Implications of the Shifting Gender Imbalance in Higher Education for the Timing and Likelihood of First Union Formation

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Abstract: A major social trend of the past decades has been the reversal of the gender gap in education: while women were a minority in higher education in the past, the situation has gradually turned around. The increasing number of highly educated women entering the mating market relative to men is expected to have implications for mating patterns. Here, using data from the third round of the European Social Survey, we investigate whether and how the shifting gender imbalance among the highly educated is associated with rates of first union formation and first marriage in the cohorts born between the 1950s and 1970s in 20 European countries. On top of modelling overall transition rates, we also address the two underlying dimensions of these rates separately, namely the likelihood of first union formation and the timing of it. Our basic expectation, derived from the marriage squeeze theory, is that the oversupply of highly educated women compared to highly educated men would lead to a lower likelihood of union formation for highly educated women and a higher age at union formation. We also derive two competing hypothesis for highly educated men. Following marital search theory (Oppenheimer 1988), an oversupply of highly educated women compared to highly educated men should lead to a higher likelihood of union formation and a lower age at union formation. Following the sociocultural theory (Guttentag and Secord 1983) an oversupply of highly educated women compared to highly educated men should lead to a lower likelihood of union formation and a higher age at union formation. However, we hardly find support in our results for the marriage squeeze perspective and the derived hypotheses.

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1 Introduction

In Europe, college education has expanded rapidly since the 1960s and has done so more for women than for men. An important consequence of this development is that differences in the relative educational attainment of men and women have changed. In the past, men were typically more highly educated than women, but from the 1970s the gender gap in higher education began to shrink and turned to the advantage of women in the mid-1990s (Vincent-Lancrin 2008; Schofer and Meyer 2005). This implies that in Europe, as in the United States, there are more highly educated women than highly educated men entering today's marriage market (Esteve, García-Román and Permanyer 2012; Grow and Van Bavel 2015). Following Van Bavel (2012), we expect that this will affect the timing and likelihood of union formation in Europe.

When maintaining the traditional pattern of assortative mating, i.e. men marrying women who are at most as highly educated as themselves and women marrying men who are at least as highly educated as themselves, the shifting gender balance in higher education implies that highly educated women will find less eligible partners on the marriage market and increasingly suffer a marriage squeeze. The reversal of the gender balance in higher education would on itself lead to a negative relationship between education and marriage for women and a positive relationship for men (Van Bavel 2012).

Yet, research in the United States has found no decline in the likelihood of marriage among highly educated women. In the United States, it appears that a shift in patterns of assortative mating has allowed the marriage market to absorb the increasing number of highly educated women (Rose 2004; Schwartz and Han 2014; Schwartz and Mare 2005). A higher education is associated with a later age at marriage – and nowadays a later age at first union formation-, but not with a lower chance to form a union (Manning, Brown and Payne 2014; Qian and Preston 1993).

A similar concern about the marriage prospects of highly educated women recently appeared in East Asia, where traditional patterns of assortative mating still dominate and gender specialization remains a basic feature of marriage. In Japan and China marriage rates for highly educated women are low and the shifting gender balance in higher education contributes to the negative educational gradient in marriage for women (Qian and Qian 2014; Raymo and Iwasawa 2005). However, in Taiwan and South Korea, highly educated women became more likely to marry despite facing a smaller pool of eligible men. In Taiwan and South Korea the positive educational gradient in marriage is accompanied with a strong increase in homogamy among the highly educated, causing a trend toward more social closure among the highly educated (Cheng 2014; Park and Smits 2005).

In this paper we examine for Europe whether and how effects of the educational levels of men and women on rates of first union formation interact with the shifting gender balance in higher education. On the one hand, we observe for Europe that couples where the woman is more highly educated than the man are becoming more prevalent than couples where the man is more highly educated than the woman (Esteve et al. 2012; Grow and Van Bavel 2015). On the other hand, for several European countries the effect of women's education on the chance of forming a union is (still) negative (Dykstra and Poortman 2010; Kalmijn 2013; Wiik and Dommermuth 2014), suggesting that the relative improvements in women's educational attainment are not accompanied by convergence in the criteria that men and women use to evaluate the educational attainment of potential partners. If this is the case, the shifting gender balance in higher education will result in a mating squeeze for highly educated women and enhance the negative educational gradient in union formation for women and the positive educational gradient in union formation for men (Van Bavel 2012).

We estimated semiparametric survival models with country fixed effects to test whether the shifting gender balance in higher education, as a macro-level condition, is associated with rates of entry into a first union at the individual level. Since event history models address both the timing and likelihood question jointly, we also investigated both components separately using linear and binary logistic regression. Given the spread of cohabitation in many countries, our focus is on first union formation. However, considering that first union and marriage may represent qualitatively different types of partnership formation (Wiik and Dommermuth 2014), we also conducted a parallel analysis of first marriages. Throughout the paper, when we talk about union formation, it is meant to include both unmarried cohabitation and marriage.

The data come from the third round of the European Social Survey (ESS3 - 2006) which include information on first union formation and first marriage for 20 European countries. The IIASA/VID Educational Attainment Model is used to reconstruct the gender balance in higher education by cohort and country. Before we formulate hypotheses about the influence of the gender gap reversal in higher education on union formation, we introduce the concept of marriage squeeze and discuss the educational gradient in union formation in Europe. Next, we describe data and methods. The result section presents extensively descriptive results and findings coming from models applied. Finally, conclusions and suggestions for further research in the field are provided.

2 Background and hypotheses

2.1 The marriage squeeze: the concept and earlier studies

The phrase marriage squeeze was coined by Glick, Heer and Beresford in 1963 to describe an imbalance between the numbers of males and females in the prime marriage ages. They observed that a sharp rise in birth rates during the postwar period combined with the fact that women marry men who are on average two or three years older resulted twenty years later in a disproportion between the number of potential brides and the number of potential grooms. This shortage of suitably aged men placed women in a marriage squeeze. As a result, they speculated that some women would have to postpone marriage and eventually marry a man of a less suitable age or not marry at all.

In the first marriage squeeze studies suitability of potential partners was only defined by age (Akers 1967; Muhsam 1974; Schoen 1983). In the 1980s also race came into the picture when Spanier & Glick (1980) and Guttentag and Secord (1983) stated that differences in marriage behaviour between black and white Americans partly resulted from black-white differences in marriage market opportunities. Especially in the 1970s, the shortage of black men was acute and brought on lower marriage rates for black women and higher divorce and illegitimacy rates (Crowder and Tolnay 2000; Lichter, Leclere and Mclaughlin 1991; Lloyd and South 1996). Wilson (1987) took this a step further and added that high black male mortality rates, combined with high black male unemployment rates, compromised the proportion of black men who are in the position to support a family. A shortage of economically attractive black men caused black women to postpone or even to forgo marriage.

Most marriage squeeze studies focus on marriage outcomes for women. In first instance, the concept of marriage squeeze was used to clarify declining marriage rates of women in the 1960s, but along the line it has been updated according to new research findings. In the United States, the link between changes in the availability of suitable spouses and the decline in marriage among minority and low-income populations has been most often investigated. A shortage of economically stable men, measured by their social characteristics such as labour force participation, income and educational attainment (Goldman, Westoff and Hammerslough 1984; Qian and Preston 1993; Schoen and Kluegel 1988; South and Lloyd 1992) is found to play a significant role in widening racial and socioeconomic differences in marriage rates (Fossett and Kiecolt 1993; Guzzo 2006; Lichter et al. 1992; South and Lloyd 1992).

In the literature, two explanations can be found to shed light on the effect of marriage market opportunities on marriage behaviour. The first explanation, marital search theory

(Oppenheimer 1988), postulates that the delayed timing of marriage stems mainly from the difficulties people encounter in mating assortatively. When and if a mate is found depends on the efficiency of the selection or search process. This efficiency is determined by the numbers of potential suitable partner available on the marriage market and by a person's minimum acceptance level. Oppenheimer (1988) presumes that men and women equally value and seek out marriage. For both sexes, it is the case that when few potential partners are available, the transition to marriage will be delayed and perhaps forgone entirely.

A second explanation, known as the sociocultural theory or imbalanced sex ratio theory (Guttentag and Secord 1983), emphasizes men's and women's conflicting familial goals brought on by the structural power that is held by men. Guttentag and Secord (1983) argue that members of the sex in short supply have a stronger position because a greater number of alternative relationships are available to them. This power, referred to as dyadic power, allows to bargain more favourable outcomes within the dyad. Because of this enlarged availability, members of the scarcer sex will be less committed to existing relationship, choosing to end them more frequently for alternative relationships. In this conception of dyadic power, the social consequences of high and low sex ratios are the same for both sexes. To explain the historical observed gender differentials in responses to sex ratio imbalances the authors look for another source of control, called structural power. Structural power incorporates the political, economic, and legal power in a society and shapes moral values and practices. In nearly all societies, men have been in possession of this forceful source of control and used their structural power to modify women's use of dyadic power by constraining women's access to alternative mates. Guttentag and Secord (1983) hypothesize that when women outnumber men, the latter have the bargaining power and can secure sexual relationships without commitment. As a result, marriage rates for women and for men will be low. When men outnumber women, women use their bargaining advantage to marry. Because of women's relative scarcity, men are motivated to commit to marriage. As a result, women's and men's marriage rates will be high.

Research on the effect of sex ratios on men's marriage behavior is scarce, but all the more interesting, since it sheds light on the alternative theoretical frameworks that guide research on the impact of sex ratios on family formation. Lower male marriage rates in case of a high supply of women were indeed found by several scholars in the United States (Angrist 2002; Schmitt 2005; Uecker and Regnerus 2010; Warner et al. 2011). However, Lloyd and South (1996) who studied the effect of sex ratios on men's marriage behavior at the individual level, reported that an oversupply of women had increased men's marriage chances. Cready, Fossett and Kiecolt

(1997) and Albrecht and Albrecht (2001) found a curvilinear effect of the sex ratio on men's chances of marriage, with low marriage odds when women are plentiful or scarce and high marriage odds in a balanced marriage market.

Not only for men but also for women empirical research on the marriage squeeze presents a mixed picture about the influence of unbalanced sex ratios on marriage. Results are often inconsistent, depending on how mate availability was computed, what the framework of the analyses was, and which questions were addressed (see De Hauw, Piazza & Van Bavel 2014).

Since we focus on changes in marriage market opportunities caused by the reversal of the gender gap in higher education, we adopt the education-specific mating squeeze concept introduced by Van Bavel (2012). The education-specific mating squeeze is an upgrade of the marriage squeeze concept which incorporates besides age and sex also education and union status, two important characteristics for studying partnership and family formation today. Given that unmarried cohabitation is on the rise and has attained a status similar to marriage in many European countries, we will look at the effect of the shifting gender balance in higher education for union formation (married and unmarried couples together).

2.2 Preferences and the educational gradient in union formation

Marriage market arguments focus on the demographic conditions on the marriage market. Yet preferences also play a role in union formation. Becker's (1981) economic approach has been extremely influential for theorizing about partner preferences in demographic research. According to Becker (1981) the gains from marriage are maximized when partners are alike for complementary traits like physical capital, religion, social origin and education, and different for substitutable traits. It follows from the household division of labour that market work of men and household work of women are substitutable traits. Becker categorizes education as a complementary trait, but given its connection with labour market opportunities and income, education has commonly been considered a substitutable trait. As women prefer men with good labor market prospects, they compete for men with high levels of education. Men, on the other hand, are looking for a wife who can take care of the household and family. Thus, in this framework, a strong labor market position and a high education hardly represent trading value on the marriage market for women (Blossfeld 2009; Eeckhaut et al. 2011; Schwartz 2013).

Becker's gender role specialization is losing its explanatory power for behavior related to union formation. Instead pooling resources is argued as an adequate strategy of couples' adaptation to new challenges in the labour market (Oppenheimer 1997). This is expected to change the association between education and union formation. Increasing women's role as an economic provider defines the importance of women's economic potential as a spouse selection criteria, which should lead to a positive relationship between women's educational attainment and marriage (Oppenheimer 1997; Sweeney 2002). In addition, with women's growing economic independence, men's earning potential and education may have become relatively less important for their chances on the marriage market. If women place less weight on men's education, women's preferences for highly educated men should decrease (Buss et al. 2001).

Several studies confirmed that in the United States a reversal in the effect of women's educational attainment on the likelihood of marriage has taken place (Goldstein and Kenney 2001; Torr 2011). While in the past highly educated women were the least likely to marry, they are the most likely to marry today. Highly educated men are still the most likely to marry, as was already the case in the past. However, Sassler and Goldscheider (2004) observed a decline in the positive effect of education on marriage chances for men.

Less empirical findings exists for other Western countries and on the likelihood of ever forming a coresidential union. A study conducted in the Netherlands (Dykstra and Poortman 2010) shows that education still has a negative effect on the likelihood to ever form a union for women and a positive effect on the likelihood to ever form a union for men. Better educated women and less educated men were the most likely to remain single, with the exception of university educated men. The latter's chances of remaining single were similar to men with only primary education. The effects of education did not change over time or when analyzing marriage instead of union formation. Results for Norway by Wiik and Dommermuth (2014) are similar to those of the Netherlands. Highly educated women and low educated men were the least likely to ever form a union formation or marriage. In Norway, the positive effect of education on men's likelihood to form a union has decreased over the cohorts, suggesting that highly educated men are increasingly more likely to remain single. A change across cohorts was not found for women.

Kalmijn (2013) examined the educational gradient of being in a union during midlife (ages 40-49) among 25 European countries and showed that differences in the educational effects on union formationare related to several societal characteristics. In countries where gender roles are traditional, highly educated women are the least likely to be in a union at age 40-49, while for men, the educational gradient is absent. In countries where gender roles are more egalitarian, highly educated women and highly educated men are more likely to be in a union.

In most countries education leads to a delay in marriage for both men and women. The highly educated postpone marriage because they have been in school longer (Blossfeld and

Huinink 1991). The effect of education on the timing of first union formation, thus including unmarried cohabitation as well as marriage, is less marked (Liefbroer and Corijn 1999). In general, union formation is often less strongly associated with education than marriage (Kravdal 1999; Wiik and Dommermuth 2014).

2.3 Hypotheses on the education-specific mating squeeze

Our analysis will test a number of hypotheses that are related to the education-specific mating squeeze. The overall concept behind the hypotheses formulated is that as the gender balance in higher education changes, it will influence union formation rates in the population. Below we listed hypotheses for first union formation, which will be tested separately for union formation in general and for marriage specifically. Based on the marital search theory, it is hypothesised that:

- ❖ Hypothesis 1: An increase in the gender balance in higher education in favour of women is negatively related to first union formation rates of highly educated women. Since increased numbers of highly educated women are looking for a partner with the same educational level, the relatively lower number of potential partners on the mating market may result in lower rates of first union formation for highly educated women.
    - ➢ Hypothesis 1a: Additionally, we expect that lower rates of union formation among highly educated women are the result of postponement of union formation. Therefore, sub-hypothesis H1a says that an increase in sex ratio among the highly educated is positively associated with the age of union formation of highly educated women.
    - ➢ Hypothesis 1b: Lower rates of union formation may also be due to lower proportions of women entering a union. As opposed to the effects of timing, this will result in fewer highly educated women ever establishing a partnership. Thus, H1b claims that an increase in the sex ratio among the highly educated is negatively associated with the probability that highly educated women ever form a union.
- ❖ Hypothesis 2: Analogously to H1, but now for men, we hypothesize that an increase in the gender balance in higher education in favour of women is positively associated with highly educated men's union formation rates. In this case we expect that among the highly educated there is a tendency towards homogamy and the increasing numbers of highly educated women, on the one hand, become a "supply" for highly educated men, but on the other hand there is also an increasing demand for highly educated men as the numbers of highly educated women go up.

