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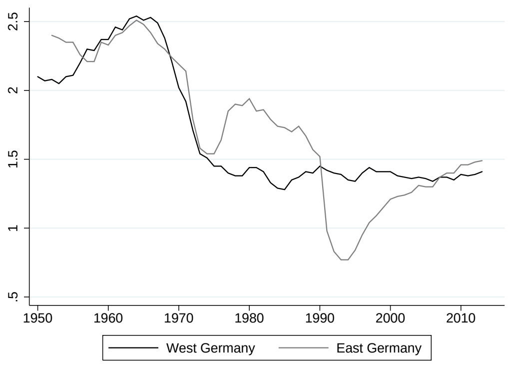
Preface

Germany has a fertility rate far below replacement level, meaning that its population is projected to shrink. Thus, fertility behavior is of highest interest for both policymakers and researchers. Figure 1 shows the development of the total fertility rate (TFR) for West and East Germany. The TFR is a hypothetical measure of how many children a woman would have over her lifetime if she were to experience the current age-specific fertility pattern. Since the baby boom in the middle of the 1960s, the West German TFR has been falling. The replacement rate of 2.1 children per woman was reached for the last time in 1970. Now, the TFR is stagnating at a very low level of around 1.4 children per woman, making it one of the lowest in the developed world. The long-run trend for East Germany looks similar – negative and far below the replacement level. However, East Germany's short-run fertility trend is different than that of West Germany. It has been argued that the tremendous fall in the eastern part of the country's TFR to 0.8 children per woman in the period 1991 to 1994 is partly due to the uncertainty surrounding German Reunification and the transition from a socialist to a market-based economy. Indeed, the German Reunification is an example, among others, of *economic uncertainty*.

Figure 1 : Total fertility rate from 1950 to 2013 in West and East Germany



Source: Destatis (1950-2013)

Apart from demographers and also sociologists who have been interested in fertility behavior for decades, even economists have started to focus on this topic. The economic literature identifies several determinants of the fertility level, including preferences and values (Easterlin, 1973; Easterlin, Pollak, and Wachter, 1980), socioeconomic factors (Becker, 1991; Schultz, 1974), and institutional settings (Lundberg and Pollak, 2007). As pointed out by Becker (1960) early on, economic circumstances and, especially, labor market conditions are of major importance in fertility decisions since they determine the opportunity costs of childbearing. In a very simple framework, couples of reproductive age devote their disposable time either to working in the labor market or to childrearing at home. Therefore, labor market conditions can influence the time allocation within households and, consequently, fertility. The contribution of this dissertation is that it explicitly investigates the socioeconomic consequences arising from economic uncertainty on the labor market.

Measuring economic uncertainty, however, is non-trivial. Situations that some individuals assess as uncertain in the sense of unpredictable future developments might be interpreted by other individuals in a more positive way as either challenging or even beneficial. At the individual level, uncertainty can be measured subjectively or objectively. For instance, individuals can subjectively assess the economic situation in general as well as the security of their own particular employment. However, it is not clear whether self-assessed uncertainty is comparable across individuals. Thus, an often preferred way of measuring uncertainty is to use observable and objective characteristics. More precisely, spells of unemployment and atypical employment, like marginal or fixed-term employment, are considered. Unemployment – and, to some extent, marginal employment – leads to career interruptions and uncertainty about future job and income prospects, whereas the economic uncertainty aspect of fixed-term employment involves the unemployment risk once the contract ends. At the aggregate level, economic uncertainty is basically measured by national or local economic conditions. One way of measuring economic uncertainty at the aggregate level is the unemployment rate in a certain area. For employed individuals in that area, it represents the risk of becoming unemployed. Thus, not the realized shock, but the inherent risk of becoming unemployment causes the uncertainty.

This dissertation is comprised of four stand-alone research papers in which I analyze the socioeconomic consequences of economic uncertainty. At the aggregate level, I focus on local labor market conditions and how they affect period fertility measures in the short run (Chapter 1) and cohort fertility in the long run (Chapter 2). At the individual level, I empirically investigate the effects of starting a career with a fixed-term contract on the quantum and timing of fertility (Chapter 3) and on health conditions and well-being (Chapter 4). The remainder of this preface contains nontechnical summaries of the four chapters.

Chapter 1: Fertility and Local Labor Market Opportunities

In Chapter 1, I investigate, at the aggregate level, how local labor market conditions affect birth rates in German travel-to-work areas.

Graphical evidence suggests a correlation between current labor market dynamics and fertility trends. However, very little is known about the true relationship between unemployment and fertility rates at an aggregate level. I use German administrative birth and unemployment data from 1997 to 2011 to clarify the role played by local labor markets in fertility levels. My approach is based on fixed effects regressions as well as instrumental variable estimations making use of a shift-share index that models changes in the local labor demand. The key findings are as follows. There is a significant negative impact of local unemployment on fertility. I find strong indications that the negative effects on fertility are permanent and not driven by the postponement of births. Consistent with economic theory, increases in the gap between male and female unemployment are associated with lower fertility rates. Moreover, heterogeneity analyses show that the findings hold particularly in the West German travel-to-work areas and for the subset of native Germans. The main estimate suggests that in the period from 2001 to 2004, when unemployment rose on average by 1.9 percentage points, Germany lost approximately 50,000 children due to unfavorable developments on the labor market.

This chapter makes several contributions to the field. First, I use monthly data on a very precise regional level that allows me to link labor market conditions and fertility behavior more closely than has been done in previous work. Second, the use of an instrumental variable approach in the German context is unique. Finally, extant research ignores the fact that the short-run labor market effects could be entirely due to changes in the timing of fertility. I present extensive evidence that a mere tempo effect on childbearing is very unlikely in the German context.

Chapter 2: The Long-Run Consequences of Unemployment Experience on Fertility

In Chapter 2, I empirically assess how state-level unemployment is related to completed fertility and childlessness for the female birth cohorts from 1954 to 1967.

Extant fertility research mainly focuses on the short-run relationship between local labor market conditions and birth rates; not much, if any, work has been done on the long-run effects of unemployment on fertility. Chapter 2 goes some distance in addressing this oversight by examining the consequences of experiencing high levels of local unemployment over the fertile age on the number of children and the incidence of childlessness at age 40. Applying standard estimation techniques to a sample of women from the birth cohorts 1954 to 1967 (data from the Microcensus 2008 and 2012) shows that unemployment experience averaged over five-year age intervals does indeed matter for women's fertility behavior: increasing female unemployment rates during early career years significantly increase fertility, whereas rising male unemployment rates have the opposite effect. This relationship is mainly driven by changes in the probability of remaining childless. For instance, if the average female unemployment rate for the age range of 20 to 24 increases by 1 percentage point, the likelihood that a woman remains childless decreases by 1.6 percentage points. However, if the male unemployment rate rises by the same amount in the same period, childlessness increases by 1.2 percentage points. Since most couples do not make decisions about having children before age 25, we argue that our results represent a "scaring" effect of unemployment that influences future fertility behavior. Two mechanisms may explain the findings: first, unemployment rates have a substantial impact on marriage market outcomes, that is, the likelihood of marrying, and, second, the level of unemployment has an impact on household income.

The combined results of Chapters 1 and 2 have an important message for policymakers. We show in detail that labor market conditions interfere with fertility behavior. Adverse labor market conditions influence the level of fertility in the short as well as the long run. However, men and women react differently to changes in unemployment rates. If the job market prospects for women worsen, they increase fertility, whereas poorer conditions for men reduce fertility. Therefore, in the event of rising unemployment rates, a well-designed family-oriented labor market policy should attempt to minimize the negative consequences of reduced income and reduce the opportunity costs of childbearing.

Chapter 3: Fixed-Term Employment and Fertility: Evidence from German Micro Data

In the third chapter, my co-author, Natalia Danzer, and I study the short- to medium-run effects on subsequent fertility of starting a career with a fixed-term contract.

Fixed-term employment has become tremendously popular in the German labor market. By 2012, almost every second new employment contract was of limited duration. Previous research often discusses the employment and income effects of fixed-term employment but ignores possible spill-over effects to other domains of life. Therefore, we close this gap by analyzing the effects of starting a career with a fixed-term contract on timing of first birth and number of children. We focus on career start since we expect that fixed-term contracts and their inherent economic uncertainty imply a path dependence, setting individuals on career paths that are characterized by repeated spells of temporary employment, lower income progression, and higher risk of unemployment.

Descriptive evidence suggests that fixed-term employment is indeed associated with economic uncertainty and that having children requires secure economic conditions. In our multivariate analysis we compare women with either a permanent or a temporary first contract in regard to their fertility behavior during the first 10 years after they entered the labor market. Based on rich data from the German Socio-Economic Panel, which provides comprehensive information about individuals' labor market history as well as their fertility, our main results are the following. Women tend to postpone first birth when they enter the labor market with a fixed-term contract and reduce the number of children in the first 10 years after graduation. These associations are strongest in the subsample of native women with at least vocational training but no university degree. In contrast, we find no significant correlations for men. Results are robust to the inclusion of a large set of control variables and a number of sensitivity checks. In addition, based on observable characteristics, we find no evidence that certain women tend to select into fixed-term employment.

The main contribution of this chapter is its explicit focus on the type of first contract and the inherent path dependence caused by starting the career with a fixed-term contract. We conclude that fixed-term employment disproportionately affects the young generation (i.e., women of reproductive age). Therefore, policymakers should strive for a more equal distribution of the costs associated with flexible labor markets across population subgroups.

Chapter 4: Health Consequences of Starting a Career with a Fixed-Term Contract

Chapter 4 is a follow-up to Chapter 3 in which I study the short- to medium-run effects on subsequent health outcomes of starting a career with a fixed-term contract.

Official health insurance statistics provide evidence that mental health issues are a major concern in Germany. The 2014 report of the company health insurance fund (BKK) contains statistics suggesting that absenteeism due to mental illness has increased rapidly – since the 1970s, absence days per insured person have quintupled. Therefore, I investigate in this last chapter how the type of first employment contract affects health and well-being during the first five years of an individual's career.

Again, I make use of the German Socio-Economic Panel, which has provided information about mental and physical health conditions since 2002. The main analysis shows that women whose first employment contract is of the fixed-term type tend to report worse mental health in the short run compared to women who start their career with a permanent contract. This relationship is driven by the subjective perception of stress and pressure in these jobs, fades out over time, and is strongest in the sample of women with secondary education. However, economic uncertainty due to fixed-term employment has the opposite effect on men's mental health. At the beginning of the careers, men do not appear to be affected by economic uncertainty, but starting in their third year in the labor market, men report significantly better health outcomes when their first contract was of limited duration and not a permanent one. The path dependence consequent to starting a career with a fixed-term contract is the main mechanism explaining our findings. Men's and women's physical health is not affected at all. The results are robust to the inclusion of a large set of control variables and a number of sensitivity checks. In addition, we find no evidence that certain women or men tend to select into fixed-term employment based on observable characteristics.

If economic uncertainty due to fixed-term employment at the beginning of the career is associated with poor mental health conditions for women, it means that such contracts are accompanied by unintended costs. Since mental health problems are a major reason for absenteeism from work, these costs are also incurred by health insurance companies and employers. Similar to the policy implications of Chapter 3, it appears that making the labor market more flexible implies unintended costs for young women that should be taken into consideration when evaluating the benefits of fixed-term employment.

Keywords: Fertility behavior, economic uncertainty, panel data, instrumental variable, cohort fertility, fixed-term contracts, career start, timing of first birth, mental health, life satisfaction, SOEP, bias correction

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1 Fertility and Local Labor Market Opportunities

1.1 Introduction

Nearly every developed country is now suffering from the problems that accompany fertility rates far below the replacement level, perhaps one of the most troublesome of which is that in an ageing society, shrinking population threatens stable growth and a sustainable social security system. Economic factors play an important role in fertility decisions and economic uncertainty is seen as an obstacle to young couples having children. During the recent global financial crisis, the relationship between business cycles and fertility became of renewed interest to both the public and the scientific community. Based on the latest cross-country fertility trends, *The Economist* concluded that the current recession is having a dampening effect on birth rates, as witnessed by its article entitled "Europe's Other Crisis: Recession is Bringing Europe's Brief Fertility Rally to a Shuddering Halt" (*The Economist*, June 30, 2012). Given the already very low levels of fertility in many European countries, this is not good news. However, will the assertion that economic downturns have a negative impact on fertility pass a thorough empirical examination? That is what this chapter intends to analyze.[1]

Standard microeconomic theory of fertility, which dates back to Becker (see, e.g., Becker, 1960, 1965, 1991) does not predict an unambiguous negative effect of increased economic uncertainty on fertility. In Becker's work, children are modeled as normal consumption goods and fertility decisions are based on the relative costs and benefits of having children. The overall effect on fertility of a recession characterized by increasing unemployment rates is the result of two opposing effects and can be positive or negative. On the one hand, demand for children will fall if unemployment leads to a permanent reduction of wages and family income (income effect). On the other hand, lower wages reduce the opportunity costs of time required for childrearing, which should thus increase the demand for children during spells of unemployment (substitution effect). The overall effect of income on fertility depends on the relative size of these opposing income and costs of time effects. In many countries women traditionally devote more time to childrearing than do men, and hence the opportunity costs argument applies mainly to women. As a result, worsening economic conditions for men due to increased unemployment risk and lower wages are expected to lower fertility rates, whereas worse labor market conditions for women are expected to affect fertility through a negative income and a positive substitution effect. However, the extent to which birth rates respond to changing labor market conditions is ultimately an empirical question that we will address in this chapter.

[1] This chapter is based on a research proposal by Wolfgang Auer, Natalia Danzer, and Helmut Rainer that was submitted to the Fritz-Thyssen-Stiftung in 2013.

The question of whether and, if so, how labor market opportunities affect fertility has attracted scientific interest for many decades. In general, the findings of most – but not all – existing empirical studies suggest that economic booms are associated with higher birth rates, whereas economic downturns are associated with lower birth rates (Sobotka, Skirbekk, and Philipov, 2011).

The extant research on this topic can be divided in two types of studies. The first examines the effect of national unemployment rates on fertility outcomes in a cross-country or a cross-region framework. The results for Europe (e.g., Adsera, 2005, 2011 for 13 European countries without Germany) as well as for Latin America (e.g., Adsera and Menendez, 2011 for 18 Latin American countries) show that higher levels of unemployment are associated with decreases in fertility and a delay in childbearing. For the United States, Schaller (2016) explicitly links birth rates to aggregate unemployment rates in a causal way. Using gender-specific shift-share instruments, she shows that birth rates rise when labor market conditions for men improve, and fall when such conditions become better for women. Currie and Schwandt (2014) look at the short- and long-run consequences of state-level unemployment rates for fertility outcomes. They conclude that higher unemployment implies lower fertility in both the short and long term that is mainly driven by increased levels of childlessness. In contrast, Dehejia and Lleras-Muney (2004) do not find any significant effect of state unemployment rates on regional birth rates in the United States.

The second type of study investigates individual-level fertility effects of economic uncertainty. Kreyenfeld (2010) examines whether German women postpone childbearing in response to economic uncertainty. While she does not find any significant effects for the pooled sample, a subgroup analysis by educational level reveals that low-educated women who are unemployed have higher birth rates than do employed women. A more reliable identification strategy is applied to Austrian data for the years 1990 to 1998 by Del Bono, Weber, and Winter-Ebmer (2012). They find a significant and robust reduction in fertility due to career interruptions. In a follow-up paper, Del Bono, Weber, and Winter-Ebmer (2015) confirm that displacement from a career-oriented job is detrimental for fertility but find that unemployment spells per se do not cause a drop in fertility. In a broader context, Schmitt (2012) analyzes the impact of unemployment and precarious employment (fixed-term contracts, part-time work) on individuals' fertility choices using German and U.K. data. His results suggest negative effects of atypical employment on fertility in Germany but not in the United Kingdom, and positive effects of female unemployment on fertility in both countries. However, most of the papers discussed above do not account for either the fact that fertility and labor supply decisions are interrelated, that changes in the unemployment rate might be caused by fertility-induced changes in labor supply, or that changes in the timing of births might explain the estimated effects.

Hence, we intend to answer the following questions. First, is there a causal effect of local labor market conditions on fertility rates? Second, is the impact of local labor market conditions on fertility temporary or persistent? Third, are fertility rates differentially affected by male and female labor market opportunities? In particular, we advance the field by using detailed administrative vital statistics data from Germany to investigate the causal link between local labor market conditions and fertility rates and by analyzing whether there are differences in fertility responses with respect to changing male versus female labor market opportunities as economic theory would suggest. Finally, this is the first paper that explicitly examines whether the estimated effect is temporary or persistent.

Since unemployment might be endogenous, the empirical identification of a causal effect on fertility requires exogenous shifts in labor demand. Our identification is closely related to that employed by Schaller (2016), but expands her approach. Using birth and employment data aggregated on the level of 244 travel-to-work areas (TTWAs) from 1997 to 2011, we instrument for the local unemployment rate by using an industry shift-share indicator of labor demand – the Bartik IV, named after Timothy J. Bartik who proposed the use of labor demand indices (Bartik, 1991). Since, potentially, births are postponed in recessions and then pursued in the subsequent recovery, we explicitly discuss whether we are estimating an effect on the timing of births or an actual quantum effect. In addition, we test existing theory on gender-specific effects.

Our results suggest that local labor market conditions have a negative impact that reduces, on average, monthly births per 1,000 women by 0.5 percent (FE) to 0.9 percent (IV). According to this finding, the burst of the dot-com bubble after 2000 prevented the birth of more than 50,000 babies in a four-year period. We also find that the extensive margin as well as the intensive margin is affected: first, second, and higher-order births become less likely if economic circumstances worsen. To check potential changes in birth timing we investigate the effects on age-group-specific birth rates as well as age at birth and find no evidence that births are completely postponed. In support of the above-discussed theory, we find that a rise in the gap between male and female unemployment rates reduces local fertility rates, a strong indicator for different effects of male and female unemployment rates. For women, the substitution effect seems to be more pronounced than the income effect. Heterogeneity analyses show that the results are driven by native women in West Germany.

The remainder of the chapter is organized as follows: Section 1.2 describes the data and the empirical approach. Section 1.3 shows results for quantum as well as tempo effects of changes in local labor market opportunities. Gender-specific effects are presented in Section 1.4 and further heterogeneity in Section 1.5. Section 1.6 concludes.

1.2 Data and Method

1.2.1 Data

To conduct the empirical analysis we construct a regional panel dataset for the period 1997 to 2011 in which we match monthly regional birth register data with indicators of local labor market conditions and other relevant region-specific characteristics. All data (except for the vital statistics) can be found on the regional database of the German Statistical Office (regionalstatistik.de) or the statistic website of the Federal Employment Agency (statistik.arbeitsagentur.de). The unit of analysis is local labor markets as classified by the Federal Institute for Research on Building, Urban Affairs and Spatial Development (BBSR), hereafter called travel-to-work areas (TTWAs). A complete list (Table A.1) and map (Figure A.1) of all 258 German TTWAs can be found in the Appendix.

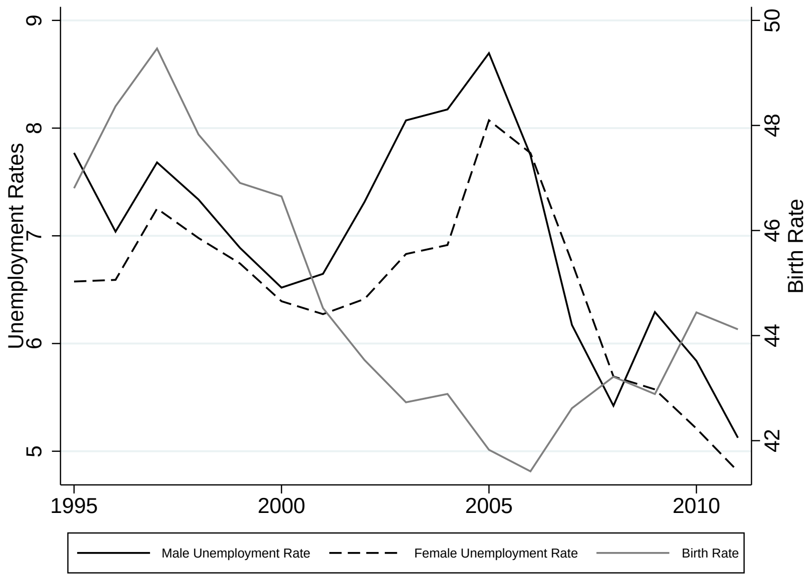
Fertility data are from the German Birth Register, which contains information on all birth certificates in one calendar year, covering more than 650,000 annual births. We collapse the individual birth data on TTWA-year-month cells and merge the sociodemographic information. The birth certificates contain information about county of residence, age of mother, and birth order within a given marriage. Using this information, we calculate regional fertility rates by age of mother and by birth order. Our main dependent variable is the birth rate per 1,000 women, that is, the sum of births in a TTWA during a month per 1,000 women of reproductive age (15 to 44 years). Similarly, the age-specific birth rates are defined as the sum of births per 1,000 women of a specific age relative to all women in the specific age group. Within a given marriage, the data contains a count variable that identifies the birth order. Dividing the sum of first, second, and third and higher-order births by the number of women of reproductive age ($\times$ 1,000) allows investigating the effects of local unemployment risk on fertility at the extensive as well as the intensive margin.

The explanatory variable of main interest is the (local) unemployment rate. The unemployment rate is a good proxy for general economic activity since unemployment directly affects individuals and is less likely to be endogenous, unlike, for example, GDP growth or individual wages. We use administrative unemployment data from the labor statistics of the Federal Employment Agency. To obtain a measure of local labor market conditions we divide the number of unemployed individuals in the TTWA by the local working-age population (i.e., men and women between 15 and 64 years of age). Given the availability of the data (regional level and long time horizon), this is the best approach for constructing the unemployment rate even if it is not completely in line with the definition used by the Federal Employment Agency.[2] Since both the birth rates and the unemployment rates are available on a monthly basis, we can lag the unemployment rate by nine months to control for the fact that the decision to have a child is made at conception rather than at birth.

As a first step, Figure 1.1 shows the development of unemployment rates (by gender) as well as the birth rate over time – each series aggregated by year and Germany as a whole. The figure reveals that the birth rate was relatively high in 1997, with almost 50 births per 1,000 women of reproductive age, but dropped drastically by 2006. Since then, there has been an

[2] The German Federal Employment Agency sets the number of unemployed relative to the total work force defined as the sum of unemployed and the employed individuals. Inactive individuals are not counted. Thus, our definition underestimates the true level of unemployment. This could be problematic if the share of inactive individuals varies systematically across regions.

Figure 1.1 : Development of gender unemployment rates and birth rate



Source: German Birth Register (1996-2011), Labor statistics of the Federal Employment Agency (1996-2011).
Notes: Germany, 1996-2011. Unemployment rates are defined as number of unemployed women or men divided by the respective working-age population (15-64) multiplied by 100. Birth rate is defined as number of births per 1,000 women of reproductive age (15-44). All rates are yearly averages for Germany.

upward trend interrupted by a kink in the most recent year. From 1995 to 2000, the birth rate follows the development of the unemployment rates. However, since 2000, the graph shows that unemployment and fertility rates go in the opposite direction. In the year after unemployment rates peak at 9 percent, fertility hits the bottom. In the subsequent period of economic recovery, unemployment rates decrease and the level of fertility goes up again.

In the main analysis, we focus on birth and unemployment rates at the regional level, which gives us several advantages. Compared to cross-country studies employing aggregate national unemployment rates that might be confounded by general country-specific time trends, our analysis exploits regional variation in economic uncertainty across space and over time within Germany. This not only allows assessing the importance of local labor market conditions on regional fertility levels, but also provides more variation and statistical power for the empirical analysis. In contrast to micro-level studies that focus on individual unemployment or job loss incidence, regional unemployment rates are theoretically better proxies for more general economic uncertainty (Dehejia and Lleras-Muney, 2004).

A TTWA combines one or more counties into a regional unit based on certain prerequisites, including that one-way commuting time is less than 45 minutes as well as that there have to be jobs for at least 65 percent of the labor force in a TTWA. Thus, a TTWA is a restricted, albeit not exclusive, area where economic activity is concentrated. Even though an analysis at the county level would give more regional variation and hence more statistical power, the main advantage of our approach is that the unemployment rate is expected to be far more representative on the regional than on the county level. For instance, a significant share of workers live in the countryside and commute to urban centers where jobs are located. Assigning the unemployment rate of their county of residence rather than their county of work would bias the relevant measure of labor market conditions. There are 258 TTWAs in Germany but due to several local government reorganizations (mainly in Eastern Germany), we have to aggregate some of them to ensure a consistent definition of the regional units of observation.[3] The full sample is comprised of 244 TTWAs, that we observe monthly for 15 years, summing to 43,920 observations. Due to the nine-month lag between measurement of unemployment and fertility, the final sample size is 41,710 TTWA-month observations.[4]

Obviously, there is some variation in the sociodemographic characteristics within and across the TTWA. These factors are likely to affect the level of fertility in an area as well as the level of unemployment and therefore confound our estimates. Thus, it is essential to include them in our regression to eliminate this source of endogeneity. Control variables are age structure of population, population density, and share of migrants of reproductive age (15-44 years). All controls are available only as yearly averages. To avoid jumps in the covariates only from December to January each year, we interpolate the variables over all 12 months assuming a linear development in the demographic controls.[5] Table 1.1 shows descriptive statistics for the outcome as well as the control variables.

On average, there are 3.7 monthly births per 1,000 women of reproductive age, which adds up to 44.4 births per year and 1,000 women. Age-specific birth rates are low for very young women and women at the end of their childbearing years. The highest rate of 7.9 monthly births per 1,000 women occurs in the age group of 25 to 29 years. Mean age at birth is a little above 29 years. The average monthly unemployment rate is 6.8 percent, with a slightly higher level for men and a slightly lower level for women. Some West German regions exhibit unemployment rates far below 5 percent (with a minimum of 1.7 percent), but some regions, particularly in East Germany, face unemployment rates higher than 20 percent.

[3] Figure A.1 in the Appendix is a map of all TTWAs with markers for those TTWAs that are affected by local government reorganizations.

[4] The actual number is 41,724 but 14 observations have missing unemployment rates. Düsseldorf and Mönchengladbach did not report unemployment numbers for February 2010, and nor did Eisenach for the year 1997.

[5] Both linear interpolation as well as constant values over the year might create some structure in the error term that in turn might be correlated with the unemployment rate and therefore might cause biased coefficients. However, we are convinced that not controlling for demographic factors (measured annually) increases the threat of confounded estimates even more drastically.

Table 1.1 : Descriptive statistics of outcome and control variables

	N	Mean	SD	Min	Max
A. Dependent variables					
Birth rate per 1,000 women 15-44	41,710	3.693	0.599	1.527	7.191
Birth rate per 1,000 women 15-19	41,710	0.927	0.514	0.000	5.204
Birth rate per 1,000 women 20-24	41,710	4.347	1.364	0.282	12.450
Birth rate per 1,000 women 25-29	41,710	7.937	1.759	1.385	18.024
Birth rate per 1,000 women 30-34	41,710	7.010	1.655	1.044	15.816
Birth rate per 1,000 women 35-39	41,710	2.793	1.000	0.000	8.381
Birth rate per 1,000 women 40-44	41,710	0.472	0.302	0.000	2.802
Age at birth women 15-44	41,710	29.183	0.964	24.696	33.150
B. Control variables					
Unemployment rate	41,710	0.068	0.032	0.017	0.213
Male unemployment rate	41,710	0.070	0.032	0.013	0.224
Female unemployment rate	41,710	0.066	0.035	0.016	0.218
Labor demand index	41,710	0.031	0.027	-0.047	0.158
Population density	41,710	299.1	432.8	38.22	3919.9
Share of migrants 15-44	41,710	0.094	0.047	0.009	0.250
Share of women 15-19	41,710	0.146	0.019	0.083	0.194
Share of women 20-24	41,710	0.143	0.016	0.107	0.232
Share of women 25-29	41,710	0.145	0.016	0.095	0.208
Share of women 30-34	41,710	0.166	0.025	0.108	0.225
Share of women 34-39	41,710	0.194	0.017	0.136	0.235
Share of women 40-44	41,710	0.206	0.021	0.144	0.264
Population fraction 45-49	41,710	0.076	0.008	0.057	0.104
Population fraction 50-54	41,710	0.066	0.009	0.040	0.098
Population fraction 55-59	41,710	0.061	0.009	0.038	0.089
Population fraction 60-64	41,710	0.060	0.009	0.036	0.083
Population fraction 65-74	41,710	0.106	0.014	0.072	0.153
Population fraction 75+	41,710	0.081	0.013	0.047	0.132

Notes: Full sample of 244 TTWAs, 15 years, and 12 months; unemployment data are missing for Düsseldorf and Mönchengladbach for February 2010 and for Eisenach for all of 1997.

The demographic control variables differ largely across regions. First, we include the year-of-age shares of women over all women between 15 and 44 years,[6] which ensures that our estimations are not confounded by changes in the size of the relevant female cohort. The younger cohorts are smaller than the older ones in the majority of areas. The declining fertility rates over the last decades explain why cohorts are shrinking over time. Second, the age structure of population changes steadily, resulting in higher shares of people close to or at retirement age. We define the age structure as the number of people of every single age from 45 to 74, and above 75, divided by the total population in the respective TTWA. These variables control for the fact that regions with a higher fraction of older residents might have different preferences and therefore different fertility behavior. Third, TTWAs vary widely in

[6] For a more readable presentation of summary statistics we only show population controls in five-year age intervals. In the empirical analyses we use these controls as year-of-age shares.

population density, which is defined as number of residents per square kilometer. On average, 299 people live on a square kilometer in a German TTWA, but the values range from 38 to over 3,900. Population density is a good measure for the degree of urbanization in a region. Fourth, the share of foreign population possibly influences the level of fertility (Fernández and Fogli, 2006). Hence, we control for the share of migrants of reproductive age, 15 to 44 years. On average, 9.4 percent of our sample has a foreign background. The values range from less than 1 percent to 25 percent of non-native population in a TTWA.

1.2.2 Method

In a first step, we examine the relationship between (local) labor market conditions and fertility by exploiting the variation in the unemployment rate over time within regions. The underlying empirical model is a simple linear regression of the following form:

$$log(FRT_{rt}) = \beta UR_{rt} + \gamma' X_{rt} + \mu_t + \phi_r + \epsilon_{rt}. \tag{1.1}$$

$log(FRT_{rt})$ denotes the natural logarithm of the fertility measures, that is, number of births per 1,000 women, in region r at time t. On the right-hand side of the equation, UR_{rt} is the unemployment rate at time of conception. Consequently, β captures the effect of local labor market conditions on the different fertility measures. Time-varying, region-specific characteristics are combined in the vector of controls, X_{rt}, to account for changes in the demographic composition of a TTWA. Time fixed effects, μ_t, control for dynamics in the fertility rates that are common to all regions. These time controls comprise a full set of dummies for the interaction of years and months to precisely capture the general dynamics as well as seasonal patterns in fertility and unemployment. TTWA fixed effects, ϕ_r, capture all time-invariant differences in birth rates that are unique to any region, r. Finally, ϵ_{rt} is an idiosyncratic error term. Under the assumption that based on the observable characteristics as well as the year and region fixed effects, unemployment rates are exogenous to fertility, we can estimate a consistent β. In other words, there must not be any unobservable time-varying and region-specific characteristics that are correlated with both the birth rate and the unemployment rate.

In a second step, we augment our linear model by implementing a region-specific linear time trend, $\omega_r \times T$:

$$log(FRT_{rt}) = \beta UR_{rt} + \gamma' X_{rt} + \mu_t + \phi_r + \omega_r T + \epsilon_{rt}. \tag{1.2}$$

This approach nets out correlation between unobservable characteristics and birth rates that follow a linear trend over time within region. The identifying assumption changes slightly: we estimate consistent effects of local labor market conditions when time-varying confounders follow a linear time trend after controlling for all regional characteristics and fixed effects.

However, there are still sources of endogeneity that might cause biased estimates, even after controlling for all observable characteristics and including a variety of fixed effects. For example, there could be reverse causation if, for example, a higher propensity to give birth results in lower unemployment since women are no longer registered as unemployed. Second, unobserved heterogeneity across TTWAs might cause an omitted variable bias. Assuming that an unobserved change in preferences induces fertility to fall and labor market attachment to rise, then, mechanically, the fertility rate as well as the unemployment rate (due to larger denominator) decline, resulting in an underestimation of the true (negative) relationship. For these reasons, we look for a source of exogenous variation in regional unemployment rates. Such an instrumental variable must not be correlated with the outcome variable in any way, except through the channel "unemployment rate." This condition excludes the instrument from causal model of interest (exclusion restriction) and requires the instrument to be sufficiently correlated with the endogenous variable (relevance condition). While the earlier condition cannot be tested, we show that the instrument is relevant in the first-stage results (see Table 1.2).

