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Analysis of the determinants of fertility decline in the Czech Republic

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Abstract In this paper, we analyze the decline in the total fertility rate (TFR) in the Czech Republic during the economic transition. To identify transition-specific features of this decline, we estimate a Heckman–Walker multistate model of the birth process using data from the 1998 Family and Fertility Survey. We find that the negative effect of transition on TFR is mostly driven by a sharply increased influence of higher education, limited ability to combine employment with childbearing and lack of adequate childcare facilities. We also detect a significant role of the increased use of contraception, motivated by both economic and demographic reasons.

Keywords Economic transition · Fertility decline · Czech Republic · Multistate model of birth process

JEL Classification J13 · J11

1 Introduction

The large fertility decline in the transition countries of Central and Eastern Europe is a well-known phenomenon in contemporary population economics. Fertility rates in most countries of Central and Eastern Europe have declined from close to replacement level to levels far below that, ensuring that populations in Eastern Europe will shrink considerably in coming decades. Increased mortality in some countries and rising emigration further accelerate this population decline. According to Sobotka

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(2001, 2004), it took just 10 years for the post-communist countries as a group to move from the region with the highest fertility levels in Europe to the one with the lowest fertility levels.

A series of research papers has attempted to classify this phenomenon and look into its possible sources. In particular, Sobotka (2001) identifies two different scenarios that depend on the success of the transition process. He shows that in relatively successful countries, fertility decline is characterized by a strong postponement of the first birth and a decline in teenage fertility rates, which implies a change to a new reproductive regime that is more similar to Western European countries. In less successful countries, fertility decline occurs with moderate postponement of the first birth and higher teenage fertility—the features of the old reproductive regime, although with higher exit rate after the first birth. Billari and Kohler (2002) complement this study by demonstrating that the trends toward lowest-low fertility in Central and Eastern European countries are accompanied by relatively early household independence and union formation. The authors also point to a large diversity in the evolution of fertility patterns in the European countries, arguing that despite the commonality of the total fertility decline, transition countries have their own distinct paths in this process and will exhibit little convergence with the rest of Europe in the near future.

According to Sobotka (2001), the evolution of fertility patterns in the Czech Republic has proceeded along the first scenario. This also conforms to the findings of Billari and Kohler (2002), who additionally find low progression probability once the first child is born. The results of Sobotka (2001) for the Czech Republic are further supported by Kantorová (2003), who shows that during the transition, education acquires an increasing role in the postponement of the first birth. Finally, Mašková and Stašová (2000) provide evidence of the changes in contraception use and point to the end of the "abortion culture" of the socialist period. Burcin and Kučera (2000) demonstrate that the decline of the total fertility is higher if the woman lives in an urban area.

Reviewing this literature, one notices, however, that the analytical methods used in the above studies mainly consist of descriptive statistics and comparisons (with Kantorová 2003 being the only exception). Although this descriptive approach provides a very good illustration of the phenomenon, it lacks formal structure, it is not linked to explicit models of fertility behavior, and thus, it does not allow to assess the impact of transition on the determinants of fertility such as education, employment childcare facilities, and the like. Furthermore, given still incomplete birth histories of the youngest women, who are in fact those most affected by the transition, it is critical to use methods that include their yet incomplete birth histories in an assessment of fertility behavior. The present paper aims at filling these gaps. In our study, we use an approach that formalizes the birth process as a probabilistic model, in which frequency and number of births depend on a set of individual characteristics and environmental/institutional features.

To the best of our knowledge, three different statistical frameworks are at our disposal. First, there are count data models that specify the number of born children as a point process with nonnegative outcomes. Among applications of this class of regressions in population economics research, there are papers of Famoye and Wang (1997), who estimated a negative binomial model, and Covas and Santos Silva (2000), who worked with a modified hurdle extension of the preceding specification. The second framework is a "sequential probability model" of Barmby and Cigno

(1990). The approach models the birth history as a discrete time process with a binary "birth"– "no birth" outcome at each point in time. Finally, the third framework is a multistate model of the birth process developed by Heckman and Walker (1987, 1990a,b). It is a discrete space continuous time model which concentrates on jointly modeling waiting times between adjacent births and progression probabilities once the birth is realized.

To study the transition-specific features of the fertility decline, we rely on a multistate model of birth process. Several reasons justify this choice. Firstly, in the continuous time environment, it becomes much easier to model the dependence of the next birth on the previous birth history and introduce unobserved heterogeneity. Secondly, we will be able to study the transition-specific changes in both the timing of births (tempo effect) and the probability of early exit from childbearing (quantum effect), which is not possible in the two alternative settings above. Having a possibility to analyze the quantum effect is especially important in view of the results of Sobotka (2004), who shows it to be a significant determinant in the contemporary trend toward lowest-low fertility in Europe. Finally, the chosen specification easily handles incomplete birth histories.

The rest of the paper is organized as follows. In the second section, we present the problem of declining fertility in the Czech Republic, show its specific features, and state our goals. The next section contains a description of the data used to explain the driving forces of fertility decline. In the fourth section, we describe the Heckman–Walker statistical model of birth process. In this section, we also discuss the estimation strategy. Section five contains the discussion of the estimation results and possible policy implications. The last section summarizes our main findings.

2 Fertility decline in the Czech Republic

Like the rest of the transition countries, the Czech Republic has undergone a massive fertility decline in the past decade. Although the country-specific trends of this decline seem to vary (see Billari and Kohler 2002), there still exist at least two facts that are uniformly true for all transition countries. The first one is a reduction in the total fertility rate (TFR), and the second one is a postponement of the first birth. In what follows, we will refer to them as "stylized facts" in the contemporary evolution of fertility patterns in Central and Eastern Europe.

An indication of the first stylized fact for the Czech Republic can be found in Billari and Kohler (2002) and Sobotka (2001) among many others. As to the postponement of births, Kantorová (2003), Mašková and Stašová (2000), and Sobotka (2001, 2004) also present convincing evidence of the phenomenon. To demonstrate the magnitude of the fertility decline and birth postponement observed in our data, we also calculate a TFR (see Fig. 1) and the mean age at the first birth (Fig. 2) during the 1988–1997 period. As expected, our figures also show a drastically falling TFR and a sharply rising mean age at first birth. From Fig. 1, we see that it took only 10 years for TFR to fall from 2.25 to 1.25 children per woman. Simultaneously, Fig. 2 reveals an upward trend in the postponement of the first

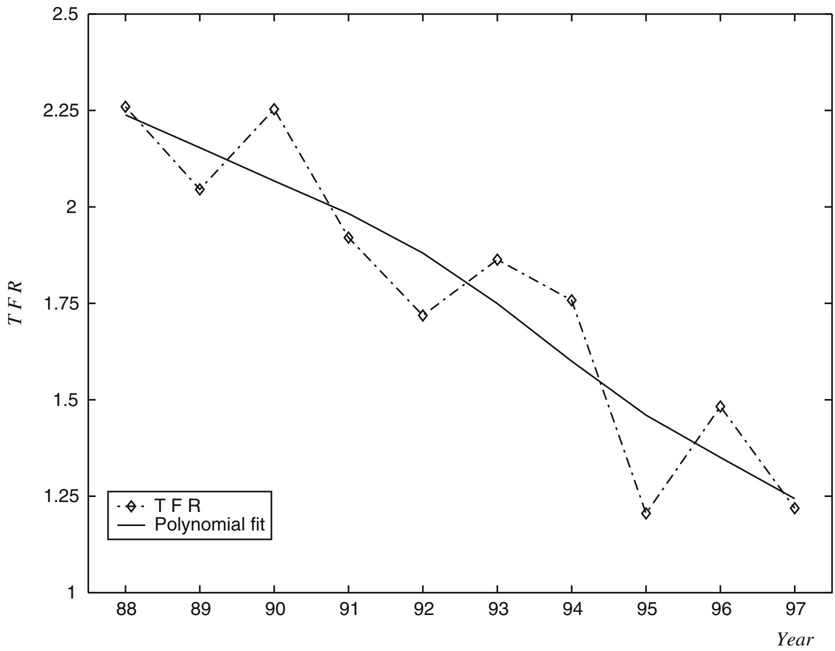

Fig. 1 Total fertility rate in the Czech Republic in 1988–1997. *Source*: UNECE (2003), own calculations

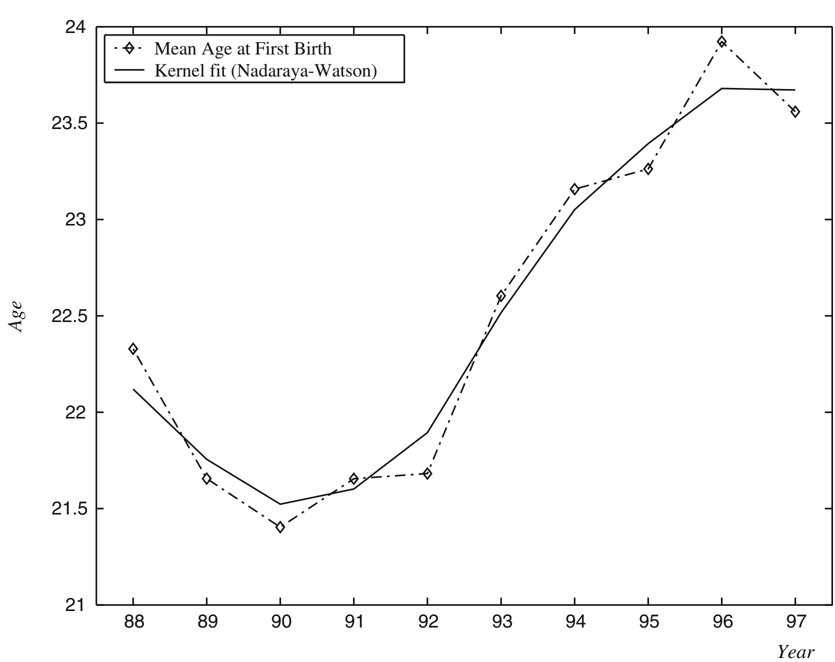

Fig. 2 Mean age at first birth in the Czech Republic in 1988–1997. *Source*: UNECE (2003), own calculations

birth. Remarkably, despite containing the simplest descriptive statistic measures, Figs. 1 and 2 convey the same patterns as those obtained from the application of more comprehensive ones.[1]

