Institut für Wirtschaftsforschung Halle *Halle Institute for Economic Research*

Human Capital and Fertility in Germany after 1990: Evidence from a Multi-Spell Model

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December 2009 No. 22

IWH-Diskussionspapiere
IWH-Discussion Papers

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Abstract

We analyze the timing of birth of the first three children based on German panel data (GSOEP) within a hazard rate framework. A random effects estimator is used to accommodate correlation across spells. We consider the role of human capital – approximated by a Mincer-type regression – and its gender-specific effects on postponement of parenthood and possible recuperation at higher-order births. An advantage of the use of panel data in this context consists in its prospective nature, so that determinants of fertility can be measured when at risk rather than ex-post, thus helping to reduce the risk of reverse causality. The analysis finds evidence for strong recuperation effects, i.e., women with greater human capital endowments follow, on average, a different birth history trajectory, but with negligible curtailment of completed fertility.

Keywords: fertility, human capital, event history analysis

JEL classification: J13

Humankapital und Fertilität: Ergebnisse eines Multispell-Modells für Deutschland nach 1990

Zusammenfassung

Der Beitrag untersucht mit Daten des Sozio-Oekonomischen Panels (SOEP) und einem Verweildauermodell die Zeitpunkte der Geburt der ersten drei Kinder. Mit einem Random-Effects-Schätzer wird der Korrelation der einzelnen Übergänge Rechnung getragen. Besonderes Augenmerk gilt den geschlechtsspezifischen Effekten von Humankapital – approximiert durch die vorhergesagten Werte einer Lohnschätzung – im Hinblick auf die Aufschiebung des Übergangs zur Elternschaft sowie auf möglicherweise aufholendes Geburtsverhalten im Anschluss. Die longitudinalen Umfragedaten ermöglichen eine prospektive Analyse, d.h. mögliche Determinanten des Fertilitätsverhaltens können zum Zeitpunkt der Fertilitätsentscheidung abgebildet werden statt nachträglich gemessen zu werden. Dies senkt das Risiko, Ursache und Wirkung zu vertauschen. Die Untersuchung zeigt deutliche Hinweise auf aufholendes Geburtsverhalten: Frauen mit höheren Investitionen in Humankapital folgen einem anderen zeitlichen Fertilitätsverhalten, weisen aber keine oder nur geringe Unterschiede in der endgültigen Kinderzahl auf.

Schlagwörter: Fertilität, Humankapital, Verweildaueranalyse

JEL-Klassifikation: J13

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1 Introduction

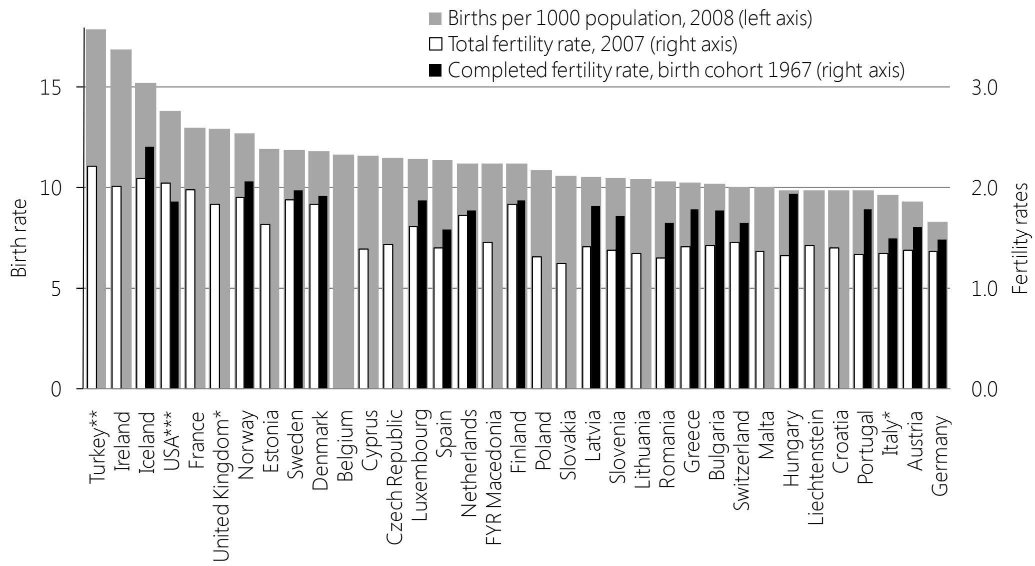
In the majority of developed countries fertility rates are below replacement level now.[1] If not compensated for by net-immigration, this results in demographic change towards older and smaller societies. The process cannot easily be reversed, as smaller birth cohorts now imply fewer potential mothers in the following decades. The ramifications of this process include pressure on social security systems because they typically rely on a balanced age structure. Factors underlying the low fertility rates in rich countries have thus received increasing attention. A particular feature of recent decades is the trend towards ever later family formation. This trend is considered as a potential problem, as a postponement in the onset of childbearing among women is associated with shorter remaining fertile periods and potentially lower completed fertility. Several forces have been discussed as contributors to postponement, including changing values ("second demographic transition"), increasing uncertainty e.g. in the form of fixed-term work contracts with the consequence of volatile income streams, but also rising educational attainment and career opportunities of women (Billari et al. 2006).

However, postponement of family formation may in theory be counteracted by a more rapid progression to higher-order births, which is referred to as "recuperation." To what extent postponement and recuperation go hand in hand, is a question that has to be addressed mainly from an empirical perspective, especially against the background of an expansion of tertiary education and corresponding career aspirations of women that are difficult to reconcile with early parenthood (Brewster and Rindfuss 2000). Differences in institutional settings can have considerable impact on the extent of recuperation, so that empirical associations may not be the same between different countries. In economics, greater career opportunities of women are usually believed not only to raise the demand for children through an income effect but also to raise the opportunity

[1] In 2002, 278 million Europeans lived in countries with total fertility rates (TFR) below 1.3 children per woman, which is regarded as "lowest-low" fertility (Kohler et al. 2002, Kohler 2006). This group comprises Mediterranean countries like Greece, Italy and Spain, and most of the Eastern European countries in the wake of the political and socio-economic transformation process.

costs associated with raising children. Hence, there are concerns that couples with high investments in human capital may favor a smaller family size, so that expansion of human capital might paradoxically produce smaller populations.[2]

Figure 1: Indicators of fertility in Europe and the US.



Remarks: * Total fertility rate refers to 2006. ** Total fertility rate from CIA world fact book 2009. *** Total fertility rate from CIA world fact book 2009, completed fertility rate refers to women aged 40–44 in 2006 (source: US Bureau of the Census).

