Gender Differences in Union Formation in France: Is There Convergence Over the Recent Period?

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Draft

Abstract

Numerous studies have shown that educational attainment and labour force status have a strong impact on the timing of union formation for men and women. The effects of educational level and enrolment as well as employment appear to be different for men and women. The aim of this paper is to investigate how these gender differences of the effects of education and employment on union formation changed over time, particularly, whether these differences are vanishing in the recent years. We use a large-scale survey (more than 250,000 men and women born after 1940) conducted within the French 1999 census. The convergences hypothesis is supported for the effects of educational attainment and working status. But there is less evidence for the convergence of the effects of enrolment and work experiences for men and women. Moreover, we will test our hypotheses also for specific groups through more sophisticated models including several interactions, which could not be tested with a smaller sample size.

1 Introduction

Educational attainment and employment have shown to be the main predictors of the timing of union formation for males and females, both on theoretical reasoning (Becker 1981, Oppenheimer 1988) and on an empirical basis (Goldscheider and Waite 1986, Hoem 1986, Blossfeld and Huinik 1991, Thornton et al. 1995, Liefbroer and Corijn 1999). However, these studies reveal differences in the impact of educational level and enrolment and employment for men and women. For instance, based on the New Home Economics theory, marriage rates are expected to be increasing with increasing levels of education among men but decreasing among women (Becker 1981), since higher educational levels are associated with higher expected earning levels. While the latter increases the attractiveness on the marriage market for men, they imply greater opportunity costs of housework and child rearing and greater

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economic independence for women (Becker 1981). Furthermore, Goldscheider and Waite (1986), Blossfeld and Huinik (1991) and Coppola (2004) find that the negative effect of school enrolment is greater for women than for men, which seem to imply that school continuation is usually more difficult within marriage for women than for men (Goldscheider and Waite 1986). Moreover, difficulties in becoming employed hampers the start of union formation among men (Goldscheider and Waite 1986, Oppenheimer et al. 1997, Liefbroer and Corijn 1999), while being unemployed for a women does not necessarily delay the entry into union (Liefbroer and Corijn 1999, Blossfeld and Huinik 1991).

However, educational attainment and employment for males and females were subject to tremendous changes in the last decades. In fact, the educational level attained was rising for both sexes, but more dramatically for women, "so that [in France] at university, girls are now more numerous than boys" (Leridon and Toulemon 1995). The educational expansion was accompanied by a rise in female employment. In particular, Leridon and Toulemon (1995) report that the rate of employment for women aged 25 to 39 years, where the "competition between maternity and economic occupation is crucial", almost doubled from 41.5 percent to 74 percent between 1962 and 1989. In 2002, the labour force participation rate of women aged 25 to 34 years even amounts to about 79 percent (ILO, 2003).

The pattern of union formation was also subject to dramatic changes since the 1960s in most of the Western countries. In particular, the mean age at first marriage rose rapidly from about 22 years in the early 1970s to about 28 years in the year 2000 while the period measure of the proportion of women ever-marrying fell from 0.90 in 1970 to 0.65 in 1995 in France (Goldstein 2002). Moreover, since the late-1960s, the number of consensual unions has risen dramatically, so that in 1995, about 90 percent of first unions began outside marriage (Toulemon 1997). The spread of cohabitation may have changed gender differences in union formation as well, since Leridon and Toulemon (1995) find that in France, cohabitation may be more strongly associated with new—less sex-differentiated—roles, whilst married couples adhere to a more traditional pattern."

The focus of this paper is to investigate the differences between men and women in union formation in France. In particular, we aim to analyse the differences between males and females of the impact of educational attainment and employment and how these gender differences of the effects of education and employment changed over time, particularly, whether these differences are vanishing in the recent periods.

The remainder of the paper is structured as follows. In section 2, we discuss in detail the theories of union formation in relation to education and employment and their empirical relevance. Section 3 introduces the data and methods and section 4 presents the results. Finally, section 5 concludes and outlines further analysis.

2 Theoretical framework and hypotheses

Several theories have been developed about why people marry and what factors influence the timing of marriage. In the New Home Economics theory, marriage is seen as a rational choice made by individuals for whom the gain from marriage outweighs the benefits from remaining single. Given the complementarity of men and women in the household production of commodities, these individuals would be more productive in a joint household than they would be if the remained single. If each sex specializes in its comparative advantage[1], then the sexual division of labour within households creates gains from marriage (Becker 1976, 1981).

[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 in the market is higher for him/her than for the other household member(s) when all supply the same amount of time to the household and when all invest in the same human capital (Becker 1981).

Thus, the gains from marriage stem from the assumption of complementarity between men and women but also from their market opportunities.

Higher wages imply, on the one hand, higher total income of the joint household, which increases the gain from marriage (*income effect*). But 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, sex-specific division of labour within the family becomes less advantageous for the individual specializing in household tasks, which reduces the gains from marriage. If traditional division of labour prevails, the income effect is expected to dominate among men, whereas the price effect is supposed 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 that a positive impact of education among males and a negative one among females. Indeed, Huinik (1995) finds that males' marriage chances increase with earning levels. Goldscheider and Waite (1986) report a positive effect of men's education on marriage rates, albeit only before age 25. However, Brüderl and Diekmann (1994) report that the effect of men's educational attainment changed from positive to negative across cohorts.

Empirical evidence of the impact of education on union formation for females is mixed. Some studies have estimated positive effects of female education on marriage (Goldscheider and Waite 1986), others have found insignificant effects (Blossfeld and Huinik 1991, Hoem 1986) and yet others have found negative effects (Preston and Richards 1975). However, the specialisation model of Becker has been heavily criticised of several reasons. In particular, Oppenheimer (2000) argues that "sex role specialisation is essentially a high risk and inflexible family strategy in an independent nuclear family system." Thus, highly educated women with greater labour market potential are more attractive to their future spouses than less educated women with poor employment prospects.

Moreover, Liefbroer and Corijn (1999) find that the impact of educational attainment and labour force participation depends on the degree of incompatibility of employment and family life. By using Dutch and Flemish survey data, they empirically verify their 'societal contingency hypothesis' that the more gender "equality is a dominant cultural value and the better structural opportunities to combine work and family life, the weaker the impact of educational attainment and labour force participation." Moreover, Liefbroer and Corijn (1999, p. 50) argue that in "the past, the lack of efficient contraceptives resulted in a very close connection between [marriage and parenthood]. Either initialisation of sexual intercourse at marriage led to pregnancy soon after marriage, or the initialisation of sexual intercourse before marriage led to a "forced" marriage soon after the women became pregnant. As a result, women who preferred to postpone childbearing were well-advised to postpone marriage as well." However, in many countries, the incompatibility of female labour force participation and family life has weakened in the last half of the century (Gauthier 1996), and with the provision of efficient contraceptives, the link between first union and the occurrence of a first child is broken[2], and the negative impact of educational attainment on first union may have been weakened.