- Hypothesis 2a: We hypothesise that the mechanism given in H2 influences the timing of men's union formation. For highly educated men, since they are in "higher demand", the search period is shortened and this increases the rates of union formation. Therefore, H2a says that an increase in the sex ratio is negatively associated with highly educated men's age at first union formation.
- Hypothesis 2b: It is also possible that higher rates of union formation in H2 are the result of increasing proportion of highly educated men who form a union. To test this, H2b states that an increase in the sex ratio among the highly educated is positively associated with the probability that a highly educated man has formed a union.

The socio-cultural theory (Guttentag and Secord 1983) suggests that men react differently to mating market imbalances, because of the unequal division of structural power in favour of men. When mating opportunities are high, union formation rates for men are expected to be low because the numerical abundance of women discourages men to commit to one women as there is sufficient supply of potentially attractive alternatives. Hence, more men and women will remain single and when they partner, they partner later in life. Based on the sociocultural theory we formulate an extra hypothesis for men, which is competing with Hypothesis 2:

- Hypothesis 3: An increase in the gender balance in higher education in favour of women will result in lower union formation rates for highly educated men, due to an increase in the age at union formation (H3a) and/or a decrease in the proportion of highly educated men who ever formed a union (H3b)

3 Data, Measures and Method

3.1 Data

The data come from the third round of the European Social Survey (2006),[1] which contains a module called 'the timing of life'. Respondents were asked the following questions: 'Have you ever lived with a spouse or partner for three months or more?', 'In what year did you first live with a spouse or partner for three months or more?', 'Are you or have you ever been married?', and 'In what year did you first marry?'. This information allowed us to examine entry into first union formation and first marriage.

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[1] ESS Round 3: European Social Survey Round 3 Data (2006). Norwegian Social Science Data Services, Norway – Data Archive and distributor of ESS data for ESS ERIC.

The data cover 20 countries from different regions of Europe (Austria, Belgium, Bulgaria, Denmark, Estonia, Finland, France, Germany, Hungary, Ireland, the Netherlands, Norway, Poland, Portugal, Slovakia, Slovenia, Spain, Sweden, Switzerland and the United Kingdom). We selected respondents born between 1950 and 1975, aged 31 to 57 years old. Age 31 as a minimal age has the advantage that the majority of men and women have completed their education and formed a union by then. We deleted respondents who were younger than 16 when they first formed a union and respondents for whom information on gender was missing (N=16). After this selection, the weighted data set contained 7921 male and 9087 female respondents[2]. To investigate the timing and likelihood question separately we raised the minimum age of the respondents to 40 years and, as a result, narrowed the cohort range to 1950-1967.

To compose the gender balance in higher education, the IIASA/VID data is used (K.C. et al. 2010; Lutz et al. 2007). IIASA/VID provide reconstructions (for the period 1970–2000) and projections (for the period 2005–2050) of the distribution of educational attainment in five-year intervals for five-year age groups in a large number of countries. Following De Hauw, Grow and Van Bavel (2015), we linearly interpolated the numbers of individuals for the different levels of educational attainment to obtain yearly measures.

3.2 Independent variables

In ESS, educational attainment is harmonized across countries based on the International Standard Classification of Education (ISCED). ESS3-2006 used five categories to measure respondent's highest educational level. We recoded educational level into three larger categories. This somewhat reduces the amount of detail in measuring educational attainment, but facilitates comparison of countries with different educational systems. First we collapsed less than lower secondary education (ISCED 0-1) and lower secondary education completed (ISCED 2) into *low educated*. The lower secondary are included in the low education category to do more justice to the fact that this educational level is part of basic education in many countries. Second, individuals were classified as *medium educated* when they completed upper or post-secondary education (ISCED 3 and 4). Post-secondary education has been included in the medium education category since this category is too small to stand on its own. Third, *highly educated* consist of respondents who completed tertiary education (ISCED 5 and 6).

Our key explanatory variable represents the gender balance in higher education in the country and cohort of the respondent. It is measured in the year when the respondent turned 30

[2] The design weights provided by the ESS were used to adjust for unequal probabilities of selection in the survey sampling design.

years of age, i.e., at an age when the vast majority of individuals has usually completed fulltime education and the cohort-specific gender distribution by educational attainment can be determined. Using IIASA/VID data, we calculated for each respondent the sex ratio among highly educated women and highly educated men by dividing the number of highly educated women who were 25–34 years old (F_{High}) by the number of highly educated men who were 27–36 years old (M_{High}) for the year in which the respondent was 30 years old.[3] We opted for a ten-year age interval instead of the five-year age interval that has often been used in earlier research (Fossett & Kiecolt, 1991). This larger age interval is more robust to erratic fluctuations caused by sampling errors. In addition, five-year age intervals may fail to account for the fact that people may look in adjacent age categories when they do not find a mate in their own age group (De Hauw, Piazza and Van Bavel 2014). We took the log of this sex ratio (i.e. $\log(F_{High}/M_{High})$) to make the measure symmetric around the value of zero, which represents a balanced mating market. Because we divided the number of women by the number of men, a positive value means that highly educated women are more numerous than highly educated men. A negative value, by contrast, represents a mating market where highly educated men outnumber highly educated women. For brevity, we refer to this measure also simply as ‘the sex ratio’.[4]

Note that our sex ratio measure only focuses on the gender imbalance in tertiary education and that we examine how low, medium, and highly educated respondents are affected by this aspect of the mating market. The reason is that in the European context, the important changes in the relative educational attainment of men and women have occurred in the distinction between the college educated and those with less education. In addition, sex ratios for the highly educated correlate strongly with sex ratios for the medium and the low educated (De Hauw, Grow and Van Bavel, 2015).

We included information about respondents’ birth cohort in the analysis to control for possible cohort effects. The cohort variable is dummy coded based on respondents’ year of birth in five-year intervals between 1950–1976. Furthermore we controlled for the age of the respondent at the time of interview to capture any monotonous cohort changes that are not

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[3] The IIASA/VID data is based on five-year age groupings (e.g., 25–29 years, 30–34, etc.). We therefore had to approximate the number of highly educated men who were 27–36 years old in a given year. We did so by taking the number of highly educated men of men who were 30–34 years old in a given year and added to this 60% of the number of men who were 25–29 years old and 40% of the number of men who were 35–39 years old.

[4] To examine the possibility of a curvilinear relationship, we initially included a quadratic variable for the sex ratio in our models. Since this variable proved to be non-significant and did not alter the results, we excluded this variable from the analyses.

captured by the cohort dummies and we controlled for those individuals who are still enrolled in education.

3.3 Methods

Three distinctive types of regression analysis are presented. In the first we employed an event history approach and estimated Cox proportional hazards models of entry into first union and first marriage. A limitation of event history models is that they mix the timing and the quantum, i.e. we cannot distinguish whether some of the covariates act more towards postponement of union or clearly limit the number of events that would ultimately happen (Bernardi 2001). This could be problematic since change in the gender balance in higher education may have diverging effects on these two components: for example a positive effect on the eventual probability of union formation but a negative effect on the speed of making the transition (Van Bavel 2012).

To disentangle the timing and the quantum from rates of union formation, we addressed these components separately. The second type of regression analysis focused solely on the probability that a person had ever formed a union or entered a marriage. In this case, timing of an event was ignored and an ordinary logistic regression was employed on the binary outcome variable, the latter indicating whether a person had ever formed a union by age 40 at the latest. We estimated these models for men and women who were at least 40 years old at the time of interview. Only unions and marriages before age 40 were counted as events. Unions formed at higher ages were censored in order to allow the same amount of exposure time to all cohorts.

In the third type of regression analysis we focused solely on the timing aspect of union formation and marriage. That is, only individuals who had ever formed a union before age 40 entered the analysis and the time to event in continuous scale was the dependent variable. The absence of censoring allowed us to use simple linear regression modeling. To obtain a congruent dataset as used in the second part of analysis, we excluded respondents who are younger than 40 and considered only time to event that had happened before age 40.

To control for the potentially confounding influence of unobserved country characteristics, we included country fixed effects in all regression models. Taking into consideration the hierarchical nature of the data, we adjusted standard errors for the non-independence of observations nested within countries.

We modeled men and women in separate models. The gender-specific models are more straightforward to interpret than pooled models, as it is not necessary to account for a different educational gradient in union formation between men and women by means of complex interaction effects. The central point in the regression models is the association between educational level and the macro-level sex ratio of the highly educated. For this reason, an interaction term between the sex ratio variable and individuals' own educational attainment was included.

4 Results

4.1 Descriptive results

The change in the educational gender balance has developed over many birth cohorts and from country to country this process has not developed simultaneously. The data used in this analysis cover 20 European countries and birth cohorts since the 1950s. In addition to international differences in the gender balance in education, the included countries are not homogeneous in their background of union formation and marriage. One of the main differences is that in Western and Northern Europe the retreat from marriage started earlier. Marriages were postponed or foregone in favour of non-marital cohabitation. Other regions of Europe have later followed this process (Lesthaeghe 2010). It is therefore expected that across countries we observe varying discrepancies between ages at first union formation and first marriage. While our regression analyses focus on the dynamics over birth cohorts, this international heterogeneity cannot be ignored. The differences manifest themselves mostly in the patterns of non-marital cohabitation (see Sobotka and Toulemon 2008; Wiik 2009).

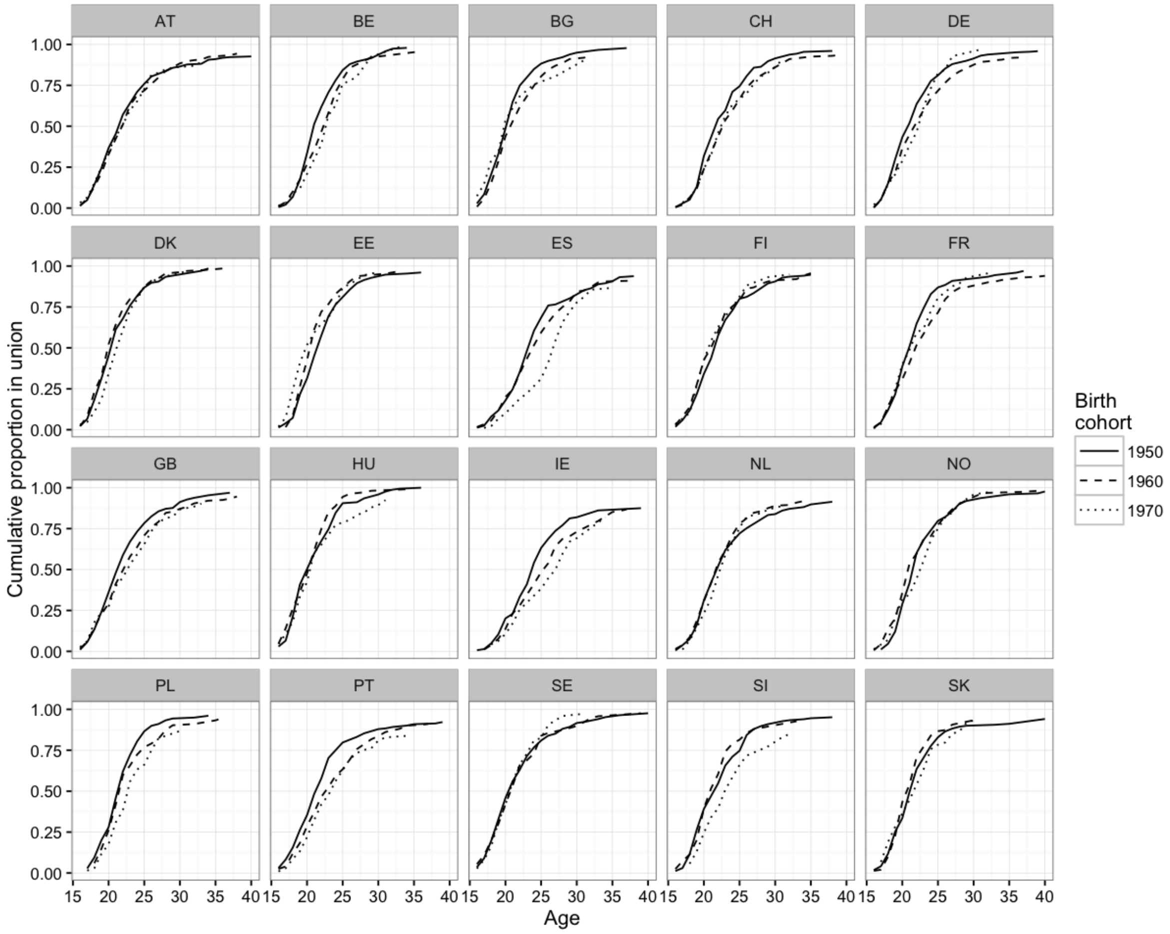
In this section, we describe the cross-country differences in cohort patterns of entry into first union formation and first marriage, and cross-country differences in the timing and quantum of both union formation and marriage. As in the subsequent regression analysis, the timing and quantum of events are assessed for the subsample that is at least 40 years old (born 1950s – 1967) and we only take into account events that have occurred until age 40. At the end of the section, descriptive statistics on the changing gender balance in higher education are presented.

4.1.1 Cohort patterns in first union formation and first marriage

Figure 1 shows the age-cumulative proportions of first union for women by 10-year birth cohorts. The general pattern is that there is a slight postponement of first union formation in the later cohorts. This trend is especially noticeable in Spain, Ireland, and Portugal. As a contrast, in Estonia, women in successive cohorts actually exhibit a decreasing age at first union formation. The latter is in line with previous findings (Katus et al. 2007), so it does not indicate a problem with the data. Also in some other Central and East European (CEE) countries, such as Hungary, Slovenia, and Slovakia we can observe some decreases in the age at first union formation. As of the proportions of women that have ever experienced a union by age 40, there are no big variations across countries and across cohorts. For some countries like Great Britain and Poland we observe a lower proportion of women ever in a union, but for most countries the difference between cohorts is negligible.