Source: EUROSTAT, unless noted otherwise.

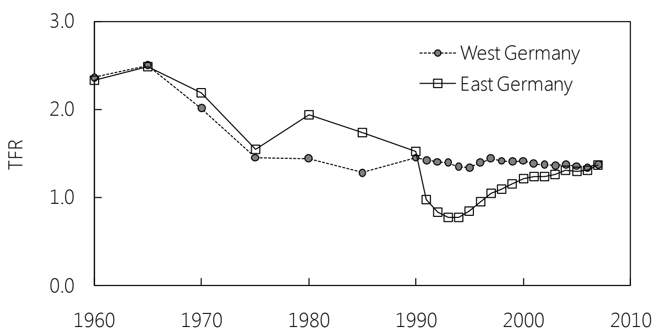
In the present paper we consider the German experience after unification from individual-level data.[3] The panel data available for Germany offer some unique opportunities to study fertility transitions, whereas data used in for many other countries are derived from cross-sectional surveys with retrospective birth histories. Such data usually do not provide accurate time-varying covariates, whereas with panel data it is, e.g., possible to assess an individual's stock of human capital at each point in time when making decisions about fertility rather than at the end of the fertility career. We extend existing models of German fertility by considering a multi-spell fertility process with control for unobserved heterogeneity (Heckman and Walker 1990a, Kravdal 2001, Rondinelli et al. 2006). Our focus lies on the association between the stock of human capital on the one hand and postponement and recuperation in birth histories on the other hand.

[2]See Myrsklylä et al. (2009) for a counter-argument.

[3]See Kreyenfeld (2001) for a very thorough analysis of German micro-data on fertility.

Germany as a country is an interesting case as it currently holds the European record of the lowest crude birth rate, and also has consistently low fertility rates. Germany and Italy are among the European countries with the lowest completed fertility rate in the cohort born 1967, whose female members arguably had almost completed their birth histories by 2008 (Figure 1). East Germany in 1993/1994 was the region with record-low period fertility rates at 0.77 children born to a hypothetical woman living all her life in that particular year (Witte and Wagner 1995, Conrad et al. 1996). The average age of married women at the age of birth of their first child increased rapidly in East Germany, from 24.9 to 28.4 years between 1991 and 2000. In the meantime, the rate has caught up with that of West Germany that had been undulating at a level of 1.3–1.4 (Figure 2). That is to say, even though there are currently European countries with transitorily lower period fertility rates mainly due to birth tempo distortions, Germany is one of the countries with a most chronic shortage of births, warranting analysis. Only very recently has the German Federal Statistical Office pointed out that German women born in the 1930s-60s with better educational background had given birth to fewer children throughout their lives than women on average, fueling concerns about differential fertility patterns by investment in human capital (Statistisches Bundesamt 2008a).

Figure 2: Total fertility rate in Germany, by region.



Remark: The western part of Berlin is counted as part of "West Germany" until 2000 and as part of "East Germany" thereafter.

Source: German Federal Statistical Office, *Fachserie 1 Reihe 1.1.*

2 Related literature

Several studies have considered micro-econometric modeling of determinants of fertility decisions focusing on education or female wages as factors influencing the opportunity cost associated with children. Many of these studies trace back to a seminal series of articles by Heckman and Walker (1990) who analyze the transition to the first three births in a Swedish context. Their merit lies in proposing the technical framework to such models and highlighting that neoclassical economic reasoning can improve the fit of previously "purely demographic" models: they documented strong negative effects of female wages and positive effects of the male wage on fertility. They also highlight the importance of allowing for unobserved heterogeneity, as child preferences or fecundity of a woman may be correlated across her births. In fact, if such components play a sizable role, then a separate estimation of each transition (e.g. to the birth of the second child) is a risky endeavor, as coefficients may be biased. This is a point that Kravdal (2001) stresses. While Heckman and Walker (1990) in fact did not obtain significant results for unobserved heterogeneity, Kravdal's analysis of Norwegian birth transitions shows that single-transition estimates deviate considerably from the multi-spell model that accounts for unobserved heterogeneity. While the former tend to give the impression that Norwegian women with higher formal education where more likely to have children, the latter model shows the opposite.

Heckman and Walker (1990) did not have actual wages at their disposal in the 1981 Swedish Fertility Survey data set. Hence, they proxied wages by annual data on the average wages of female and male workers to reflect income effects and opportunity costs of children. These macro wages are derived from aggregate personal tax returns of selected years within the period considered. While this approach is able to capture the strong increase of female wages relative to those of men in Sweden in the 1960s and 1970s, it obviously does not consider cross-sectional variation of human capital among the women in the sample. While Walker (2002) maintains that the use of aggregate wages circumvents issues of endogeneity bias that might arise with individual-level wages, a reduction in fertility may still account for some growth in relative female wages through a positive effect on female labor market experience. This may not have been the case in Sweden where political action contributed towards an equalization of female and male wages, but in general such a concericro wage data. After this change, the strong and negative association between female wages and fertility found by Heckman and Walker becomes much weaker. Walker (2002) identifies measurement problems in the data used by Tasiran and suggests to use predicted wages in the fertility regression model rather than actual wages.

Kreyenfeld (2002) analyzes the role of education on the transition to the second child in Germany, using register data (*Mikrozensus*). Apart from age and duration time, her model considers woman's education and the education of cohabiting partner. While the effect of partner's education exert a consistently positive effect, the role of female education to the second transition is less obvious. The role of female education only becomes clearer when adding the transition to the first birth, which is dramatically lower among college-educated women in the sample analyzed (by about 60%). Kreyenfeld suggests that the modest effect of women's education in the second transition may be explained by a self-selection process, inasmuch as college women at risk of having a second child have revealed their family preference in the first transition before. Once arrived in the risk set for the second birth, these women do not differ much from women without higher education. This result implies that neglecting previous transitions in fertility analyses may give rise to misleading interpretations. Kreyenfeld's model takes a shortcut for first births, though, by analyzing a dichotomous outcome only without the timing.

More recently, Rondinelli et al. (2006) investigated postponement and recuperation effects in Italy. Their model is close to the one in the present paper in that it assumes a discrete-time process with multiple transitions and unobserved heterogeneity. Furthermore, they use predicted wages as a measure of earnings capacity, or human capital. Their results suggest that, while there is considerable postponement of fertility among women with high earnings potential (evaluated at c. age 40), recuperation effects are stronger than previously thought, so that in spite of differences in the tempo of fertility, the quantum of fertility did not differ much across human capital strata.

