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Research Article

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Tamás Bartus

Lívia Murinkó

Ivett Szalma

Bernadett Szél

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Tamás Bartus[1]

Lívia Murinkó[2]

Ivett Szalma[3]

Bernadett Szél[4]

Abstract

BACKGROUND
In recent years, several studies have reported a positive effect of women's education on the transition to second births. This finding contradicts the economic theory of fertility. Three explanations were proposed: the selection, the time-squeeze, and the partner effect hypotheses.

OBJECTIVE
We propose a modification of the economic theory to account for the positive educational gradient with regard to second births. We empirically examine the effect of women's education on the timing of second births.

METHODS
We use a sample of women born between 1946 and 1983 from all three waves of the Hungarian Generations and Gender Survey (GGS) data. We estimate lognormal survival models of the timing of second births.

[1] Associate professor, Corvinus University of Budapest; Research fellow, Demographic Research Institute, HCSO. E-mail: tamas.bartus@uni-corvinus.hu.
[2] Research fellow, Demographic Research Institute, HCSO. E-mail: murinko@demografia.hu.
[3] Researcher, Corvinus University of Budapest. E-mail: iszalma@gmail.com.
[4] Lecturer, Corvinus University of Budapest. E-mail: bernadett.szel@uni-corvinus.hu.

RESULTS
We find that female education reduces the waiting time to second conception in Hungary. The results remain robust after controlling for sample selection and cannot be explained away in terms of time-squeeze and the partner's education.

CONCLUSIONS
We conclude that the relationship between women's education and spacing behavior might be a causal one.

1. Introduction

In recent years, several studies have examined the relationship between education and the transition to second births (Kreyenfeld 2002; Gerster et al. 2007; Kravdal 2001; 2007; Klesment and Puur 2010; Mureşan and Hoem 2010; Billingsley 2011). In most of these studies, women's education was found to have a significant positive effect on the transition to second births. This finding contradicts the economic theory of fertility, which argues that both *quantum* and *tempo* fertility is negatively related to education. The negative quantum effect arises because the shadow price of raising (high-quality) children is relatively high for educated women (Becker and Lewis 1973; Jones, Schoonbroodt, and Tertilt 2008). The negative tempo effect is due to the fact that wages rise with experience and postponement of childbearing minimizes the lifetime opportunity costs of career interruption (Mincer and Polachek 1974; Happel, Hill, and Low 1984; Montgomery and Trussel 1986; Taniguchi 1999).

In order to account for the surprising positive effect of female education on the transition to second births, three explanations were proposed (Kreyenfeld 2002). The selection effect hypothesis states that women with a strong unobserved preference for children are over-represented among those who postpone the first birth, and this unobserved characteristic is responsible for the fast transition to second birth. The time-squeeze hypothesis argues that women who postpone the first birth are closer to the end of their reproductive span, which reduces the waiting time to the second birth. Finally, the partner effect hypothesis states that the positive effect of family income on childbearing, known as the income effect, suppresses the opposite effect of the shadow price of raising high-quality children (Becker and Lewis 1973; Jones, Schoonbroodt, and Tertilt 2008). The argument is supplemented with the empirically realistic assumption that highly educated women tend to marry (or live with) educated men, a phenomenon known as educational homogamy or assortative mating (Becker 1981; Kalmijn 1998). If wages rise with experience, especially among educated men, the postponement of the first child minimizes not only the lifetime foregone earnings of women but also helps the educated partner reach a high income level, reducing the cost of raising high-quality children.

The objective of this paper is to describe and explain the relationship between education and the transition to second birth in Hungary. Our research questions are: (1) Do women with higher education space births closer together? (2) Can the relationship between education and spacing of births be explained in terms of self-selection, time-squeeze, or the partner's education? The relevance of the second question is related to the fact that while the selection, the time-squeeze, and the partner effect hypotheses assume that the relationship between education and spacing of births is a spurious one, there is some evidence (Gerster et al. 2007; Klesment and Puur 2010) that the relationship in question might be a causal one.

The Hungarian setting is interesting for two reasons. First, less effort has been made to examine the relationship between education and the transition to second birth in Central and Eastern European countries, which are infamous for their low level of total fertility (Spéder 2006; Thornton and Philipov 2009; Gerber and Berman 2010; Kapitány and Spéder 2010). Second, previous research into fertility in Hungary found that the fertility of college and university graduates exceeds that of high-school graduates (Husz 2006). A similar pattern was found for the relationship between education and the probability of delivering a second child within a five-year interval after the first birth (Spéder 2006).

The paper is organized as follows. In Section 2, we discuss the implications of the economic theory of birth timing as it relates to the relationship between education and the spacing of births. Since available evidence suggests that the relationship might be a causal one, we examine the mechanisms that might underlie a positive causal effect of education on the transition to second births. Section 3 situates the argument in the Hungarian context. In Section 4, we describe the data and methods we use to examine the relationship between education and fertility. The results of the analyses are presented in Section 5. Section 6 concludes.

2. Spacing of births: An economic perspective

The standard economic theory of the timing of births, also known as the life-cycle model, relies on three assumptions: individuals maximize their well-being over the life-cycle; human capital and thereby wages increase with experience at a decreasing rate; and human capital depreciates while being out of the labor force (Mincer and Polachek 1974; Happel, Hill, and Low 1984; Montgomery and Trussel 1986; Gustafsson 2001). The last assumption implies that the annual wages of mothers should be lower than that of childless women. In other words, mothers pay a wage penalty for career interruption. If human capital rises with experience at a decreasing rate, the wage penalty is especially large if the career is interrupted at a young age (Taniguchi 1999). The postponement of childbearing is therefore a rational solution to the problem of minimizing the wage penalty. Educational differences in postponing behavior can be explained with the additional assumption that the rate at which wages are growing is relatively large in jobs typically filled in with highly educated individuals (Ross 1974). The postponement behavior of highly educated women is an attempt to avoid career interruption at the early stages of career when human capital is accumulated at a relatively fast rate. In contrast, the economic incentives to delay childbearing are much weaker among women with low education: there is no room for wage growth in unskilled occupations, thus the wage penalty is virtually zero for women with lower education.

The life-cycle model is predominantly concerned with the timing of first births. Without further assumptions, the wage penalty argument predicts that mothers are interested in postponing the second birth as well – a prediction which is at odds with previous research findings. This contradiction justifies the emergence of the time-squeeze, the self-selection, and the partner-effect hypotheses. However, the life-cycle model might explain the negative educational gradient if one takes into account the changes in the annual costs of raising children over the life-cycle and economies of scale; that is, the reduction of the costs of childrearing arising from spacing births close together (Newman 1983; Happel, Hill, and Low 1984).