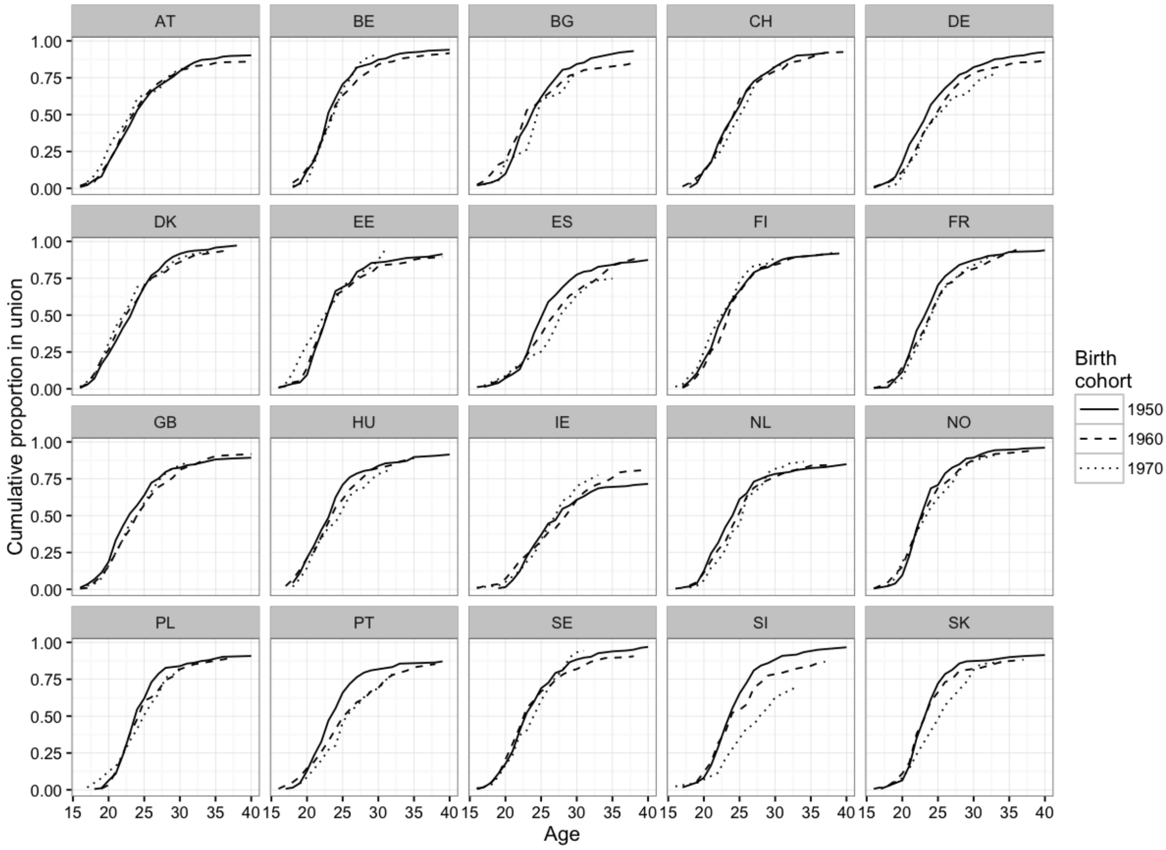
Age-cumulative proportions of men’s first union are shown in Figure 2. Postponement of first union formation across birth cohorts is more present in Spain, Portugal, Slovenia and Slovakia. Compared to women’s respective figures, one of the characteristics of male first union formation is the rectangular shape of the 1950s curve, as seen for instance in Estonia, Poland, Slovenia and Slovakia. A high proportion of first union formation occur within a narrow age range in the first half of the twenties. Later cohorts seem to introduce more variability in the timing of union formation and the rectangular shape is replaced with a less steep curve of cumulative proportions.Thus, depending on the country, the timing of first union may or may not be responsive to cohort-to-cohort changes.

Figure 1 Age-cumulative proportions of first union, women



Source: ESS3-2006, sampling weights, own estimation
Note: “1950” refers to the cohorts born in the 1950s, etc.

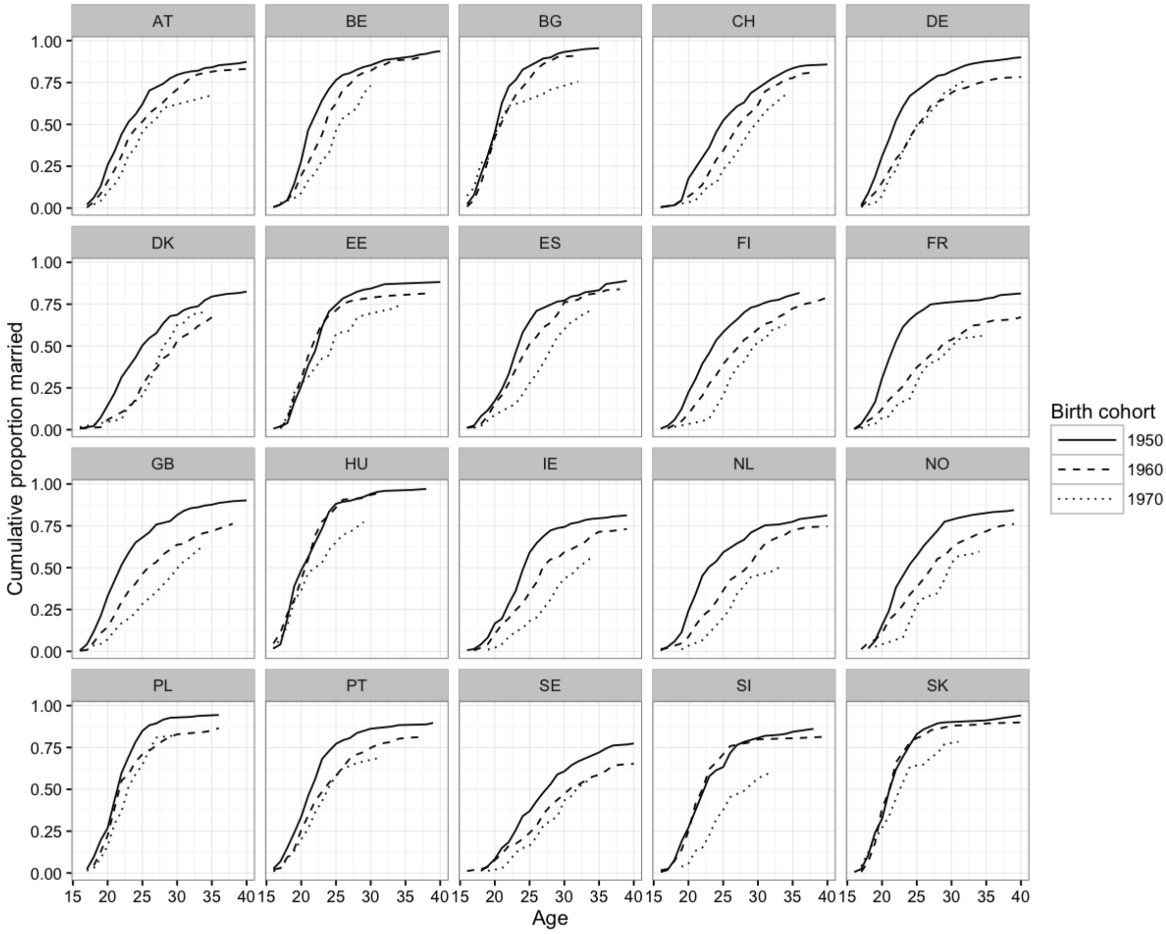
Figure 2 Age-cumulative proportions of first union, men



Source: ESS3-2006, sampling weights, own estimation
Note: “1950” refers to the cohorts born in the 1950s, etc.

Turning now to first marriage formation, we notice more variability across countries and across cohorts. Figure 3 depicts the age-cumulative proportions of first-married women by 10-year birth cohorts. In all countries we observe postponement of first marriage formation and in most countries there is a decline in the proportion of ever-married women. As an example of postponement, in Belgium the age when 50% of women have married has shifted by about five years between the cohorts of 1950s and 1970s. The lowering proportions of ever-married, together with increasing age at marriage, can be well seen in France, Great Britain, Norway, and Sweden. Most Western and Northern European countries show strong postponement and declining levels of marriage across the cohorts. Among the CEE countries, these tendencies appear mostly in the 1970s cohort. Before 1970, marriage was widespread in these countries, which is illustrated by an almost indistinguishable difference between the 1950s and 1960s cohorts in some of the CEE countries.

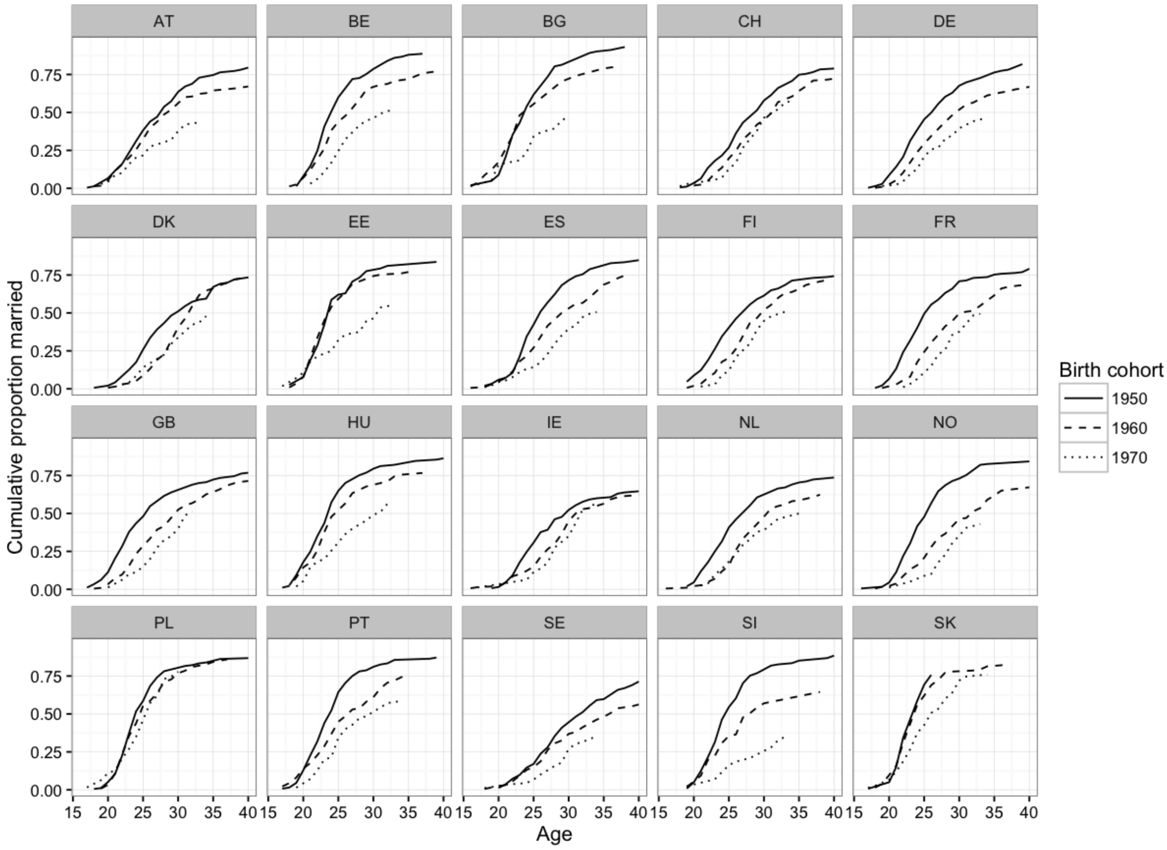
Figure 3 Age-cumulative proportions of first marriage, women



Source: ESS3-2006, sampling weights, own estimation
Note: “1950” refers to the cohorts born in the 1950s, etc.

The marriage patterns of men (see Figure 4) follows largely the cross-cohort trend of women. But in some countries, the contrast between successive cohorts is higher than it was seen for women. In several countries the proportion of men ever married drops to around 50% in the 1970s birth cohort. Only in Poland, the age-cumulative marriage pattern remains relatively unchanged and there is only a small drop in the levels of married men. This corresponds to low levels of Polish non-marital cohabitation that is observed also in earlier studies (Sobotka and Toulemon 2008; Matysiak 2009).

Figure 4 Age-cumulative proportions of first marriage, men



Source: ESS3-2006, sampling weights, own estimation
Note: “1950” refers to the cohorts born in the 1950s, etc.

We conclude from this subsection that first union formation is relatively stable across cohorts. The slight variations in the timing of first union formation seem to hardly influence the proportion of the population that will end up in a partnership at all. However, major changes have occurred in first marriage. In most countries, there has been postponement of marriage and a decline in proportions ever married, which has been slightly more visible for men. Yet, so far we have looked at the whole population by gender, without making any difference by educational levels. The stability or non-stability, shown in this section, may not apply to all educational levels equally. In the following, we will detail the timing and the quantum components of first union formation and first marriage, and link them with educational level of women and men.

4.1.2 Mean age at first union formation and first marriage

Table 1 shows basic descriptive statistics of ages at first union formation and first marriage for women and men by country. Mean age at first union is between 21 and 24 for women and between 24 and 26 for men. Age at first marriage is more spread out across countries, ranging from 21 to 27 for women and from 24 to 29 for men. Note that in some countries, like several CEE countries, there is very little difference between age at first union and age at first marriage. As a contrast, the gap is much bigger in Northern European countries (for example Denmark and Sweden). This is the result of the fact that non-marital cohabitation was more common in North and West Europe. In CEE, where direct marriage prevailed, the difference between first union and marriage timing is much smaller.

In addition, a relatively high mean age at first marriage is not necessarily indicative of a relatively higher mean age at first union formation. For example Denmark shows one of the highest mean age at first marriage, but the age at first union formation is among the lowest. This may be due to processes such as a long premarital cohabitation period or high selectivity into marriage (and hence a longer waiting time until marriage).

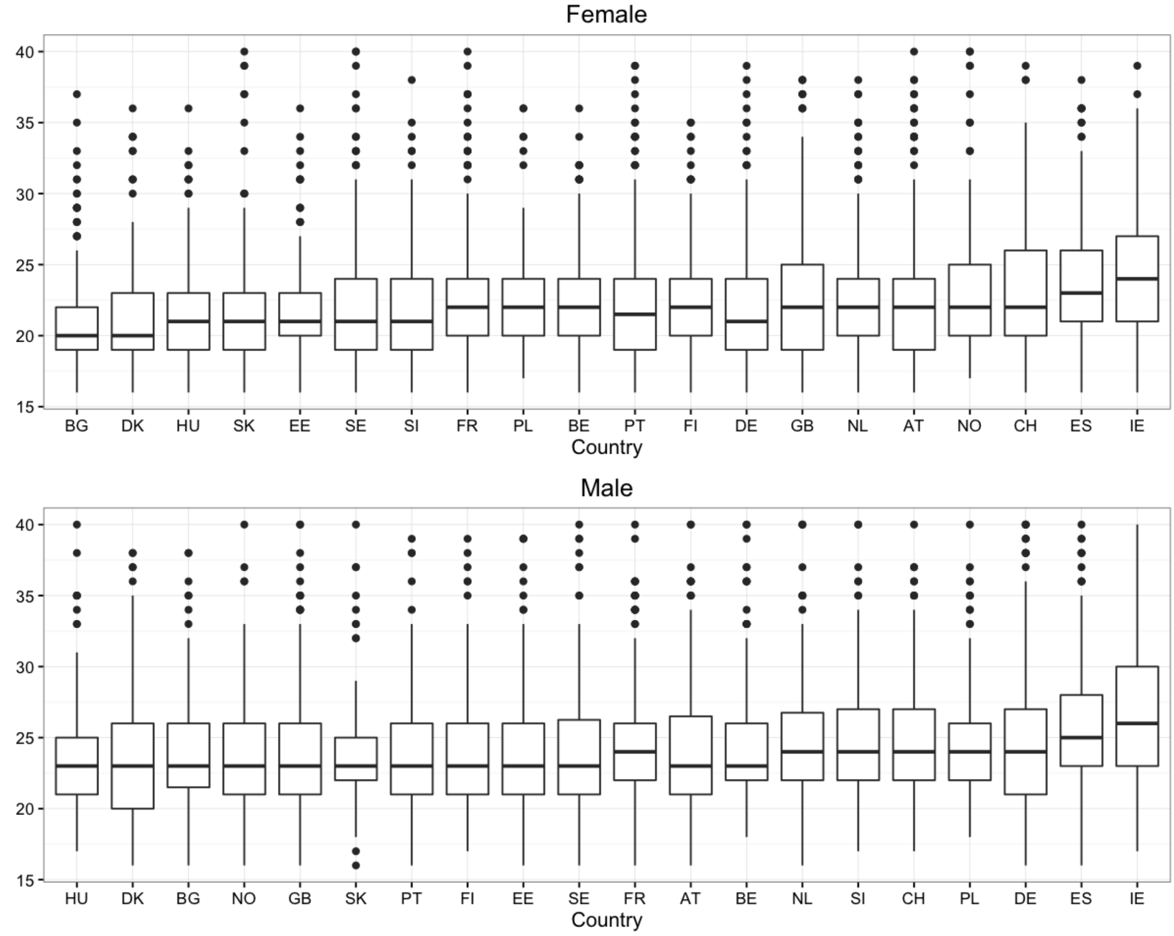
To examine the country differences in the distribution of mean age at union formation, Figure 5 shows the respective boxplot by gender for each country. For women and men, the countries are ordered by the mean age at first union formation, not by the median which is at the centre of each boxplot. For women, the order of the countries indicates generally lower ages in CEE. Also Denmark appears among countries with relatively early mean age at first union formation. Women's mean ages are the highest in Ireland and Spain, where the median age is 23–24. It is only in the countries of relatively high mean age at first union (Ireland, Spain, Switzerland, and Great Britain) where the upper quartiles reach and exceed age 25. In all other countries, three fourths of the first unions were formed before women reached age 25. Men's age at first union (lower part of Figure 5) are generally higher than women's. The first quartile for men is above age 20 in all countries except Denmark. Also, there are no countries where the upper quartile is below age 25 and in the countries on the right side of the graph the median age is 25 or higher. For women and men, some countries exhibit a larger range between quartiles than others. For instance, among women in Estonia and men in Slovakia the mean age at first union formation is distributed over a relatively narrow range, while in other countries this age range is more spread out.