The illustration above leaves no doubt that the transition brought about serious changes to the reproductive behavior of the population in the Czech Republic. These changes may be either temporary or permanent (Sobotka 2004). Intuitively, the reason for the temporary fall of TFR could be the increased importance of a consumption-smoothing motive in a more uncertain transition environment (see Hotz et al. 1997). Permanent changes may also be influenced by economic considerations such as, for instance, stronger career planning motives (Hotz et al. 1997) brought about by the increasing necessity to invest in human capital (Gustafsson 2001). Additionally, demographic reasons such as emerging new modes of fertility regulation or changes in views on partnership and cohabitation may come into play. Finally, changes of both types can be triggered by the transforming social environment and the replacement of old social values.

In this study, we investigate the nature of the changes to the reproductive behavior during the transition in the Czech Republic. Within a set of individual characteristics that contains both economic and demographic determinants, we seek to determine the channels through which the ongoing socioeconomic changes negatively affect fertility decisions. To do so, we first learn about the effect of any given determinant during the socialist time and then check for the changes of this effect during transition.

3 Data

We used the data from the Family and Fertility Survey (FFS) of 1998, undertaken with support from the UN Economic Commission of Europe (see UNECE, 2003). It is a cross-sectional survey with detailed information about fertility and family formation behavior of women between 15 and 45 years old. We assume that for each woman, the reproductive age begins at the time when she turns 16. Since one and the same covariate can have different effects on fertility decisions of a woman at different ages, we define three birth cohorts in the sample. The first cohort includes women who were 16–26 years old in 1998, the second cohort includes those who were 27–35 years old, and the third one includes women aged 36–44. With this division, we note that birth decisions in the second and third cohorts were mostly determined by conditions prevailing under socialism. The first cohort, however, came of age after the transition process had begun, and their birth decisions are already influenced by the transition.

First, we restore the birth history of every woman. From the data on the timing between births (see Table 1), we can see that for about 30% of all women in the first cohort, there has been no first birth by 1998. Furthermore, for almost 64% of all women in this cohort, we cannot observe the second birth, and there is a negligible number of three or more births (naturally, for older cohorts, the problem of incomplete birth histories is less severe). This implies that in general, we cannot make any

[1] For richer summary statistics calculating age- and cohort-specific fertility rates, refer to Billari and Kohler (2002). An overview of tempo-adjusted Bongaarts–Feeney rates can be found in Sobotka (2001).

Table 1 Descriptive statistics of the variables used in the study

Variable	First cohort (born 1972–1982) Mean (SD)	Second cohort (born 1963–1971) Mean (SD)	Third cohort (born 1954–1962) Mean (SD)
Time spell between			
Zero and one birth	4.59 (2.42)	6.03 (3.64)	6.43 (4.58)
Percentage of incomplete observation	30.97	6.36	2.29
One and two births	2.92 (1.85)	4.41 (3.26)	5.21 (5.31)
Percentage of incomplete observation	63.78	27.60	14.99
Education (years)	10.72 (1.60)	11.56 (2.32)	11.30 (0.67)
Employment intensity[a]			
Between zero and one birth	2.33 (1.02)	2.32 (1.06)	2.53 (1.21)
Percentage of nonparticipants[b]	15.30	14.71	10.53
Between one and two births	3.08 (1.36)	3.27 (1.11)	3.56 (0.93)
Percentage of nonparticipants[b]	66.79	38.17	23.34
	Percentage of the sample		
Rented dwellings	0.3657	0.4612	0.5309
Urban area	–	0.1292	0.1373
Partner education-medium	0.4142	0.3956	0.3341
Partner education-high	0.0746	0.1431	0.1419
Beliefs about the negative effect of			
Housing conditions	0.4664	–	–
Childcare facilities	0.1343	–	–
Desire for personal advancement	0.1119	–	–
Increased availability of contraception	0.1343	–	–

[a] Excludes nonparticipants and those who have not been employed between births at least once
[b] Includes those who are permanently unemployed

sensible inference for the number of births higher than two. However, given that TFR fell from slightly above 2 to about 1.3, most of the fertility change during the transition will be driven by a greater postponement of the first birth and the increasing number of childless and single-child families. Therefore, the available data should still give us a good grasp of the fertility decline phenomenon.

To explain the variation in birth histories, we use a set of variables that can be divided into two groups, which we call socioeconomic and belief variables.

The first group contains socioeconomic characteristics of a woman such as education level, employment history, housing ownership, and place of residence. The education variable represents years of formal education. It is constructed using the information on the length of education spells available in the survey.[2] Additionally, we use data on the partner's education. Since education spells are available only for the respondent, partner education is represented by two dummy variables that correspond to medium (post-secondary) and high (graduate) levels of attainment, respectively. The employment variable is a summary of employment intensity (average hours of work per week) between any two adjacent births. This variable is constructed with the help of three components—the length of job spells between births (T_j), the indicator of hours per week spent at each of these jobs (I_j), and the whole length of waiting time between births (T). The variable represents a weighted average of weekly working hours between two adjacent births, where weights are T_j/T, and the weighted variable is a discrete indicator, I_j. I_j is "0" when weekly hours were zero, "1" when weekly hours were less then 10 and so forth up to "5" if a woman worked more than 45 h per week. Due to the discreteness of I_j, the resulting "employment" variable is a quasi-intensity, although it should represent the actual work intensity fairly well. Ownership refers to a dummy variable, which is "1" if the dwelling is rented by a couple. Finally, the residence indicator is a dummy that equals to 1 whenever the responding woman lives in an urban area (more than 100,000 inhabitants).

Income data are not available in the survey for either respondent or partner. To remedy this problem, the selected Heckman–Walker model accounts for unobserved changes in real income over time (see Section 4 for the discussion of correlation of unobservables across birth spells).

The group of belief variables contains variables that reflect respondents' opinion on possible sources of the fertility decline. However, belief questions are not specifically asking about the respondents' own fertility.[3] The question is rather set as "... women nowadays have fewer children than in previous generations. Do you think that the following circumstances have played [...] a role?" After that, a list of reasons follows, and the respondent is to assign "1" to a reason which she finds very important ("0" otherwise). To ensure that these general beliefs correctly reflect one's own beliefs (and thus can be treated as a determinant of one's own fertility decisions), we do not use the individual response as the covariate. Instead, for each respondent, we define a reference group and use the mean importance of the belief for the group, arguing that beliefs prevalent within the group largely affect the formation of respondents' own beliefs and thus affect her own fertility decisions. We propose constructing such reference groups on the basis of residence up to the age of 15 and current residence. Former and current residence indicators consist of five different levels that represent population density (less then 2,000 inhabitants, 2,000–9,999 inhabitants and so on). Altogether, this implies 25 possible groups. Intuitively, a woman who lived in a rural area before 15 and still lives in this area observes mostly the other rural dwellers who did not migrate. This "localizes" the information set on which this woman forms her own views on having children, so as a result, her personal views are not expected to be sharply different from the views common to this group. Having calculated the mean importance of the belief for each group, we

[2] To avoid measurement error bias in the first cohort, we consider only those females who have completed their education. This simultaneously reduces the efficiency loss due to censoring.

[3] There are also some questions about the reasons for own fertility decisions, but very high nonresponse rates render these data useless.

also notice the considerable between-group variation of this variable—for all the "localized" variables, the standard deviation is almost as high as the mean.

When selecting the reasons for the contemporary fertility decline, we follow the concept of Sobotka (2001) of the "socialist greenhouse environment" and the dissolution of "socialist greenhouse" during the transition. With the "socialist greenhouse" Sobotka (2001) refers to the effectively pronatalist policies of the socialist governments, few opportunities for personal advancement, low returns to education, low access to consumption goods, and poor access to contraception, all of which facilitated women to have children. In brief, Sobotka (2001) considers five domains of this phenomenon: "1. Education," "2. Career," "3. Social security support and population policy," "4. Private and family life, leisure, lifestyle and quality of life," and "5. Reproduction and intimacy," suggesting that with the onset of transition, these have undergone the most crucial changes. On the basis of this classification, we pick up the most important feature from each domain (excluding education, which we already dealt with) and get the four essential belief variables that explain fertility decline:

1. Poor housing conditions (social security and population policy domain)
2. Insufficient childcare facilities (social security and population policy domain)
3. Growing desire among men and women for independence and personal advancement (private, family life, and lifestyle domain; career domain)
4. Increased availability of contraception (reproduction and intimacy domain)

We also notice that belief variables are the opinions of the people on the date of sampling, which means that they must be important for the current birth decision only. Our analysis is restricted to the first two births. At the same time, the majority of those who belong to the second and third age cohorts have already had their first and second births by the date of sampling. Therefore, in these two cohorts, belief variables are relevant only for the third birth and more. Since these lie beyond the scope of our analysis, we do not use belief variables when estimating the model for the second and third cohorts. Given the way we personalize beliefs, we also exclude the residence indicator from the set of explanatory variables for the youngest cohort.