Changes in the costs of raising children. The annual cost of raising a child can be assumed to increase with the age of the child. For simplicity, we assume that the annual cost rises at a constant rate. Moreover, the annual costs might be larger for mothers with higher education, provided that educated parents prefer educated children and the annual costs of raising children are proportional to the investment into their human capital (Gustafsson 2001). Since wages grow at a decreasing rate but the annual costs of raising children grow at a constant rate, the rising annual costs of childbearing works against the postponement of childbearing, especially among women with higher education. The reason is simple: educated women who postpone childbearing do not experience much wage growth upon returning to the labor market, but they do experience a high-level and constant growth in the costs of raising children. Educated women who postpone motherhood might even run the risk of retiring before their offspring would complete the desired level of education. If retired women cannot afford the annual costs of investing in the education of children, postponement of the second child is at odds with the objective of smoothing one's own consumption over the life course (Happel, Hill, and Low 1984). The strategy in question is not acceptable for educated parents since it works against the goal of avoiding the downward mobility of the children.

This extension of the standard wage penalty argument offers the following explanation for the positive educational gradient. When timing any birth, the life-cycle difference between annual earnings and annual losses matters, the latter including the wage penalty and the costs of raising children. When timing the first birth, young educated women face an upward sloping age-earnings profile which induce them to postpone the first birth. Given that the first birth was postponed, educated mothers who re-enter the labor market face a relatively flat age-earnings profile. Since they prefer highly educated children, the spacing decision will be more sensitive to the growth in the annual costs of raising children than to the wage penalty resulting from another career interruption. Educated women therefore find it rational to space births closer together.

Economies of scale. The costs of raising children can be divided into a fixed and a variable component: the size of the former is independent of and the size of the latter is proportional to the number of children. The larger the share of the fixed component, the smaller the increase in the total annual costs of raising more children upon the birth of an additional child.

There are several aspects of child-care activities which exhibit large fixed costs and small, even zero, variable costs. For instance,

> "… one baby-sitter can care for several children, nursery schools often charge less for a second child from the same family, it is easier to make use of hand-me-downs, the mother can chauffeur two children to the same activity as easily as she can one, and so on." (Ross 1974: 35)

The above examples also suggest that the share of fixed costs is large if the children are of similar age: nursery schools would not charge less for a second child if the first one is already in high-school, and parents might need to chauffeur young and older children to different activities. Spacing of births close together, therefore, can substantially reduce the costs of raising children.

The argument from the economies of scale offers the following solution to the problem of why highly educated women might space births closer together. Child care is especially time-intensive when the child is young; it is therefore reasonable to assume that the annual time-costs decrease with the age of the child (Hotz and Miller 1988). Spacing births closer together reduces especially the time costs of raising children. Since the shadow price of leisure is widely assumed to be increasing with education, it is the educated women who can gain the most from spacing births closer together. Note that the economies of scale argument thus reverses the traditional argument that the time costs of raising children force women to postpone motherhood.

3. The Hungarian context

The dramatic decline in fertility in the former socialist countries is often explained in terms of institutional changes that have occurred during the transition to capitalism such as the increase in income inequalities, consumption aspirations and labor market uncertainties, the expansion of higher education, and the decline in free access to childcare (Kantorová 2004; Klasen and Launov 2006; Thornton and Philipov 2009). In this section, we situate the theoretical arguments presented in the previous section in the Hungarian context. First, we present some evidence that education indeed is a proxy for the steepness of age-earnings profiles. Second, we examine the implications of Hungarian family policy for the spacing of births.

Education and wage growth. A central assumption in our paper is that education reflects career opportunities, in general, and the rate at which earnings grow, in particular. The higher the rate at which wages are growing, the larger the incentives to postpone motherhood. The economic literature relies on the stylized fact that age-earnings profiles are relatively steep for educated women and relatively flat for women with poor education. We believe that these educational differences in age-earnings profiles apply to Hungary as well.

In the Appendix we present some evidence on age-earnings profiles of men and women by education and birth cohort. The figures presented in the Appendix clearly show that men and women with higher education face an upward-sloping age-earnings profile after the transition. The graphs show that the largest wage disparities can be found between people with higher education and people with secondary education, especially among men. The wage differentials become especially pronounced during the late 1990s and the 2000s. It is also striking that there is virtually no average wage difference between people with vocational education and people with primary education. People with secondary education enjoy a small wage advantage of about 90 Euro per month over people with lower education. The wage differences between educational groups are affected by gender. On average, men with secondary or even lower education do not earn more than women with similar education. In contrast, college and university educated men born between 1966 and 1975 do enjoy a substantial wage advantage over women with the same educational level. The advantage of men over women seems to diminish as we move from older cohorts to younger ones.

These observations support the theoretical assumption that the steepness of age-earnings profiles increases with female education. They also support the assumption of the partner-effect hypothesis that the partner's education might be more important in fertility decisions than that of women. Among the highly educated, the age-earnings profiles are steeper for men than women. As a consequence, highly educated couples who postpone the first birth are likely to be able to raise two or even more high-quality children.

Hungarian family policy and the spacing of births. The economies of scale argument suggests that spacing births closer together can substantially reduce the time costs of raising children. The Hungarian family policy as well as the institutional constraints on reconciling family and work create additional incentives to space births closer together.

The Hungarian family support system is one of the most generous ones in international comparison if cash benefits are concerned, but falls behind regarding the availability of child-care and part-time or flexible work arrangements. The system of subsidies is complicated; our simplified overview follows Blaskó (2010).[5] Maternity leave is 24 weeks with 70% of the average daily earnings and no ceiling on payments. After maternity leave, mothers and fathers alike are entitled for parental leave until the second birthday of the child. The cash benefits of this parental leave amounts to 70 percent of what the parent earned in the previous year. However, parental leave cannot exceed 140 percent of the minimum wage. This means that the highest available cash benefit amount is close to the average earnings of women with secondary and primary education (see Figure 1 in the Appendix). After parental leave, insured parents are also entitled to a flat rate benefit until the child's third birthday (uninsured parents are entitled to only this amount from the birth of the children until their third birthday). The total length of paid leave is thus 36 months, which is similar to the length of paid leave in the Czech Republic, Poland, and France, but is substantially longer than paid leave in Germany, Denmark, and the Netherlands (Korintus 2010).

The generous cash benefits women receive until the second birthday of the child might compensate for foregone earnings, apart from the wage penalty. However, the real price which mothers have to pay for this generosity is the difficulty of reconciling family and work. In international comparison, Hungarian women have limited access to either child care or flexible forms of employment. Although the vast majority of children aged three to six attend kindergarten, the rate of children aged zero to three attending nurseries is very low, about ten percent in 2006 (Blaskó 2010). This figure is similar to the attendance rate in some former socialist countries like the Czech Republic, Slovakia, and Poland, but substantially lower than the attendance rate in

[5] Our exposition focuses on people with social security insurance; for details, consult Blaskó (2010) and the references cited therein.