In addition, Thornton, Axinn, and Teachman (1995) suggest that the opportunity costs of cohabitation are lower than those of marriage. Similarly, Oppenheimer (1988) argues that the low opportunity costs of unmarried cohabitation would make this living arrangement especially attractive to better-educated women. From the perspective of bargaining theory, Cherlin (2000) finds that women are incorporating premarital cohabitation into their search processes because cohabitation provides a better opportunity to observe men's earnings

[2] Robert-Bobée and Mazuy (2003) reveal that in France the average duration between union formation and arrival of first child is increasing over birth cohorts from about two years for women born before 1940 and men born before 1935 to about 3.5 years among men and women born in the early 1960s.

potential and willingness to share household and child-raising tasks. In fact, Leridon and Toulemon (1995) even find a positive effect of education on cohabitation intensities but negative on marriage intensities.

These considerations above lead to the formulation of our first hypothesis.

Hypothesis 1: The effect of educational level attained is positive for men, but negative or U-shaped for women. But with the spread of cohabitation, and the introduction of institutions, which ease the combination of family life and employment, the negative link of union formation and educational level has been weakened over time. Moreover, in the recent periods the union formation rates for low-educated women are declining implying a less decreasing or even increasing impact of the educational level on union formation rates.

However, higher educational attainment implies a prolonged period of educational enrolment, which itself has a substantial impact on the timing of union formation. In particular, when attending school or university, students are normally not economically independent, and they have to rely to a large extent on parental financial support (Blossfeld and Huinik 1991, Blossfeld and Jaenichen 1992, Coppola 2004). Moreover, Blossfeld and Huinik (1991, p. 158) state that "from a sociological point of view, there exists normative expectations in the society that young people who attend school are not `at risk´ of entering marriage." The incompatibility of student roles with adult family activities thus delays family formation until the educational career has been completed.

However, as cohabitation generally involves a lower commitment than marriage, the opportunity costs of cohabitation may be lower than those of marriage (Oppenheimer 1988). Therefore, the roles of cohabitation and being a student may be less conflicting and "students may be more willing to enter cohabitation than marriage" (Thornton et al. 1995).

There exists overwhelming empirical evidence that being in education has a strong negative impact on entry into first union (e.g. Hoem 1986, Blossfeld and Huinik 1991, Blossfeld and Jaenichen 1992, Leridon and Toulemon 1995, Coppola 2004). However, several studies reveal gender differences in the impact of educational enrolment on union formation. In particular, they find that the effect of school enrolment is bigger for women than for men (Goldscheider and Waite 1986, Blossfeld and Huinik 1991, Coppola 2004), which seem to imply that school continuation is usually more difficult within marriage for women than for men (Goldscheider and Waite 1986). In contrast, Robert-Bobée and Mazuy (2003) find, by examining recent French data, that women form more often a union while being in education than men, as they are entering their first union at a lower age and complete their studies later than men.

Motivated by the considerations above, we aim to test the following hypothesis:

Hypothesis 2: The effect of educational enrolment is bigger for women than for men, but this effect is decreasing in the recent periods. Since the completion of education normally can be foreseen in the short run, the negative impact of educational enrolment on union formation is significantly lower shortly before leaving school or university.

Finishing education is one of the markers in the transition to adulthood; entry into employment is another. An individual's current labour market position affects his or her ability to form a union because it is a mean to achieve the necessary economic independence to set up an independent household (Oppenheimer 1988). According to the New Home Economics, the latter will speed up union formation for men, but enables women to avoid union formation (Becker 1976, 1981). But women might perceive the economic independence also as a mean to share the costs of setting up a common household and, thus, a way to accelerate the entry into a union (Oppenheimer 1994). Nevertheless, Oppenheimer and Lewin (1999, p. 13) find that it is still unlikely that "women's familial roles are normatively defined in terms of their ability to make a major and long-term stable income contribution to the family to the same degree as men's." Hence, men's career and career maturity are playing a more important role than women's for the timing of marriage of both men and women.

Indeed, several studies empirically verify that unemployment hampers the start of family formation for men (Goldscheider and Waite 1986, Oppenheimer et al. 1997, Liefbroer and Corijn 1999). In contrast, there is empirical evidence that being unemployed for a woman does not necessarily delay the entry into union (Blossfeld and Huinik 1991, Liefbroer and Corijn 1999), but in some cases women's economic independence has been found to accelerate the entry into union (McLaughlin et al. 1993, Oppenheimer 1994).

However, it may not only be the employment status but also the work experience, which affects the timing of union formation. The more work experience collected, the more stable the employment might be. Indeed, Oppenheimer and Lewin (1999) find that young men's career maturity has a strong positive impact on their marriage formation. Furthermore, Kravdal (1999) obtains that Norwegian single women with less than one year of work experience have significantly lower union formation rate than those with more than two years of experience. However, Kravdal (1999) finds that marriage requires a stronger economic underpinning than informal cohabitation. As mentioned earlier, Cherlin (2000) claims that women are incorporating premarital cohabitation into their search processes because cohabitation provides a better opportunity to observe men's earnings potential. Therefore, the effect of work experience on union formation may have decreased in the recent periods due to the spread of cohabitation.

The considerations above lead to our third set of hypotheses.

Hypothesis 3a: The effect of working status matters more for men than for women, but the difference is more pronounced in earlier periods than recently, while the effect may have remained for fertility behaviours.

Hypothesis 3b: The time elapsed since first job matters more for men than for women, but the difference is more pronounced in earlier periods than recently. Moreover, the effect may be weaker in the recent period for both sexes.

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, Héran, Toulemon 2000). In this study, 235 000 women and 145 000 completed an additional questionnaire on their origin, children, partnerships, working life, social origin and languages spoken in the family. We restricted our sample to birth cohorts after 1940. Immigrants were only included if they arrived in metropolitan France before they reached age 15, i.e. they underwent their transition to adulthood in France. Moreover, we excluded observations, where the event took place before the age of 15. Finally, about 146 000 women and about 96 000 men remained in our sample, where about 85 percent of women and about 78 percent of men experienced a first union.