Table 1 Mean and standard deviation of age at first union formation and first marriage, women and men who are at least 40 years old

		Women				Men			
		First union		First marriage		First union		First marriage	
		Mean	SD	Mean	SD	Mean	SD	Mean	SD
German speaking	AT	22.5	4.2	24.1	4.5	24.3	4.5	26.4	4.7
	CH	23.2	3.9	25.6	4.5	24.7	4.1	27.7	4.7
	DE	22.3	4.0	23.8	4.8	24.9	4.8	26.8	5.1
West Europe	BE	22.2	3.4	23.2	4.4	24.3	4.2	25.2	4.4
	FR	22.1	3.9	23.2	4.9	24.3	4.1	26.2	4.8
	NL	22.5	3.8	24.2	5.0	24.5	4.0	26.5	4.5
Nord Europe	DK	21.4	3.8	26.2	5.4	23.6	4.4	29.0	4.8
	FI	22.2	3.8	24.6	5.1	23.9	4.1	26.6	4.8
	SE	22.0	4.6	26.6	5.1	24.1	4.4	29.0	5.4
	NO	22.7	4.1	24.5	4.5	23.7	3.7	26.3	4.6
South Europe	PT	22.2	4.3	22.2	4.1	23.8	3.9	23.8	3.6
	ES	24.0	4.4	24.3	4.7	25.9	4.6	26.7	4.7
British Isles	GB	22.4	4.4	23.2	4.8	23.7	4.2	25.5	5.0
	IE	24.2	4.2	24.6	4.4	26.5	4.5	27.0	4.3
Central and East Europe	PL	22.2	3.3	22.2	3.3	24.8	3.6	24.9	3.7
	SI	22.0	3.7	22.6	4.2	24.7	4.2	25.4	4.3
	SK	21.8	3.7	22.1	4.0	23.8	3.4	24.0	3.4
	EE	21.9	3.3	22.4	3.6	24.0	4.1	24.1	3.6
	HU	21.4	3.5	21.6	3.6	23.5	4.2	24.1	4.2
	BG	21.2	3.4	21.2	3.3	23.7	4.3	23.9	4.3

Source: ESS3-2006, sampling weights.

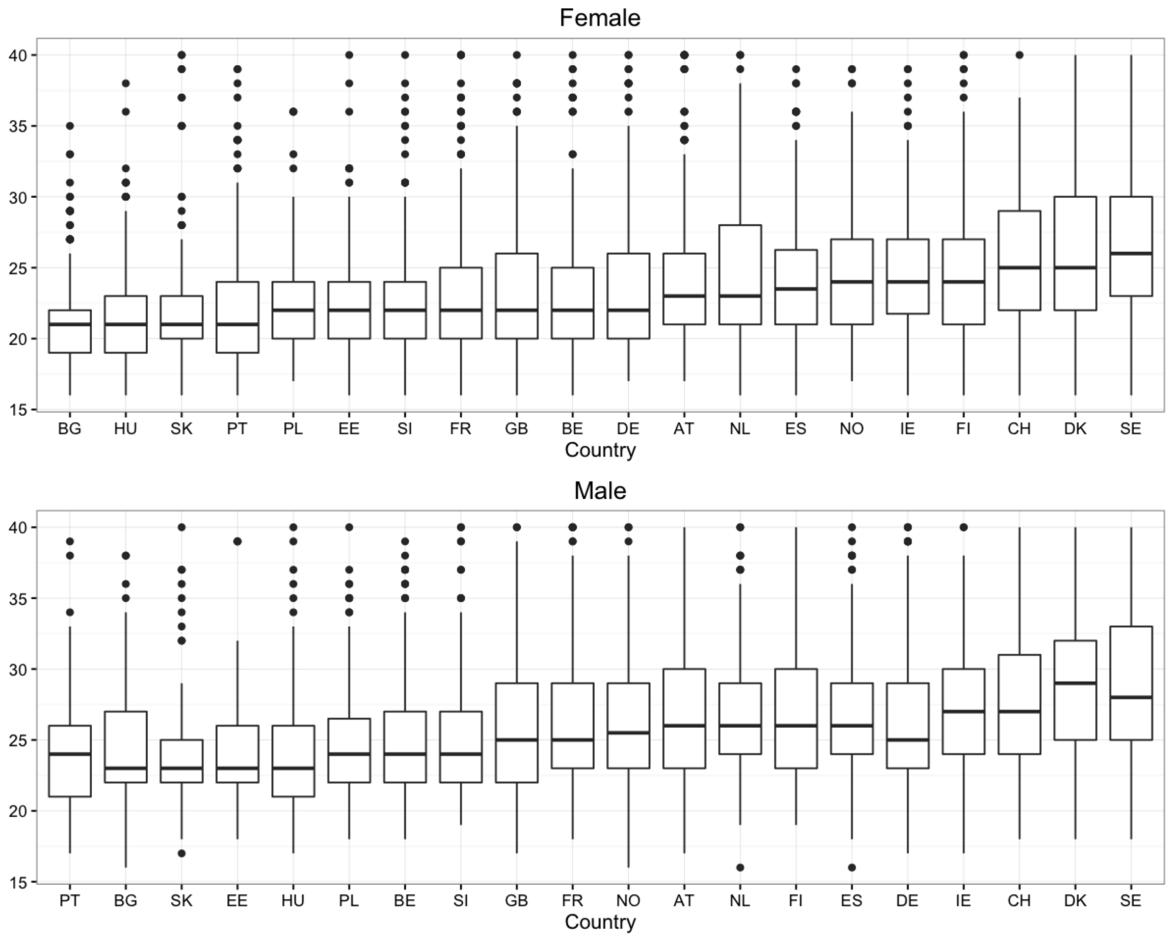
Figure 5 Boxplot of age at first union formation, women and men who are at least 40 years old



Source: ESS3-2006. Countries ordered by mean age at union formation, only unions up to age 40 considered.

Figure 6 shows the distribution of age at first marriage for women and men in each country. Compared to age at first union formation, age at first marriage is more heterogeneous across countries and shows more variability within countries the age at first marriage is relatively high. Especially among women, there is an increasing variability as the mean age at first marriage increases. For example in Great Britain, France, the Netherlands and Sweden the range of quartiles is around 6–7 years, whereas among countries with a low age at first marriage the range of quartiles is much narrower. Since we are pooling cohorts over several decades, a wide distribution of mean age at first marriage (or first union formation) may also be due to cohort changes. The association of such change with the gender balance in higher education is central to our research questions. A cohort view of the descriptive statistics is presented later in this section.

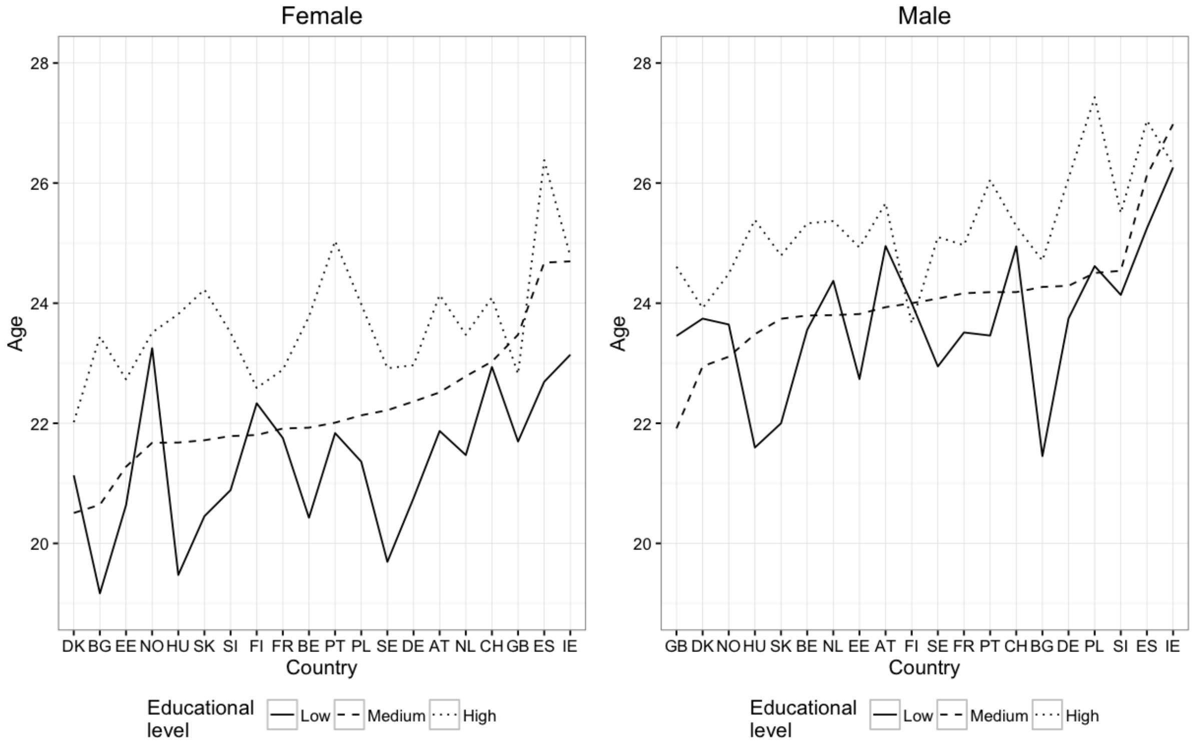
Figure 6 Boxplot of age at first marriage, women and men at who are at least 40 years old



Source: ESS3-2006. Countries ordered by women's mean age at first marriage.

Next, we examine how age at union formation and marriage depend on education. As previously, we use the age at first union or first marriage of those individuals that have experienced the event before age 40, including only those who are at least 40 years old at the time of the interview. Figure 7 shows that in general, for both men and women, the mean age at first union formation is positively associated with education. There are exceptions to this general gradient, primarily in how the age at union formation among the low educated differs from that of the medium educated. For several countries, the data show that low educated women and men have a higher age at first union formation compared to medium educated women and men. In addition to that, a couple of countries do not exhibit the positive gradient in the age at first union formation. Among Finnish and Irish men, the pattern is mixed and there is hardly any sign that tertiary educated men would have their first union at later ages than medium educated men.

Figure 7 Mean age at first union formation by level of education for women and men born in 1950s – 1975

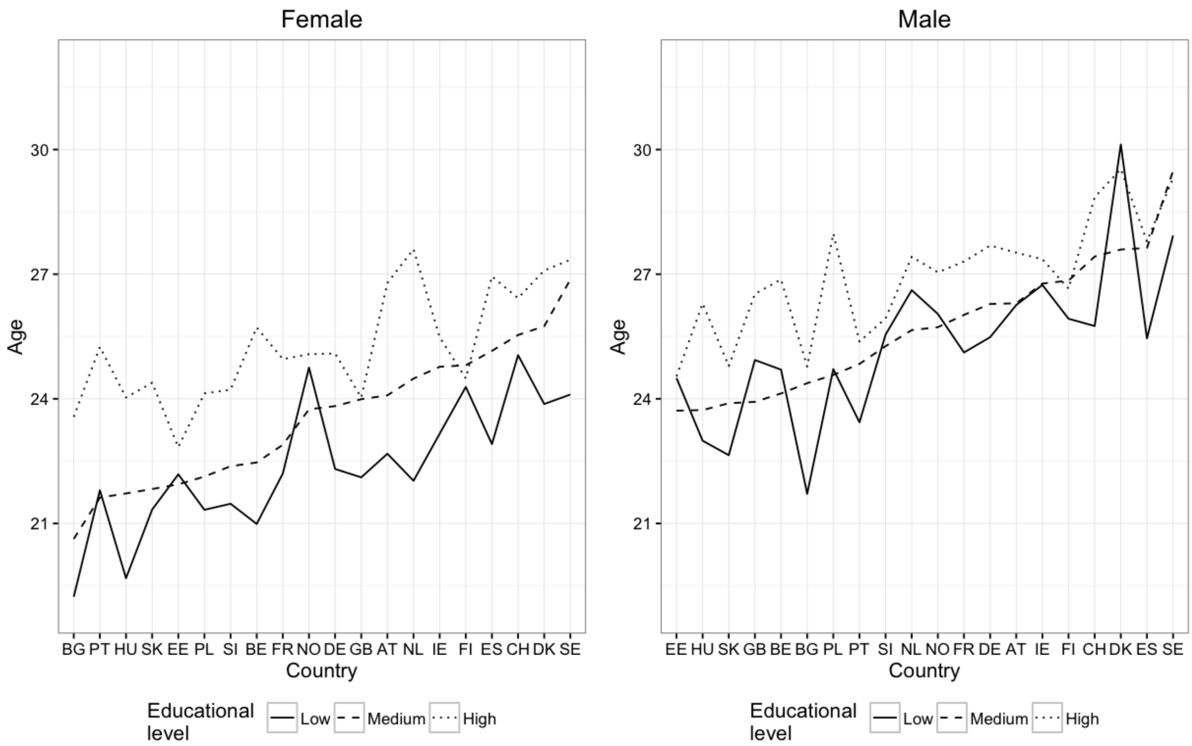


Source: ESS3-2006. Countries ordered by mean age at first union of persons with a medium education

In Figure 7, countries are ordered by the mean age at first union formation of persons with a medium educated group. We notice that the difference between medium and highly educated women tends to be larger in countries where the age at first union formation of the medium educated is relatively low. It is possible that the age at exit from the tertiary education causes the larger gap between the two in the countries with lower age in the medium educated group. In countries where the medium educated have relatively high age at first union, the additional schooling years in the tertiary education have less potential to create the gap between the medium and highly educated.

