7	0.97*	14.6	15.4	0.89***
1985–1989		15.8	14.1	1.01	16.0	14.9	0.88***
1990–1994		16.4	13.9	1.05***	16.8	14.8	0.90***
1995–1998		11.5	9.5	0.99	12.2	10.2	0.82***
School enrolment in education and employment status interaction							
Not working	More than 1 year to graduation	32.2	2.5	0.17***	25.4	1.7	0.31***
	In final year of school	7.2	3.3	0.74***	5.7	1.2	0.69***
	Left school less than 3 years ago	2.1	0.8	1.33***	2.2	1.0	0.81***
	Left school more than 3 years ago	0.7	0.6	1.53***	0.7	0.5	0.66***

Table 4 continued

Covariate		Women			Men		
		Exposure	Events	Hazard	Exposure	Events	Hazard
Working	More than 1 year until graduation	4.7	4.9	0.45***	3.8	1.3	0.67***
	In final year of school	1.9	2.9	0.92	1.0	0.3	0.99
	Left school; working 1st year (reference)	6.8	6.0	1	6.0	2.9	1
	Left school; working 2nd year	6.6	8.5	1.25***	6.1	4.4	1.23***
	Left school; working 3rd year	6.0	10.1	1.41***	6.0	5.7	1.33***
	Left school; working 3 and more years	31.8	60.6	1.62***	43.1	81.0	1.64***
Person-years of exposure		1,459,650			1,159,670		
Events			101,005			56,506	
Log-likelihood				−80,849			−36,697
Bayesian information criteria				162,608			74,287

*$p < 0.05$, **$p < 0.01$, ***$p < 0.001$

The effect of the level of education is U-shaped for females; women with a primary education or no qualifications and university graduates have the highest rates of entry into motherhood. Male university graduates also have significantly higher first conception rates compared to their peers with other educational levels. However, being enrolled in education significantly lowers the risk of entering parenthood for men and women, with the effect being stronger for females than for males.

Among those who are no longer enrolled in education, working status has a negative effect for women, but a positive one for men. Hence, non-working women show higher rates of having a first child than their peers during their first year of work. However, for men, entering the labour market seems to be a prerequisite for their transition to fatherhood. The time spent on the labour market speeds up the transition into parenthood for those who are working, for both men and women.

As in the analysis of union formation, we modelled the calendar period for the interaction using a spline with a node in the period 1975–1979. The estimated slope of spline segments yields the average change of the effect over time for the respective period relative to 1975–1979.

Table 5 depicts the educational level effects for men and women for all 5-year periods, while the contrasts by education for the three periods 1955–64, 1975–79 and 1995–98 are shown on Fig. 4. Here again, the likelihood ratios are significantly different, but according to the BIC the interaction may be neglected for men. Nevertheless, the effects of educational level for men and women may be considered similar, and changed in the same way over time contrary to our hypothesis. In 1955–1964, highly educated men and women quickly progressed into parenthood. Since the mid-1970s, there has no longer been any specific effect of educational level for men and women. Men and women with a tertiary education increasingly want to enjoy life as a couple before having a first child (Table 6).

Figure 5 shows the interaction of period with enrolment in education and employment status effects on first birth rates. Both likelihood ratio test and BIC indicate that the interaction improves the models for men as well as for women, despite the fact that, for women, the contrasts appear to be almost stable with time. The main significant change concerns the effect of work experience: in the recent period, working women tend to delay the arrival of their first child, while in the 1960s they used to delay their first union. Similarly, enrolment in education is increasingly associated with lower fertility. All in all, these interactions show that young women increasingly tend to live as a couple without having a first child: the delay in the end of studies and first job leads to a delay in first unions; in addition, controlled for all the other covariates, young women delay the arrival of their first child until they have some working experience, probably linked to a more stable working status.

A similar interaction was found for men's entry into parenthood, with a larger increase of the contrasts with time: first birth rates decrease for students, compared to men working for the first year, and increase with work experience, in the most recent period.

Our fourth hypothesis is thus confirmed: union rates are less and less sensitive to enrolment in education and working status, and the contrary is true for first birth

Table 5 Estimated effects of the interaction between educational level and time period on entry into first parenthood (relative risks compared to the same reference category for each 5-year period)