Summary statistics of explanatory variables are reported in Table 1. One fact about employment intensity after the first birth needs to be mentioned here. Although the mean intensity across cohorts seems to be roughly the same, the number of nonparticipants during the transition is drastically higher (so in fact, employment intensity calculated, including nonparticipants and the permanently unemployed, in this cohort is considerably smaller). Furthermore, while nonparticipation in the first cohort amounts to almost 67%, only around 40% are within the limits of their statutory maternity leave. The rest are either nonparticipants on extended maternity leave or unemployed. This is likely to be related to increasing unemployment in the transition period, particularly for females (Klasen 1993) and the growing incompatibility between child-rearing and full-time employment in the transition period, where childcare facilities have fallen away or became much more costly.

Regarding the belief variables, surprisingly, the housing conditions appear to be the most important stated reason for the decline in fertility. Lacking childcare facilities, desire for personal advancement, and increased contraception are also cited by a minority of respondents, although to see a clearer picture, one should rather address the group-specific measures. The regularity here is that, first, mean importance of all beliefs increases with the population density of the current location. Furthermore, for each current location, the importance of the beliefs is positively correlated with the population density of the previous location. Second, for those who have either one or no realized birth, the values of all beliefs are higher than for the others.

4 The model and estimation technique

4.1 Statistical model of birth process

To formalize the birth process, we follow Heckman and Walker (1987, 1990a,b). The birth process is suggested to be a continuous time stochastic process, $\Im$, with a finite set of possible realizations $\{0, 1, 2, ..., J\}$. The process starts at a certain calendar date, $\tau(0)$, after which, events occur at subsequent calendar dates $\{\tau(j)\}_{j=1}^{J}$, spaced by time intervals of random length $\{t_j\}_{j=1}^{J}$. Each waiting time between two adjacent births, t_j, is a realization of a random variable, T_j, with probability measure $G(t)$ and density function $g(t)$ defined on the set $(0, \infty)$. Furthermore, at each point in time $\tau(j-1)+t_j\}$, there exists an information set $H(\tau(j-1)+t_j)$ based on which, the decision of having a(nother) child is made. In general, these information sets may not be perfectly observed by the analyst.[4] Assume that the unobserved component is a random variable Θ with a set of possible realizations $\{\underline{\Theta} : \Theta = \theta\}$ and probability measure $M(\theta)$, supp $(M(\theta))=\Theta$. Then given the respective information set $H(\tau(j-1)+t_j)$ the density of a time spell $g(t_j \,|\, H(\tau(j-1)+t_j),\theta)$ will imply a marginal density

$$g(t_j|\mathrm{x}) = \int_{\underline{\Theta}} g(t_j|H(\tau(j-1)+t_j),\theta)dM(\theta),$$

which conditions exclusively on the observables.

Dependence of information sets Despite taking account of unobserved heterogeneity, the formalization presented above overlooks the fact that conditioning sets may be dependent on the prior birth history. In real life, this is the case when the probability of having the second child given the first one is not the same as the probability of having the first child. The differences may arise, for instance, due to acquired experience in childbearing, use of contraception, or the fact that one has already reached or is close to reaching the desired family size.

To introduce this feature into the model, Heckman and Walker (1990a,b) suggest that unobserved heterogeneity (as seen from the researcher's side) can be split in two parts, the first of which is observable for the woman after the occurrence of the event and the second is unobservable for both. In the example above, contraception, childbearing experience, and having reached the desired family size will fall into the first category. To the second category, we can attribute, for instance, the woman's fecundity.

[4] This assumption is particularly appropriate in our case since we do not have income data, and individuals are surely heterogeneous with respect to income.

Such a division of the unobserved component implies that for any number of realized births, j, the distribution of waiting time until the j+1−th birth $g(t_{j+1} \mid H(\tau(j)+t_{j+1}),)$ will not be the same as the distribution of waiting time until the previous one because when the j−th birth obtains the unobserved heterogeneity, distribution $m\,(\mid H(\tau(j-1)+t_j))$ becomes revised. Invoking Bayes' theorem, Heckman and Walker (1990b) obtained

$$m\left(\theta \middle| H(\tau(j-1)), t_j\right) = \frac{g\left(t_j \middle| H\left(\tau(j-1)+t_j\right), \theta\right) m\left(\theta \middle| H\left(\tau(j-1)+t_j\right)\right)}{g\left(t_j \middle| H\left(\tau(j-1)+t_j\right)\right)}.$$

Under the assumption that $m\,(\theta \mid H(\tau(0)+t_1))=m(\theta)$, repeated application of the above Bayesian updating to $m(\theta \mid \cdot)$ starting from the entry date $\tau(0)$ leads to a distribution

$$\begin{aligned} f(t_1, \ldots, t_K) &= \prod_{j=1}^{K} g\left(t_j \middle| H\left(\tau(K-1)+t_j\right)\right) \\ &= \int_{\underline{\Theta}} \left(\prod_{j=1}^{K} g\left(t_j \middle| H\left(\tau(K-1)+t_j\right), \theta\right)\right) m(\theta) d\theta \end{aligned} \tag{1}$$

which describes the entire birth history of a woman with K realized births given a correlation of unobservables across spells.

Stopping of childbearing The problem of declining fertility is not only a question of birth postponement formalized in Eq. 1 but also a question of early exit from childbearing. To integrate early exit into the model, Heckman and Walker (1990a,b) introduced an additional parameter that places a positive probability mass on the event of no next birth. First, rewrite Eq. 1 in terms of the survivor function, denoted by $S(t_j \mid H(\tau(j-1)+t_j))$:

$$f = \int_{\underline{\Theta}} \left[\prod_{j=1}^{K} \frac{\partial}{\partial t_j}\left(-\ln S\left(t_j \middle| H\left(\tau(j-1)+t_j\right)\right)\right) S\left(t_j \middle| H\left(\tau(j-1)+t_j\right)\right)\right] m(\theta) d\theta \tag{2}$$

Then, assume that $p^{(l)}$ fraction of the current group of women will not continue childbearing after l-th birth, $l \in [1, ..., K]$. Since exiting from childbearing is equivalent to the hazard being zero, the survivor function can be represented as a mixture

$$S^{*}(t_{l+1} | H(\tau(l)+t_{l+1})) = p^{[l]} + \left(1-p^{[l]}\right) \exp\left\{-\int_{0}^{t_{l+1}} h(u | H(\tau(l)), \theta) du\right\} \tag{3}$$

where $p^{(l)}$ is a probability that $\Theta=-\infty$, which sets the hazard function $h(u \mid H(\tau(l)), \theta)$ to zero. Thus, for any woman, the probability of surviving the next time spell t_{l+1} is either the probability of stopping $p^{(l)}$ or the probability of surviving t_{l+1} given that the decision to continue with childbearing (i.e., exclusion of $-\infty$ from the sample space of Θ) is made.

The substitution of Eq. 3 into Eq. 2 implies the joint density that simultaneously accounts for the influence on the timing of births and quit from childbearing. This completes the formalization of the birth process.

4.2 The likelihood function and estimation method

Imposing arbitrary parametric assumptions on the distribution of unobserved heterogeneity may substantially influence estimation results (see Heckman and Singer 1984). To minimize the impact of distributional assumptions, we take a semiparametric approach to the estimation of the parameters of interest. In other words, it is only assumed that the distribution of unobservables is discrete and has a flexible form so that its mass points and corresponding probability values are estimated along with the structural parameters of the model. Let the distribution of the unobserved component have I points of increase $\{\theta_i\}_{i=1}{}^{I}$ with corresponding probability values $\{\pi_i\}_{i=1}{}^{I}$. Using Eqs. 1, 2, 3 and accounting for incomplete birth histories, the individual contribution to the likelihood function for a woman with K realized births can be written down as

$$\ln \ell = \ln \left(\sum_{i=1}^{I} \left[\prod_{j=1}^{K} \frac{\partial}{\partial t_j} \left(-\ln \left(S_j^* \left(t_j | \xi_j, \theta_i \right) \right) \right) S_j^* \left(t_j | \xi_j, \theta_i \right) \right]^{I_1} \times \left[S_{K+1}^* \left(\bar{t}_{K+1} | \xi_j, \theta_i \right) \right]^{I_2} \pi_i \right) \tag{4}$$

where ξ_j is the parameter vector that determines the survivor function $S^*(t)$, and I_1, I_2 are indicator functions such that I_1=0 if K=0 and I_2=1 if the time spell after the last birth is incompletely observed (i.e., we observe only $\bar{t}_{K+1}$ part of the true length t_{K+1}).

To complete the formulation of the likelihood function, we assume that the waiting time between births is Weibull-distributed, which implies a survivor function $S_j^* \left(t_j | \xi_j, \theta_i \right) = p^{[j-1]} + \left(1 - p^{[j-1]} \right) \exp \left\{ - \left(\exp \left\{ \mathbf{x_j} \beta_\mathbf{j} + \theta_\mathbf{i} \right\} t_j \right)^{\gamma_j} \right\}$. Furthermore, following Heckman and Walker (1987), we allow quit probabilities to depend on the covariates and introduce the parameterization $p^{[j-1]}=1/(1+\exp\{-\mathbf{x}_j\omega_j\})$.