Figure 8 shows how timing of first marriage differs by educational level. As can be seen on the left side, there is a clear gap between medium and highly educated women's age at first marriage. Only in a few countries there is hardly a difference between the two categories. Age gaps between educational levels tend to be larger for women than for men. The average age gap between low and highly educated brides is 4 years in Bulgaria, 4.3 years in Hungary, 4.9 in Denmark and 3.9 in Portugal. However, as with age at first union formation, there is no educational gradient in age at first marriage for Finnish women and men. Among men, the same can be said also for Ireland.

Figure 8 Mean age at first marriage by level of education for women and men born in 1950s – 1975



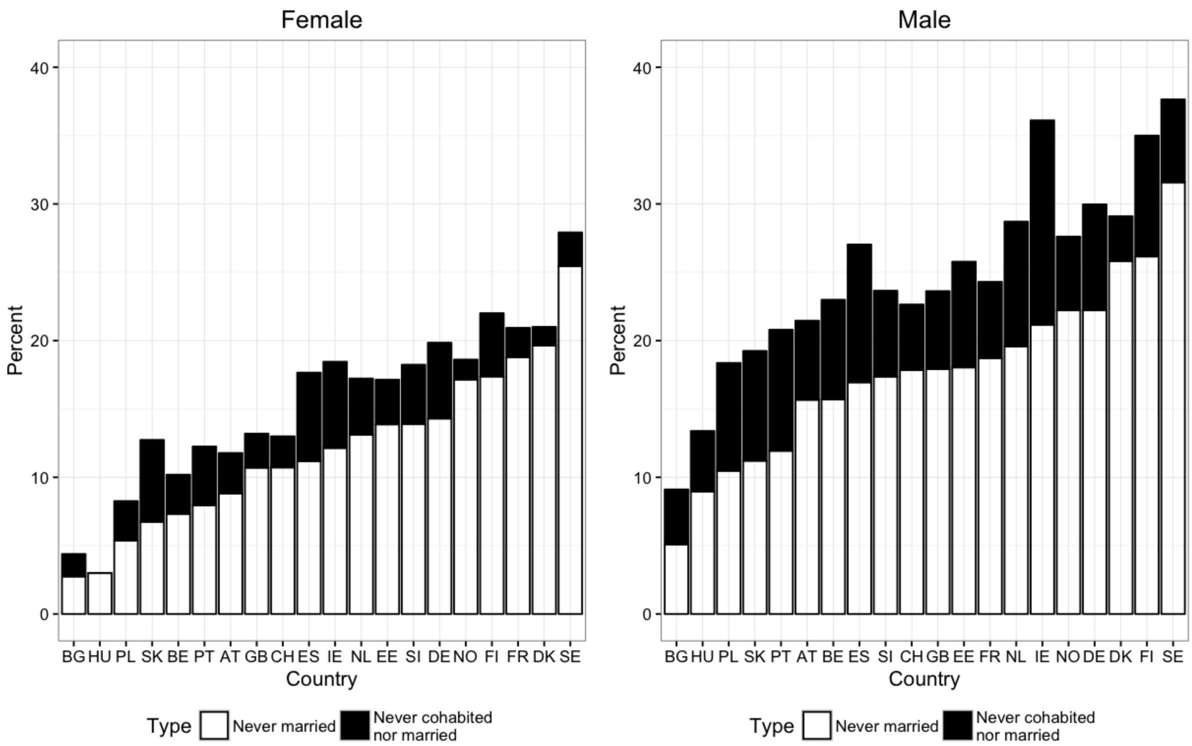
Source: ESS3-2006. Countries ordered by mean age at first marriage for persons with a medium education

4.1.3 Proportions never in a union and never married

In this section we present descriptive results about the outcomes of union formation and marriage processes. Namely, what is the likelihood that a person has been in a union or has been married by a certain age (Table 2 and Figure 9). Among women aged 40 to 57 the proportion of those who report to have never been in a cohabiting union before age 40 ranges from 0% in Hungary to 6.8% in Ireland. In Denmark, Norway and Bulgaria the percentage of women remaining single is below 2%, while in Germany, Spain and Slovakia the percentage of women remaining single is above5%. For men, the proportion of those never in a union before age 40 has much more variety across countries: it ranges from 3.7% in Denmark to 14.4% in Ireland. In Hungary, Bulgaria and Switzerland the percentage of men remaining single is around 4%, while in Spain, Portugal and Slovakia the percentage of men remaining single is above 8%.

Looking at the proportions of never married men and women by age 40, cross-country patterns become more clear due to differences in the spread of unmarried cohabitation between European countries. Among women aged 40 to 57, the proportion of never married individuals ranges from only 2.8% in Bulgaria to 25.4% in Sweden. However, it are not only the Central and Eastern European countries that stand out with high proportion of marriage. Also in Belgium and Portugal the proportion of never married is relatively low with 7.3% and 8% respectively. In Northern Europe, Germany, and France the proportion of women never married by age 40 is at least double of these figures, being around 15%. The same applies to men. The proportion of men never married ranges from 5% in Bulgaria to 31.5% in Sweden. In North Europe, Germany and Ireland over one-fifth of men are not married by age 40, while in Hungary, Slovakia and Poland this number is around 10% and lower.

Figure 9 Percentages of men and women born in 1950s – 1967 and aged 40-57 who never cohabited nor married (never in a union) and never married before age 40



Source: ESS3-2006

Table 2 Weighted counts and percentages of men and women who never formed a union and never married, men and women born in 1950s – 1967, aged 40-57

		Male					Female				
		Never in a union		Never married		N	Never in a union		Never married		N
		N	%	N	%	N	N	%	N	%	N
German speaking	AT	24	5.8	65	15.7	415	16	3.0	48	9.0	532
	CH	13	4.4	52	17.7	294	7	2.2	35	10.8	324
	DE	37	7.7	108	22.5	481	28	5.6	71	14.3	497
West Europe	BE	19	7.3	41	15.6	262	9	2.9	23	7.3	314
	FR	17	5.1	62	18.7	331	8	2.3	66	18.9	350
	NL	26	9.2	56	19.7	284	15	4.2	46	13.0	353
Nord Europe	DK	9	3.7	65	26.4	246	4	1.8	45	20.1	224
	FI	23	8.8	68	26.2	260	12	4.6	45	17.4	259
	SE	15	6.0	78	31.5	248	7	2.5	72	25.4	283
	NO	17	5.7	66	22.2	297	4	1.5	47	17.1	275
South Europe	PT	21	8.7	29	12.0	242	17	4.3	32	8.0	398
	ES	25	10.0	42	16.8	250	18	6.6	30	11.0	273
British Isles	GB	19	5.9	59	18.3	322	9	2.5	39	10.6	367
	IE	35	14.4	50	20.6	243	20	6.8	35	11.9	295
Central and East Europe	PL	19	7.6	27	10.8	251	9	3.5	14	5.4	259
	SI	12	6.3	33	17.3	191	11	4.3	35	13.8	253
	SK	20	8.5	26	11.1	235	17	6.2	21	7.6	275
	EE	15	7.7	35	18.0	194	8	3.3	34	13.9	245
	HU	8	4.3	17	9.2	185	0	0.0	9	3.2	279
	BG	7	3.9	9	5.0	181	5	1.8	8	2.8	282
All countries		381	7.0	988	18.3	5412	224	3.5	755	11.9	6337

Source: ESS3-2006, sampling weights.

Due to the low number of cases who never experienced a cohabitation or marriage we show in Table 3 the educational gradient in the probability of union formation and marriage for all countries pooled. There is a positive educational gradient in the probability of first union formation and first marriage for women and a negative educational gradient in the probability of first union formation and first marriage for men. Highly educated women are the most likely to remain single (3.4%) and low educated women are the least likely to remain single (2.8%). Similarly, the percentage of women aged 40-57 who never married before age 40 is the highest for highly educated women (14.1%) and the lowest for low educated women (8.3%). Among men the educational gradient in marriage is more pronounced than the educational gradient in union formation. Low educated men are nearly twice as likely to remain single (9.7%) as men with a high educational attainment level (5%). Low educated men are also the least likely to marry and medium educated men are the most likely to marry. 17.8% of low educated men aged 40-57 never married and 15.5% of medium educated aged 40-57 men never married.

Table 3 Weighted numbers of men and women born in 1950s – 1967, age 40-57 by level of education and the percentage of those never in a union and never married by level of education for 20 European countries

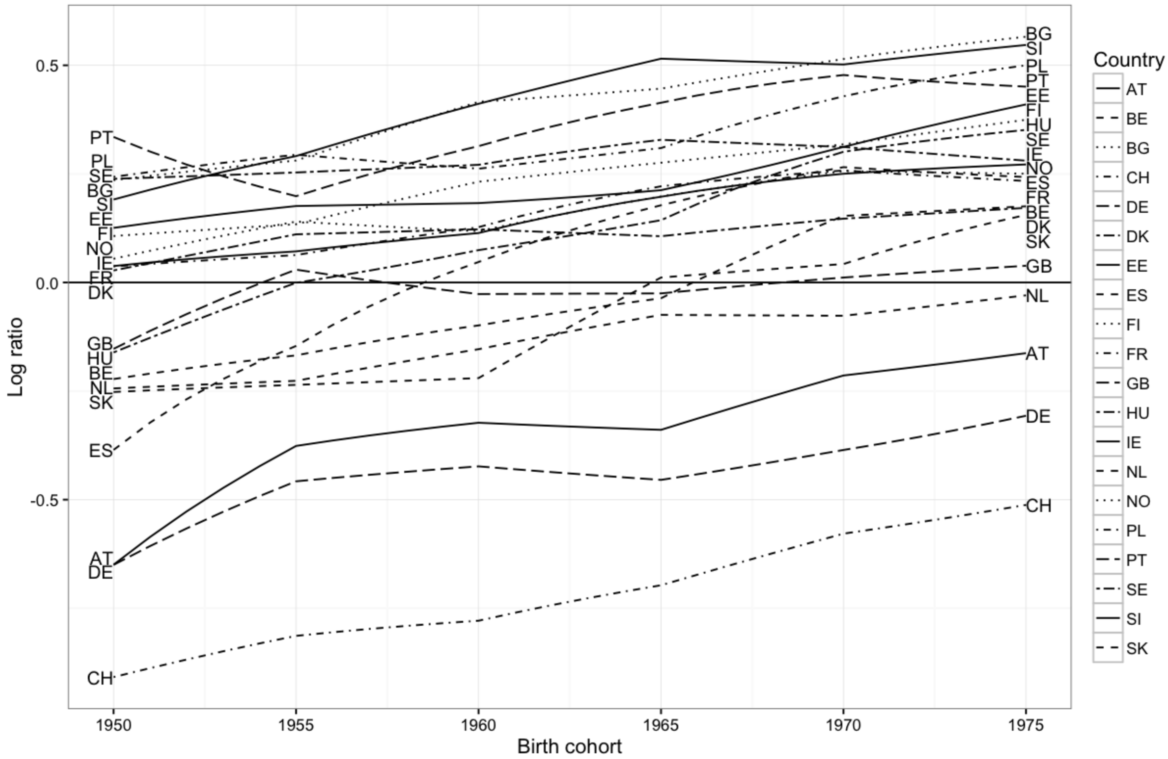
	Men			Women		
	N	% of never in a union	% of never married	N	% of never in a union	% of never married
Low	1222	9.7	17.8	1648	2.8	8.3
Medium	2654	5.3	15.5	2912	3.1	10.3
High	1513	5.0	15.9	1757	3.4	14.1
Total	5399	6.2	16.1	6317	3.1	10.8

Source: ESS3-2006, sampling weights.

4.1.4 Educational attainment

Figure 10 illustrates the changes in the educational composition of the mating market as a result of changes in the relative educational attainment of men and women. The figure plots the country-specific development of the log of the sex ratio for the highly educated for the cohorts born between 1950 and 1975 based on the IIASA/VID data. As indicated in the measurement section, a value above zero means that there are more highly educated women than highly educated men on the mating market, whereas a value below zero means that there are more highly educated men than highly educated women on the mating market. The sex ratio among the highly educated increased in all countries, such that by 2005 the gender imbalance in higher education had turned around in all countries, except for Switzerland, Germany, Austria and the Netherlands. In most eastern and northern European countries a reversed gender imbalance in higher education among 25-to 34-year old women and 27- to 36-year old men was already reached in 1980.

Figure 10 Sex ratio among the highly educated (women to men)



Source: IIASA/VID.

Table 4 presents the gender balance in higher education for the oldest cohort (men and women born in 1950) and the youngest cohort (men and women born in 1975) and the proportion of low, medium and highly educated men and women in the sample. To facilitate the interpretation of the gender balance in higher education, we converted the log of the sex ratio for the highly educated to the percentage of women among the highly educated (% Female). As in Figure 10, we notice that Austria, Germany, Switzerland and the Netherlands did not reach the gender parity in higher education by 2005. For all the Northern European countries, most of the CEE countries, Portugal, Ireland and France gender parity was already reached by 1980. When we look at the percentages of low, medium and highly educated men and women in the sample we observe large differences between countries. In the German speaking countries the bulk of men and women are medium educated. Also in CEE and West Europe the medium educated form the largest educational group. In North Europe and the British Isles the highly educated are the largest educational group, especially among women. Whereas in Southern Europe the low educated are the largest educational group.

Table 4 Percentages of women among the highly educated (% Female) in 1980 and 2005 and percentages of low, medium and highly educated men and women born in 1950s-1975

	% Female		Male			Female		
	1980	2005	Low	Medium	High	Low	Medium	High
AT	34.3	45.9	14.7	73.4	11.9	13.3	80.2	6.5
CH	28.7	37.5	15.2	47.9	36.9	19.0	57.1	23.8
DE	34.3	42.4	3.6	62.4	34.0	11.1	71.4	17.4
BE	44.5	54.4	19.9	44.4	35.8	22.8	37.6	39.6
FR	50.9	55.8	18.1	52.1	29.8	23.0	46.5	30.5
NL	43.9	49.3	27.9	40.0	32.0	34.8	37.1	28.1
DK	50.7	54.3	14.0	35.1	50.9	11.0	33.4	55.5
FI	52.7	59.3	19.2	45.4	35.4	10.4	37.9	51.8
SE	55.9	57.0	16.7	52.8	30.5	14.7	43.5	41.8
NO	51.4	56.2	9.5	49.5	40.9	9.9	38.0	52.1
PT	58.3	61.1	69.3	19.6	11.0	71.7	12.9	15.4
ES	40.5	56.0	41.4	33.7	24.9	46.2	28.8	25.1
GB	46.2	51.0	37.8	14.1	48.1	38.9	10.9	50.2
IE	51.0	56.8	32.0	23.6	44.4	28.6	21.5	49.9
PL	56.1	62.2	14.2	74.6	11.2	16.1	68.0	15.9
SI	54.8	63.3	16.2	63.5	20.3	21.0	48.7	30.3
SK	43.7	53.9	6.1	78.7	15.2	13.1	74.1	12.8
EE	53.1	60.1	12.3	54.7	33.0	6.0	49.4	44.6
HU	46.0	58.7	17.0	67.1	15.9	27.6	54.8	17.6
BG	55.9	63.8	22.9	59.6	17.6	19.7	49.9	30.4
All countries			15.2	47.9	36.9	19.0	57.1	23.8

Source: IIASA/VID and ESS3-2006, sampling weights.

