Education and the Transition to Fatherhood: the Role of Selection into Union

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Abstract

While advanced education has been found to be consistently associated with a later transition to parenthood for women, findings about education and the transition to parenthood have been much less consistent for men, and no stylized fact has emerged from the literature. We argue that the inconsistency of findings for men is due to the fact that the selection process involved in union formation has been disregarded in earlier studies. We hypothesize that men's educational attainment consistently and positively affects the transition to fatherhood via higher rates of union formation. We apply multi-process event history analysis to data from the Generations and Gender Surveys for 10 European countries. Our results show indeed a consistent positive effect of education on the transition to fatherhood, but it operates chiefly through selection into union. Failing to account for this selection process, leads to a major underestimation of the salience of education for the transition to fatherhood.

Keywords: fatherhood, union formation, education**,** selection

Introduction

A major fertility trend of the past decades in the West has been the postponement of parenthood. Chief explanations of postponement include the expansion of women's enrolment in advanced education and their increased participation in the labor market. More highly educated women, who are also more likely to be active in the paid labor market, tend to make the transition to parenthood at a later age than their lower educated peers – even if the former often catch up at later ages (Mills et al. 2011; Sobotka 2004; Sweeney 2002).

The role played by men's education in the transition to parenthood has received much less attention. It is important to focus on the role of men's education because major changes have taken place in the relative education of men and women. While men were typically more educated than women in the past, the gender gap in education has turned around in most Western countries. In recent years, the number of highly educated women reaching the reproductive ages is exceeding the number of highly educated men (DiPrete and Buchmann 2006; Van Bavel 2012). This has affected educational assortative mating: while educational homogamy remains dominant, the female partner's education now typically exceeds the male partner's (hypogamy) in case of differential attainment levels, whereas the reverse (hypergamy) has always been true in the past (Esteve et al. 2012; Grow and Van Bavel 2015; Schwartz and Mare 2005). This reversal has potentially far-reaching consequences for family formation (Van Bavel 2012).

Parenthood implies parental investments and resources from both women and men. The decline of the male breadwinner–female homemaker model has turned obsolete the "separate spheres" argument for focusing only on women's characteristics. Women's increased earning potential and activity in the labor market may put pressure on men to be more actively involved in household work and childcare activities (Huinink and Kohli 2014; Martin-Garcia 2009; McDonald 2000; Sweeney 2002). Although time-use studies have shown that men still spend much less time than women on household chores and childcare (Craig and Mullan 2010; 2011), men and women have on average become more equally involved in daily parenting activities compared to earlier generations. In the context of the dual-earner family, education plays an important role in shaping gender relations in family formation processes, as well as gender relations in established households (Carlson et al. 2013; Goldscheider et al. 2014; Martin-Garcia 2009; Van Bavel 2012).

So far, the findings about the effect of men's education on the transition to parenthood have been inconsistent. We argue that this is because earlier studies of men's education and family formation have typically looked at it either in the context of an established couple or without simultaneously considering the education of the female partner. Studies investigating fertility from the couple's perspective fall into the first category (see, e.g., Begall 2013; Corijn et al. 1996; Jalovaara and Miettinen 2013; Thomson 1997; Vignoli et al. 2012). A limitation of such studies is that they suffer from selection bias: since only partnered men are analyzed, these studies lose sight of how some men are selected into unions while others are not. The second type of studies that have looked at the effect of male characteristics do not suffer from this bias, but at the expense of failing to control for the effect of female characteristics. In the literature about the transition to adulthood, scholars have investigated the effect of education both for men and women, but in order to be able to include singles, the effect of the partner's education had to be left out of the equation (Billari and Liefbroer 2007; Corijn and Klijzing 2001). Such study design fails to account for strong educational homogamy and, as a result, one cannot tell to what extent the estimated effect of his education really reflects his rather than her education.

In this paper, we propose to combine the advantages of both approaches, namely to include both single and partnered men while taking into account the education of the female partner, if any. We do this by simultaneously modeling union formation and the transition to parenthood as two interrelated processes. We hypothesize that men's education has a consistent and positive effect on fatherhood rates, but that this acts chiefly through the process of union formation: men with higher educational attainment tend to be more attractive on the mating market and therefore exhibit higher rates of union formation. As a result, they also exhibit higher fatherhood rates. If only men with a partner are considered – after the selection effect has played out its role – the effect of his education on top of the effect of the female partner's education becomes much more uncertain, which is why estimates from couple-level studies may be inconsistent. Within couples, it is typically her rather than his education that affects fertility outcomes.

In order to test the selection-into-union hypothesis for men, we fit a multi-process model to account for the endogeneity of union formation and parenthood. So far, the literature considering the endogenous relations between family events has mostly focused on women's characteristics (see e.g. Baizan et al. 2003; Brien et al. 1999). To date, there is a lack of studies focusing on the link between men's education and the transition to fatherhood, and the relationship between union formation and fatherhood. This paper aims to fill that gap. We replicate our multi-process model in 10 different European countries, using data from the Gender and Generations Surveys (GGS), to explore the sensitivity of our hypothesis to different European contexts. Our findings do indeed show consistently positive effects of education on the transition to fatherhood in diverse European contexts for men, but they operate mainly through union formation.

The following sections discuss theoretical insights about the relationship between men's education and family formation and then review lessons learned from earlier empirical studies. Next, we explain more about our empirical strategy and report on our findings.

Education and Men's Family Formation

Given the multidimensional nature of education, education may affect union formation and the transition to parenthood through several mechanisms (Kravdal and Rindfuss 2008; Lappegard and Ronsen 2005; Tesching 2012; Van Bavel 2010). These may differ between women and men and will play a different role depending on the social context. Our theoretical framework focuses on the transition to fatherhood within the context of cohabiting or married unions, as is the case for the vast majority of childbirths in Europe and the US (Cherlin 2010; Manlove et al. 2010; Perelli-Harris et al. 2012). Our framework may be less adapted to analyze the transition to fatherhood among men who are not co-residing with (the mothers of) their children. Such situation occurs more often at younger ages and is more common among men with low socio-economic status and education, particularly in the UK and the US (Berrington et al. 2005; Rendall et al. 1999; Zhang 2011). It often leads to a situation where the biological fathers are hardly actively fathering their children (Berrington et al. 2005; Eggebeen and Knoester 2001; Goldscheider and Kaufman 1996). For a recent discussion of relevant theoretical perspectives about fatherhood outside unions, we refer to Carlson et al. (2013).

In the literature about the interplay between union formation and parenthood, two types of economic mechanisms are usually distinguished that relate his and her education with childbearing decisions: the positive income effect and the negative price effect. The income effect accounts for the fact that the more educated people tend to earn a higher income and they are therefore more likely to afford the monetary costs of having (additional) children. The price effect, on the other hand, acts through opportunity costs: highly educated people have high opportunity costs because they have more to lose when they have to devote more time to non-paid activities like childcare and household chores after becoming a parent (Becker 1991).

In the male breadwinner model, opportunity costs predominate for women while the income effect is more important for men. In this model, the expected educational gradient in union formation and fertility is negative for women and positive for men, since the opportunity costs of family formation are larger for college educated women while the positive income effect predominates for their male peers. Even before graduating, women have more difficulty to balance the role of wife/mother with that of student, so the negative effect of educational enrolment is expected to be stronger for women than for men (Blossfeld and Huinink, 1991; Liefbroer and Corijn 1999).

To a large extent, the male breadwinner model has now given way to a dual earner model (Sweeney 2002). Since the last decades of the twentieth century and the first decades of the twenty-first, gender inequalities at macro and micro levels have changed. In education, the gender gap has reversed: since about the 1990s, in most of the OECD countries, female enrollment in college level education exceeds male enrollment, and women also complete their education more successfully (Vincent-Lancrin 2008). Female labor market participation has increased dramatically and, to a much lesser extent, men have also become more involved in household chores and childcare (England 2010; Oppenheimer 1994; Raley et al. 2012). Women's earnings in the labor market are increasingly considered an essential part of the family budget (Oppenheimer 1994).

When the dual earner model prevails, the expected relationship between educational attainment and family formation is different. Women with high earning potential increasingly become more attractive partners on the mating market (DiPrete and Buchmann 2006; Oppenheimer 1988; 1994;

Sweeney 2002). For them, the positive income effect may start to prevail, while they may dampen opportunity costs either by outsourcing child care and household chores (Kravdal 2007) or by sharing household work more equally with the male partner (Sullivan et al. 2014). For men, in turn, the opportunity costs may rise, because men are under pressure to increase their engagement in parenting and housework (Huinink and Kohli 2014).

Not only an individual's own education matters for family formation, also the partner's education is relevant. In a large majority of couples, both partners have the same level of education. While this is still the case today, the reversal of the gender gap in education has changed the situation in couples where there is a difference in education: before, the male partner typically had as much as or more education as the female partner; after the reversal, women typically are equally or more educated than their partners (DiPrete and Buchmann 2006; Esteve et al. 2012; Grow and Van Bavel 2015). This is associated with an increased proportion of families where the woman is the main breadwinner, even if a gender pay gap to the disadvantage of women persists (Klesment and Van Bavel 2015). Under these new circumstances, the significance of the educational attainment of the male partner may change. The income of the male partner may become a less crucial selection criterion who themselves have a high income. High earning women may become more interested in the social fathering skills of potential partners. Men with less education may compensate a limited income potential by exhibiting the will and ability to be involved in household chores and child-rearing tasks. In such a way, they may enhance their attractiveness to college educated women who want both a career and a family (Van Bavel 2012).