Unfortunately, the analyses presented in the literature so far were often limited by the fact that data originated from cross-sectional surveys that do not allow to re-construct a person's trajectory of human capital investments. The stock of human capital at the time under risk is difficult to infer in such situations. Kravdal (2007) is the first to consider current – rather than final – educational attainment as a determinant of fertility in a multi-spell fertility model with unobserved heterogeneity. He argues that such an approach limits the risk of reverse causality, i.e. that women who realize they do not have children invest more into their career as a consequence. The present study seeks to connect to the works of Rondinelli et al. (2006), and Kravdal (2007) by applying slight variations of their approaches to German panel data.

3 Data and method

3.1 Data

The micro-level data used in the analysis are taken from the scientific use file of the German Socio-Economic Panel (GSOEP) conducted by the German Institute for Economic Research in Berlin (Wagner et al. 1993).[4] The GSOEP is an annual survey of private households in Germany, conducted for the first time in 1984 in West Germany, covering all persons aged 17 and above living in a surveyed household. Special attention is attached to the labor market status of respondents: interviews include recall questions on jobs and unemployment spells in each month of the preceding year.

The present study only considers data from waves 1990–2007 pertaining to unified Germany. GSOEP collects information on the birth histories of female participants. The history file administered by new survey participants is updated with every annual interview. In addition, birth histories of male participants have been collected since 2001. Thus, the analysis of male fertility below requires that the male respondent did not exit the panel before 2001. The fertility information is not restricted to the present marriage or union, and encompasses time of birth and gender of each child. While "historical" births are provided at annual resolution, births occurring between surveys are recorded by calendar month. Hence, we choose (calendar) month as the resolution of analysis time. The event of interest in the analysis is the timing of the birth of the first, second, or third child. Yet, we date births 10 months back in time so that explanatory variables reflect conditions prevailing at about the time of conception.[5] This condition requires that the months before an individual's most recent survey wave are excluded from the analysis. We also exclude the months following a conception up to the month in which the child was born from the risk set, as women are not at risk of becoming pregnant again during this period. Conceptions leading to a twin birth are considered as a single transition, but the parent then skips the following spell in our analysis. That is to say, spell 1 is followed by spell 3 if the first two children are twins. While information on fertility and labor market status are available at a monthly basis, the majority of GSOEP variables pertain to the time of annual interviews. In most cases, we extrapolate this information until newer information becomes available.[6]

All money values used in the analysis are expressed in EURO and are deflated. Monthly consumer prices are taken from the German Federal Statistical Office. Until December 1999, the price index is specific to the eastern and western parts of the country (Statistisches Bundesamt 2008b). After that date, we assume equal prices in both parts of the country. The original values were seasonally adjusted with the Census X-12 ARIMA procedure. The combined price series is then re-scaled to an average of 1.0 in 2007.

[4] URL: http://www.diw.de/english/soep/soepoverview/27908.html

[5] Births with missing information on month of birth are assumed to have occurred in January.

[6] Exceptions are the siblings and religion variables, for which the underlying information is not collected each year. In these cases, we also back-cast data.

Apart from potential effects of duration time and human capital (see below), we control for a set of variables often considered in the literature. These include household income and current activity in the labor market. Apart from non-participation in the labor market as the reference category, the activity variables comprise "working", being unemployed (short term or long term), and enrolment in education or training. These activity indicators were adjusted from the original data such that they are mutually exclusive. We expect that women who just became unemployed face lower opportunity costs of children, which should result in elevated transition rates. The opposite should be the case for men, at least in a traditional male breadwinner configuration. In this context we also consider the regional unemployment rate as as a factor that might raise concerns about future income.[7] We also include a set of socialization variables: number of siblings, foreign birth, and membership in a religious organization (Heineck 2006). Financial incentives by the government are difficult to model entirely as they depend, to some extent, on future labor market activities. We only include child allowances in the model, measured as the potential growth of household income due to the birth of the next child.[8] These allowances are paid until the child enters the labor market, and the values paid per month changed several times (by parity status) during the period under consideration, thus providing some variance. Finally we include a measure of life satisfaction, in the spirit of the Easterlin hypothesis, i.e. people being very satisfied with their standard of living relative to their own expectations may engage earlier in family formation.[9] Table 1 provides summary statistics for the estimation sample.

3.2 Duration model

Analysis of the timing of births requires us to study event histories rather than the number of births only. Event history methods allow us to include cases where fertility has not been completed by the end of the observation period. We treat time as discrete as our data exhibit monthly resolution at most. Thus, time period t in our model would correspond to the interval $[t-1, t)$ in a model with continuous time. A "spell" encompasses the time periods at which a person i is "at risk" of conception leading to her j^{th} child. During the (inferred) pregnancy period the person is excluded from the risk set, and after the month of child birth the person advances to spell $j+1$ (or $j+2$ in case of a twin birth). Thus, a

[7]Regional refers to states (NUTS1 level) with the exception of Saarland and Rhineland-Palatine which are combined to one unit. Monthly unemployment rates provided by the Federal Statistical Office were seasonally smoothed with the Census X-12 ARIMA procedure.

[8]We have deliberately top-coded this variable at 1 to remove the influence of a few outliers.

[9]See Kreyenfeld (2005) for modeling a related concept, economic worries.

Table 1: Descriptive statistics for regression data sets used in Table 3.

	WOMEN		MEN	
	mean	s.d.	mean	s.d.
age	32.267	7.666	32.409	7.589
human capital proxy	2.322	0.257	2.539	2.539
log real household income per capita	6.801	0.552	6.900	0.509
income growth due to child allowances	0.061	0.061	0.056	0.049
satisfied with life, range [-1 to +1]	0.404	0.347	0.402	0.338
number of siblings	1.840	1.646	1.855	1.711
regional unemployment rate	11.406	4.620	11.491	4.654
activity: in education/training	0.110		0.115	
activity: working	0.630		0.787	
activity: unemployed (up to 6 months)	0.025		0.026	
activity: unemployed (6+ months)	0.037		0.034	
activity: not in labor force	0.199		0.038	
member of religious organization	0.696		0.658	
no member of religious organization	0.304		0.342	
foreign born	0.134		0.124	
born in Germany	0.866		0.876	
has no partner	0.312		0.388	
has a partner	0.688		0.612	
persons	8,815		7,025	
person-months	618,770		534,490	
average period at risk in regression sample (years), all spells	6.2		6.7	
1st births	1,247		1,005	
2nd births	1,027		841	
3rd births	327		248	