We follow individuals from their 15th birthday until the time of the event, i.e. union formation. There are no competing events. Furthermore, we right-censor by the 1st of January 1999 or by their 40th birthday. We model the intensity to form a union by using a piecewise constant exponential model (Blossfeld and Rohwer 2002). Let t denote the time since the 15th birthday, then the hazard rate of entry into a first union, $h(t)$ is given by

$$h(t) = \exp\left\{\alpha_l + \sum_{k=1}^{p} \beta_k Z_k(t)\right\} \quad \text{if } t \in I_l \qquad \textbf{(1)},$$

where

$$I_l = \{t \mid \tau_l \le t < \tau_{l+1}\}, \quad l = 1, \ldots, L$$

indicates the L time intervals of a constant age effects α_l. We assume that the effect of age is constant over single years of age in order to achieve maximum flexibility. Hence, $\tau_1, \tau_2, \tau_{3,\ldots,} \tau_{26}$ represent the 15th, 16th, 17th, …, 40th birthday. Moreover, β_k denote the coefficient of the (time-dependent) covariates $Z_k(t)$, which correspond to social origin, educational level, enrolment and employment status, pregnancy, and calendar period.

Explanatory variables

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

Concerning the individual characteristics, first the effect of age is controlled for. The process of union formation is supposed to be highly dependent on age, where union formation rates usually display a bell-shaped pattern with increasing age (Blossfeld and Huinik 1991, Coppola 2004). By assuming a piecewise constant exponential model as outlined above, we incorporate age in single-year steps.

Furthermore, we control for the presence of pregnancy or a child (time-varying covariate). The conception of a birth out of a union may accelerate first union formation because of a desire to offer the child the social and economic protection in 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 1997, Brien et al. 1999, Baizán et al. 2004). Since, there is a time lag between conception and detection of pregnancy of about one month, we follow pregnancies from one month after the conception.

However after the birth, the presence of children may impose constraints on resources and time, which may hamper union formation (Baizán et al. 2004). In fact, Brien et al. (1999) find that the hazard of entering a marriage or cohabitation falls approximately to or even below the level before conception for US women. Baizán et al. (2004) obtain similar findings for German and Swedish women.

Apart from individual characteristics we also control for characteristics from the family of origin. These are the parents' socio-economic position and the number of siblings (both time-constant variables). Indeed, several studies show that the transition of first union is strongly influenced by some characteristics of the family of origin. In particular, the parents' socio-economic position affect the timing of union formation not only by income positions, properties, consumption styles, and economic strategies of families, which create social opportunities for children, but also their social orientations, values, and beliefs, which influence family, educational and career decisions (Blossfeld and Huinik 1991). Therefore, controlling for socio-professional status of the father is common in analysis of family formation (Goldscheider and Waite 1986, Blossfeld and Huinik 1991, Leridon and Toulemon 1995, Thornton et al. 1995). In particular, we distinguish between farmers, self-employed, unskilled worker, skilled worker, low-level white-collar, medium level white-collar, and high-level white-collar.

We tested also to incorporate the socio-professional status of the mother. However when controlling for the other characteristics, the effects of the socio-professional status of the mother turned out to be minor and mainly insignificant and were, therefore, not longer included due to parsimony.[3]

Moreover, we incorporate the number of siblings, because there exists empirical evidence that individuals in large households enter a first union earlier (Billari and Philipov 2004). We consider zero, one, two, three, four or five and more siblings.

Our main explanatory variables are the educational level, the enrolment and the working status (all time-varying covariates). Concerning the educational level, we distinguish between primary or no degree, BEPC (Brevet d'Études du Premier Cycle) including brevet élémentaire and brevet des collèges, BEP (Brevet d'Études Professionels), CAP (Certificat d'Aptitude Professionel), Baccalauréat, and university. In France, schooling is mandatory between ages 6 and 16 since 1967, and was compulsory from 6 to 14 before 1967. Primary schooling lasts from 6 to 11, and lower secondary education is between ages 11 and 15. At the end of lower secondary education, pupils are eligible to take the brevet exam (BEPC and its predecessor brevet élémentaire and brevet des collèges). Before the extension of compulsory schooling, the vast majority of children were attending an extended primary schooling containing eight grades, which lead to the Certificat d'Études Primaires (Grenet 2004). In the upper secondary, children either enter a lycée, which ends with the Baccalauréat, or start a vocational training leading either to BEP or CAP (Eurydice 2005).

We do not have the complete educational history, but only the highest degree attained and the age at the end of studies. Therefore, we assumed that BEPC is reached by the age of 15 and the Baccalauréat by the age of 18 for constructing the time-varying educational level.

Concerning the enrolment status, we extend the commonly used dichotomous variable of being enrolled or not by taking into account the years since left school, since this has been shown to have a significant effect at least on entry into motherhood (Buber 2002). Since, the end of studies is a foreseeable event, at least in the final year of studies, we distinguish between more than 1 year to finish school and the last year of school. Indeed from a theoretical perspective, Thornton et al. (1995) find that "[f]or young people who decide to combine marriage and student roles, marrying near the end of one's schooling minimizes the time spent in conflicting roles." Therefore, we distinguish between enrolled and more than one year to finishing education, final year in education, first year after leaving school, second year after leaving school, third year after leaving school, more than three years after leaving school.

The third main explanatory variable is employment status, i.e. working versus not working. In order to incorporate the pace and difficulties to a stable work career, we incorporate the time since start of the first job in the employment status. Hence, we distinguish between not working, first year working, second year working, third year working and three and more years working.

Finally, we include the calendar period effects into our model (time-varying covariate), where we split the calendar time into five-year groups. Since we consider only birth cohorts born after 1940, there are only few events in the years 1955 to 1960. Therefore we combine the last half of the 1950s with the first half of the 1960s. Hence we distinguish between the calendar periods 1955-1964, 1960-1964, 1970-1974, 1975-1979, 1980-1984, 1985-1989, 1990-1994, and 1995-1998.

In order to observe the change of the hypothesized effects over time, we model interactions with the period variable and the other explanatory variables. We model the latter by employing a spline with a node at 1975-1979, in order to lower the number of coefficients to be estimated. By taking the calendar period 1975-1979 as a reference category, we estimate the average change of the effect of the covariates from 1955-1964 to 1975-1979 and the average change from 1975-1979 to 1995-1998.

[3] When leaving out the socio-professional status of the mother, the Bayesian Information Criteria improved.

The size of the dataset even allows us to test whether the effects of our main explanatory variables changed particularly for specific groups, i.e. for children of high-level white-collar workers, individuals with a university degree, etc. We model, therefore, first the interactions of the educational and employment variables with the covariates, which define the group membership, and, then, explore the change over time of these interactions.

The models are estimated via maximum likelihood, using the statistical software package STATA (StataCorp 2004). Model selection is based on the Bayesian Information Criteria (BIC), since it also allows comparing non-nested models.