4.2 Event history modelling results

4.2.1 First union formation and first marriage rates

We fitted semi-parametric transition rate models to estimate the effect of macro-level sex ratios of highly educated individuals on the transition to first union and first marriage. Table 5 presents the hazard ratios of first union formation for women (left side) and men (right side). Model 1 in Table 5 shows how respondent's education is related to the transition to first union while controlling only for his/her age and school enrolment. Since our hypotheses concern primarily the highly educated population, we chose tertiary education as the reference category of educational attainment. In Model 2 we added the country-level sex ratio within the highly educated population and its interaction with educational attainment. Recall that the sex ratio is defined, unconventionally but conveniently for the purpose of this paper, as the log of the number of highly educated women divided by the number of highly educated men. Also note that including interaction effects in a regression model affects the meaning of the slope coefficients for the interacted variables (Jaccard 2001). As a result, the significance test of the so-called main effect of the interaction term characterizes the influence of the sex ratio when education is set at the reference category, and conversely, the effect of education when the logarithm of the sex ratio is zero. As the last step, in Model 3 we additionally controlled for the birth cohort of the respondent. In all three models we included country dummies and use country clustered robust standard errors.

The results on the left side of Table 5 are about women. The point estimate for the sex ratio suggests that, in line with hypothesis 1, an increase in the gender balance in higher education to the advantage of women is associated with a lower hazard of first union formation for highly educated women (hazard ratio values for this coefficient are below one in all models where the sex ratio was included). However, this relationship is not statistically significant in neither case. The only statistically significant coefficient related to the sex ratio is the interaction with medium education (HR 1.225** in model 3). This indicates that, as the gender balance in education turns towards an advantage for women, union formation rates of medium educated women increase compared to union formation rates of highly educated women – and note that rates of union formation are already higher for medium educated women in a balanced mating market, as indicated by the hazard ratio for medium educated women compared to the highly educated reference category (1.298 in model 3 for women). A similar pattern emerges for low educated women, but the interaction with the sex ratio is statistically not significant. All in all, these results contain only weak indications that the rates of union formation for highly educated women would deteriorate with the reversal of the gender balance in education, but they do clearly indicate that they decrease compared to the rates for women with less education. In sum, Hypothesis 1 is hardly supported.

The results for men are on the right hand side of Table 5. The picture looks different for men compared to women. The estimates for the sex ratio effect, not interacted so referring to the reference category of highly educated men, are similar to the ones for women. Also here, they are not statistically significant. A unit change in the logged sex ratio variable would lower highly educated men's transition rate to first union by about 25–30%. However, in contrast to what we just observed for women, this reduction seems to apply for all three levels of education; among men the difference in the effect of the sex ratio between the highly educated and the rest is much smaller, if existent at all. So, all in all, the results could be seen as in line with Hypothesis 3, implying lower union formation rates for highly educated men as the sex ratio goes up. However, the evidence supporting this is very thin, as the effect is not statistically significant and should only hold for highly educated men, according to the rationale behind Hypothesis 3.

Furthermore, Table 5 indicates a negative educational gradient of first union formation in a balanced market for women, but not for men. Highly educated women have the lowest hazard of first union formation and low educated women have the highest hazard. Among men, it is the medium educated group who have the highest rates of first union formation, producing an inverted U-shape pattern. Low and highly educated men have statistically significant lower rates of union formation compared to medium educated men.

Control variables in Table 5 adjust for differences due to the age cohort of the respondent and school enrolment. Age cohort of the respondent is positively associated with the transition rate to first union in the baseline model. That is, the model would predict a higher transition rate for older respondents, i.e. those from earlier birth cohorts. After we add the interaction effects between the sex ratio and education, the age coefficient loses its statistical significance, which may be due to the fact that age and the sex ratio are correlated. The school enrolment variable suggest a consistent difference between individuals who are out of schooling and those who are still enrolled in education. Being “in education” has a similar influence on both women and men; it decreases the hazard of first union formation. As of the differences between countries, these are illustrated by the respective indicator variables. For example, in the final model we can observe relatively higher transition rates for both sexes for Denmark, Estonia, and Sweden (compared to the baseline which is Austria). On the other hand, when only women are considered, Spain and Ireland show relatively lower hazard of first union compared to the reference country.

Table 5 Hazard ratios and standard errors from Cox regression models of first union formation for women and men born in 1950s–1975 (aged 31–57)

	Women			**Men**		
	(1)	(2)	(3)	(1)	(2)	(3)
Education (ref. = High)						
Low	1.519***	1.498***	1.498***	1.059	1.055	1.054
	(0.050)	(0.048)	(0.048)	(0.039)	(0.039)	(0.039)
Medium	1.317***	1.300***	1.298***	1.172***	1.173***	1.172***
	(0.032)	(0.031)	(0.031)	(0.031)	(0.031)	(0.031)
Sex ratio		0.768	0.756		0.715	0.699
		(0.168)	(0.166)		(0.167)	(0.164)
Sex ratio*Low		1.217	1.225		1.061	1.064
		(0.132)	(0.133)		(0.135)	(0.136)
Sex ratio*Medium		1.222**	1.225**		0.947	0.949
		(0.091)	(0.092)		(0.073)	(0.073)
Control variables						
Age cohort	1.004**	1.003	0.995	1.008***	1.004	1.005
	(0.002)	(0.003)	(0.007)	(0.002)	(0.003)	(0.008)
In education	0.789***	0.789***	0.791***	0.707**	0.707**	0.706**
	(0.055)	(0.055)	(0.055)	(0.077)	(0.077)	(0.077)
Cohort (ref. = 1950–1955)						
1955–1959			0.978			1.072
			(0.049)			(0.057)
1960–1964			0.962			1.025
			(0.077)			(0.088)
1965–1969			0.895			1.023
			(0.100)			(0.124)
1970–1975			0.843			1.038
			(0.123)			(0.163)

Source: ESS3-2006 and IIASA/VID, sampling weights.
Clustered robust standard errors
**p*<0.05; ***p*<0.01 ; ****p*<0.001

(continued on next page)

Table 5 Hazard ratios and standard errors from Cox regression models of first union formation for women and men born in 1950s–1975 (aged 31–57), continued

	Women			Men		
	(1)	(2)	(3)	(1)	(2)	(3)
Country (ref. = AT)						
BE	1.123*	1.142	1.146	1.072	1.190	1.201
	(0.065)	(0.097)	(0.098)	(0.075)	(0.115)	(0.116)
BG	1.458***	1.577**	1.585**	1.055	1.373	1.399
	(0.104)	(0.266)	(0.269)	(0.091)	(0.254)	(0.260)
CH	0.965	0.910	0.903	0.999	0.871	0.863
	(0.052)	(0.091)	(0.091)	(0.064)	(0.097)	(0.096)
DE	0.977	0.957	0.952	0.859*	0.819**	0.821**
	(0.056)	(0.061)	(0.061)	(0.053)	(0.056)	(0.056)
DK	1.542***	1.602***	1.612***	1.316***	1.543***	1.559***
	(0.106)	(0.184)	(0.186)	(0.096)	(0.193)	(0.196)
EE	1.346***	1.422*	1.438**	1.114	1.361*	1.383*
	(0.092)	(0.195)	(0.198)	(0.085)	(0.201)	(0.205)
ES	0.681***	0.703***	0.706***	0.711***	0.820	0.829
	(0.039)	(0.072)	(0.073)	(0.046)	(0.091)	(0.092)
FI	1.158*	1.226	1.232	1.104	1.351*	1.373*
	(0.080)	(0.168)	(0.170)	(0.080)	(0.197)	(0.201)
FR	1.147*	1.196	1.202	1.085	1.294*	1.308*
	(0.070)	(0.144)	(0.145)	(0.073)	(0.167)	(0.169)
GB	1.090	1.107	1.107	1.136	1.271*	1.283*
	(0.064)	(0.095)	(0.096)	(0.080)	(0.126)	(0.127)
HU	1.478***	1.529***	1.542***	1.068	1.255	1.271
	(0.105)	(0.172)	(0.174)	(0.091)	(0.162)	(0.164)
IE	0.744***	0.778*	0.784*	0.694***	0.824	0.834
	(0.043)	(0.092)	(0.093)	(0.048)	(0.107)	(0.108)
NL	1.022	1.028	1.029	0.949	1.022	1.029
	(0.060)	(0.074)	(0.074)	(0.066)	(0.083)	(0.083)
NO	1.196**	1.253	1.259	1.195**	1.427**	1.446**
	(0.072)	(0.150)	(0.152)	(0.079)	(0.185)	(0.188)
PL	1.024	1.085	1.093	0.885	1.124	1.141
	(0.065)	(0.165)	(0.168)	(0.061)	(0.184)	(0.187)
PT	0.883	0.939	0.951	0.900	1.136	1.157
	(0.057)	(0.151)	(0.154)	(0.070)	(0.196)	(0.201)
SE	1.280***	1.361*	1.366*	1.124	1.399*	1.426*
	(0.084)	(0.196)	(0.198)	(0.078)	(0.212)	(0.217)
SI	1.075	1.158	1.167	0.902	1.179	1.201
	(0.074)	(0.196)	(0.199)	(0.067)	(0.213)	(0.218)
SK	1.112	1.128	1.133	0.973	1.064	1.075
	(0.076)	(0.096)	(0.097)	(0.069)	(0.096)	(0.098)
Observations	8782			7688		

Source: ESS3-2006 and IIASA/VID, sampling weights.
Clustered robust standard errors
**p*<0.05; ***p*<0.01; ****p*<0.001

We now turn to a sub-category of first union formation and focus on first marriage. This process is analysed in the same way as first union formation. The results are presented in Table 6. However, we should bear in mind that the accordance between union formation and marriage differs by country and cohort, depending on how widespread unmarried cohabitation is. Considering the birth cohorts analysed, we have higher levels of non-marital cohabitation (and lower rates of marriage, respectively) in Nord and West European countries and high rates of marriages in East European countries, as shown in the descriptive part. This must be kept in mind, as high rates and early age of marriage coincided with a gender balance in higher education that was already in favour of women in the 1950s birth cohorts. In Table 6, models for women show a higher relative risk of marriage associated with an increase in the logged sex ratio of the highly educated. For highly educated women the transition rate to marriage goes up very strongly as the sex ratio increases (the hazard ratio is 1.995** in model 2 and 1.939** in model 3). This finding is strongly at odds with Hypothesis 1, according to which women's union formation rates, including marriage, should go down as women become a larger majority among the highly educated.

Although not shown in the table, the statistical significance of these ratios disappears when we do not control for the age of the respondent, which again highlights the fact that the age cohort and the sex ratio variables are strongly correlated. Apart from that, we also note that there seems to be a positive correlation between our sex ratio variable and marriage rates across countries. This suggests that, in countries where marriage rates are relatively high, they are particularly high among the highly educated (compared to highly educated in other countries), and that such is particularly the case in countries where women have a high advantage in education. Again, this is clearly at odds with the marriage variant of Hypothesis 1.

The models for men, on the right hand side of Table 6, indicate that an increase in the logged sex ratio of the highly educated population reduces the rate of transition to marriage for the low educated compared to the highly educated (hazard ratio figures 0.737* and 0.736*). For the highly educated, if anything, there is an increase in the marriage rate as the sex ratio increases (which would be in line with Hypothesis 2), but the coefficients do not reach the level of statistical significance. All in all, we find hardly support for Hypothesis 2 and no support for Hypothesis 3. The most important finding here is that marriages rates of low educated men decrease as women gain an advantage in higher education.

Regarding the other variables in the models in Table 6, level of education yields relatively similar results to first marriage as to first union formation. For women there is a clear and consistent negative gradient. The low educated show hazard rates that are increased by about half compared to highly educated women. The difference in rates between highly educated and medium educated women is about a third of the level of the former. On the right side of the table, for men, only a medium level of education shows a statistically significant difference from the tertiary educational level, producing a similar pattern that was observed for first union formation.

Among the control variables, age remains a significant predictor of marriage even after including country-level sex ratio of the highly educated and 5-year birth cohort. This may be an indication that in our data, marriage rates are more sensitive to the birth cohort of the respondent, as marriage was more common in older cohorts.

Table 6 Hazard ratios and standard errors from Cox regression models of first marriage for women and men born in 1950s–1975 (aged 31–57)

	Women			**Men**		
	(1)	(2)	(3)	(1)	(2)	(3)
Education (ref. = Highly educated)						
Low educated	1.515***	1.523***	1.525***	1.012	1.023	1.019
	(0.055)	(0.055)	(0.055)	(0.041)	(0.041)	(0.041)
Medium educated	1.316***	1.323***	1.319***	1.103**	1.101**	1.099**
	(0.037)	(0.037)	(0.037)	(0.034)	(0.034)	(0.034)
Sex ratio*Education						
Sex ratio		1.995**	1.939**		1.178	1.122
		(0.492)	(0.484)		(0.314)	(0.303)
Sex ratio*Low		0.840	0.847		0.737*	0.736*
		(0.103)	(0.103)		(0.105)	(0.105)
Sex ratio*Medium		0.986	0.989		0.919	0.919
		(0.084)	(0.085)		(0.078)	(0.079)
Control variables						
Age cohort	1.030***	1.038***	1.034***	1.038***	1.039***	1.036***
	(0.002)	(0.003)	(0.008)	(0.002)	(0.003)	(0.010)
In education	0.894	0.892	0.892	0.718**	0.710**	0.714**
	(0.065)	(0.065)	(0.064)	(0.091)	(0.090)	(0.092)
Cohort (ref. = 1950-1954)						
1955-1959			1.081			1.092
			(0.061)			(0.066)
1960-1964			1.082			1.070
			(0.097)			(0.105)
1965-1969			0.989			1.033
			(0.125)			(0.144)
1970-1975			0.922			0.913
			(0.153)			(0.165)

Source: ESS3-2006 and IIASA/VID, sampling weights.
Clustered robust standard errors.
**p*<0.05; ***p*<0.01

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Table 6 Hazard ratios and standard errors from Cox regression models of first marriage for women and men born in 1950s–1975 (aged 31–57), continued

