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Essays on Transition Economies

Kovacheva, Penka

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ESSAYS ON TRANSITION ECONOMIES

Penka Kovacheva

A DISSERTATION

PRESENTED TO THE FACULTY

OF PRINCETON UNIVERSITY

IN CANDIDACY FOR THE DEGREE

OF DOCTOR OF PHILOSOPHY

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Advisor: Anne Case

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Abstract

This thesis consists of three chapters, which analyze how labor market and policy uncertainty affect education, employment, fertility, and well-being in the context of three transition economies.

The first chapter examines the cause of the unprecedented, and still unexplained, 40% decline in total fertility rates across the former Communist Bloc during the 1990s. Using individual-level panel data from Russia and Ukraine, I show that a substantial increase in wage uncertainty in the 1990s triggered the fertility decline, and that a moderate decline in wage uncertainty in the 2000s prompted a partial fertility recovery. I build a dynamic structural model of the joint decisions over fertility and job sector: private, state, and non-work. In the model, wage uncertainty affects fertility because children are illiquid assets. I estimate the key parameters of the model exploiting panel variation in wage uncertainty across job sectors, regions, and time. I construct a measure of wage uncertainty which is a function of the incidence of job dismissal, administrative unpaid leave, and wage arrears. Counterfactual exercises show that lowering wage uncertainty in the private sector to the level of the state sector reduces the fertility gap between sectors by 51-57%. Lowering the levels of the 1990s uncertainty in both sectors to the levels of the late 2000s increases overall fertility by 26-30%. By contrast, I estimate much smaller effects from a traditional fertility stimulus in the form of more generous maternity benefits. These results suggest that in economies with poor labor law enforcement, targeting employment violations (such as wage arrears) can be part of effective fertility policies. More broadly, labor market uncertainty can impose significant negative externalities on irreversible investments such as children.

The second chapter, co-authored with Xiaotong Niu, investigates the effects of a 1996 pension crisis in Russia on subjective well-being. We use a difference-in-difference strategy and an individual fixed-effects model, which isolate the pension crisis from other changes in the economic environment and account for individual heterogeneity in reported well-being. We find that the failure to receive one's pension during the crisis not only significantly reduces current life satisfaction, but also depresses expectations for the future. Additionally, the effect of the crisis extends to non-pensioners who live in households with pensioners in arrears: they experience an equally strong

and significant decline in life satisfaction, even after accounting for their personal income. We also find that the crisis, despite being a purely monetary shock, affects well-being in ways that go beyond the monetary size of pension loss. Policies aimed to fully compensate for such disruptions in the redistribution system would need to take these externalities into account.

The third chapter investigates the dynamics of human capital returns and their effect on wage inequality during Bulgaria's transition to a market economy. I use regression analysis to estimate the monetary returns to different individual characteristics, while accounting for the selection of individuals into the new market economy. I also use semi-parametric analysis to examine counterfactual earnings distributions under several scenarios for how individual characteristics could have evolved over the transition. I find that returns to human capital are not substantially affected by gender, ethnicity, or sector of employment but differ by age. Returns to work experience are stagnating as the market economy places almost no value to experience acquired under communism. Returns to education start low relative to both Western Europe and some of Central Europe but increase throughout the 1990s. These rising returns to education are the single most important driver behind the growth in wage inequality, accounting for over 75% of it. This leads to an interesting policy prescription: little if any policy action is needed since the rise in inequality can be expected to level off as subsequent generations get educated under the new system.

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Contents

Abstract **iii**

Acknowledgements **v**

1 Transition, Wage Uncertainty, and Fertility **1**

1.1 Introduction 1

1.2 Background 6

1.2.1 Fertility trends in Russia and Ukraine 6

1.2.2 Labor market transformation after the fall of the Soviet Union 7

1.3 The model 8

1.3.1 Preferences 9

1.3.2 Budget constraints 10

1.3.3 The woman's planning problem 14

1.3.4 Expectations about uncertainty 14

1.3.5 Problem solution 15

1.3.6 Implications of the model 15

1.4 Data and empirical analysis 16

1.4.1 The data 16

1.4.2 Sample characteristics 18

1.4.3 Wage uncertainty measures 19

1.4.4 Cross-sectional estimation 22

1.4.5 Panel estimation . . . 24
1.5 Structural estimation . . . 25
1.5.1 Estimation method . . . 26
1.5.2 Results . . . 27
1.5.3 Counterfactual exercises . . . 28
1.6 Discussion . . . 30
1.6.1 Other potential factors affecting fertility . . . 30
1.6.2 The role of husband's job sector . . . 32
1.7 Conclusion . . . 34
1.A Construction of uncertainty and child cost measures . . . 65
1.A.1 Wage uncertainty measures . . . 65
1.A.2 Child cost measures . . . 68
1.B Mean and variance of the wage distribution . . . 68
1.C Additional Results . . . 72