Denmark, the Netherlands, and France. The low attendance rate is related to the fact that the capacity of nurseries in the 2000s was half of that in 1990 (Blaskó 2010). In Hungary the proportion of employees who work either part-time or home-based is one of the lowest in Europe (for details, consult Bajnai, Hámori, and Köllő 2009). Not surprisingly, the employment rate of mothers with children aged zero to three are also very low in international comparison (Blaskó 2010). Finally, although the difference between men and women in their rate of labor-force participation is not particularly high in comparison to other European countries (Frey 2011), attitudes towards the importance of women's income and labor force participation exhibit the characteristics of the male-breadwinner model (Blaskó 2005; Szalma 2010).

Generous cash transfers together with limited access to flexible work and nurseries create substantial incentives for mothers to exhaust the three years of paid leave. This leads to a substantial depreciation of human capital and results in a male-female wage gap among qualified employees. Highly educated women therefore are likely to prefer leaving the labor force only once and spacing births close together over experiencing repeated spells of being in and out of the labor force.

4. Data and methods

4.1 The sample

The panel survey *Turning Points of the Life Course* was launched in 2001 and data collection was repeated in 2004 and 2008.[6] The second wave of the survey corresponds to the first harmonized wave of the Generations and Gender Survey (see Spéder 2001 for more information about the Hungarian survey and Vikat et al. 2007 about the GGS). For simplicity, we refer to this dataset as the three waves of the Hungarian GGS. The survey includes retrospective information on fertility, partnership, labor market, and education histories as well as cross-sectional information on the characteristics of partners. The target population included people aged 18-74 in 2001. Individuals were selected using a stratified two-stage sampling procedure: the strata were defined in terms of settlement size and gender and the primary sampling units were settlements. The number of participants dropped from 16,363 to 10,641 between the first and the third wave. Although weights adjusting for dropout are available, the subsequent analyses will not use them.

[6] The first wave of data collection was conducted between November 2001 and March 2002, the second between November 2004 and July 2005, and the third took place between November 2008 and February 2009.

The main sample for the subsequent empirical analyses include 5,890 women who were born between 1946 and 1983. Sample inclusion is independent of the number of completed waves. Since our objective is to examine the relationship between education and the time to second conception, we constructed a dataset of the dates of conceptions, partnership formations, and partnership dissolutions. The sample does not include respondents who got pregnant before turning 14 or respondents with incomplete or inconsistent life histories.[7]

The test of the partner effect hypothesis requires the matching of cross-sectional information on partner's education to retrospective birth histories, provided that partnership status and the identity of the partner is the same at the time of conception and at the time of interview. Out of the 5890, we identified 2830 women who were partnered at the time of the first wave interview and who were partnered to the same partner when exposed to the risks of second conception. We label this subsample the partner sample. Due to its construction, the partner sample is selective. It excludes women who were not partnered when being at risk of a second conception. It also excludes women who were partnered at the time of first birth but the partnership dissolved subsequently, even if the disruption was followed by another cohabitation or marriage. The selective nature of the sample will be taken into account during data analysis.

4.2 Method

We will use event history or survival analysis to examine the effect of respondents' education and partners' education on the transition to second births. The separate modeling of this transition is subject to sample-selection bias: if education has a negative effect on the transition to first birth, education will be positively correlated with unobserved causes of fertility in samples of mothers (Kravdal 2007). For instance, if highly educated people face better career opportunities and therefore postpone first birth, family-oriented values or preferences for children must be, on average, stronger among educated mothers than among mothers with poor education. Otherwise, highly educated women would not enter the sample of mothers. The comparison of the fertility outcomes across educational categories in the sample of mothers therefore measures not only the true effect of education but also the effect of unobserved preferences or personality traits (Kravdal 2001).[8]

[7] There are 6092 women born between 1946 and 1983 in the sample, out of which 202 were dropped.

[8] We are also aware of the endogeneity problem. Partnership formation, for instance, is endogenous since the intention to have children or an unintended conception belong to the main causes of marriage (for evidence based on the GGS project, see Hoem et al. 2009). If the unobserved intention to have children matters, then

In this paper, we will use the Stata module cmp (Roodman 2011) to estimate survival models with sample selection.[9] The cmp module allows one to estimate recursive systems of equations under the assumption that the equation-specific disturbances are correlated and follow a multivariate normal distribution.[10] Because of this distributional assumption, we estimate lognormal survival models. The lognormal model has two distinctive features (Klein and Moeschberger 2003). First, the model is formulated only in the accelerated failure time metric. As a consequence, the dependent variable is the natural logarithm of time to event and not the hazard rate. The second feature is that the hazard rate implicit in the lognormal model exhibits a non-monotonic duration dependence: the implicit baseline hazard first increases then decreases over time. Our choice for the lognormal model can be criticized on the grounds that we impose a specific form of duration dependence on our data instead of estimating it. However, lognormal duration dependence is consistent with the theoretical assumption that the timing of births is a utility-maximizing decision. Since the decision depends on several unobserved and possibly random factors, observed waiting times can be assumed to be normally distributed around the unknown optimal waiting time. Moreover, studies using the piecewise-exponential model often reported observed and baseline hazards which are first increasing then decreasing with process time (Blossfeld, Golsch, and Rohwer 2007; Kreyenfeld 2002; Kulu and Vikat 2007; Mureşan and Hoem 2010; Oláh and Fratczak 2004; Gerster et al. 2007).

4.3 Variables

The dependent variables in our study are related to the waiting time to second conception. Waiting time begins when delivering the first child and is measured in months. As explained above, this paper will model log durations instead of hazard rates. More specifically, we estimate interval-censored regression models on log durations. Interval-censored regression models require two dependent variables, indicating the lower and upper bounds of time intervals. Let t denote the number of months that elapses since the first birth until the second conception or the end of the observation period (censoring date). For uncensored durations, the lower and upper bounds are

partnership dissolution should be endogenous as well. Even education can be endogenous, since preferences for children, which were shaped during socialization, might affect educational decisions (Baizan and Martin-Garcia 2006; Jones et al. 2008). In this paper, we do not address these issues and we assume education and partnership formation to be exogenous.

[9] We are, of course, aware of the widespread use of aML among demographers.

[10] No assumptions are imposed on the variance-covariance matrix of disturbances. Systems involving more than two equations are estimated using simulated maximum likelihood, in general, and the Geweke, Hajivassiliou, and Keane simulator, in particular. For details, consult Roodman (2011).

ln(t-1) and ln(t), respectively, meaning that conception occurred between time points t-1 and t (Lillard and Panis 2003). For right-censored durations, the respective lower and upper bounds are ln(t) and ln(∞); meaning that the event did not occur during the observation period.