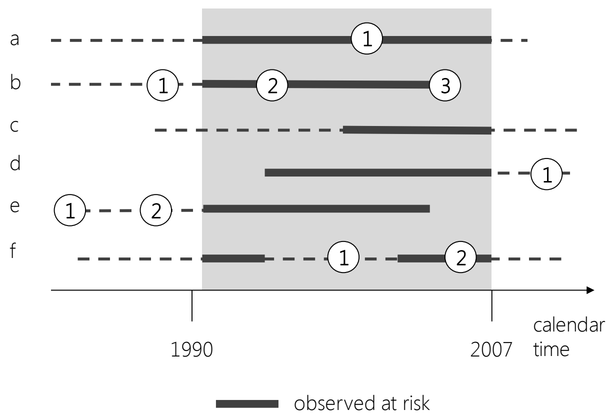
persons's fertility history consists of a sequence of non-overlapping spells. For the present analysis, fertility histories are censored at 9 months before the last survey interview, when the third child is born, or upon reaching age 45.[10]

Figure 3 illustrates a few hypothetical event histories. Transitions cannot be observed outside the shaded interval covered by our data. In case *a* the person was "at risk" for the first conception leading to birth already before we observe her (dashed line). After the first birth –which is observed– she is at risk for a second transition. This continues after the end of the observation period. Person *b* is observed for the first time when she is already at risk for a second transition. According to our criteria, we do not follow her birth history after the third transition took place. Individual *c* contributes with a censored first spell and

[10] Age 45 was chosen as it is about the mean age of onset of female sterility (Leridon 2004). Even among men, this age filter discards only very few recorded births.

may remain childless. As a result of age constraints or non-participation in surveys, persons may enter or exit the risk set during the period considered (cases *d* or *e*). Non-participation or missing values may furthermore give rise to interrupted spells, as in case *f*. We do not observe the first transition, but when the individual can be observed again after this transition, we know that she is in the risk set for the second transition due to the birth history information.

Figure 3: Hypothetical event histories.



For notational simplicity, we set analysis time to 1 at the beginning of a person's new spell. Thus, $t_i^{(j)}$ gives the time elapsed since the start of the jth spell of person i. The final (observed) time period of the spell is denoted as $T_i^{(j)}$. Calendar time may be recovered as $\tau = B_i + t_i^{(j)} + \sum_{h=0}^{j-1} T_i^{(h)}$, where B_i is i's date of birth in months and $T_i^{(0)}$ is the time period after which i is at risk for the very first time. The key element to describe the transition from one spell to the next is the hazard rate. It gives the probability of a conception leading to the birth of person i's j^{th} child within time period $t_i^{(j)}$, conditional on the fact that this child has not been born before:

$$\lambda_i^{(j)}\left(t_i^{(j)} \mid \mathbf{x}_{it}^{(j)}, \theta_i\right) = \Pr\left[T_i^{(j)} = t_i^{(j)} \mid T_i^{(j)} \geq t_i^{(j)}, \mathbf{x}_{it}^{(j)}, \theta_i\right] \tag{1}$$

The hazard rate may vary across observed and potentially time-varying characteristics that are captured by the vector $\mathbf{x}_{it}^{(j)}$. However, additional, unobserved factors may influence the hazard as well. Omitted variables in duration models tend to produce spurious negative duration dependence: as individuals with high transition probabilities exit the risk set early on, the distribution of types of individuals changes with duration time. If this composition effect is ignored, the effect would entirely be attributed to the effect of duration time. This is the more important as we are sampling from a stock because individuals may enter in the panel at various stages of the fertility process. The random effect θ_i summarizes the impact of unobserved heterogeneity in our model. It may reflect, e.g., the variation in fecundity or family preferences. As in Newman and McCulloch (1984), we assume that unobserved heterogeneity is specific to the individual but constant over analysis time and spells. We also assume that θ_i is independent of observed characteristics and can be approximated by a normal distribution with zero mean and unknown variance. Monte Carlo simulations by Nicoletti and Rondinelli (2009) suggest that the bias arising from misspecification of the random effect distribution is modest compared to the bias from ignoring random effects in the presence of unobserved heterogeneity.

The term $1-\lambda_i^{(j)}\left(t_i^j \mid \mathbf{x}_{it}^{(j)}, \theta_i\right)$ is then probability that the respective child is not born in the period considered, given that it had not been born until the beginning of that period. This implies that the probability of "survival" in the risk set (i.e. no birth) from the beginning of the spell until the end of period $t_i^{(j)}$ amounts to $\prod_{k=1}^{t_i^{(j)}}\left[1-\lambda_i^{(j)}\left(k \mid \mathbf{x}_{it}^{(j)}, \theta_i\right)\right]$. The likelihood of observing a particular spell depends on right-censoring. Let the binary indicator $d_i^{(j)} \in \{0,1\}$ recode whether or not the sample ends before the end of an active spell, i.e. whether the spell is censored. Right-censored spells contribute only with the survivor function, whereas completed spells contribute with the survivor function until the penultimate period and the hazard rate in the final period of the spell. For a given value of θ_i, these two cases may be expressed in a single equation using the censoring indicator (Allison 1982):

$$\ell_i^{(j)} = \left[\frac{\lambda_i^{(j)}\left(T_i^{(j)} \mid \mathbf{x}_{iT_i^{(j)}}^{(j)}, \theta_i\right)}{1-\lambda_i^{(j)}\left(T_i^{(j)} \mid \mathbf{x}_{iT_i^{(j)}}^{(j)}, \theta_i\right)}\right]^{d_i^{(j)}} \times \prod_{k=1}^{T_i^{(j)}}\left[1-\lambda_i^{(j)}\left(k \mid \mathbf{x}_{ik}^{(j)}, \theta_i\right)\right] \tag{2}$$

This expression assumes that the sample covers all periods in the beginning of a spell, i.e. there is no left truncation. When sampling from a stock, though, the starting points of the sample and individual i's first spell observed may not coincide.[11] If a spell is already under way in the first period covered by the data ($t_{0,i}^{(j)} < T_i^{(j)}$), the likelihood function in (2) has to be conditioned on survival through the pre-sample interval (Tsai et al. 1987, Uzunogullari and Wang 1992,

[11] We treat our sample as a stock sample – despite the availability of birth history data for the pre-panel period – due to lack of information on time-varying covariates.