	Women			Men		
	(1)	(2)	(3)	(1)	(2)	(3)
Country (ref. = AT)						
BE	1.243***	1.034	1.051	1.121	2.289***	1.338**
	(0.081)	(0.098)	(0.100)	(0.089)	(0.207)	(0.141)
BG	1.972***	1.229	1.266	1.684***	9.093***	2.197***
	(0.157)	(0.232)	(0.242)	(0.163)	(1.294)	(0.452)
CH	0.862**	1.115	1.110	0.945	0.393***	0.849
	(0.048)	(0.124)	(0.124)	(0.063)	(0.033)	(0.101)
DE	0.918	0.999	0.998	0.833**	0.662***	0.837*
	(0.055)	(0.068)	(0.068)	(0.057)	(0.046)	(0.062)
DK	0.735***	0.551***	0.562***	0.721***	2.088***	0.874
	(0.050)	(0.069)	(0.071)	(0.052)	(0.203)	(0.117)
EE	1.192*	0.825	0.850	1.191	4.521***	1.546**
	(0.096)	(0.129)	(0.135)	(0.111)	(0.560)	(0.257)
ES	0.887	0.691**	0.703**	0.896	2.290***	1.140
	(0.056)	(0.078)	(0.081)	(0.065)	(0.218)	(0.136)
FI	0.806**	0.560***	0.574***	0.793**	3.016***	1.007
	(0.056)	(0.084)	(0.087)	(0.060)	(0.339)	(0.161)
FR	0.802**	0.587***	0.599***	0.890	2.874***	1.127
	(0.056)	(0.079)	(0.082)	(0.066)	(0.301)	(0.160)
GB	1.059	0.874	0.883	1.017	2.222***	1.240*
	(0.068)	(0.083)	(0.085)	(0.079)	(0.203)	(0.134)
HU	1.807***	1.376*	1.414**	1.360**	3.960***	1.719***
	(0.147)	(0.174)	(0.181)	(0.129)	(0.453)	(0.245)
IE	0.873*	0.640***	0.654**	0.831*	2.687***	1.062
	(0.059)	(0.085)	(0.088)	(0.065)	(0.287)	(0.153)
NL	0.826**	0.730***	0.739***	0.858*	1.423***	0.988
	(0.052)	(0.056)	(0.057)	(0.066)	(0.121)	(0.091)
NO	0.822**	0.597***	0.608***	0.825**	2.742***	1.054
	(0.055)	(0.080)	(0.083)	(0.061)	(0.286)	(0.151)
PL	1.438***	0.937	0.967	1.441***	6.814***	1.877***
	(0.104)	(0.162)	(0.169)	(0.113)	(0.833)	(0.341)
PT	1.129	0.740	0.766	1.263**	6.916***	1.910***
	(0.081)	(0.134)	(0.140)	(0.114)	(0.930)	(0.362)
SE	0.567***	0.380***	0.388***	0.556***	2.358***	0.727
	(0.038)	(0.061)	(0.064)	(0.042)	(0.272)	(0.124)
SI	1.009	0.631*	0.650*	0.919	5.349***	1.277
	(0.081)	(0.120)	(0.125)	(0.083)	(0.722)	(0.257)
SK	1.476***	1.270**	1.298**	1.504***	2.774***	1.786***
	(0.108)	(0.116)	(0.120)	(0.122)	(0.254)	(0.184)
Observations	8756			7664		

Source: ESS3-2006 and IIASA/VID, sampling weights.
Clustered robust standard errors.
**p*<0.05; ***p*<0.01

4.2.2 Probability of first union formation and first marriage

Rates of union formation and entry into marriage consist of two components: the probability of ever making the transition and the timing of the event. In this section we focus on the probability that the first union or marriage takes place at all before age 40. The results of binary logistic regressions of union formation are shown in Table 7.

Both for women and men, the estimates for the effect of the sex ratio as well as for the interaction with educational level remain below the level of statistical significance. The results do suggest that an increase in the sex ratio may lower the probability of union formation for highly educated women and men, thus supporting the sociocultural theory (Guttentag and Secord 1983) or H1b and H3b, but the standard errors are too large compared to the point estimates to make any reliable claims about this.

If there is a negative educational gradient in the likelihood of union formation for women, is does not appear as statistically significant in our data. Among men, however, it is the low educated group who clearly exhibit the lowest likelihood of union formation, and this differences appears statistically significant.

The control variables age and 5-year birth cohort (the last birth cohort has been dropped due to age limitations of the subsample) produce no statistically significant results. When individuals are still enrolled in education at the time of interview, their likelihood to have experienced a co-residential union is lower, especially for men. Country coefficients for Hungarian women are missing due to lack of observations for those never been in a union.

Table 7 Logistic regression of first union formation, women and men born in 1950s–1967 (age 40-57), odds ratios and standard errors

	Women		Men	
	(1)	**(2)**	**(1)**	**(2)**
Education (ref. High)				
Low	1.319	1.291	0.502***	0.507***
	(0.287)	(0.281)	(0.092)	(0.092)
Medium	1.267	1.250	0.920	0.945
	(0.193)	(0.177)	(0.174)	(0.186)
Sex ratio		0.180		0.573
		(0.186)		(1.080)
Sex ratio*Low		2.793		1.820
		(2.102)		(1.102)
Sex ratio*Medium		2.200		1.903
		(0.989)		(1.092)
Age cohort	0.975	0.961	1.078	1.077
	(0.044)	(0.047)	(0.046)	(0.053)
In education	0.441**	0.450**	0.184***	0.185***
	(0.110)	(0.113)	(0.060)	(0.060)
Cohort (ref. 1950-1954)				
1955-1959	0.974	0.982	1.383	1.385
	(0.284)	(0.284)	(0.358)	(0.359)
1960-1964	0.611	0.613	1.662	1.664
	(0.364)	(0.369)	(0.679)	(0.685)
1965-1969	0.654	0.652	2.144	2.152
	(0.499)	(0.506)	(1.189)	(1.201)
Observations	5772		5193	5193

Source: ESS3-2006 and IIASA/VID, sampling weights.
Clustered robust standard errors.
**p*<0.05; ***p*<0.01; ****p*<0.001

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Table 7 Logistic regression of first union formation, women and men born in 1950s–1967 (age 40-57), odds ratios and standard errors, continued

	Women		Men	
	(1)	**(2)**	**(1)**	**(2)**
Country (ref. = AT)				
BE	1.088	1.357	0.763***	0.754
	(0.101)	(0.394)	(0.045)	(0.367)
BG	1.794***	3.893	1.446***	1.454
	(0.186)	(2.775)	(0.085)	(1.978)
CH	1.316***	0.830	1.246***	1.148
	(0.0942)	(0.299)	(0.071)	(0.877)
DE	0.513***	0.444***	0.629***	0.592*
	(0.031)	(0.052)	(0.039)	(0.125)
DK	2.441***	3.966**	1.762***	1.810
	(0.261)	(1.900)	(0.150)	(1.600)
EE	0.992	1.808	0.685***	0.696
	(0.090)	(0.991)	(0.029)	(0.714)
ES	0.435***	0.617	0.633***	0.633
	(0.046)	(0.249)	(0.048)	(0.431)
FI	0.679***	1.246	0.635***	0.650
	(0.076)	(0.700)	(0.041)	(0.689)
FR	1.412***	2.332	1.037	1.049
	(0.107)	(1.154)	(0.051)	(0.975)
GB	1.288*	1.741	1.142	1.142
	(0.166)	(0.663)	(0.111)	(0.744)
HU			1.314***	1.304
			(0.048)	(0.987)
IE	0.489***	0.810	0.385***	0.392
	(0.050)	(0.403)	(0.027)	(0.369)
NL	0.756**	0.900	0.640***	0.628
	(0.066)	(0.228)	(0.040)	(0.246)
NO	2.275***	3.850**	1.053	1.079
	(0.231)	(1.930)	(0.071)	(1.018)
PL	0.836*	1.590	0.714***	0.707
	(0.065)	(1.021)	(0.029)	(0.850)
PT	0.647**	1.240	0.871	0.872
	(0.098)	(0.886)	(0.082)	(1.165)
SE	1.323**	2.659	0.943	0.962
	(0.121)	(1.681)	(0.055)	(1.165)
SI	0.733**	1.628	1.019	1.027
	(0.071)	(1.182)	(0.044)	(1.432)
SK	0.436***	0.533**	0.589***	0.585
	(0.022)	(0.121)	(0.023)	(0.226)
Observations	5772		5193	

Source: ESS3-2006 and IIASA/VID, sampling weights.
Clustered robust standard errors.
**p*<0.05; ***p*<0.01; ****p*<0.001

The results for the likelihood of ever been married before age 40 are shown in Table 8. A major difference compared to union formation is found in the effect of the logged sex ratio for the reference category of highly educated. An increase in the sex ratio among the highly educated results in much higher odds of ever been married before age 40 for highly educated women. This goes against our hypothesis about the levels of ever married and sex ratio (H1b).

Perhaps the explanation of this lies in the large country differences in how the marriage rates evolved over the cohorts. In Eastern Europe, where a reversal of the gender imbalance in higher education was already observed for the 1950 cohorts, there are high rates of marriage and low numbers of non-marital cohabitation. Very high odds-ratios of marriage in some CEE countries, such as a Bulgaria, Hungary, Poland, and Slovakia in Table 8 confirm the tendency towards a universality of marriage. Highly educated women in these countries are much more likely to marry than to cohabiting. In Western countries, on the other hand, non-marital cohabitation was more likely to be chosen by highly educated women, thus resulting in lower proportions of ever married college educated women. It may be that our sex ratio variable is picking up cohort- and country specific trends that are not accounted for by our cohort and country control variables.

The gradient by education follows more closely that of union formation: low educated women have the highest probability to ever marry and low educated men have the lowest probability to ever marry.

As regards the control variables in Table 8 we see that only for men there are statistically significant results for the age and the schooling variable. The age variable suggests higher proportions of ever married among older respondents. The schooling variable suggests that men who are enrolled in education at the time of interview (thus at later ages) have a lower likelihood of ever been married.

Table 8 Logistic regression of marriage, women and men born in 1950s–1967 (age 40-57), odds ratios and standard errors

	Women		Men	
	(1)	(2)	(1)	(2)
Education (ref. High)				
Low	1.426*	1.460*	0.730*	0.727**
	(0.236)	(0.221)	(0.091)	(0.087)
Medium	1.318*	1.341**	0.960	0.944
	(0.173)	(0.147)	(0.089)	(0.091)
Sex ratio		3.816*		0.316
		(2.574)		(0.447)
Sex ratio*Low		0.479		0.609
		(0.186)		(0.232)
Sex ratio*Medium		0.464**		0.732
		(0.118)		(0.199)
Age cohort	1.047	1.058	1.080***	1.062**
	(0.038)	(0.037)	(0.025)	(0.024)
In education	0.826	0.818	0.397**	0.391**
	(0.144)	(0.142)	(0.120)	(0.119)
Cohort (ref. 1950-1954)				
1955-1959	1.132	1.122	1.197	1.200
	(0.262)	(0.264)	(0.204)	(0.206)
1960-1964	1.019	1.015	1.227	1.213
	(0.446)	(0.444)	(0.324)	(0.315)
1965-1969	0.941	0.938	1.234	1.220
	(0.503)	(0.495)	(0.465)	(0.453)
Observations	6001	6001	5190	5190

Source: ESS3-2006 and IIASA/VID, sampling weights.
Clustered robust standard errors.
**p*<0.05; ***p*<0.01; ****p*<0.001

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Table 8 Logistic regression of marriage, women and men born in 1950s–1967 (age 40-57), odds ratios and standard errors, continued

	Women		Men	
	(1)	**(2)**	**(1)**	**(2)**
Country (ref. = AT)				
BE	1.254***	1.097	1.003	1.460
	(0.069)	(0.205)	(0.026)	(0.538)
BG	3.276***	1.920	3.478***	10.13*
	(0.185)	(0.934)	(0.101)	(10.720)
CH	0.801***	1.188	0.850***	0.492
	(0.031)	(0.319)	(0.025)	(0.288)
DE	0.573***	0.646***	0.620***	0.552***
	(0.019)	(0.053)	(0.023)	(0.084)
DK	0.418***	0.301***	0.528***	1.018
	(0.029)	(0.097)	(0.022)	(0.676)
EE	0.626***	0.416*	0.819***	1.798
	(0.041)	(0.155)	(0.018)	(1.404)
ES	0.743***	0.590*	0.983	1.684
	(0.034)	(0.147)	(0.031)	(0.936)
FI	0.483***	0.317**	0.506***	1.135
	(0.033)	(0.123)	(0.015)	(0.918)
FR	0.411***	0.293***	0.800***	1.628
	(0.017)	(0.098)	(0.017)	(1.153)
GB	0.863*	0.713	0.911*	1.465
	(0.063)	(0.167)	(0.037)	(0.713)
HU	2.926***	2.244**	1.828***	3.327*
	(0.118)	(0.623)	(0.044)	(1.971)
IE	0.751***	0.529	0.716***	1.446
	(0.047)	(0.175)	(0.022)	(1.023)
NL	0.647***	0.587***	0.773***	1.040
	(0.026)	(0.088)	(0.020)	(0.306)
NO	0.496***	0.346**	0.651***	1.326
	(0.032)	(0.118)	(0.023)	(0.950)
PL	1.577***	1.026	1.574***	4.050
	(0.064)	(0.449)	(0.040)	(3.766)
PT	0.994	0.622	1.558***	4.373
	(0.063)	(0.284)	(0.080)	(4.485)
SE	0.294***	0.182***	0.385***	0.977
	(0.016)	(0.078)	(0.011)	(0.904)
SI	0.592***	0.343*	0.912***	2.698
	(0.027)	(0.168)	(0.023)	(2.905)
SK	1.256***	1.106	1.380***	1.910*
	(0.029)	(0.166)	(0.032)	(0.599)
Observations	6001	6001	5190	5190

Source: ESS3-2006 and IIASA/VID, sampling weights.
Clustered robust standard errors.
**p*<0.05; ***p*<0.01; ****p*<0.001

4.2.3 Timing of first union and first marriage

In this section we focus on the timing component of first union formation and first marriage. Only individuals who experienced union formation or marriage before age 40 are considered in this part, i.e. we get rid of the censoring that was present in the event history models. The time to first union or first marriage is modelled using simple linear regression. The results for union formation are shown in Table 9 and the coefficients in the table can be interpreted as change in years in the age at first union formation.

Both for women and men, none of the coefficients of the terms related to the sex ratio are statistically significant because the standard errors are very large. Nonetheless, the direction of the coefficients suggests that a unit increase in the logged sex ratio would be associated with about half a year postponement of first union formation for highly educated women, while the medium and low educated would not be similarly affected. As opposed to that, for highly educated men a unit increase in the logged sex ratio would mean a slight decrease in the age of first union formation. While the interaction results for women and men may indicate some support for the proposed association between the sex ratio and the timing of unions (H1a and H2a), the great uncertainty of the estimated coefficients implies that we do not find support in our data for any of these hypotheses.

Education shows a consistent positive gradient both for women and men. Low educated women form their first union more than two years earlier than highly educated women. The gap between medium educated women and highly educated women is over one year. Similar association holds for men, although the timing gap between low and highly educated men and low and medium educated men is much smaller. Except the country indicators, none of the control variables in Table 9 show statistically significant differences.