The key explanatory variables are the respondent's and the partner's education. Education is measured with the help of four educational levels: primary, vocational, secondary, and higher.[11] Primary education refers to the first stage of compulsory education, which typically begins at age six and traditionally lasts eight years. In this paper, the category "primary education" also includes people who have no formal education or who have incomplete primary education. After completing primary education, children would follow either the vocational, the academic secondary, or the vocational secondary track (by law, enrollment was compulsory until age 18 during data collection). In our analyses, secondary education embraces both vocational secondary and academic secondary education. Traditionally, primary and secondary education lasts eight and four years, respectively, but, during the transition, other forms like the "6+6" and "4+8" year combinations also emerged. Secondary education is completed by passing the Matura (or A-level) exam which is a necessary condition for college or university admission. In contrast, vocational schools do not offer the Matura exam. Throughout this paper, higher education refers to college and university graduates.

For reasons of simplicity, we treat education as a time-constant variable. When the first wave of the Hungarian GGS was administered in 2001, 15 percent of the women in our sample were enrolled in education. We treat this subset as if they had the degree provided by the current studies. On the basis of this assumption, enrolled women are treated as if they completed the current studies and then entered the labor market. For example, secondary education is assigned to students enrolled in secondary schools.

In regression analyses, we will also include indicator variables for birth cohort and several variables related to the age of the mother at first delivery. Birth cohorts are grouped into seven categories, the first including women born between 1946 and 1950 and the last including women born between 1976 and 1983. In order to control for the time-squeeze mechanism, we will include the age at first delivery centered around 30[12], the square thereof, and the interactions between the age variables and the education dummies. Since the fear of reaching the biological limits of fertility must be stronger for those who spend more time in education *and* postpone childbearing, the time-

[11] For a good explanation of the Hungarian educational system, consult Kézdi, Köllő, and Varga (2009).

[12] Subtracting 30 years from age at previous delivery resulted in a substantial decrease in the correlation between the two age variables. Although choosing 30 instead of another number is admittedly arbitrary, 30 is close to the mean age when highly educated women deliver their first child.

squeeze mechanism implies a negative effect of the interaction between higher education and the age at first birth.

Table 1 presents the means and standard deviations of the variables in the full sample as well as in the partner sample. Recall that the latter sample is used to test the partner effect hypothesis. In the full sample, approximately three out of four women have become mothers and half of the mothers delivered two or more children. The probability of giving birth to the second child, conditional upon being a mother, is about 70 percent. The average waiting time to second conception is exp(3.341) = 28.25 months. The sample is relatively balanced in terms of education: the most frequent educational level is secondary (36 percent), the relative frequencies of the other educational levels range between 20 and 22 percent.

Table 1: Descriptive statistics for the main sample and the partner sample

Variables	Main sample (N=5890)		Partner sample (N=2830)	
	Mean	S.D.	Mean	S.D.
Dependent variables				
Exposed to second conception	0.764	0.425	0.983	0.128
Second conception observed	0.542	0.498	0.754	0.431
Lower limit of log waiting time to conception	3.718	1.294	3.669	1.223
Upper limit of log waiting time to conception	3.341	0.866	3.312	0.817
Independent variables				
Partner's education				
not partnered	0.291	0.454	-	-
primary	0.086	0.281	0.128	0.334
vocational	0.288	0.453	0.422	0.494
secondary	0.198	0.398	0.272	0.445
higher	0.137	0.344	0.178	0.382
Education				
primary	0.205	0.403	0.205	0.404
vocational	0.211	0.408	0.240	0.427
secondary	0.357	0.479	0.369	0.483
higher	0.228	0.419	0.186	0.389
Age at first delivery	-6.926	4.024	-6.870	3.711
Age at first delivery squared	64.162	47.439	60.962	43.699

Table 1: (Continued)

Variables	Main sample (N=5890)		Partner sample (N=2830)	
	Mean	S.D.	Mean	S.D.
Birth cohort				
1946-1950	0.138	0.345	0.158	0.365
1951-1955	0.151	0.358	0.171	0.376
1956-1960	0.129	0.335	0.168	0.374
1961-1965	0.108	0.310	0.143	0.351
1966-1970	0.119	0.324	0.142	0.349
1971-1975	0.146	0.353	0.146	0.353
1976-1983	0.210	0.407	0.072	0.258

Note: Age at first delivery is centered around 30.

5. Empirical analyses

The empirical analyses proceed in three steps. First, we present a simple description of the relationship between education and the waiting time to second conception. These analyses ignore the censoring of fertility histories of younger women. In the second step, we correct for this problem by estimating survival models of transitions to conceptions. Finally, we use survival models to examine the partner effect hypothesis.

5.1 The relationship between education and time to second conception

Table 2 presents the means of the waiting times to second conceptions by education and birth cohort. The figures are just raw means and they are not adjusted for censoring. Censoring is substantial for members of the younger cohorts. For instance, the youngest respondents, born in 1983, were only 18-25 years old at the time of the first wave and 25-33 at the time of the third wave of the Hungarian GGS.

Table 2: Mean number of years between first birth and second conception by education and birth cohort

Birth cohort	Primary	Vocational	Secondary	Higher
1946-1950	3.46	3.76	3.60	2.92
1951-1955	3.00	3.19	3.40	2.99
1956-1960	3.25	3.30	3.96	3.19
1961-1965	3.10	3.90	3.07	2.80
1966-1970	3.01	3.36	3.34	2.96
1971-1975	3.06	3.51	3.23	2.66
1976-1983	2.49	3.25	2.70	1.99

The first thing to note is that the average time between first and second births ranges between two and four years. The relationship between the time to second conception and education nevertheless exhibits an inverted U-shaped pattern: within each cohort, women with primary and higher education space births closer together than women with either vocational or secondary education. The fact that it is highly educated women who space births the closest together – the same group which postpones the transition to motherhood – seems to be consistent with the time-squeeze hypothesis. It is also noteworthy that within each educational level, younger cohorts tend to space births closer together than older ones. The reduction of waiting time is about half a year for women with vocational education and one year for other educational levels.

Before generalizing the results from our sample to the population, one should keep in mind that data on waiting times are calculated using a probably selective sample of women. The sample is selective since it includes only mothers, and the factors affecting the probability of becoming a mother is not randomly distributed in our sample. It is well-known that women with higher education wait more to the first conception than their lower educated counterparts. For instance, out of 100 highly educated women born between 1976-1983, only 44 became a mother, as opposed to the 90 percent which characterizes members of the same cohort with primary education. This difference is due to the fact that the highly educated members of this cohort were too young to become mothers; they were only 18-25 years old at the time of the first wave and 25-33 at the time of the third wave of the Hungarian GGS.