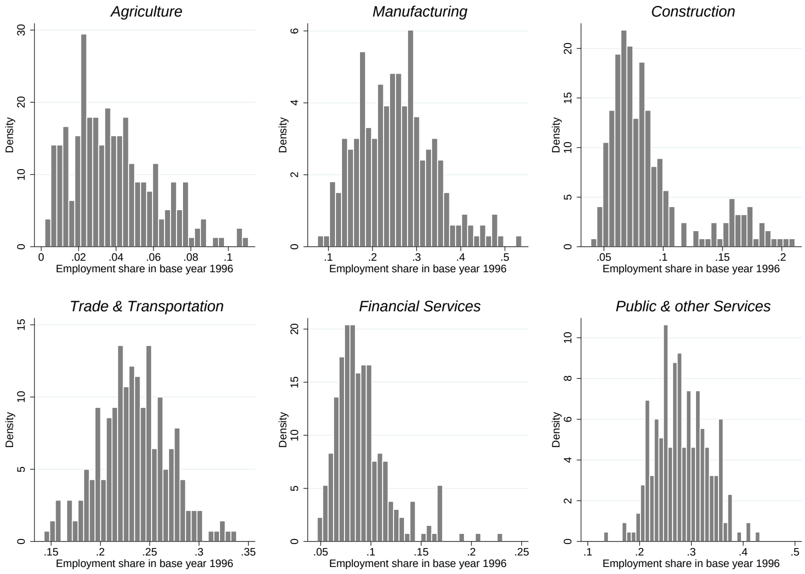
Following Schaller (2016), we propose a shift-share index of labor demand (LDI), which takes advantage of differences in the regional industry structure and differences in employment trends across industries. Traditionally, these shift-share indices are used to instrument for local labor market opportunities if supply as well as demand shifts may influence, for example, employment and unemployment rates (Bartik, 1991; Blanchard and Katz, 1992). Bertrand, Pan, and Kamenica (2013) analyze the consequences of relative income within households, Aizer (2010) studies the influence of the gender wage gap on domestic violence, Gould, Weinberg, and Mustard (2002) investigate the local unemployment effect on crime rates, and, in a comparable setup, Bound and Holzer (2000) look at how labor demand shifts affect employment rates and earnings.

The main idea is that these differences in employment trends are mainly due to changes on the labor demand side (e.g., changes in production technology or product demand) and therefore do not influence the fertility decisions except through changes in the level of unemployment. However, not all industries experience the same employment trajectory. Some industries, such as financial services, are growing faster than others, whereas manufacturing exhibits a negative employment trend, at least before the year 2000.

To construct our instrumental variable, we exploit the variation that arises due to substantial differences in industry mix across German TTWAs and hence the fact that the region of residence might influence the local labor market conditions.[7] Figure 1.2 illustrates regional variation in initial industry composition. For instance, the shares of the construction sector range from below 5 percent to almost 25 percent of the local economy. This is probably due

[7] We use six broad industry categories defined by the German Statistical Office (WZ03): (1) agriculture, (2) manufacturing, (3) construction, (4) trade, transportation, and communication, (5) financial services, and (6) public and other services. In 2008 there was a change in classification (WZ08) with the consequence that industries are not defined consistently over time. For the years 2000 to 2009 we have information about WZ03 as well as WZ08 employment. Thus, we are able to compute weights such that the WZ08 classification after 2009 and WZ03 data are comparable over time.

Figure 1.2 : Histogram of regional employment shares by industry in 1996



Source: Destatis (1996).

to the construction boom in East Germany after Reunification. Trade and transportation, which encompasses also the hospitality industry, is particularly important in regions that have a strong tourism industry, while financial services determine more than 20 percent of employment in Frankfurt and Munich.

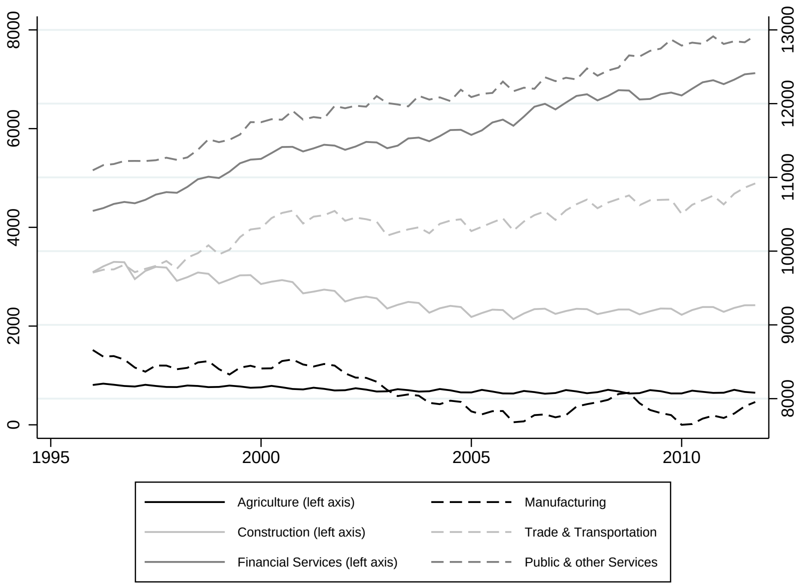
Furthermore, we exploit variation over time that is due to differences in national employment trends across industries. Figure 1.3 illustrates the development of overall employment between 1996, our base year, and 2011. The service sector, which includes "Trade and Transportation," "Financial Services," and "Public and Other Services," exhibits increasing employment throughout the whole period of observation, whereas the share of employment in agricultural and the construction sector steadily shrinks. Employment in manufacturing remains quite stable. Changes in labor demand, induced by technological change or changes in product demand, have a more substantial impact on those TTWAs in which the affected industry has a high share of total employment. Therefore, changes in national employment are likely to also influence the level of local unemployment.

Following Bartik (1991) and Schaller (2016), we construct a variable representing the predicted employment growth to the base period as follows:[8]

(1.3) $$LDI_{rt} = \sum_i G_{it} * \frac{E_{ir0}}{E_{r0}}$$

[8] A very clear derivation of the measure of (changes in) labor demand is provided by Maestas, Mullen, and Powell (2013). The authors also show why the Bartik IV can be regarded as an exogenous measure of labor market conditions and labor market dynamics.

Figure 1.3 : Development of national employment by industry (in 1,000s)



Source: Destatis (1996-2011).
Notes: Classification of industries is based on WZ03. Data after 2009 is based on WZ08 weighted such that numbers are comparable over time. Employment includes both dependent employment and self-employment.

where i is a subscript for industry, r for the regional level, and t denotes the respective time period. G_{it} is the national growth rate to the base period 1996 of employment in industry i in period t. This change is weighted by the relative importance of the industry in the region in the base year 1996, $\frac{E_{ir0}}{E_{r0}}$, and summed over all industries. Summing over all industries minimizes the threat of different people sorting into different industries (Aizer, 2010). Since national trends for these six industries are available only on a quarterly basis, we again interpolate the values for all months within a quarter.[9] The variation we make use of comes from differences in the initial industry structure across regions and differences in the national employment trends over time.[10]

Table 1.2 reports the first stage estimates predicting the local unemployment rate by labor demand shocks. Columns 1 and 2 focus on the overall unemployment rate; Column 3 on the gender gap in unemployment rates, that is the difference between male and female unemployment rates. We return to the gender-specific LDI in Section 1.4. Without TTWA-specific trends, the shift-share indicator does not predict the unemployment very well. Apparently, the change

[9] Total employment includes dependent employment and self-employment. The data come from the national accounts statistics for Germany and are available through the website of the Federal Statistical Office (destatis.de).
[10] The literature uses several variations of the initial labor demand index proposed by Bartik (1991). For instance, we interacted the change in labor demand with the initial employment rate. The coefficients are very similar to what we show in the following section but in some cases not as precisely estimated. Moreover, an interaction of oil price shocks and regional employment in the manufacturing industry is not a good instrument for local unemployment rates (Raphael and Winter-Ebmer, 2001).

Table 1.2 : First stage results for different unemployment measures

Dependent variable	Unemployment rate		Gender gap in unemployment
	(1)	(2)	(3)
Labor demand index	-1.7855***	-2.0945***	
	(0.1949)	(0.1934)	
Gender gap in labor demand index			-3.9511***
			(0.9163)
1st stage F-stat	83.92	117.5	18.59
prob>F	0.000	0.000	0.000
Controls	No	Yes	Yes
TTWA time trends	Yes	Yes	Yes
Observations	41,710	41,710	41,710

Notes: All regressions contain TTWA and year×month fixed effects as well as the year-of-age shares of 15- to 44-year-old women over all women aged 15 to 44. Control variables include population density, share of migrants of reproductive age, and the year-of-age shares of 45- to 74-year-old and 75+ years old people over the population in each TTWA. TTWA time trends are region-specific linear trend variables. Robust standard errors clustered on TTWA level in parentheses, *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

to the base year in employment and the unemployment rate follow a common trend, resulting in a spurious correlation. If we account for this trend and include region-specific time effects, as in Columns 1 and 2, the labor demand index seems to be a strong and relevant instrument for unemployment rates. As expected, the sign of the correlation is negative, meaning that a positive labor demand shock reduces unemployment and vice versa. On the one hand, our instrument is valid only conditional on the regional trend variables. On the other hand, the change to the base year is less vulnerable to sorting of individuals into industries. For ease of interpretation we standardize the instrumental variable. Thus, a one standard deviation shock in labor demand reduces the local unemployment rate by 1.8 to 2.1 percentage points, depending on the specification.

To estimate the relationship between local birth rates and local labor market opportunities, we apply standard fixed-effects estimation (FE) techniques with robust standard errors clustered at the level of TTWA to account for correlation within TTWAs. The correct specification of the variance-covariance matrix is a major challenge in this setup since there is potential correlation in the error term in a panel with a large time dimension. First and foremost, the large set of time controls should capture national dynamics. Second, including TTWA-specific effects models the time-constant part in the error terms, and third, the TTWA-specific trend accounts for regional dynamics that would otherwise have been a systematic part of the error term.

Finally, following Cameron and Miller (2015), we run regressions with clustered as well as HAC standard errors to avoid misleading inferences due to autocorrelated error terms.[11]

1.3 Main Results

1.3.1 Quantum Effects: Birth Rate

Results from regressions of the log birth rate on the local unemployment rate[12] are reported in Table 1.3. Panel A shows estimates from FE models; Panel B the IV estimates. The first column shows the coefficient from a specification without demographic controls but year$\times$month and TTWA fixed effects. Moving right across the table, we first add control variables, then the TTWA-specific linear time trend, and finally, in the full specification, both control and trend variables. To reflect differences in cohort size that may influence the birth rate, we control for the year-of-age shares of 15- to 44-year-old women over all women between 15 and 44 years. The set of regional controls consists of population density and the share of migrants of reproductive age, as well as a set of variables modeling the age structure, that is, the year-of-age shares of 45- to 74-year-old and 75+ year-old people over the total population in each TTWA.

Across all specifications the correlation between the birth rate and the local unemployment rate is negative and significantly different from zero. The estimate in the first column of Table 1.3 is quite large but shrinks when we include convariates in Column 2 or region-specific trend variables in Column 3. Thus, parts of the initial correlation can be explained by time-varying characteristics of the TTWAs, such as changing size of female cohorts. The preferred specification – that with trend and control variables – in the last column allows for unobserved characteristics that follow a linear trend. Compared to the specification in Columns 2 and 3 the coefficient is larger in magnitude, meaning that leaving these controls in the error term causes a bias toward zero. The coefficient suggests that a 1 percentage point increase in the local unemployment rate leads to an approximately 0.5 percent decrease in the birth rate. To net out a possible endogeneity bias, we run IV regressions and present the estimates as well as the first-stage F-statistic for the labor demand index in Panel B of Table 1.3. We refrain from showing results for Columns 1 and 2 since the instrument is not valid without TTWA trends. The effects are somewhat larger in magnitude than in the FE setup, as expected given the predicted direction of bias toward zero from reverse causation and unobserved heterogeneity. The relationship weakens when we add the control variables. Thus, at least parts of the effects are due to differences in regional fertility patterns over time. Our preferred specification in

[11] Stata's xtivreg2 by Schaffer (2005) allows for clustering as well as heteroskedasticity and autocorrelation-robust (HAC) standard errors using a kernel estimation for the variance-covariance matrix. Since HAC standard errors tend to be smaller, we report the more conservative estimates and show clustered standard errors in all tables. For more information about standard error issues, see also Angrist and Pischke (2008).

[12] To simplify interpretation of the estimates we multiply the unemployment rate by 100. Thus, a marginal change in the unemployment rate implies a 1 percentage point increase instead of an increase from 0 to 100.

Table 1.3 : Effects on birth rate

Dependent variable	(log) births per 1,000 women			
	(1)	(2)	(3)	(4)
Panel A: FE				
Unemployment rate	-0.0296***	-0.0022*	-0.0038***	-0.0051***
	(0.0035)	(0.0012)	(0.0011)	(0.0011)
Panel B: IV				
Unemployment rate			-0.0125**	-0.0086**
			(0.0049)	(0.0042)
1st stage F-stat			83.92	117.5
prob>F			0.000	0.000
Controls	No	Yes	No	Yes
TTWA time trends	No	No	Yes	Yes
Observations	41,710	41,710	41,710	41,710

Notes: All regressions contain TTWA and year×month fixed effects. Control variables include the year-of-age shares of 15- to 44-year-old women over all women aged 15 to 44 as well as population density, share of migrants of reproductive age, and the year-of-age shares of 45- to 74-year-old and 75+ year-old people over the population in each TTWA. TTWA time trends are region-specific linear trend variables. Robust standard errors clustered at the TTWA level in parentheses, *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Column 4 shows that a 1 percentage point increase in the local unemployment rate lowers the birth rate by almost 0.9 percent.

1.3.2 Quantum Effects: Birth Rate by Birth Order

Next, we investigate whether fertility responds to changes in unemployment at the extensive or the intensive margin. Therefore, we construct birth-order-specific fertility rates, defined as the number of first, second, and third and higher-order births divided by the number of women of reproductive age (multiplied by 1,000). Table 1.4, Panel A presents the standard FE results of the specification with demographic control variables and TTWA time trends; Panel B shows the corresponding IV estimates.

The effect of local unemployment on fertility is the strongest among formerly childless women entering parenthood in the FE specification: a 1 percentage point increase in the unemployment rate decreases the number of first births per 1,000 women of reproductive age by 1.2 percent. The estimated association is highly significant and very similar to the causal estimator in Panel B. The coefficient for second births is somewhat smaller but still significant. For third and higher-order births, the sign is reversed and the coefficient imprecisely estimated in the FE estimations. Looking at the IV results in Panel B changes the findings since the estimates suggest that both the extensive and the intensive margin are affected. Higher levels of local unemployment seem to decrease fertility rates at the extensive margin: first

Table 1.4 : Effects on birth-order-specific birth rate

Dependent variable	(log) births per 1,000 women		
	(1) 1st births	(2) 2nd births	(3) 3rd and higher order births
Panel A: FE			
Unemployment rate	-0.0116***	-0.0066***	0.0013
	(0.0024)	(0.0025)	(0.0028)
Panel B: IV			
Unemployment rate	-0.0155*	-0.0332***	-0.0246**
	(0.0087)	(0.0083)	(0.0122)
Controls	Yes	Yes	Yes
TTWA time trends	Yes	Yes	Yes
Observations	41,710	41,710	41,710

Notes: All regressions contain TTWA and year×month fixed effects. Control variables include the year-of-age shares of 15- to 44-year-old women over all women aged 15 to 44 as well as population density, share of migrants of reproductive age, and the year-of-age shares of 45- to 74-year-old and 75+ year-old people over the population in each TTWA. TTWA time trends are region-specific linear trend variables. Robust standard errors clustered at the TTWA level in parentheses, *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

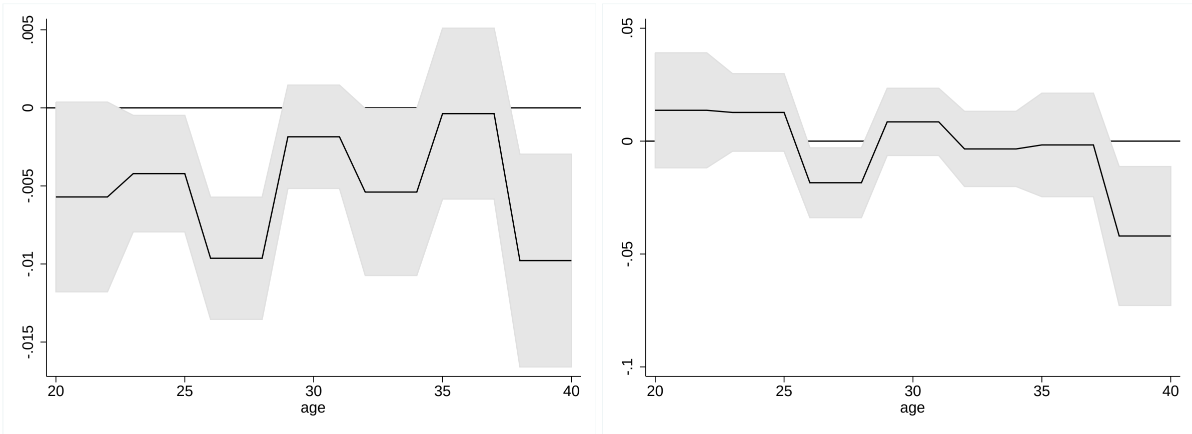
births are reduced by 1.6 percent. However, higher-order births respond even more strongly to changes in local unemployment: the coefficient of -3.3 percent on second births is more than twice as large compared to the effect on first births. Since not only childless women seem to reduce fertility but also mothers who already have one or more children, we interpret these results as first evidence against the postponement hypothesis. Nevertheless, we take a closer look at the timing of births in the next sections.

1.3.3 Tempo Effects: Age-Group-Specific Birth Rates

So far, we have shown that increases in unemployment cause birth rates to fall. However, our measure of fertility, the birth rate per 1,000 women of reproductive age, is a period measure rather than a cohort measure of fertility (Bongaarts and Feeney, 1998). As such, it is suitable for analyzing the influence of business cycles on fertility but, in contrast to cohort fertility measures that give the actual number of births per woman measured after the reproductive-age period, it can be distorted by tempo effects, that is, by changes in the timing of births (Bauernschuster, Hener, and Rainer, 2015). The substantial impact of changes in the unemployment rate at both the extensive and intensive margins is only a weak indication that there is no tempo effect.

Thus, we look for empirical evidence that allows us to distinguish actual quantum effects from mere tempo effects. To this end, we make use of both the standard FE specification and the IV

Figure 1.4 : Effects on birth rate by age groups



(a) FE

(b) IV

Notes: Coefficients and confidence intervals for regressions of (log) birth rate on unemployment rate by age group of mothers. All regressions contain TTWA and year×month fixed effects. Control variables include the year-of-age shares of 15- to 44-year-old women over all women aged 15 to 44 as well as population density, share of migrants of reproductive age, and the year-of-age shares of 45- to 74-year-old and 75+ year-old people over the population in each TTWA. Region-specific linear trend variables are included. The gray area marks the 90 percent confidence interval calculated using standard errors clustered at the TTWA level.

strategy, and adjust the outcome variable to allow for heterogeneous effects along the age distribution of mothers. Specifically, we split the dependent variable into separate variables in such a way that each of them measures the birth rate for three-year age groups from 20-22 up to 38-40. More precisely, the age-group-specific birth rates are defined as the number of births by women in each of the age groups, divided by the total female population in the respective age group. If our results are not just driven by tempo effects, induced by women postponing childbearing, we should observe negative effects of adverse local labor market conditions on fertility across all cohorts. Figure 1.4 presents the coefficients and confidence intervals from seven separate regressions for the age-group-specific birth rates.

The effects differ across age groups of mothers. In both the FE (Panel A) and IV (Panel B) regressions, women at the lower and upper ends of the age distribution appear to react more strongly to changes in the local unemployment rate. As expected, the two panels are similar but, as was previously noticed in Table 1.3, the magnitude of the IV coefficients is somewhat larger. For women below 29 years, the FE estimates indicate a significant negative effect, with the exception of women 20 to 22 years old, for which the coefficients just fail to reach significance. The strongest effect is found for women between 26 and 28 who reduce fertility by almost 1 percent (IV in Panel B). Both graphs show that fertility decreases for older women at the end of their fertile age. That both older and younger women are affected is some evidence against a postponement of births. However, in the prime fertility age between 29 and 34, point

IV estimates are positive albeit not significant. This can be explained by a catching-up effect of births postponed earlier in life. Women older than 35 years have lower completed fertility because, due to biological reasons, postponing births is not as possible for them as it is for younger women. However, since birth rates among women above 35 years are very low, the overall effect on births is not substantial. One reason why women below 30 years respond more strongly to higher unemployment rates than do women above 30 years might be that older women are already better integrated in the labor market and thus react less sensitively to changes in local labor market conditions since their own employment situation is more stable relative to younger women. We look at the heterogeneous effect by the employment status of mothers in Section 1.5. In the end, the age-group-specific analysis does not completely solve the timing puzzle even if the results do suggest rather a substantial fertility reduction.

1.3.4 Tempo Effects: Birth-Order-Specific Age at Birth

To look more deeply into this problem, we now estimate the effect of changes in local unemployment on mothers' age at birth. A plausible reason for a decrease in the total number of children in a cohort is a decline in higher-order births, which makes the average age at birth decrease as well. However, this kind of age effect would not be considered as a mere tempo effect. Thus, we want to test whether local labor market conditions affect mothers' age at first, second, and third and higher-order births in the same way. If women postpone childbearing in response to unemployment shocks, we would expect the coefficients in Table 1.5 to be significantly positive. FE results in Panel A do not show any differences in age at birth due to changes in local unemployment rates either for all births or by birth order. In contrast, Panel B suggests that the overall age at birth marginally decreases on average by 0.06 years. This effect is small in size and unexpected in sign since a postponement should lead to higher age at birth. Looking at birth-order effects reveals no significant change in the birth age even if the sign for the first birth coefficient is positive. Thus, instead of decelerating fertility, rising unemployment seems to have inspired younger women to have children even though we do not see any effect at the intensive margin.

Since we are not able to follow all women over their fertile lifecycle, we cannot infer that the estimated effect eventually leads to reduced completed fertility. Moreover, unemployment typically follows cyclical patterns, with increasing unemployment in downturn periods and falling unemployment in upswing periods. With the linear estimator we employ, it is not possible to distinguish between boom and bust periods; hence, the effect is the same independent of how the unemployment rate changes. Nevertheless, all evidence shown in this section points at a quantum effect of local unemployment on birth rates.

Table 1.5 : Effects on birth-order-specific age at birth

Dependent variable	Age at birth			
	(1) All births	(2) 1st births	(3) 2nd Births	(4) 3rd and higher order births
Panel A: FE				
Unemployment rate	-0.0074	-0.0046	-0.0150	0.0075
	(0.0051)	(0.0095)	(0.0091)	(0.0160)
Panel B: IV				
Unemployment rate	-0.0633***	0.0610	-0.0108	-0.0022
	(0.0220)	(0.0373)	(0.0346)	(0.0584)
Controls	Yes	Yes	Yes	Yes
TTWA time trends	Yes	Yes	Yes	Yes
Observations	41,710	41,710	41,710	41,710

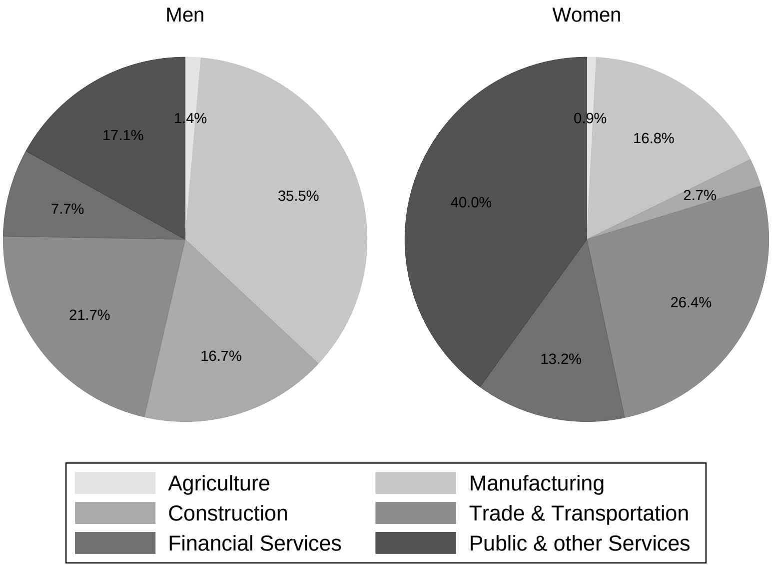
Notes: All regressions contain TTWA and year$\times$month fixed effects. Control variables include the year-of-age shares of 15- to 44-year-old women over all women aged 15 to 44 as well as population density, share of migrants of reproductive age, and the year-of-age shares of 45- to 74-year-old and 75+ year-old people over the population in each TTWA. TTWA time trends are region-specific linear trend variables. Robust standard errors clustered at the TTWA level in parentheses, *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

1.4 Gender-Specific Effects of Labor Market Conditions on Fertility

In a last step we are interested in discovering whether there are heterogeneous effects by gender. Based on our theoretical considerations, we expect men and women to respond differently to unemployment at the individual level. To this point, we have shown that there is a robust negative effect of local labor market conditions on fertility rates at the aggregate level, meaning that the negative income effect outweighs the positive substitution effect or, in other words, the reduction in opportunity costs that is probably more relevant for women plays only a minor role in fertility decisions. Now, we want to test the presumption that even on an aggregate level, male and female unemployment has differential effects on fertility rates.

We therefore replace the overall unemployment rate with the gender-specific unemployment rates UR^m_{rt} and UR^f_{rt}. Since we suppose that male and female are both endogenous regressors and, in addition, highly correlated, we follow Anderberg, Rainer, Wadsworth, and Wilson (2015) and use the gender gap in unemployment, $UR^m_{rt} - UR^f_{rt}$, to analyze gender-specific unemployment effects. Using the gender gap in unemployment has the advantage that the male unemployment rate reacts more strongly to business cycle fluctuations, meaning that in

Figure 1.5 : Share of male and female employees by industry in 1996 (in percent)



Source: German Microcensus (1996).

recessions the gender gap widens, whereas in economic upswings, it often closes (Albanesi and Sahin, 2013). We make use of this source of variation in the updated estimation equation:

$$log(FRT_{rt}) = \theta(UR^{m}_{rt} - UR^{f}_{rt}) + \gamma' X_{rt} + \mu_t + \phi_r + \omega_r * T + \epsilon_{rt}. \tag{1.4}$$

If standard micro models are also valid at the aggregate level, we expect that the coefficient on the gender gap in the unemployment rates, θ, will be significant and negative, implying a differential effect of male and female unemployment rates. We expect that increases in male unemployment widen the gap and reduce fertility due to a negative income effect. If decreasing female unemployment is the reason for a larger gender gap, theory predicts higher fertility when the substitution effect outweighs the income effect and lower fertility vice versa.

To analyze the gender-specific effects of local unemployment in the IV framework, we take advantage of the fact that male and female employment is concentrated in particular industries (see Figure 1.5). Albanesi and Sahin (2013) find that around half the gender differences in unemployment growth can be explained by differences in industry composition. In our base year 1996, men are overrepresented in construction and manufacturing, whereas a large majority of public-sector employees are female. We exploit this variation by constructing gender-specific labor demand indices based on employment trends in either the male- or the female-dominated industries. We expect the male unemployment rate to react strongly to changes in manufacturing labor demand. Female unemployment is expected to be highly correlated with changes in public-sector and private services employment. Thus, men are more affected in TTWAs where manufacturing represents a larger proportion of total employment. Vice versa, women suffer more from employment shocks in regions with a high workforce concentration in public services.

We follow the previous work by Albanesi and Sahin (2013) and Anderberg et al. (2015) in constructing a gender-specific shift-share index,

(1.5) $$LDI^g_{rt} = \sum_j G_{jt} * \frac{E_{jr0}}{E_{r0}},$$

where g is a superscript for gender and can be either m (= male) or f (= female). If we are interested in the demand shocks for males, j comprises the manufacturing sector, whereas the public sector is employed for the female index. The additional identifying assumption is that it is not possible to change from male- to female-dominated industries, or vice versa, to avoid negative labor market shocks. To simplify interpretation of the gender gap, we multiply the gender-specific LDI by the initial employment rate. The intuitive meaning of the LDI is then: How would the initial employment rate have evolved given the national trend in female- or male-dominated industries and the initial industry composition within the TTWAs? Finally, the instrumental variable for the gender gap in unemployment is the difference between the male and the female labor demand index, $LDI^m_{rt} - LDI^f_{rt}$. Again, we use the standardized measure of the gender gap in LDI for the first-stage regression. Column 3 in Table 1.2 confirms that the gender difference in our measure of labor demand is a valid predictor of the gender unemployment gap.

Table 1.6 shows FE as well as IV results for the gender-specific local unemployment rates as well as the gender gap in unemployment rates. From theory, we expect male unemployment to reduce fertility due to a negative income effect and female unemployment to either enhance or reduce fertility since women experience both lower opportunity costs of childbearing but also lower income due to higher probability of unemployment. FE results in Panel A suggest that male and female unemployment rates almost equally contribute to the overall effect of -0.5 percent, implying no gender differences in the effect of unemployment on fertility. Looking at birth order effects reveals that the negative effect for female unemployment is purely driven by a strong reduction in third births. Concerning family enlargement the female income effect seems to play a major role. At the extensive margin, that is, the decision to have a first child, male unemployment rates matter much more which may be because men are expected to be the main breadwinner after the birth of the first baby and thus refrain from entering parenthood until they can afford to maintain a family. However, we should be careful not to rely too heavily on the results in Panel A since they may suffer from an omitted variables bias.

To this point, we have not found any evidence for differential effects of male and female unemployment rates. Therefore, we substitute the gender-specific unemployment rates for the gender gap in unemployment. The estimates of the coefficients suggest gender differences as a 1 percentage point increase in the male-female unemployment gap leads to a reduction in fertility of 0.2 percent (FE, Panel B, Column 1) and 1.9 percent, respectively (IV, Panel C, Column 1). However, both coefficients fail to reach significance at the 10 percent level. In contrast, IV estimates for the intensive margins are negative and significant. Increases in the

Table 1.6 : Gender-specific effects on birth rate (by birth order)

Dependent variable	(log) births per 1,000 women			
	(1)	(2)	(3)	(4)
	All births	1st births	2nd births	3rd and higher order births
Panel A: FE				
Unemployment rate men	-0.0026**	-0.0089***	-0.0042*	0.0089***
	(0.0010)	(0.0022)	(0.0024)	(0.0031)
Unemployment rate women	-0.0026*	-0.0014	-0.0021	-0.0113***
	(0.0014)	(0.0032)	(0.0032)	(0.0043)
Panel B: FE				
Gap in unemployment rate	-0.0016	-0.0069***	-0.0030	0.0093***
	(0.0010)	(0.0020)	(0.0023)	(0.0031)
Panel C: IV				
Gap in unemployment rate	-0.0192	-0.0177	-0.0981***	-0.0606*
	(0.0125)	(0.0184)	(0.0321)	(0.0322)
1st stage F-stat	18.59	18.59	18.59	18.59
prob>F	0.000	0.000	0.000	0.000
Controls	Yes	Yes	Yes	Yes
TTWA time trends	Yes	Yes	Yes	Yes
Observations	41,710	41,710	41,710	41,710

Notes: All regressions contain TTWA and year$\times$month fixed effects. Control variables include the year-of-age shares of 15- to 44-year-old women over all women aged 15 to 44 as well as population density, share of migrants of reproductive age, and the year-of-age shares of 45- to 74-year-old and 75+ year-old people over the population in each TTWA. TTWA time trends are region-specific linear trend variables. Robust standard errors clustered at the TTWA level in parentheses, *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

gender gap strongly reduce birth rates for second and higher-order births. For instance, if the male unemployment rate goes up by 1 percentage point and the female rate remains unchanged, the rate of second births per 1,000 women falls by 9.8 percent and for third and higher-order births by 6.1 percent. If female unemployment rises, however, the sign of the effect is reversed. Finally, a surprising finding that is hard to explain and contrary to theory is the positive association between male unemployment (Panel A, Column 4) as well as the gender gap (Panel B, Column 4) and higher-order birth rates. When we apply the instrumental approach, the positive correlation vanishes and the coefficient becomes negative and significant. Thus, we conclude that the FE results, which imply a predominance of the substitution effect also for men, are biased.

For childless couples, neither income nor substitution effects seem to play much of a role in fertility decisions or cancel out each other. Thus, childlessness is not the channel through which the gender gap in unemployment reduces birth rates. It is, instead, the decision to extend a family that is affected. Rises in male unemployment reduce birth rates due to a negative income effect and increased female unemployment fosters higher-order fertility since the positive substitution effect dominates. In couples with offspring, the male partner usually works full-time while the mother divides her time between childrearing and working. If the risk of unemployment increases for the male partner, that is, the unemployment gap widens, the couple decides against having more children because of the imminent loss of income. For couples without children, the gender difference is less pronounced since usually both partners work full-time. Thus, the loss of income is comparable for men and women, resulting in a predominance of the income effect for women also.

To sum up, higher male unemployment relative to female reduces fertility, whereas higher female unemployment induces women to have children. The main mechanism behind this finding is the effect of the gender gap in unemployment on higher-order fertility. The effects at the intensive margins are negative and large, suggesting that the substitution effect of increased labor market uncertainty is predominant for women.

1.5 Heterogeneous Effects

Above, we looked at the effects of local labor markets on fertility by age of mothers. In this section we run more heterogeneity tests to discover the extent to which various population subgroups react to labor market uncertainties. Table 1.7 shows FE and IV results separately for East and West Germany (Columns 1 and 2), German and non-German mothers (Columns 3 and 4), and by mothers' employment status just before the birth of their child (Columns 5 and 6). The regional differences in the first two columns imply that there are fewer observations after splitting the sample. There are 204 TTWAs in West Germany and 40 TTWAs in the East (including Berlin). FE results for West Germany are in magnitude, sign, and significance level fairly similar to the combined coefficients, but much smaller and insignificant for the East German TTWAs. Applying the IV strategy changes the sign but not the level of significance for East Germany. Thus, in the former socialist part of Germany women do not reduce fertility in response to higher unemployment rates. The effect for West Germany is large and highly significant. Looking at the citizenship of the mothers reveals that uncertainty due to increased local unemployment is mainly an issue for German women. We divide births by German and non-German mothers by the respective population of all German or non-German women of reproductive age. The FE coefficient of -0.3 percent for German mothers is slightly lower than in the overall regression but still significant. In the IV case, it increases, implying a -1.3 percent reduction in fertility in response to a 1 percentage point increase in the local unemployment rate. Foreign mothers do not respond significantly to changes in local labor market conditions. Finally, we focus on mothers' employment status right before giving birth. The idea is that for employed mothers, local labor market conditions might matter more since they may fear losing their jobs when pregnancy and recessions coincide. However, neither FE nor IV regressions reveal a significant effect on fertility rates in the TTWAs. The FE coefficient of employed mothers at least shows a negative sign, but is not precisely enough estimated.