To estimate the parameters of interest, we take the advice of Heckman and Singer (1982), and instead of direct maximization of Eq. 4 subject to $\sum_{i=1}^{I} \pi_i = 1,\ \pi_i \in (0, 1)$, we use an expectation-maximization (EM) algorithm, which avoids constrained maximization and is numerically more stable for high dimensions. To set up the algorithm, we follow Lancaster (1990). Consider the joint log-density of the time spells which build up the entire birth history of a woman given the parameters (ξ, ω, θ):

$$\ln \widetilde{\ell} = \sum_{i=1}^{I} \delta_i \ln \left(\left[\prod_{j=1}^{K} \frac{\partial}{\partial t_j} \left(-\ln \left(S_j^* \left(t_j | \xi_j, \omega_j, \theta_i \right) \right) \right) S_j^* \left(t_j | \xi_j, \omega_j, \theta_i \right) \right]^{I_1} \times \left[S_j^* \left(t_j | \xi_j, \omega_j, \theta_i \right) \right]^{I_2} \right) + \delta_i \ln \pi_i, \tag{5}$$

where δ_i is an element of the unobserved indicator vector that takes value one at certain θ_i and zero otherwise. On the E step of the algorithm, the unknown values of δ_i are substituted with their expected values

$$E(\delta_i) = \varphi\left(t\middle|\xi_j, \omega_j, \theta_i\right) / \sum_{i=1}^{I} \pi_j \varphi\left(t\middle|\xi_j, \omega_j, \theta_i\right), \tag{6}$$

where $\varphi\ (t \mid \xi_j, \omega_j, \theta_i)$ is a joint density of the time spells that build up the whole birth history of a woman given unobservables (i.e., integrand in Eq. 1, properly modified for quit probabilities and censoring in case the history is incomplete). Given Eq. 6 on the M step, Eq. 5 is maximized with respect to $\{\xi, \omega, \theta, \pi\}$, and the solution is used to form a new expectation of δ_i.[5] Dempster et al. (1977) demonstrate that for the general class of problems, the sequence of EM iterations converges to a solution that satisfies $\underset{\{\xi,\omega,\theta,\pi\}}{\arg\max}\left(\sum_{n=1}^{N} \ln \widetilde{\ell}_n\right) = \underset{\{\xi,\omega,\theta,\pi\}}{\arg\max}\left(\sum_{n=1}^{N} \ln \ell_n\right)$, where N is the sample size.

As long as the EM sequence may converge to a local maximum for any given I, we repeat estimation from the variety of starting values and detect the superior maximum. We set off with no heterogeneity case (I=1) and add a new point to the support of $M(\theta)$ until the total likelihood $\mathcal{L} = \sum_{n=1}^{N} \ln \ell_n$ fails to improve in terms of information criteria [Akaike Information Criterion (AIC) and Schwartz-Bayesian Criterion (SBC)].

To make the inference robust to possible misspecification of underlying density, we compute a quasi-maximum likelihood (QML) estimator of the covariance matrix introduced by White (1982)

$$\widehat{Cov}(\zeta) = \left(\sum_{n=1}^{N} \frac{\partial^2 \ln \ell_n}{\partial\widehat{\zeta}\partial\widehat{\zeta}'}\right)^{-1} \left[\sum_{n=1}^{N} \frac{\partial \ln \ell_n}{\partial\widehat{\zeta}} \frac{\partial \ln \ell_n}{\partial\widehat{\zeta}'}\right] \left(\sum_{n=1}^{N} \frac{\partial^2 \ln \ell_n}{\partial\widehat{\zeta}\partial\widehat{\zeta}'}\right)^{-1} \tag{7}$$

$\zeta = \left\{\beta, \gamma, \omega, \theta, \{\pi_i\}_{i=1}^{I-1}\right\}$ and $\ln\ \ell_n$ as in Eq. 4. This will minimize our error in case the true time spells are not Weibull-distributed.

In the next section, we also discuss the direction and the marginal size of the effect of the covariate on the length of waiting time and the magnitude of the quit probability. Details on the computation of both can be found in the Appendix.

5 Estimation results and discussion

This section contains the estimation results and their interpretation. As we have mentioned in our discussion of the data, only the first two births can sensibly be analyzed and are most important for the observed fertility decline. Therefore, we estimate the model, which contains information up to the second birth (or the end

[5] Identification of θ requires setting one of its elements to zero. Alternatively, one can suppress intercept term in $\mathbf{x}$ (see Heckman and Singer 1982, p.64).

of observation period) only. Estimation results are presented in Tables A1, A2, A3 of the Appendix.

5.1 Quantum effect of fertility decline

First, we discuss the quantum effect of the transition. As it was argued before, the chosen econometric model is capable of retrieving the progression probabilities which are otherwise noisily observed from the data with incomplete birth histories. This feature contributes to the illustration of the magnitude of fertility decline initially presented in Section 2 (see Figs. 1, 2).

Consider first the probability of a woman remaining childless. During the early and late socialist times, this probability was 3.4 and 5.7%, respectively. These estimates correspond fairly well to rates of biological infertility. However, when transition comes into play, the estimated probability of having no children increases from 5.7 to 30.6%, far beyond the rate of biological infertility (see Table A1). Keeping in mind that this change has happened in just 10–12 years, we see that the increase is immense. To demonstrate how large the effect of the transition is, we refer to Gustafsson et al. (2001), who show that for Britain, West Germany, the Netherlands, and Sweden, probabilities of remaining childless are normally fluctuating between 18 and 25% (with West Germany being the absolute leader with 25%).

Consider now the estimates of quit probability after the first birth. Again, in the early and late socialist times, this probability was 25.0 and 32.9%, respectively, showing not much of a discrepancy. However, when transition comes along, the estimated probability almost doubles and becomes 60.9%. As before, the increase in quit probability during the transition period is highly significant. Taken together, huge increases in the probability of having no births and quitting from childbearing after the first birth demonstrate that the first 10 years of the transition brought about permanent shifts into the fertility behavior of the population in the Czech Republic. This ensures that, induced by the transition, a considerable decrease of the population size in the next few decades is inevitable.

5.2 Fertility decline: socioeconomic determinants

Education level The first interesting result here is a uniform negative influence of education on the timing of the first birth. As is plausible, higher education maps into a later first birth. This is what we can see from the estimation results presented in Tables A1, A2, A3 of the Appendix: for all birth cohorts, an increase in education implies a longer waiting time until the first child is born. Furthermore, we see that with the start of the transition, the strength of the education effect goes up. While during the socialist time the magnitude of the effect appears to be quite stable, with the onset of transition, it becomes more than 30% higher. This observation can be formalized as a hypothesis of no increase in the effect of education (i.e., equality of "education" coefficients) across all the cohorts and tested by means of the standard Wald test. The test results (see Table A4) clearly show that in the youngest cohort, the effect of education on postponing the birth of the first child is significantly larger, whereas any change within the socialist time is rejected. This suggests that during the transition process, the ability (and willingness) to combine education with early childbearing has been reduced, which has led to ever-greater education-induced postponements.

To further understand this result, it would be particularly useful to consider the changes that happened to the structure of the labor market in the Czech Republic during the transition. In the socialist times, returns to education were relatively small because all wages were paid according to the administrative wage setting system rather than as a result of the interaction of labor market forces. With the transition to a market economy, returns to education acquire much higher importance. Following Kantorová (2003), the first reason for that is that the amount of accumulated human capital is positively correlated with the probability of being employed in the emerging modern sector that pays higher wages. The second reason is that human capital accumulation appears to act in the Czech Republic as an insurance from economic uncertainty—it can be observed that during the transition, unemployment rates of the most educated labor market participants were the lowest. In view of these changes, it is no surprise that the socioeconomic transition should lead to an increase in the postponement of the first birth, motivated by the increased opportunity cost of childbearing for women with higher human capital.

We also examine the marginal increase in waiting time induced by the additional year spent in education (see Fig. 3). The results demonstrate that in the transition period, the marginal effect of education on the postponement of the birth is again higher than during socialism. Furthermore, the significant departure of the marginal effect during the transition is observed only when the length of schooling crosses the 11- to 12-year threshold. This implies that only higher education has become increasingly important. In view of this observation, it could be quite instructive to calculate the difference in postponement between socialist and transition periods.

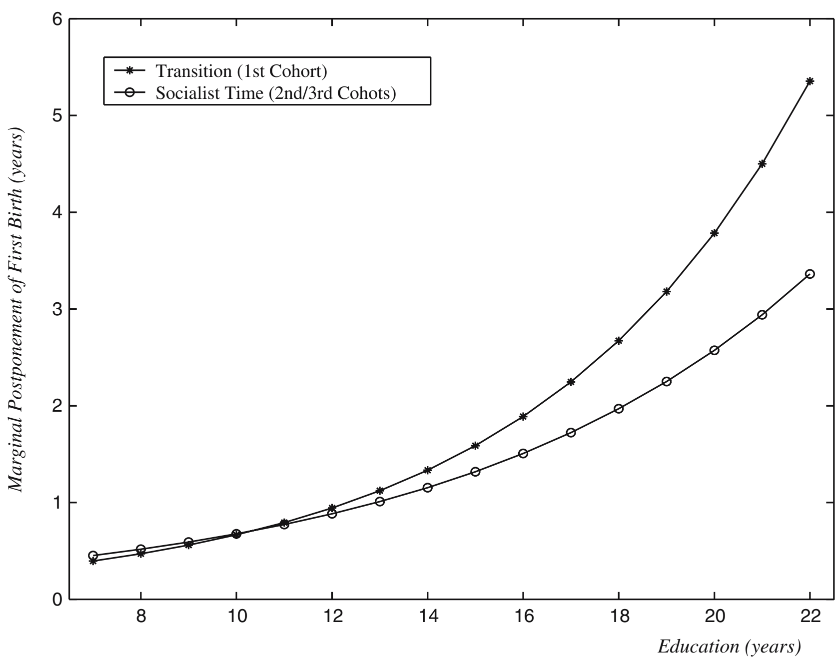

Fig. 3 Marginal effect of education on the timing of first birth

Assume that a representative of the first cohort decides to get a bachelor-equivalent degree, which corresponds to an increase in education length from 12 to 16 years. Then, after graduation, the postponement of her first birth will become 1.1 years longer than it would have been in the socialist times. The same applies to a person determined to complete a masters degree, i.e., increase one's length of schooling from 12 to at least 17 years. In this case, the postponement during the transition will be at least 1.7 years longer than it would have been before. These calculations strongly support the argument of Kantorová (2003) that during the transition, human capital accumulation considerations generate especially sharp postponement effects on the first birth.