As a consequence, unobserved variables which affect childbearing and are correlated with education should be, on average, more pronounced among mothers with higher education than among mothers with lower education. The relationship between education and the waiting time to second conception therefore might be due to the presence of such factors. We will consider this issue in the next subsection.

5.2 Education and the transition to second conception

We now proceed to estimating two duration models for the waiting time to second conceptions. The first one is an interval-censored regression of log failure time to second births using the sample of mothers with one child. The explanatory variables include indicator variables for educational level and birth cohort, the age of mothers at first delivery centered around 30 and the square thereof, as well as interactions between education and the age variables. The interaction terms between educational levels and the age variables allows us to assess the time-squeeze mechanism, which implies that those women who spend more time in education *and* postpone childbearing tend to space births closer together. We have chosen higher education as the reference category. As a consequence, the time-squeeze mechanism implies that the interaction terms between education and age at previous delivery will be positive, meaning that women who spend less time in education and did not postpone the first birth much will wait longer until the second conception.

The first model addresses the issue of time squeeze but it does not address that of sample selection. To minimize selection bias, the conception equation, described in the previous paragraph, is estimated jointly with a selection equation. The dependent variable in the selection equation is an indicator variable marking mothers of one child. The independent variables include education and birth cohort. The link function for the equation is probit. The residuals of the conception and selection equations are assumed to follow a bivariate normal distribution.

Table 3 presents the estimation results. One should keep in mind that coefficients reflect partial changes in log durations instead of changes in hazard rates, and a positive effect on duration is equivalent to a negative effect on the hazard rate. The positive main effects of the education variables in Model 1 thus mean that, compared to women with higher education, women with lower education tend to *postpone* the second birth. In other words, the positive coefficients are evidence for the hypothesis that highly educated women space births closer together than women with lower education.

Compared to Model 1, the coefficients of the same education dummies remain positive as well as statistically significant in Model 2. This finding indicates that differences in spacing behavior across educational levels cannot be attributed to sample selection. Note that the residuals of the conception and selection equations are negatively correlated and the correlation is significant. That is, there are unobserved factors common to both models which have the opposite effect on being a mother and the waiting time to second conception. Given the inverse relationship between hazards and waiting times, the unobserved factors have the same effect on the transition to first and second conceptions. The selection bias resulting from the presence of common unobserved factors is, however, very small, and the results from Model 1 are robust against selection bias.

Table 3: Education and waiting time to second conception

Variables	Model 1		Model 2			
	Conception		Conception		Selection	
Education[a]						
primary	0.893***	(3.60)	0.824***	(3.32)	0.62***	(9.52)
vocational	1.422***	(5.56)	1.349***	(5.28)	0.605***	(10.01)
secondary	0.634***	(4.13)	0.601***	(3.91)	0.24***	(5.00)
Age at first delivery[b]	0.118**	(3.12)	0.118**	(3.12)		
Age at first delivery squared[b]	0.011**	(2.69)	0.011**	(2.69)		
Age at first delivery X …						
primary education	0.048	(0.76)	0.049	(0.78)		
vocational education	0.209**	(3.01)	0.209**	(3.01)		
secondary education	0.015	(0.31)	0.015	(0.31)		
Age at first delivery squared X …						
primary education	-0.008	(1.58)	-0.008	(1.57)		
vocational education	0.005	(0.83)	0.005	(0.84)		
secondary education	-0.006	(1.14)	-0.006	(1.14)		
Birth cohort[c]						
1946-1950	-0.122	(1.26)	-0.355***	(3.47)	1.459***	(20.08)
1951-1955	-0.201*	(2.14)	-0.425***	(4.29)	1.379***	(20.29)
1956-1960	-0.239*	(2.57)	-0.483***	(4.91)	1.571***	(20.73)
1961-1965	-0.354***	(3.69)	-0.607***	(5.87)	1.683***	(19.65)
1966-1970	-0.305***	(3.30)	-0.515***	(5.25)	1.241***	(18.02)
1971-1975	-0.079	(0.86)	-0.219*	(2.30)	0.737***	(12.78)
Constant	4.363***	(35.59)	4.711***	(34.60)	-0.524***	(11.54)
log SD of the residual[d]	0.394***	(26.79)	0.402***	(25.97)	0	
Correlation of residuals[e]					-0.236***	(5.31)
N[f]	4495				5890	

Notes: Coefficients from interval regressions of log durations until conceptions. See the text for details of the estimation method. Numbers in parentheses are *t* statistics. * p<0.05, ** p<0.01; *** p<0.001; two tailed tests assumed.

a The reference category is higher education.

b Age at first delivery is centered around 30. The square of age at first delivery is the square of that centered variable.

c The reference category is 1976-1983.

d The standard deviation of the residual in the selection equation is normalized to unity.

e Correlation coefficients are in fact the Firsher's transformations, calculated as 0.5[ln(1+r) - ln(1-r)], where r is the correlation coefficient.

f Log-likelihoods in Models 1 and 2 are -17466.38 and -19998.66, respectively.

Now we address the question whether differences in spacing of second births across educational categories can be explained in terms of time-squeeze. The inclusion of interaction terms was motivated by this hypothesis. Due to the presence of the interaction terms and centering age at first delivery around 30, the main effects of education measure the effect of education in a population of women who became mothers at age 30. By the same token, the effects of age at first delivery and the square thereof are meaningful in a population of women with higher education. In both models, the coefficient of age at first delivery is about ten times larger than the coefficient of the squared term. Both coefficients are positive and significant. Therefore, there is a U-shaped relationship between the waiting time to second conception and age at first delivery, and waiting time is minimal if the first child was delivered at age 25.[13] Since women with higher education tend to deliver the first child when older than 25, they also tend to wait a longer time until the second birth. The pattern contradicts the time-squeeze hypothesis.

Additional evidence against the time-squeeze hypothesis is that the interaction terms between education and age at first delivery are not significant, with one exception: the interaction between vocational education and age at previous delivery is positive and statistically significant. This means that there is quite limited evidence that education would modify the effect of postponing the first birth on the spacing of births. Nevertheless, the time-squeeze mechanism is at work among women with vocational education. The sum of that interaction term and the main effect of age at first delivery are roughly 30 times larger than the main effect of the squared age term. This implies that women with vocational education would space the births closest together if they were about 30/2+30=45 years old when delivering the first child. Since there is a U shaped relationship between waiting time and age at first delivery, the postponement of childbearing reduces the waiting time among women with vocational education.