4 Results

Table 1 gives the parameter estimates for men and women, where we control for age, number of siblings, socio-professional status of father, pregnancy, level of education, enrolment status, working status, and calendar period (model A). The numbers given in Table 1 are the estimated values for the coefficients β_k in Equation (1). In order to derive the relative risk of the specific category, one has to evaluate the exponential function at the estimated coefficient. In addition, Figure 1 displays the estimates for men and women of model A, which are given in Table 1.

The majority of the covariates have the expected effects. The union formation rates exhibit a bell-shaped pattern with increasing age, both for men and women, where the female union intensities reach their highest values about two years earlier than men's (age 21 versus age 23). Moreover, young women display higher union intensities than young men, while starting from the mid-twenties, men and women show similar values for the hazard of entry into first union. Previous studies show that men enter a union significantly later than women (e.g. Coppola 2004). By estimating the models for males and females separately, we reveal that the gender difference is particularly pronounced before age 25. For ages above 25 years, the difference between men and women is only minor.

Growing up with at least on sibling has also a positive impact on union formation. This is true for men and women, where being the single 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 minor and hardly significant effects.

The profession of the father almost does not have any significant impact on union formation among men and women, except for men if the father is a farmer. Being the son of a farmer reduces union formation intensities by 12 per cent! This may be due to the fact that sons of farmers have a higher probability to become farmers themselves compared to children with another social background and that farmers have lower marriage chances than other professionals. (Courgeau and Lelièvre 1986). Furthermore, if the father is a medium or high-level white-collar worker, then men show slightly higher union formation rates (about 4 per cent), 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 union formation (about 3 per cent).

The presence of a pregnancy strongly accelerates union formation. In this case, men and women show 12-fold and 8-fold higher union formation rates than before the conception. After the birth of the child, union formation intensities steeply decrease. In fact, men and women only have higher union formation rates than before the conception of about 76 and 17 per cent, respectively. 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, neglected to report about this child. In this case, the estimates are biased upwards.

The educational degree attained has a significant, positive effect on union formation intensities for men and women. But for low education (primary or no degree), there is a significant, negative impact on union formation for men, while there is no significant effect for women. In particular, low educated men have about 16 per cent lower union formation rates than those with a BEPC. In contrast, men and women with a CAP degree show about 4 per cent higher union formation intensities than with a BEPC degree. Furthermore, passing a BEP degree increases the union formation rates for women by about 9 per cent, while there is no significant effect for men. Finishing the baccalaurèat raises the hazard of entry into a first union by about 13 per cent, while a university degree even increases union formation rates by about 54 per cent, both for men and women. Summing up, we find, differently to our hypothesis, an increasing impact of the educational level on union formation rates for men and women, and therefore, no gender differences except for the lowest educational level.

Being in education has a significant negative impact on union formation for both men and women, where the effect is lower in the final year in education. According to Thornton et al. (1995) students entering a union in the final year spend less time in conflicting roles. Therefore, the graduation in the near future diminishes the negative impact of enrolment on union formation rates. However, it may also be the case that the conflict between union and student roles causes a soon drop out of education, which biases the estimate of the effect of enrolment upwards shortly before leaving school.

In line with our hypothesis, we find that the effect of enrolment is bigger for women than for men. In particular, enrolment decreases union formation rates by 71 and 43 per cent for women and men, respectively, when there is still more than one year until graduation. In the final year of education, union formation rates are still lower by about 27 and 11 per cent for female and male students, respectively.

Surprisingly, the effect of the time elapsed since leaving school is negative for women and positive for men. Moreover, the longer the period since graduation, the stronger is the negative impact for the women! In contrast, the time since the start of first job has a significant positive impact on union formation. In particular in the second year of employment, women exhibit higher union formation rates of about nine per cent compared to the first year of working. In the third year of employment, the female union formation intensities rise to a difference of about 21 per cent and further increase to about 31 per cent for three and years of work experience. Men show similar values, although, the rise in the union formation rates with increasing working experience is less steep. In contrast, the status of not working decreases the hazard of entry into a first union much stronger for men than for women (36 versus 16 per cent).

Hence, we demonstrate that the work experience rather than the time elapsed since graduation speeds up union formation, by additionally controlling for the time elapsed since the start of first job. In order to study the relation between the effects of work experience and enrolment status (including the time since graduation) more in depth, we explicitly model the interaction between the two explanatory variables in the next step.

Finally, we find a hump-shaped pattern of calendar period effects on union formation for men and women, where the estimated effects are rather stable since the 1985.

As mentioned above, we explicitly model the interaction between enrolment and employment status and their effect on union formation. However, the incorporation of all the interactions between the different categories of enrolment and employment status would imply the estimation of further 20 coefficients. Therefore, we combined some interaction categories in a meaningful manner, which is also confirmed by an improvement in the BIC relative to model A as well as relative to the model including all the interactions (result not shown). In particular, we distinguish for non-working individuals according to the enrolment status between enrolled and more than one year to graduation, final year in school, left school less than three years ago, and left school 3 and more years ago. For the working, we differentiate between enrolled and more than one year until graduation, final year in school, left school and working 1st year, left school and working 2nd year, left school and working 3rd

year, and left school and working three and more years. The estimated parameter values of the resulting model, which is called model B, are given in Table 1.

In particular, we find that being enrolled significantly decreases the union formation rates for working and non-working individuals. In particular if there is still more than one year until graduation, the hazard of entry into a first union decreases by about 77 per cent for non-working women and 65 per cent for non-working men compared to those who just finished school and entered employment. In contrast, working students have by about 48 and 35 per cent lower union intensities for women and men, respectively, in this case. Similar to model A, we find that the negative impact of enrolment on union formation rates is lower in the final year of education, i.e. non-working students exhibit union formation intensities, which are only by about 35 and 41 per cent lower for women and men, respectively, while working students are entering by about 17 and 23 per cent less frequent a first union for women and men, respectively, in the final year of school than those who just finished school and started their first job. By distinguishing between working status for the enrolled, we show that those who already combine student and working roles are entering more often a union than non-working students but less frequent than not enrolled, working individuals. However due to data limitations we do not distinguish neither 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 them.

Concerning the difficulties in the transition to employment, we reveal that women who did not yet start working, although they left the school less than three years ago, only show slightly lower union formation rates than those who just began to work (-6 per cent). However, if the period between leaving school and entering employment extends over three years, then the negative effect increases to 21 per cent. In contrast, the status of "left school but not yet working" has a much stronger negative impact on union formation intensities for men. Indeed, for non-working males who left the school less than three years ago, show by about 30 per cent reduced union formation rates, and for those who left education more than three years ago, even have about 61 per cent lower union formation intensities. The latter effect is of similar magnitude than the effect of non-working students!