2 The Mental Cost of Pension Loss 74

2.1 Introduction . . . 74
2.2 Data . . . 79
2.3 The pension crisis of 1996 . . . 82
2.4 Empirical strategy . . . 84
2.4.1 Main estimates . . . 84
2.4.2 Channels of the pension crisis . . . 87
2.4.3 Intra- and inter-household externalities . . . 89
2.4.4 Specification checks . . . 90
2.5 Results . . . 92
2.5.1 The effect of the crisis on SWB of pensioners . . . 92
2.5.2 The effect of the crisis on SWB of non-pensioners . . . 94
2.5.3 Robustness checks . . . 96

2.6 Additional considerations 97
2.7 Discussion and conclusions 100

3 Human Capital and Wage Inequality **127**

3.1 Introduction 127
3.2 Data 129
3.3 Empirical methods 130
3.3.1 Estimation of returns to human capital 130
3.3.2 Estimation of wage inequality 133
3.4 Results 136
3.4.1 Returns to human capital 136
3.4.2 Wage inequality 140
3.5 Conclusions 144

Bibliography **158**

Chapter 1

Transition, Wage Uncertainty, and Fertility: Evidence from Russia and Ukraine

1.1 Introduction

This chapter examines how wage uncertainty affects women's choices of where to work and when to have children throughout the childbearing years. In the past two decades the countries of the former Communist Bloc experienced unprecedented changes in fertility that remain unexplained to this day. Across the region, total fertility rates (TFR) dropped from slightly above 2 children per woman in 1989 to below 1.3 children in 1999 and back up to 1.5 in 2010. Russia and Ukraine in particular saw a TFR drop of over 40% during the 1990s, which was then followed by an increase of over 30% during the 2000s.[1] During this same period, unemployment rates climbed from virtually zero at the end of the 1980s to around 12 – 13% in the late 1990s before going back down to single digits. Two other sources of wage uncertainty, wage arrears and compulsory unpaid leave, appeared in the period after the fall of the Soviet Union and followed similar trajectories to the unemployment rate. In addition to the substantial time variation in wage uncertainty, there was also large variation in uncertainty across job sectors – state and private – and across geographic

[1]The data source for these statistics is the TransMONEE database, which can be found at http://www.transmonee.org.

regions. Overall, the state sector had lower wage uncertainty than the private sector, and the gap between the two sectors was larger during the first few years of the transition period. I analyze the effects of this variation in wage uncertainty on women's fertility behavior using individual-level panel data from Russia and Ukraine that span the period 1994 – 2010.

For the purpose of this analysis, I build a dynamic structural model of women's joint choices of where to work and when to have children. In my model, wage uncertainty affects fertility not only because individuals are risk-averse but also because children are illiquid assets and borrowing and saving are prohibited (reflecting high inflation and undeveloped financial markets during the transition years). I use a dynamic framework because the decision to have a child in any period is associated with child costs in every subsequent period (Becker, 1960) and the decision to work in a certain sector influences subsequent job choices through accumulated job sector experience. In addition, the model traces women's decisions from the start of childbearing until more than a decade later in order to capture long-term effects. In finance theory (Pyndick 1991, Dixit 1992), persistent uncertainty only causes postponement of irreversible investments, but because women are constrained by a biological clock, completed fertility can also be affected. Lastly, I structurally estimate the model's parameters to quantify the long-term fertility response to counterfactual scenarios for wage uncertainty and to a traditional fertility stimulus in the form of more generous maternity leave. This analysis can be informative for the design of fertility policies in economies struggling with low fertility and high wage uncertainty.