To summarize, we found evidence for the positive effect of higher education on the transition to second birth. Women with higher education space births closer together than women with lower education. This effect cannot be explained away with the help of the selection and the time-squeeze hypothesis. The former hypotheses can be rejected on the grounds that the estimates in Models 1 and 2 do not differ substantially, thus the educational effect cannot be attributed to unobserved factors which affect the waiting times to first and second births. The time-squeeze hypothesis is at odds with two findings. First, the relationship between age at first birth and waiting time to second conception exhibits a U-shaped pattern, and the time-squeeze effect applies only to

[13] The minimum is obtained by dividing the coefficient of age at first delivery by -2 times the coefficient of the squared variable. The value of this ratio is about -5. Since the variables are centered around 30, -5 corresponds to 25 on the natural age scale.

women who become mothers at a below-average age. Second, there is only limited evidence that the postponement of first birth would be conditional on education.

Another possible explanation for the positive effect of education on the transition to second birth is the partner effect hypothesis (Kreyenfeld 2002). We now turn to testing this hypothesis.

5.3 The partner effect hypothesis

The partner effect hypothesis was examined using the partner sample (see Section 4.1). Before testing the effect of the partner’s education on spacing behavior, we first examined the extent of assortative mating. The distribution of partner’s education is presented in Table 4. The strength of educational homogamy depends on education: the chances of being partnered to a men with the same education is about 40 percent among women with either primary or secondary education, and about 60 percent among women with either vocational or higher education. The chances of living with highly educated men shows considerable variation. While 60 percent of college or university educated women were partnered to highly educated men, this proportion drops to 15 percent among women with secondary education, and to 2 percent among women with even lower education.

Table 4: Percentage distribution of partner's education in the estimation sample

Partner’s education	Female education			
	primary	vocational	secondary	higher
primary	38.61	14.42	3.71	0.58
vocational	50.36	61.91	39.20	14.42
secondary	8.72	21.40	42.13	25.38
higher	2.31	2.28	14.96	59.62
total	100.00	100.00	100.00	100.00
N	562	659	1023	520

The fact that approximately 40 percent of the college and university educated women married “downwards” is associated with the unequal distribution of educational attainment between the two sexes. There are simply more highly educated women than men, and, similarly, there are more secondary education graduates among women than among men. As a consequence, some of the highly educated women must be married to or live with men with a lower education. For instance, in the first wave of the

Hungarian GGS, the ratio of the number of female college and university graduates to the number of men with the same educational level is three to two in the birth cohorts 1966-1983. This means that perfect educational homogamy cannot be achieved and half of the highly educated women must choose men with secondary education. This leads to an increasing competition among women with secondary education to find partners with higher or the same level of education. The first wave data suggests that the ratio of the number of female competitors to the number of male partners will again be three to two in the marriage market for men with secondary education. Therefore, half of the female competitors would most likely marry men with a lower education. The lesson is that women with tertiary and secondary education are disadvantaged in the marriage market since some of them must choose men with lower education.

We now proceed to examining the effect of partner's education on birth transitions. Again, we specify two models. Model 1 adds the partner's education to the model which was labeled Model 1 in the previous subsection. Model 2 consists of two equations, which are estimated simultaneously. The first equation, labeled Conception, is the same equation as Model 1. The second equation, labeled Selection, is a probit model of being in the partner sample on female education and birth cohort. Similarly to the previous subsection, this equation is included in order to control for unobserved factors which affect both transitions to first and second births. Additionally, this equation also controls for possible sample-selection biases. Recall that the partner sample is not a random sample of retrospective event histories; due to its construction mechanism, it excludes unions that failed to survive until the first wave of data collection. The equations which constitute Model 2 are estimated simultaneously by allowing the equation-specific residuals to be correlated.

The estimation results are presented in Table 5. The coefficients of the two models are quite similar to each other, meaning that the results from Model 1 cannot be attributed to sample selection. The variables capturing the partner's education have positive and significant effects, with the exception of primary education. Since higher education is the reference category, there is evidence that women partnered to highly educated men space births closer together than women partnered to men with secondary or vocational education. However, there is no significant difference in spacing behavior between women partnered to men with primary education and women partnered to highly educated men. In short, there is an inverted U-shaped pattern between the partner's education and the waiting time to second birth.

Table 5: Partner's education and waiting time to second conception

Variables	Model 1		Model 2			
	Conception		Conception		Selection	
Partner's education						
primary	0.090	(0.71)	0.012	(0.10)	0.090	(0.71)
vocational	0.258**	(2.72)	0.221*	(2.49)	0.258**	(2.72)
secondary	0.231*	(2.52)	0.202*	(2.41)	0.231*	(2.52)
Education						
primary	0.638*	(2.09)	0.51	(1.96)	0.048	(0.92)
vocational	1.004***	(3.73)	1.447***	(6.02)	0.308***	(6.16)
secondary	0.441*	(2.42)	0.626***	(3.69)	0.181***	(4.17)
Age at first delivery[b]	0.146**	(3.26)	0.100**	(2.81)		
Age at first delivery squared[b]	0.011*	(2.34)	0.009*	(2.45)		
Age at first delivery X …						
primary education	-0.003	(0.04)	0.006	(0.10)		
vocational education	0.136	(1.92)	0.170**	(2.88)		
secondary education	-0.005	(0.09)	0.021	(0.46)		
Age at first delivery squared X …						
primary education	-0.008	(1.29)	-0.006	(1.27)		
vocational education	0.002	(0.34)	0.004	(0.83)		
secondary education	-0.005	(0.91)	-0.002	(0.55)		
Birth cohort[c]						
1946-1950	-0.280*	(2.22)	1.338***	(8.05)	1.153***	(18.06)
1951-1955	-0.391**	(3.20)	1.213***	(7.47)	1.104***	(18.06)
1956-1960	-0.392**	(3.27)	1.479***	(9.12)	1.281***	(20.36)
1961-1965	-0.551***	(4.56)	1.353***	(8.21)	1.287***	(19.70)
1966-1970	-0.534***	(4.54)	1.212***	(7.64)	1.107***	(17.74)
1971-1975	-0.307**	(2.60)	1.143***	(7.41)	0.907***	(15.14)
Constant	4.352***	(26.33)	1.097***	(5.69)	-1.120***	(23.70)
log SD of the residual[d]	0.312***	(16.26)	0.736***	(26.15)	0	
Correlation of residuals[e]					1.631***	(28.63)
N[f]	2716				5890	

Notes: Coefficients from interval regressions of log durations until conceptions. See the text for details of the estimation method. Numbers in parentheses are *t* statistics. * $p<0.05$, ** $p<0.01$; *** $p<0.001$; two tailed tests assumed.

a The reference category is higher education.

b Age at first delivery is centered around 30. The square of age at first delivery is the square of that centered variable.

c The reference category is 1976-1983.

d The standard deviation of the residual in the selection equation is normalized to unity.

e Correlation coefficients are in fact the Firsher's transformations, calculated as $0.5[\ln(1+r) - \ln(1-r)]$, where *r* is the correlation coefficient.

f Log-likelihoods in Models 1 and 2 are -11124.81 and -14728.69, respectively.