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

Concerning the impact of work experience of non-students on entry into a first union, we find the more work experience, the higher the union formation rates. In particular, being employed for the second year already increases union formation hazard by about seven and five per cent for women and men, respectively. The third year of work experience raises the union formation rates already by about 16 and 12 per cent for females and males, respectively, while working for three and more years, the union formation intensities increase by about 22 and 26 per cent for women and men, respectively, compared to those in the first year of employment. In contrast to our hypothesis, there is almost no difference in work experience between male and female non-students.

In the next step, we intend to investigate the trend in the gender differences of the effects of level of education, and enrolment and employment status. Therefore, we additionally take the interaction of these explanatory variables with the calendar period into account. As mentioned earlier, we model the calendar period for the interaction by 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 of the corresponding period relative to 1975-1979.

Figure 2 shows the estimated effect of the level of education attained for the periods 1955-1964, 1975-1979, and 1995-1998. As evident from Figure 2, the effect of low and medium levels of education clearly changed for women over time, while the relative effect of a university degree for women is pretty stable. In particular, a primary or no degree even had a positive effect on female union formation compared to a BEPC in the earliest period. Hence, the effect of educational level was slightly U-shaped during 1955-1964. But between 1955-1964 and 1975-1979, the effect of primary or no degree turned negative relative to a BEPC, and since the latter period an increasing effect of educational on the entry into a first women can also be observed for women.

The effect of educational level on male union formation exhibited an increasing shape for all the periods, except BEP, although, the difference between primary or no degree and BEPC is minor for the earliest period. However, similar to the female union formation rates, the impact of the lowest level of education on male union formation strongly decreased between 1955-1964 and 1975-1979, and since then, the effect of union formation is clearly increasing for males. Summing up, we find evidence that the gender differences with respect to educational level diminished over time, although, the pattern was rather similar already in the earliest period.

Concerning the interaction of enrolment and working status, we found more pronounced gender differences in the analysis above. Since enrolment and labour force participation were subject to tremendous changes in the last decades, we expect also more pronounced changes in the effect of enrolment and working status on union formation rates. Figure 3 displays the effects of the interaction of enrolment and employment status for the periods 1955-1964, 1975-1979, and 1995-1998. Indeed, female union formation rates changed a lot since 1955, particularly for non-working women who finished their education, while for non-working female students the pattern was rather stable. During 1955-1964, women who finished the school but did not yet enter a job even formed more frequently a union than those who started their work after graduation. Moreover, the longer the duration since graduation, the higher are the female union formation rates during 1955-1964. This result is in line with earlier studies (Blossfeld and Huinink 1991, Liefbroer and Corijn 1999), which show that unemployment does not necessarily hamper union formation for women.

However from 1955-1962 to 1975-1979, the female union formation intensities of non-working non-students decreased relative to those who just started working and having left school. While non-working women who left school less than three years ago show only slightly higher union formation rates, those who left school more than three years ago already display significantly lower values than working women after finishing school. The trend of increasing importance of working status continued also during 1975-1979 to 1995-1998, where non-working women show significantly lower values of union hazards in the last period. In contrast, male union formation intensities of non-working non-students were in all the periods less than of the working non-students, where the working status increasingly gained importance, particularly, if the graduation was more than three years ago. Hence, the convergence hypothesis is confirmed with respect to the working status for people not involved in education.

As mentioned earlier the union formation rates of non-working female students hardly changed compared to working school graduates from 1955 to 1998. In contrast, the male rates for non-working students in the final year significantly decreased over time, showing already lower relative values of union formation than women in 1975-1979 and 1995-1998. This may reflect the growing importance of working status for men. Moreover, Robert-Bobée and Mazuy (2004) argue that this norm is particularly binding for men, since women now more often form a union before finishing their studies, and, most notably, if they enter a union with an elder man, who has reached already a stable working position. Hence, we even find evidence for increasing divergence of union formation for non-working female and male students.

Figure 3 also illustrates the change in the effect of work experience for non-students over time. In 1955-1964, work experience had a positive effect on union formation rates for both sexes. While an additional year of work experience increased the union formation hazard between 25 to 32 per cent for women, the increase for the first two additional years of work experience is less steep for men (around 12 per cent per year). However, if men collected already more than three years of work experience, their union formation rates are by about 90 per cent higher than at the start of their first job. The less steep increase in the initial years of employment may reflect the pace and difficulties to set up a stable working position, and their stronger importance of the latter for men compared to women as outlined by Oppenheimer and Lewin (1999).

However, the impact of increasing work experience diminished over time, as evident from Figure 3. While for men, there is still a positive effect of increasing work experience in the recent period, there is no effect visible any longer for women. Hence, there are still persisting differences of the effect of work experience in the entry to a first union.

5 Summary

In this paper, we have investigated the effects of educational level attained, and enrolment and employment status on the hazard of entry into a first union for men and women in France using data from the EHF 1999. Based on theoretical arguments and existing empirical studies we formulated three hypotheses about the gender differences of the impact of educational attainment, enrolment and working status on union formation. Moreover, this gender-specific impact was assumed to have changed over time. In particular, we hypothesise that the gender differences have narrowed over time and there is convergence in the impact of educational attainment, enrolment and employment on union formation for males and females.

Concerning the educational level attained, we assumed that the effect is positive for men, but negative or U-shaped for women. However, we find an increasing impact of the educational level on union formation rates for men and women. Only for the lowest educational level, we find a negative impact on the entry into a first union for men, but no significant effect for women. Hence, there is no evidence of gender differences from the data, except for the lowest educational level. But when investigating the changes of the effects over time, we reveal that the effect of the lowest educational level similarly turned negative between 1955-1964 and 1975-1979. Therefore, the convergence hypothesis of the effect of educational level attained was supported by the data, although, the pattern was already rather similar for men and women in the earliest period.

In our second hypothesis, we assumed that the effect of educational enrolment is bigger for women than for men, and lower in the final year of education for both sexes. Moreover, we assume that the negative impact of enrolment is decreasing in the recent period. Indeed, the gender difference of the negative impact is confirmed by the data, as well as the smaller effect in the last year before graduation. However, the male union formation rates for non-working students in the final year of education significantly decreased over time, showing even lower values of union formation than women in 1975-1979 and 1995-1998. Hence, the gender difference of the impact of educational enrolment in the final year of school even reversed. This may reflect a greater importance of working status for male students shortly before graduation, while for women this norm may be less binding in the recent period, since women now more often form a union before finishing their studies, in particular, if they enter a union with an elder man, who has reached already a stable working position (Robert-Bobée and Mazuy 2004).