Analyzing marginal effects further, we notice that the more educated the woman is, the greater is the marginal postponement of the first birth. The homogeneity of this increase in postponement by education is driven by the assumption that waiting time is Weibull-distributed (see Table A1 in the Appendix). However, if we examine the theory that links investment in preparental human capital with the timing of the first birth, such a distributional assumption seems plausible. Consider the Gustafsson (2001) model of the optimal age of motherhood. Gustafsson (2001) demonstrates that under the convex earnings profile, the capital cost of a later birth is smaller than the capital cost of an earlier one. Furthermore, the direct wage loss at the later birth is lower. Both of these facts imply that the higher is the preparental education level, the bigger is the incentive to postpone the first birth. Finally, our distributional assumptions are also in line with empirical findings of Gustafsson and Wetzels (2000) about the postponement of the first birth in West Germany, Britain, Sweden, and the Netherlands. Gustafsson and Wetzels (2000) show that higher educated women tend to postpone their first birth more than less educated ones.

It is also interesting to notice that the transition brought no large changes to the influence of education on the timing of the second birth. From Tables A1, A2, A3, A4, we see that the education coefficients are insignificant and small in absolute size. This is, again, fully consistent with the general view conveyed by Sobotka (2001) and others.

Consider now the influence of education on the exit from childbearing. First, let us discuss the exit from childbearing with no realized births. As we have already mentioned, the transition brought about a very sharp increase in the probability of no births. This probability has risen from about 5% in the socialist time to about 30% in the period of socioeconomic changes. From Table A1, we readily see that there exists a positive significant relationship between the amount of time spent in education and the resulting probability of no birth. It is also important to notice that the effect is again exclusively transition-specific because no such dependence can be observed throughout the whole socialist period (in fact, it was reversed). Analyzing the marginal increase in quit probability with an additional year of schooling (see Fig. 4), we also find that this increase is the highest at around 12–13 years, i.e., exactly at the point when the decision about going into the higher education is being made. The relationship essentially implies that lately, people have started considering that accumulating human capital is more important than having children. However, if there was a possibility of combining education at the higher stages with childbearing, such a relationship would not have existed because a woman would not have faced a trade off between having a child or continuing education. Therefore, our finding actually points at the limited ability to combine education with childbearing during the transition. This leads to an increased probability of not having a first child, showing

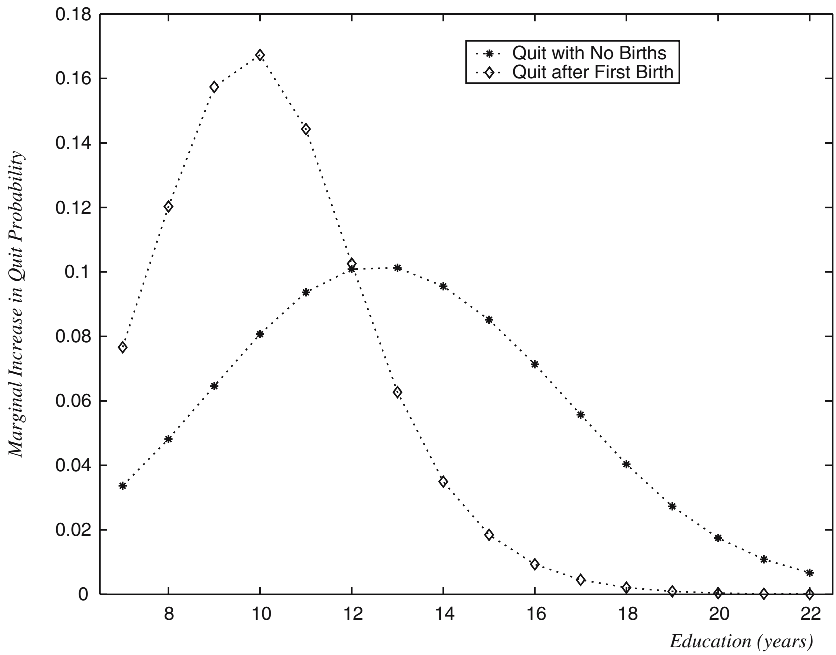

Fig. 4 Marginal effect of education on the quit probability (first cohort)

thereby not only significant tempo but also significant quantum effect. Moreover, this also points at the inability to combine childbearing with prospective employment. It is true that higher education implies higher returns, so there is an incentive to study longer. However, if employment and childbearing were compatible, the longer education would just postpone the birth without influencing the quit probability. So our findings suggest that women anticipate difficulties in having a good job and raising children, which fosters quit behavior even before entering the labor market.

Proceeding with the analysis, let us look into to the influence of education on the quit probability with only one realized birth. Again, during the transition, this probability increases up to almost 60%, which is approximately twice as much as it used to be in the late socialist period. Again, a positive significant relationship between years of education and quit probability can be observed only during the transition. In the two older cohorts, this variable is insignificant (Tables A2, A3). Looking at the curve that represents marginal quit probabilities (see Fig. 4), we can also see that compared to the same curve for no single birth, the marginal probability of quitting with one child only is much more skewed to the right. After 16 years of education, it almost touches the *x*-axis, demonstrating that further increases in years of schooling do not contribute to the increase in quit probability. Integrating this marginal effect over years of schooling, we discover that for those who stay in education 16 years or more, the probability of having two children is almost zero. This means that as a result of the changing role of education in the transition environment, a family with a mother having completed higher education (bachelor or master) is expected to be a one-child family only. We know that by the second birth, the formal education process is already over for almost everybody. Therefore, this finding just reinforces the discussed above implication of the difficulties that educated women would face if they try to combine work with childbearing.

Employment intensity The role of the next variable, female employment intensity, is equally important. As before, one may expect that higher total hours cause later births and earlier quits. Considering the first birth, this expectation is based on the theoretical model of Happel et al. (1984). In their model of the birth decision and premarital work experience, Happel et al. (1984) show that the optimal waiting time until the first birth must increase with experience. Because our measure of employment intensity mirrors the magnitude of work experience fairly well (see Section 2 for the discussion on the construction of the variable), we can easily use our results to test this prediction. From Tables A1, A2, A3, we see that for all cohorts, there is a significant positive dependence between employment intensity and postponement of the first birth, which strongly supports the predictions of Happel et al. (1984). However, we also notice that during the transition period, the negative role of employment starts falling. Sequential testing of the hypothesis of the equality of employment coefficients across cohorts (see Table A4) reveals time-constant effect under socialism and its subsequent significant decrease during the transition. Moreover, we can also state that this result is not driven by the fact that nonparticipation rates of childless females in the transition time have risen a lot (see Table 1 for cross-cohort comparison of nonparticipation rates before first birth). The result is remarkable since it implies that the dependence between work experience and first birth is more complex than the one described by Happel et al. (1984). To provide a more refined theoretical explanation of this finding, consider the alternative model formulated by Cigno (1991). In a more general framework, Cigno (1991) shows that higher premarriage work experience implies higher income, which generates a positive income effect on the timing of the birth. From this, he concludes that women who used to work more intensively may actually have their first birth faster. With this result in mind, consider now the transition economy. It is true that under socialism, with its administratively planned wage payment system, the key rewarding factors were age and seniority (see Kantorová 2003). With the transition, however, wage levels and promotion possibilities in the emerging modern sector become rather dependent on workers' abilities, productivity, and work intensity. This implies that at a sufficiently high level of intensity, the income effect described by Cigno (1991) may dominate the postponement effect of Happel et al. (1984). Therefore, there would exist a possibility that a woman may have her first child earlier than it would have happened in the socialist environment. This interpretation would be consistent with our estimation results.