Summing up the argument so far, to the extent that the dual earner family becomes the norm, the impact of education on family formation will tend to become more similar for men and women, with opportunity costs of parenthood for men and income effects for women becoming more salient. At the same time, the impact of men's own education will also depend on the education of the chosen spouse, and vice versa. All this implies that the effect of education on the transition to parenthood cannot be understood without considering the selection effects involved in union formation.

On top of the selection effects, the interdependencies between union formation and the transition to parenthood complicate the distinction between cause and effect. For men in particular, accounting for the interrelatedness of union formation and parenthood is crucial. Finding a suitable partner is a necessary prerequisite to become a father. After finding a partner, the transition to fatherhood strongly depends on the stability of the union and the characteristics of the spouse. Co-residence typically implies an acceleration of the family formation process. Conversely, looking in the other causal direction, a pregnancy may expedite co-residence (Baizan et al. 2003). The interrelationship between union formation and parenthood is strengthened by the fact that individuals are heterogeneous in factors that may simultaneously affect both kinds of events. Some of such factors are observed, like education, but many are typically unobserved, like personality traits and physical characteristics. A proper empirical analysis should account for these interdependencies.

Earlier Empirical Findings

Studies addressing men's transition to parenthood tend to fall into one of two categories. First, life course research about the transition to adulthood typically looks at men and women separately, investigating variability in the occurrence, order, and timing of events. Second, studies that focus on fertility from a couple's perspective typically look at the influence of male characteristics after controlling for female characteristics.

Studies about the transition to adulthood consistently show that school and college enrolment delay union formation and parenthood, for men as well as for women (Blossfeld and Huinink 1991; Corijn and Klijzing 2001). Enrolment delays parenthood more than union formation, and the effect is found to be weaker for men than for women (Corijn and Klijzing 2001; Liefbroer and Corijn 1999; Winkler-Dworak and Toulemon 2007). The effect of educational attainment is much less clear than the enrolment effect. Some studies found that high attainment accelerates men's union formation and marriage (Corijn and Klijzing 2001; Goldscheider and Waite 1986; Winkler-Dworak and Toulemon 2007). Kalmijn (2011; 2013) showed that, in Europe, men with better career prospects and positions on the labor market have higher chances of forming a union and getting married, while unmarried cohabitation was related to a lower socioeconomic position. As to the effect on the transition to parenthood, Corijn and Klijzing (2001) found that the effect of educational attainment was negative both for men and women in several Western European countries, but weaker for men than for women. Yet, in France, the effect was found to be positive for men, while for women a U-shaped effect was found – both low and highly educated women showing higher first birth rates compared to the medium educated (Winkler-Dworak and Toulemon 2007). Perhaps some of these inconsistencies relate to the fact that the effect of education may change over the life course. Previous studies suggest that the association between educational attainment and the transition to parenthood may depend on time since graduation (Brien et al. 1999; Martin-Garcia and Baizan 2006; Martin-Garcia 2009; Winkler-Dworak and Toulemon 2007). The economic rationale for this is that differences in earning potential may show up only a couple of years after graduation, after a professional career has been established and people get ready for family formation. This would hold most for people with advanced degrees.

In the literature on fertility from a couple's perspective, scholars have been looking at the relative influence of partners' characteristics on the transition to parenthood (Begall 2013; Corijn et al. 1996; Gustafsson and Worku 2006; Jalovaara and Miettinen 2013; Martin-Garcia 2009; Thomson 1997; Vignoli et al. 2012). These studies include individuals who were in a co-residential union at the time of data collection. This implies that those less likely to enter or stay in a union are more likely to stay out of the picture. This is a crucial shortcoming because it disregards the effects of mate selection on fertility. More specifically, we expect that men's education affects their fathering rates chiefly through its effect on forming cohabiting or marital unions with women. If this is true, studies from a couple's perspective might wrongly conclude that the male partner's education matters less for fertility than the female partner's, while in fact his education may matter as much or more, but only during a different stage of the family formation process.

Few studies account for the interrelationships between union formation and fertility, applying the simultaneous equations approach for hazards developed by Lillard and colleagues (Brien et al. 1999; Lillard 1993; Lillard and Waite 1993; Lillard et al. 1995; Lillard and Panis 2003). A couple of studies have analyzed the interrelationship between first union formation and the transition to parenthood for women (Baizan et al. 2003; Brien et al. 1999). Both studies concluded that the two processes share unobserved factors that jointly affect the experience of events. For men, the relationship between union formation and the transition to fatherhood has hardly been studied. Winkler-Dworak and Toulemon (2007) considered the role of men's selection into unions when analyzing the transition to fatherhood in France. Their results suggest that part of the positive effect of men's educational attainment on the transition to fatherhood was driven by the higher rate of union formation among highly educated men. In Finland, Jalovaara (2012) and Jalovaara and Miettinen (2013) found that socio-economic resources for Finnish men and women were important to be selected into unions as well as to become parents. Still, the authors find that the female partner's education has a stronger impact on the transition to parenthood than the male partner's education (Jalovaara and Miettinen 2013). The latter finding suggest, in line with our hypothesis, that his education matters more for being selected into a union rather than for becoming a father once selected into a union.

The Selection-into-Union Hypothesis

Based on the theoretical arguments and earlier empirical studies summarized above, we expect that the level of educational attainment has a consistently positive effect on men's transition to fatherhood, but that this effect is largely indirect, namely through its positive effect on the rate of union formation. The underlying assumption is that highly educated men tend to be attractive on the mating market. An economic reason for their attractiveness is their relatively high earning potential. Another attractive feature, at least for some women, may be that they are more likely to hold egalitarian gender-role attitudes, and thus may be more prone to share household chores with their partners (Coltrane 2000; Sullivan et al. 2014). Lower educated men have more difficulty finding a committed partner and therefore, all else equal, are expected to experience lower fatherhood rates (Kravdal and Rindfuss 2008; Lappegård and Rønsen 2013; Winkler-Dworak and Toulemon 2007).

These expectations hold for men who have completed their studies and who are no longer enrolled in education. The effect of enrolment in education is expected to be negative throughout. So, even if men who pursue a college degree will have their first child later, we are predicting consistently higher fatherhood rates for them once they have obtained their higher degree. Our hypothesis implies that the higher fatherhood rates for the college educated can be explained by the fact that they are able to match with a committed female partner more quickly than their low educated counterparts. When we model union formation and fatherhood jointly, we expect to find a consistent positive effect of educational attainment on union formation but no consistent effect on the transition to fatherhood.

A competing hypothesis, hence, is that the educational attainment level still has a positive effect on fatherhood rates after accounting for selection into union. At least three potential reasons for such a competing hypothesis could be mentioned. First, higher educated men may be more likely to enter parenthood even after accounting for union formation because they have been found to be more likely to marry (Jalovaara 2012). Second, some studies on couple level fertility have suggested that men's socioeconomic resources, associated with their education, may stimulate the couple's childbearing behavior (Jalovaara and Miettinen 2013; Kreyenfeld and Konietzka 2008; Vignoli et al. 2012). Third, partnered men with advanced education may experience higher fatherhood rates due to lower divorce and separation risks (Jalovaara 2011).

Data and Methods

Sample selection

To test our hypothesis, we have used survey data of the Generations and Gender Surveys (GGS, see http://www.ggp-i.org/) for 10 European countries that provide the information needed: Austria, Belgium, Bulgaria, Estonia, France, Hungary, Lithuania, Norway, Poland and Romania. We chose to replicate our empirical tests in these 10 countries rather than focusing on just one or a couple of countries. We included these particular countries because the GGS-data needed to test our hypothesis are available for them.[1] The number of countries is insufficient, however, to test the role played by country-level factors using multilevel models.

The GGS surveys include men and women between 18 and 79 years old and deal with topics such as fertility and partnership histories, the transition to adulthood, economic activity, care duties and attitudes (Vikat et al. 2007). Response rates are above 60% in all countries except Belgium (41.8%) and Lithuania (35.6%) (Fokkema et al. 2016), so caution is needed in particular for these two countries. For this study, we selected men born after 1949. Men were censored at age 45 for both union formation and first birth, because both first unions and first births very rarely occur at older ages even among men. We used information about the month and year of events. If the month was missing, we randomly imputed it. From an initial sample of 51224 men (for all countries), we excluded from the analysis men involved in same-sex relationships (n=163) and those born before 1950 (n=14881). Then we dropped cases with missing or misreported information on the date of first union (n=703) as well as date of first birth (n=28), cases where it was not possible to determine whether or not the event of interest occurred and cases for whom the event occurred before the 15th birthday (n=125 for first union and n=29 for first birth). After these selections, our sample totaled 35295 men. We distinguish between three birth cohorts: 1950-1959, 1960-1969, and 1970-1990[2]. Table 1 gives descriptive statistics for the samples and variables used.

[1] We excluded Russia and Georgia because of the very different cultural and institutional backgrounds of these countries. The data for the Czech Republic have become available by now, but not when we were conducting our analyses.

[2] The higher limit of this birth cohort differs among countries: 1983 (Estonia, Hungary); 1988 (Norway); 1989 (Lithuania); 1990 (Austria, Belgium); 1993 (Poland).

Measures

The date of first partnership formation has been coded using information on the month and year of the first reported co-residential partnership, distinguishing between unmarried cohabitation and marriage and keeping track of any changes in marital status. The GGS surveys collected information only on partnerships which lasted for at least three months (Vikat et al. 2007). To focus on the relationship between *first* union as well as first birth, respondents who experienced more than one co-residential union have been censored at the end of the first one, so that first births happening in higher order union are not considered. Hence, if no transition to fatherhood is observed, men are censored either (1) at age 45, (2) at the time of interview, or (3) at the break-up of the first union, whichever comes first. Only in Norway and Austria did the proportions of first births in higher order unions exceed 10% (namely 12 and 11%, respectively). In France, it was almost 9%, in Belgium and Hungary 7%, around 1% in the rest of the countries. As a robustness check, for countries where the proportion of first births in higher order unions exceeded 6%, we ran a version of our models that was modified to include higher order unions. The results do not deviate substantially form the ones reported below.