The third set of hypotheses deals with the gender differences in the effect of employment. We particularly consider two different effects of employment on union formation: first, the impact of employment status *per se*, which means the difference between working and not working and, secondly, the effect of work experience. For both effects, we assume that the impact is greater for men than for women, but the difference is more pronounced in the earlier periods than recently.

The hypothesis of gender differences with respect to the difficulties in the transition of employment after having completed education are confirmed by the data. Indeed, we find a very strong negative impact of the status of not working on union formation rates for males, which significantly increases the longer the duration since graduation, while the effect is considerably smaller for women and even positive in the earliest period. Our results are in line with the arguments of Oppenheimer and colleagues that men's working position are playing a more important role than women's for the timing of a marriage (Oppenheimer and Lewin 1999). Moreover, our findings agree with the results of Blossfeld and Huinink (1991) and Liefbroer and Corijn (1999) that being unemployed does not necessarily delay the entry into a union for women. However, since 1955-1964, the working status also increasingly gained importance for the entry into a union for women, where non-working women show significantly lower union formation rates in the last period. Hence, the convergence hypothesis is confirmed with respect to the working status.

Furthermore, our study reveals that the impact of work experience on first union formation is increasing for a longer duration of employment. For all the periods together, we find, contrary to our hypothesis, almost no differences in work experience between males and females. However, by additionally taking the change of the impact of work experience over time into account, we obtain that the effect became weaker for men but almost disappeared for women in the recent period. Hence, there is no evidence for the convergence hypothesis with respect to working status.

In a further step, we aim to test whether the effects of educational attainment, enrolment and employment status changed particularly for specific groups, i.e. for children of high-level white-collar workers, individuals with a university degree, etc. Therefore, we will model first the interactions of the educational and employment variables with the covariates, which define the group membership, and, then, explore the change over time of these interactions.

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Tables and Figures

Table 1: Model estimates for log hazards; selected models.

Covariate	Model A		Model B	
	Females	Males	Females	Males
Age				
15	**-6.78**	**-8.74**	**-6.77**	**-8.72**
16	**-5.85**	**-8.09**	**-5.86**	**-8.07**
17	**-5.02**	**-6.96**	**-5.05**	**-6.94**
18	**-4.47**	**-6.09**	**-4.51**	**-6.05**
19	**-4.21**	**-5.54**	**-4.24**	**-5.50**
20	**-4.07**	**-5.00**	**-4.11**	**-4.95**
21	**-4.03**	**-4.59**	**-4.08**	**-4.54**
22	**-4.08**	**-4.44**	**-4.13**	**-4.39**
23	**-4.18**	**-4.37**	**-4.23**	**-4.32**
24	**-4.31**	**-4.39**	**-4.36**	**-4.34**
25	**-4.46**	**-4.48**	**-4.51**	**-4.42**
26	**-4.64**	**-4.60**	**-4.70**	**-4.54**
27	**-4.75**	**-4.69**	**-4.80**	**-4.63**
28	**-4.85**	**-4.84**	**-4.90**	**-4.78**
29	**-4.96**	**-4.95**	**-5.02**	**-4.89**
30	**-5.21**	**-5.00**	**-5.26**	**-4.93**
31	**-5.29**	**-5.21**	**-5.34**	**-5.14**
32	**-5.35**	**-5.36**	**-5.40**	**-5.30**
33	**-5.44**	**-5.47**	**-5.49**	**-5.41**
34	**-5.76**	**-5.55**	**-5.81**	**-5.48**
35	**-5.72**	**-5.58**	**-5.77**	**-5.52**
36	**-5.80**	**-5.81**	**-5.86**	**-5.74**
37	**-5.91**	**-5.86**	**-5.96**	**-5.80**
38	**-6.04**	**-5.85**	**-6.09**	**-5.79**
39	**-6,17**	**-5,97**	**-6,22**	**-5,91**
Siblings				
None	**-0.15**	**-0.21**	**-0.15**	**-0.21**
1	*0.00*	*0.00*	*0.00*	*0.00*
2	**0.03**	0.02	0.03	0.02
3	0.02	0.02	0.02	0.02
4	0.01	**0.04**	0.00	**0.04**
5 and more	0.01	0.01	0.00	0.01
Socio-professional status of father				
Farmer	-0.02	**-0.13**	-0.02	**-0.13**
Self-employed	0.00	0.03	0.00	0.03
Unskilled worker	0.02	-0.02	0.02	-0.02
Skilled worker	*0.00*	*0.00*	*0.00*	*0.00*
Low-level white-collar worker	**-0.03**	0.02	**-0.03**	0.02
Medium-level white-collar worker	0.00	**0.04**	0.00	**0.04**
High-level white-collar worker	-0.03	**0.04**	-0.02	**0.04**
Pregnancy				
Not pregnant/no child	*0.00*	*0.00*	*0.00*	*0.00*
Pregnant	**2.09**	**2.47**	**2.09**	**2.47**
Child	**0.16**	**0.57**	**0.15**	**0.56**
Level of education attained				
Primary/no diplom	-0.01	**-0.17**	-0.02	**-0.17**
BEPC	*0.00*	*0.00*	*0.00*	*0.00*
CAP	**0.04**	**0.04**	**0.05**	**0.04**
BEP	**0.09**	0.03	**0.09**	0.02
Baccalaurèat	**0.12**	**0.12**	**0.13**	**0.11**
University	**0.43**	**0.43**	**0.44**	**0.40**

Remark: Boldface numbers indicate that the specific value is significantly different at the five per cent level from the baseline.

Table 1 (continued): Model estimates for log hazards; selected models.