The contribution of this chapter is fourfold. The first contribution is establishing that the empirical relationship between wage uncertainty and fertility holds outside developed economies. As such, this chapter adds to a large OECD literature on this subject (e.g. Mira and Ahn 2002 and Da Rocha and Fuster 2005).[2] Transition economies are particularly suitable for studying the rela-

[2]In the European Union, Adsera (2011) shows that income uncertainty due to institutional features, such as short-term contracts, can delay birth. De La Rica and Izza (2005) find that fixed-term contracts delay entry into motherhood in Spain, and Andersson (2000) finds that stable employment is associated with higher fertility in Sweden. In the United States, Dehejia and Lleras-Muney (2004) find that higher unemployment is associated with lower fertility for credit-constrained Black women. Sommer (2010) shows that American occupations with lower earnings variability (teach-

tionship between uncertainty and fertility. Properly identifying the role of macro-level uncertainty is generally difficult due to the lack of a control group. The transition from a state- to a market-driven economy provides an opportunity to exploit variation in uncertainty across job sectors based on firm ownership. The inter-sector gap in uncertainty has both a time and a regional dimension that allows for clean identification.

The second contribution of the chapter is resolving the inconclusive evidence from the transition economy literature. Despite the above-mentioned advantages of a transition economy context, transition studies find mixed effects of wage uncertainty. One possible reason is the wide variety of uncertainty measures. For example, studies have measured uncertainty based on self-reported feelings of uncertainty (Bhaumik and Nugent 2005), on a recent job change (Chase 2003), and, *separately*, on regional unemployment or wage arrears (Kohler and Kohler 2002). Such measures are problematic for various reasons. It is unclear whether subjective feelings capture the effects of economic conditions or the respondents' pessimism.[3] Also, having changed jobs in the past does not necessarily reflect uncertainty. In this chapter, I show that a job change could in fact be linked to future fertility intentions. Likewise, a separate analysis of each source of uncertainty would fail to identify an effect when the various sources of uncertainty counteract one another. For this purpose, I build an aggregate measure of uncertainty to incorporate *all available* sources of wage non-payment. Indeed, I find that in Russia state-sector wages appear more uncertain in times of wage arrears, but when the likelihood of job dismissal and compulsory unpaid leave are also considered, state-sector wages are less uncertain overall.

The third contribution of the chapter is measuring uncertainty in several ways to account for endogeneity bias and time-invariant heterogeneity across individuals. Measuring uncertainty at the individual level poses a problem of endogeneity. It is conceivable that unobservable individual characteristics underlie the relationship between individual experience of uncertainty and fertility

ers, health-care professionals, engineers) are associated with higher fertility than occupations with higher earnings variability (sciences, arts, entertainment).

[3]Recent evidence from Russia shows that the information elicited from such self-reports is only weakly correlated to individuals' actual economic situation (Gimpelson and Oshchepkov 2012).

preferences. I address this issue in several ways. First, I create measures of the probability of wage non-payment at the sector and region level. Second, I also use community-level data from Russia on closures of state enterprises as an alternative measure of state sector wage uncertainty in the community. This strategy is similar to the literature that uses plant closures as exogenous variation (e.g. Huttunen and Kellokumpu 2010 and Lindo 2010).[4] Third, I exploit the panel nature of the data to control for time-invariant heterogeneity and trace the association between changes in woman's fertility and changes in her job sector over time.

The fourth contribution of the chapter is modeling the joint nature of women's decisions over job sector and fertility and structurally estimating the parameters of these decision processes. To date, no dynamic structural model of joint labor supply and fertility, with endogeneous work experience, has been estimated in a context outside developed economies.[5] As a result, existing models have viewed the link between labor supply and fertility choices as being based on intertemporal trade-off between consumption and leisure. Situating my model in an economy with uncertain labor and undeveloped credit markets allows me to investigate an additional, previously unexplored, link between labor supply and fertility choices: sector- and time-variant wage uncertainty. In addition, most other studies that investigate the fertility response to policy changes in child subsidies or parental leave rely on natural experiments and use difference-in-difference models that capture short-term effects.[6] By contrast, I use structural estimation of my model to capture the long-term fertility response over the childbearing years to counterfactual changes in wage uncertainty and in maternity benefits.

I use data from the Russia Longitudinal Monitoring Survey (RLMS) Phase II, which consists of 15 survey waves from 1994 – 2010, and the Ukrainian Longitudinal Monitoring Survey (ULMS),

[4]However, they are not exactly the same. My aim is to identify the effect of uncertainty resulting from working in a sector in which jobs become more uncertain. This literature, on the other hand, aims to identify the effect of an individual's own experience of job loss resulting from plant closure.

[5]To the best of my knowledge, the only existing studies are Francesconi (2002) and Keane and Wolpin (2010), and a recent working paper by Adda et al. (2011), all of which analyze labor markets in developed economies (the United States and Germany).