Table 1 Descriptive statistics by country: percentage distributions of variables used, sample size, and number of events

	Austria	Belgium	Bulgaria	Estonia	France	Hungary	Lithuania	Norway	Poland	Romania
Cohort %										
1950-1959	NA	27.55	20.10	28.52	28.09	29.40	25.16	25.66	29.20	31.30
1960-1969	29.84	28.07	29.59	27.22	30.71	22.76	24.84	28.00	20.07	29.65
1970-1990	70.16	44.38	50.31	44.26	41.20	47.84	50.00	46.35	50.72	39.05
Education %										
Low	10.54	26.25	23.79	17.80	21.54	12.97	14.47	19.70	12.20	21.44
Medium	72.35	38.19	62.20	62.39	50.27	73.38	64.79	48.41	67.98	66.62
High	17.11	34.99	13.96	19.81	28.20	13.65	20.75	30.80	19.34	11.94
Unknown	0.00	0.56	0.05	0.00	0.00	0.00	0.00	1.09	0.47	0.00
Parents' education %										
Both low	24.13	44.85	42.07	32.14	49.56	34.72	35.71	21.25	31.89	60.62
Only father medium-high	25.36	18.43	9.51	12.12	17.25	21.99	7.39	22.42	14.41	16.59
Only mother medium-high	9.73	11.85	12.21	20.02	11.55	7.20	21.19	18.86	10.92	4.03
Both medium-high	36.00	19.51	32.93	35.39	12.50	35.38	27.03	32.78	38.91	17.07
Both unknown	4.79	5.36	3.27	0.32	9.14	0.71	8.68	4.69	3.87	1.69
Partner's education %										
Low	12.12	11.63	15.16	8.60	11.16	8.44	4.09	7.13	6.59	20.46
Medium	48.22	13.02	33.89	55.41	20.76	30.72	38.28	17.26	47.05	38.51
High	9.06	18.17	14.18	16.02	14.67	8.39	16.13	17.49	18.46	7.34
Unknown	0.15	39.49	0.59	0.00	31.42	22.60	14.20	34.71	0.42	10.35
Not in union	30.45	17.69	36.18	19.97	22.00	29.85	27.29	23.41	27.48	23.34
Siblings %										
No siblings	10.13	9.90	13.66	14.50	6.59	11.67	16.89	4.90	8.07	15.12
1	33.30	29.37	55.54	46.32	25.86	48.58	41.06	29.15	29.38	31.91
2	24.34	21.84	14.57	21.32	25.54	21.33	21.74	32.05	24.92	22.05
3+	32.23	38.88	16.22	17.86	42.01	18.42	20.31	33.90	37.63	30.91
Number of events by type										
Number of first births	809	1186	2273	1170	1359	1675	1873	2690	3833	2457
Number of first unions	1366	1903	2597	1479	2202	2651	2488	3857	4458	2986
Sample size	1964	2312	4069	1848	2823	3779	3422	5036	6147	3895

The date of the transition to fatherhood was back-dated by 8 months to avoid anticipation bias (Baizan et al. 2003), based on the date of birth reported by the respondent for his first biological child, if any. It is known that men may underreport their fertility (by a major margin in the US and the UK according to Rendall et al. 1999; Joyner et al. 2012; only minor underreporting according to the estimates by Alich 2009 for Russia). The underreporting of male fertility is likely to be selective with respect to union status and education: births outside marital or cohabiting unions and births to men with low educational attainment are expected to be disproportionally at risk to remain underreported. Given that low educated men are likely to be underrepresented in our data anyway due to non-response (Fokkema et al. 2016), this implies that births to low educated men are most likely underreported in our data, particularly births to men who are not living with the mother of their first child.

Our analysis looks at two dimensions of education: enrolment and attainment. Both variables are constructed as time-varying covariates. To capture the enrolment dimension, time since graduation is included as a categorical variable with three categories: (1) still enrolled in education; (2) up to two years after graduation; (3) more than two years after graduation. We distinguish between the very recent graduates and those who left school or college more than two year ago because it takes some time to find a job after graduation (typically 1 to 2 years, Quintini and Manfredi 2009), earn a living and be ready for family formation (cf. Martin-Garcia 2009). The end of enrolment is based on the time of actual graduation as reported by the respondent in most cases (93.6%). In case this information was missing, we imputed the end of enrolment as the standard age at graduation for the relevant attainment level and country. We had to do this for 6.4% of the cases overall. For men who completed a college level degree, where study duration is more varying, the percentage of missing data on time of graduation was only 0.4%, so we hardly had to impute for the latter group. Still, as a robustness check, we re-ran our models after dropping cases where information on the actual date of graduation was missing. The results remained basically the same and would not affect our conclusions.

For educational attainment, we grouped men into three levels (low, medium, high), collapsing categories from the International Standard Classification of Education (ISCED 1997). The first group includes those who completed primary plus lower secondary school (at least 8 years of schooling, ISCED 0, 1, and 2). The medium category consists of men who attained the upper-secondary and those who also got a post-secondary level (ISCED 3 and 4). Finally, highly educated men are those who got a bachelor/master/PhD degree (ISCED 5 and 6).

Corijn and Klijzing (2001) argue that the effect of education on the transition to parenthood may not be constant but changing over the life course, which would represent a violation of the proportionality assumption. One way to try to account for this time-dependency is to interact age with education. However, earlier studies suggest that the time-dependence is a function of time since graduation rather than age per se (Brien et al. 1999; Martin-Garcia and Baizan 2006; Martin-Garcia 2009) – with age at graduation obviously correlating with the level of the degree obtained. We have therefore included interaction terms between men's educational attainment and the three categories of years since graduation in all models. Apart from that, we also include five-year age splines to accommodate any nonlinear relationship between age and the events of interest (see below for more technical details).

The female partner's education is categorized in the same way as for the male respondent. To catch the effect of a long-term dimension of the social status, we included parents' educational attainment, coded into 4 categories ("both parents low educated", "only the father at least medium educated", "only the mother at least medium educated", "both parents at least medium educated"). Since it has been showed that individuals with more siblings are more prone to start a family (Murphy 2013), we included the number of siblings as a time constant variable.

All variables mentioned so far have been included in both the model of first birth and first union. Next, the model of first birth includes the time-varying endogenous variable “union status”, indicating whether the male respondent is living in a co-residential union or not. Once a man is living with a female partner, we distinguish between those partnered with a low, medium or highly educated woman. We added a category “not available” to accommodate men in union but with missing information on the partner’s education. We also included a time varying dummy variable indicating whether the union is a formal marriage rather than unmarried cohabitation. Finally, in the model of first union formation, a time-varying dummy variable for the conception of the first child (birth date backdated by 8 months) is included.

Analytical Strategy

We test the selection-into-union hypothesis in two steps. First, we model first union formation and the transition to parenthood separately. In the second step, we model the two processes jointly. For the first step we fitted both a piecewise exponential hazard model (see, e.g., Blossfeld et al. 2007) using the STATA software and a piecewise linear hazard model using the aML package (Lillard and Panis 2003). We compared the results from both approaches to check whether the results would be similar, which was indeed the case. We present the results from the piecewise linear approach, which was also the one applied in the second step. A general formulation of the piecewise linear hazard model is:

$$\ln h(t) = \gamma\,'T(t) + \beta\,'X(t)$$

$\ln h(t)$ is the log-hazard of occurrence at time t, $\gamma\,'T(t)$ captures the baseline hazard duration dependence, and $\beta' X\ (t)$ represent the covariates (both fixed and time-varying) which shift the baseline hazard up or down. In the piecewise linear specification $\gamma\,'T(t)$, we implement five year age intervals to parameterize the baseline log-hazard. The duration dependence is characterized by a pattern of nodes and slopes as well as an origin (Panis 1994). We set the latter at the 15th birthday of the respondent.

To address the endogeneity of union formation and parenthood, with both processes affecting each other, we jointly model the two processes and estimate the correlation between residuals to represent unobserved heterogeneity. In doing so, we account for unmeasured, time-constant factors that simultaneously affect union formation and parenthood. One advantage of the joint model is that we account for the fact that individuals with a higher probability of experiencing the two events will leave the population at younger ages. As a result, the observed hazard at older ages for both events strongly reduces due to selection. When we do not account for this, the baseline hazard represents also this selection effect rather than only the actual effect of age (Baizan et al. 2003). Our statistical estimations follow the framework developed by Lillard (1993). In formal terms, we have:

$$\ln h(t)^{F} = \gamma\,'T(t) + \beta\,'X(t) + \varepsilon$$

$$\ln h(t)^{U} = \gamma\,'T(t) + \beta\,'X(t) + \delta$$

The superscripts *F* and *U* refer to the equation for fatherhood and union formation, respectively. The random variables ε and δ represent unobserved heterogeneity terms, which are assumed to have a joint bivariate standard normal distribution:

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

Since we consider only *first* unions and births, we deal with non-repeatable events. Aassve et al. (2003) showed that the estimates for hazard models for non-repeatable, correlated events may be sensitive to the variance of the unobserved heterogeneity term. To check the sensitivity in our case, we ran our models in two versions: one with fixing the variance to 1 (the variance of the standard normal distribution) and one where we estimated the variance empirically. The results tend to be robust as to the direction of the effects and their significance, changing slightly with regard the magnitude of the effects. In France and Estonia, the estimated correlation between unobserved factors changes both in sign and statistical significance, but the estimates for the fixed effects of education – the ones of substantive interest – remain stable also in these two countries.