This finding also holds for unemployed mothers. Heterogeneity analyses show that the most affected group is German women in West German TTWAs.

Table 1.7 : Heterogeneity of effects on birth rate

Dependent variable	(log) births per 1,000 women					
	(1) West Germany	(2) East Germany	(3) German mothers	(4) Foreign mothers	(5) Employed mothers	(6) Unemployed mothers
Panel A: FE						
Unemployment rate	-0.005***	-0.0025	-0.0030**	-0.0050	-0.0025	-0.0007
	(0.0014)	(0.0029)	(0.0015)	(0.0041)	(0.0063)	(0.0043)
Panel B: IV						
Unemployment rate	-0.0248**	0.0128	-0.0131**	0.0147	0.0103	0.0169
	(0.0107)	(0.0285)	(0.0057)	(0.0165)	(0.0355)	(0.0251)
Controls	Yes	Yes	Yes	Yes	Yes	Yes
TTWA time trends	Yes	Yes	Yes	Yes	Yes	Yes
Observations	34,882	6,828	41,710	41,710	41,710	41,710

Notes: All regressions contain TTWA and year$\times$month fixed effects. Control variables include the year-of-age shares of 15- to 44-year-old women over all women aged 15 to 44 as well as population density, share of migrants of reproductive age, and the year-of-age shares of 45- to 74-year-old and 75+ year-old people over the population in each TTWA. TTWA time trends are region-specific linear trend variables. Robust standard errors clustered at the TTWA level in parentheses, *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

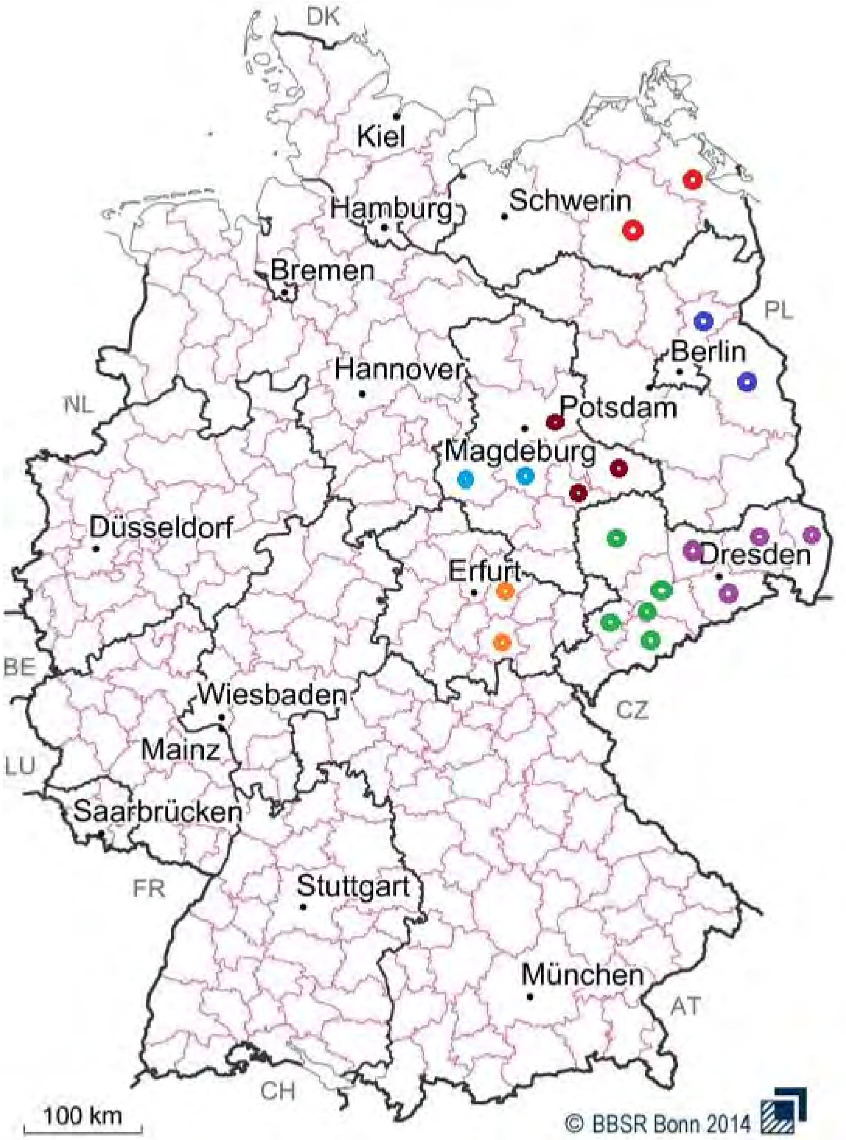
1.6 Conclusion

Germany has experienced fertility rates far below replacement level ever since the 1970s. If this trend continues, it will result, in the long run, in an ageing and even shrinking society with possibly negative consequences for economic growth and the public budget. In this chapter, we shed light on how adverse local labor market conditions influence birth rates. A high unemployment rate in the area of residence may imply a high level of unemployment risk. Theory suggests that this economic uncertainty can have consequences for couples' fertility plans. In fact, our main results from FE as well as IV regressions suggest a negative impact of increased local unemployment on birth rates. The size of the effect is substantial: in our main specification, a 1 percentage point increase in the local unemployment rate leads to a 0.9 percent decrease in the birth rate per 1,000 women. For instance, after the burst of the dot-com bubble, the average annual unemployment rate went from 6.4 percent in 2001 to almost 8.5 percent in 2005. This 1.9 percentage point increase in unemployment led to an average decrease in births of roughly 1.6 percent over this four-year period. Given 767,000 births in the pre-crisis year, this means that there were 50,000 fewer babies born because of this recessionary period.

Since changes in unemployment are cyclical – periods of higher unemployment are followed by periods of lower unemployment – we test whether the chapter's main finding can be explained by changes in the timing of births. Specifically, we run regressions for age-group-specific birth rates as well as for mothers' age at first, second, and third birth. We find no evidence either for changes in age at first to third birth or for younger women being particularly affected. Thus, we argue that we are not measuring a timing effect due to postponement but an actual decline in birth rates. Based on theoretical considerations and empirical results, policymakers interested in increasing the birth rate should strive to mitigate the negative impact of increased local unemployment and its associated uncertainty. Since reduced income appears to be the main obstacle to fertility, compensating for loss of income or improving men's labor market conditions to close the gender unemployment gap would appear to be appropriate measures for increasing or at least maintaining current birth rates.

Appendix A.1 Supplementary Figures

Figure A.1 : Map of Travel-to-Work areas in Germany 2012



Source: University of Kassel, Prof. Kosfeld; Laufende Raumbeobachtung des BBSR. Geometrische Grundlage: BKG, Kreise 31.12.2012. Editing: P. Kuhlmann. *Notes:* TTWA in East Germany with dots of the same color have to be combined for the analysis due to local government reorganisations. The number of TTWAs is reduced from 258 to 244 due to aggregating.

Appendix A.2 Supplementary Tables

Table A.1 : Travel-to-Work areas in Germany 2012

(1) TTWA	(2) Name	(3) TTWA	(4) Name	(5) TTWA	(6) Name	(7) TTWA	(8) Name	(9) TTWA	(10) Name
1	Husum	53	Schwelm	105	Trier	157	Weilheim	209	Eberswalde
2	Heide	54	Remscheid	106	Bernkastel-Wittlich	158	Landsberg	210	Luckenwalde
3	Itzehoe	55	Kleve	107	Daun	159	München	211	Finsterwalde
4	Flensburg	56	Aachen	108	Bitburg	160	Ingolstadt	212	Oranienburg
5	Lübeck	57	Köln	109	Kaiserslautern	161	Kelheim-Mainburg	213	Neuruppin
6	Kiel	58	Leverkusen	110	Landau	162	Landshut	214	Perleberg
7	Ratzeburg	59	Bonn	111	Mainz	163	Dingolfing	215	Prenzlau
8	Hamburg	60	Düren	112	Alzey-Worms	164	Eggenfelden/ Pfarrkirchen	216	Rostock
9	Braunschweig	61	Euskirchen	113	Pirmasens	165	Passau	217	Schwerin
10	Salzgitter	62	Gummers-bach	114	Ludwigshafen	166	Freyung	218	Mecklenburgische Seenplatte
11	Wolfsburg	63	Gelsenkirchen	115	Germersheim	167	Regen-Zwiesel	219	Nordvorpommern
12	Göttingen	64	Münster	116	Merzig	168	Deggendorf	220	Südvorpommern
13	Goslar	65	Borken	117	St. Wendel	169	Straubing	221	Chemnitz
14	Helmstedt	66	Steinfurt	118	Saarbrücken	170	Cham	222	Erzgebirgskreis
15	Einbeck	67	Bielefeld	119	Homburg/Saar	171	Regensburg	223	Mittelsachsen
16	Osterode	68	Gütersloh	120	Stuttgart	172	Schwandorf	224	Vogtlandkreis
17	Hannover	69	Detmold	121	Göppingen	173	Amberg	225	Zwickau
18	Sulingen	70	Minden	122	Heilbronn	174	Neumarkt	226	Dresden
19	Hameln	71	Paderborn	123	Schwäbisch Hall	175	Weiden	227	Bautzen
20	Hildesheim	72	Bochum	124	Tauberbischofsheim	176	Marktredwitz	228	Görlitz
21	Holzminden	73	Dortmund	125	Heidenheim	177	Hof	229	Meißen
22	Nienburg	74	Hagen	126	Aalen	178	Bayreuth	230	Leipzig
23	Stadthagen	75	Lüdenscheid	127	Baden-Baden	179	Bamberg	231	Dessau-Roßlau
24	Celle	76	Meschede	128	Karlsruhe	180	Kulmbach	232	Halle
25	Lüneburg	77	Siegen	129	Heidelberg	181	Kronach	233	Magdeburg
26	Zeven	78	Olpe	130	Mannheim	182	Coburg	234	Salzwedel
27	Soltau	79	Soest	131	Mosbach	183	Lichtenfels	235	Anhalt-Bitterfeld
28	Stade	80	Korbach	132	Pforzheim	184	Erlangen	236	Burgenlandkreis
29	Uelzen	81	Kassel	133	Calw	185	Nürnberg	237	Harz
30	Verden	82	Eschwege	134	Freudenstadt	186	Weißenburg-Gunzenhausen	238	Mansfeld-Südharz
31	Emden	83	Schwalm-Eder	135	Freiburg	187	Ansbach	239	Salzlandkreis
32	Westerstede	84	Hersfeld	136	Offenburg	188	Neustadt/Aisch	240	Stendal
33	Oldenburg	85	Marburg	137	Rottweil	189	Kitzingen	241	Wittenberg
34	Osnabrück	86	Lauterbach	138	Villingen-Schwenningen	190	Würzburg	242	Erfurt
35	Wilhelms-haven	87	Fulda	139	Tuttlingen	191	Schweinfurt	243	Gera
36	Cloppenburg	88	Wetzlar	140	Konstanz	192	Haßfurt	244	Jena
37	Lingen	89	Gießen	141	Lörrach	193	Bad Neustadt/Saale	245	Suhl
38	Nordhorn	90	Limburg	142	Waldshut	194	Bad Kissingen	246	Weimar
39	Leer	91	Wiesbaden	143	Reutlingen/ Tübingen	195	Lohr am Main	247	Eisenach
40	Vechta	92	Frankfurt/ Main	144	Balingen	196	Aschaffenburg	248	Eichsfeld
41	Nordenham	93	Hanau	145	Ulm	197	Donauwörth-Nördlingen	249	Nordhausen
42	Bremen	94	Darmstadt	146	Biberach	198	Dillingen	250	Mühlhausen
43	Bremerhaven	95	Erbach	147	Friedrichshafen	199	Günzburg	251	Sondershausen
44	Höxter	96	Altenkirchen	148	Ravensburg	200	Augsburg	252	Meiningen
45	Düsseldorf	97	Montabaur	149	Sigmaringen	201	Memmingen	253	Gotha
46	Duisburg	98	Neuwied	150	Bad Reichenhall	202	Kaufbeuren	254	Arnstadt
47	Essen	99	Ahrweiler	151	Traunstein	203	Kempten	255	Sonneberg
48	Krefeld	100	Koblenz	152	Burghausen	204	Lindau	256	Saalfeld
49	Viersen	101	Bad Kreuznach	153	Mühldorf	205	Berlin	257	Pößneck
50	Mönchen-gladbach	102	Idar-Oberstein	154	Rosenheim	206	Potsdam-Brandenburg	258	Altenburg
51	Heinsberg	103	Cochem	155	Bad Tölz	207	Cottbus		
52	Wuppertal	104	Simmern	156	Garmisch-Partenkirchen	208	Frankfurt/Oder		

2 The Long-Run Consequences of Unemployment Experience on Fertility

2.1 Introduction

In 1986, the seminal paper by Blanchard and Summers introduced the concept of *hysteresis* to characterize labor market dynamics in Europe during the last few decades. *Hysteresis* is the phenomenon that increases in unemployment due to negative shocks have a persistent effect on the *natural unemployment rate*. For instance, the oil crises in the 1970s permanently raised the unemployment rate despite the subsequent economic recovery. Researchers have tried to discover whether increased unemployment is responsible for the steady decline in birth rates. However, most studies, such as Schaller (2016) and Dehejia and Lleras-Muney (2004), as well as the preceding chapter, focus on only the short-run effects of adverse labor market opportunities and ignore possible catching-up effects if births are entirely postponed. Thus, we analyze the long-run consequences for fertility of unemployment rates experienced during the reproductive age.

To the best of our knowledge, we are the first to explicitly investigate the long-run effects of unemployment rates on completed fertility and childlessness in the German context. We further contribute to the literature by investigating differences across gender and educational groups, and by providing evidence regarding mechanisms that have the potential to explain our findings.

The last several decades have seen increased interest in fertility behavior by economists, demographers, and sociologists. The extant research offers a variety of explanations for the low fertility rates in the developed world. Among demographers and sociologists, preference-based explanations prevail, whereas economists focus on the price of childbearing. In recent years, the role played by family policies has received increasing attention from researchers.

According to the theory of the "second demographic transition," recent fertility trends are due to changes in individual values and preferences (see, e.g., Van de Kaa, 1987; Lesthaeghe, 1995). Specifically, the theory argues that the introduction of the birth control pill and the women's rights movement had substantial influence on fertility. Women became able to live more self-determined lives and control their fertility. As women now had more choices, including those involving self-realization, leisure, and consumption, they began to spend more time on attaining education and in the labor market than on household production, that is, childrearing (Lesthaeghe, 2010). A similar approach employing endogenous preferences was proposed by Easterlin (1973, 1987). He explains the baby boom and the subsequent baby bust by changing preferences for children. Willis (1987) posits that it is changes in the intergenerational relative income across cohorts that cause these shifts in preferences. If the parents' generation is relatively poor, individuals believe that they can afford to enter parenthood early and have large families.

The economic approach to fertility dates back to Gary S. Becker (1960). In his work, preferences are assumed to be exogenously fixed and demand for children is ultimately altered by changes in the price of children. Children are modeled as normal consumption goods and fertility decisions are based on the relative costs of having children, including direct costs (e.g., clothing, nutrition) and the costs of foregone income (i.e., opportunity costs). Becker argues that the increased earning power of women and the concomitant higher opportunity costs enhance female labor supply and, ultimately, lower fertility (Becker, 1991). In addition, with increasing levels of income and education, parents tend to invest more in the quality, rather than quantity, of their children.

Inspired by Becker's theory, a large body of economic literature investigates how increased female education affects fertility levels. Concerning the timing of births, empirical evidence suggests that higher educational attainment causes a decrease in teenage motherhood and leads to a postponement of first birth (Black, Devereux, and Salvanes, 2008; Cygan-Rehm and Maeder, 2013; Duflo, Dupas, and Kremer, 2015), whereas the findings regarding completed fertility are inconclusive. Although women with higher education tend to delay first births, they catch up later in life. Thus, education influences only the timing of fertility but not its completion (Breierova and Duflo, 2004; Monstad, Propper, and Salvanes, 2008; Grönqvist and Hall, 2013). In contrast, Amin and Behrman (2014) show that the higher the level of education, the fewer the number of children. Lavy and Zablotsky (2015), who also find a reduction in completed fertility if education rises exogenously, provide several explanations for this phenomenon, including increased child quality and changes in fertility preferences. In their recent work, Baudin, de la Croix, and Gobbi (2015) develop a model of extensive- and intensive-margin fertility. They show that childlessness among very low- and very high-educated women in the United States is a main reason for low fertility. Women at the lower tail of the educational distribution suffer from "social sterility," i.e., they cannot afford having children because of their precarious income. In contrast, women at the upper tail of the distribution refrain from having children due to high opportunity costs.

In their pioneer work, Becker and Lewis (1973) and Willis (1973) established an inverse relationship between female labor force participation and fertility. Testing the model with U.S. data, Butz and Ward (1979) show that this inverse relationship also holds empirically. More recent studies challenge the direction of causality and find that lower fertility encourages women to participate in the labor market (see, e.g., Angrist and Evans, 1998; Bailey, 2006). Ward and Butz's (1980) dynamic model of fertility behavior reveals the role labor market conditions play in determining the opportunity costs of having children. They find strong evidence for an intertemporal substitution effect between market work and fertility since higher female wage expectations lower fertility, whereas for male income the effect goes in the opposite direction. They conclude that current economic conditions alter opportunity costs and thus determine the fertility level. Ward and Butz (1980) and Bongaarts and Feeney (1998), point out that a reduction in period fertility measures, such as the birth rate or the total fertility rate, need not cause a reduction in cohort fertility if the period measure is confounded by changes in fertility timing.

Recently, the role played by institutions and family policies has attracted the attention of researchers. The basic idea is that improvements in a family's financial resources or a better compatibility of work and family life can reduce the opportunity costs of childbearing and thus induce couples to have more children. Lalive and Zweimüller (2009) investigate the effect of a parental leave reform. The authors discover positive effects on fertility in response to an extension of parental leave. In contrast, a reduction in parental leave lowers fertility in the years right after the reform (Cygan-Rehm, 2015). This effect is driven by low-income women who cannot afford to catch up postponed births. Similarly, expansion of public childcare raises fertility by enabling women to combine children and a successful career (Rindfuss, Guilkey, Morgan, and Kravdal, 2010; Bauernschuster et al., 2015). Bailey (2012) tests the effectiveness of a poverty-reducing policy in the United States and finds that among low-income families public subsidies reduce fertility. However, despite the positive effects of family policies, the negative relationship between education and fertility persists (Björklund, 2006).

The literature dealing with the link between labor market conditions and fertility tends to focus on short-run effects; very little of this work distinguishes between tempo and quantum effects of changes in local labor market opportunities (see, e.g., Schaller, 2016; Dehejia and Lleras-Muney, 2004; Karaman Örsal and Goldstein, 2010). However, there are a few studies evaluating the effects of individual or aggregate unemployment on completed fertility. Due to the generous welfare system in France, Pailhé and Solaz (2012) find only weak evidence that fertility decreases in response to spells of unemployment. Only long-term unemployed men significantly reduce completed fertility; women delay childbearing but do not adjust the number of births. In the Norwegian setting, Kravdal (2002) discovers a negative relationship between unemployment and fertility only at the aggregate level. Again, individual unemployment does not seem to matter in couples' fertility decisions. The most relevant study for our research is that by Currie and Schwandt (2014) who estimate the long-run relationship between fertility and unemployment for the US. Following state-year birth cohorts over time, they find a significant negative effect of state-level unemployment on completed fertility rates induced by an increase in the likelihood of staying childless. Observing where women give birth to their children enables them to control cohort migration patterns. Employing an instrumental variable approach, corrects for a potential selection bias.

Based on this literature we investigate the effect on women's completed fertility of the unemployment rates experienced during reproductive age. We make use of data from the 2008 and 2012 German Microcensuses, which provide information about the number of children born to each woman. We focus on the female birth cohorts from 1954 to 1967 since these women are at least 40 years old in both waves. Four state-specific unemployment measures are merged with the individual data. We average the experienced unemployment rates over five-year age intervals starting with 20-24 up to 35-39. Standard OLS and probit estimation results suggest that increases in average female unemployment rates during early career years significantly increase fertility, whereas rising male unemployment rates have the opposite effect. This relationship is mainly driven by changes in the probability of remaining childless. For instance, if the average female unemployment rate during the age 20 to 24 increases by 1 percentage point, the likelihood that a woman remains childless decreases by 1.6 percentage points. However, if the male unemployment rate rises during the same age span, childlessness increases by 1.2 percentage points. The heterogeneity analysis reveals that the findings are mainly driven by women with low education. Two mechanisms may explain the findings: first, unemployment rates have a substantial influence on marriage market outcomes and, second, they impact household income. Since we are not able to explicitly control for the mobility of a cohort, we argue graphically and in a regression framework that the results are robust to potential distortions caused by women migrating to states with more favorable labor market conditions.

The remainder of the chapter is organized as follows: Section 2.2 describes the data and Section 2.3 the empirical approach. Section 2.4 sets out the main results while Sections 2.5 and 2.6 present the heterogeneity analysis as well as the potential mechanisms. Robustness checks are presented in Section 2.7; Section 2.8 concludes.

2.2 Data and Descriptives

2.2.1 Completed Fertility

To conduct the empirical analysis, we need information on women's completed fertility. The completed fertility rate (CFR) is a measure of cohort fertility at the end of the reproductive age. Unlike period measures, such as the total fertility rate (TFR) or the crude birth rate, it is not distorted by changes in the timing of births. Thus, delayed childbearing reduces, and accelerated childbearing raises, the period measures even if completed fertility is not affected (Bongaarts and Feeney, 1998).

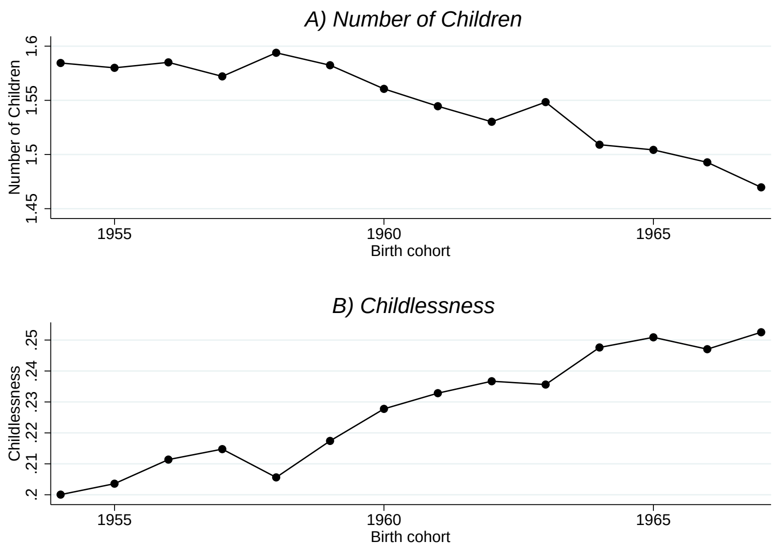
The German Microcensus (GMC) – the largest household survey in Europe – annually collects information on about 1 percent of the German population. In 2008 and 2012, the GMC asked whether a woman between the ages 15 and 75 had born children and, if so, how many. The corresponding questions were: "Have you born children?" and "How many children have you born?" From the first question we obtain a measure of childlessness and from the second a measure of completed fertility, that is, the number of children ever born to a woman. We restrict the sample to women from the West German federal states born in the birth cohorts 1954 to 1967 since they are at least 40 years old in both waves.[1] The sample restriction ensures

[1] We do not consider East Germany in our research because reliable unemployment data are missing for the time prior to Reunification. Moreover, differences in East German fertility behavior compared to that of West Germans are well documented in the literature (see, e.g., Kreyenfeld, 2004). Since we do not have information

that all cohorts are equally represented and the number of observations is sufficient to obtain precise estimates. To avoid distortions caused by outliers, we drop women with more than 10 children, which is only in 0.03 percent of all observations. The main sample consists of 14 birth cohorts including 90,756 women from 10 West German states.

Figure 2.1 shows the development of cohort fertility and childlessness for our sample cohorts 1954 to 1967. Completed fertility is stable for women born between 1954 and 1958. They have, on average, 1.6 children by the end of their reproductive years. Until 1970, the corresponding TFR is above 2.0; after 1970, the gap between TFR and CFR begins to narrow. This suggests substantial changes in the timing of births: during the baby boom years, older women catch up births and younger women enter parenthood earlier in life. Since the beginning of the 1960s, cohort fertility has continuously decreased and is around 1.46 for the latest cohort. The development of childlessness shows an opposite trend. In the earlier cohorts, less than 20 percent of women above 40 years remain childless, whereas for women from the birth cohort 1967, this share increases to almost 25 percent. However, the increasing trend in childlessness appears to flatten for the younger cohorts. The increase in childlessness among women may explain the drop in completed fertility.

Figure 2.1 : Development of completed fertility and childlessness



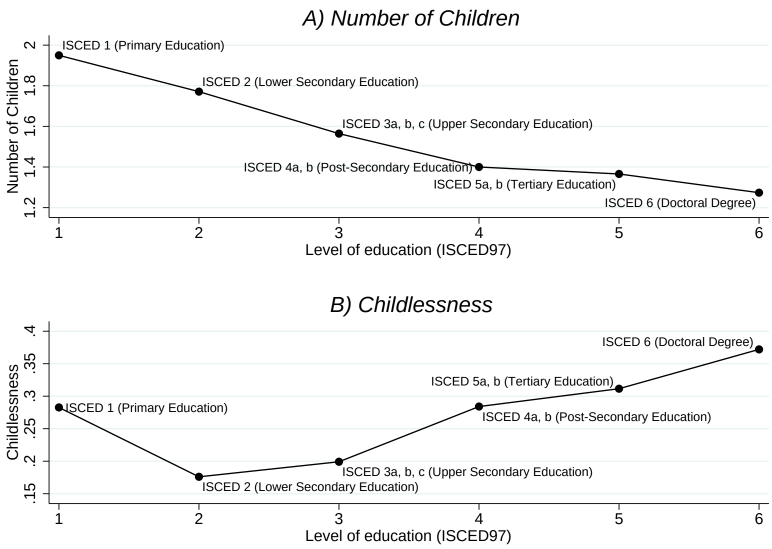
Notes: Female birth cohorts from 1954 to 1967 in West Germany.
Source: Microcensus (2008, 2012).

Fertility and childlessness also vary widely across educational subgroups. Figure 2.2 shows the average number of children and the share of childless women for six different educational

about birthplace, we drop East Germans based on their current state of residence and their level of education. Educational degrees obtained before Reunification differ between East and West Germany and allow us to identify individuals who are from the former socialist part of Germany. Moreover, we drop women who lived abroad during their reproductive years.

groups. Women with low to medium education (ISCED 1-3) have, on average, between 1.6 and 2.0 children. Compared to other developed countries, this number is fairly high. In addition, these women comprise almost 70 percent of the female population between 20 and 39 years. Thus, for a significant share of women we do not observe lowest-low fertility rates. Women with upper secondary or tertiary education (ISCED 4-6) exhibit fertility rates less than 1.4 children per woman. Particularly, the high level of childlessness seems to explain why completed fertility is so low among highly qualified women (ISCED 5: 32 percent, ISCED 6: 40 percent). However, childlessness is not only noticeable among highly educated women but also occurs in the lowest educational group. Women without a secondary degree either have far more children than average or are childless.

Figure 2.2 : Completed fertility and childlessness by education

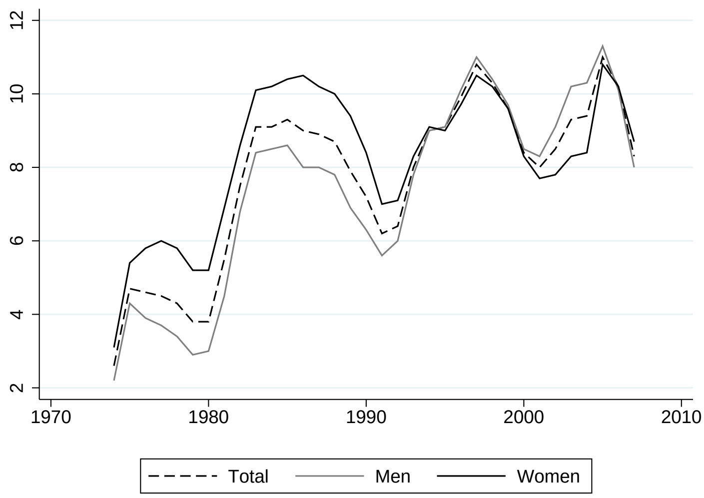


Notes: Female birth cohorts from 1954 to 1967 in West Germany.
Source: Microcensus (2008, 2012).

2.2.2 State-Level Unemployment Rates

For our analysis, we match the Microcensus data with data from the annual statistics of the German Federal Employment Agency (BA) that are known as "Amtliche Nachrichten der Bundesagentur für Arbeit (ANBA)". Since 1974, the BA has published annual unemployment rates by gender for all West German states. For every single birth cohort, we compute the average state-specific unemployment rate for the five-year age intervals starting at 20 to 24 years up to 35 to 39 years. Ideally, we would like to know in which federal state a woman gave birth to her children. However, we have information on state of residence only ex-post, at

Figure 2.3 : Development of national unemployment rate (by gender)



Notes: West Germany from 1974 to 2007.
Source: ANBA (1974-2007).

the end of the reproductive age. Thus, we are forced to assume that women do not move between federal states during their reproductive age.[2]

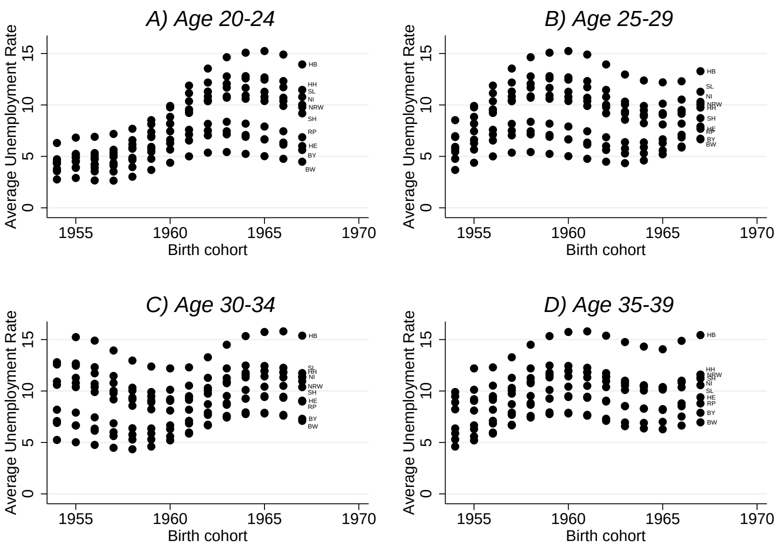
Figure 2.3 shows the overall, as well as the gender-specific, unemployment rates from 1974 to 2007 – the period during which our sample birth cohorts are between 20 and 40 years of age. The development is characterized by a steady upward trend interrupted only by short periods of recovery. This phenomenon is exactly described in Blanchard and Summers (1986), who explain that increases in unemployment may be persistent since they directly impact the natural level of unemployment. For instance, the sharp increases in unemployment due to the recessions in 1975 and 1982 were not followed by equivalent decreases in the subsequent boom period. Instead, the stock of unemployed individuals is larger after the recession. The same applies to the recession after the Reunification boom and the burst of the dot-com bubble in the early 2000s. The Hartz reforms implemented in the middle of the last decade initially increased unemployment rates, reaching a peak of 11 percent in 2005. Since then, there has been a slight decrease in unemployment.

Figure 2.4 focuses on the average state-level unemployment rates to which the birth cohorts 1954 to 1967 were exposed during their reproductive years. Early cohorts suffered most from high unemployment when they were between 30 and 34 – rather late in their period of fertility, meaning that decisions whether to have children quite likely have already been made. Later cohorts were hit by several recessions and at different stages of the life course. For instance, the birth cohort 1965 experienced high levels of unemployment between the ages of 20 and 24 due to the second oil crisis and the subsequent recession in 1982, and

[2] Section 2.7 presents descriptive evidence that for the vast majority of women there is very little year-by-year mobility across states.

also after the Reunification boom when they were between the ages of 30 and 34. Note that there is a fair degree of variation in average unemployment rates not only across cohorts but also across states. After the 1982 recession, the highest unemployment rates occurred in Bremen at around 15 percent, whereas at the same time, rather low rates prevailed in Bavaria, Baden-Württemberg, and Hesse, ranging from 5 percent to 7 percent. In addition, the peaks of the recessions hit different cohorts across states. For example, average unemployment in Bavaria for the age group 20-24 peaks for the birth cohort 1963; Bremen exhibits the highest average unemployment two cohorts later, that is, for the women born in 1965.