Table 9 Linear regression of timing of first union, women and men born in 1950s–1967 (aged 40-57), unstandardized coefficients and standard errors

	Women		**Men**	
	(1)	**(2)**	**(1)**	**(2)**
Education (ref. High)				
Low	-2.269***	-2.253***	-1.396***	-1.348***
	(0.282)	(0.275)	(0.226)	(0.208)
Medium	-1.291***	-1.275***	-1.168***	-1.131***
	(0.183)	(0.175)	(0.169)	(0.161)
Sex ratio		0.528		-0.293
		(1.798)		(2.381)
Sex ratio*Low		-0.876		-1.083
		(0.553)		(0.623)
Sex ratio*Medium		-0.799		0.708
		(0.460)		(0.436)
Age cohort	0.0233	0.0237	0.043	0.041
	(0.030)	(0.042)	(0.046)	(0.051)
In education	0.247	0.236	-0.016	-0.057
	(0.300)	(0.297)	(0.638)	(0.667)
Cohort (ref. 1950-1954)				
1955-1959	0.181	0.180	-0.053	-0.052
	(0.196)	(0.195)	(0.221)	(0.224)
1960-1964	0.0504	0.0551	0.318	0.310
	(0.382)	(0.380)	(0.505)	(0.503)
1965-1969	0.359	0.370	0.507	0.483
	(0.576)	(0.578)	(0.740)	(0.746)
Constant	22.64***	22.52***	23.13***	23.21***
	(1.673)	(1.785)	(2.469)	(2.434)
Observations	5699	5699	4707	4707

Source: ESS3-2006 and IIASA/VID, sampling weights.
Clustered robust standard errors.
**p*<0.05; ***p*<0.01; ****p*<0.001

(continued on next page)

Table 9 Linear regression of timing of first union, women and men born in 1950s–1967 (aged 40-57), unstandardized coefficients and standard errors, continued

	Women (1)	(2)	Men (1)	(2)
Country (ref. = AT)				
BE	-0.574***	-0.487	-0.227***	-0.285
	(0.073)	(0.447)	(0.055)	(0.597)
BG	-1.588***	-1.482	-0.527***	-0.540
	(0.073)	(1.229)	(0.054)	(1.693)
CH	0.597***	0.596	0.171**	0.0523
	(0.051)	(0.713)	(0.059)	(0.975)
DE	-0.388***	-0.369	0.365***	0.395
	(0.043)	(0.201)	(0.041)	(0.260)
DK	-1.732***	-1.652	-1.141***	-1.145
	(0.100)	(0.828)	(0.078)	(1.089)
EE	-1.208***	-1.136	-0.444***	-0.478
	(0.096)	(0.974)	(0.033)	(1.290)
ES	1.679***	1.766**	1.663***	1.597
	(0.084)	(0.607)	(0.098)	(0.821)
FI	-0.797***	-0.723	-0.592***	-0.593
	(0.098)	(0.974)	(0.064)	(1.307)
FR	-0.530***	-0.434	-0.174**	-0.192
	(0.053)	(0.846)	(0.046)	(1.149)
GB	-0.372**	-0.276	-0.824***	-0.874
	(0.102)	(0.597)	(0.114)	(0.779)
HU	-1.112***	-1.022	-0.778***	-0.819
	(0.066)	(0.678)	(0.053)	(0.939)
IE	1.369***	1.452	1.915***	1.946
	(0.089)	(0.855)	(0.088)	(1.138)
NL	-0.076	-0.004	0.002	-0.077
	(0.073)	(0.364)	(0.0645)	(0.474)
NO	-0.336**	-0.257	-0.847***	-0.869
	(0.093)	(0.869)	(0.058)	(1.173)
PL	-0.335***	-0.193	0.593***	0.525
	(0.051)	(1.108)	(0.048)	(1.526)
PT	0.257*	0.417	-0.276	-0.030
	(0.123)	(1.182)	(0.146)	(1.594)
SE	-0.869***	-0.786	-0.331***	-0.316
	(0.079)	(1.112)	(0.057)	(1.513)
SI	-0.651***	-0.542	0.337***	0.302
	(0.075)	(1.246)	(0.054)	(1.700)
SK	-0.761***	-0.706	-0.564***	-0.576
	(0.034)	(0.353)	(0.024)	(0.482)
Observations				

Source: ESS3-2006 and IIASA/VID, sampling weights.
Clustered robust standard errors.
**p*<0.05; ***p*<0.01; ****p*<0.001

The results of the timing of first marriage are shown in Table 10. Again none of the coefficients of the terms related to the sex ratio are statistically significant because the standard errors are very large. Compared to the timing of first union formation, the direction of the coefficients for first marriage tends to be different for highly educated women. An increase in the sex ratio in favour of women decreases the age at first marriage for highly educated women. This goes against our hypothesis H1a. However, we notice that an increase in the sex ratio in favour of women decreases age at first marriage for all women and men (regardless of their educational level). Again, the explanation to this lies perhaps in the large country differences in the timing of marriage, and how they evolved over the cohorts. In Eastern Europe, where a reversal of the gender imbalance in higher education was already observed for the 1950 cohorts, men and women marry not only more often but also early in life. In Western, Southern and German-speaking countries, the reversal of the gender balance in higher education was much later and age at marriage is much higher. Such long term, traditional characteristics of marriage may be overriding the potential influence of skewed sex ratios among the highly educated population.

The educational gradient in the timing of first marriage is similar to the educational gradient in the timing of first union formation. For marriage, however, the timing gap between highly educated women and low educated women is almost three years. For men, the timing gap between highly educated men and low educated men is less than two years.

Table 10 Linear regression of timing of first marriage, women and men born in 1950s–1967 (aged 40-57), unstandardized coefficients and standard errors

	Women		**Men**	
	(1)	**(2)**	**(1)**	**(2)**
Education (ref. High)				
Low	-2.859***	-2.855***	-1.633***	-1.610***
	(0.350)	(0.349)	(0.228)	(0.219)
Medium	-1.638***	-1.642***	-1.246***	-1.204***
	(0.254)	(0.258)	(0.180)	(0.181)
Sex ratio		-1.134		-3.169
		(2.502)		(2.696)
Sex ratio*Low		0.273		0.591
		(1.044)		(0.755)
Sex ratio*Medium		-0.061		0.687
		(0.656)		(0.389)
Age cohort	-0.065*	-0.079	-0.060	-0.094
	(0.028)	(0.049)	(0.044)	(0.049)
In education	0.618	0.623	-0.712	-0.734
	(0.385)	(0.389)	(0.792)	(0.790)
Cohort (ref. 1950-1954)				
1955-1959	-0.050	-0.045	-0.001	0.006
	(0.138)	(0.138)	(0.240)	(0.238)
1960-1964	-0.099	-0.103	0.194	0.172
	(0.276)	(0.285)	(0.508)	(0.495)
1965-1969	0.153	0.147	0.579	0.549
	(0.447)	(0.461)	(0.633)	(0.637)
Constant	28.81***	29.05***	30.39***	31.05***
	(1.379)	(1.723)	(2.302)	(2.223)
Observations	5131	5131	4049	4049

Source: ESS3-2006 and IIASA/VID, sampling weights.
Clustered robust standard errors.
**p*<0.05; ***p*<0.01; ****p*<0.001

(continued on next page)

Table 10 Linear regression of timing of first marriage, women and men born in 1950s–1967 (aged 40-57), unstandardized coefficients and standard errors, continued

	Women (1)	Women (2)	Men (1)	Men (2)
Country (ref. = AT)				
BE	-1.039***	-0.740	-1.487***	-0.834
	(0.090)	(0.607)	(0.060)	(0.687)
BG	-2.969***	-2.154	-2.573***	-0.689
	(0.089)	(1.703)	(0.047)	(1.964)
CH	1.489***	1.039	0.966***	-0.224
	(0.052)	(0.981)	(0.051)	(1.117)
DE	-0.306***	-0.429	-0.025	-0.365
	(0.046)	(0.264)	(0.036)	(0.287)
DK	1.503***	2.044	1.983***	3.237*
	(0.140)	(1.136)	(0.076)	(1.264)
EE	-2.300***	-1.658	-2.630***	-1.166
	(0.131)	(1.331)	(0.035)	(1.503)
ES	0.599***	1.000	0.211	1.113
	(0.081)	(0.832)	(0.111)	(0.921)
FI	-0.036	0.605	-0.053	1.448
	(0.128)	(1.333)	(0.060)	(1.517)
FR	-0.878***	-0.317	-0.432***	0.865
	(0.059)	(1.168)	(0.049)	(1.329)
GB	-1.103***	-0.715	-1.294***	-0.414
	(0.119)	(0.814)	(0.124)	(0.894)
HU	-2.378***	-1.924	-2.343***	-1.306
	(0.079)	(0.943)	(0.049)	(1.094)
IE	0.148	0.694	0.276**	1.580
	(0.117)	(1.167)	(0.093)	(1.305)
NL	0.252**	0.499	-0.181*	0.325
	(0.072)	(0.485)	(0.066)	(0.545)
NO	-0.155	0.419	-0.510***	0.832
	(0.124)	(1.193)	(0.050)	(1.372)
PL	-1.731***	-0.983	-1.464***	0.220
	(0.055)	(1.540)	(0.041)	(1.782)
PT	-1.030***	-0.309	-2.302***	-0.516
	(0.125)	(1.662)	(0.158)	(1.803)
SE	2.074***	2.805	2.544***	4.266*
	(0.107)	(1.532)	(0.053)	(1.754)
SI	-1.501***	-0.683	-1.019***	0.883
	(0.089)	(1.725)	(0.046)	(1.968)
SK	-1.922***	-1.680**	-2.518***	-1.979**
	(0.048)	(0.491)	(0.031)	(0.572)
Observations				

Source: ESS3-2006 and IIASA/VID, sampling weights.
Clustered robust standard errors.
**p*<0.05; ***p*<0.01; ****p*<0.001

5 Conclusions

In recent decades a reversed gender gap in tertiary education has emerged in the majority of European countries. Women have overtaken men in enrollment and completion of tertiary education. As a result, there are more highly educated women than highly educated men entering today’s mating markets. As people have education-specific partner preferences, we expect that changes in the gender balance in higher education will influence the timing and likelihood of partnership formation. Following Van Bavel (2012) we posit that the reversal of the gender imbalance in higher education results in an education-specific mating squeeze for highly educated women.

We have approached this question using two theoretical frameworks. Marital search theory (Oppenheimer 1988) and sociocultural theory (Guttentag and Secord 1983) offer explanations to how imbalances on the mating market may affect union formation rates. Based on marital search theory, which rests on the simple rules of the supply of potential partners, we expected lower rates of union formation for highly educated women (H1) and higher rates of union formation for highly educated men (H2). In line with the sociocultural theory, which holds that men react differently to sex ratio imbalances, it was expected that a higher number of available partners discourages men from making a commitment to form a union. Accordingly, we hypothesized that the reversal of the gender imbalance in higher education lowers the rates of union formation for highly educated women and highly educated men (H3).

The study made use of the European Social Survey data, coming from the third survey round in 2006. The study sample included respondents from 20 European countries, born between 1950-1975. For some parts of the analysis we had to narrow the cohort range to 1950-1967. The descriptive results confirmed that there is a substantial cross country heterogeneity in the timing of entry into first union at different levels of education. In most countries, there was some postponement of union formation across cohorts. Unlike with first union formation, we saw substantial postponement of first marriage formation over cohorts and changes in the proportions of people who ever got married. The observed differentials by country and cohort are much larger for first marriages than for first union formation. For most countries, the educational gradient of both first union formation and marriage was positive, but with less consistent differences between the low and medium educated. Among the cohorts born in the 1950s, the gender imbalance in higher education had already turned to the advantage of women in 11 countries of the 20 in our sample. For the cohorts born in 1975, the gender imbalance in higher education had turned to the advantage of women in all countries expect Austria, The Netherlands, Switzerland and Germany. We proceeded to test whether these changes at the macro level had the hypothesized association with union formation rates at the micro level.

We analysed the association between the shifting gender balance in higher education and the rates of first union formation by means of survival modelling. To further disentangle the timing and quantum components in the transition rate models, linear and binary logistic regressions were used to model separately the timing and the probability of first union formation. The primary focus of the analysis was on entry into first union formation, either marriage or non-marital cohabitation. In parallel, we selected only marriages and tested our hypotheses on this subset of first unions. It was statistically not feasible to address unmarried cohabitation separately due to low sample sizes.

The modelling results provided us with several insights to the association between the gender balance in higher education and the rates of union formation. However, our data do not lend full support to any of the hypotheses derived from the marriage squeeze perspective. Sometimes, the point estimates were in line with what was hypothesized, but the standard errors were too large to reach statistical significance, suggesting an imprecise estimation of the sex ratio effects. Nevertheless, despite the weak statistical power of the estimated coefficients, the direction of the associations lends some support to marital search theory when looking at the timing of first union formation and to sociocultural theory when looking at the probability of first union formation. In line with hypothesis H1, an increase in the number of highly educated women relative to highly educated men was associated with a higher age at first union formation and with a lower probability of first union formation among highly educated women. For highly educated men, the presence of a relatively high number of highly educated women in the mating market was associated with a lower age of first union formation (H2) but, on the other hand, also with a lower probability of first union formation (H3). Again, these associations are statistically not significant.

Similarly, we did not find support for our hypotheses about the association between the shifting gender balance in higher education and the timing and likelihood of first marriage. In general, changes in the gender balance in higher education in favour of women show a positive effect on the timing and likelihood of entry into first marriage, but only the sex ratio effects for the likelihood of first marriage of highly educated women are significant. Remarkably, the results for highly educated women go in opposite direction as formulated in hypothesis H1. An increase in the number of highly educated women relative to highly educated men was associated with a higher probability to ever marry for highly educated women, and the association was statistically significant. Results for highly educated men are supportive of marital search theory but insignificant. A potential explanation for the positive effects of the sex ratio on marriage could be the positive correlation between the sex ratio and marriage rates across countries. The descriptive results showed high percentages of married and early marriage for most of the Central and East European countries, where the gender balance in higher education was already in favour of women in the 1950s birth cohorts, and lower rates of married and late marriages for Western, Southern and German-speaking countries. While our models include controls for country and cohorts (fixed effects), we did not include their interaction. Our sex ratio measure, in contrast, is specific for combinations of country and cohort. So perhaps our sex ratio measure was picking up cohort- and country-specific trends that may have nothing to do with the sex ratio as such, but were just correlated with them. This remains to be investigated in the future.

The results for the effect of education on the likelihood of first union formation and first marriage corroborates earlier research findings by Dykstra and Poortman (2010) and Wiik and Dommermuth (2014). For men, we found a positive educational gradient of the probability to ever form a union and to ever marry. For women, we found an insignificant negative educational gradient of the probability to ever form a union and a significant negative educational gradient of the probability to ever marry. Furthermore, high educational attainment was positively associated with the age at entry into first union and first marriage for both men and women.

The notion of the education-specific mating squeeze is based on the assumption of a certain rigidity in partner preferences. The negative educational gradient of union and marriage entry for women and the positive educational gradient for men suggest to some extent that preferences have not changed. Nevertheless, we do not find evidence in this paper that highly educated women suffer an education-specific mating squeeze. In an earlier paper, De Hauw, Grow and Van Bavel (2015) observed that as the gender balance in higher education turned to the advantage of women, highly educated women partner more often with less educated men, suggesting that on average, in Europe, highly educated women tend to adjust their union formation behaviour to the demographic reality on the mating market (see also Esteve et al. 2012), and that some modification in mating preferences are in place. Furthermore, as to the timing and likelihood of union formation, we may speculate that mating market conditions set by the shifting gender balance in higher education have a relatively weak influence compared to other processes, such as those characterizing the Second Demographic Transition, for example (Lesthaeghe 2010).

Still, our results suffer from imprecise estimation of sex ratio effects. Future work should further attempt to improve on this. One way to try, if data were available, is to use panel model approaches where country-specific populations are followed over time. Alternatively, exogenous (random) shocks in the distribution by gender and education could be a useful instrument to improve the estimation of the effect of the reversal of the gender balance in education on union formation and marriage.

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