The joint modelling approach is able to account for the correlation between the processes of union formation and entry into fatherhood for various reasons, including causal effects running in both directions. Still, we were concerned about how “shotgun” union formation may affect our estimates, i.e. situations where pregnancy of the female partner induces union formation. As a robustness check, we re-ran our models where we backdated all births by 18 months, which implies that only children conceived about 9 months after union formation are counted as born within the union. The results are in line with the ones reported below.

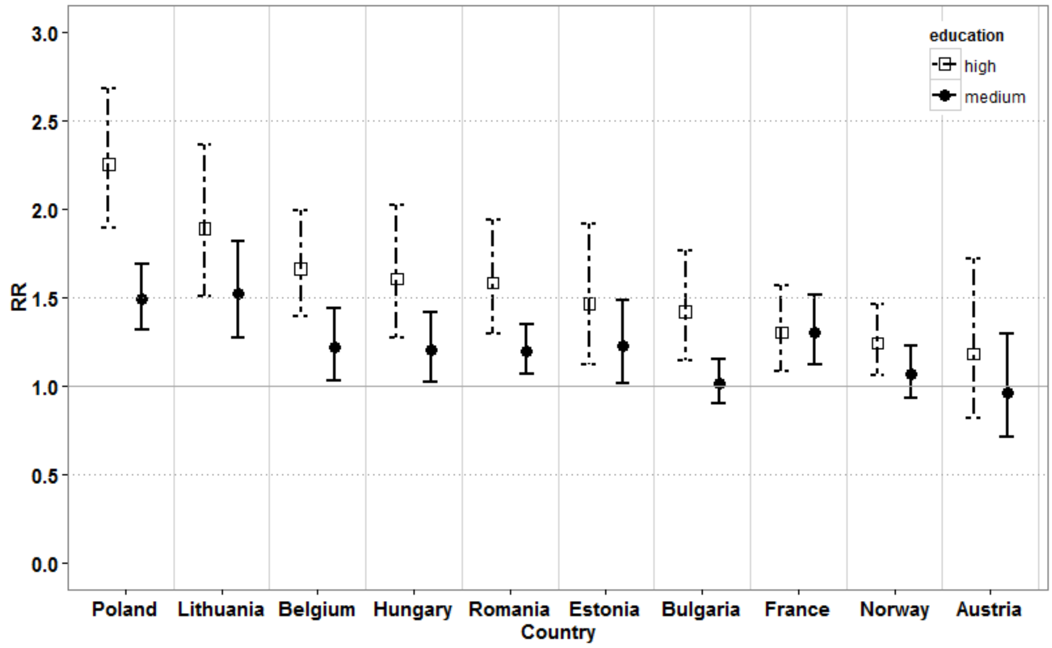
The models are estimated and replicated for each country-sample separately. We did not apply multilevel modelling because our focus is not to test the effect of country characteristics. Rather, we want to replicate the same hypothesis testing in different contexts. Anyway, the number of countries would be too small to apply multilevel modelling and test hypotheses about the role of country characteristics. In an additional analysis not reported below, we pooled our 10 countries. The results we then got are an averaged, stylized summary of our main finding, consistent with what we do report below for countries separately.

Results

We have fitted 3 models for each of the 10 countries. Appendix 1 reports the estimates for all 30 models. For each country, Model 1 is the model of the transition to fatherhood that does not control for union status. Controls that are included, as well as in all subsequent models, are age, cohort, time since graduation, own educational attainment of the male respondent, parental educational attainment, and number of siblings. Model 2 adds the control for union status, type of union, and the educational attainment of the female partner, if any. Figure 1 plots the effects of own educational attainment on the transition to fatherhood, estimated from Model 1, along with their 95% confidence intervals; Figure 2 is doing the same for Model 2. All figures compare the rates for men with high or medium educational attainment with the reference category with low educational attainment (represented by the 1.0 line), two years or more after graduating from school or college.

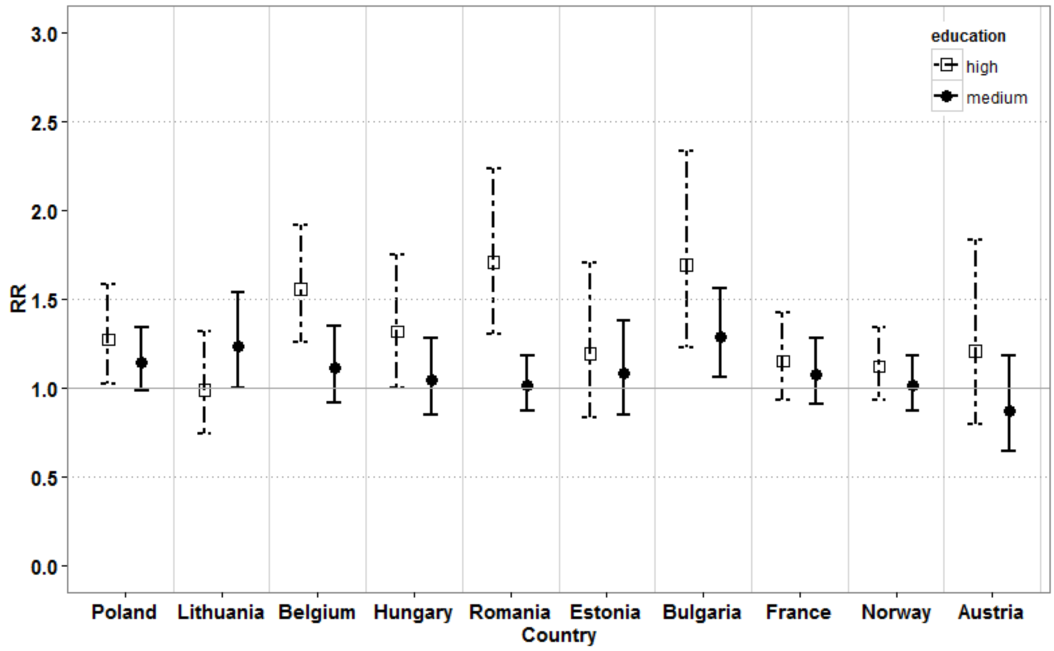
As can be seen in Figure 1, educational attainment has a consistently positive effect on the transition to fatherhood when union status is not controlled for. In all countries except Austria, the difference between low educated and highly educated men is statistically significant. In the majority of countries, the fatherhood rates are more than 50% higher for college educated men compared to men with low educational attainment. The relative rate for medium educated men is in most countries in between. Figure 2 shows that once we control for union status, the effect of education is strongly reduced in most of the countries. Thus, most of the positive effect of education on the transition to fatherhood appears to be driven by union formation.

Fig. 1 Estimates for the effect of educational attainment on the transition to fatherhood without controlling for union status: relative risks for highly educated (square with dashed bars) and medium educated men (circles with solid bars) compared to low educated men (1.0 reference line)



Note: displayed relative risks apply for men who have been out of school for more than two years

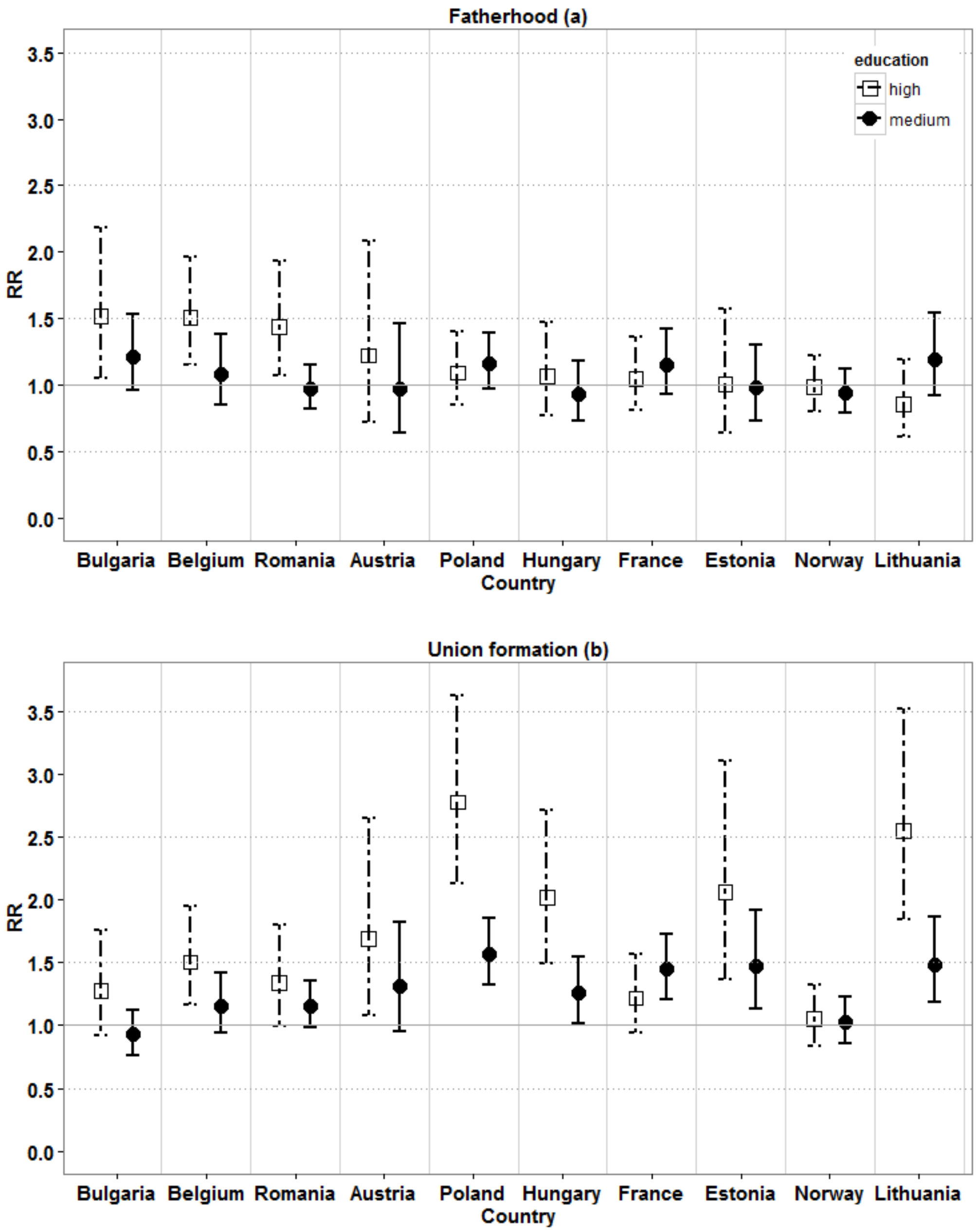
Fig. 2 Estimates for the effect of educational attainment on the transition to fatherhood after controlling for union status: relative risks for highly educated (square with dashed bars) and medium educated men (circles with solid bars) compared to low educated men (1.0 reference line)