Figure 2.4 : Development of average state-level unemployment rates



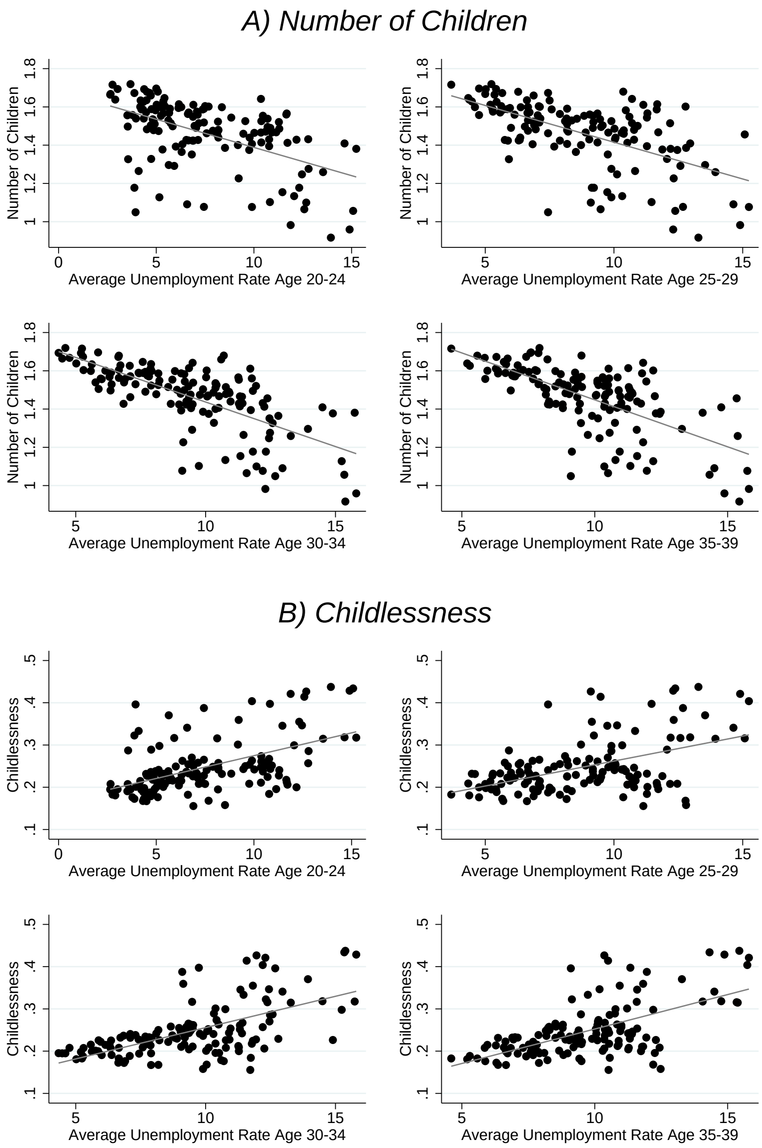
Notes: Birth cohorts from 1954 to 1967; West German states are: BW: Baden-Württemberg; BY: Bavaria; HB: Bremen; HE: Hesse; HH: Hamburg; NI: Lower Saxony; NRW: North Rhine-Westphalia; RP: Rhineland-Palatinate; SH: Schleswig-Holstein; SL: Saarland.
Source: ANBA (1974-2007).

2.2.3 Graphical Evidence

To investigate the relationship between unemployment and completed fertility, we first look graphically at the correlation on the state-cohort level. Therefore, we merge the ANBA unemployment rates with the GMC data, making use of the unique combination of federal states and birth cohorts, and obtain 140 state-cohort observations. Figure 2.5 illustrates the correlation between average unemployment rates in different age groups and the number of children in Panel A and the share of childless women in Panel B.

If we focus on the overall unemployment rates, low unemployment coincides with high fertility and high unemployment with low fertility (Panel A of Figure 2.5). Thus, the fitted line shows first evidence of a negative relationship between experienced unemployment and completed

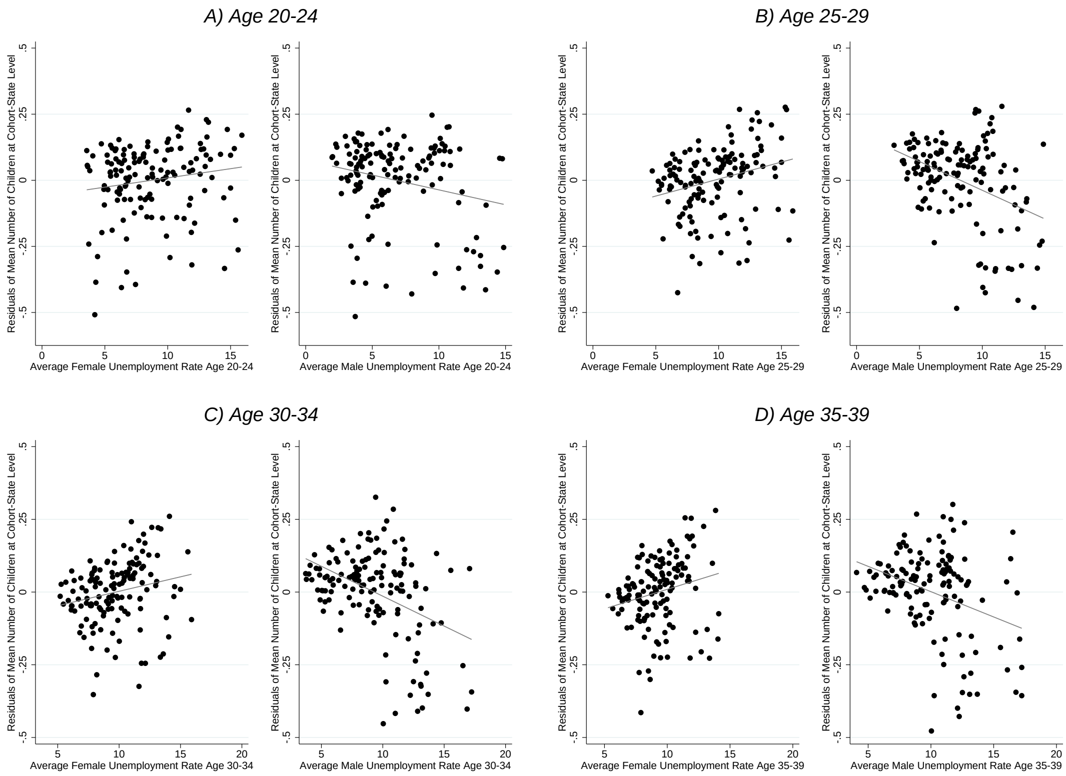
Figure 2.5 : Relationship between unemployment, childlessness, and fertility



Notes: The upper-left graph of both panels is a scatter plot of the fertility measures and unemployment experienced between the age of 20 and 24 and the upper-right graph shows the same for those between 25 and 29. The lower-left graph is a scatter plot of the fertility measures and unemployment experienced between ages 30 and 34 and the lower-right graph shows the same for those between 35 and 39. Each data point represents a cohort-state cell.
Sources: Microcensus (2008, 2012), ANBA (1974-2007).

fertility. For the age group 30 to 34, the negative association seems very pronounced. In combination with the low dispersion around the fitted line, we expect a significant correlation in this age group. Similarly, the share of childless women in a state-cohort cell increases with the level of unemployment (Panel B). Thus, the negative correlation with fertility corresponds to a positive correlation between experienced unemployment rates and childlessness. At the right end of each graph, where unemployment rates are high, there are very few observations, mainly Bremen, and variation is very large, which could potentially confound the fit of the line and estimation results later in the chapter.

Figure 2.6 : Gender-specific unemployment and completed fertility

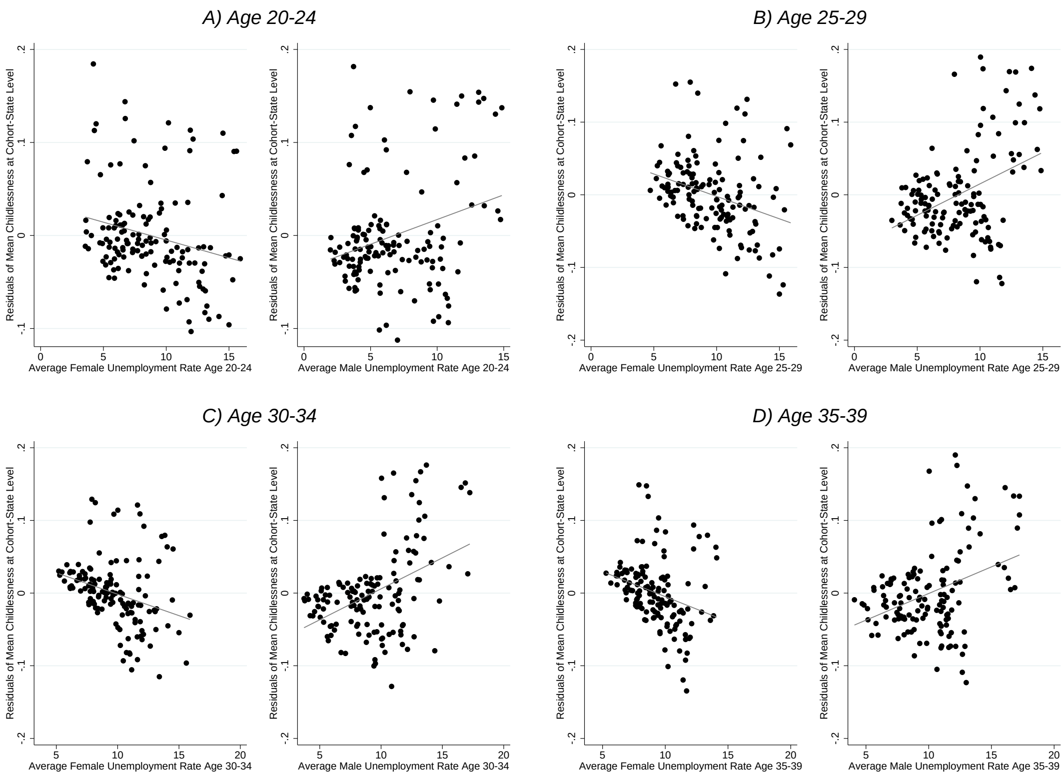


Notes: Each data point represents a cohort-state cell.
Source: Microcensus (2008, 2012), ANBA (1974-2007)

To discover gender-related differences in response to the experience of unemployment rates, we now replace the overall unemployment rates with the gender-specific rates. Since male and female average unemployment rates might be highly correlated, we look at the correlation holding constant the other gender's level of unemployment. Figure 2.6 reveals a negative correlation between average male unemployment rates and completed fertility for all age groups. However, for female unemployment rates, the pattern reverses. Holding male unemployment constant, women tend to have more children if the unemployment rate they experience during their reproductive years increases. A possible explanation is the decline in opportunity costs of childbearing if labor market conditions worsen. The negative relationship for men seems to be stronger than the positive relationship for women. Although, graphically, it is hard to identify differences across age groups, the slope of the fitted lines seems to increase for higher age categories.

In Figure 2.7 we plot the same relationship substituting number of children by incidence of childlessness. The basic pattern remains unchanged but the sign of the relationship reverses:

Figure 2.7 : Gender-specific unemployment and childlessness



Notes: Each data point represents a cohort-state cell.
Source: Microcensus (2008, 2012), ANBA (1974-2012)

higher female unemployment rates are accompanied by lower shares of childlessness, whereas higher male unemployment is associated with increases in childlessness. Particularly, the correlation with the male unemployment rate shows differences across age groups. Between 25 and 29, as well as between 30 and 34, an increase in the level of male unemployment substantially increases the share of childlessness. The negative relationship for experienced female unemployment rates is weak and quite stable across age groups. Again, we find a high level of variation in each of the plots that might confound the fit of the line.

To sum up, graphical evidence suggests that fertility outcomes and unemployment rates experienced throughout the reproductive years are related at the state-cohort level. Using male and female unemployment rates reveals differences across gender: male unemployment lowers, whereas female unemployment increases, fertility. The following section explores the extent to which these findings hold in a multivariate set-up.

2.3 Method

In our main analysis we apply standard OLS and probit methods to estimate the experience effects of state-level unemployment rates on fertility outcomes at age 40. The basic regression implements a stepwise experience function of unemployment rates since it allows for different coefficients along women's reproductive age. The estimation equation is:

$$y_{ics} = \beta_1\overline{U(2024)}_{cs}+\beta_2\overline{U(2529)}_{cs}+\beta_3\overline{U(3034)}_{cs}+\beta_4\overline{U(3539)}_{cs}+\gamma' X_{ics}+\mu_c+\phi_s+\epsilon_{ics} \tag{2.1}$$

where y_{ics} is the fertility outcome of woman i of cohort c in state s. $\overline{U(2024)}_{cs}$ is the average unemployment rate that all women in cohort c faced between the ages of 20 and 24 in their current state of residence. The other unemployment rate variables are defined equivalently. Since time-varying variables might be affected by the state unemployment rates and therefore cannot be treated as valid controls, we do not include them in our model. Instead, we use time-constant covariates, X_{ics}, an indicator variable for migratory background, and four dummy variables to control for level of education. The unemployment rates a woman experiences during her reproductive years, can have an influence on her educational choices. For instance, adverse labor market conditions may induce her to invest in higher education. Thus, parts of the correlation between unemployment rates and fertility are due to changes in educational outcomes and not to labor market conditions. To reduce the threat of these feedback effects, we do not include the ISCED-level in our model. Instead, level of schooling proxies for level of education since schooling mainly takes place before age 20 and is not affected by the experienced unemployment rates. In addition, μ_c and ϕ_s represent cohort- and federal-state-fixed effects. Cohort- and state-fixed effects account for systematic differences in fertility behavior of women across birth years and state. For instance, a change in fertility preferences is captured by the cohort-fixed effect if these changes affect some cohorts more than others. Permanent differences between unemployment rates and fertility are captured by the state-fixed effects as long as they are constant over time.

Standard microeconomic theory of fertility does not predict an unambiguous effect of increased unemployment on fertility (see, e.g., Becker, 1991). On the one hand, demand for children will fall since unemployment leads to a reduction in wages and family income (income effect). On the other hand, lower wages reduce the opportunity costs of childrearing, which should increase demand for children (substitution effect). However, as in Germany women traditionally devote more time to childrearing than do men, the opportunity costs argument applies mainly to women. As a result, unfavorable economic conditions for men through increased unemployment risk and lower income are expected to reduce fertility rates, whereas adverse labor market conditions for women affect fertility through a negative income and a positive substitution effect. Thus, we look more closely at gender-specific effects by augmenting the equation with two sets of unemployment rate variables, one containing average male unemployment rates (e.g., $\overline{U(2024)}_{cs}^{m}$) and one average female unemployment rates (e.g., $\overline{U(2024)}_{cs}^{f}$). This approach enables us to answer the question of how male and female labor market conditions affect completed fertility as well as the incidence of childlessness.

$$
\begin{aligned}
y_{ics} = {} & \beta_1^m \overline{U(2024)}_{cs}^m + \beta_2^m \overline{U(2529)}_{cs}^m + \beta_3^m \overline{U(3034)}_{cs}^m + \beta_4^m \overline{U(3539)}_{cs}^m \\
& + \beta_1^f \overline{U(2024)}_{cs}^f + \beta_2^f \overline{U(2529)}_{cs}^f + \beta_3^f \overline{U(3034)}_{cs}^f + \beta_4^f \overline{U(3539)}_{cs}^f \\
& + \gamma' X_{ics} + \mu_c + \phi_s + \epsilon_{ics}
\end{aligned} \tag{2.2}
$$

Table 2.1 reports summary statistics for the fertility measure as well as the unemployment rates and the control variables. On average, women of the birth cohorts 1954 to 1967 have 1.53 children during their reproductive years and 23 percent remain childless. Experienced unemployment rates increase over the fertile lifecycle, starting with 7 percent between 20 and 24 and peaking at almost 9 percent for the ages 35 to 39. This reflects the overall upward trend in unemployment rates over the sample period 1974 to 2007. 5 percent of the women in the sample have a migratory background, meaning that either they or their parents are foreign born. Sample statistics reveal that over 37 percent of women from the birth cohorts 1954 to 1967 have a lower secondary school degree (*Hauptschulabschluss*), 35 percent graduate from *Realschule* and the remaining 27 percent finish school with a university entrance degree – either for a university (21 percent) or a university of applied sciences (6 percent).

Table 2.1 : Sample summary statistics: women from birth cohorts 1954 to 1967

	Mean	SD	Min	Max
Number of children	1.534	1.149	0	10
Childlessness	0.228	0.419	0	1
Average unemployment rate 20-24	0.069	0.026	0.026	0.152
Average unemployment rate 25-29	0.079	0.022	0.037	0.152
Average unemployment rate 30-34	0.085	0.022	0.043	0.158
Average unemployment rate 35-39	0.089	0.020	0.046	0.158
Migratory background	0.047	0.211	0	1
Schooling				
Lower secondary degree	0.377	0.485	0	1
Upper secondary degree	0.354	0.478	0	1
Entrance degree for university of applied sciences	0.063	0.242	0	1
Entrance degree for regular university	0.207	0.405	0	1

Notes: Main sample consisting of women from birth cohorts 1954 to 1967 in the West German states who lived in Germany during their reproductive years.
Source: Microcensus (2008, 2012).

2.4 Main Results

To estimate the coefficients of the linear equation, we employ standard OLS techniques with robust standard errors clustered at the state level.[3] For the regressions of the incidence of childlessness at age 40 we make use of a probit estimator and report average marginal effects, again with clustered standard errors in parentheses. The estimated βs tell us to what extent an increase in the average unemployment rate experienced during a certain age interval leads to a change in completed fertility.

Table 2.2 : Long-run effects of average unemployment rate on completed fertility

Dependent variable	Number of children above age 40		
	(1)	(2)	(3)
Effect of average unemployment rate at			
Age 20-24	-0.0007	-0.0013	0.0018
	(0.0054)	(0.0051)	(0.0054)
Age 25-29	-0.0035	-0.0048	-0.0035
	(0.0062)	(0.0062)	(0.0067)
Age 30-34	-0.0140	-0.0140	-0.0111
	(0.0143)	(0.0142)	(0.0139)
Age 35-39	-0.0078	-0.0075	-0.0065
	(0.0163)	(0.0161)	(0.0163)
Migration	No	Yes	Yes
Education	No	No	Yes
Observations	90,756	90,756	90,756

Notes: Marginal effects from OLS regressions of number of children on average unemployment rates in age groups. Sample consists of women from West Germany born between 1954 and 1967. All regressions contain cohort- and state-fixed effects. Robust standard errors clustered at the state level in parentheses, *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

We start with showing the results from a regression of the number of children on the overall unemployment rates experienced in the age groups 20-24, 25-29, 30-34, and 35-39 as in Equation 2.1. Column 1 of Table 2.2 presents the estimated coefficients from a specification with cohort- and state-fixed effects. In Column 2 we add a dummy variable indicating whether a woman has a migratory background. Our preferred specification, Column 3, also includes indicator variables for level of schooling. None of the estimates is significantly different from zero meaning that the overall unemployment rates, that women experience during their reproductive years, do not affect completed fertility. However, Becker's theory predicts two opposing effects of an increase in the risk of unemployment – a negative income and a positive substitution effect. If these effects are similar in magnitude, they quite likely cancel each other

[3] Since the number of clusters is quite small, we also make use of a cluster bootstrap procedure proposed by Cameron, Gelbach, and Miller (2008). Inference based on the bootstrap procedure is very similar. Therefore, we present standard errors that are robust to any form of heteroskedasticity and clustered at the state level.

out. To take a closer look at this issue, we replace the overall unemployment rates with the gender-specific rates.

Table 2.3 : Long-run effects of average gender-specific unemployment rate on completed fertility

Dependent variable	Number of children above age 40		
	(1)	(2)	(3)
Effect of average female unemployment rate at			
Age 20-24	0.0356**	0.0346**	0.0414**
	(0.0144)	(0.0141)	(0.0135)
Age 25-29	-0.0056	-0.0051	-0.0067
	(0.0086)	(0.0085)	(0.0092)
Age 30-34	0.0045	0.0040	0.0034
	(0.0160)	(0.0160)	(0.0160)
Age 35-39	0.0116	0.0118	0.0171
	(0.0117)	(0.0112)	(0.0112)
Effect of average male unemployment rate at			
Age 20-24	-0.0307**	-0.0303**	-0.0338***
	0.0098)	(0.0097)	(0.0096)
Age 25-29	0.0006	-0.0012	0.0013
	(0.0062)	(0.0060)	(0.0069)
Age 30-34	-0.0288	-0.0281	-0.0276
	(0.0170)	(0.0168)	(0.0167)
Age 35-39	-0.0101	-0.0103	-0.0128
	(0.0102)	(0.0099)	(0.0088)
Migration	No	Yes	Yes
Education	No	No	Yes
Observations	90,756	90,756	90,756

Notes: Marginal effects from OLS regressions of number of children on average unemployment rates in age groups. Sample consists of women from West Germany born between 1954 and 1967. All regressions contain cohort- and state-fixed effects. Robust standard errors clustered at the state level in parentheses, *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

The estimation results in Table 2.3 show the relationship between completed fertility and gender-specific unemployment rates that women have faced in different periods of their fertile lifecycle. The average female unemployment rate between age 20 and 24 has a significant influence on the number of children. The preferred specification in Column 3 indicates that a 1 percentage point increase in the average unemployment rate increases the number of children by more than 0.04. Relative to the average number of children of 1.53 this effect boosts completed fertility by 2.7 percent. However, the number of children is reduced almost to the same extent if the male unemployment rate experienced between the ages of 20 and 24 increases by 1 percentage point. In Germany, the minority of births take place with a mother between the ages of 20 and 24. We thus interpret the significant effects in this age group as "scarring" effects. If women face periods of increasing unemployment along with its inherent economic uncertainty in their early career years, they adjust their current as well as their future fertility behavior. Based on Becker's theoretical considerations, we propose a potential explanation for these results: the positive estimates of the female unemployment rate at age 20 to 24 suggest that the positive substitution effect outweighs the negative income effect for women. Difficulties with starting a career due to unfavorable labor market conditions entail diminishing opportunity costs that in turn induce women to have children. In contrast, in periods of rising male unemployment, women reduce fertility. Since men are still the main breadwinners, the reduced income has a negative effect on completed fertility. Nevertheless, we cannot say whether couples with children adjust their optimal family size nor whether women tend to remain childless when economic conditions worsen.

Table 2.4 : Long-run effects of average gender-specific unemployment rate on intensive and extensive margin

Dependent variable		Childlessness			Number of children (only mothers)	
	(1)	(2)	(3)	(4)	(5)	(6)
Effect of average female unemployment rate at						
Age 20-24	-0.0130**	-0.0128**	-0.0161***	0.0152	0.0142	0.0148
	(0.0051)	(0.0050)	(0.0045)	(0.0124)	(0.0123)	(0.0128)
Age 25-29	0.0004	0.0003	0.0013	-0.0061	-0.0054	-0.0052
	(0.0024)	(0.0025)	(0.0026)	(0.0081)	(0.0080)	(0.0079)
Age 30-34	-0.0027	-0.0027	-0.0025	0.0005	-0.0004	-0.0008
	(0.0038)	(0.0038)	(0.0037)	(0.0129)	(0.0129)	(0.0131)
Age 35-39	0.0002	0.0002	-0.0029	0.0179	0.0180	0.0176
	(0.0051)	(0.0052)	(0.0050)	(0.0149)	(0.0142)	(0.0145)
Effect of average male unemployment rate at						
Age 20-24	0.0102*	0.0102*	0.0119**	-0.0143	-0.0139	-0.0141
	(0.0055)	(0.0056)	(0.0051)	(0.0101)	(0.0099)	(0.0103)
Age 25-29	0.0007	0.0011	-0.0004	0.0040	0.0021	0.0019
	(0.0035)	(0.0035)	(0.0035)	(0.0093)	(0.0090)	(0.0089)
Age 30-34	0.0104**	0.0103**	0.0103***	-0.0145	-0.0136	-0.0130
	(0.0042)	(0.0041)	(0.0039)	(0.0157)	(0.0157)	(0.0161)
Age 35-39	0.0073**	0.0073**	0.0088***	0.0049	0.0046	0.0045
	(0.0035)	(0.0034)	(0.0030)	(0.0068)	(0.0068)	(0.0067)
Migration	No	Yes	Yes	No	Yes	Yes
Education	No	No	Yes	No	No	Yes
Observations	90,756	90,756	90,756	70,097	70,097	70,097

Notes: Columns 1-3: average marginal effects on remaining childless from probit regressions; Columns 4-6: marginal effects from OLS regressions conditional on having children. Sample consists of women from West Germany born between 1954 and 1967. All regressions contain cohort- and state-fixed effects. Robust standard errors clustered at the state level in parentheses, *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

To shed more light on this issue, Table 2.4 focuses on the extensive (childlessness; Columns 1-3) and the intensive margin (number of children conditional on having children; Columns 4-6) of the fertility decision. The later regressions are based on a reduced sample conditional on having children. Estimated coefficients of this reduced sample are small in magnitude and imprecisely estimated. Looking at the incidence of childlessness reveals that experienced unemployment rates are associated with substantial changes in the probability of remaining childless. Higher female unemployment rates in early years of the fertile lifecycle lower the probability of remaining childless, whereas increased male unemployment is positively associated with childlessness for almost all age groups. Suppose a woman faces a 1 percentage point increase in male unemployment rates for all ages from 20 to 39. This would lead to an average increase of more than 3 percentage points in her probability of remaining childless throughout her fertile life. In contrast, the experienced female unemployment rates significantly influence the decision to remain childless only between the ages of 20 and 24. A marginal increase in the average unemployment rate during this age interval lowers childlessness by almost 1.6 percentage points. Again, the significant relationship in the age group 20-24 signals an experience effect that influences future behavior since the decision to remain childless is typically not taken so early in life.

Similar to the argument made above, we argue that adverse conditions in the labor market reduce the male breadwinner's income, inducing couples to remain childless. At the same time, the costs of foregone earnings are the lowest at an early stage of a woman's career. Thus, if women substitute labor market participation with childrearing activities due to higher experienced unemployment rates, they are more likely to become mothers. However, this explanation is not expected to hold for women with a tertiary degree. These women typically do not finish their education before age 25, meaning that the labor market conditions prevailing while they are between the ages of 20 and 24 should not matter for their fertility behavior. Heterogeneity in the relationship across levels of education is examined in the following section.

2.5 Heterogeneous Effects by Level of Schooling

When we analyze the effect of unemployment experienced during a woman's reproductive years, we run up against a substantial problem. In response to increases in the unemployment rate (especially the gender-specific unemployment rate), women not only change their fertility behavior, they also change their education decisions. For instance, if labor market opportunities are unfavorable, women may decide to invest more time in their education until economic conditions improve. As a consequence, the level of fertility and the level of education are simultaneously affected by unemployment rates and the estimates are probably biased. Thus, we run the heterogeneity analysis by level of schooling since school education typically ends before age 20. More precisely, we estimate the regressions for completed fertility and childlessness separately for women with lower secondary education, with upper secondary education, and for women with a secondary degree that permits university entrance.[4]

[4] Lower and upper secondary education terminate with degrees from *Hauptschule* and *Realschule* that basically allow their recipients to start an apprenticeship or vocational training. A university entrance degree allows to study either at a regular university or at a university of applied sciences. Three-quarters of the women in our sample belong to the former two categories, one-quarter to the latter (see Table 2.1).

Table 2.5 reports the regression results of completed fertility and childlessness for three different educational groups. Columns 1 and 4 show that women with lower secondary schooling respond to the unemployment rate throughout their reproductive years. Similar to earlier results, female unemployment at the beginning of the career reduces childlessness and hence promotes completed fertility. The correlations are almost twice as large as in the pooled estimations. The female unemployment rates women experience between the ages of 30 and 34 have an effect on the number of children but leave the probability of remaining childless unaffected. In contrast, higher female unemployment experienced between age 25 and 29, when women are already established in the labor market, reduces fertility. An increase in male unemployment in the age groups 20-24 and 30-34 is associated with a significant reduction in fertility. Contrary to the theoretical predictions, male unemployment at age 25-29 has a positive effect on the number of children – thus a negative effect on remaining childless.

If women hold an upper secondary degree, female unemployment rates are negatively associated with remaining childless for all age groups (Columns 2 and 4 of Table 2.5).[5] This translates into an increase in the number of children that is statistically significant only before age 30. Similarly, higher experienced male unemployment rates before age 30 reduce fertility and increase the likelihood of remaining childless. The coefficient at age 25-29 is differently signed for women with lower and upper secondary schooling. Only for low-educated women is the experience of increasing male unemployment rates in this age group fertility enhancing. A potential explanation is a positive substitution effect for men that occurs only in low-educated couples where the opportunity costs of having children are relatively low.

Women with a university entrance degree do not react to the experienced unemployment rates (see Columns 3 and 6 of Table 2.5): it seems that increasing unemployment does not pose a threat to highly educated women. However, when women with a high level of schooling are between the ages of 25 and 29 – which may be the same time period during which their partners are graduating from university – the male unemployment rate has a negative effect on the number of children. Again, the labor market conditions for men at the start of their careers affect fertility behavior.

The heterogeneity analysis shows that the effects of unemployment experience on fertility vary strongly with women's educational attainment. Women without lower and upper secondary degrees are particularly affected by changing labor market conditions, whereas women with a university entrance degree are seemingly unaffected. Moreover, estimations suggest that a substantial part of the overall effect on completed fertility can be attributed to changes in the probability of remaining childless. To shed more light on the mechanism that explains these findings, potential channels are discussed in the following paragraphs.

5 Only female unemployment rates experienced between the ages of 30 and 34 are statistically not significant.

Table 2.5 : Long-run effects of average gender-specific unemployment rate by level of schooling

Dependent variable	Childlessness			Number of children		
Level of schooling	Low	Upper	High	Low	Upper	High
	(1)	(2)	(3)	(4)	(5)	(6)
Effect of average female unemployment rate at						
Age 20-24	-0.0148***	-0.0291***	-0.0023	0.0759***	0.0633**	-0.0188
	(0.0054)	(0.0092)	(0.0115)	(0.0176)	(0.0232)	(0.0248)
Age 25-29	0.0101**	-0.0075***	-0.0027	-0.0523**	0.0161*	0.0333
	(0.0046)	(0.0025)	(0.0080)	(0.0223)	(0.0079)	(0.0191)
Age 30-34	-0.0033	-0.0100	0.0122	0.0483*	0.0141	-0.0526
	(0.0050)	(0.0074)	(0.0137)	(0.0219)	(0.0290)	(0.0313)
Age 35-39	-0.0055	-0.0199**	0.0181	0.0170	0.0463	-0.0010
	(0.0082)	(0.0101)	(0.0120)	(0.0366)	(0.0328)	(0.0253)
Effect of average male unemployment rate at						
Age 20-24	0.0118**	0.0150*	0.0112	-0.0601**	-0.0368*	0.0027
	(0.0059)	(0.0090)	(0.0103)	(0.0213)	(0.0173)	(0.0234)
Age 25-29	-0.0160***	0.0117***	0.0107	0.0735***	-0.0345***	-0.0462*
	(0.0035)	(0.0042)	(0.0085)	(0.0194)	(0.0086)	(0.0245)
Age 30-34	0.0107*	0.0134	0.0080	-0.0756***	-0.0350	0.0385
	(0.0059)	(0.0097)	(0.0118)	(0.0217)	(0.0357)	(0.0316)
Age 35-39	0.0094	0.0094	0.0083	-0.0016	-0.0063	-0.0191
	(0.0076)	(0.0091)	(0.0115)	(0.0291)	(0.0347)	(0.0187)
Migration	Yes	Yes	Yes	Yes	Yes	Yes
Education	Yes	Yes	Yes	Yes	Yes	Yes
Observations	34,199	32,091	24,466	34,199	32,091	24,466

Notes: Columns 1-3: average marginal effects on remaining childless from probit regressions; Columns 4-6: marginal effects from OLS regressions. Sample consists of women from West Germany born between 1954 and 1967. All regressions contain cohort- and state-fixed effects. Robust standard errors clustered at the state level in parentheses, *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

2.6 Potential Mechanism

In this section we investigate two mechanisms that may be driving our main results. First, adverse labor market conditions might have an influence on marriage market outcomes, including the probability of finding a suitable partner or the stability of the partnership. If a recession raises the threat of unemployment, men and women may be more willing to accept low-quality matches on the partner market. Having a partner ensures financial support in the event one of the partners becomes unemployed. In the short run, this insurance motive of a marriage outweighs the potential negative consequences of being in an inferior partnership. However, in the long run, an increase in the probability of marriage dissolution may also imply a reduction in fertility. Unfortunately, the GMC does not provide a proper measure for the quality of a partnership. Thus, we make use of the current marital status to proxy for the quality and duration of the partnership. More precisely, we construct a dummy variable that indicates whether a woman above age 40 has never been married throughout her fertile lifecycle. Second, income is an important determinant of fertility behavior. Assuming that a higher threat of unemployment is associated with lower income, we expect that parts of the effect of unemployment on fertility can be explained by the income channel. Ideally, we would like to analyze how cumulative earnings vary with changes in unemployment rates. Since the GMC reports only contemporary (net) household income, we assume that past unemployment rates affect the income path and thus the current income. Using the household income rather than the woman's own income, accounts for the prevalence of the male breadwinner model where men make the larger monetary contribution to the household.