Jenkins 1995):

$$\begin{aligned}\tilde{\ell}_i^{(j)} &= \ell_i^{(j)} \Big/ \prod_{k=1}^{t_{0,i}^{(j)}-1} \left[1-\lambda_i^{(j)}\left(k \mid \mathbf{x}_{ik}^{(j)},\theta_i\right)\right] \\ &= \left[\frac{\lambda_i^{(j)}\left(T_i^{(j)} \mid \mathbf{x}_{iT_i^{(j)}}^{(j)},\theta_i\right)}{1-\lambda_i^{(j)}\left(T_i^{(j)} \mid \mathbf{x}_{iT_i^{(j)}}^{(j)},\theta_i\right)}\right]^{d_i^{(j)}} \times \prod_{k=t_{0,i}^{(j)}}^{T_i^{(j)}} \left[1-\lambda_i^{(j)}\left(k \mid \mathbf{x}_{ik}^{(j)},\theta_i\right)\right] \end{aligned} \tag{3}$$

The likelihood of the observed sequence of spells associated with person i is the product of likelihoods of her spells. If a spell was completed before the start of the sample, the entire spell does not enter the likelihood function. We denote the first and last spell of i observed in the sample as $\underline{J}_i$ and $\overline{J}_i$, respectively. The likelihood of the sample is then the product of spell sequences over all individuals in the sample. As the random effect θ is unknown, the sample likelihood has to be integrated over its domain $\underline{\theta}$ and becomes[12]

$$L = \int_{\underline{\theta}} \Big(\prod_{i=1}^{n} \prod_{j=\underline{J}_i}^{\overline{J}_i} \tilde{\ell}_i^{(j)}\Big) d\theta \,. \tag{4}$$

In order to estimate the model, a logit parametrization of the hazard rate is assumed (Allison 1982):

$$\lambda_i^{(j)}\left(t_i^{(j)} \mid \mathbf{x}_{it}^{(j)},\theta_i\right) = \left\{1+\exp\left[-(\beta^{(j)\prime}\mathbf{x}_{it}^{(j)} + \gamma^{(j)\prime}\mathbf{y}_{it}^{(j)} + \theta_i)\right]\right\}^{-1} , \tag{5}$$

where $\mathbf{y}_{it}^{(j)}$ may contain terms allowing for estimation of duration dependence, age at previous birth, the effect of calendar time, and an intercept. Maximization of the sample likelihood with respect to the parameters requires numerical integration over the distribution of the random effect. A feasible implementation consists in the approximation of the normal distribution by a discrete distribution with a finite number of mass points, using Gauss-Hermite quadrature. Thus, the variance of θ can be estimated as an additional parameter.

3.3 Human capital

The literature offers some variances as to how human capital is implemented into fertility regressions. Studies based on cross section with retrospective fertility information often use education as a summary measure. While not explicitly taking work experience as another source of human capital into account,

[12] If θ is ignored the resulting model would be equivalent to a set of spell-specific models, which Heckman and Walker (1990a) label "piecemeal approach."

this approach may be justified for Germany as the propensity to work full time is positively associated with education among German women (Kreyenfeld and Geisler 2006). Heckman and Walker (1990a) came up with the idea of assigning mean wages by age and year groups from external sources to the individual-level data. Their procedure has been called into question by Tasiran (1995, 2002) as structural breaks in the external data used and a coarse grid of years used for interpolation of the series. As the analysis was stratified by birth cohort, the macro-wage approach may not pick up much of the cross-sectional heterogeneity within a year but instead reflect calendar year trends in general. Tasiran (2002) suggests imputing micro-level wages, using variables in the dataset that predict wages. However, such cross-sectional data do not offer information on the timing of investments in formal education nor on the length of labor market experience at each point in time.

Despite the availability of observed wages at the individual level, we do not consider them directly in the fertility model as this would limit the analysis to working women and possibly evoke sample selection issues. A popular alternative in the demographic literature consists in controlling for educational attainment at the end of the fertility career of an individual. A virtue of such an approach might consist in the possibility to straightforwardly interpret coefficients, and often such approaches can be employed when the original data are cross-sectional and birth histories can be reconstructed from household composition. However, final educational attainment as a regressor may give rise to problems if backcast to periods at risk when the individual still attended school or college (Kravdal 2007).

We employ the association between human capital and wages to predict earnings capacity. While the data permit construction of wage data at the individual level for the fertility analysis, we refrain from using these directly in the fertility model as such a procedure would restrict the analysis to working individuals. Instead our indicator of human capital is the predicted logarithm of the (real) hourly wage rate obtained from a Mincer-type wage regression. This approach is related to Rondinelli et al. (2006). However, due to the panel information available, we may use the individual's current education (ed) rather than final education, and actual years of work experience (ex) instead of age. Furthermore, the annual updates of the survey allow us to control for trends in the returns to human capital. Lupo and Anger (2009) document that such trends exist in post-unification Germany. While their analysis allows for very flexible trends, we estimate a simplified version of the wage equation, stratified by gender:

$$\ln w_{it} = \sum_{p=0}^{P} \sum_{q=0}^{1} t^p G_{it}^q \cdot \{\alpha_{pq,0} + \alpha_{pq,1} ED_{it} + \alpha_{pq,2} EX_{it} + \alpha_{pq,3} (EX_{it})^2\} + u_{it} \quad (6)$$

The dummy variable G is set to 1 if an individual lives in the territory of the former GDR at the time of the interview and 0 otherwise. If $P > 0$ the specification allows for drift in the returns to human capital through interactions with calendar time (t). Such drift has been documented by Lupo and Anger (2009) for Germany. Estimation is carried out on the basis of the GSOEP data. We restrict the samples to individuals aged 17–45 who worked for at least 10 hours per week at the time of the survey. Observations are taken from survey years 1990 (East Germany: 1991) to 2006.[13]

4 Results

Before addressing the factors associated with fertility we turn to results of our wage regression. Table 2 presents OLS coefficients with standard errors adjusted for clustering within individuals. Note that the results are not adjusted for sample selection, as the usual instruments to identify labor force participation involve fertility choices, which would introduce problems of interpretation when using wages in the fertility equation later on. There are also no industry dummies or firm size effects included in the model because these would complicate wage predictions for individuals outside the labor force.

Our estimates document some drift in parameters over time, so we use linear time interactions in our preferred model. The trend suggests that real wages of workers with a low stock of human capital stagnated or even declined. At the same time, returns to human capital increased, especially in the East. An additional year of formal education was rewarded with a wage premium of 4% among East German men in 1991, and with 9% in 2006, whereas the increase amounted to only about 2 percentage points in the West. Female returns to formal education were also 9% in 2006 but started from a higher level than male ones. To be sure, wages of the GDR era are not included in the estimation sample. It would seem very likely, though, that the widening of wage dispersion after the demise of the command economy was even larger than what our model implies. Even at the end of the period considered, wages in the east at given human capital stock were more than 20% below western levels. Notice that formal education is not the number of years the individual actually spent in education but the number of years it would at least take to obtain the person's degree (up to university degrees), so that repeating grades at school does not inflate the measure. Years of work experience are taken from employment history information and are not imputed from age. Thus, individuals with the same age and formal education may have quite different (predicted) wages. As these covari-

[13] Thus, individuals can enter the regression data set more than once, which we take into account by reporting clustered standard errors as in Lupo and Anger (2009).