[6]For example, see Cohen et al. (2010) or Lalive and Zweimueller (2009).

which consists of 3 waves in 2003, 2004, and 2007. The latter survey also contains detailed information on fertility and complete employment history since 1997, which I use to create a panel dataset for the period $1997-2007$. I use these data to construct a time-, region-, and sector-specific measure of wage uncertainty. I measure wage uncertainty as the risk of wage non-payment, which is a function of the incidence of job dismissal, administrative unpaid leave, and wage arrears. I use this measure in both a reduced-form and structural analysis.

Cross-sectional regression analysis shows that, in both Russia and Ukraine, the likelihood of having a baby is $5-6$ percentage points higher for women who work in state jobs in a region and time period when state jobs have 1 percentage point lower chance of wage non-payment. This relationship is also evident in the time series. I find that a fertility gap of about 0.1 children emerges between women who leave for the state sector and women who remain in the private sector. This gap emerges as soon as one year after the sector switch while there is no gap prior to the switch. In addition, every 1 percentage point increase in the relative probability of wage non-payment in the private sector is associated with $0.4-0.6$ increase in the fertility gap after the sector switch.

The results from the structural analysis closely follow those of the regression analysis. The model fits well the observed trends of employment and fertility by job sector. In addition, the simulated dynamics of the fertility response to a sector switch matches the observed dynamics. I then use the estimated structural parameters to evaluate how fertility behavior would change in several counterfactual regimes. Under the first scenario, I lower the wage uncertainty in the private sector to equal that of the state sector for a given year and geographical region. This counterfactual exercise reduces the fertility gap between the two sectors by $51-57\%$. Under the second scenario, I lower the uncertainty in both sectors from the high levels of the 1990s to the much lower levels of the late 2000s. This counterfactual produces an increase in overall fertility of $26-30\%$. By contrast, a counterfactual policy that increases the duration of fully paid maternity leave by 20% increases overall fertility by only 12% in Russia and by only 5% in Ukraine. These results have important policy implications for policymakers in transition economies, who are concerned with low fertility levels. In particular, my findings suggest that in the presence of poor labor law en-

forcement, targeting employment violations can be an important component of an effective fertility stimulus.

The remainder of the chapter is organized as follows: Section 1.2 provides an overview of the trends in fertility and wage uncertainty in Russia and Ukraine since the fall of the Soviet Union. Section 1.3 presents the dynamic model. Section 1.4 introduces the data and describes the regression analysis. Section 1.5 presents the structural analysis and results from counterfactual analysis of fertility under different regimes of wage uncertainty and a policy with more generous maternity leave. Section 1.6 discusses additional evidence, and Section 1.7 summarizes the results and discusses their policy implications.

1.2 Background

Russia and Ukraine experienced large fluctuations in fertility after the collapse of the Soviet Union; fertility decreased in the 1990s and moderately increased in the 2000s. These changes coincided with a massive labor market transformation. The socialist model of full employment and lifetime tenures at state firms was replaced by widespread experience of unemployment as private owners – of privatized state firms or newly created private firms – reduced the size of their labor force. What is more, wages were not guaranteed even if one was formally employed, and wage arrears and compulsory unpaid leave emerged. This section presents an overview of these trends for the two largest economies in the former Soviet Union, Russia and Ukraine.

1.2.1 Fertility trends in Russia and Ukraine

If wage uncertainty during the transition explains the fertility trend, it is necessary to first establish that the fertility trend did not pre-date the economic transition.

Figure 1.1 shows the historic trend in TFR in Russia and Ukraine beginning in 1960. The fertility rates are reported in 5-year intervals. As evidenced from the figure, aside from a slight decline in the 1960s, there had been no major change in fertility until the 1990s. Throughout

the 1970s and 1980s fertility hovered around 2 children per woman. When the economic transition began in the early 1990s, fertility also started to dramatically descend. The fertility decline stopped in the early 2000s, and the trend reversed. The fertility growth achieved in the 2000s was not as rapid as the decline in the previous decade but it was still substantial.

1.2.2 Labor market transformation after the fall of the Soviet Union

The U-shape of the fertility trends in Russia and Ukraine since the fall of the Soviet Union matches the inverse U-shape of the unemployment trends in these countries (Figures 1.2 and 1.3). From levels of virtually zero, unemployment rates in both countries reached double digits in the late 1990s. The peak of the unemployment rate roughly coincides with the trough of the fertility rate in each country.