Educational level		1955–1964	1965–1969	1970–1974	1975–1979	1980–1984	1985–1989	1990–1994	1995–1998
Women									
Primary		1.02	1.04	1.05	1.07	1.08	1.09	1.10	1.11
Secondary	Short (BEPC) (reference)	1	1	1	1	1	1	1	1
	Long apprenticeship (CAP)	1.08	1.05	1.03	1.01	1.02	1.04	1.06	1.08
	Long voc. studies (BEP)	1.28	1.17	1.06	0.97	1.00	1.04	1.07	1.11
	Completed (Baccalauréat)	1.27	1.14	1.02	0.92	0.93	0.94	0.95	0.96
Tertiary		1.53	1.40	1.29	1.18	1.15	1.12	1.09	1.06
Men									
Primary		1.04	1.01	0.99	0.96	0.97	0.98	0.99	0.99
Secondary	Short (BEPC) (reference)	1	1	1	1	1	1	1	1
	Long apprenticeship (CAP)	1.14	1.08	1.03	0.98	0.98	0.98	0.99	0.99
	Long voc. studies (BEP)	1.08	1.07	1.06	1.05	1.04	1.02	1.00	0.99
	Completed (Baccalauréat)	1.31	1.17	1.05	0.94	0.93	0.93	0.93	0.92
Tertiary		1.73	1.56	1.40	1.25	1.22	1.18	1.15	1.12
Time trend for reference category									
Women with short secondary studies (BEPC)		1.25	1.20	1.22	1	0.97	1.01	1.05	0.99
Men with short secondary studies (BEPC)		1.29	1.37	1.21	1	0.90	0.89	0.91	0.84

Note: The BIC of the models including the interaction between educational level and time period amounts to 162,108 for women and 73,871 for men. The improvement is significant for women, but not for men

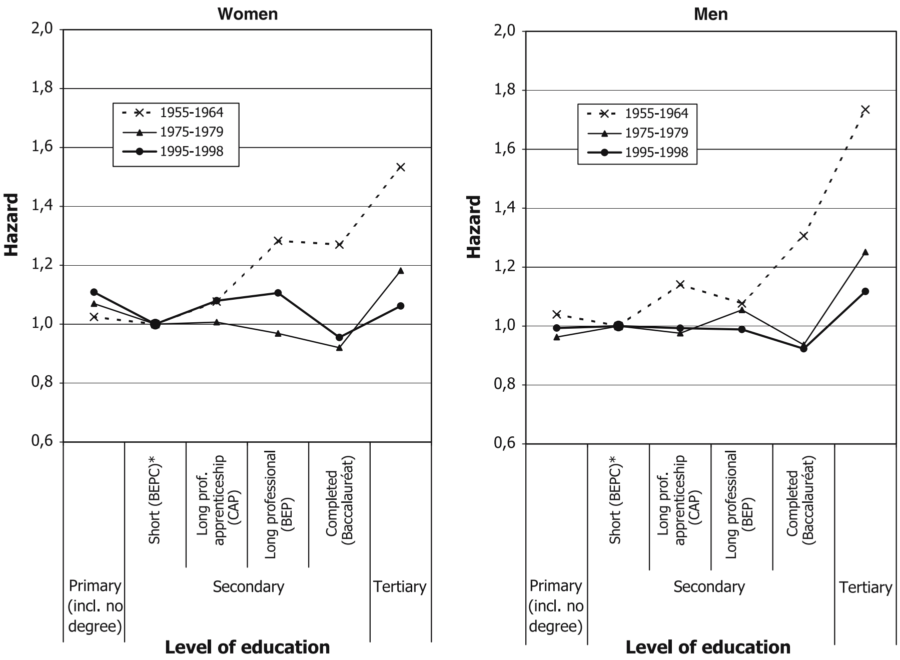

Fig. 4 Estimated effects of level of education attained on entry into first parenthood for selected years (relative risks without main period effect). The asterisk denotes the baseline category

rates. The spread of unmarried cohabitation, and the possibility of living as a couple without having a child, has moved the stakes relative to family formation from entering a first union to having a first child. These two transitions are not so closely related as they used to be, especially among highly educated men and women, whose relative propensity to have a first child, controlled for all other covariates, has declined.

As first births out of any union remain rare in France (Toulemon 1995), entering a first union and having a first child may be considered as two related events, and entering a first union may be considered as endogenous to the desire to have a first child. We thus checked whether removing union status from the covariates led to a change in the apparent effect of many of the covariates. The main result, for men as well as for women, is that the time trend changed: part of the decline in first birth rates in the recent period may be "attributed" to the decline in union rates, and the stability shown in Table 4 after 1975–79 in first birth rates is replaced by an ongoing decline when union status is not controlled for. The contrast between men and women still enrolled in education or already working is also increasing, and a negative impact of primary education appears for men. This shows that for men, part of the higher first birth rates of working men and men with a higher level of education comes from their higher rates of first union formation. For women, on the contrary, having a first job is linked to a delay in first union and in parenthood, especially during the first working years.