Table 2: Wage regression estimates.

	WOMEN		MEN	
	coef.	s.e.	coef.	s.e.
constant	1.51983***	0.05149	1.83214***	0.03819
ED	0.05760***	0.00419	0.04993***	0.00299
EX	0.03195***	0.00458	0.03676***	0.00351
EX^2	-0.00075***	0.00019	-0.00092***	0.00013
$East$	-0.37676***	0.11068	-0.07676	0.08903
$East \times ED$	0.01558*	0.00839	-0.01779***	0.00669
$East \times EX$	-0.01543*	0.00838	-0.01353*	0.00773
$East \times EX^2$	0.00032	0.00033	0.00006	0.00029
t	-0.01704***	0.00486	-0.02806***	0.00393
$t \times ED$	0.00100***	0.00038	0.00108***	0.00030
$t \times EX$	0.00077*	0.00045	0.00201***	0.00038
$t \times EX^2$	-0.00001	0.00002	-0.00005***	0.00001
$t \times East$	-0.00803	0.01074	-0.02479**	0.00992
$t \times East \times ED$	-0.00007	0.00078	0.00240***	0.00073
$t \times East \times EX$	0.00137	0.00087	-0.00021	0.00086
$t \times East \times EX^2$	-0.00002	0.00003	0.00002	0.00003
R^2	0.22		0.29	
person-years	37,583		47,910	
persons	7,492		8,324	

Remarks: The dependent variable is the natural logarithm of the real hourly wage in 2007' EURO. Only individuals aged 17–44 years working for at least 10 hours per week at the time of the interview are considered. The variable t is defined as calendar year minus 1990. Standard errors are corrected for clustering within subjects. */**/*** indicates that a coefficient is statistically different from zero at the 10/5/1 percent level.

Source: GSOEP waves 1990–2006.

ate values are also available for persons not working, we can predict log real wages from these equations for the entire fertility sample as a proxy for human capital. In the case of women, this summary measure may be a more appropriate indicator of opportunity costs associated with raising children than a woman's education observed at the end of her fertile period because it avoids reverse causality as only information known up to the period at risk is used in construction; in addition, it acknowledges that formal education is not an exclusive source of earnings capacity.

Results of the fertility model are presented in Tables 3 and 4. In both cases, we focus on individuals aged 17–45 at risk of conception leading to the birth of one of the first three children. The former table considers the "entire" sample (with nonmissing variable values), whereas the latter is restricted to individuals with cohabiting partner. In such cases, the partner is included in the GSOEP, so that additional information can be used. Let us first look at results for the entire sample and the effect of our "control variables", though. We restrict some explanatory variables to have the same influence for all transitions while others may have spell-specific effects (see, e.g. Barmby and Cigno 1990). Information on partner is limited to a dummy variable recoding whether the person has no steady partner, independent of his or her survey participation status.[14] Not surprisingly, those without partner are much less likely to have children. We do not include marital status, though, as marriages may be formed with the intention to have children, giving rise to endogeneity concerns.

A further strong obstacle is enrolment in education or training programs, even when controlling for age, with the effect being stronger among women. This result is very common in the literature and does not come as a surprise, but it is nevertheless important to keep in mind when analyzing the joint effect of human capital. Other activities considered are "working", "unemployed" and neither being in the labor force nor in education. Among women, the positive effect of being unemployed makes sense because unemployment reduces the opportunity cost of children. That long-term unemployment is associated with elevated fertility of men is less consistent with theoretical considerations. However, these effects are estimated net of household income, so that in a male breadwinner partnership arrangement, the relatively strong effect of (net) household income on fertility also has to be considered in the equation, and a 50% reduction in income associated with long-term unemployment would nullify the positive coefficient of unemployment.[15]. Alternative definitions of income considered – such as gross unearned income – produced quantitatively similar results. This specification also yields a significant association between fertility and the relative magnitude of child allowances paid by the government for the next child. The effect has the positive sign that one would expect, whereas the large coefficients are somewhat misleading because of the small income growth rates associated with this component (0.06).[16] Even though we already control for income, there is an independent positive effect of life satisfaction, which suggests that quality of life beyond its monetary component is conducive to population growth. The socialization variables considered in the model are all statistically significant: religious attachment, birth in a

[14]This information is only updated annually, so there is some room for mismeasurement. We also cannot determine for how long partners (without being married) had cohabited at the start of the sampling period, so we have to leave out length of partnership.

[15]$0.561 \cdot \Delta \ln income + 0.395 \cdot 1 \stackrel{!}{=} 0$

[16]While child allowances are paid until the child entered the labor market, there existed additional means of financial support to parents with a shorter payment period that have not been considered in this variable (such as *Erziehungsgeld* by the federal government and some states within Germany.)

Table 3: Fertility regression results (random effects logit models).

	WOMEN		MEN	
	coef.	s.e.	coef.	s.e.
1st birth:				
constant	-3.908***	0.801	-10.880***	1.076
duration	-0.294***	0.052	0.040	0.059
duration squared	-0.018***	0.001	-0.012***	0.001
East Germany	-0.571**	0.249	-0.365	0.331
human capital	-1.947***	0.339	0.207	0.429
duration × human capital	0.254***	0.027	0.074***	0.027
2nd birth:				
constant	-9.203***	0.682	-12.632***	0.880
duration	0.808***	0.138	1.162***	0.199
duration squared	-0.035***	0.004	-0.050***	0.005
age at previous birth	-0.103***	0.012	-0.085***	0.013
1st child is a boy	0.046	0.070	0.108	0.078
East Germany	-0.895***	0.270	-0.845**	0.355
1st child born in GDR	-0.126	0.238	-0.138	0.292
human capital	2.004***	0.304	2.192***	0.343
duration × human capital	-0.261***	0.057	-0.325***	0.072
3rd birth:				
constant	-7.719***	0.995	-10.588***	1.398
duration	0.862***	0.190	0.848***	0.269
duration squared	-0.014***	0.004	-0.011**	0.005
age at previous birth	-0.154***	0.021	-0.139***	0.022
first two children have same sex	0.190	0.117	0.383***	0.135
East Germany	-1.042***	0.361	-0.738	0.455
1st child born in GDR	-0.020	0.325	-0.723*	0.401
human capital	1.539***	0.498	1.598***	0.595
duration × human capital	-0.379***	0.079	-0.335***	0.096
All births:				
in education/training	-0.907***	0.114	-0.481**	0.200
working	-0.123**	0.063	0.121	0.158
unemployed (short term)	0.285**	0.122	0.190	0.211
unemployed (long term)	0.270**	0.120	0.395*	0.202
log real household income (per capita)	0.305***	0.063	0.561***	0.074
income growth due to child allowances	1.684***	0.470	3.121***	0.728
satisfied with life, range [-1 to 1]	0.194***	0.067	0.214***	0.078
member of religious organization	0.395***	0.064	0.332***	0.067
foreign born	0.204***	0.065	0.361***	0.073
number of siblings	0.074***	0.014	0.083***	0.014
has no partner	-1.225***	0.063	-1.864***	0.082