Table 2.6 : Mechanism of the long-run effect of average unemployment rate

Dependent variable	Never married			(Net) Household income		
	(1)	(2)	(3)	(4)	(5)	(6)
Effect of average female unemployment rate at						
Age 20-24	-0.0122**	-0.0121**	-0.0146***	0.2099***	0.2126***	0.1280**
	(0.0041)	(0.0041)	(0.0036)	(0.0570)	(0.0570)	(0.0550)
Age 25-29	-0.0005	-0.0006	0.0003	-0.0882***	-0.0896***	-0.0857***
	(0.0021)	(0.0021)	(0.0022)	(0.0186)	(0.0183)	(0.0255)
Age 30-34	-0.0054	-0.0053	-0.0051	0.0282	0.0296	0.0278
	(0.0037)	(0.0037)	(0.0036)	(0.0756)	(0.0752)	(0.0743)
Age 35-39	0.0004	0.0004	-0.0021	0.2277***	0.2272***	0.1678***
	(0.0089)	(0.0088)	(0.0080)	(0.0529)	(0.0503)	(0.0468)
Effect of average male unemployment rate at						
Age 20-24	0.0118**	0.0117**	0.0130***	-0.1823***	-0.1834***	-0.1392**
	(0.0040)	(0.0040)	(0.0038)	(0.0491)	(0.0491)	(0.0452)
Age 25-29	-0.0038*	-0.0035	-0.0047*	0.0071	0.0122	-0.0092
	(0.0021)	(0.0021)	(0.0022)	(0.0367)	(0.0362)	(0.0375)
Age 30-34	0.0073*	0.0072*	0.0071*	-0.1452**	-0.1472**	-0.1500**
	(0.0038)	(0.0038)	(0.0035)	(0.0626)	(0.0622)	(0.0656)
Age 35-39	0.0044	0.0044	0.0055	-0.1432**	-0.1425***	-0.1141**
	(0.0072)	(0.0072)	(0.0063)	(0.0443)	(0.0435)	(0.0433)
Migration	No	Yes	Yes	No	Yes	Yes
Education	No	No	Yes	No	No	Yes
Observations	90,756	90,756	90,756	90,756	90,756	90,756

Notes: Columns 1-3: average marginal effects on remaining unmarried from probit regressions; Columns 4-6: marginal effects from OLS regressions of (net) household income on average unemployment rates in age groups. Sample consists of women from West Germany born between 1954 and 1967. All regressions contain cohort- and state-fixed effects. Robust standard errors clustered at the state level in parentheses, *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table 2.6 reports average marginal effects for the probability of remaining unmarried (Columns 1 to 3) and marginal effects on the net household income (Columns 4 to 6).[6] Again, the preferred specifications control for the level of education and the migratory background. The female unemployment rate at ages 20 to 24 lowers the risk of not marrying until age 40. Hence, if, at the beginning of their careers, women experience uncertainty in the labor market, they are more likely to become married (the insurance motive). In contrast, higher male unemployment rates during early ages reduce men's attractiveness and women prefer to remain single. Both findings are in line with the results from Table 2.3: the increasing likelihood of getting married due to shifts in female unemployment rates explains increases in fertility. Positive effects on remaining unmarried due to changes in male unemployment rates ultimately reduce fertility. The negative association between *never married* and the male unemployment rate between age 25 and 29 implies that unfavorable labor market conditions for men induce women to get married.

In Columns 4 to 6 of Table 2.6 we present the coefficients on household income. Contrary to the presumed negative consequences, an increase in the female unemployment rate between the ages of 20 and 24, as well as between the ages 35 and 39, increases household income. A possible explanation is that women spend more time working in the labor market to compensate for the threat of unemployment. At the same time, a higher labor market attachment minimizes the probability of actually being affected by unemployment. Since previous results show that women tend rather to have children than to increase their labor supply in response to rising (female) unemployment rates, a more realistic explanation implies an overcompensating of the male partner for the lost income. Household income is negatively associated with female unemployment rates experienced between the ages of 25 and 29 – a crucial period during which women's income path seems to be determined. Negative shocks due to higher labor market uncertainty reduce household income significantly. The coefficients of experienced male unemployment rates are negative and significant (except for age 25-29). A higher risk of unemployment for men implies a reduction in household income and, hence, can explain the negative effect on our fertility measures.

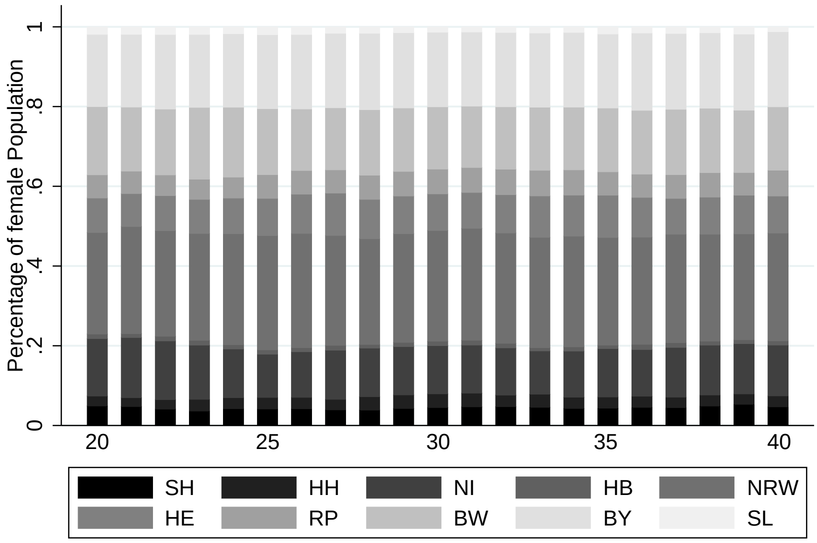
2.7 Robustness Check

To this point, our findings suggest that experienced female unemployment rates at the state level induce women to increase fertility, whereas rising male unemployment rates reduce fertility. In this section we check whether migration of women across states is a threat to the identification of the coefficients. If mobility is high, the unemployment rates measured in the current state of residence do not reflect the actual labor market conditions women face during their reproductive years. Moreover, if women with very low or very high preferences

6 Net household income consists of 24 categories with an average bin size of some 300 Euro. Since the size is not equal for all bins, it is difficult to interpret the coefficients quantitatively. However, the sign and the significance level retain their intuitive meaning.

for children move to other states to benefit from better conditions in the labor market, our estimates are biased. Therefore, we use information in the GMC about state of residence to illustrate the distribution of women of a certain cohort across federal states. As presented in Figure 2.8, the distribution of the exemplary cohort 1965 across states is very stable over time. We find no evidence that women move from states where unemployment is very high, for example, Bremen (HB),[7] to states such as Bavaria (BY) or Baden-Württemberg (BW) that have more favorable labor market conditions.[8]

Figure 2.8 : Distribution of birth cohort 1965 across states



Notes: Female birth cohort 1965; West German states are: BW: Baden-Württemberg; BY: Bavaria; HB: Bremen; HE: Hesse; HH: Hamburg; NI: Lower Saxony; NRW: North Rhine-Westphalia; RP: Rhineland-Palatinate; SH: Schleswig-Holstein; SL: Saarland.
Source: Microcensus (1985-2005).

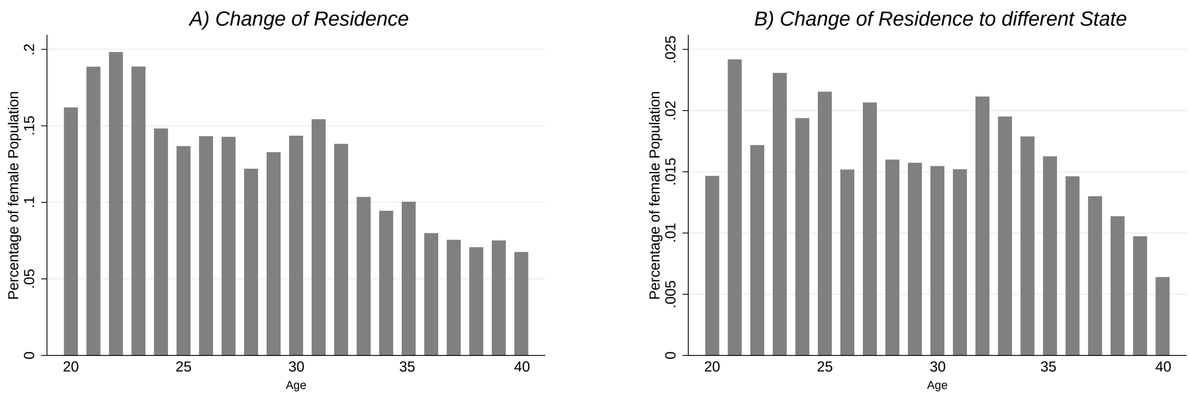
Since the GMC is a repeated cross-section, we are not able to follow individuals over time. Instead, we observe wave by wave which fraction of women from a birth cohort change residence and also to which federal state they move. The GMC allows us to plot these shares for the waves 1985 to 1997 and 2005 to 2012. For the cohort 1965, Figure 2.9 shows the share of women who have changed residence within the last 12 months, both within and across federal states. Panel A reveals that young women exhibit the highest mobility. In their early 20s, almost 20 percent of the cohort 1965 changes residence. This is not surprising, since women typically start their own household around this age. Some women might also change residence due to education, partnership, or for work-related reasons. At the end of the reproductive age,

[7] Even if the graphical analysis shows that some observations from *Bremen* might be outliers, regressions without these women provide very similar results.

[8] Due to data restrictions we are not able to monitor the mobility of cohorts born before 1965. Since mobility is an increasing phenomenon over time, we assume that older cohorts are even less likely to change their state of residence during their reproductive years.

the share of women changing residence reduces to approximately 7 percent. Nevertheless, these findings are not problematic for our empirical analysis if women move within the same state or at least do not move systematically across state borders. In Panel B we graphically assess what share of the cohort born in 1965 moves to another federal state during the fertile lifecycle. No clear pattern emerges since younger as well as older women both seem to move across federal state borders. Moreover, the share of movers is vanishingly low: on average, less than 2 percent of the female population born in 1965 finds a new residence outside their home state.

Figure 2.9 : Mobility of birth cohort 1965 within and across states



Notes: Panel A: percentage of women born in 1965 who changes residence by age; Panel B: percentage of women born in 1965 who changes federal state of residence by age. Values for ages 33 to 39 interpolated due to data restrictions.
Sources: Microcensus (1985-2005).

Finally, in a regression framework replacing the state-specific unemployment rates with the national unemployment rates, the threat that our estimates are confounded by systematic migration is reduced. Table 2.7 presents OLS and probit estimates for the initial regressions where state-level unemployment is substituted by national unemployment. Note that in this setting the unemployment rate varies across cohorts, but no longer across regions. Thus, the remaining variation in the unemployment rates is based on differences in the labor market conditions of the 14 birth cohorts in the sample. The estimates reveal that the sign of the coefficients is in line with previous results; the magnitude is even higher for almost all age groups, but the standard errors are up to 10 times larger than in Table 2.3 or 2.4. Most likely, the variation in the national unemployment measures is not sufficient to achieve reasonable standard errors. Hence, we are not able to estimate the coefficients precisely enough to statistically distinguish them from zero. Nevertheless, we do not interpret the estimates as contradicting the main results.

The robustness checks aim at providing evidence against selective migration across federal states. We show graphically, as well as in a regression framework, that our main estimates are not confounded by the mobility patterns of women of reproductive age. Indeed, using national instead of state-specific unemployment rates confirms the main insights from the previous analysis.

Table 2.7 : Long-run effects of national gender-specific unemployment rate

Dependent variable	Childlessness			Number of children		
	(1)	(2)	(3)	(4)	(5)	(6)
Effect of average female unemployment rate at						
Age 20-24	-0.0410	-0.0386	-0.0402	0.0238	0.0118	0.0107
	(0.0397)	(0.0396)	(0.0416)	(0.1198)	(0.1183)	(0.1237)
Age 25-29	-0.0068	-0.0082	-0.0088	0.0706	0.0785	0.0767
	(0.0201)	(0.0197)	(0.0204)	(0.0543)	(0.0527)	(0.0551)
Age 30-34	-0.0211	-0.0223	-0.0237	0.0969	0.1022	0.1053
	(0.0325)	(0.0327)	(0.0363)	(0.0930)	(0.0945)	(0.0974)
Age 35-39	-0.0358**	-0.0351**	-0.0384**	0.0345	0.0311	0.0331
	(0.0165)	(0.0169)	(0.0158)	(0.0559)	(0.0562)	(0.0548)
Effect of average male unemployment rate at						
Age 20-24	0.0446	0.0424	0.0435	-0.0329	-0.0220	-0.0187
	(0.0387)	(0.0387)	(0.0414)	(0.1121)	(0.1111)	(0.1168)
Age 25-29	0.0146*	0.0154*	0.0135	-0.0677***	-0.0720***	-0.0652**
	(0.0087)	(0.0086)	(0.0088)	(0.0194)	(0.0192)	(0.0208)
Age 30-34	0.0049	0.0051	0.0009	-0.0192	-0.0187	-0.0130
	(0.0135)	(0.0135)	(0.0137)	(0.0318)	(0.0320)	(0.0312)
Age 35-39	0.0074	0.0065	0.0054	0.0331	0.0374	0.0415
	(0.0162)	(0.0162)	(0.0171)	(0.0402)	(0.0399)	(0.0436)
Migration	No	Yes	Yes	No	Yes	Yes
Education	No	No	Yes	No	No	Yes
Observations	90,756	90,756	90,756	90,756	90,756	90,756

Notes: Columns 1-3: average marginal effects on remaining childless from probit regressions; Columns 4-6: marginal effects from OLS regressions. Sample consists of women from West Germany born between 1954 and 1967. All regressions contain cohort- and state-fixed effects. Robust standard errors clustered at the state level in parentheses, *** p<0.01, ** p<0.05, * p<0.1.

2.8 Conclusion

The labor market conditions women experience throughout their reproductive years are expected to influence completed fertility as well as childlessness. Standard microeconomic theory suggests that unemployment spells and the wage reductions that accompany increases in the unemployment rate have two opposing effects on fertility: a negative income effect and a positive substitution effect. Which of the two dominates remains, ultimately, an empirical question. In addition, short-run studies that analyze the contemporaneous relationship between unemployment and birth rates are not able to distinguish a mere tempo effect from a quantum effect.

Thus, we intend to answer in this chapter how the labor market conditions women face during their reproductive years affect the fertility behavior. To this end, we make use of the German Microcensus from 2008 and 2012, focusing on the incidence of childlessness and the number of children born to each woman from the birth cohorts 1954 to 1967. Applying standard OLS and probit regression methods, the results suggest that increases in the five-year averages of state-specific unemployment rates have a significant impact on fertility decisions. More precisely, experiencing higher female unemployment rates during early career years increases fertility, whereas rising male unemployment rates have the opposite effect. This relationship is mainly driven by changes in the probability of remaining childless. For instance, if the average female unemployment rate at age 20 to 24 increases by 1 percentage point, the likelihood that a woman remains childless decreases by 1.6 percentage points. However, if the male unemployment rate rises in the same period, childlessness increases by 1.2 percentage points. Given the baseline of 23 percent childlessness in our sample, the estimates correspond to a 7 percent decline or a 5.3 percent growth in childlessness. The changes in the probability of remaining childless translate into lower completed fertility. The number of children is reduced by 2.2 percent and raised by 2.7 percent if male, respectively, female unemployment increases in the age group 20-24. Since most fertility decisions are taken after age 25, we argue that our results represent a "scarring" effect of unemployment that influences future fertility behavior. In addition, the heterogeneity analysis reveals that women with lower and upper secondary degrees respond most strongly to changes in unemployment rates. Two mechanisms may explain the findings: first and foremost, experienced unemployment rates have a substantial influence on marriage market outcomes and, second, they also impact household income. We argue that the results are robust to potential distortions caused by women migrating to states with more favorable labor market conditions.

From a policy perspective, our research opens a new way of thinking about labor market and fertility policy. Since labor market experience significantly affects fertility behavior – especially during the early years of a career – policy measures regarding male and female unemployment need to be evaluated for the extent to which they may interfere with family policies. A narrow perspective focused solely on the labor market may have severe consequences for the fertility decisions of couples.

3 Fixed-Term Employment and Fertility: Evidence from German Micro Data[1]

3.1 Introduction

One of the most striking facts about labor market development in many European countries over the last decades is the tremendous increase in fixed-term employment. By 2012, the average share of temporary[2] employees among all 25- to 54-year-old employees was, on average, 12 percent in all European OECD countries (OECD, 2014). Germany has witnessed a particularly strong rise in fixed-term employment in recent years. By 2012, almost 50 percent of new employment contracts were of limited duration. Fixed-term employment is particularly concentrated among young adults in their early careers – a period in life that is crucial both for career progression and family formation. Recent evidence shows that adverse labor market conditions at the beginning of the career can lead to severe and persistent earning losses (e.g., Oreopoulos, von Wachter, and Heisz, 2012). Temporary employment might cause a similarly negative labor market path dependence via repeated episodes of temporary employment, decelerated wage progression, and higher likelihood of future unemployment (Hagen, 2002; Bruno, Caroleo, and Dessy, 2012; Booth, Francesconi, and Frank, 2002; Pavlopoulos, 2009). Previous studies mainly link contemporaneous temporary employment and fertility responses at different stages of the lifecycle and produce mixed evidence. Their approach neglects the potential endogeneity of fixed-term contracts as well as any path dependence. The empirical literature on whether and, if so, how increased levels of economic uncertainty due to unstable working contracts at the beginning of a career have spill-over effects on other domains of life is rare.

The main objective and contribution of this chapter is to fill this gap by empirically assessing the implications for subsequent fertility of entering the labor market with a fixed-term contract. To this end, we focus on several cohorts of graduates from vocational training or tertiary education and follow them for their first 10 years in the labor market. We analyze the timing of first birth (tempo effect) as well as the number of children (quantum effect) in the short to medium run. We also contribute to the literature by discussing and addressing the potential selection of individuals into different types of contracts. To reduce possible omitted variable bias, we exploit our rich and unique data set and include a large set of novel control variables.

[1] Joint with Natalia Danzer, published as "Fixed-Term Employment and Fertility: Evidence from German Micro Data" in CESifo Economic Studies 62 (4), 2016. This and the last chapter were part of a project funded by IBS (Warsaw) by grant number 3_SG/IBS/2014

[2] Throughout the third and fourth chapter of this dissertation, the terms "fixed-term contract" and "temporary contract" are used interchangeably.

Based on the survey years 1995 to 2012 of the German Socio-Economic Panel (SOEP), we apply probit and poisson estimation methods on a pooled sample of women who are childless when they finish their education and enter the labor market. Our results for natives confirm that starting a career with a fixed-term contract is negatively associated with subsequent fertility: we find an increased postponement of first birth and a reduction in the number of children in the first 10 years after graduation. These results also hold when we expand the sample and include migrants in the analysis; however, the effects in the full sample are slightly less pronounced. Furthermore, we show that fixed-term employment seems to particularly affect the fertility decisions of women with secondary education. We find no significant correlations between job uncertainty and fertility for the subsample of men. As fertility decisions, as well as holding a fixed-term contract, may be driven by unobserved heterogeneity, we address potential endogeneity concerns on two fronts: first, by including many important and previously neglected control variables and, second, by showing that entering the labor market with a fixed-term contract is related neither to family nor to career preferences. Against this background, we reckon that our results actually reflect a robust, negative relationship between job uncertainty in the early career and the timing and number of children. Even though we are not able to examine the effect on completed fertility in this empirical setup, our results suggest that completed fertility might be negatively affected as well.

Our research contributes to the growing literature on the relationship between economic uncertainty and fertility. Several empirical studies suggest that fertility reacts pro-cyclically to macroeconomic conditions: higher unemployment rates are generally associated with reduced fertility rates and vice versa (e.g., Adsera, 2005; Adsera and Menendez, 2011; Goldstein, Kreyenfeld, Jasilioniene, and Örsal, 2013). Analyses of how individual unemployment affects fertility yield mixed evidence (e.g., Del Bono et al., 2012; Kreyenfeld, 2010). Focusing on perceived economic uncertainty using German data, Bhaumik and Nugent (2011) and Hofmann and Hohmeyer (2013)[3] find a reduction in fertility and a study by Kreyenfeld (2010) confirms this result for a subsample of highly educated women.

Temporary employment is considered to be one form of economic uncertainty. Unfortunately, evidence on the relationship between fixed-term employment and fertility is scarce and inconclusive. For Germany, Gebel and Giesecke (2009) find no evidence that fixed-term contracts influence the fertility decisions of young couples, whereas the results by Schmitt (2012) suggest a negative impact. Tölke and Diewald (2003) find evidence for a postponement of first birth due to economic uncertainty for young men. Kind and Kleibrink (2013) disagree, concluding that time-limited contracts postpone childbearing only for women, not for men. Studies from other European countries report less ambiguous results. For Spain, studies by Ahn and Mira (2001) and De la Rica and Iza (2005) conclude that fixed-term employment has a negative effect on the decision to marry and delays childbearing. Similarly, Sutela (2012) reports that

[3] The study by Hofmann and Hohmeyer (2013) stands out from the other studies in its attempt to correct for the potential endogeneity of subjective economic uncertainty.

in Finland, fixed-term employment is negatively associated with entering parenthood. All these studies have in common that they focus mostly on empirical associations between holding a fixed-term contract and fertility. They neither consider the potential endogeneity problems of fixed-term contracts (which might be increasing in labor market experience as well as previous number of children) nor the potential path dependence of entering the labor market with a contract of limited duration.

The remainder of the chapter is organized as follows. The following section discusses the theoretical background. Section 3.3 motivates our empirical analysis by descriptively showing the relationship between fixed-term employment, economic uncertainty, and the fertility decisions of young couples. Section 3.4 introduces our data and our empirical approach. The main results, as well as several sensitivity and subgroup analyses, are presented in Section 3.5. Section 3.6 concludes and discusses potential policy implications.

3.2 Theoretical Background

The main microeconomic theory of fertility dates back to Becker (1960, 1965, 1991). In his work, children are modeled as normal consumption goods and fertility decisions are based on the relative costs and benefits of having children. These models are also referred to as opportunity cost models or price-of-time models since wage increases not only induce a positive income effect (raising the demand for children), but also a negative opportunity cost effect (substitution effect). Direct opportunity costs arise due to foregone earnings during the time that parents take off from work or reduce their working hours to care for their children. Additionally, childrearing incurs indirect opportunity costs or future career costs through human capital depreciation during employment interruptions, which in turn negatively impact the future earnings profile. The overall effect of changes in the wage on fertility depends on the relative size of these opposing income and opportunity costs of time effects. However, as women in many countries traditionally devote more time to childrearing than do men, the opportunity costs argument is mainly applied to women. In contrast, wage increases of men are expected to have a positive income effect.[4]

Based on this theoretical framework, fixed-term employment could affect fertility in several ways. First, demand for children should be reduced as the wages of fixed-term employees are, on average, lower than those of their colleagues with permanent contracts (income effect). Second, lower wages also imply smaller direct opportunity costs of childbearing, thus fostering the demand for children. Third, it is likely that fixed-term contracts further exacerbate the future career costs of children through increased economic uncertainty: temporary employment is generally associated with a higher risk of future unemployment Hagen (2002). In addition, unemployed women with children might be disadvantaged in the labor market

[4] Becker and Lewis (1973) extend this framework by incorporating the possibility that parents trade off the quantity and quality of children. In their model, a rise in income does not necessarily increase the number of children, but can instead raise the quality per child (e.g., through additional investments in education).

and might find it more difficult to find a job than childless women (Del Bono et al., 2012). Taken together, having children while in a temporary job is likely to put women in an even more unfavourable situation. Moreover, fixed-term employment might hamper the success of future job search as human capital accumulation in temporary jobs is generally decelerated due to smaller investments in firm- or task-specific skills (Albert, García-Serrano, and Hernanz, 2005). Hence, these additional future career costs of children exclusively related to fixed-term contracts and their associated economic uncertainty should reduce the demand for children and might deter women from entering motherhood while on a fixed-term contract.

Overall though, the standard economic theory of fertility does not predict an unambiguous effect of fixed-term employment on fertility. Moreover, this static framework neither allows drawing any conclusions about the optimal timing of childbearing, nor does it explicitly account for the potential role of economic uncertainty.[5] These two aspects are jointly captured in the economic models of fertility proposed by Ranjan (1999) and Iyer and Velu (2006):[6] in both models, childbirth decisions are considered irreversible and parents have the option to postpone childbearing to future periods. The intuitive implication in both cases is that in light of future uncertainties (about own income or the net benefit of children), it might be worthwhile for parents to postpone their childbearing decision to the (next) period when the uncertainty is resolved. This way parents can avoid entering parenthood and incurring its associated irreversible costs in a bad state of the world when having children is not optimal. Against this background, fixed-term contracts should increase the option value of postponing the childbearing decision and thus cause a delay in parenthood. The main driver will be the economic uncertainty associated with fixed-term contracts due to a more unstable future employment and income path.

3.3 Descriptive Evidence

This section descriptively motivates economic uncertainty as the main channel through which starting a career with a fixed-term contract affects fertility decisions. We employ two large-scale and nationally representative German micro-data sources – the German Socio-Economic

[5] The timing of fertility is addressed in so-called lifecycle or dynamic models of fertility. However, these models do not yield clear predictions and are also difficult to test econometrically (Hotz, Klerman, and Willis, 1997). Hotz et al. (1997) and Gustafsson (2001) provide comprehensive reviews of the theoretical and empirical literature on the postponement of maternity, but do not address the potential role of economic uncertainty.

[6] Iyer and Velu (2006) incorporate a real options approach in their theoretical framework. Their model aims at explaining demographic processes in developing countries in which economic uncertainty increases the insurance motive for children (positive fertility effect) and at the same time the option to wait (negative fertility effect). However, as the insurance motive should be of little, if any, relevance in developed countries, the model predicts an unambiguous delay of childbearing due to economic uncertainty in the German context (Bhaumik and Nugent, 2011).

Panel (SOEP)[7] and the Panel Analysis of Intimate Relationships and Family Dynamics (pairfam)[8] – to shed light on (a) the degree of economic uncertainty and the path dependence associated with a career start in a fixed-term employment and (b) the role played by economic security in the decision to have children.

3.3.1 Fixed-Term Contract and Economic Uncertainty

We employ several subjective as well as objective measures to capture economic uncertainty. Controlling for various covariates, Table 3.1 shows predicted probabilities for individual perception of job security and own economic situation by type of contract. Over 48 percent of women and almost 47 percent of men with a regular, permanent contract report that they are not concerned at all about job security, whereas only 22 percent of the temporary employed women and 28 percent of men do not worry about job security. Moreover, almost one-third of female workers with a fixed-term contract are very concerned about job security, whereas this is true of only 13 percent of workers on permanent contracts. A similar picture emerges if we look at assessments of the general economic situation. While only 20 (16) percent of all female (male) permanent employees report that they are very concerned about the general economic situation, almost 29 (23) percent of their temporary colleagues are worried. Thus, self-reported job and economic uncertainty is indeed much more pronounced among temporary than among permanent employees.

Table 3.1 : Worries about job security and economic situation (in percent)

	Men		Women	
	Very concerned	Not concerned at all	Very concerned	Not concerned at all
A. Are you concerned about job security?				
Permanent contract	11.3	46.9	12.8	48.1
Fixed-term contract	23.7	27.5	33.2	22.1
B. Are you concerned about your own economic situation?				
Permanent contract	15.7	28.0	20.1	24.3
Fixed-term contract	23.0	19.5	28.8	16.7

Notes: SOEP 1995-2012, employed men and women, 18-49 years, predicted probabilities controlling for age, education, migratory background, net wage, occupation, industry, year, and federal state; residual category *"somewhat concerned"*.

We find a similar pattern when using several objective measures of economic uncertainty, namely, income volatility, future unemployment risk, and wage progression. Our first measure, income volatility, reflects the degree of uncertainty in wages attached to fixed-term

[7] For more information regarding the SOEP, see Section 3.4.

[8] The pairfam study (Huinink, Brüderl, Nauck, Walper, Castiglioni, and Feldhaus, 2011) covers the complex processes of partnership development, family formation, and childrearing, as well as intergenerational relations. It was first conducted in 2008/2009, and consists of three birth cohorts. The first wave of the birth cohort 1981-1983, which is used in this section, comprises 1,238 childless women and 1,659 childless men.

employment. Following Bonin, Dohmen, Falk, Huffman, and Sunde (2007), we analyze the variance of the residual part of a Mincer wage regression using individual net and gross labor income. If the variance of the unexplained part for temporary employees exceeds that for permanent workers, uncertainty is higher for the former.

Table 3.2 : Variance ratio test for unexplained part of wages by type of contract

	Men		Women	
	Net wages	Gross wages	Net wages	Gross wages
A. Mean values				
Permanent contract	1748.2	2743.1	1223.4	1905.4
Fixed-term contract	1453.5	2252.0	985.3	1554.8
B. Variance of unexplained part				
Permanent contract	0.258	0.281	0.223	0.268
Fixed-term contract	0.391	0.430	0.312	0.354
C. Variance ratio test				
F-statistic	0.660	0.654	0.714	0.757
p-value	0.000	0.000	0.000	0.000

Notes: SOEP 1995-2012, employed men and women, 18-49 years, controlling for gender, age, education, migratory background, experience, tenure, net wage, occupation, industry, year, and federal state.

Table 3.2 shows variances, test statistics, and p-values for the variance ratio test by gender. On average, wages of temporary employed men and women are lower than those of their colleagues with permanent contracts. Furthermore, the earnings are more volatile and therefore more uncertain for temporary workers. The formal test confirms this result since the F-statistic leads to a rejection of the null hypothesis of equal variances (p-value < 0.001). Individuals with a fixed-term contract experience significantly higher earnings uncertainty compared to individuals in permanent jobs.

Descriptive evidence regarding future unemployment risk and future wages related to fixed-term employment is provided in Table 3.3. We present future labor market outcomes for men and women whose first job is on a temporary or on a permanent basis. The picture that emerges supports the notion of a negative path dependence of starting a career with a fixed-term contract. The risk of subsequent unemployment is substantially higher if the first job has a limited duration. During the first 10 years after labor market entry, these individuals are more likely to have had at least one unemployment spell than their colleagues who started with a permanent contract. On average, they have also experienced more periods of unemployment. This finding holds for men as well as for women (Panel A and B). In contrast, conditional on employment, the net wages of both groups are only slightly different at the beginning of the career and converge over time (Panel C).

However, when including unemployed individuals in the wage calculations, the earnings gap widens for men but turns around for women (Table 3.3, Panel D).[9] Hence, while we do not find

[9] We included unemployed and inactive individuals in these earnings calculations by assigning them a zero labor market income.

Table 3.3 : Path dependence of type of the first contract

	Men			Women		
	Years after graduation					
Starting a career with	3	5	10	3	5	10
A. Incidence of at least one unemployment spell (in percent)						
Permanent contract	9.6	14.3	18.9	11.3	20.3	34.4
Fixed-term contract	23.2	26.5	29.3	15.9	26.2	39.7
B. Number of unemployment spells						
Permanent contract	0.111	0.212	0.447	0.137	0.313	0.834
Fixed-term contract	0.293	0.518	1.061	0.182	0.392	0.944
C. Net wages						
Permanent contract	1548.9	1716.9	2154.0	1198.3	1227.9	1349.2
Fixed-term contract	1530.2	1612.0	2200.5	1123.2	1274.1	1419.4
D. Net wage (UE = 0)						
Permanent contract	1462.6	1620.7	2061.0	1020.6	929.2	998.2
Fixed-term contract	1284.6	1417.8	1905.3	884.9	989.9	1104.0

Notes: SOEP 1995-2012, men and women, 18-49 years.

consistent evidence for differences in actual wage profiles (see Booth et al., 2002; Pavlopoulos, 2009), earning stability is much lower among those employees who entered the labor market with a fixed-term contract.

To sum up, descriptive statistics suggest that holding a fixed-term contract is indeed associated with a high degree of uncertainty and negative future career consequences. This holds for subjective as well as objective measures of economic uncertainty.

3.3.2 Job Security and First Birth

Table 3.4 : Conditions for having children (in percent)

	Men	Women
A. Financial affordability must be satisfied before first birth		
Permanent contract	76.9	77.8
Fixed-term contract	77.9	78.8
B. Compatibility with work life must be satisfied before first birth		
Permanent contract	63.1	68.3
Fixed-term contract	64.0	69.2

Notes: pairfam 2009, childless men and women, 24-29 years, predicted probabilities controlling for gender, age, education, migratory background, parental education, importance of career and family, and federal state.

How does the economic situation affect the fertility decisions of young couples? Table 3.4 lists predicted probabilities by gender for the two most often mentioned conditions for having children (pairfam data). Both conditions are related to work life: parenthood has to be financially affordable and has to be compatible with the work situation.[10] The numbers differ only

[10] Other, less important, conditions are the availability of childcare or leisure-time pursuits.

marginally by type of contract: for instance, 77 (78) percent of the male (female) respondents in permanent employment report that financial affordability must be satisfied before having a first child, while 78 (79) percent with fixed-term contracts do so. Hence, the groups are not different in their desire for economic security and stability before having children. This suggests that young people do not self-select into fixed-term contracts with respect to these observable family- and work-related attitudes.

The answers as to whether these conditions are satisfied are in striking contrast to this similarity in desires (Table 3.5). The differences between individuals with fixed-term and those with permanent contracts are substantial. Male respondents with a fixed-term contract are 15 (13) percentage points less likely to rate the financial (job-related) situation as good enough to become parents. Women with a fixed-term contract assess their financial preconditions for entering motherhood even worse than do men: only 48 percent report that financial conditions allow them to have a baby. This is almost 15 percentage points less than women with a regular contract. This descriptive evidence indicates that job-related factors play a major role in young couples' decisions to have children. Independent of the type of employment contract, individuals prefer to be economically secure before having children. However, this condition is significantly less often satisfied for temporary than for permanent employees.