Note: displayed relative risks apply for men who have been out of school for more than two years

Next, we account for the selection-into-union effect and the endogeneity of union formation and fatherhood by simultaneously estimating joint models of both processes. This is Model 3 in Appendix 1. Table 2 reports for each country the estimates for the fatherhood equation of Model 3, Table 3 reports estimates for the union formation equation. To facilitate interpretation, Figure 3 plots the effects of male educational attainment on (a) fatherhood and (b) union formation. It turns out that there is generally no effect of educational attainment on fatherhood rates after accounting for the selection-into-union process in most countries. Belgium, Romania and Bulgaria are the only countries where we still see a direct, statistically significant effect: in these three countries, college educated men exhibit higher fatherhood rates compared to men with low educational attainment. In Austria, the point estimate for highly educated men is also a bit above the 1.0 reference line of the low educated men, but the difference is not statistically significant. For medium educated men, there are no significant differences in any country. In contrast, educational attainment positively influences union formation in most countries, with statistically significant differences in all countries except Bulgaria and Norway. In France, unexpectedly, medium educated men appear to exhibit higher union formation rates than highly educated men, but the difference is not statistically significant. All-in-all, the estimates quite consistently point to a positive educational gradient in union formation rates.

Fig. 3 Estimates for the effect of educational attainment on the transition to fatherhood (a) and union formation (b) from a joint model simultaneously estimating the parameters for both processes: relative risks for highly educated (square with dashed bars) and medium educated men (circles with solid bars) compared to low educated men (1.0 reference line)



Note: displayed relative risks apply for men who have been out of school for more than two years

The control variables in all models (Appendix 1, and Tables 2 and 3 for the joint Model 3) tend to have the expected effects – discussing these is beyond the focus of this paper. The joint model also yield estimates of the correlation between residual, unobserved terms in the fatherhood and union formation equations (see Appendix 2). Both processes are obviously positively correlated. Indeed, in all countries, the estimated correlation between the unobserved factors in the joint model is positive as long as we exclude any of the two endogenous variables (union status in the equation of first birth or conception in the equation of first union). Interestingly, however, once we include in the system of equations both endogenous variables, the positive correlation between unobserved factors disappears in almost all countries. Only in Belgium the correlation between unobserved factors remains positive. In Austria it is not significant (probably due to the small sample size), in Norway it is positive but not significant, while in France the significance and sign of the correlation term are sensitive to the values of the unobserved heterogeneity factors.

In all eastern European countries (Bulgaria, Hungary, Poland and Romania as well as in Lithuania), the correlation between unobserved factors in both processes turns negative (the estimated correlation is not robust for Estonia). Such negative correlation implies that, after accounting for the effect of union formation on fatherhood and vice versa, there are some unobserved factors that enhance the experience of one event but delay the other. For instance, it could be that men with unobserved personality traits who are particularly eager to enter a romantic relationship decide to live with a partner relatively early while they are at the same time not eager to actually start fathering a child. Testing this interpretation is beyond the focus of this paper.

Table 2 Estimated regression coefficients: full joint model, first birth, 10 countries

	Austria		Belgium		Bulgaria		Estonia		France		Hungary		Lithuania		Norway		Poland		Romania	
	b	se	b	se	b	se	b	se	b	se	b	se	b	se	b	se	b	se	b	se
Duration splines																				
15-19	**0.54**	0.08	**0.54**	0.13	**0.30**	0.04	**0.51**	0.06	**0.51**	0.10	**0.46**	0.06	**0.71**	0.06	**0.68**	0.06	**0.75**	0.05	**0.51**	0.06
20-24	0.06	0.04	**0.13**	0.04	0.04	0.02	-0.08	0.05	0.00	0.03	0.00	0.02	0.00	0.03	**0.05**	0.02	**0.07**	0.02	0.03	0.02
25-29	-0.04	0.04	**0.10**	0.03	**-0.18**	0.02	**-0.13**	0.04	**0.12**	0.02	-0.04	0.02	**-0.14**	0.02	**0.09**	0.02	**-0.13**	0.02	**-0.19**	0.02
30-34	0.01	0.05	**-0.13**	0.03	**-0.14**	0.04	**-0.15**	0.06	**-0.13**	0.03	**-0.10**	0.03	**-0.22**	0.04	-0.01	0.02	**-0.13**	0.02	**-0.14**	0.03
35-39	**-0.26**	0.09	**-0.12**	0.05	-0.12	0.07	**-0.35**	0.12	**-0.15**	0.06	**-0.24**	0.06	-0.07	0.07	**-0.24**	0.05	**-0.18**	0.05	**-0.45**	0.08
40+	-0.06	0.29	**-0.41**	0.15	-0.30	0.18	-0.24	0.24	**-0.23**	0.14	0.05	0.10	**-0.59**	0.22	-0.21	0.11	**-0.42**	0.14	-0.02	0.16
Constant	**-7.91**	0.47	**-8.39**	0.64	**-6.48**	0.23	**-7.02**	0.34	**-7.75**	0.49	**-7.13**	0.34	**-8.20**	0.35	**-8.16**	0.31	**-8.08**	0.27	**-7.74**	0.31
Cohort (Ref. = 1970-1990)																				
1950-59	-	-	-0.03	0.12	**0.22**	0.10	**0.56**	0.14	0.10	0.11	**-0.41**	0.10	0.17	0.09	**0.66**	0.08	**0.25**	0.07	0.14	0.08
1960-69	**0.48**	0.10	0.16	0.10	**0.34**	0.08	**0.66**	0.12	**0.18**	0.09	**0.38**	0.09	**0.34**	0.09	**0.40**	0.06	**0.44**	0.06	**0.35**	0.08
Education (Ref.= Low educated with at least 2 years after leaving school)																				
Enrolled	**-0.53**	0.24	**-0.60**	0.17	-0.19	0.14	**-0.44**	0.18	**-0.49**	0.16	**-0.46**	0.16	0.03	0.15	**-0.44**	0.09	**-0.30**	0.10	-0.19	0.12
Low0-2	**1.13**	0.42	0.24	0.42	0.39	0.22	0.02	0.34	0.59	0.66	0.61	0.34	-0.04	0.23	0.15	0.30	-0.85	0.64	-0.24	0.49
Medium0-2	-0.22	0.26	-0.43	0.31	-0.03	0.15	**-0.65**	0.20	-0.42	0.24	-0.13	0.18	0.08	0.15	-0.09	0.12	0.03	0.11	0.02	0.13
Medium2+	-0.04	0.21	0.08	0.12	0.19	0.12	-0.02	0.15	0.14	0.11	-0.08	0.12	0.18	0.13	-0.07	0.09	0.15	0.09	-0.03	0.09
High0-2	-0.28	0.32	**-0.59**	0.21	**0.55**	0.19	-0.36	0.25	-0.20	0.19	0.02	0.21	0.01	0.19	**-0.32**	0.12	**-0.48**	0.14	-0.04	0.18
High2+	0.20	0.27	**0.41**	0.14	**0.42**	0.19	0.00	0.23	0.05	0.13	0.07	0.16	-0.16	0.17	-0.01	0.11	0.09	0.13	**0.36**	0.15
Unknown	-	-	0.63	0.79	-	-	-	-	-	-	-	-	-	-	-0.23	0.34	-0.43	0.36	-	-

(Continued on next page)

Table 2 (continued)