We also plot the marginal effects of employment intensity on the postponement of the first birth (see Fig. 5). To make the correct inference from these plots, we need to remember that the average employment intensity was 2.33 in the youngest cohort, 2.32 in the medium-aged, and 2.35 in the eldest one (see Table 1). Here, we are interested in averages because such cases as employment intensity greater or equal to 5, which means that the woman is childless and works full-time with no unemployment spells, are quite rare in socialist times and almost nonexisting during the transition. Considering the average employment intensity, we see that during the transition, the increase in its level will map for the majority of respondents into a 1-year postponement. In the late socialist times, this postponement would have been as

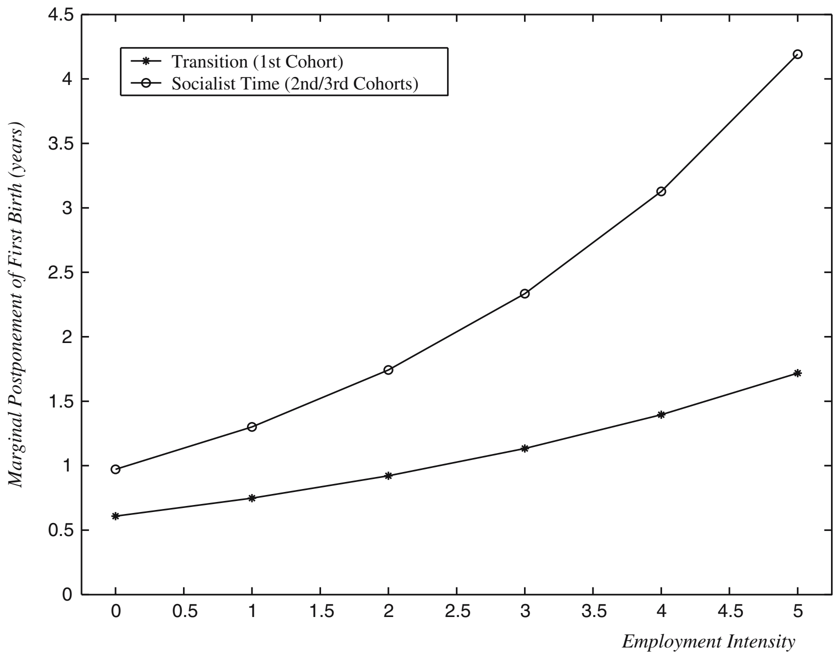

Fig. 5 Marginal effect of employment intensity on the timing of first birth

much as 2 years. Therefore, we may suggest that the income effect described by Cigno (1991) actually brought a relative reduction of waiting time by approximately 1 year.[6]

Consider now the influence of female employment intensity on the timing of the second birth. From Tables A1, A2, A3, we see that the relationship between employment intensity and postponement of the second birth is positive as we could have expected given the results of Happel et al. (1984). However, the influence of this variable in absolute terms is very small, and the amount of postponement turns out to be negligible. Testing the hypothesis of no change of negative effect of employment intensity on the timing of the second birth (Table A4), we again find that under the socialist regime, the effect is rather stable. With the transition, it falls and eventually becomes insignificant (Table A1). We should note, however, that our capacity of making the correct inference about the dependence between employment intensity and the timing of the second birth in the youngest cohort is quite limited. The reason is the incomparably high nonparticipation and permanent unemployment after the first birth, which makes it difficult to retrieve the true effect for those who are employed (see Table 1 for cross-cohort comparison).

We proceed with the influence of employment intensity on the likelihood of exiting from childbearing. We find that in the first cohort, there emerges a significant positive relationship between employment intensity and probability of not having children. The insignificance of the employment variable during the socialist time (see Tables A2, A3) also suggests that the discovered result is a purely transition-specific feature of the reproductive behavior. The interpretation is basically similar to the one

[6] Keep in mind, though, that the effect of employment intensity is still negative and significantly contributes to the delay of first birth as described by Happel et al. (1984).

already described in the case of education. It implies that the transition process made women to choose employment over children. On one hand, relatively high and increasing frictions in the transition labor market may have increased the importance of a career planning motive (see Hotz et al. 1997) for the employed prospective mothers. On the other hand, this relationship may be an implicit result of the inability to combine employment with childbearing. Arguing in the way we have already done before, if a degree of flexibility in working and childcare was high enough, no prospective mother would have faced a trade off between employment and the birth of a child. Consequently, the discovered relationship would have been insignificant as it used to be under socialist regime with its very strong publicly provided system of free childcare facilities (in fact, one of the main features of "socialist greenhouse" according to Sobotka 2001). Therefore, summarizing the evidence above, we may conclude that the transition-specific strengthening of career planing preferences along with the inability to combine childbearing with employment at high levels of working hours has led to the significant negative quantum effect of employment on Czech fertility.

Looking at the marginal increase in the probability of not having children induced by an increase in employment intensity (Fig. 6), we also notice that this increase is the highest for those who try to move from a part-time equivalent employment to a full-time employment. This result provides additional support to our conclusion about the inability to combine high total hours with childbearing.

Thinking about policy implications in this respect, one can turn to an Italian example analyzed by del Boca (2002). In a somewhat similar situation of low fertility and low participation equilibrium, del Boca (2002) empirically demonstrated that the promotion of part-time employment policies significantly increases the probability of being employed and having a child. Taken together, our discussion and the results of

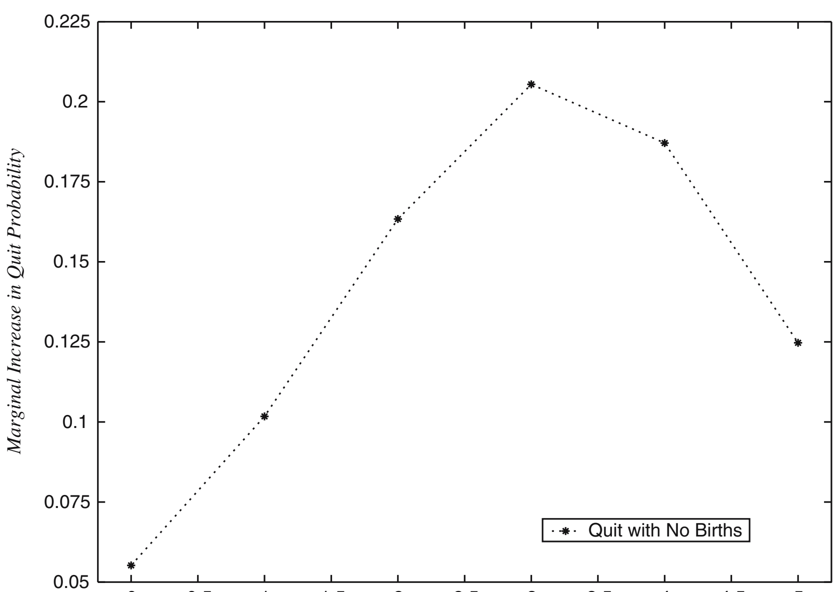

Fig. 6 Marginal effect of employment on the quit with no birth (first cohort)

del Boca (2002) could indeed make part-time work (as well as measures to combine childbearing with employment such as widely available and subsidized childcare) a valuable option for potential pronatal measures.

Finally, we briefly discuss the influence of employment intensity on the probability of quitting after the first child has been born. From Tables A2, A3 and A5, we see that in the socialist era, employment intensity had a stable positive effect on exiting from childbearing. With the beginning of transition, however, this effect ceases to be significant. In view of the very high nonparticipation rate and permanent unemployment, the reliability of this result for those who work is, however, questionable (see also the discussion above). Our intuition would be that employment intensity should have a positive significant effect on stopping after the first birth also during the transition.

Other socioeconomic determinants These include partner education, information on whether the apartment is owned or rented, and information about the current residence area of the respondent. Our findings about partner education demonstrate that this determinant is mostly relevant for the postponement of the first birth only. The pattern of dependence is also quite expected—the higher is the partner's education, the longer is the shift of the first birth into the future. It is also interesting to notice that the importance of partner education increases throughout the socialist time. But when the transition comes into play, this determinant loses its significance (see Tables A1, A2, A3).

Considering next the ownership information, we find that this variable fails to provide us with any clear pattern. Back in the socialist time, it was playing no role. The only exception was the statistically significant contribution to the reduction of quit probability with no births in the second cohort. With the start of the transition, this effect totally fades out. Women living in rented apartments are less likely to have the second, but not the first, birth. The difference in quit probability is quite high (about 10%).

Finally, an indicator of whether a woman resides in an urban (more than 100,000 inhabitants) area can be interesting only from the historical perspective because it does not enter the set of explanatory variables for the first cohorts (see Section 2). The estimated results demonstrate that there exists a stable postponement effect of urban residence on the first birth. This result also goes in line with Burcin and Kučera (2000), who report that among the total of six Czech cities with a population of more than 100,000 inhabitants, three—Prague, Plze, and Brno permanently stay in the top five locations with the lowest TFR. It is also worth mentioning that inclusion of this variable into the set of regressors for the first cohort when belief variables are not "personalized" as described in Section 3 shows that urban residence during the transition does not have any explanatory power whatsoever.[7]

[7] This finding is more of a demographic nature and is quite similar to the results of Bonneuil (1997) on the dependence between fertility and urbanization. Analyzing cointegrating relationships between fertility and urbanization in France, Bonneuil (1997) demonstrates that with the fall of fertility rates, the positive effect of urbanization on birth postponement disappears.

5.3 Fertility decline: beliefs

Every belief variable reflects a group-specific opinion about the importance of a certain reason in the contemporary fertility decline. The set of possible reasons consists of poor housing conditions, insufficient childcare facilities, desire for independence and personal advancement, and increased availability of contraception. Because all belief variables contain the information about the transition period only, we do not include them in the set of regressors when estimating the model for older cohorts. Any significant effect of the belief is, by construction, a transition-specific feature of the fertility decline.[8]

From the estimation results for the youngest cohort (Table A1), we see that two such variables do matter. First, we observe that increased availability of contraception significantly contributes to the probability of not having any birth. Results of a similar kind are not uncommon in the literature. For instance, Mašková and Stašová (2000) demonstrate that there exists a positive correlation between contraception expansion and the currently observed fertility decline in the Czech Republic. They argue that this fact can be regarded as an indicator of the end of the abortion culture that was the main instrument of birth control in the socialist time and switching to the new mode of fertility regulation (see also Hotz et al. 1997). Our results bring additional insight into the problem, suggesting that the steady use of contraception among young women becomes observationally equivalent to permanent postponement of childbearing (i.e., exit from it). To visualize the magnitude of the effect, we split the distribution of contraception belief into four 25% quantiles and calculate the quit probability at the mean belief value of each quantile. The results plotted in Fig. 7a demonstrate that roughly half of all respondents have below-average quit risk, and half of all respondents are above the average.