Table 3.5 : Satisfaction of conditions for having children (in percent)

	Men	Women
A. Financial affordability is satisfied		
Permanent contract	60.0	62.3
Fixed-term contract	45.5	47.9
B. Compatibility with work life is satisfied		
Permanent contract	59.7	59.5
Fixed-term contract	46.6	46.4

Notes: pairfam 2009, childless men and women, 24-29 years, predicted probabilities controlling for gender, age, education, migratory background, parental education, importance of career and family, and federal state.

The descriptive analysis in this section suggests that (a) fixed-term contracts are indeed associated with increased economic uncertainty and (b) economic uncertainty seems to deter young couples from entering parenthood. The resulting hypothesis that temporary employment induces a postponement of first birth (or even a negative fertility effect) will be empirically assessed in a regression framework in the next section.

3.4 Data and Empirical Strategy

3.4.1 Data, Sample Restrictions, and Variables

We employ the German Socio-Economic Panel (SOEP), which has provided annual and nationally representative panel data since 1984 (Wagner, Frick, and Schupp, 2007). In 2012, the

SOEP covered more than 20,000 individuals living in over 12,000 households. SOEP contains detailed information on a variety of individual as well as household-specific socioeconomic characteristics. Moreover, the respondents provide information about their labor market history as well as their current labor force status. Most importantly, we know whether their employment contract was permanent or temporary.

We focus on the waves 1995 to 2012 since consistent information on the type of employment contract for all working individuals was collected only from 1995 onward.[11] Our main sample consists of native women who are childless, 18 to 30 years old at the time of graduation, and for whom information on subsequent fertility for at least 10 years after graduation is available.[12] Furthermore, we restrict the main analysis to women who have obtained at least a vocational degree (i.e., ISCED codes 3 to 6).[13] The restriction concerning age at graduation is imposed because we want to ensure that the biological preconditions for becoming pregnant and giving birth are not too different in the 10 years following graduation. Women who finish their education after their 31st birthday have a comparatively narrower biological time interval in which to postpone the birth of their first child. Furthermore, for these older women it seems more likely that fertility and education choices are made simultaneously. We end up with a balanced sample of 267 women from the graduation cohort 1995 to 2003 whom we observe at the start of their career and at least once more 10 years after graduation.

The outcome variables are the following. First, to measure the timing of first birth, we create a set of dummy variables taking the value 1 if a woman has had a first child in year z after graduation or labor market entry[14] (with $z = 4, ..., 10$) and remains 1 from this point onward. The dummy variable is 0 if the woman is still childless in that particular year z. For instance, for a woman who has remained childless until her fifth year on the labor market and has a child from year 5 on, we code the outcome variable as 0 for years 1 to 4 and as 1 for years 5 to 10. Second, to analyze the quantum effect, that is, whether a postponement of first birth also translates into a decline in the total number of children, we generate a set of variables indicating the number of children, again in each of years 4 to 10 after career start.[15] The

[11] Respondents who did not report a job change were excluded from this question. Thus, switching from temporary to permanent employment at the same employer was not part of the questionnaire up to 1995.

[12] To be precise, these women are childless and not pregnant when they enter the labor market. For those individuals who did not participate in each wave of the survey, we filled in the missing fertility information retrospectively using the birth history reported in year 10 after graduation.

[13] We employ these restrictions to increase the homogeneity of our small sample and to drop outliers (e.g., there are only nine observations with less than secondary education). However, our main results hold when we relax all sample restrictions.

[14] We use the expressions “year of graduation” and “year of labor market entry” interchangeably even though, technically speaking, we measure the information on the first job in the calendar year after graduation. The main reason for doing so is that we do not have information on the exact date of graduation. Our approach ensures that the job information is indeed measured after graduation.

[15] Due to the low number of first births two and three years after graduation (for the number of children estimations, also four years after graduation), there is not enough variation to estimate the regression until this point.

dependent variables thus reflect the proportion of women who have had their first child after a certain amount of time after graduation, as well as the average number of children.

The main explanatory variable is a binary variable indicating whether the first job after graduation had a fixed-term employment contract or a permanent contract (we also control for whether the respondent is unemployed after graduation). A great advantage of the SOEP data is the variety of unique information provided about the respondent. The data allow us to include a large set of controls for individual, background, personality, and first job characteristics, as well as for partnership status at labor market entry. All control variables are either predetermined (i.e., determined before labor market entry) or measured in the year of career start. Individual control variables are age at graduation, years of education, and being born in East Germany.[16] As a proxy for the respondent's predetermined family- and career-related background we include variables indicating whether her mother has tertiary education, whether the respondent's mother was employed when the respondent was 15 years old, whether the respondent has siblings, and mother's age at the respondent's birth. Personality traits and self-reported attitudes are captured by the "Big 5", locus of control, and Kluckhohn's importance of life area measures. More specifically, five variables on a scale from 1 (low) to 7 (high) reflect the respondent's *openness to new experience*, *agreeableness*, *conscientiousness*, *extraversion*, and *neuroticism*. The locus of control variable takes on high values if the respondent is convinced that her own actions determine her success in life. Four binary variables indicate individual career- and family-related attitudes and values. The latter take the value 1 if a woman claims that having children, being in a happy partnership/marriage, being able to afford something, or having a career is important or very important in her life.[17] In addition, a dummy variable indicates whether a woman is risk averse, that is, reports a (very) low subjective willingness to take risks. As regards the characteristics of the first job, we include only very rough indicators, namely, dummy variables for blue- versus white-collar occupations and five industry dummies for the main economic sectors.[18] The prevalence of fixed-term employment differs across industries and occupational groups. Similarly, women with particular fertility preferences might self-select into particular industries and occupational groups. We control for these job characteristics to ensure that our results do not reflect spurious correlations between temporary jobs and fertility.[19] Finally, we include a dummy variable indicating whether the person is in a partnership at career start.

[16] In the robustness checks we include migrants in the sample and add a control variable for migration background.
[17] Questions on personality traits and attitudes are not included in every wave of the survey. To exclude possible feedback effects of labor market or partnership experiences on personality traits and family and career attitudes, we use only the first available observation. According to the psychological literature, personality traits are stable in adulthood. The majority of women answered this question at around the age of 21.
[18] The five main industries are generated according to the classification of the Federal Statistical Office (destatis). These are manufacturing, construction, trade and transportation, financial services, and public and other services. We dropped the only respondent working in the agricultural sector.
[19] However, our main results are robust to excluding these industry and occupational dummy variables.

Table 3.6 : Descriptive statistics by employment status

	First job permanent contract			First job fixed-term contract			Unemployed after graduation		
	Mean	SD	N	Mean	SD	N	Mean	SD	N
A. Outcome variables									
First birth after 3 years	0.136	0.379	177	0.171	0.343	76	–	–	14
First birth after 5 years	0.322	0.469	177	0.263	0.443	76	0.071	0.267	14
First birth after 7 years	0.446	0.499	177	0.421	0.497	76	0.286	0.469	14
First birth after 10 years	0.605	0.490	177	0.553	0.501	76	0.643	0.497	14
Number of children after 5 years	0.379	0.592	177	0.329	0.598	76	0.071	0.267	14
Number of children after 7 years	0.588	0.726	177	0.579	0.753	76	0.286	0.469	14
Number of children after 10 years	0.927	0.892	177	0.868	0.929	76	0.714	0.611	14
B. Control variables									
Age at graduation	23.48	2.682	177	24.71	3.382	76	23.43	3.368	14
Years of education	13.11	2.483	177	13.97	2.940	76	12.21	2.463	14
Born in East Germany	0.367	0.483	177	0.434	0.499	76	0.571	0.514	14
In partnership after graduation	0.684	0.466	177	0.737	0.443	76	0.786	0.426	14
Married after graduation	0.056	0.232	177	0.092	0.291	76	0.286	0.469	14
Cohabiting after graduation	0.395	0.490	177	0.408	0.495	76	0.357	0.497	14
Openness	4.529	1.183	177	4.553	1.199	76	4.357	1.017	14
Agreeableness	5.471	0.842	177	5.360	0.955	76	5.500	0.894	14
Conscientiousness	5.914	0.778	177	5.893	0.811	76	6.333	0.692	14
Extraversion	5.053	1.143	177	4.989	1.075	76	5.238	1.505	14
Neuroticism	4.339	0.775	177	4.364	0.820	76	4.643	0.891	14
Risk aversion	0.079	0.271	177	0.039	0.196	76	0.071	0.267	14
Locus of control	4.141	0.614	177	4.107	0.626	76	4.357	0.924	14
Importance of having children	0.650	0.478	177	0.750	0.436	76	0.714	0.469	14
Importance of career	0.910	0.288	177	0.895	0.309	76	1.000	0.000	14

Notes: Main sample, including native women with at least secondary education, younger than 31 years, and childless at graduation. Note that except for age at graduation and years of education, all differences between temporary and permanent workers are not significant at the 10 percent level.

Table 3.6 contains summary statistics of the fertility measures as well as the covariates by type of first job contract and employment status after graduation. The share of women in regular jobs who enter parenthood increases from almost 14 percent within the first three years after graduation to more than 60 percent after 10 years. As soon as four years after graduation there is a greater likelihood that women starting work with a permanent contract (vs. a temporary contract) have become mothers. The gap remains constant at between 2 and 6 percentage points. A similar, albeit much weaker, pattern emerges when we consider the total number of children. The lower panel of Table 3.6 shows differences in characteristics of temporary and permanent employed women. Women who start their career with a contract with limited duration are significantly older at labor market entry and have more education. The personality traits are highly similar across groups but the proxy for child preferences (self-rated importance of having children) is substantially higher and almost significant at the 10 percent level for women in fixed-term jobs. Women with different types of contracts do not seem too different, yet controlling for observable characteristics should improve the precision of our estimates.

3.4.2 Empirical Strategy

We more thoroughly examine the effects on fertility of starting a career with a fixed-term contract in a regression framework. Our basic empirical strategy is to compare women entering the labor market with a fixed-term contract with their counterparts on permanent contracts in terms of their short- to medium-run fertility. The empirical setup is comparable to that employed in papers studying future effects of adverse labor market conditions at the beginning of the career (see, e.g., Kahn, 2010; Liu, Salvanes, and Sørensen, 2012; Stevens, 2008). We take advantage of the fact that even though fixed-term employment increased tremendously over the last 15 years, not all regions or industries were equally affected. Thus, a substantial part of the variation in starting a career with a fixed-term contract is caused by this exogenous, labor-demand driven increase in temporary employment. The underlying empirical model can be described in a very simple linear regression form as follows:

$$Y_{it_0+z} = \beta FT_{it_0} + \gamma UE_{it_0} + \delta' X_{it_0} + \varphi_{st_0} + \mu_{t_0} + \varepsilon_{it_0} \tag{3.1}$$

Y_{it_0+z} denotes the outcome of interest for woman i in period $t_0 + z$, where z indicates the year after graduation or end of vocational training. FT_{it_0} is an indicator variable for starting a career with a fixed-term contract and UE_{it_0} indicates whether an individual experiences an unemployment spell after graduation. Therefore, the base category in our regressions will be starting a career with a permanent contract. Further, X_{it_0} are observed predetermined individual and job characteristics measured in t_0, φ_{st_0} is the federal state of the first job, μ_{t_0} is year of graduation, and ε_{it_0} is the unobserved error term.

It is crucial to include variables that influence the probability of holding a fixed-term contract and might also and simultaneously correlate with the fertility decision. Not controlling for these variables may leave them in the error term as confounding factors, which may cause spurious correlations between fertility and holding a fixed-term contract at labor market entry. If workers with particular characteristics or preferences for children self-select into particular types of contracts, our estimates would be biased. This aspect is usually ignored in previous studies analyzing the role of fixed-term employment on fertility outcomes. For example, one might think of an individual who is strongly risk averse and therefore will be most reluctant to work with a fixed-term contract. Most likely, this person will keep looking for a job until she finds a permanent position. At the same time, her risk aversion might make her less likely to have a child since entering parenthood obviously involves a great deal of uncertainty. The presence of such individuals in our sample would cause a positive bias and our results would underestimate the true effect. In contrast, we can expect a negative bias if, say, a freedom-loving, flexible woman is more likely to accept a fixed-term contract and also less likely to have a strong preference for children. Fortunately, the SOEP data allow us to control for a variety of individual characteristics and preference indicators. Thus, all regressions include controls for the degree of risk aversion as well as for personality traits and general attitudes. For instance, family preferences are controlled for by Kluckhohn's importance of life area measures. Furthermore, we test whether any of the predetermined observable characteristics significantly affects the likelihood of starting a career with a fixed-term contract (Table B.1 in the Appendix): Almost none of the coefficients are significantly different from zero; the exception being risk aversion (significantly negative coefficient). This result is reassuring and important as it provides further evidence against the possibility of fertility-related self-selection into fixed-term contracts at labor market entry. Summing up, we cannot claim to estimate the causal effect of fixed-term employment on fertility as we lack truly exogenous variation in temporary contracts. However, controlling for a large set of personality traits and attitudes and given the insignificance of predetermined characteristics to type of first job contract, we believe that our results are robustly estimated associations.

We run separate regressions for all outcome variables using a standard probit model to estimate the association between starting a career with a fixed-term contract and the probability of entering parenthood. Since women can only have a nonnegative integer number of children, we employ a maximum likelihood procedure with an underlying poisson distribution for the estimations of number of children. We use robust standard errors to account for potential heteroskedasticity.[20]

3.5 Main Results

3.5.1 Probability of Entering Motherhood

In this section we present the results of the regression analysis. Table 3.7 shows the main results of separate probit regressions. Each cell reports the average marginal effect of starting a career with a fixed-term contract on the probability of having a first birth during the first z years after graduation.[21] The first column reports the results from the specification including individual, background, and job characteristics. In the second column we add personality traits and attitudes. Finally, in the last column we add a control for partnership status. Column III is our preferred specification since it contains all relevant control variables.

The first finding is that the estimates are quite stable across the different specifications, suggesting that the results are not purely driven by selection based on observable characteristics, personality traits, and attitudes. We proceed further in time when going from the top of the

[20] Basically, we use the same sample of 267 women in all estimations. In practice, the number of observations varies slightly between the estimations in the main table since the maximum likelihood procedure cannot use all observations if the outcome is predicted perfectly.

[21] Strictly speaking, we estimate the correlation between starting a career with a fixed-term contract and the probability of having had a first child within z years after graduation.

table to the bottom: that is, the first row reports the average marginal effects on first birth probability four years after graduation; the last row reports the average marginal effects after 10 years. The association between the first birth probability and starting a career with a fixed-term contract is zero up to three years after graduation since in these years the vast majority of women are still childless and are working in their first job. But four years after graduation, the size of the marginal effect increases considerably and becomes significantly different from zero. For women entering the labor market with a fixed-term contract, the probability of having entered motherhood within five years after graduation is smaller by 20 percentage points than it is for those starting work with a permanent contract. This difference 10 years after graduation is still notable at 15 percentage points. Hence, after a period in which all women work and none have children, women who started their career with a fixed-term contract are significantly less likely to have become mothers compared to women on permanent first job contracts. We interpret this finding as a postponement effect due to temporary jobs.

Table 3.7 : Probability of first birth 4 to 10 years after graduation

	(I)	(II)	(III)
After 4 years	-0.102*	-0.127***	-0.147***
	(0.052)	(0.045)	(0.043)
After 5 years	-0.152***	-0.182***	-0.195***
	(0.052)	(0.044)	(0.043)
After 6 years	-0.123**	-0.148***	-0.160***
	(0.059)	(0.051)	(0.050)
After 7 years	-0.117*	-0.138**	-0.156***
	(0.064)	(0.059)	(0.055)
After 8 years	-0.147**	-0.159**	-0.172***
	(0.066)	(0.062)	(0.059)
After 9 years	-0.116*	-0.124*	-0.137**
	(0.067)	(0.064)	(0.061)
After 10 years	-0.127*	-0.139**	-0.152**
	(0.068)	(0.064)	(0.061)
First job characteristics	YES	YES	YES
Personality traits & attitudes	NO	YES	YES
Partnership status	NO	NO	YES
Observations	267	267	267

Notes: Average marginal effects from probit regressions for starting the career with a fixed-term contract; female sample, no migrants; all regressions contain controls for individual characteristics and background characteristics, and federal state of first job and year of graduation dummies; robust standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.1.

All coefficients of the other explanatory variables not reported in Table 3.7 show the expected sign (see Table B.2 in the Appendix).[22] For instance, all else equal, older graduates are more

[22] Table B.2 in the Appendix provides an example of a complete regression table on first birth probability five years after graduation.

likely, and the better educated less likely, to enter parenthood within five years. Family background seems to play no role in the decision to have a child, but personality does: open-minded and conscientious women are less likely to have a child five years after they finish their education. Furthermore, attitudes and values seem to matter. Women for whom the family is very important are significantly more likely to have entered motherhood during the five-year period, whereas women for whom a career is very important are significantly less likely to do so. Finally, having a partner at the time of labor market entry increases the probability of becoming a mother in the first five years after graduation. According to economic theories of fertility, temporary jobs could affect fertility decisions via reduced first job income. In the main regression we do not explicitly control for an individual's income since it might be endogenous. However, we include the most important predictors of average income, such as education, age, occupation, industry, personality traits, and attitudes. Hence, we implicitly control for an individual's earning potential, but omit all remaining idiosyncratic variation in earnings, which is probably highly endogenous. For completeness, we later present results controlling for net labor income (wages) at labor market entry.

3.5.2 Number of Children

Does a delay in entering motherhood result in having fewer children? The evidence presented in Table 3.8 reveals significantly negative effects of entering the labor market with a fixed-term contract on number of children up to 10 years after graduation. Compared to previous estimates, however, these results are slightly weaker and less significant for all specifications. Again, the effect does not kick in before year 4 after graduation. In the full specification (Column III), the estimated coefficient remains significantly different from zero and increases continuously. For instance, starting a career in a fixed-term job reduces fertility five years after graduation on average by almost 0.22 of a child and after 10 years by more than 0.28 of a child. This indicates that the observed postponement does indeed translate into lower fertility and cumulates over time. Since we do not observe women all the way until the end of their reproductive age, our analysis does not allow making any statements about total fertility. However, the significant reduction in the number of children 10 years after graduation points to a potential reduction in total fertility as well.

Table 3.8 : Number of children 4 to 10 years after graduation

	(I)	(II)	(III)
After 4 years	-0.082	-0.121**	-0.155**
	(-)†	(0.061)	(0.075)
After 5 years	-0.134	-0.190**	-0.215**
	(0.089)	(0.090)	(0.092)
After 6 years	-0.069	-0.125	-0.150**
	(0.091)	(-)†	(0.074)
After 7 years	-0.127	-0.176*	-0.208**
	(0.099)	(0.097)	(0.088)
After 8 years	-0.202**	-0.240**	-0.247***
	(0.103)	(0.101)	(0.095)
After 9 years	-0.187*	-0.237**	-0.251**
	(0.114)	(0.112)	(0.107)
After 10 years	-0.224*	-0.278**	-0.286**
	(0.123)	(0.119)	(0.115)
First job characteristics	YES	YES	YES
Personality traits & attitudes	NO	YES	YES
Partnership status	NO	NO	YES
Observations	267	267	267

Notes: Average marginal effects from poisson regressions for starting the career with a fixed-term contract; all regressions contain controls for individual characteristics and background characteristics, and federal state of first job and year of graduation dummies; † standard error cannot be estimated consistently; interpretation of coefficient with caution; female sample, no migrants; robust standard errors in parentheses; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

3.5.3 Robustness Checks

In this section we test the sensitivity of our main results in several ways. First, we test whether our main results, which were based on a homogenous population subsample (natives with at least secondary education), are affected when we include individuals with migratory background and without secondary education (full sample). Second, we relax the age at graduation limitation to see whether our main results are robust to including women who finish their education or training after age 30. The results for both tests are shown in Table 3.9. The first column reveals that the negative association between starting a career with a fixed-term contract and entering motherhood also holds for the full sample. The results seem particularly robust for years 5 to 8 and even for year 10. Overall, the estimated coefficients are slightly smaller than those in Table 3.7 and the significance levels for the early and late years are somewhat reduced. For instance, the average marginal effects on having had a first birth after four and five years after graduation decline by around 5 percentage points but remain statistically significant.

Table 3.9 : Sensitivity analysis: full sample and different age-at-graduation cut-offs

Dependent variable		First birth			Number of children	
	(I)	(II)	(III)	(IV)	(V)	(VI)
Sample	Full sample	Age at grad. < 35	Age at grad. < 40	Full sample	Age at grad. < 35	Age at grad. < 40
After 4 years	-0.083**	-0.125***	-0.113**	-0.072	-0.131*	-0.101
	(0.041)	(0.046)	(0.047)	(0.058)	(0.068)	(0.071)
After 5 years	-0.143***	-0.172***	-0.162***	-0.153**	-0.206***	-0.185**
	(0.043)	(0.044)	(0.045)	(0.066)	(0.066)	(0.086)
After 6 years	-0.130***	-0.138***	-0.121**	-0.114	-0.126*	-0.087
	(0.045)	(0.049)	(0.050)	(0.081)	(0.075)	(0.078)
After 7 years	-0.147***	-0.139**	-0.128**	-0.172*	-0.189**	-0.157*
	(0.049)	(0.055)	(0.055)	(0.088)	(0.088)	(0.088)
After 8 years	-0.135***	-0.160***	-0.138**	-0.166*	-0.241***	-0.198**
	(0.052)	(0.058)	(0.058)	(0.095)	(0.091)	(0.093)
After 9 years	-0.103*	-0.106*	-0.087	-0.156	-0.208**	-0.168
	(0.054)	(0.060)	(0.061)	(0.104)	(0.105)	(0.105)
After 10 years	-0.123**	-0.117*	-0.096	-0.208*	-0.244**	-0.199*
	(0.053)	(0.060)	(0.061)	(0.108)	(0.111)	(0.113)
Observations	363	287	294	363	287	294

Notes: Average marginal effects from probit (Columns I-III) and poisson (Columns IV-VI) regressions for starting the career with a fixed-term contract; female sample; all regressions contain controls for individual characteristics and background characteristics, and federal state of first job and year of graduation dummies; robust standard errors in parentheses; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

It appears that fixed-term employment does not influence the fertility decisions of migrants to the same degree as it does the decisions of non-migrants, possible due to cultural difference in fertility behavior (Fernández and Fogli, 2006). Turning our attention to the effects on the number of children (Table 3.9, Column IV), we see that the estimated marginal effects are smaller than in our main specification and are less precisely estimated; the only significant coefficient at the 5 percent level is on the number of children in year 5 after graduation. Hence, in the full sample that includes migrants and women with less than secondary education, the significant postponement effect of fertility continues to have an impact on the number of children 10 years after graduation, but is a weaker effect than found in the sample previously investigated.[23]

Our main results are also generally robust to relaxing the age at graduation restriction. The average marginal effects for the main sample including childless women graduating up to age

[23] However, even though the estimates for the later years are about half the size of our main results, they are not very close to zero. We cannot rule out the possibility that the estimates become insignificant as we lack precision due to our small sample size.

34 and 39 are reported in Table 3.9, Columns II and V respectively Columns III and VI.[24] Ten years after labor market entry these women are 44 and 49 years old and have probably completed their fertility plans. Furthermore, compared to younger women, these older graduates might have a lower fecundity and might find it increasingly difficult to realize their fertility intentions.[25] Nevertheless, for both samples having a fixed-term contract in the first job is associated with a lower probability of first birth up to the eighth year after graduation. The results for year 9 and 10 after graduation are not significantly different from zero when we include women who are relatively old (age 35+) at labor market entry. This could indicate that older graduates are not able to postpone childbearing too long as they are closer to the end of their reproductive age. Overall, the postponement effects are slightly smaller than those in our main regressions but remain qualitatively almost equal. The results on number of children are also quite robust for the sample including women graduating up to age 34, but are smaller and less often significantly different from zero when including women graduating up to age 39. This could also be related to the reduced time window during which older graduates can realize their fertility intentions.

In a third robustness test we analyze whether our results are biased by sample attrition. Recall that our main analysis is based on a (generated) balanced sample of women whom we observe for at least 10 subsequent years after they finish their education. If dropping out of the survey is correlated with starting a career with a particular type of contract, our results may be confounded. Therefore, we construct a balanced sample including all women who stay in the survey for at least five years after graduation. This sample condition is less strict and substantially increases the size of the sample. If our main results are driven by a confounding change in sample composition, the marginal effects in Table 3.10 should diverge from those in Table 3.7 and 3.8.

However, as the results in Columns I and II of Table 3.10 show, the estimated negative relationship between labor market entry with a fixed-term contract and our fertility outcomes remains very robust. In fact, the larger sample size increases the precision of the estimates. Three years after graduation, the likelihood of entering motherhood is around 6 to 7 percentage points lower among women who started with a fixed-term contract, compared to those whose first job was permanent. This difference increases to more than 14 percentage points in year 5 after graduation. We observe a similar pattern for the results on the number of children: the marginal effects remain very similar to our main results, but the standard errors become smaller. Hence, these results reveal a significant reduction in the number of children due to fixed-term employment at labor market entry.

[24] In our sample, about 4 percent and 2 percent of women who are childless at graduation finish their education when they are older than 34 and 39, respectively.

[25] For example, in our sample, none of the childless women graduating at age 39 or older give birth after entering the labor market.

Table 3.10 : Sensitivity analysis: five years balanced sample

Dependent variable	First birth		Number of children	
	(I)	(II)	(III)	(IV)
After 3 years	-0.056*	-0.066**	-0.056*	-0.072**
	(0.030)	(0.028)	(0.034)	(0.033)
After 4 years	-0.084**	-0.091***	-0.087**	-0.097**
	(0.034)	(0.033)	(0.043)	(0.043)
After 5 years	-0.140***	-0.146***	-0.145***	-0.156***
	(0.036)	(0.034)	(0.049)	(0.048)
Personality traits & attitudes	YES	YES	YES	YES
Partnership status	NO	YES	NO	YES
Observations	490	490	490	490

Notes: Average marginal effects from probit (Columns I-II) and poisson (Columns III-IV) regressions for starting the career with a fixed-term contract; female sample, no migrants; all regressions contain controls for individual characteristics and background characteristics, and federal state of first job and year of graduation dummies; robust standard errors in parentheses; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

In Table 3.11 we run three more robustness checks. First, partner's labor market status might be important in the decision to have children. Even if the own contract is only fixed-term, a permanent job arrangement for the partner could enable a couple to have children. Second, we argue that economic uncertainty is the main explanation for the negative estimates. Hence, specifying uncertainty explicitly in the regression equation should have an effect on the coefficients. As an objective measure of economic uncertainty we include local labor market conditions in the regressions (see Columns II and V). Finally, Columns III and VI directly control for the perceived level of job security at labor market entry.

Male partner controls are measured at the career start of the women and contain a dummy variable for not working and one for fixed-term employment. The unemployment rate is measured at the level of "Raumordnungsregionen", an aggregation level between federal states and counties. Assuming that the initial type of contract has no influence on future unemployment rates, we include the unemployment rate lagged by one year (which means in our notation $t = z - 1$). Finally, perceived job security takes the value 1 if the individual is very concerned about job security; 0 otherwise. Note that the number of observations drops slightly since SOEP does not provide regional identifiers for the first year (1995) of our sample. Similarly, sample size is reduced in Columns I and IV since partner information is available only for cohabiting couples and job uncertainty is reported solely by working individuals (Columns III and VI).

Without any formal test the estimated coefficients in Table 3.11 seem similar in sign and size to the estimates in Table 3.7 and 3.8. On average, the associations between starting a career with a fixed-term contract and the fertility outcomes are even stronger when controlling for partner's labor market status when entering the labor market and including the lagged

Table 3.11 : Sensitivity analysis: additional measures of economic uncertainty

Dependent variable		First birth			Number of children	
	(I)	(II)	(III)	(IV)	(V)	(VI)
Additional Controls	Partner's LM status	Lagged unemploy-ment rate	Perceived job uncertainty at LME	Partner's LM status	Lagged unemploy-ment rate	Perceived job uncertainty at LME
After 5 years	-0.215***	-0.241***	-0.214***	-0.200**	-0.260***	-0.227***
	(0.067)	(0.045)	(0.060)	(0.098)	(0.081)	(0.077)
After 6 years	-0.154**	-0.163***	-0.158***	-0.143	-0.136	-0.161*
	(0.063)	(0.056)	(0.060)	(0.105)	(0.083)	(0.088)
After 7 years	-0.194***	-0.204***	-0.139**	-0.237**	-0.283***	-0.177*
	(0.059)	(0.057)	(0.065)	(0.101)	(0.097)	(0.102)
After 8 years	-0.225***	-0.213***	-0.153**	-0.209*	-0.317***	-0.242**
	(0.064)	(0.061)	(0.065)	(0.112)	(0.106)	(0.108)
After 9 years	-0.196***	-0.159**	-0.119*	-0.272**	-0.279**	-0.238*
	(0.069)	(0.063)	(0.068)	(0.124)	(0.111)	(0.123)
After 10 years	-0.228***	-0.186***	-0.146**	-0.317**	-0.324***	-0.305**
	(0.064)	(0.063)	(0.067)	(0.125)	(0.125)	(0.133)
Observations	181	237	248	181	237	248

Notes: Average marginal effects from probit (Columns I-III) and poisson (Columns IV-VI) regressions for starting the career with a fixed-term contract; female sample; all regressions contain controls for individual characteristics and background characteristics, and federal state of first job and year of graduation dummies; robust standard errors in parentheses; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

regional unemployment rate. Possibly, uncertainty lessens when the partner has a permanent contract, or that opportunity costs shrink when labor market conditions are unfavourable. Even more similar to the main specifications are the estimates in Columns III and VI, implying that the uncertainty is not directly due to the type of first contract. Otherwise, including job uncertainty should have alleviated the correlation significantly. We argue that the mechanism depends more on the uncertainty due to the path dependence of starting a career with a fixed-term contract. Repeated spells in temporary employment and the associated economic uncertainty induces women to postpone childbearing.

3.5.4 Mechanism

We now test whether the negative association of fixed-term employment at career start and fertility is driven by the lower wage of temporary jobs and a lower attractiveness on the marriage market. From descriptive evidence we know that starting a career with a fixed-term contract is associated with lower wages, which in turn may cause reduced fertility. To this point, we have controlled for many wage predictors. Now, we re-estimate our main specification controlling explicitly for monthly net wages of the first job (Columns I and IV of Table 3.12).[26] If wages and their profile over time are the main channel through which fixed-term jobs affect fertility, the coefficient of fixed-term employment should become much smaller and even insignificant.

Table 3.12 : Mechanism: income and marriage market effects

Dependent variable	First birth			Number of children		
	(I)	(II)	(III)	(IV)	(V)	(VI)
Additional Controls	Wage of first job	Married at labor market entry	Cohabiting at labor market entry	Wage of first job	Married at labor market entry	Cohabiting at labor market entry
After 4 years	-0.116**	-0.120***	-0.137***	-0.119	–	–
	(0.047)	(0.045)	(0.044)	(0.144)	–	–
After 5 years	-0.168***	-0.178***	-0.183***	-0.178**	-0.183**	-0.205**
	(0.047)	(0.045)	(0.043)	(0.071)	(0.086)	(0.104)
After 6 years	-0.130**	-0.142***	-0.142***	-0.091	-0.111	-0.124
	(0.054)	(0.051)	(0.051)	(0.087)	(0.085)	(0.158)
After 7 years	-0.143**	-0.137**	-0.141**	-0.141	-0.160*	-0.166*
	(0.058)	(0.058)	(0.058)	(0.095)	(0.096)	(0.090)
After 8 years	-0.147**	-0.136**	-0.136**	-0.187*	-0.191*	-0.175*
	(0.062)	(0.062)	(0.062)	(0.102)	(0.102)	(0.096)
After 9 years	-0.122*	-0.110*	-0.115*	-0.212*	-0.213*	-0.208*
	(0.063)	(0.064)	(0.062)	(0.112)	(0.111)	(0.109)
After 10 years	-0.129**	-0.122*	-0.124**	-0.245**	-0.261**	-0.243**
	(0.063)	(0.064)	(0.062)	(0.119)	(0.117)	(0.116)
Observations	267	267	267	267	267	267

Notes: Average marginal effects from probit (Columns I-III) and poisson (Columns IV-VI) regressions for starting the career with a fixed-term contract; female sample, no migrants; all regressions contain controls for individual characteristics and background characteristics, and federal state of first job and year of graduation dummies; robust standard errors in parentheses; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

However, the results in Table 3.12 are very similar to our main estimates in Table 3.7 and 3.8, thus indicating that it is not the lower income of the first job per se that induces women to postpone childbearing, but more likely the economic uncertainty associated with temporary contracts.

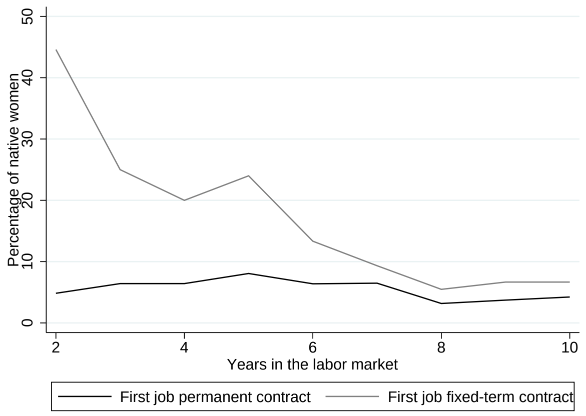
Another potential channel is lower attractiveness on the marriage or partner market. Women in fixed-term employment are supposed to be less attractive to men and therefore less likely to have a suitable partner with whom to form a family. In the standard specifications (last columns of Table 3.7 and 3.8), we control for partnership status at labor market entry to account for this channel. Replacing having a partner by being married or living together at labor market entry appears to have no substantial effect on the correlation between starting

[26] Unemployed individuals are assigned a zero wage.

a career with a fixed-term contract and fertility outcomes. Columns II, III, V, and VI show that the coefficients are only slightly different from the ones estimated previously. In particular, Column II of Table 3.7 is very similar to the results here, meaning that cohabitation or marriage does not alleviate the uncertainty caused by temporary employment. If anything, having a partner seems to matter for the way economic uncertainty influences fertility decisions.