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	WOMEN		MEN	
	coef.	s.e.	coef.	s.e.
Macro level:				
regional unemployment rate	-0.010	0.010	-0.012	0.011
linear time trend (1990 = 0)	0.062***	0.024	0.058**	0.027
time trend squared	-0.004***	0.001	-0.004***	0.001
East Germany × time trend	0.129**	0.056	0.155**	0.070
East Germany × time trend squared	-0.002	0.003	-0.005	0.004
σ	0.561	0.021	0.571	0.026
LR-test unobs. het. (p-value)	0.000		0.004	
person-months	618,770		534,490	
persons	8,815		7,025	
births	2,601		2,094	

Remarks: */**/*** indicates that a coefficient is statistically different from zero at the 10/5/1 percent level.
Source: GSOEP waves 1990–2007.

foreign country, and the number of own siblings are positively associated with fertility among both genders. While we hypothesized that regional unemployment should be negatively associated with fertility, we fail to find a statistically meaningful relationship in this case. Instead, our model fails to explain much of the fertility rebound in East Germany after the transformation trough, as the significant trend variables indicate. Furthermore, strong statistical significance is attached to the random effects component, i.e. the explanatory ignore some important factors for fertility transitions. In a case like this, it is better to estimate the model jointly than to have separate models for each spell (Heckman and Walker 1990a).

Let us now turn to spell-specific effects. Our proxy for human capital is strongly associated with the transition to parenthood and also higher-order birth transitions among women, whereas the association is weaker in men in the case of the transition into parenthood. As expected, men with higher earnings potential are able to have the first child earlier, whereas the effect in women is more difficult to interpret quantitatively: earnings capacity tends to reduce the first transition intensity at first and then lifts it at higher waiting times.

Higher-order transitions include the age at previous birth, i.e. cumulative lagged duration. If postponement of first birth would generally be associated with recuperation, one would expect that "late" parents try to reduce birth spacing for subsequent births. Empirically, the opposite seems to be the case, with older parents having lower higher-order transition intensities. This pattern corresponds to the "engine of fertility" notion, where more transition rates are correlated within subjects across spells. However, the association is modified through human capital. In contrast to the pattern of women's first birth, human capital at higher-order births is associated with increased transition intensities at short waiting times. This implies that recuperation may be selective with respect to earnings capacity.

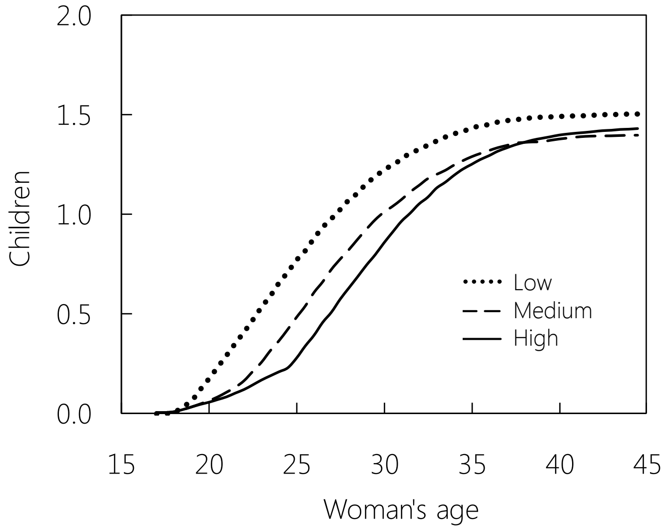
The implication is difficult to read off the coefficients, though, inasmuch as human capital in our framework may evolve over time instead of being fixed at its final period value. Furthermore, greater human capital formation typically involves longer formal education, which we know acts as a deterrent to childbirth. We may illustrate the course of average number of children born to a woman with certain characteristics by means of simulation. We simulate transitions for a hypothetical woman living in West Germany in the year 1998 throughout her life between ages 17 and 45.[17] We consider three cases: a "middle"-range human capital trajectory (roughly following the sample average of female human capital and associated with leaving formal education at the 21st birthday), a "low" trajectory (25% lower predicted wage, leaving education at age 18), and the "high" road (25% higher predicted wage, leaving education at age 24). The exercise involves drawing a random effect for each replication from the normal distribution with the estimated standard error and drawing random numbers from the extreme value distribution to obtain logit probabilities. Once a probability crosses the 0.5 marker in the first (second) spell, the woman enters the second (third) spell with the respective coefficients, without considering the possibility of twin births. Each of the three types of trajectories is simulated with 2000 replications. Figure 4 plots the average number of children for each type by age (at birth rather than conception). Notice that, as is the case in the underlying model, no births beyond the third one are considered in the simulation. While the simulation predicts considerable differences in fertility by age 25 (0.5 children) between the "low" and the "high" type, they end up with very similar completed fertility.[18] This would suggest that postponement of motherhood related to human capital investments is almost entirely offset by recuperation at higher ages.

A caveat against this model consists in the lack of information on partners. In particular, in the presence of assortative mating chances are that human capital is correlated in couples. Thus, recuperation among women with high scores on the human capital proxy might rather be related to the earnings capacity of the husband. Table 4 presents the results that pertain only to individuals whose partner is also covered by the GSOEP. We can thus augment the list of covariates by the education of the partner, expressed as a difference in years of

[17]Further assumptions: all variables in "all birth" section set to zero except for income and child allowances (sample means), regional unemployment rate set to 7%.

[18]Notice, however, that the simulation follows the underlying model in disregarding any births after the third birth and any births before age 17.