Covariate	Model A		Model B	
	Females	Males	Females	Males
Enrolment status				
More than 1 year to graduation	**-1.23**	**-0.56**		
Final year of school	**-0.31**	**-0.12**		
Left school less than 1 year ago	*0.00*	*0.00*		
Left school between 1 and 2 years ago	**-0.03**	**0.08**		
Left school between 2 and 3 years ago	**-0.05**	**0.08**		
Left school more than 3 years ago	**-0.13**	**0.08**		
Employment status				
Not working	**-0.17**	**-0.45**		
1st year working	*0.00*	*0.00*		
2nd year working	**0.09**	0.01		
3rd year working	**0.19**	**0.07**		
3 and more years working	**0.27**	**0.20**		
Calendar period				
1955-1964	**-0.23**	**-0.25**	**-0.24**	**-0.24**
1965-1969	**-0.13**	**0.02**	**-0.13**	**0.02**
1970-1974	**-0.04**	**0.04**	**-0.04**	**0.04**
1975-1979	*0.00*	*0.00*	*0.00*	*0.00*
1980-1984	**-0.06**	-0.06	-0.06	-0.06
1985-1989	**-0.14**	**-0.21**	**-0.14**	**-0.21**
1990-1994	**-0.09**	**-0.22**	**-0.09**	**-0.22**
1995-1998	**-0.12**	**-0.25**	**-0.12**	**-0.25**
Enrolment and employment status interaction				
Not working				
More than 1 year to graduation			**-1.45**	**-1.06**
Final year of school			**-0.43**	**-0.53**
Left school less than 3 years ago			**-0.06**	**-0.35**
Left school more than 3 years ago			**-0.23**	**-0.95**
Working				
More than 1 year to graduation			**-0.66**	**-0.43**
Final year of school			**-0.19**	**-0.26**
Left school; working 1st year			*0.00*	*0.00*
Left school; working 2nd year			**0.07**	0.05
Left school; working 3rd year			**0.15**	**0.11**
Left school; working 3 and more years			**0.20**	**0.23**
Log-likelihood	-107684	-56323	-107464	-56250
Bayesian Information Criteria	216264	113528	215829	113382

Remark: Boldface numbers indicate that the specific value is significantly different at the five per cent level from the baseline.

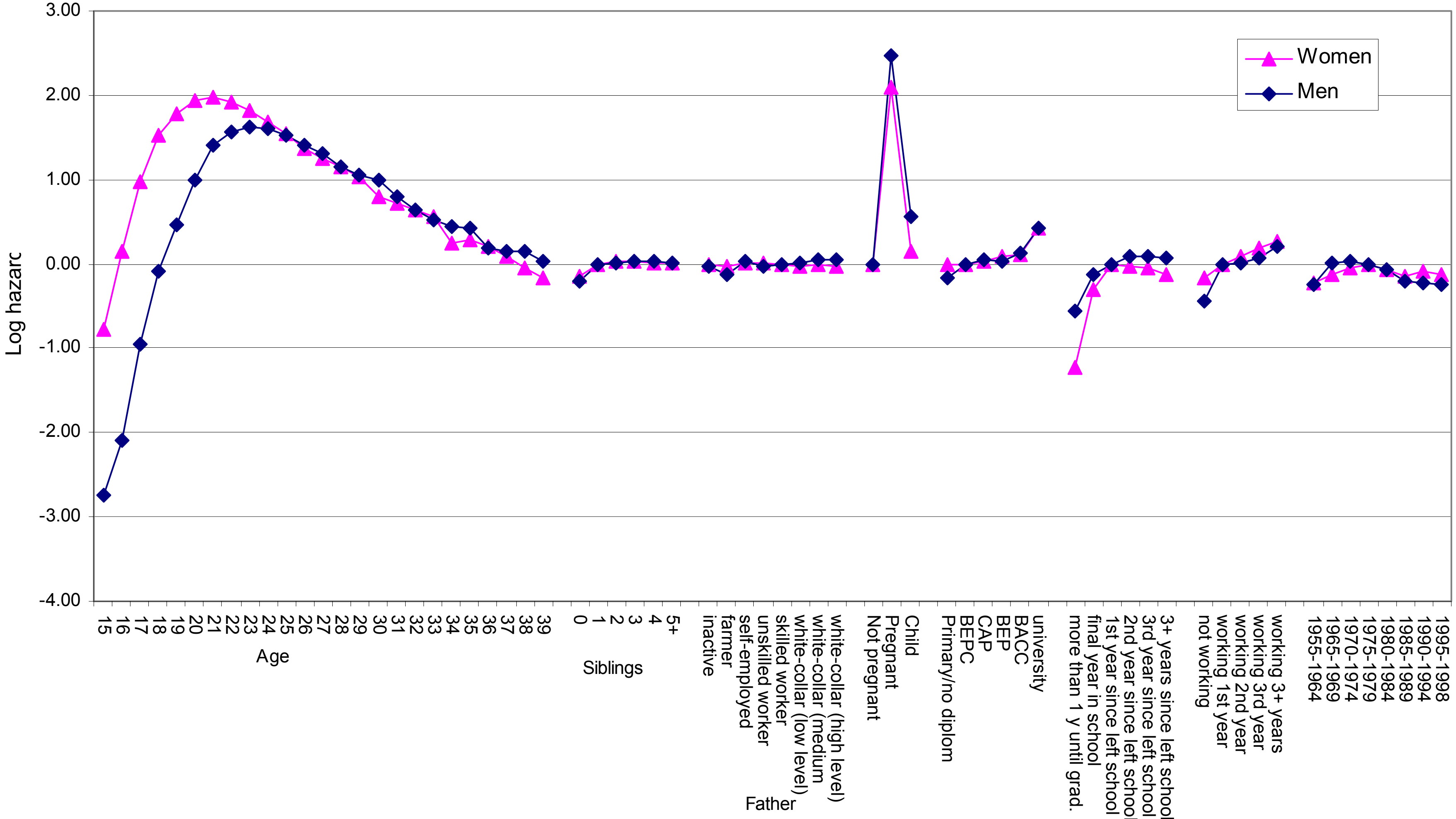

Figure 1: Estimated effects on first union intensities. The estimated values corresponding to age are shifted by a value of 6 in order to enhance visibility.