We propose that neither lower income nor lower attractiveness on the partner market are the driving mechanisms for the negative association but that, instead, it is the economic uncertainty attached to the path dependence of starting a career with a fixed-term contract. Figure 3.1 shows the stability of an initial fixed-term contract over the first 10 years in the labor market. In the second year in the labor market, almost 50 percent of women, who started in a temporary job, have a contract of limited duration, but this number declines steadily over time. Nevertheless, in the first five years, repeated spells in fixed-term employment seem to be common. Women who start their careers with a regular contract have for the whole period a much lower probability of working with a fixed-term contract.

Figure 3.1 : Stability of initial employment by contract type



Notes: The graph is based upon the main sample of native women. To correct for feedback effects of childbirth, we restrict the sample to women without children.

Table 3.13 reports marginal effects of the type of initial contract on the type of contract in year z after graduation. The inclusion of first job characteristics in Column II has no substantial effect on the size of the estimates. The development indicates a strong persistence in temporary employment for at least the first five years in the labor market. Controlling for first job characteristics, we estimate a 20 percentage point higher probability of holding a fixed-term contract even five years after the first job.

We argue that this path dependence is the main force that creates economic uncertainty for the affected women and induces them to postpone entering motherhood and, as a result, to have on average fewer children than women with a permanent first job. Since the large

Table 3.13 : Path dependence of starting the career with a fixed-term contract

Dependent variable	Future fixed-term contract	
	(I)	(II)
After 2 years	0.464***	0.448***
	(0.036)	(0.037)
After 3 years	0.329***	0.299***
	(0.068)	(0.071)
After 4 years	0.196***	0.221***
	(0.058)	(0.056)
After 5 years	0.184***	0.242***
	(0.064)	(0.069)
First job characteristics	NO	YES

Notes: Average marginal effects from probit regressions for starting the career with a fixed-term contract; female sample, no migrants; all regressions contain controls for individual characteristics and background characteristics, and federal state of first job and year of graduation dummies; robust standard errors in parentheses; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

majority of women have found a permanent job after five years, estimations are not possible beyond this point.

3.5.5 Further Heterogeneity Analysis and Results for Men

Results by Education

In this section we investigate whether starting a career with a fixed-term contract affects fertility outcomes of women differently depending on their level of education. Specifically, we compare women with secondary education to those with tertiary education. Therefore, we split our main sample into two groups: (a) all women who attained middle vocational training or vocational training and "Abitur" (ISCED codes 3 and 4) and (b) all women with higher vocational training or a university degree (ISCED codes 5 and 6).[27] The main rationale for doing this is that women with high educational attainment, such as a university degree, enter the labor market relatively late but face the same "biological age restrictions" as women who finish their education at younger ages. Hence, the scope for postponing having a child is much more restricted for more highly educated women. Furthermore, for older women it becomes comparatively more risky, both for health and biological reasons, to postpone childbearing. Thus, conditional on a particular intended number of children, we would expect a smaller postponement effect of fixed-term employment for women with higher education.

Table 3.14 reports the average marginal effects by educational subgroup for selected years. As hypothesized, the postponement effects are stronger for women with secondary education:

[27] In the residual group, that is also not part of the main sample, are women who dropped out of school or do not have any vocational training at all.

the coefficients are more negative and the estimates are highly significant. For instance, five years after graduation, the first birth probability is reduced by almost 24 percentage points if the first contract was of limited duration. The magnitude of the effect declines over time but remains statistically significant, even 10 years after graduation. In contrast, the point estimates for women with tertiary education are smaller and only weakly significant. Regarding the number of children, we find a similar picture: starting the career with a fixed-term contract significantly reduces the realized number of children for women with secondary education. For women with a university degree, the economic uncertainty associated with starting a career with a fixed-term contract does not seem to play such a crucial role in the timing of the first child or for the number of children in the first 10 years after their graduation. The estimates are sizeable, but much smaller, and we are not able to estimate the coefficients precisely enough to distinguish them from a zero effect. Even though a formal test is not easily applicable in this setting, the confidence intervals suggest that the coefficients are statistically different between the two groups, at least in the earlier years after labor market entry. These findings are in line with our expectations that relatively old university graduates are not able to postpone parenthood to the same extent as can younger women.

Table 3.14 : Heterogeneity analysis: effects by education

Dependent variable	First birth		Number of children	
	(I)	(II)	(III)	(IV)
Sample	Secondary education	Tertiary education	Secondary education	Tertiary education
After 5 years	-0.236***	-0.095*	-0.277*	-0.147
	(0.050)	(0.053)	(0.148)	(0.152)
After 7 years	-0.335***	-0.014	-0.308***	-0.110
	(0.091)	(0.031)	(0.107)	(0.203)
After 10 years	-0.279**	-0.112	-0.348**	-0.092
	(0.112)	(0.115)	(0.145)	(0.168)
Observations	204	141	204	141

Notes: Average marginal effects from probit (Columns I-II) and poisson (Columns III-IV) regressions for starting the career with a fixed-term contract; note that due to sample size problems, the ML method cannot find a maximum in the standard specification; therefore, partnership status is removed from the specification; female sample; all regressions contain controls for individual characteristics and background characteristics, and federal state of first job and year of graduation dummies; robust standard errors in parentheses; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Results for Men

It is possible that entering the labor market with a fixed-term contract affects the fertility outcomes of men, too. The corresponding results are reported in Table 3.15. In the male sample, the age at graduation cut-off is two years later than in the female sample since men, on average, graduate two years later and are not exposed to biological constraints regarding fertility. The male sample consists of 225 observations. The estimated association between temporary jobs and subsequent fertility is close to zero and never statistically distinguishable from zero. Hence, the results indicate that men do not react as sensitively as women to economic uncertainty. One possible explanation is that men do not suffer from fixed-term contracts in the long run since they can quite easily find a permanent job even if they already have a child. In contrast, women, who are responsible for the majority of childrearing, first want to gain a foothold in the labor market before deciding to enter motherhood.

Table 3.15 : Heterogeneity analysis: effects on men

Dependent variable	First birth		Number of children	
	(I)	(II)	(III)	(IV)
After 5 years	0.077	0.070	0.178	0.116
	(0.070)	(0.064)	(0.155)	(0.122)
After 7 years	-0.012	-0.038	0.158	0.103
	(0.079)	(0.072)	(0.165)	(0.140)
After 10 years	-0.023	-0.052	0.078	-0.003
	(0.080)	(0.068)	(0.171)	(0.136)
Personality traits & attitudes	YES	YES	YES	YES
Partnership status	NO	YES	NO	YES
Observations	225	225	225	225

Notes: Average marginal effects from probit (Columns I-II) and poisson (Columns III-IV) regressions for starting the career with a fixed-term contract; male sample, no migrants; all regressions contain controls for individual characteristics and background characteristics, and federal state of first job and year of graduation dummies; robust standard errors in parentheses; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

3.6 Conclusion

In countries with strong employment protection legislation, fixed-term contracts are intended to increase the flexibility of the recruitment process and are thought to foster employment in general. However, existing evidence suggests that fixed-term contracts lead to a dual labor market and casts doubt on the notion that temporary contracts foster employment and reduce aggregate unemployment in the long run (Cahuc and Postel-Vinay, 2002; Bentolila and Dolado, 1994; Boeri and Garibaldi, 2007).

Our analysis sheds light on potential spill-over effects of fixed-term employment on fertility. Using German data on young female graduates, we find a significant postponement of first birth and a reduction in the number of children in the first ten years after graduation. These results are robust to several sensitivity tests. A possible explanation is the economic uncertainty associated with a career start in fixed-term employment. The likelihood of repeated spells in precarious jobs is significantly higher when entering the labor market on a temporary contract. Thus, we measure an indirect effect of the initial contract on fertility, which works via recurrent contemporaneous economic instabilities. Furthermore, we show that fixed-term employment appears to particularly affect the fertility decisions of women with secondary education. In contrast, our results reveal no significant correlations between job uncertainty at the beginning of a career and fertility for young men. We address potential endogeneity threats by including a large set of controls and by showing evidence against fertility-related self-selection into temporary contracts at the beginning of a career. Hence, we believe that the results suggest a negative relationship between fixed-term employment and fertility that is robust to a variety of sensitivity checks.

Our study has important implications for policymakers in low-fertility countries as our findings highlight negative spill-over effects of temporary employment on demographic outcomes. Fixed-term contracts might facilitate the labor market entry of older persons and the long-term unemployed (stepping-stone hypothesis), but they seem to impede the integration of young graduates into the labor market and to negatively affect fertility outcomes. As such, this labor market policy imposes a disproportionate burden on the young generation. Against this background, policymakers should possibly reconsider the costs and benefits of this policy instrument and strive for a more equal distribution of the costs associated with employment protection across population subgroups. A possible approach is broader reform of the employment protection legislation, that is, a reduction in dismissal costs for all workers (Blanchard and Landier, 2002).

Appendix B.1 Supplementary Tables

Table B.1 : Probability of starting a career with a fixed-term contract

Sample	(I) Native women	(II) All women
Age at graduation	0.023	0.018
	(0.015)	(0.012)
Years of education	0.012	0.002
	(0.016)	(0.014)
Born in East Germany	0.082	0.029
	(0.111)	(0.106)
High education mother	-0.107	0.033
	(0.085)	(0.093)
Employment mother	-0.119	-0.135
	(0.118)	(0.092)
Age at birth mother	0.006	0.001
	(0.007)	(0.005)
Number of siblings	0.067	0.034
	(0.080)	(0.077)
Openness	0.016	0.030
	(0.032)	(0.026)
Agreeableness	-0.042	-0.032
	(0.036)	(0.030)
Conscientiousness	-0.013	-0.029
	(0.044)	(0.033)
Extraversion	-0.029	-0.015
	(0.030)	(0.025)
Neuroticism	0.006	0.007
	(0.041)	(0.035)
Risk aversion	-0.173**	-0.165**
	(0.084)	(0.072)
Locus of control	0.000	-0.035
	(0.054)	(0.045)
Importance of having children	0.090	0.022
	(0.077)	(0.068)
Importance of partnership	0.121	0.156
	(0.146)	(0.100)
Importance of career	0.002	0.008
	(0.110)	(0.081)
Importance of affording something	0.003	-0.032
	(0.099)	(0.092)
In partnership after graduation	0.078	0.148*
	(0.123)	(0.090)
Observations	267	363

Notes: Average marginal effects from probit regressions; female working sample, all regressions contain federal-state-of-first-job and year-of-graduation dummies; robust standard errors in parentheses; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table B.2 : Probability of first birth 5 years after graduation

	(I)	(II)	(III)
First job fixed-term contract	-0.152***	-0.182***	-0.195***
	(0.052)	(0.044)	(0.043)
Unemployment after graduation	0.649***	0.645***	0.623***
	(0.027)	(0.028)	(0.045)
Age at graduation	0.034***	0.038***	0.037***
	(0.012)	(0.011)	(0.011)
Years of education	-0.036***	-0.041***	-0.040***
	(0.013)	(0.012)	(0.012)
Born in East Germany	0.071	0.083	0.065
	(0.100)	(0.093)	(0.094)
High education mother	-0.007	-0.002	-0.026
	(0.092)	(0.082)	(0.078)
Employment mother	0.112	0.064	0.069
	(0.127)	(0.119)	(0.115)
Age at birth (mother)	-0.004	-0.007	-0.007
	(0.006)	(0.005)	(0.005)
Number of siblings	0.035	0.066	0.064
	(0.077)	(0.075)	(0.072)
Openness		-0.018	-0.007
		(0.022)	(0.021)
Agreeableness		-0.017	-0.031
		(0.029)	(0.028)
Conscientiousness		-0.039	-0.035
		(0.032)	(0.031)
Extraversion		0.027	0.014
		(0.024)	(0.023)
Neuroticism		-0.019	-0.033
		(0.031)	(0.030)
Risk aversion		0.117	0.108
		(0.087)	(0.081)
Locus of control		-0.031	-0.025
		(0.040)	(0.038)
Importance of having children		0.215***	0.225***
		(0.060)	(0.060)
Importance of partnership		0.143	0.053
		(0.166)	(0.171)
Importance of career		-0.157*	-0.151*
		(0.083)	(0.080)
Importance of affording something		-0.019	0.001
		(0.086)	(0.083)
First job blue collar	0.121	0.116	0.125
	(0.104)	(0.098)	(0.096)
In partnership after graduation			0.190***
			(0.051)
Observations	267	267	267

Notes: Average marginal effects from probit regressions; female sample, no migrants; all regressions contain federal-state-of-first-job and year-of-graduation dummies; robust standard errors in parentheses; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

4 Health Consequences of Starting a Career with a Fixed-Term Contract

4.1 Introduction

Official health insurance statistics provide evidence that mental health issues are a major concern in Germany. The 2014 report of the company health insurance fund (BKK) contains statistics suggesting that absenteeism due to mental illness has increased rapidly – since the 1970s, absence days per insured person have quintupled and currently account for more than 15 percent of all absenteeism. In addition, mental illness causes the longest periods of absence with, on average, 38 days (Knieps and Pfaff, 2014).

Previous studies mainly aim at finding the contemporaneous link between health conditions and fixed-term employment at different stages of the lifecycle and have produced mixed evidence. Their approach neglects the potential endogeneity of fixed-term contracts due to path dependence of starting the career in a temporary job. The empirical literature on whether and, if so, how increased levels of economic uncertainty due to unstable working contracts at the beginning of the career have spill-over effects on other domains of life is rare. The main objective and contribution of this chapter is to fill this gap by empirically assessing the implications for subsequent physical and mental health outcomes of entering the labor market with a fixed-term contract.

From a theoretical perspective, the effect of fixed-term employment on health is not clear a priori. First, lower opportunity costs might allow individuals to devote more time to healthy behavior and therefore induce better health outcomes (Grossman, 1972). On the other hand, higher levels of stress, uncertainty, and financial instability might negatively affect mental health and health investments. Empirically, we focus on several cohorts of graduates from vocational training or tertiary education and follow them up to five years after entering the labor market. We analyze the effects of starting a career with a fixed-term contract on subsequent health outcomes in the short run. Another contribution of this chapter is to carefully discuss and address the potential endogeneity due to path dependence of starting a career with a fixed-term contract. To reduce possible omitted variable bias, we exploit a rich and unique data set and include a large set of new control variables (e.g., personality traits, attitudes, family background, and ex-ante health status). In addition, we check to what extent men and women select into fixed-term employment based on observable characteristics and employ an approach to assess the resulting bias.

Based on the survey years 1995 to 2012 of the German Socio-Economic Panel (SOEP), the results suggest strong gender difference in response to economic uncertainty at the beginning of the career. For women, starting a career with a fixed-term contract is associated with negative subsequent health outcomes in the short run. However, men respond with a higher status of mental health as of their third year on the labor market. In contrast, neither men's nor women's physical health outcomes are affected by the type of the contract held at the beginning of the career. These findings are robust to a variety of sensitivity checks and can be explained, at least for women, by a negative path dependence of starting a career with a fixed-term contract. As health status, as well as holding a fixed-term contract, may be driven by unobserved heterogeneity, we address potential endogeneity concerns on two fronts: first, by including pre-graduation health status as well as a set of previously neglected control variables and, second, by showing that entering the labor market with a fixed-term contract is related neither to family nor to career preferences.

In general, this research contributes to the growing literature on the relationship between economic uncertainty and health. Several empirical studies focus on aggregate unemployment and how it affects mortality rates (see, e.g., Ruhm, 2000, 2003, 2005). Mortality rates seem to follow a pro-cyclical pattern at the state level, which is surprising since it suggests an adverse health effect of reduced unemployment rates. On the individual level, the relationship seems to be reversed. For instance, Sullivan and von Wachter (2009) show a particularly pronounced increase in the annual probability of dying immediately after job loss. Theodossiou and Vasileiou (2007) find that the effect of perceived risk of job loss on job satisfaction is significantly negative and large.

Temporary employment is considered to be one particular form of economic uncertainty. Unfortunately, evidence on the relationship between fixed-term employment and health is scarce and inconclusive. Several studies find negative effects of fixed-term employment on job satisfaction (e.g., Booth et al., 2002; Chadi and Hetschko, 2013; Dawson, Veliziotis, Pacheco, and Webber, 2015), but show that this does not translate into lower well-being (Dawson and Veliziotis, 2013) or worse health status (Bardasi and Francesconi, 2004). In the study most similar to our research, Rodriguez (2002) analyzes British and German micro data and finds that German workers with fixed-term contracts have a significantly higher probability of reporting worse health than their permanently employed colleagues. In contrast, the effect of fixed-term employment is not significant among British workers, a finding in line with the results of Bardasi and Francesconi (2004) for the United Kingdom. All these studies have in common that they focus mostly on empirical associations between holding a fixed-term contract and health outcomes. They neither consider the potential endogeneity problems nor the potential path dependence of entering the labor market with a contract with limited duration.

The remainder of this chapter is organized as follows. Section 4.2 introduces the data and Section 4.3 the empirical approach. The main results and mechanisms as well as several sensitivity and subgroup analyses are presented in Section 4.4. Section 4.5 concludes.

4.2 Data and Sample Restrictions

We employ the German Socio-Economic Panel (SOEP), which has provided annual and nationally representative panel data since 1984 (Wagner et al., 2007). In 2012, the SOEP covered more than 20,000 individuals living in over 12,000 households. SOEP contains detailed information on a variety of individual as well as household-specific socioeconomic characteristics. Moreover, the respondents provide information about their labor market history as well as their current labor force status. Most importantly, we observe when the individuals finish their education and enter the labor market and whether their first employment contract is permanent or temporary. We focus on the waves 1995 to 2012 since consistent information on the type of employment contract for all working individuals was collected only from 1995 onward. Respondents who do not report a job change are excluded from this question before 1995. Thus, switching from temporary to permanent employment at the same employer is not part of the questionnaire up to 1995. The questions regarding life and health satisfaction are available for the whole observation period. Since the health questions have been part of the questionnaire since 2002, but only every second year, we are not able to look at the same sample period as for the satisfaction analysis. To make the sample more homogenous, we restrict the sample to women and men who entered the labor market with their highest degree before their 36th birthday and have at least a secondary school degree. In the robustness checks we show that the results are not sensitive to the choice of the age-at-graduation cut-off. By restricting the sample to men and women who answer the questionnaire in five subsequent years after their labor market entry, we obtain a balanced sample. The sample for the health analysis comprises 1442 observations of 297 women and 1180 observations of 245 men from the graduation cohorts 2002 to 2007.

To measure the health status of the individuals in the sample, we use various outcome variables. First, the SOEP data provide an overall index of mental and physical health that are standardized for the survey year 2004 to a mean of 50 and a standard deviation of 10.[1] Since both summary scales are not available annually, we fill up the gaps in the main analysis: assuming a linear development of health status over time, we use the average value of the health measures between two observations. On the one hand, we obtain an increased number of observations and thus more precise estimates. On the other hand, the linearity assumption might be too restrictive and in some cases inappropriate and misleading. For instance, in years when we do not observe the health status of a respondent the linearity assumption causes a change by construction. In this case we are not able to ensure that the type of the first contract and not the computation is the main reason for the change in the health status. Since this would cause an overestimation of the true relationship we address this issue in the sensitivity analyses. Second, the data also contain information on each of the subcategories of the health indices, which we use in a refinement of the analysis. Third, we use self-rated information about satisfaction with life and health status. Although these variables are measured on a

[1] For more detail on how the cardinal measures for mental and physical health are constructed from the survey items, see Nübling, Andersen, Mühlbacher, Schupp, and Wagner (2007).

scale from 0 (worst outcome) to 10 (best outcome), we assume continuity and do not further manipulate the scales. In the analysis of life and health satisfaction we are able to look at the graduation cohorts 1995 to 2007. Thus, the number of women in the sample increases to 672 and the number of men to 607.

Table 4.1 : Development of outcome variables by employment status

	First job permanent contract			First job fixed-term contract			Unemployed after graduation		
	Mean	SD	N	Mean	SD	N	Mean	SD	N
A. Women									
Mental health 1st year	48.62	8.71	148	47.52	8.46	85	49.06	8.75	64
Mental health after 2 years	49.25	8.34	152	47.45	7.54	84	49.59	8.55	59
Mental health after 3 years	49.88	7.76	143	47.28	7.46	82	49.29	9.61	59
Mental health after 4 years	49.58	8.10	144	47.66	8.33	81	49.24	9.20	61
Mental health after 5 years	48.54	8.93	142	46.68	9.09	81	49.48	9.20	57
Physical health 1st year	56.44	5.67	148	55.54	5.53	85	55.53	5.78	64
Physical health after 2 years	55.94	6.00	152	56.01	5.76	84	55.59	5.90	59
Physical health after 3 years	55.76	6.65	143	55.93	6.47	82	55.34	6.82	59
Physical health after 4 years	55.21	7.32	144	55.53	5.66	81	54.94	6.82	61
Physical health after 5 years	55.09	7.25	142	55.26	5.36	81	54.55	6.04	57
B. Men									
Mental health 1st year	51.45	7.89	150	50.18	7.76	48	50.85	7.47	47
Mental health after 2 years	51.63	7.70	146	51.00	8.39	53	50.41	8.03	41
Mental health after 3 years	50.49	7.75	139	52.41	7.09	51	48.86	10.01	42
Mental health after 4 years	50.11	8.20	142	51.72	6.78	52	48.97	8.90	39
Mental health after 5 years	50.02	8.43	145	51.08	6.65	46	48.75	10.54	39
Physical health 1st year	56.26	5.46	150	57.25	4.97	48	56.21	5.42	47
Physical health after 2 years	55.86	6.01	146	56.51	4.73	53	55.82	5.96	41
Physical health after 3 years	56.27	4.88	139	56.21	4.61	51	56.18	6.44	42
Physical health after 4 years	56.32	4.83	142	56.44	4.78	52	55.56	7.13	39
Physical health after 5 years	55.98	5.44	145	56.01	5.10	46	54.33	6.90	39

Notes: Summary statistics of the main sample including women/men with at least secondary education, younger than 36 years at labor market entry and at least for 5 years in the sample.

Table 4.1 contains summary statistics of the two main health measures by type of first contract and employment status after graduation, respectively. Three facts are worth noting: First, fixed-term contracts are more prevalent among young women than among young men: 29 percent of the women in the sample start their career with a fixed-term contract, whereas only 20 percent of men are affected. Second, if men start their career with a fixed-term contract, their mental health appears to be slightly higher – at least after three years. However, if women's first contract is of limited duration, they tend to report a lower mental health status. The differences are statistically significant only for women three years after their labor market entry. Third, physical health conditions in the male and the female sample seem to be very similar across employment status.

The main explanatory variable is a binary variable indicating whether the first job after graduation has a fixed-term or a permanent employment contract. Since we also control for whether the respondent is unemployed after graduation, the reference group is entering the labor market with a permanent contract. A great advantage of the SOEP data is the variety of unique information about the respondents. The data allow us to include a large set of controls for individual, background, personality, and job characteristics, as well as health-related information.

Individual control variables are age, years of education, a dummy variable for migratory background and being born in East Germany as well as the partnership status. As a proxy for the respondent's predetermined family background we include variables indicating whether his or her mother has tertiary education, whether the mother was employed when she was 15 years old, whether the respondent has siblings and his or her mother's age at the respondent's birth. Personality traits and self-reported attitudes are captured by the "Big 5" and Kluckhohn's importance of life areas. More precisely, five variables reflect the respondent's *openness to new experience*, *agreeableness*, *conscientiousness*, *extraversion*, and *neuroticism*. Four binary variables indicate individual career and family-related attitudes. They take the value 1 if a woman claims that having children, being in a happy relationship, the ability to afford something or having a career is important or very important in her life. Questions on personality traits and attitudes are not included in every wave of the survey. However, in order to exclude possible feedback effects of personal labor market or partnership experience on personality traits and family and career attitudes, we only use the first available observation. The majority of women answer this question around the age of 21 or younger. In addition, a dummy variable indicates whether an individual is risk averse, that is, reports a (very) low subjective willingness to take risks. Our last measure of personality is locus of control, where higher values imply a higher level of internal locus of control.

As regards the characteristics of the job, we include only very rough indicators, namely, dummy variables for blue- versus white-collar occupations and for part-time and self-employment. Moreover, the weekly working hours are supposed to capture the labor market attachment. Six industry dummies are generated according to the classification of the Federal Statistical Office (destatis). These are agriculture, manufacturing, construction, trade and transportation, financial services, and public and other services. The prevalence of fixed-term employment differs across industries and occupational groups. Similarly, women and men with particular health conditions might self-select into particular industries and occupational groups. By controlling for these job characteristics we want to make sure that the results do not reflect spurious correlations between temporary jobs and the health status. To allow for different health conditions before entering the labor market, we add indicators for baseline health conditions (pre-graduation), such as the number of doctor visits and days in the hospital before labor market entry. Tables C.1 and C.2 in the Appendix presents summary statistics by gender for the majority of control variables.

4.3 Empirical Strategy

We examine the health effects of starting a career with a fixed-term contract more thoroughly in a regression framework. The basic empirical strategy is to compare women and men entering the labor market on fixed-term contracts with their counterparts on permanent contracts in terms of short- to medium-run health outcomes. The empirical setup is comparable to that used in papers studying future effects of adverse labor market conditions at the beginning of the career (see, e.g., Kahn, 2010; Liu et al., 2012; Stevens, 2008). The underlying empirical model can be described in a very simple linear regression form as follows:

$$Y_{i,t} = \alpha Y_{i,t=0} + \beta_t FT_{i,t=1} + \gamma UE_{i,t=1} + \delta' X_{i,t} + \varphi_s + \mu_t + \varepsilon_{i,t} \quad (4.1)$$

$Y_{i,t}$ denotes the health outcome of individual i in period t. We add $Y_{i,t=0}$, that is, the health status in the year before labor market entry. In so doing, we intend to equalize the baseline health status and control for ex-ante health conditions. Moreover, changes in health due to a fixed-term contract become more comparable across the individuals in the sample. Since the original data provide bi-yearly information only, we use health status either one or two years before labor market entry. This allows us to avoid biased estimates due to the linearity assumption. $FT_{i,t=1}$ is an indicator variable for starting a career with a fixed-term contract and $UE_{i,t=1}$ indicates whether an individual experienced an unemployment spell after graduation. Therefore, the reference category in the regressions is starting a career with a permanent contract. β_t measures the relationship we are interested in: How does starting a career with a fixed-term contract affect the health status in each of the first t years on the labor market. Furthermore, $X_{i,t}$ are observed individual and job characteristics, φ_s represents dummies for the federal state of residence, μ_t are year fixed effects, and $\varepsilon_{i,t}$ is the unobserved error term. We run OLS regressions with standard errors robust to any form of heteroskedasticity and clustered at the individual level.

In the regressions, it is crucial to include variables that influence the probability of holding a fixed-term contract and might simultaneously correlate with ex-ante health conditions. Not controlling for these variables, may leave them in the error term as confounding factors, which may cause spurious correlations between health and holding a fixed-term contract at labor market entry. If workers with particular characteristics self-select into particular types of contracts, the estimates would be biased. This aspect has been mainly ignored in previous studies analyzing the role of fixed-term employment on health outcomes. We discuss this issue in detail later on. Controlling for a large set of covariates, we believe that the estimates express robust associations. However, in Section 4.4.5 we assess the potential bias due to selection into fixed-term employment.

4.4 Results

4.4.1 Main Results for Health Outcomes by Gender

The estimated gender-specific coefficients of starting a career with a fixed-term contract are reported in Table 4.2. For the main indices of mental and physical health, the first row presents the contemporaneous relationship followed by the short- to medium-run effects up to the fifth year on the labor market. Similar to the summary statistics, the regression results confirm that physical health is not affected by the type of the first contract. However, women's mental health appears to respond to spells of temporary employment at the beginning of the career. While in the first job or during the first years of their careers, economic uncertainty reduces the mental well-being of women. Starting a career with a fixed-term contract is associated with a decline in women's mental health status by about 1.7 to 2.2 index points in the second and third year after labor market entry. Compared to the average pre-graduation health of 47.5 index points and a standard deviation of 10.4, female mental health is reduced by more than 4 percent or 20 percent of a standard deviation. However, as soon as four years after labor market entry, the association becomes smaller and insignificant. According to these findings, the economic uncertainty associated with fixed-term contracts is detrimental for women's mental condition. For the male sample, we find a different pattern. Right at the start of the career, the coefficient is basically zero and remains small in the second year. Thus, men do not respond to uncertainties on the labor market in the short run. Probably, men see the fixed-term contract as an extended probation period and do not worry much about the security of the job. However, in the third year men's mental well-being even increases if they started a career with a fixed-term contract. Thus, the way men and women handle economic uncertainty at an early career stage is very different. The following analyses aim at finding a proper explanation for these results.

The way SOEP provides health information allows us to have a closer look at the subcategories of the mental health index to analyze in particular what causes the reduction or the rise in the mental health index. The four categories are: 1. *Mental health* that measures whether respondents have experienced time pressure or have felt depressed in the last four weeks. 2. *Vitality* takes on high values if respondents have felt calm and peaceful.[2] 3. *Social functioning* is high if respondents have not experienced limitations of social contacts due to mental health problems. 4. *Role emotional* is a variable that attains a high value if the respondent does not feel limited in work or other activities due to mental and emotional problems. Table 4.3 reports the coefficients from regressions of the subcategories of mental health on starting a career with a fixed-term contract for the female subsample. As regards *Mental health* and *Vitality*, a similar pattern as for the overall index emerges. The subcategory *Mental health* does not show a significant correlation with the mental health status except for the third year on the labor market. The pattern for the measure *Vitality* is more conclusive: all coefficients suggest

[2] To simplify the interpretation of the regression coefficients we invert the category *vitality* such that high values imply good health conditions.

Table 4.2 : Starting a career with a fixed-term contract and subsequent mental and physical health

Dependent variable	Mental health		Physical health	
	(I)	(II)	(III)	(IV)
	Women	Men	Women	Men
At labor market entry (1st year)	-1.071	-0.257	-0.146	0.704
	(1.027)	(0.980)	(0.728)	(0.809)
2nd year on the labor market	-1.716*	0.386	0.615	0.602
	(0.950)	(1.151)	(0.810)	(0.820)
3rd year on the labor market	-2.185**	3.106***	0.714	-0.136
	(0.968)	(1.015)	(0.943)	(0.780)
4th year on the labor market	-1.144	2.769**	0.735	0.388
	(1.073)	(1.068)	(0.901)	(0.803)
5th year on the labor market	-1.191	2.757**	0.559	0.496
	(1.126)	(1.161)	(0.849)	(0.895)
Observations	1,442	1,180	1,442	1,180

Notes: Marginal effects of starting a career with a fixed-term contract from OLS regressions with robust standard errors clustered at the individual level in parentheses; dependent variables interpolated assuming a linear trend; all regressions contain controls for individual characteristics, background characteristics, job characteristics, personality traits and attitudes, partnership status, ex-ante health status, federal-state and year dummies; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

a significant association that is strongest in the third year on the labor market. Women appear to suffer from economic uncertainty by reporting less time for rest and recreation. In contrast, women do not feel that they are restricted in social contacts or in their work capacity due to economic uncertainty at the start of their career (see Columns III and IV). Thus, a first channel of poorer mental conditions for women with a temporary first contract is that they feel more often depressed and under (time) pressure and less often calm and peaceful.

Looking at men, we find evidence for gender differences in the health consequences of starting a career with a fixed-term contract. Table 4.4 provides two notable insights: First, the coefficients for feeling under pressure show the opposite sign compared to the female sample, the correlations with *Vitality* completely disappear. Men with a temporary job at the beginning of the career seem to be less stressed or depressed than their counterparts with permanent first contracts. Second, parts of the positive association with the overall mental health measure can be attributed to changes in *Social functioning* and *Role emotional*. Unlike women, starting a career with a fixed-term contract motivates men such that they feel more productive at work and can maintain their social contacts. Again, this finding arises not before the third year on the labor market.