The significant role of contraception can be quite naturally expected in view of the increasing importance of career planning and consumption-smoothing motives, which lead to a desire to postpone births. In this case, contraception acts as an additional insurance that guarantees optimal planning and smoothing schedules.[9]

Another belief variable that turns out to play a role is the insufficiency of childcare facilities. From Table A1, we can see that this variable significantly increases stopping chances for those women who have already got one child. Unlike the previous belief, this one has a clear asymmetric behavior. Only the bottom quarter of all respondents have a below-average probability of ending up with a one-child family. The remaining three quarters lie above the average. Such pattern implies that even a moderate (below average) concern about childcare can induce the above-average propensity to give up childbearing after the first birth. If we consider those most worried about the existing situation with child care (top 25%) only three women of ten will decide to have another baby, whereas for the rest of the respondents, this number is four of ten.

The significance of this variable again is not unexpected. The reason is that during the transition, the costs of existing childcare facilities were partly shifted

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[8] We also ran the model for the youngest cohort without the belief variables to see whether this would change the other covariates (particularly the employment and education variables). It has very little impact on them, so that the results for these variables are robust to inclusion or exclusion of the belief variables.

[9] Apart from purely economic reasons, the significant effect of the increased use of contraception may also have behavioral and demographic roots (see Bonneuil 1997).

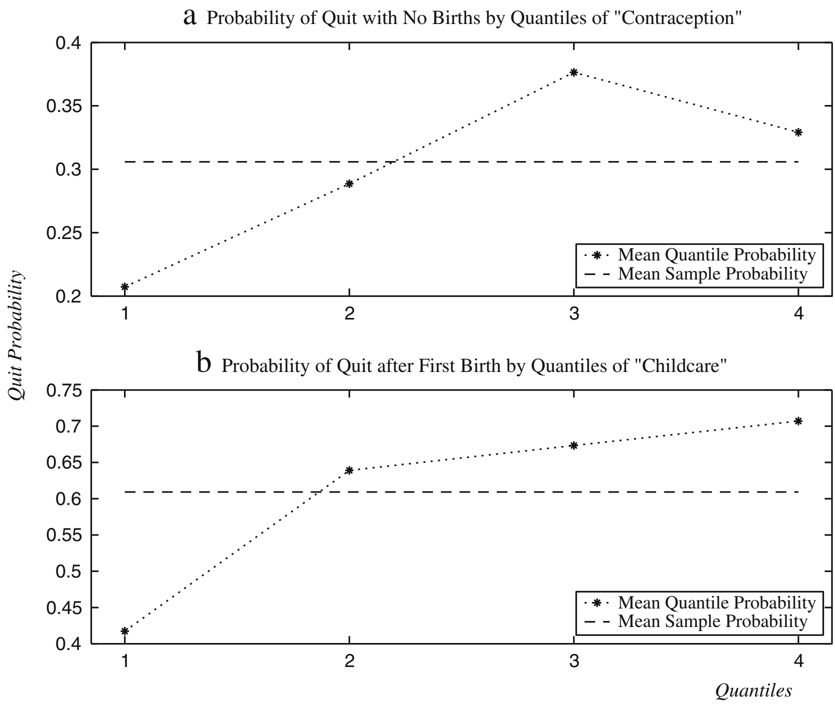

Fig. 7 Mean quit probabilities by quantiles of beliefs (first cohort) **a** Probability of quit with no births by quantiles of "contraception". **b** Probability of quit after first birth by quantiles of "childcare"

onto parents. Additionally, funding cuts could have potentially decreased the quality and availability of the childcare system. Taken together, this would imply that children have become more expensive since one and the same cost should now be paid for less adequate/sufficient childcare services. This fact would naturally lead to an upward shift in the quit probability.

In this context significant contribution of "children" to the reduction of the waiting time till the second birth could be interpreted as a desire of getting a child as quickly as possible, once a birth decision is made. Significant influence of this variable on the reduction of the danger of quit with no birth is, however, difficult to explain.

No other belief variables appear to have significant explanatory power, which positively characterizes the situation with housing conditions and implies that the emergence of new lifestyles and social norms is, at least so far, not incompatible with childbearing.

6 Summary and conclusions

In the present work, we focus on identifying the driving forces of the fertility decline in the Czech Republic during the transition period. Following the literature, we suggest that the observed fertility decline consists of two components, namely, postponement of the next birth (tempo effect) and quitting earlier from childbearing (quantum effect). The approach of this paper investigates the effect of the transition on both tempo and quantum of the observed fertility level.

We discover five transition-specific phenomena that expose a downward pressure on the TFRs. First of all, we find an increasing negative influence of education on the timing of the first birth. This means that in the transition period, women with higher levels of education will wait for the birth of their first child longer than it used to be in the socialist period. We also find that for women with undergraduate or higher degrees, this effect is especially strong. Secondly, there has appeared a positive dependence between education level and the probability of exit from childbearing. This includes both exit with no births and exit with only one realized birth. This dependence implies that during the transition, people willing to increase their education level are much more likely to end up with a one-child family—a fact that was not observed in the socialist era.

We also discover an emerging inability to combine employment with childbearing. The inability to combine employment with childbearing significantly reduces the probability of having any children, which we classify as a third harmful transition-specific effect.

Eventually, the last two factors that negatively influence fertility outcomes are the increased availability of contraception and the belief of insufficient childcare facilities. Agents who report these factors to be important for their birth decisions have the above-average risk of having no children or only one child, respectively. We argue that high importance of the use of contraceptives provides evidence of preventive measures taken to efficiently plan the career and redistribute consumption over time, which became especially difficult in the period of socioeconomic changes. The reported insufficiency of childcare facilities may be a reflection of the weaknesses in keeping up the old social security infrastructure in the early phases of transition.

We conclude that increased returns to education, the inability to combine employment with childbearing, insufficient childcare facilities, and increased use of contraception are the channels through which socioeconomic transition has negatively affected TFR in the Czech Republic. All other covariates, be they socioeconomic variables or belief variables, turn out to either not be significant at all or to not change significantly during the transition process.

The policy implication from these findings is rather clear. It appears that it has become increasingly difficult for educated women to combine higher education and employment with childbearing, so efforts to address this shortcoming are likely to be the most promising avenues for stemming the dramatic fertility decline in the Czech Republic and, likely, in other transition economies.

Acknowledgements We would like to thank Michael Grimm, Siv Gustafsson, and the participants of ESPE 2004 and EEA 2005. Comments of an anonymous referee were especially useful. Funding from the German Science Foundation (DFG) within SFB 386: “Statistical Analysis of Discrete Structures” is greatly acknowledged.

Appendix

Direction of the effect and marginal effects

Waiting time First, consider the effect of the covariates on Weibull hazard function

$$h(t|\theta_i) = \gamma(\exp\{\mathbf{x}\beta + \theta_\mathrm{i}\})^\gamma t^{1-\gamma}$$

Since the first derivative with respect to $\mathbf{x\beta}$ is positive, the sign of the effect will always be equal to the sign of the estimated parameter. For instance, a positive estimated coefficient conveys that an increase in $\mathbf{x}$ leads to an increase in the hazard of having a child and, hence, to a reduction in the waiting time between births.

To compute the size of the effect, consider the expected value of a Weibull-distributed random variable

$$E(T|\phi(\mathbf{x})) = \phi(\mathbf{x})^{-1}\Gamma\left(1+\gamma^{-1}\right),$$

where in our case, ϕ ($\mathbf{x}$)=exp $\{\mathbf{x}\beta+\theta_i\}$. In the presence of unobserved heterogeneity, this expectation modifies to

$$E(T|\mathbf{x},\theta) = \sum_i \frac{\pi_i}{\exp\{\mathbf{x}\beta+\theta_\mathrm{i}\}}\Gamma\left(1+\gamma^{-1}\right).$$

Taking the first derivative of $E(T\,|\,\mathbf{x},\theta)$ with respect to $\mathbf{x}$, we get

$$\partial E(T|\mathbf{x},\theta)\partial\mathbf{x} = \sum_i -\frac{\pi_i\Gamma(1+\gamma^{-1})}{\exp\{\mathbf{x}\beta+\theta_\mathrm{i}\}}\beta. \tag{A1}$$

As covariates enter $E(T\mathbf{x},\theta)$ nonlinearly, the marginal effect at the different levels of one and the same variable will not be the same (e.g., an additional year of schooling for a person with university education will cause a change in waiting time different from that caused by an additional year for a person with only secondary school education). To ensure correct inference, the marginal effects must be calculated for all possible levels of the explanatory variables.