The estimation results for the other variables are virtually the same as the results presented in the previous subsection and in Table 3. Women with higher education space births closer together than women with lower education; since most of the interaction effects lack significance, we have no evidence that age at first delivery would modify the differences across education levels.

Our finding is that female education matters in spacing behavior even after controlling for partner's education. This finding is consistent with those reported in Gerster et al. (2007) and Klesment and Puur (2010). While our study supports the classic income effect hypothesis, it does not support Kreyenfeld's (2002) partner effect hypothesis that the relationship between female education and the transition to the second birth is spurious and is due to educational homogamy and the effect of male education. Given the available evidence, the relationship between women's education and spacing behavior seems to be a causal one, and the partner effect hypothesis describes another mechanism which strengthens the correlation between higher education and the transition to second births.

6. Conclusions and discussion

In recent years, several studies have reported a positive relationship between education and the transition to second birth (Kreyenfeld 2002; Gerster et al. 2007; Kravdal 2001, 2007; Klesment and Puur 2010; Billingsley 2011). This finding seems to contradict the economic theory of fertility since the shadow price of raising (high-quality) children as well as the foregone earnings from career interruptions should be higher for educated women (Becker and Lewis 1973; Mincer and Polachek 1974; Happel, Hill, and Low 1984; Montgomery and Trussel 1986; Taniguchi 1999; Jones, Schoonbroodt, and Tertilt 2008). In order to account for the surprising positive effect of female education on the transition to second births, three explanations were proposed (Kreyenfeld 2002). The selection effect hypothesis states that women with a strong unobserved preference for children are over-represented among those who become mothers at a relatively late age, and this unobserved characteristic is responsible for the fast transition to second birth. The time-squeeze hypothesis argues that women who postpone the first birth are closer to the end of their reproductive span, which reduces the waiting time to the second birth. Finally, the partner effect hypothesis states that the positive effect of family income on childbearing, known as the income effect, suppresses the opposite effect of the shadow price of raising high-quality children (Becker and Lewis 1973; Jones, Schoonbroodt, and Tertilt 2008). The argument is supplemented with the empirically realistic assumption that highly educated women tend to marry (or live with) educated men, a phenomenon known as educational homogamy or assortative mating (Becker 1981; Kalmijn 1998).

In this paper, we examine the relationship between education and the transition to second birth in Hungary, using three waves of the Hungarian Generations and Gender Survey (GGS). We estimate lognormal survival models which also control for both the time-squeeze and the selection hypotheses (Kreyenfeld 2002). Our results are as follows. First, women with higher education space the first and the second births closer together than women with lower education. This result is consistent with earlier results about the positive educational gradient concerning the transition to second and higher order births. Second, the education of the partner has the expected effect: women partnered to college or university educated men space births closer together than women partnered to men with either vocational or secondary education. However, the effect of female education remains significant, thus the effect of female education cannot be explained away in terms of the partner's education. This finding is consistent with those reported in Gerster et al. (2007) and Klesment and Puur (2010), but it does not support the partner effect hypothesis. The relationship between women's education and spacing behavior is likely to be a causal one, and the partner effect hypothesis describes another mechanism which strengthens the correlation between higher education and the transition to second births.

In our paper, we make an attempt to modify the economic explanation for the timing of births; we have argued that women with higher education might be more interested in spacing births closer together than women with lower education. Unfortunately, we did not test this theory by deriving and examining new hypotheses. The support for the theory comes in the form of rejecting some alternative explanations, namely the self-selection, the time-squeeze, and the partner effect hypotheses. However, there are more alternative explanations which might account for the observed patterns. It might be the case that the education of the partner has a positive effect on marital stability: partners with low education lack the resources to support the family, or highly educated men might be more involved in housework than their counterparts with lower education. Finally, spacing births closer together might simply reflect a desire for a large family. We did not consider values and attitudes which might shape fertility intentions. Future research should offer a more detailed description of the mechanisms between education and the spacing of births.

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Appendix

In the Appendix, we present some evidence on the age-earnings profiles of men and women by education and birth cohort. We use the 1994-2008 waves of the Hungarian Wage and Earnings Survey (WES). The WES is collected annually, the sample includes 150,000-217,000 employees, depending on the year. (For a detailed description of the sampling design, consult Lovász 2008). An important feature of the survey is that no individuals are interviewed: all data on individual wages and personal characteristics are provided by the representatives of the sampled firms. Using the individual-level data, we created a quasi-panel dataset. There are 24 groups in the quasi-panel dataset, defined as the combinations of four educational levels, three birth cohorts (1966-1970, 1971-1975, and 1976-1983) and the two sexes. For each group, average net wages are observed in the period 1994-2008. The nominal earnings figures were transformed into real values using price-index data published by the Hungarian Central Statistical Office, the base year being 2011.

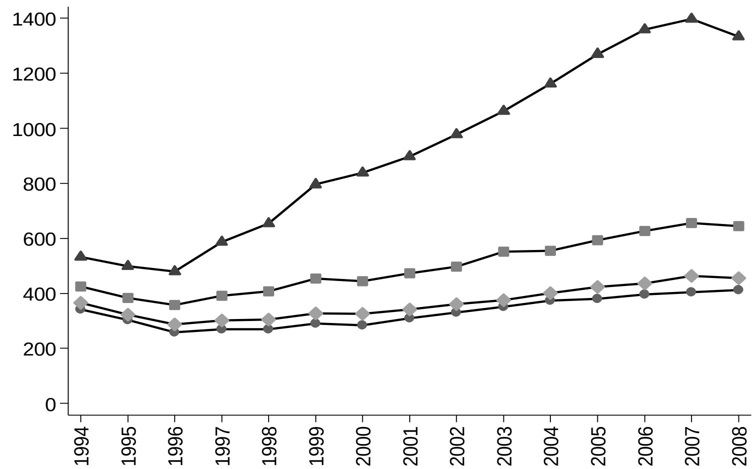
Figure 1 shows the real value of net monthly earnings in Euro by educational level for three birth cohorts and the two sexes separately.[14] In 2011, the net minimum wage was about 280 euros per month, thus each 200 unit on the Y axis represents about 70 percent of the minimum wage. The graphs show that the largest wage disparities can be found between people with higher education and people with secondary education, especially among men. The wage differentials became pronounced during the late 1990s and the 2000s. It is also striking that there is virtually no average wage difference between people with vocational education and people with primary education. People with secondary education enjoy a small wage advantage of about 90 Euro per month over people with poor education.