Figure 4: Model predictions: number of children per woman by human capital trajectory.



Remark: Simulation involves coefficients from Table 3. See text for details.

formal education between the partner and the sample person.[19] While matching information on the partner reduces sample size dramatically, interesting results emerge. Education of the partner has an additional, statistically significant (and positive) impact on the fertility of women, while higher education of a female partner does not raise (or lower) the fertility outcome of men in this sample. It is remarkable that the coefficients associated with human capital in the women sample tend to get stronger rather than weaker by the inclusion of partner characteristics. We also introduce age difference as a regressor. For both genders it indicates that cohabiting with a younger partner is conducive to fertility. Bringing additional children into a relationship – measured by the dummy variable "partner has more children" – reduces fertility aspirations of men considerably, while we find no such statistically significant effect among women. Still, these additional explanatory variables cannot remove the significance of the random effects component.

The fertility regression models also point out to regional differences in fertility transitions. These are dominated by a strong positive time trend. Even when controlling for income, the hazard rate in the east is much below the western

[19] It would be desirable to use the human capital proxy for the partner as well, but this would require discarding partners older than 45 years.

Table 4: Fertility regression results for individuals with partner present (random effects logit models).

	WOMEN		MEN	
	coef.	s.e.	coef.	s.e.
1st birth:				
constant	-3.282**	1.326	-10.430***	1.613
duration	-0.331***	0.079	-0.001	0.082
duration squared	-0.019***	0.002	-0.011***	0.001
East Germany	-0.862**	0.362	-0.404	0.390
human capital	-1.907***	0.531	0.723	0.611
duration × human capital	0.269***	0.041	0.065*	0.037
2nd birth:				
constant	-9.481***	0.939	-10.682***	1.065
duration	0.976***	0.171	1.029***	0.205
duration squared	-0.043***	0.005	-0.046***	0.005
age at previous birth	-0.139***	0.015	-0.116***	0.014
1st child is a boy	0.092	0.079	0.113	0.080
East Germany	-1.156***	0.376	-0.988**	0.406
1st child born in GDR	0.113	0.283	-0.137	0.308
human capital	2.707***	0.363	2.244***	0.383
duration × human capital	-0.314***	0.070	-0.288***	0.075
3rd birth:				
constant	-8.132***	1.242	-9.018***	1.561
duration	1.035***	0.221	0.825***	0.282
duration squared	-0.013***	0.004	-0.009**	0.005
age at previous birth	-0.201***	0.024	-0.168***	0.023
first two children have same sex	0.249*	0.132	0.326**	0.137
East Germany	-1.105**	0.462	-0.806	0.497
1st child born in GDR	0.006	0.367	-0.715*	0.412
human capital	2.495***	0.573	1.812***	0.625
duration × human capital	-0.479***	0.091	-0.336***	0.101
All births:				
in education/training	-0.680***	0.167	-0.515**	0.240
working	-0.125*	0.073	-0.252	0.173
unemployed (short term)	0.172	0.160	-0.147	0.236
unemployed (long term)	0.194	0.158	0.037	0.230
log real household income (per capita)	0.260**	0.105	0.451***	0.105
income growth due to child allowances	-0.562	1.521	0.687	1.573
satisfied with life, range [-1 to 1]	0.160*	0.084	0.186**	0.088
member of religious organization	0.464***	0.078	0.324***	0.074
foreign born	0.240***	0.080	0.281***	0.080
number of siblings	0.077***	0.017	0.067***	0.016

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	WOMEN		MEN	
	coef.	s.e.	coef.	s.e.
Partner:				
age difference	-0.041***	0.007	-0.077***	0.009
education difference	0.051***	0.012	0.020	0.013
has more children	-0.181	0.167	-0.671***	0.150
Macro level:				
regional unemployment rate	-0.018	0.012	-0.021*	0.012
linear time trend (1990 = 0)	0.082***	0.030	0.078**	0.031
time trend squared	-0.004***	0.002	-0.005***	0.002
East Germany × time trend	0.200***	0.077	0.197**	0.080
East Germany × trend squared	-0.006	0.004	-0.006	0.004
σ	0.549	0.023	0.545	0.024
LR-test unobs. het. (p-value)	0.000		0.000	
person-months	352,252		308,695	
persons	5,376		4,788	
births	1,807		1,730	

Remarks: */**/*** indicates that a coefficient is statistically different from zero at the 10/5/1 percent level.

Source: GSOEP waves 1990–2007.

level at the beginning of the period considered. About five years after unification, though, the model for women predicts equal first birth transition rates for both parts of the country had all other variables been equal. Higher-order birth transitions rates were more attenuated in the east, and we do not find evidence of higher rates among women who had given birth to their first child in the GDR era. Despite the sharp rise in mean age at first birth in East Germany after unification, East German mothers still tend to give birth to their first child at an earlier age than their western counterparts, even though female labor force participation tends to be higher in the East. Hank et al. (2003) suggest that ample provision of institutionalized child care in East Germany is a critical factor to improve the compatibility child rearing and pursuit of a career to young women.

5 Conclusion

With the lowest birth rate in Europe and consistently low fertility rates, Germany faces considerable demographic challenges in the next decades. These may be addressed by several means that encompass an increase in labor force participation, especially among women in West Germany, and incentives for women with higher educational background to raise children. Initiatives already underway include a massive expansion of institutionalized child care in West Germany and a parental leave program with replacement payments proportional to past income rather than lump sum (*Elterngeld*).[20]

The present paper addressed the question whether postponement of births, in the context of increasing human capital investments especially among females, tends to reduce completed fertility in Germany. While there exist numerous studies on determinants of fertility, few have combined a multi-spell framework with information on current human capital. As Kravdal (2007) points out, using education observed at the end of a person's fertility career is prone to reverse causality issues. Similar to studies on Norway and Italy we find postponement in the transition to parenthood among women with higher current stock of human capital (Kravdal 2007, Rondinelli et al. 2006). At the same time, recuperation forces seem to be strong enough to provide women with high human capital trajectories with similar completed fertility as those with lower human capital investments.

A drawback of the present study is that, in contrast to some other empirical studies in the literature, we cannot follow a single cohort throughout their entire fertility history, thus making it necessary to employ sampling from a stock. In addition, the number of births in East Germany is hardly large enough to warrant a detailed separate analysis even though it is clear that the particular fertility pattern in East Germany, similar to that in other Eastern European transformation countries, would require more attention.

[20] Note that this program came into effect in 2007, i.e. after the period considered in our analysis. Thus, the tendency towards recuperation may have even increased in the meantime.

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