Parents' education (Ref.=Both low)																				
Only father medium-high	-0.25	0.14	-0.04	0.12	-0.10	0.13	-0.04	0.18	-0.01	0.11	-0.07	0.09	-0.21	0.13	-0.08	0.08	-0.07	0.08	-0.04	0.08
Only mother medium-high	**-0.43**	0.19	0.12	0.15	-0.02	0.12	-0.08	0.13	**-0.28**	0.12	-0.05	0.15	-0.03	0.11	-0.06	0.09	-0.10	0.09	-0.31	0.18
Both medium-high	-0.25	0.14	-0.22	0.13	**-0.20**	0.10	**-0.31**	0.13	-0.15	0.13	**-0.22**	0.09	-0.09	0.10	**-0.21**	0.08	**-0.21**	0.07	**-0.35**	0.10
Both unknown	-0.07	0.21	0.15	0.19	-0.07	0.20	-0.89	0.94	0.12	0.15	-0.06	0.53	0.05	0.11	**-0.40**	0.14	-0.15	0.15	0.16	0.28
Siblings (Ref.= No siblings)																				
1	**0.56**	0.20	0.15	0.15	-0.01	0.10	0.19	0.13	-0.18	0.16	0.07	0.11	**0.34**	0.11	-0.02	0.14	0.15	0.10	0.16	0.10
2	**0.53**	0.21	**0.32**	0.15	0.18	0.14	0.20	0.16	-0.05	0.16	0.13	0.13	**0.34**	0.12	0.05	0.14	**0.29**	0.10	0.09	0.11
3+	**0.69**	0.20	**0.33**	0.14	**0.32**	0.14	0.15	0.16	-0.04	0.16	0.16	0.13	**0.38**	0.12	0.19	0.14	**0.43**	0.10	**0.28**	0.10
Partner's education (Ref.=Not in union)																				
Low	**2.42**	0.23	**2.52**	0.21	**4.29**	0.12	**3.81**	0.29	**3.24**	0.19	**3.24**	0.17	**3.40**	0.21	**2.66**	0.14	**2.77**	0.16	**4.55**	0.12
Medium	**2.19**	0.18	**2.62**	0.19	**4.32**	0.10	**3.76**	0.25	**3.02**	0.16	**3.29**	0.14	**3.16**	0.13	**2.70**	0.11	**2.72**	0.10	**4.47**	0.12
High	**1.89**	0.22	**2.56**	0.17	**4.24**	0.12	**3.68**	0.27	**2.73**	0.17	**3.13**	0.16	**3.07**	0.15	**2.45**	0.11	**2.53**	0.11	**4.00**	0.16
Unknown	-	-	**1.44**	0.21	**3.28**	0.37	-	-	**1.01**	0.21	**1.22**	0.17	**1.07**	0.17	**1.73**	0.12	**2.20**	0.83	**3.15**	0.15
Married (Ref. = Not married)	**1.60**	0.11	**0.87**	0.10	**0.47**	0.08	**0.67**	0.10	**1.35**	0.08	**0.92**	0.10	**1.17**	0.12	**1.13**	0.06	**1.11**	0.08	**0.33**	0.09
SigmaEps	1.00		1.00		1.00		1.00		1.00		1.00		1.00		1.00		1.00		1.00	
SigmaDelta	1.00		1.00		1.00		1.00		1.00		1.00		1.00		1.00		1.00		1.00	
Rho	-0.22	0.17	**0.35**	0.15	**-0.57**	0.03	-0.53	0.28	0.10	0.14	**-0.66**	0.09	**-0.89**	0.03	0.17	0.11	**-0.59**	0.05	**-0.69**	0.05

Note: estimates with p-values lower than 0.05 are in bold

Table 3 Estimated regression coefficients: full joint model, first union, 10 countries

	Austria		Belgium		Bulgaria		Estonia		France		Hungary		Lithuania		Norway		Poland		Romania	
	beta	se	beta	se	beta	se	beta	se	beta	se	beta	se	beta	se	beta	se	beta	se	beta	se
Duration splines																				
15-19	**0.55**	0.04	**0.27**	0.03	**0.54**	0.04	**0.71**	0.05	**0.80**	0.10	**0.68**	0.04	**0.77**	0.05	**0.83**	0.03	**0.83**	0.05	**0.66**	0.05
20-24	**0.17**	0.02	**0.20**	0.02	**0.22**	0.02	**0.24**	0.02	**0.24**	0.03	**0.26**	0.02	**0.30**	0.02	**0.29**	0.01	**0.33**	0.01	**0.37**	0.02
25-29	**0.07**	0.03	-0.05	0.03	**-0.14**	0.02	**-0.11**	0.03	-0.04	0.02	**-0.07**	0.02	**-0.08**	0.02	-0.02	0.02	**-0.07**	0.02	**-0.05**	0.02
30-34	**-0.10**	0.05	**-0.10**	0.04	**-0.13**	0.04	-0.04	0.06	-0.06	0.03	**-0.12**	0.04	**-0.11**	0.04	-0.02	0.03	**-0.17**	0.03	**-0.07**	0.03
35-39	-0.08	0.10	-0.07	0.07	**-0.19**	0.08	-0.20	0.14	**-0.17**	0.06	**-0.18**	0.09	-0.02	0.06	**-0.20**	0.05	**-0.16**	0.05	**-0.22**	0.07
40+	0.23	0.21	-0.27	0.15	-0.09	0.14	-0.16	0.31	0.01	0.14	-0.11	0.20	**-0.61**	0.20	-0.01	0.09	-0.15	0.10	-0.19	0.15
Constant	**-5.93**	0.28	**-3.64**	0.18	**-5.82**	0.20	**-6.85**	0.29	**-6.78**	0.29	**-7.06**	0.23	**-7.52**	0.27	**-7.30**	0.21	**-8.45**	0.25	**-7.00**	0.25
Cohort (Ref. = 1970-1990)																				
1950-59	-	-	-0.11	0.09	**0.48**	0.08	**0.36**	0.11	0.09	0.08	**0.68**	0.07	-0.16	0.08	0.10	0.06	**0.17**	0.06	**0.23**	0.07
1960-69	0.04	0.09	-0.14	0.09	**0.38**	0.07	**0.27**	0.10	0.05	0.08	**0.47**	0.08	-0.03	0.08	**0.15**	0.06	-0.06	0.06	**0.34**	0.07
Education (Ref.= Low educated with at least 2 years after leaving school)																				
Enrolled	0.11	0.17	**-0.44**	0.10	**-0.77**	0.11	0.01	0.15	**-0.56**	0.10	**-0.25**	0.12	0.17	0.13	**-0.25**	0.08	0.14	0.10	**-0.53**	0.10
Low0-2	**0.94**	0.34	**-0.38**	0.16	**0.55**	0.18	**0.74**	0.24	-0.26	0.44	**1.17**	0.24	**0.44**	0.19	**0.51**	0.19	0.68	0.35	0.35	0.29
Medium0-2	0.24	0.18	**-0.37**	0.14	**-0.76**	0.12	-0.21	0.16	0.01	0.12	-0.03	0.13	**0.55**	0.13	-0.08	0.10	**0.38**	0.10	-0.15	0.11
Medium2+	0.27	0.17	0.14	0.10	-0.08	0.10	**0.38**	0.14	**0.37**	0.09	**0.23**	0.11	**0.40**	0.12	0.02	0.09	**0.45**	0.09	0.14	0.08
High0-2	0.31	0.26	0.07	0.14	-0.07	0.16	0.37	0.21	-0.05	0.13	0.32	0.18	**0.68**	0.15	0.10	0.12	**0.80**	0.12	**0.38**	0.15
High2+	**0.53**	0.23	**0.41**	0.13	0.24	0.16	**0.72**	0.21	0.20	0.13	**0.70**	0.15	**0.93**	0.17	0.05	0.12	**1.02**	0.14	0.29	0.15
Unknown	-	-	-0.38	0.34	-	-	-	-	-	-	-	-	-	-	0.13	0.24	0.26	0.39	-	-

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Table 3 (continued)

Parents' education (Ref.=Both low)																				
Only father medium-high	0.13	0.11	0.10	0.10	**-0.28**	0.12	0.31	0.16	**0.18**	0.09	0.09	0.08	-0.14	0.13	0.10	0.08	0.13	0.07	-0.05	0.08
Only mother medium-high	**0.41**	0.16	0.13	0.12	**-0.48**	0.10	-0.02	0.11	0.21	0.11	-0.23	0.12	0.11	0.10	0.05	0.08	0.06	0.09	-0.16	0.15
Both medium-high	0.01	0.12	0.03	0.11	**-0.36**	0.08	**0.29**	0.12	-0.04	0.11	0.10	0.08	0.09	0.09	-0.08	0.08	**0.24**	0.06	0.05	0.09
Both unknown	0.36	0.21	0.07	0.14	-0.14	0.18	-0.16	0.42	0.05	0.13	-0.35	0.49	-0.12	0.11	**-0.29**	0.14	0.22	0.12	0.16	0.22
Siblings (Ref.= No siblings)																				
1	0.19	0.14	0.04	0.13	**0.20**	0.09	0.08	0.11	0.06	0.15	0.04	0.09	0.19	0.10	0.16	0.12	0.05	0.10	-0.01	0.09
2	0.16	0.15	0.12	0.14	**0.48**	0.12	**0.34**	0.13	0.03	0.15	**0.26**	0.10	**0.35**	0.11	**0.28**	0.12	0.16	0.10	**0.21**	0.10
3+	0.05	0.15	-0.08	0.13	**0.44**	0.12	0.14	0.14	-0.03	0.15	**0.32**	0.10	0.21	0.11	**0.24**	0.12	0.09	0.10	**0.28**	0.09
Conception (Ref.=No conception)	**2.04**	0.20	**1.43**	0.37	**3.65**	0.12	**3.49**	0.27	**2.09**	0.21	**4.02**	0.12	**3.83**	0.09	**1.97**	0.14	**3.80**	0.07	**3.73**	0.11
SigmaEps	1.00		1.00		1.00		1.00		1.00		1.00		1.00		1.00		1.00		1.00	
SigmaDelta	1.00		1.00		1.00		1.00		1.00		1.00		1.00		1.00		1.00		1.00	
Rho	-0.22	0.17	**0.35**	0.15	**-0.57**	0.03	-0.53	0.28	0.10	0.14	**-0.66**	0.09	**-0.89**	0.03	0.17	0.11	**-0.59**	0.05	**-0.69**	0.05

Note: estimates with p-values lower than 0.05 are in bold

Discussion

Several authors have recently argued that men's attitudes, intentions and behaviors are becoming increasingly important factors to understand patterns and trends in family formation (Goldscheider et al. 2014; Huinink and Kohli 2014; Van Bavel 2012). Yet, empirical research so far has not yielded clear and consistent results about how men's characteristics affect the transition to parenthood. In this paper we have argued that the explanation for this may be that men's characteristics perhaps matter most in the process of union formation, where men with more attractive features for women are more likely to be selected as their partners. Empirical analyses that look at the couple level, after unions have been formed, will then perhaps fail to see any clear effects of male characteristics, net of the effects of female characteristics.