Table 6 Estimated effects of the interaction between school enrolment × working status and time period on entry into first parenthood (relative risks compared to the same reference category for each 5-year period)

School enrolment × Working status		1955–1964	1965–1969	1970–1974	1975–1979	1980–1984	1985–1989	1990–1994	1995–1998
Women									
Not working	More than 1 year to graduation	0.20	0.21	0.22	0.22	0.19	0.17	0.14	0.12
	In final year of school	0.74	0.77	0.80	0.83	0.77	0.71	0.66	0.61
	Left school less than 3 years ago	1.22	1.28	1.35	1.42	1.39	1.36	1.33	1.31
	Left school more than 3 years ago	1.51	1.45	1.39	1.34	1.44	1.55	1.66	1.78
Working	More than 1 year to graduation	0.54	0.53	0.52	0.52	0.49	0.47	0.45	0.43
	In final year of school	1.01	0.99	0.97	0.95	0.95	0.95	0.95	0.95
	Left school; working 1st year (ref.)	1	1	1	1	1	1	1	1
	Left school; working 2nd year	1.26	1.24	1.22	1.19	1.22	1.24	1.27	1.30
	Left school; working 3rd year	1.37	1.35	1.33	1.31	1.37	1.44	1.50	1.57
	Left school; working 3+ years	1.38	1.38	1.39	1.39	1.57	1.77	1.99	2.25
Men									
Not working	More than 1 year to graduation	0.44	0.41	0.39	0.37	0.30	0.25	0.20	0.16
	In final year of school	0.99	0.86	0.74	0.64	0.62	0.60	0.58	0.56
	Left school less than 3 years ago	1.15	1.03	0.92	0.82	0.78	0.74	0.71	0.67
	Left school more than 3 years ago	0.67	0.62	0.58	0.54	0.63	0.72	0.84	0.97
Working	More than 1 year to graduation	0.67	0.67	0.66	0.66	0.68	0.71	0.73	0.76
	In final year of school	1.68	1.32	1.05	0.83	0.87	0.92	0.98	1.04
	Left school; working 1st year (ref.)	1	1	1	1	1	1	1	1
	Left school; working 2nd year	1.11	1.12	1.13	1.14	1.22	1.31	1.41	1.51
	Left school; working 3rd year	1.24	1.24	1.23	1.23	1.31	1.39	1.48	1.58
	Left school; working 3+ years	1.32	1.31	1.30	1.30	1.57	1.89	2.28	2.76

Table 6 continued

School enrolment × Working status	1955–1964	1965–1969	1970–1974	1975–1979	1980–1984	1985–1989	1990–1994	1995–1998
Time trend for reference category								
Women; left school; working 1st year	1.28	1.25	1.26	1	0.90	0.87	0.84	0.73
Men; left school; working 1st year	1.32	1.42	1.24	1	0.77	0.65	0.57	0.44

Note: The BIC of the models including the interaction between school enrolment in education × working status and time period amounts to 162,108 for women and 73,871 for men. The improvement is significant for women as well as for men

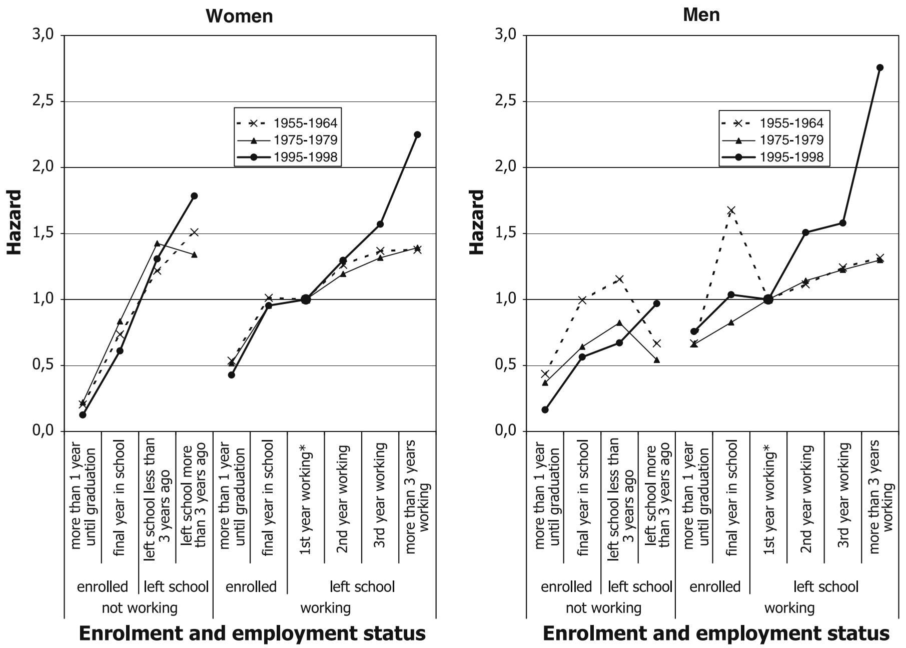

Fig. 5 Estimated effects of interaction between enrolment in education and employment status on entry into parenthood for selected years (relative risks without main period effect). The asterisk denotes the baseline level

5 Summary and Conclusion

In this article, we investigated the differences in male and female family formation over time in France. Using event history techniques and based on data from the EHF 1999, we studied how the educational level attained, enrolment in education and employment status affect the risk of entering a first union or having a first child for men and women in France. We formulated three hypotheses about the gender differences regarding the impact of educational attainment, enrolment and working status on union formation and first parenthood. Moreover, we hypothesised that the gender differences narrowed over time and that there was convergence in the impact of educational attainment, enrolment in education and employment on the union formation of men and women.