Wage differences between educational groups are affected by gender. On average, men with secondary education or less do not earn more than women with similar education. In contrast, college and university educated men born between 1966 and 1975 do enjoy a substantial wage advantage over women with the same educational level. The advantage of men seems to diminish as we move from older cohorts to younger ones.

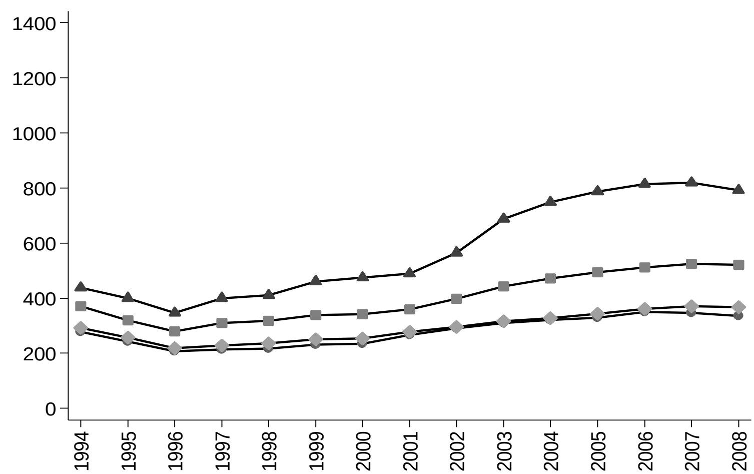
[14] We assumed that 1 Euro equals 280 HUF.

Figure 1: Net real wages in Euro by educational level, gender and birth cohort, 1994-2008

Men born 1966-1970



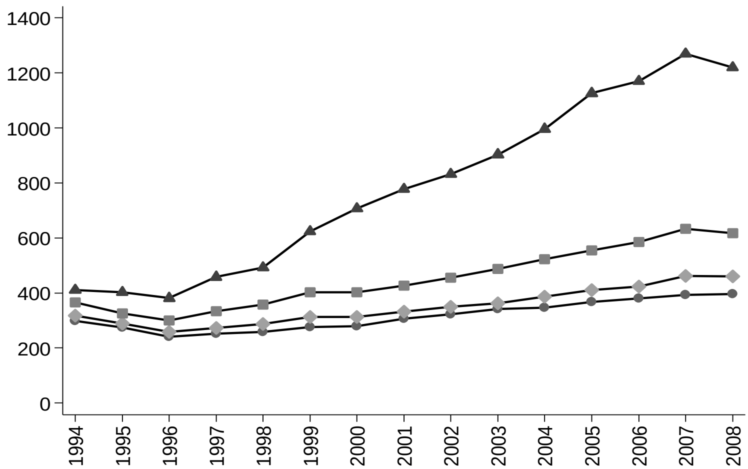
Women born 1966-1970



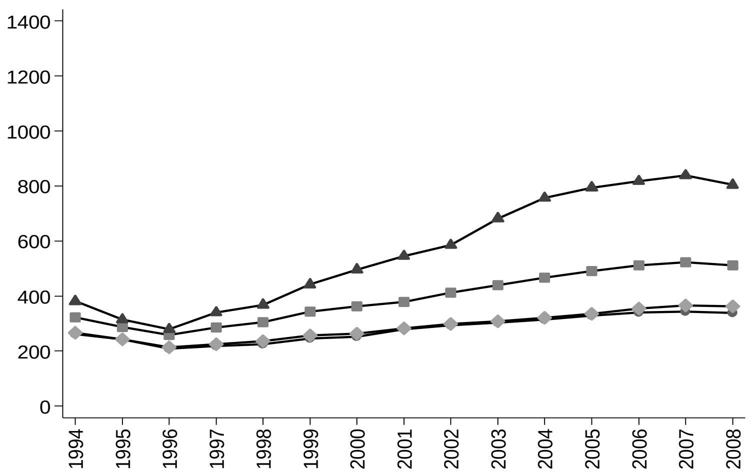
Legend: triangle = higher education, square = secondary education; diamond = vocational education; circle = primary education

Figure 1: (Continued)

Men born 1971-1975



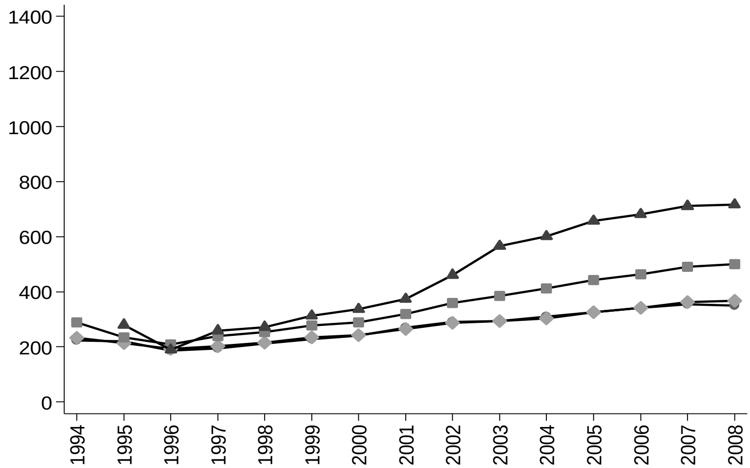
Women born 1971-1975



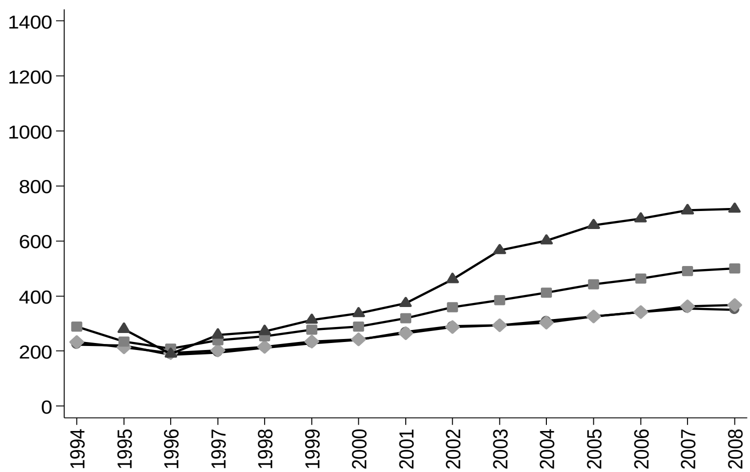
Legend: triangle = higher education, square = secondary education; diamond = vocational education; circle = primary education

Figure 1: (Continued)

Men born 1976-1983



Women born 1976-1983



Legend: triangle = higher education, square = secondary education; diamond = vocational education; circle = primary education