More specifically, in this paper we wanted to test the selection-into-union hypothesis with a focus on educational attainment. We expected that educational attainment has a consistently positive effect on men's transition to fatherhood, but that this effect is largely indirect, namely through its positive effect on the rate of union formation. Our results generally supported the hypothesis, with differences between European countries but with a clear overall pattern: there is a positive educational gradient in men's union formation but, after accounting for that, not in their transition to fatherhood. This pattern shows up particularly for men who left school more than two years ago – presumably the time needed for the majority to have gained an established position in the labor market. Before that, so just after leaving school, the results are more mixed but also then we hardly find support for a competing hypothesis that men's educational attainment would have a direct positive effect on fatherhood. Only in Belgium, Bulgaria, and Romania, we still find a positive effect of high educational attainment on fatherhood compared to low educational attainment.

Our hypothesis appears to apply particularly strongly in Central and Eastern European countries, including the Baltic countries Lithuania and Estonia. In the latter two countries, there is a clear positive educational gradient in fatherhood rates, but this gradient vanishes completely when selection into union is accounted for. The positive educational gradient in union formation came out even more clearly in the joint modeling framework.

In Western Europe, Belgium is the only country where a positive educational gradient in fatherhood remains in the joint modeling framework, after accounting for union formation. Here, we can only speculate about the reasons for the Belgian exception – and it remains to be seen whether future studies with other data can replicate our GGS-based finding. If real, a positive educational gradient in fatherhood (after accounting for selection-into-union) could signify that there are strong father role expectations, and that the highly educated are the most likely to meet these expectations. An alternative explanation could be that the effect of educational attainment represents a direct income effect on the transition to fatherhood. We speculate, however, that this would rather hold for the other two countries where a positive effect of education on fatherhood remains in the joint model – namely in Romania and Bulgaria, two economically disadvantaged Eastern European countries. In these countries, more highly educated men can perhaps more easily “afford” to father children – and perhaps low educated fail to report any children they have, if they are aware of them at all.

By jointly modeling union formation and fatherhood, this study was able to overcome major limitations of earlier work on education and the transition to fatherhood. Earlier studies with models that included the education of the male partner along with the education of the female partner only included men who were partnered already. Our argument is that this approach suffers from a crucial selection bias because men’s education chiefly influences fertility through the selection into unions. That crucial selection effect is missed in studies investigating fertility at the couple level.

Alternatively, studies in the transition-to-adulthood tradition have typically failed to control for the female partner's educational attainment, ignoring strong educational homogamy and, hence, unable to tell whether it is his or her education that matters.

This study has its limitations. First, we cannot be sure about whether or not the men in our sample may have had other, perhaps unacknowledged children with other women than the ones identified in our data. It may be the case that some men have underreported their children in GGS data, either unintentionally (they simply do not know about those children) or intentionally (Alich 2009). In this last case, it is likely that intentional underreporting will be selective with respect to union status and education. Men who are afraid to be socially sanctioned for their extra-union childbearing behavior, or men who are in trouble for paying child alimony, may tend to omit children born from dissolved unions and who do not co-reside with them anymore (Lindberg et al. 1998; Joyner et al. 2012). This implies that births to low educated men are most likely underreported in our data, particularly births to men who are not living with the mother of their first child. Most likely, however, men will not be engaged in much active fathering of their unacknowledged or unreported biological children. Second, we have disregarded the distinction between marriage and unmarried cohabitation in this study. For example, we have not modeled the selection process which would lead highly committed men to marry for having children rather than to cohabit. This could affect the correlation terms between unobserved factors: men who signal stronger commitment by marrying would probably have higher fatherhood rates than those who choose to cohabit. As a result, the positive correlation between the transition to parenthood and marriage would be stronger compared to the correlation with unmarried cohabitation. The aim of this paper, however, was not to analyze the role of different kinds of living arrangement histories, but rather to assess the role played by men's educational attainment in their transition to parenthood through the selection into union. Note that, while we do not model the selection process into marriage versus cohabitation, we do control for the distinction in the equation for the transition to fatherhood. Third, our paper focused only on two dimensions of education: enrolment and attainment. Still it would be interesting to test if the selection-into-union hypothesis holds with regard to the effect of educational field of study. Finally, future research could also address the role of mating market composition in terms of educational attainment on the selection effect at the time of union, including both individual- and macro-level indicators.

Even with all mentioned limitations in mind, we argue that the selection-into-union effect should be taken into account, particularly when comparing the role of education in the transition to parenthood between men and women. Currently, the consensus in the field seems to be that women's education matters more for fertility than men's. However, this paper has shown that earlier research may have strongly underestimated the role of men's education because the process of their selection into unions has been ignored.

With the reversal of the gender gap in education, the selection of men into unions based on their education may even become a more important factor, given that the growing group of college educated women typically are looking for partners with the same educational status (Van Bavel 2012). More generally, we speculate that role of men's characteristics, intentions and behaviors may become more and more important for future fertility trends and patterns because college educated women are increasingly "competing" for men with similar education. The relative scarcity of the latter on the marriage market may enhance their power to have they say in decision-marking about fertility. But if we fail to account for the selection-into-union process, we risk missing that point.

It would be interesting to compare the effect of education on fertility through selection-into-union between men and women. Perhaps selection on education is playing an increasingly important role for women, as their contributions to the household budgets are becoming more significant and as their own earning potential is playing an increasingly important role in mate selection by men. Several studies have reported that the positive effect of education, and its earning potential, on marriage rates has strongly increased over time (Oppenheimer 1994; 1997; Sweeney 2002). In a recent European study, low educated mothers were found to remain single more often than in earlier cohorts and compared to college educated women (De Hauw et al. 2015). Selection based on education may therefore increasingly play a role for women's transition to parenthood as well. These findings and trends highlight the relevance of investigating the selection-into-union hypothesis not only for men, as was done in this study, but also for women.

References

Aassve, A., Billari, F. C., Michielin, F., & Panis, C. (2003). A Monte Carlo study of (simultaneous) hazard models with flexible baseline and normally distributed error terms. *Studi Demografici n.14*, Istituto di Metodi Quantitativi, Universita' L.Bocconi, Milano.

Alich, D. (2009). *Fatherhood in Russia between 1970 and 2004: the male perspective of family and fertility behavior in a changing society*. Dissertation, University of Rostock.

Baizán, P., Aassve, A., & Billari, F. (2003). Cohabitation, marriage, and first birth: The interrelationship of family formation events in Spain. *European Journal of Population*, *19*(2), 147–169.

Becker, G.S. (1991). *A treatise on the family*. Cambridge, MA: Harvard University Press.

Begall, K. (2013). How do educational and occupational resources relate to the timing of family formation? A couple analysis of the Netherlands. *Demographic Research*, 29(34): 907-936.

Berrington, A. M., Hernández, M. I. C., Ingham, R., & Stevenson, J. (2005). *Antecedents and outcomes of young fatherhood : Longitudinal evidence from the 1970 British birth cohort study*. Southampton (UK), Southampton Statistical Sciences Research Institute, Applications & Policy Working Paper A05/09.

Billari, Francesco C. and Liefbroer, Aart C (2007). Should I stay or should I go? The impact of age norms on leaving home. *Demography* 44(1):181-198.

Blossfeld, H., & Huinink, J. (1991). Human capital investments or norms of role transition? How women's schooling and career affect the process of family formation. *American Journal of Sociology*, *97*(1), 143–168.

Blossfeld, H.-P., Golsch, K., & Rohwer, G. (2007). *Event history analysis with Stata*. Mahwah Erlbaum Associates.

Brien, M. J., Lillard, L. A., & Waite, L. J. (1999). Interrelated family-building behaviors: cohabitation, marriage, and non-marital conception. *Demography*, *36*(4), 535–51.

Carlson, M. J., VanOrman, A. G., & Pilkauskas, N. V. (2013). Examining the antecedents of U.S. nonmarital fatherhood. *Demography*, *50*(4), 1421–1447.

Cherlin, A. J. (2010). Demographic trends in the United States: A review of research in the 2000s. *Journal of Marriage and Family*, 72(3), 403–419.

Coltrane, S. (2000). Research on household labor : Modeling and measuring the social embeddedness of routine family work. *Journal of Marriage and Family*, *62*(4), 1208–1233.

Corijn, M., Liefbroer, A., & Gierveld, J. de J. (1996). It takes two to tango, doesn't it? The influence of couple characteristics on the timing of the birth of the first child. *Journal of Marriage and the Family*, *58*(1), 117–126.

Corijn, M., Klijzing, E. (Eds) (2001). *Transition to adulthood in Europe*, European Population Series, Kluwer Publishers, Dordrecht.

Craig, L., & Mullan, K. (2010). Parenthood, gender and work-family time in the United States, Australia, Italy, France, and Denmark. *Journal of Marriage and Family*, 72(5), 1344–1361.