Next, we turn our attention to slightly different measures of well-being: Life and health satisfaction are the dependent variables in Table 4.5. Estimation results show that women are less satisfied with their lives when they enter the labor market with a temporary instead of a permanent job. However, this negative relationship holds only contemporaneously and

Table 4.3 : Effects on subcategories of women's mental health

Dependent Variable	Mental health (I)	Vitality (II)	Social functioning (III)	Role emotional (IV)
At labor market entry (1st year)	-1.101	-1.774*	-0.532	0.034
	(0.966)	(0.991)	(1.034)	(1.025)
2nd year on the labor market	-1.478	-2.836***	-0.863	0.329
	(0.945)	(1.006)	(0.965)	(0.997)
3rd year on the labor market	-1.860**	-3.627***	-0.974	-0.259
	(0.943)	(1.079)	(1.030)	(1.079)
4th year on the labor market	-1.384	-3.103***	0.665	0.067
	(0.999)	(1.167)	(1.165)	(1.042)
5th year on the labor market	-1.389	-2.332*	0.282	-0.208
	(1.072)	(1.194)	(1.167)	(1.130)
Observations	1,442	1,442	1,442	1,442

Notes: Marginal effects of starting a career with a fixed-term contract from OLS regressions with robust standard errors clustered at the individual level in parentheses; dependent variables interpolated assuming a linear trend; all regressions contain controls for individual characteristics, background characteristics, job characteristics, personality traits and attitudes, partnership status, ex-ante health status, federal-state and year dummies; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table 4.4 : Effects on subcategories of men's mental health

Dependent Variable	Mental health (I)	Vitality (II)	Social functioning (III)	Role emotional (IV)
At labor market entry (1st year)	-0.312	-1.348	0.945	0.186
	(1.038)	(1.105)	(0.979)	(1.040)
2nd year on the labor market	1.114	-0.010	0.143	0.109
	(1.141)	(1.132)	(1.227)	(1.115)
3rd year on the labor market	3.479***	1.642	1.664	1.731*
	(1.086)	(1.271)	(1.050)	(0.958)
4th year on the labor market	2.783**	1.543	2.229**	1.316
	(1.100)	(1.336)	(0.962)	(1.027)
5th year on the labor market	1.336	1.667	2.698**	2.576**
	(1.298)	(1.351)	(1.070)	(1.094)
Observations	1,180	1,180	1,180	1,180

Notes: Marginal effects of starting a career with a fixed-term contract from OLS regressions with robust standard errors clustered at the individual level in parentheses; dependent variables interpolated assuming a linear trend; all regressions contain controls for individual characteristics, background characteristics, job characteristics, personality traits and attitudes, partnership status, ex-ante health status, federal-state and year dummies; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

becomes negligible and insignificant one year later. Turning to the second subjective measure of well-being, our results suggest no significant differences in reported health satisfaction between individuals starting their career with a fixed-term versus a permanent contract. All in all, our estimations suggest that life and health satisfaction are not the channels trough which mental health is affected.

Table 4.5 : Effects on life and health satisfaction

Dependent variable	Life satisfaction		Health satisfaction	
	(I)	(II)	(III)	(IV)
	Women	Men	Women	Men
At labor market entry (1st year)	-0.249*	-0.171	-0.177	-0.197
	(0.133)	(0.140)	(0.173)	(0.143)
2nd year on the labor market	0.011	-0.068	0.065	-0.012
	(0.131)	(0.156)	(0.153)	(0.161)
3rd year on the labor market	-0.050	0.113	0.071	-0.034
	(0.125)	(0.128)	(0.151)	(0.157)
4th year on the labor market	-0.113	-0.093	0.016	0.039
	(0.140)	(0.137)	(0.163)	(0.166)
5th year on the labor market	-0.032	-0.055	0.307*	-0.047
	(0.139)	(0.139)	(0.167)	(0.167)
Observations	3,315	2,972	3,315	2,972

Notes: Marginal effects of starting a career with a fixed-term contract from OLS regressions with robust standard errors clustered at the individual level in parentheses; all regressions contain controls for individual characteristics, background characteristics, job characteristics, personality traits and attitudes, partnership status, ex-ante health status, federal-state and year dummies; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

The subcategories have already shed some light on the question of channels driving the relationship between economic uncertainty and men's or women's health conditions. To analyze the underlying mechanisms more deeply, we now test whether the associations of fixed-term employment at the start of a career and subsequent health outcomes can be explained by 1) lower wages or 2) changes in the fertility behavior. To this point, we have controlled for many wage predictors. Now, we re-estimate our main specification controlling explicitly for monthly net income assigning a zero wage to unemployed individuals. If the income and its profile over time are the main channel through which fixed-term jobs affect health, the coefficient of fixed-term employment should become much smaller and even insignificant. Similarly, if changes in the fertility behavior account for the negative (positive) correlation, including a control variable for the number of children should at least capture parts of the initial effect. In fact, women with a temporary first job are less likely to have children within the first five years in the labor market. If this translates into lower mental conditions, not the uncertainty but the fertility accounts for the negative estimates. However, Table 4.6 assures that neither income nor fertility is able to explain why starting a career with a fixed-term contract lowers women's mental health in the short run and rises men's health in the short to medium run. The estimated correlations are very similar to our main results, indicating that we still lack a comprehensive explanation for the mechanism.

Table 4.6 : Effects on mental health controlling for income and fertility

Dependent variable	Mental health			
	Women		Men	
	(I)	(II)	(III)	(IV)
At labor market entry (1st year)	-1.068	-1.042	-0.289	-0.256
	(1.029)	(1.025)	(0.980)	(0.981)
2nd year on the labor market	-1.750*	-1.750*	0.415	0.375
	(0.954)	(0.943)	(1.151)	(1.150)
3rd year on the labor market	-2.211**	-2.278**	3.137***	3.093***
	(0.969)	(0.961)	(1.015)	(1.019)
4th year on the labor market	-1.236	-1.242	2.799***	2.771***
	(1.075)	(1.061)	(1.069)	(1.068)
5th year on the labor market	-1.185	-1.351	2.778**	2.742**
	(1.123)	(1.130)	(1.164)	(1.165)
Net income	Yes	No	Yes	No
Number of children	No	Yes	No	Yes
Observations	1,442	1,442	1,180	1,180

Notes: Marginal effects of starting a career with a fixed-term contract from OLS regressions with robust standard errors clustered at the individual level in parentheses; dependent variable interpolated assuming a linear trend; all regressions contain controls for individual characteristics, background characteristics, job characteristics, personality traits and attitudes, partnership status, ex-ante health status, federal-state and year dummies; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

4.4.2 Path Dependence as Potential Mechanism

The transmission channel we propose, is the incarceration effect of starting a career with a fixed-term contract. We call this phenomenon *path dependence* since the type of the initial employment contract prescribes the future career path. Table 4.7 shows the path dependence of starting a career with a fixed-term contract. The probability of repeated spells in fixed-term employment is positive and significant. We argue that the underlying mechanism for the negative effects on women's mental health is this path dependence and the associated economic uncertainty due to the risk of becoming unemployed when the contract ends. However, even if fixed-term employment seems to be little more persistent, men respond differently to economic uncertainty. A potential explanation is that men adjust their behavior to the long-lasting uncertain employment situation and are able to handle the uncertainty in a better way. Instead of increased levels of stress and anxiety about the future, they adapt to the situation and thus improve the mental conditions in the short to medium run. Another possible explanation is that men with a temporary first contract experience better mental health conditions, if they soon find a permanent position. Thus, switching to a permanent job can account for the positive effects on men's mental health.

Table 4.7 : Effects on future fixed-term contract

Dependent variable	Future fixed-term contract	
	(I)	(II)
	Women	Men
2nd year on the labor market	0.421***	0.440***
	(0.050)	(0.061)
3rd year on the labor market	0.247***	0.371***
	(0.045)	(0.061)
4th year on the labor market	0.151***	0.284***
	(0.038)	(0.058)
5th year on the labor market	0.195***	0.091**
	(0.042)	(0.039)
Observations	1,442	1,180

Notes: Marginal effects from OLS regressions of holding a fixed-term contract (conditional on first contract was fixed-term) on starting a career with a fixed-term contract with robust standard errors clustered at the individual level in parentheses; all regressions contain controls for individual characteristics, background characteristics, job characteristics, personality traits and attitudes, partnership status, ex-ante health status, federal-state and year dummies; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

How does the path dependence influence the relationship between mental health and starting a career with a fixed-term contract? To answer this question, we go back to the initial estimation equation replacing our main explanatory variable with two slightly different variables. The first indicator variable signals that a person changes from a fixed-term contract to a permanent contract already after the first year on the labor market. In this case, the uncertainty affects the individual only for a very short period. The second variable is a binary variable indicating that a person remains in fixed-term employment for at least one more year. Thus, we actually split the main measure of starting a career with a fixed-term contract into two variables: the first measures fixed-term employment in the short run and the second a long-lasting uncertainty due to repeated fixed-term contracts. To confirm the hypothesis of path dependence, we expect that the effect on mental health of starting a career with a fixed-term contract is stronger if the individual is exposed to fixed-term employment for more than one period. Table 4.8 reports gender-specific correlations between the type of the first contract and the summary measure of mental health conditional on the duration of the initial fixed-term contract. The estimates in Column I are based on a pooled regression of women's mental health on both explanatory variables and the full set of controls, the ones in Column II make use of the male sample.

We find a large and negative coefficient in the second year on the labor market if a woman who started with a fixed-term contract still holds a fixed-term contract. The coefficient is statistically different from the point estimate for holding a fixed-term contract only in the first year. The same applies to the estimate in the third year even if the difference is marginally not significant any more. This strongly supports our hypothesis of the negative path dependence.

However, the pattern changes in the male sample. The positive relationship seems to be mainly driven by men finding a permanent job during the first year after labor market entry. Thus, we conclude that the path dependence can explain the negative relationship only for women since the positive effect in the male sample is due to the majority of men quickly finding a permanent job.

Table 4.8 : Path dependence: effects on mental health

Dependent Variable	Mental health			
	(I)		(II)	
	Women		Men	
Duration of fixed-term contract	Only 1st year	More than 1 year	Only 1st year	More than 1 year
2nd year on the labor market	1.983	-2.844**	2.792	-1.137
	(1.693)	(1.417)	(2.120)	(1.783)
3rd year on the labor market	1.013	-2.730**	3.121*	1.628
	(1.734)	(1.355)	(1.731)	(1.502)
4th year on the labor market	1.192	-1.752	2.816	1.656
	(1.914)	(1.643)	(1.892)	(1.343)
5th year on the labor market	-0.829	-0.840	2.638	1.832
	(2.077)	(1.420)	(1.971)	(1.396)
Observations	1,442		1,180	

Notes: Marginal effects of starting a career with a fixed-term contract from OLS regressions with robust standard errors clustered at the individual level in parentheses; dependent variable interpolated assuming a linear trend; all regressions contain controls for individual characteristics, background characteristics, job characteristics, personality traits and attitudes, partnership status, ex-ante health status, federal-state and year dummies; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

To sum up, in the short run, we find a negative association between starting a career with a fixed-term contract and mental health, but only for women. The path dependence of starting a career with a fixed-term contract can explain these findings. At least in the short to medium run, affected women might have a higher likelihood of remaining in precarious employment. The prolonged uncertainty might weaken the mental health of these women. The fact that the gap in mental health closes after a few years is related to these women adapting to uncertain circumstances or finding permanent jobs. In contrast, men seem to adapt much faster to the economic uncertainty associated with a temporary first employment contract. Already in the third year after the labor market entry, men who started with a temporary contract report a significantly higher mental health status. Although the persistence of fixed-term employment is even stronger among men than among women, switching to a permanent job, potentially explains the positive effects on men's mental health.

4.4.3 Sensitivity Analysis

In this section we investigate to what extent the previous results for the mental health status are driven by the sample restrictions and the linearity assumption. First, we run the same regressions as in the previous analyses, but omit the linearity assumption. Therefore, we make use of the original data, which contain health information only bi-yearly (Table 4.9, Columns I and II). Second, we add controls describing the characteristics of the first job (Table 4.9, Columns III and IV). Third, we relax the sample restrictions by changing the age-at-graduation cut-off (Table 4.10) as well as by including men and women with less than secondary education (Table 4.11).

At first, Columns I and II of Table 4.9 discuss the sensitivity of the results by making use of the original summary scale of female mental health, that is available only biannual. The association for women between starting a career with a fixed-term contract and mental health does not change qualitatively but quantitatively. The signs remain unchanged but the size and the precision of the coefficients vary slightly: The associations are strong in Year 1, 3 and 5, whereas coefficients are small and insignificant in Year 2 and 4. Since the gaps in the data potentially account for this profile in the coefficients, we prefer the specification with the assumption of a linear trend in the health status. Using averages gives a smoother pattern in the coefficients but does not change the qualitative interpretation of the results.

Table 4.9 : Sensitivity analysis: original mental health measure

Dependent variable	Original mental health index		Mental health	
	(I)	(II)	(III)	(IV)
	Women	Men	Women	Men
At labor market entry (1st year)	-2.338	-0.364	-1.162	-0.261
	(1.785)	(1.736)	(1.038)	(1.049)
2nd year on the labor market	-1.049	-1.078	-2.005**	0.400
	(1.390)	(2.098)	(0.961)	(1.200)
3rd year on the labor market	-3.617**	4.398**	-2.228**	3.182***
	(1.534)	(1.775)	(0.996)	(1.044)
4th year on the labor market	-0.925	1.827	-1.326	2.792**
	(1.767)	(1.671)	(1.100)	(1.133)
5th year on the labor market	-3.427*	4.411***	-1.285	2.851**
	(1.750)	(1.703)	(1.172)	(1.231)
First-job characteristics	No	No	Yes	Yes
Observations	729	588	1,442	1,180

Notes: Marginal effects of starting a career with a fixed-term contract from OLS regressions with robust standard errors clustered at the individual level in parentheses; Columns I and II: original biannual mental health index; Columns III and IV: interpolated annual measure of mental health; regressions contain controls for individual characteristics, background characteristics, job characteristics, personality traits and attitudes, partnership status, ex-ante health status, federal-state and year dummies; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

The second robustness check adds time-invariant first-job characteristics to the main regression equation. These *fixed effects* revoke all variation in the mental health status due to the characteristics of the first job. The set of additional controls comprises the working hours, self-employment as well as occupation and industry dummies of the first job. Columns III and IV of Table 4.9 report the coefficients of starting a career with a fixed-term contract from the enlarged specification. The coefficients are very similar to the main results. Women's mental health is reduced in the second and third year on the labor market, whereas men whose first contract is of limited duration exhibit higher health conditions from the third year onward. Thus, we are convinced that selection based on the first-job characteristics cannot explain the significant associations from the main tables.

Next, we provide evidence that the choice of the age-at-graduation cut-offs does not alter the findings. In the main sample, we restrict the age at graduation to a maximum of 35 years. Table 4.10 shows the coefficients from regressions with a lower and higher age cut-off. The results confirm the previous findings: women tend to indicate a lower mental health status in the years after career start if the first contract is limited in duration. The size of the estimates in the sample with the oldest women being 30 at graduation is somewhat larger compared with the numbers in the main regressions. Extending the sample up to age 40 dampens the effect slightly. However, statistically, the coefficients are not different from the ones in Table 4.2.

Table 4.10 : Sensitivity analysis: different age-at-graduation cut-offs

Dependent variable	Mental health			
	Women		Men	
	(I)	(II)	(III)	(IV)
Sample	Age at grad. ≤30	Age at grad. ≤40	Age at grad. ≤30	Age at grad. ≤40
At labor market entry (1st year)	-1.297	-1.405	-0.169	-0.605
	(1.102)	(0.975)	(1.143)	(1.024)
2nd year on the labor market	-1.851*	-1.769*	0.355	0.345
	(0.984)	(0.910)	(1.325)	(1.160)
3rd year on the labor market	-2.539**	-1.981**	3.099***	2.686**
	(1.002)	(0.948)	(1.157)	(1.047)
4th year on the labor market	-1.943*	-1.234	2.939**	2.502**
	(1.089)	(1.044)	(1.208)	(1.098)
5th year on the labor market	-1.731	-1.211	3.749***	1.925*
	(1.171)	(1.104)	(1.259)	(1.142)
Observations	1,225	1,577	964	1,366

Notes: Marginal effects of starting a career with a fixed-term contract from OLS regressions with robust standard errors clustered at the individual level in parentheses; dependent variable interpolated assuming a linear trend; all regressions contain controls for individual characteristics, background characteristics, job characteristics, personality traits and attitudes, partnership status, ex-ante health status, federal-state and year dummies; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

The results for men approve that fixed-term employment at the beginning of the career positively affects men's mental health. Again, the estimates are smaller for the sample including older men but the basic insights remain unchanged. Thus, the negative relationship in the female and the positive in the male sample seem to be driven by younger individuals. Nevertheless, we can conclude that the choice of the age-at-graduation cut-off does not significantly influence our results.

Finally, we relax the education restriction and include men and women who have no secondary school degree. As Table 4.11 shows, these individuals do not substantially change the estimates. The main pattern from the previous regressions remains the same. In the very short run, women react to temporary employment at the beginning of their careers with reduced mental health. This effect becomes weaker over time but the negative sign does not vanish. The size of the association in the second and third year is slightly smaller but statistically not distinguishable from the former results. In the male sample, previous findings are confirmed as well.

Table 4.11 : Sensitivity analysis: including individuals without secondary degree

Dependent variable	Mental health	
	(I)	(II)
	Women	Men
At labor market entry (1st year)	-1.197	0.039
	(1.006)	(0.996)
2nd year on the labor market	-1.555	0.703
	(0.957)	(1.058)
3rd year on the labor market	-1.982**	2.376**
	(0.940)	(1.004)
4th year on the labor market	-1.034	2.076**
	(1.028)	(1.054)
5th year on the labor market	-1.046	2.208*
	(1.098)	(1.143)
Observations	1,562	1,320

Notes: Marginal effects of starting a career with a fixed-term contract from OLS regressions with robust standard errors clustered at the individual level in parentheses; dependent variable interpolated assuming a linear trend; all regressions contain controls for individual characteristics, background characteristics, job characteristics, personality traits and attitudes, partnership status, ex-ante health status, federal-state and year dummies; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

4.4.4 Heterogeneous Health Effects by Education

Motivated by the previous section, we analyze the effects of starting a career with a fixed-term contract for educational subgroups. Table 4.12 shows that women as well as men who have less than a university degree are responsible for the effects of starting a career with a fixed-term contract. The associations are very similar to the overall effects even if the pattern for women with secondary education differs slightly. The coefficient remains large and significant also after 5 years on the labor market. These women probably have difficulties to find a permanent job and remain in fixed-term employment. In contrast, among women with university education, estimation results suggest no significant association between the type of the first contract and mental health. An potential explanation is that highly educated women are able to change to a permanent position more quickly. For men with secondary education, we observe the opposite effects. Three years after they started a career with a temporary job, they report a significantly higher mental health status. In contrast, the mental health of men with tertiary education seems to be unaffected by the type of contract held in their first job.

Table 4.12 : Effects on mental health by education

Dependent variable	Mental health			
	Women		Men	
	(I) Secondary education	(II) Tertiary education	(III) Secondary education	(IV) Tertiary education
At labor market entry (1st year)	-0.467	-0.396	-0.120	-2.361
	(1.485)	(1.508)	(1.357)	(1.569)
2nd year on the labor market	-1.513	-0.372	0.994	-1.784
	(1.340)	(1.516)	(1.909)	(1.505)
3rd year on the labor market	-2.371*	-0.405	4.886***	-0.099
	(1.385)	(1.682)	(1.142)	(1.655)
4th year on the labor market	-1.862	0.515	4.055***	0.837
	(1.573)	(1.722)	(1.340)	(1.640)
5th year on the labor market	-2.348	0.768	3.772**	1.041
	(1.638)	(1.903)	(1.488)	(1.956)
Observations	867	542	721	433

Notes: Marginal effects of starting a career with a fixed-term contract from OLS regressions with robust standard errors clustered at the individual level in parentheses; dependent variable interpolated assuming a linear trend; all regressions contain controls for individual characteristics, background characteristics, job characteristics, personality traits and attitudes, partnership status, ex-ante health status, federal-state and year dummies; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

4.4.5 Selection into Fixed-term Employment

In this section, we present evidence that the effect of starting a career with a fixed-term contract is not purely due to selection. To this point, we are able to estimate a consistent average treatment effect on the treated (ATT), that is, the mean effect for those men and women who started their career with a fixed-term contract ($FT = 1$) conditional on all observable factors, X: $ATT = E(Y_1 - Y_o|X, FT = 1)$. If we are willing to belief that, conditional on all observable characteristics, holding a fixed-term contract at the start of the career is randomly assigned, then we are able to estimate a consistent average treatment effect (ATE). This is the expected effect on a randomly drawn person from the population of men and women entering the labor market not later than age 35 (see Wooldridge, 2010, Chapter 21).

However, the underlying problem is that starting a career with a fixed-term contract is not randomly assigned. For instance, an individual with very high ability will be more likely to find a permanent contract – at least it is more likely that her employer will offer her a permanent contract after a short probation period with a fixed-term contract. At the same time, she might have fewer mental problems due to economic uncertainty since she is convinced that her ability will be an advantage in finding an appropriate and permanent position. Such a situation would cause a positive bias and the results would underestimate the true effect. Fortunately, the SOEP data allow us to control for a variety of individual characteristics that are typically unobserved in other survey data and proxy well for unobserved ability. Thus, all regressions include controls for the degree of risk aversion, the locus of control as well as a set of personality traits and general attitudes.

Nevertheless, we further investigate the selection issue by testing whether any of the predetermined observable characteristics significantly affects the likelihood of starting the career with a fixed-term contract (Table C.3 in the Appendix). Almost none of the coefficients are significantly different from zero, the only exception in the female sample is *age*, meaning that older women are less likely to start a career with a fixed-term contract. In the male sample none of the explanatory variables helps predicting the type of the first contract. It is important to note that all variables related to the health status before labor market entry have no explanatory power for the type of first contract. Even if we include ex-ante life and health satisfaction, as in Columns II and IV, the coefficients remain insignificant. This result is reassuring and important as it provides further supporting evidence against the possibility of health-related self-selection into fixed-term contracts at labor market entry.

Finally, we investigate the robustness of our results to omitted variable bias. In so doing, we follow Oster (2013) who recently developed a novel method to assess the bias that arises from unobserved factors. Oster's method relies on the choice of the degree of proportionality between the selection based on observable and unobservable factors ($\tilde{\delta}$). We assume equal importance of observable and unobservable factors, that is $\tilde{\delta} = 1$. We are then able to compute a bias-adjusted coefficient (β^*) of starting a career with a fixed-term contract equal to

$$\beta^* \left(R^2_{max}, \tilde{\delta}\right) = \tilde{\beta} - \tilde{\delta} \left[\dot{\beta} - \tilde{\beta}\right] \frac{R^2_{max} - \tilde{R}^2}{\tilde{R}^2 - \dot{R}^2}. \tag{4.2}$$

$\dot{R}^2$ and $\dot{\beta}$ come from a regression of the health outcome on starting a career with a fixed-term contract without further controls. $\tilde{R}^2$ and $\tilde{\beta}$ arise from a controlled regression as in Table 4.2.[3] The adjustment of $\tilde{\beta}$ accounts for changes in the estimated coefficients due to the observable controls weighted by relative changes in the explained variation in mental health. R^2_{max} is the hypothetical value of R^2 controlling for all observable and unobservable characteristics.

[3] If we include control variables in the male sample, the coefficients rise. Only under the assumption $\tilde{\delta} < 0$, that is, the selection based on observable and based on unobservable characteristics go in different directions, the bias adjustment would make sense. Since the derivation of the bias-adjusted coefficient in this case is not straight forward, we refrain from showing bias-adjusted estimates for men.

Following Oster (2013), we choose $R^2_{max} = min\{2.2\tilde{R}^2, 1\}$. In the end, we obtain an identified set of the treatment effect of starting a career with a fixed-term contract $\left[\tilde{\beta}, \beta^*\left(R^2_{max}, \tilde{\delta}\right)\right]$. If this set excludes 0, the controlled coefficients can be considered robust to omitted variable bias.

Table 4.13 : Robustness to omitted variable bias of the results for women

Dependent variable	Mental Health		
	(I)	(II)	(III)
	Baseline effect $\dot{\beta}$	Controlled effect $\tilde{\beta}$	Identified set $[\tilde{\beta}, \beta^*]$
At labor market entry (1st year)	-1.231	-1.071	[-1.071, -0.870]
	(1.095)	(1.027)	
2nd year on the labor market	-1.894*	-1.716*	[-1.716, -1.493]
	(1.005)	(0.95)	
3rd year on the labor market	-2.430**	-2.185**	[-2.185, -1.878]
	(1.011)	(0.968)	
4th year on the labor market	-1.820*	-1.144	[-1.144, -0.295]
	(1.096)	(1.073)	
5th year on the labor market	-2.127*	-1.191	[-1.191, -0.018]
	(1.194)	(1.126)	
R^2	$\dot{R}^2$: 0.012	$\tilde{R}^2$: 0.277	R^2_{max}: 0.610
Observations	1,442	1,442	

Notes: Marginal effects of starting a career with a fixed-term contract from OLS regressions with robust standard errors clustered at the individual level in parentheses; Column I: baseline regression contains no controls; Column II: controlled regression contains controls for individual characteristics, background characteristics, job characteristics, personality traits and attitudes, partnership status, ex-ante health status, federal-state and year dummies; Column III: identified set gives a lower bound for the coefficients from the controlled specification assuming equal selection based on observable and unobservable characteristics, $\tilde{\delta} = 1$; *** p<0.01, ** p<0.05, * p<0.1.

Table 4.13 shows in Column I the coefficients from the baseline model without controls, whereas Column II displays the controlled coefficients. For each year after labor market entry, the estimates become slightly smaller if we include the full set of controls. In addition, the R^2 increases from 0.012 to 0.277 meaning that the controlled model is able to explain a much larger part of the variance in mental health. In the last column, we present the identified set with the controlled coefficient as upper bound and the bias-adjusted coefficient as lower bound. Since all sets exclude 0, we can conclude that under the assumption of equal selection based on observable and unobservable characteristics our controlled estimates are robust to omitted variable bias.

4.5 Conclusion

The analysis sheds light on potential spill-over effects of fixed-term employment on health outcomes. Using German data on young male and female graduates, we find a substantial reduction in female mental well-being due to fixed-term employment at the beginning of the career. This association is significant only in the short run, and has no effect after the third year on the labor market. In contrast, young male graduates are not affected in the short run but starting with the third year after labor market entry they report an improved mental health status. Detailed analyses of the subcategories of the summary measure of mental health reveal that women with a limited first employment contract experience more (time) pressure and feel more often depressed. For men, however, limitations in social interactions and job productivity are alleviated. The main results are robust to several sensitivity tests.

A possible explanation is the path dependence of starting a career with a fixed-term contract. The likelihood of repeated spells in precarious jobs is significantly higher when entering the labor market with a temporary contract. Thus, we measure an indirect effect of the initial contract on health, that works through recurrent contemporaneous economic instabilities. The heterogeneity analysis shows that women with secondary education are particularly affected. We address potential endogeneity threats by including a large set of controls and by showing evidence against health-related self-selection into temporary contracts at the beginning of a career. Among men and women, ex-ante baseline health indicators are not related to the type of contract in the first job. Hence, the results suggest a robust, negative association between fixed-term employment and women's mental health in the short run and a positive relationship with men's mental health in the short to medium run.

Subject of this study are graduates and newcomers on the labor market from the generations below age 35. As a result of their youth, they suffer less from physical problems but are rather affected by mental illness. If economic uncertainty due to fixed-term employment at the beginning of the career facilitates poor mental conditions for women, this implies costs that are probably not indented by the policymakers. Given the importance of mental health problems for absenteeism from work, further research should also take into account potential consequences of fixed-term employment for employers and health insurance companies.

Appendix C.1 Supplementary Tables

Table C.1 : Summary statistics of control variables for female sample

	N	Mean	SD	Min	Max
Age	1442	28.08	4.12	18	39
Migratory background	1442	0.21	0.41	0	1
Years of education	1442	13.96	2.68	7	18
Born in East Germany	1442	0.36	0.48	0	1
Mother tertiary education	1442	0.15	0.36	0	1
Mother employed	1442	0.39	0.49	0	1
Mother's age at birth	1442	26.38	5.21	16	43
Siblings	1442	0.90	0.29	0	1
Openness	1442	4.80	1.13	1.33	7
Agreeableness	1442	5.58	0.89	2.33	7
Conscientiousness	1442	5.77	0.88	2.67	7
Extraversion	1442	5.12	1.21	1.67	7
Neuroticism	1442	4.35	0.81	2.33	6.33
Risk aversion	1442	0.05	0.23	0	1
Internal locus of control	1442	4.11	0.67	2.56	6.22
Importance of having children	1442	0.71	0.46	0	1
Importance of partnership	1442	0.92	0.27	0	1
Importance of career	1442	0.92	0.27	0	1
Importance of affording sth.	1442	0.88	0.32	0	1
Unemployed	1442	0.22	0.42	0	1
Weekly working hours	1442	30.06	18.89	0	75
Part-time work	1442	0.14	0.34	0	1
Self-employed	1442	0.04	0.20	0	1
Blue collar job	1442	0.07	0.26	0	1
White collar job	1442	0.60	0.49	0	1
Doctor visits	1442	9.12	11.38	0	80
Hospital stays	1442	0.11	0.44	0	4
Partnership	1442	0.78	0.42	0	1

Table C.2 : Summary statistics of control variables for male sample

	N	Mean	SD	Min	Max
Age	1180	28.75	4.29	19	39
Migratory background	1180	0.15	0.36	0	1
Years of education	1180	13.82	2.89	7	18
Born in East Germany	1180	0.36	0.48	0	1
Mother tertiary education	1180	0.13	0.34	0	1
Mother employed	1180	0.41	0.49	0	1
Mother’s age at birth	1180	26.29	4.76	16	46
Siblings	1180	0.83	0.38	0	1
Openness	1180	4.61	1.12	1.67	7
Agreeableness	1180	5.44	0.92	2.67	7
Conscientiousness	1180	5.64	0.93	2	7
Extraversion	1180	4.84	1.19	1.67	7
Neuroticism	1180	4.05	0.82	2	6.67
Risk aversion	1180	0.02	0.15	0	1
Internal locus of control	1180	4.07	0.66	2.67	6.33
Importance of having children	1180	0.51	0.50	0	1
Importance of partnership	1180	0.88	0.33	0	1
Importance of career	1180	0.95	0.23	0	1
Importance of affording sth.	1180	0.93	0.26	0	1
Unemployed	1180	0.12	0.33	0	1
Weekly working hours	1180	39.04	17.12	0	80
Part-time work	1180	0.03	0.17	0	1
Self-employed	1180	0.08	0.27	0	1
Blue collar job	1180	0.26	0.44	0	1
White collar job	1180	0.46	0.50	0	1
Doctor visits	1180	4.54	7.75	0	60
Hospital stays	1180	0.06	0.25	0	2
Partnership	1180	0.64	0.48	0	1

Table C.3 : Selection into fixed-term contract at the beginning of the career

Dependent Variable	Starting the career with a fixed-term contract			
	Women		Men	
	(I)	(II)	(III)	(IV)
Ex-ante mental health	0.000	0.002	-0.002	-0.002
	(0.003)	(0.003)	(0.003)	(0.003)
Ex-ante health satisfaction		-0.005		0.001
		(0.018)		(0.018)
Ex-ante life satisfaction		-0.017		-0.008
		(0.019)		(0.017)
Number of annual doctor visits	-0.000	-0.001	-0.000	-0.000
	(0.002)	(0.003)	(0.004)	(0.004)
Number of nights in hospital	0.054	0.057	-0.010	-0.010
	(0.047)	(0.047)	(0.084)	(0.085)
Age	-0.013*	-0.013	0.002	0.002
	(0.008)	(0.008)	(0.007)	(0.008)
Migratory Background	-0.027	-0.031	-0.030	-0.032
	(0.070)	(0.069)	(0.072)	(0.074)
Years of education	0.015	0.016	0.003	0.004
	(0.013)	(0.013)	(0.014)	(0.014)
Born in East Germany	-0.034	-0.035	-0.140	-0.144
	(0.097)	(0.095)	(0.090)	(0.093)
Mother tertiary education	-0.118	-0.121	-0.026	-0.024
	(0.077)	(0.078)	(0.082)	(0.083)
Mother employed	-0.071	-0.067	-0.024	-0.023
	(0.085)	(0.084)	(0.087)	(0.088)
Mother's age at birth	0.006	0.006	-0.003	-0.003
	(0.006)	(0.006)	(0.006)	(0.006)
Siblings	0.013	0.013	-0.042	-0.041
	(0.088)	(0.088)	(0.072)	(0.074)
Openness	-0.005	-0.002	0.021	0.022
	(0.027)	(0.027)	(0.025)	(0.025)
Agreeableness	0.004	0.002	-0.009	-0.009
	(0.031)	(0.031)	(0.029)	(0.029)
Conscientiousness	-0.046	-0.046	-0.016	-0.016
	(0.037)	(0.037)	(0.027)	(0.028)
Extraversion	-0.021	-0.018	-0.033	-0.033
	(0.023)	(0.023)	(0.024)	(0.025)
Neuroticism	0.054	0.056	-0.010	-0.011
	(0.036)	(0.036)	(0.034)	(0.034)
Risk aversion	-0.055	-0.048	0.088	0.079
	(0.114)	(0.118)	(0.105)	(0.109)
Internal locus of control	0.009	0.004	-0.046	-0.051
	(0.043)	(0.043)	(0.043)	(0.045)
Weekly working Hours	-0.004	-0.004	0.003	0.003
	(0.003)	(0.003)	(0.004)	(0.004)
Part-time employed	-0.016	-0.023	0.245	0.238
	(0.089)	(0.088)	(0.174)	(0.177)
Self-employed	-0.240	-0.241	-0.243	-0.238
	(0.157)	(0.158)	(0.159)	(0.161)
Blue collar job	0.181	0.176	0.249	0.253
	(0.164)	(0.163)	(0.163)	(0.164)
White collar job	0.102	0.103	0.108	0.112
	(0.117)	(0.117)	(0.150)	(0.151)
Partnership	0.021	0.020	-0.087	-0.083
	(0.061)	(0.062)	(0.055)	(0.056)
Observations	297	297	245	245
R^2	0.325	0.328	0.365	0.365

Notes: Marginal effects from OLS regressions with robust standard errors clustered at the individual level in parentheses; dependent variables interpolated assuming a linear trend; all regressions contain industry, federal-state, and year dummies; *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

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