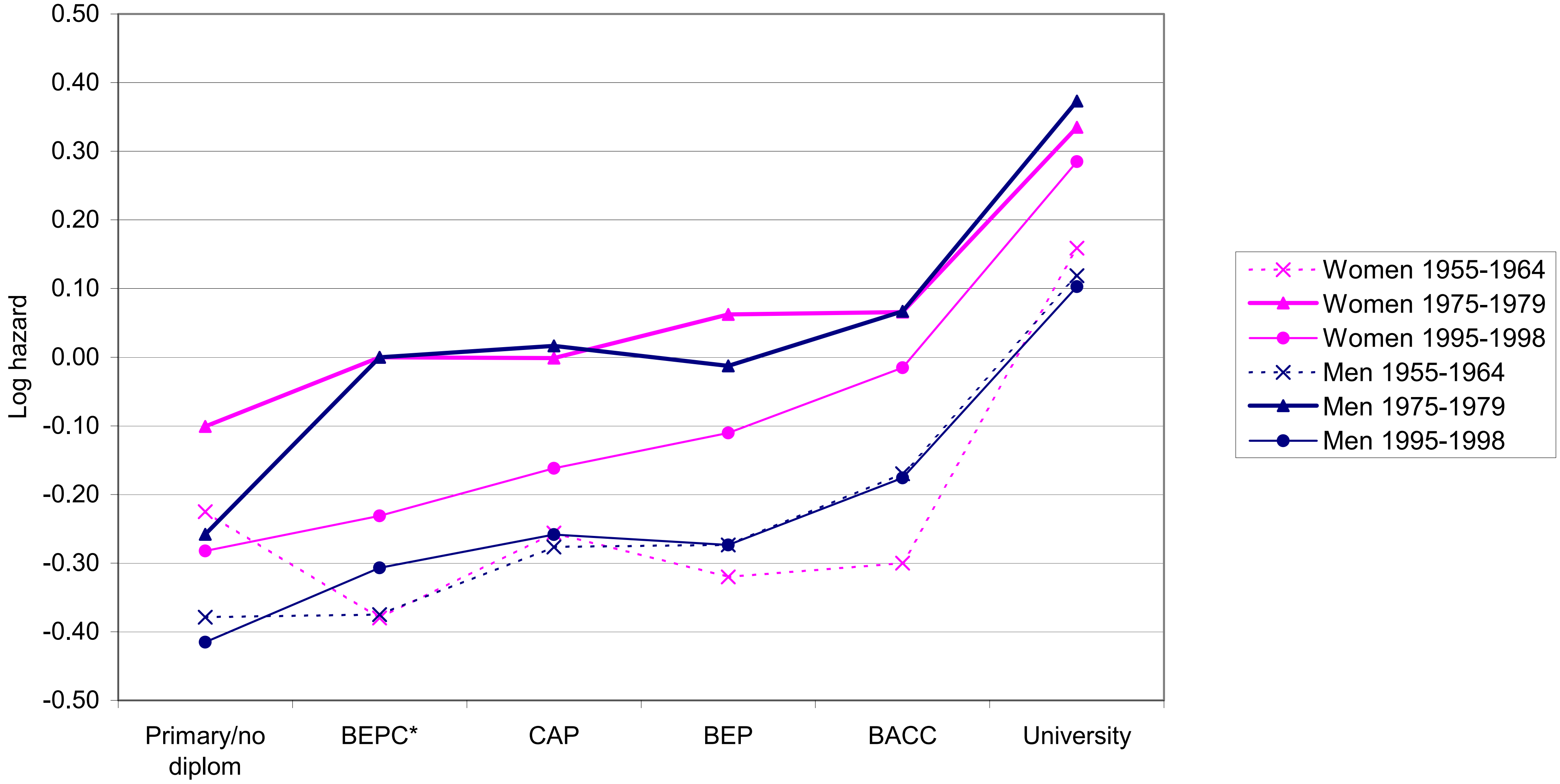

Figure 2: Estimated effects of level of education attained for various years. The asterisk denotes the baseline category.

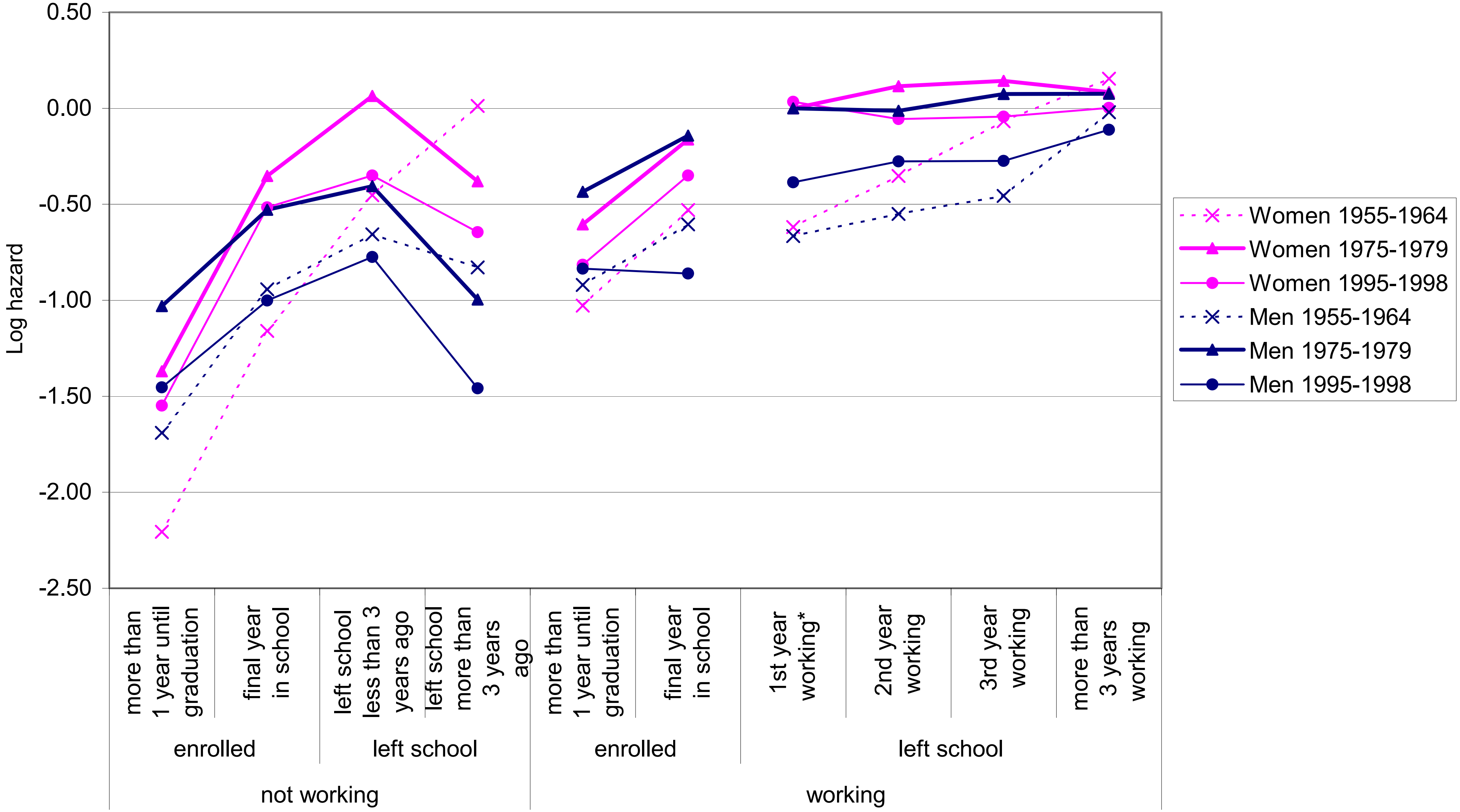

Figure 3: Estimated effects of interaction of enrolment and employment status for various years. The asterisk denotes the baseline level.