Craig, L., & Mullan, K. (2011). How mothers and fathers share childcare: A cross-national time-use comparison. *American Sociological Review*, 76(6), 834–861.

De Hauw, Y., Grow, A., and Van Bavel, J. (2015). *The reversal of gender inequality in education and assortative mating in Europe*. Paper presented at the Annual Meeting of the Population Association of America, April 30-May2, 2015, San Diego, CA.

DiPrete, T. A., & Buchmann, C. (2006). Gender-specific trends in the value of education and the emerging gender gap in college completion. *Demography* 43(1), 1–24.

Eggebeen, D. J., & Knoester, C. (2001). Does fatherhood matter for men? *Journal of Marriage and Family*, *63*(2), 381–393.

Esteve, A., García-Roman, J., & Permanyer, I. (2012). The gender-gap reversal in education and its effect on union formation: the end of hypergamy? *Population and Development Review*, *38*(3), 535–546.

England, P. (2010). The gender revolution: Uneven and stalled. *Gender & Society*, *24*(2), 149–166.

Fokkema, T., Kveder, A., Hiekel, N., Emery, T., & Liefbroer, A. C. (2016). Generations and Gender Programme Wave 1 data collection. *Demographic Research*, 34, 499–524.

Goldscheider, F., & Waite, L. (1986). Sex differences in the entry into marriage. *American Journal of Sociology*, *92*(1), 91–109.

Goldscheider, F., & Kaufman, G. (1996). Fertility and commitment: Bringing men back in. *Population and Development Review*, 22(1996), 87–99.

Goldscheider, F., Bernhardt, E., & Lappegard, T. (2014). Studies of men’s involvement in the family--part 1: Introduction. *Journal of Family Issues, 35*(8), 995–999.

Grow, A., & Van Bavel, J. (2015). Assortative mating and the reversal of gender inequality in education in Europe: An agent-based model, *PLoS ONE* 10(6), 1-24.

Gustafsson, S., & Worku, S. (2006). Assortative mating by education and postponement of couple formation and first birth in Britain and Sweden. In Gustafsson, S.; Kalwij, A. (Eds.), *Education and postponement of maternity. Economic analysis for industrialized countries*, (pp. 259–284). Dordrecht: Springer.

Huinink, J., & Kohli, M. (2014). A life-course approach to fertility. *Demographic Research*, *30*(April), 1293–1326.

Jalovaara, M. (2011). Socioeconomic resources and the dissolution of cohabitations and marriages. *European Journal of Population*,167–193.

Jalovaara, M. (2012). Socio-economic resources and first-union formation in Finland, cohorts born 1969-81. *Population studies, 66*(1), 69–85.

Jalovaara, M., & Miettinen, A. (2013). Does his paycheck also matter? *Demographic Research*, *28*,881–916.

Joyner, K., Peters, H. E., Hynes, K., Sikora, A., Taber, J. R., & Rendall, M. S. (2012). The quality of male fertility data in major U.S. surveys. *Demography*, 49(1), 101–124.

Kalmijn, M. (2011). The influence of men’s income and employment on marriage and cohabitation: Testing Oppenheimer's theory in Europe. *European Journal of Population, 27*(3), 269–293.

Kalmijn, M. (2013). The educational gradient in marriage: a comparison of 25 European countries. *Demography, 50*(4).

Klesment, Martin and Jan Van Bavel. 2015 . *The reversal of the gender gap in education and female breadwinners in Europe*. Stockholm: Stockholm University - Department of Sociology. (FamiliesAndSocieties Working Paper 26).

Kravdal, Ø., & Rindfuss, R. (2008). Changing relationships between education and fertility: A study of women and men born 1940 to 1964. *American Sociological Review*, 73(5), 854–873.

Kreyenfeld, M., & Konietzka, D. (2008). Education and fertility in Germany. In I. Hamm, H. Seitz, & M. Werding (Eds.), *Demographic Change in Germany* (pp. 165–188). Springer Berlin Heidelberg New York.

Lappegård, T., & Rønsen, M. (2005). The multifaceted impact of education on entry into motherhood. *European Journal of Population*, *21*(1), 31–49.

Lappegård, T., & Rønsen, M. (2013). Socioeconomic differences in multipartner fertility among Norwegian men. *Demography* 50 (3), 1135–53.

Liefbroer, A., & Corijn, M. (1999). Who, what, where, and when? Specifying the impact of educational attainment and labor force participation on family formation. *European Journal of Population*, *15*(1), 45-75.

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

Lillard, L. a, & Waite, L. J. (1993). A joint model of marital childbearing and marital disruption. *Demography*, 30(4), 653–81.

Lillard, L. a, Brien, M. J., & Waite, L. J. (1995). Premarital cohabitation and subsequent marital dissolution: a matter of self-selection? *Demography*, *32*(3), 437–57.

Lillard, L. A. and C. W. A. Panis. 2003. *AML Multilevel Multiprocess Statistical Software*, Release 2.0. Los Angeles: EconWare.

Lindberg, L. D., Sonenstein, F. L., Martinez, G., & Marcotte, J. (1998). Completeness of young fathers' reports of fertility. *Journal of Economic and Social Measurement*, 24, 15–23.

Manlove, J., Ryan, S., Wildsmith, E., & Franzetta, K. (2010). The relationship context of nonmarital childbearing in the U.S. *Demographic Research*, 23, 615–654.

Martin-Garcia, T. , Baizan, P. (2006). The impact of the type of education and of educational enrolment on first births. *European Sociological Review*, *22*(3), 259–275.

Martin-Garcia, T. (2009). "Bring men back in": a re-examination of the impact of type of education and educational enrolment on first births in Spain. *European Sociological Review*, *25*(2), 199–213.

McDonald, P. (2000) Gender equity in theories of fertility transition, *Population and Development Review*, 26(3):427-439.

Murphy, M. (2013). Cross-national patterns of intergenerational continuities in childbearing in developed countries. *Biodemography and social biology*, 59(2):101-126.

Oppenheimer, V. (1988). A theory of marriage timing. *American Journal of Sociology*, 94(3), 563–591.

Oppenheimer, V. K. (1994). Women's rising employment and the future of the family in industrial societies. *Population and Development Review*, 20(2), 293–342.

Oppenheimer, V. K. (1997). Women's employment and the gain to marriage : The specialization and trading model. *Annual Review of Sociology*, 23, 431–453.

Panis, Constantijn W.A. (1994). The piecewise linear spline transformation with an application to age at first marriage, *Stata Technical Bulletin* 18: 27-29.

Perelli-Harris, B., Kreyenfeld, M., Sigle-Rushton, W., Keizer, R., Lappegård, T., Jasilioniene, A., Di Giulio, P. (2012). Changes in union status during the transition to parenthood in eleven European countries, 1970s to early 2000s. *Population Studies*, 66(2), 167–82.

Quintini, G., & Manfredi, T. (2009), Going separate ways? School-to-work transitions in the United States and Europe, *OECD Social*, *Employment and Migration Working Papers*, No. 90. Paris: OECD.

Raley, Sara, Suzanne M. Bianchi, and Wendy Wang. 2012. When do fathers care? Mothers' economic contribution and fathers' involvement in child care, *American Journal of Sociology* 117(5): 1422-1459.

Rendall, M. S., Clarke, L., Peters, H. E., Ranjit, N., & Verropoulou, G. (1999). Incomplete reporting of men's fertility in the United States and Britain: a research note. *Demography*, *36*(1), 135–44.

Schwartz, C. R., & Mare, R. D. (2005). Trends in educational assortative marriage from 1940 to 2003. *Demography*, *42*(4), 621–646.

Sobotka, T. (2004). *Postponement of childbearing and low fertility in Europe*, PhD Dissertation, University of Groningen

Sullivan, O., Billari, F. C., & Altintas, E. (2014). Fathers' changing contributions to child care and domestic work in very low-fertility countries: The effect of education. *Journal of Family Issues*.

Sweeney, M. M. (2002). Two decades of family change: The shifting economic foundations of marriage. *American Sociological Review*, 67(1), 132–147.

Tesching, K. (2012). *Education and fertility: Dynamic interrelations between women's educational level, educational field and fertility in Sweden*. Stockholm: Stockholm University Demography Unit (Dissertation Series 6).

Thomson, E. (1997) Couple childbearing desires, intentions, and births. *Demography*, 34(3), 343–54.

Van Bavel, J. 2010. Choice of study discipline and the postponement of motherhood in Europe: the impact of expected earnings, gender composition and family attitudes, *Demography* 47(2), 439-458.

Van Bavel, J. (2012). The reversal of gender inequality in education, union formation and fertility in Europe. *Vienna Yearbook of Population Research*, 1–40.

Vignoli, D., Drefahl, S., & De Santis, G. (2012). Whose job instability affects the likelihood of becoming a parent in Italy? A tale of two partners. *Demographic Research*, 26, 41–62.

Vikat, A., Spéder, Z., Beets, G., Billari, F. C., Bühler, C., Désesque lles, A., Fokkema, T., Hoem, J.M., MacDonald, A., Neyer, G., Pailhé, A., Pinnelli, A., and Solaz, A. (2007). *Generations and Gender Survey: Concept and Design*. In: United Nations (eds.). Generations and Gender Programme: Concepts and Guidelines. Geneva: United Nations: 3–32.

Winkler-Dworak, M., & Toulemon, L. (2007). Gender differences in the transition to adulthood in France: Is there convergence over the recent period? *European Journal of Population*, *23*(3).

Zhang, L. (2011). *Male fertility patterns and determinants* (Vol. 27). Dordrecht: Springer Netherlands.

Appendices: see separate document