Concerning the educational level attained, we found the expected increase in union formation rates with educational level for men. For women, the U-shaped effect was transformed into an increase with educational level during the 1970s. Regarding entry into first parenthood, we found a similar decline in the variations with educational level for men and women. Our data support the convergence hypothesis regarding the effect of educational level attained, although the pattern was already rather similar for men and women in the earliest period.

In our second hypothesis, we postulated that the effect of educational enrolment is stronger for women than it is for men, and that it is smaller for both sexes in the final year of education. Moreover, we posited that the negative impact of enrolment has decreased in the recent period. Our data confirmed the gender difference regarding the negative impact as well as the weaker effect in the last year before graduation. The gender difference regarding the impact of enrolment in education on the risk of entering a first union for female and male students remained stable. However, family formation rates for non-working male students in their final year of education significantly decreased over time. Hence, the impact of educational enrolment in the final year of schooling reversed for both sexes. This may reflect a greater importance of working status for male students shortly before graduation, while it may have become less binding for women in the recent period, because more women now tend to form a union before finishing their studies, in particular if they enter a union with an older man, who has already established himself in his job (Robert-Bobée and Mazuy 2003).

The third set of hypotheses deals with gender differences regarding the effect of employment. We focused on two different effects of employment on family formation, namely the impact of employment status *per se* (i.e. working and not working) and the effect of work experience. For both effects, we posited that the impact was greater for men than for women, and found that the difference was more pronounced in the earlier periods.

Our data confirm the hypothesis of gender differences with respect to difficulties in the transition from school to work. We found that the status of not working has a very strong negative impact on family formation rates for males, which significantly increases with the time elapsed since graduation. The effect was considerably smaller for women and even positive for entry into first union in the earliest period. For non-working women the risk of having a first child is even higher than for their peers in their first year of work. During the period under study, the working status has increasingly gained importance for female entry into a union, with non-working women showing significantly lower union formation rates in the last period. This confirms the convergence hypothesis with respect to the working status, as well as the crucial role of working status compared to educational level found by many authors, such as Blossfeld and Huinink (1991) for Germany and Kravdal (1994) for Norway.

Furthermore, our study shows that the impact of work experience on first union formation and first parenthood increases with a longer duration of employment. Contrary to our hypothesis, we found almost no gender differences in the impact of work experience for all periods taken together.

Our attempt to identify time trends through interaction effects is debatable, as our covariates are strongly correlated. Taking advantage of the large sample size of the EHF survey, we used three timeclocks, age, duration since the end of studies (including the last year in educational enrolment), duration since first job, and two time-varying covariates, level of education and a combination of school enrolment and working status. Adding other interactions, e.g. between educational level, age and duration since the end of the studies was not possible because of spurious correlations. Our models are dominated by age effects, as many events occur in one of our "enrolment in education and working status" covariate categories, namely "working for 3 years or more". They are thus comparable with other studies.

Nevertheless, our models allowed us to show that the time trends in the contrasts by educational and working status are different for entry into a first union and first parenthood. When taking into account how the impact of work experience changed over time, we saw that, in the most recent period, the effect has become weaker for men and has almost disappeared for women regarding their entry into a first union. For the transition to first birth, on the contrary, the effect of work experience has become stronger for men as well as for women. This is in line with our last hypothesis that the spread of unmarried cohabitation has transformed entering into a first union and having a first child into two separate events, the former being less dependent on working status than the latter.

The increasing age at the end of studies has led to a potentially increasing proportion of young men and women entering their first union before ending their studies. This appears to have been the case for women, who more often enter their first union as a student. Nevertheless, the relative risk of union formation among students, compared to working women and men, has decreased, and working experience has become increasingly important for entry into parenthood. These changes in the effects of work experience on family formation are similar for men and women.

Our main result is that for both our family formation behaviours, entry into first union and entry into parenthood, we found an increasing similarity between the effects of level of education, enrolment in education, working status and working experience for men and women. It is consistent with the spread of the "new paradigm" of union formation described by Cherlin (2000) and others based on increasing bargaining power of women and the increasing importance of gender symmetry as a value among young men and women. This convergence between forces driving the family formation behaviour of young men and women is proven by the use of hazard models, regardless of the actual sequence of transitions, which depends on the timing of events used here as covariates, namely the end of studies and the first working experience.

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