Quit probabilities Identical inference can be made about changes in quit probabilities. With parameterization

$$p = (1+\exp\{-\mathbf{x}\omega\})^{-1},$$

an increase in the value of the variable will cause an increase in probability of quitting after the realized birth, provided that the corresponding parameter is positive. The marginal effect of the covariates on the expected quit probabilities is

$$\partial E(P|\mathbf{x})\partial\mathbf{x} = \frac{\exp\{-\mathbf{x}\omega\}}{(1+\exp\{-\mathbf{x}\omega\})^2}\omega.$$

Table A1 Estimated coefficients with parameterized quit probabilities: first birth cohort (born 1972–1982)

	Coefficient	SE	z statistic	p value		Coefficient	SE	z statistic	p value
First child					Second Child				
Education*	−0.1737	0.0490	−3.5458	0.0004	Education	−0.0272	0.0454	−0.5986	0.5495
Employment*	−0.2077	0.0432	−4.8046	0.0000	Employment	−0.0023	0.0223	−0.1041	0.9171
Rented	0.0165	0.1535	0.1077	0.9142	Rented	−0.1882	0.0996	−1.8904	0.0587
Housing	0.2833	0.2637	1.0742	0.2827	Housing	0.1411	0.2888	0.4888	0.6250
Childcare	−0.3902	0.5133	−0.7602	0.4471	Childcare*	1.2448	0.6102	2.0400	0.0414
Advancement	−0.3253	0.3814	−0.8528	0.3938	Advancement	0.0299	0.4749	0.0630	0.9498
Contraception	−0.0519	0.4826	−0.1076	0.9142	Contraception	−0.1671	0.3930	−0.4253	0.6706
Partner education 1	−0.0055	0.0780	−0.0704	0.9439	Partner education 1	−0.0516	0.1226	−0.4207	0.6740
Partner education 2	−0.0525	0.1333	−0.3941	0.6935	Partner education 2	−0.1413	0.1642	−0.8602	0.3897
Intercept	0.1283	0.5305	0.2419	0.8088	Intercept*	−1.4885	0.4977	−2.9907	0.0028
γ_1*	2.8232	0.4906	5.7535	0.0000	γ_2*	3.3120	0.3332	9.9407	0.0000
Estimated parameterized "stayer" probability, no births					Estimated parameterized "stayer" probability, first birth only				
Education*	0.5743	0.1736	3.3073	0.0009	Education*	0.8842	0.3207	2.7573	0.0058
Employment*	1.0112	0.2034	4.9717	0.0000	Employment	0.2229	0.1481	1.5051	0.1323
Rented	0.3230	0.3334	0.9687	0.3326	Rented*	−1.6167	0.6650	−2.4307	0.0151
Housing	1.4650	1.4333	1.0221	0.3067	Housing	2.0863	2.0401	1.0227	0.3065
Childcare*	−6.2670	2.9832	−2.1008	0.0357	Childcare*	15.5978	5.5926	2.7890	0.0053
Advancement	−6.1623	3.3570	−1.8356	0.0664	Advancement	−2.4911	3.3871	−0.7355	0.4620

Table A1 (continued)

	Coefficient	SE	*z* statistic	*p* value		Coefficient	SE	*z* statistic	*p* value
Contraception*	9.5981	3.9243	2.4458	0.0145	Contraception	4.4560	4.4044	1.0117	0.3116
Partner education 1	−0.1538	0.3907	−0.3938	0.6937	Partner education 1	0.4382	0.5847	0.7495	0.4535
Partner education 2	0.5065	0.6865	0.7378	0.4606	Partner education 2	−0.4450	1.1108	−0.4006	0.6887
Intercept*	−9.9253	2.2028	−4.5058	0.0000	Intercept*	−11.4330	3.2928	−3.4722	0.0005
	Mean	SE	95% confidence interval			Mean	SE	95% confidence interval	
$E[p(j0=0)]$	0.3058	0.0356	0.2361, 0.3755		$E[p(j=1)]$	0.6093	0.0539	0.5035, 0.7150	

Log (likelihood)=−630.458
Number of observations=268
Number of support points in m (μ)=3
*Significant at the 5% level

Table A2 Estimated coefficients with parameterized quit probabilities: second birth cohort (born 1963–1971)

	Coefficient	SE	*z* statistic	*p* value		Coefficient	SE	*z* statistic	*p* value
First child					Second child				
Education*	−0.1314	0.0070	−18.6541	0.0000	Education	−0.0023	0.0200	−0.1129	0.9101
Employment*	−0.3067	0.0218	−14.0427	0.0000	Employment*	−0.0582	0.0205	−2.8478	0.0044
Rented	−0.0314	0.0478	−0.6568	0.5113	Rented	0.0301	0.0765	0.3942	0.6934
Urban*	−0.1731	0.0737	−2.3478	0.0189	Urban	−0.1244	0.1114	−1.1172	0.2639
Partner education 1*	−0.1466	0.0516	−2.8391	0.0045	Partner education 1	−0.0030	0.0402	−0.0748	0.9404
Partner education 2*	−0.2996	0.0663	−4.5223	0.0000	Partner education 2	−0.0476	0.1079	−0.4418	0.6586
Intercept	−0.0031	0.0424	−0.0724	0.9423	Intercept*	−1.7408	0.2178	−7.9931	0.0000
γ_1*	3.1135	0.2182	14.2691	0.0000	γ_2*	2.0513	0.0885	23.1701	0.0000
Estimated parameterized “stayer” probability, no births					Estimated parameterized “stayer” probability, first birth only				
Education*	−0.6148	0.1741	−3.5319	0.0004	Education	0.0631	0.0629	1.0032	0.3158
Employment	0.0264	0.1524	0.1734	0.8624	Employment*	0.3726	0.0545	6.8370	0.0000
Rented*	−2.3482	0.4725	−4.9699	0.0000	Rented*	0.4257	0.2139	1.9896	0.0466
Urban	0.5391	0.5243	1.0283	0.3038	Urban	0.4903	0.3250	1.5086	0.1314
Partner education 1	−0.5925	0.4421	−1.3402	0.1802	Partner education 1	0.2720	0.2243	1.2129	0.2252
Partner education 2	0.5407	0.4951	1.0922	0.2748	Partner education 2	−0.2130	0.3939	−0.5407	0.5887
Intercept*	4.3461	2.0909	2.0786	0.0377	Intercept	−2.6617	0.7703	−3.4553	0.0005
	Mean	SE	95% confidence interval			Mean	SE	95% confidence interval	
E [*p* (*j*=0)]	0.0553	0.0087	0.0383, 0.0723		*E* [*p* (*j*=1)]	0.3295	0.0103	0.3090, 0.3493	

Log (likelihood)=−1,971.655
Number of observations=503
Number of support points in m (μ)=2
*Significant at the 5% level

Table A3 Estimated coefficients with parameterized quit probabilities: third birth cohort (born 1954–1962)

	Coefficient	SE	*z* statistic	*p* value		Coefficient	SE	*z* statistic	*p* value
First child					Second child				
Education*	−0.1356	0.0151	−8.9357	0.0000	Education	−0.0306	0.0232	−1.3196	0.1870
Employment*	−0.2768	0.0271	−10.2082	0.0000	Employment*	−0.0899	0.0214	−4.2020	0.0000
Rented	0.0428	0.0404	1.0585	0.2898	Rented	−0.1159	0.0770	−1.5049	0.1323
Urban*	−0.1386	0.0673	−2.0611	0.0393	Urban	−0.1036	0.1184	−0.8747	0.3817
Partner education 1	−0.0426	0.0424	−1.0034	0.3157	Partner education 1	−0.0899	0.1000	−0.8992	0.3686
Partner education 2*	−0.1847	0.0618	−2.9874	0.0028	Partner education 2	−0.2226	0.1295	−1.7181	0.0858
Intercept*	−0.6480	0.2050	−3.1609	0.0016	Intercept*	−1.6414	0.2865	−5.7289	0.0000
γ_1*	2.5400	0.1263	20.1155	0.0000	γ_2*	1.7811	0.1012	17.5947	0.0000
Estimated parameterized "stayer" probability, no births					Estimated parameterized "stayer" probability, first birth only				
Education*	−0.3732	0.0529	−7.0525	0.0000	Education	0.0725	0.0508	1.4275	0.1534
Employment	−0.0297	0.1128	−0.2633	0.7923	Employment*	0.3392	0.0580	5.8441	0.0000
Rented	−0.2907	0.3915	−0.7427	0.4577	Rented	0.1736	0.2183	0.7955	0.4263
Urban*	1.9907	0.4267	4.6653	0.0000	Urban	0.2496	0.3027	0.8244	0.4097
Partner education 1	0.5761	0.5515	1.0446	0.2962	Partner education 1*	−0.6939	0.2341	−2.9642	0.0030
Partner education 2	0.7775	0.6572	1.1831	0.2368	Partner education 2	−0.3775	0.3482	−1.0842	0.2783
Intercept	−0.1133	0.2643	−0.4286	0.6682	Intercept*	−2.7878	0.6057	−4.6026	0.0000
	Mean	SE	95% confidence interval			Mean	SE	95% confidence interval	
$E\ [p\ (j=0)]$	0.0344	0.0021	0.0303, 0.0385		$E\ [p\ (j=1)]$	0.2501	0.0081	0.2342, 0.2661	

Log (likelihood)=−1,977.399
Number of observations=437
Number of support points in $m\ (\mu)$=2
*Significant at the 5% level

Table A4 Test of no increase in the negative effect on timing of births

Timing of births	First birth		Second birth	
	$\chi_{(1)}^2$	p value	$\chi_{(1)}^2$	p value
Education III=education II	0.0781	0.7799	–	–
Education II=education I	36.1546	0.0000	–	–
Employment III=employment II	1.2139	0.2706	2.1861	0.1393
Employment II=employment I	20.5446	0.0000	7.4756	0.0063

Table A5 Test of no increase in the negative effect on exit from childbearing

Exit from childbearing	Exit after the first birth	
	$\chi_{(1)}^2$	p value
Employment III=employment II	0.3321	0.5645
Employment II=employment I	7.5416	0.0060

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