After the fall of the Soviet Union, wage uncertainty in Russia and Ukraine manifested itself not only in greater probability of job loss, but also in delayed wage payment, i.e. wage arrears, and in forced unpaid or partially paid leave (Figures 1.4 and 1.5). Peaking in 1998, wage arrears amounted to more than 70 million average monthly wages in Russia and more than 40 million average monthly wages in Ukraine. The scale of Figure 1.4 makes it difficult to see, but the recent financial crisis of 2009 put a stop to the decline in wage arrears which had been uninterrupted since the early 2000s. In 2009, wage arrears increased by nearly 60% in Russia and 20% in Ukraine (from a relatively small base). The incidence of compulsory unpaid or partially paid leave has followed a similar trend to that of wage arrears and also peaked in the late 1990s. At its peak, compulsory leave affected about 13% and 15% of employees in medium and large enterprises in Russia and Ukraine respectively. In another similarity to wage arrears, the incidence of compulsory unpaid leave rose during the 2009 financial crisis as well.

In the aftermath of the Soviet Union, the profile of the employed changed as well. The share of state employment has been consistently falling in both countries, and the latest available figures from 2010 show that around 30-35% of employed individuals still work for state-owned companies (Figure 1.6).

Unfortunately there are no official data available that decomposes the trends of fertility or wage uncertainty by job sector. Instead, I have to rely on the RLMS and ULMS data, which I introduce in more detail in Section 1.4.

1.3 The model

To identify the effect of wage uncertainty on fertility, I develop a dynamic model of women's joint decisions about where to work and when to have children. The model builds upon the large literature on dynamic female labor supply with endogenous work experience, which dates back to Eckstein and Wolpin (1989). It contributes to the much smaller literature that examines the *joint* relationship between female labor supply and fertility in a structural, forward-looking model with endogenous work experience. To the best of my knowledge, this is the first study to examine this relationship in developing or transition economies, where wage uncertainty plays an important role.

I focus on the job sector and fertility choices of women only and assume that their partners' job sector choices are realized after women have already made their own choices about fertility and job sector. This assumption, common in the literature on dynamic female labor supply, is prompted by the fact that increasing the size of the choice set increases the dimensions of the integrations that are necessary to solve the dynamic problem. Following the literature (for example, Van der Klaauw 1996, Francesconi 2002), I let partner's expected wage enter the decision process as a function of a woman's own characteristics.

The model captures the behavior of a young woman who is either officially married or cohabitating with a partner. In each year t over her childbearing years, a woman i decides on the job sector j_{it} in which to work ($j_{it} = \{$private sector (p),state sector (s), or not working $(nw)\}$) and whether to have a baby b_{it} ($b_{it} = \{0,1\}$).

1.3.1 Preferences

A woman derives utility from consumption, leisure, and children. Whereas consumption and leisure both depend on the choice of job sector (due to different wages and hours of work across sectors), the utility from having children is independent of job sector. Following the literature on fertility choice (for example, de la Croix and Doepke (2003) and Jones et al. (2008)), I assume that utility is separable in consumption and fertility. Period utility is

$$U(c_{it}^{j}, j_{it}, n_{it}) = u(c_{it}^{j}) + \delta^{p}\mathbf{1}[j_{it} = p] + \delta^{s}\mathbf{1}[j_{it} = s] + \varphi(n_{it}) \tag{1.1}$$

where c_{it}^{j} denotes consumption when working in sector j at time t, δ^{p} and δ^{s} denote disutility from working in the private and the state sector respectively ($\delta^{p} < 0$, $\delta^{s} < 0$), $\mathbf{1}[.]$ is the indicator function, and n_{it} denotes the number of children at time t.

Note that the choice of job sector effectively determines leisure. This assumption reflects the inflexibility of labor markets in Russia and Ukraine where working hours are predominantly controlled by the employer. I do not impose any a priori restrictions on the difference in disutilities between private and sector jobs. However, I expect the disutility from working in a private job to be larger in absolute value than the disutililty from working in a state job due to longer working hours in the former.[7]

Conception at time t is governed by an exogenous probability f_t:

$$\begin{aligned} n_{it} &= b_{it} + n_{it-1} \quad \text{with probability } f_t \\ n_{it} &= n_{it-1} \qquad\quad \text{otherwise.} \end{aligned} \tag{1.2}$$

The utility from consumption is constant relative risk aversion (CRRA):

[7] Both Russian and Ukrainian women in state jobs work significantly shorter hours than women in private jobs. In Ukraine, women's average work week in the state sector is 38.5 hours versus 44.7 hours in the private sector. In Russia, women's average work week in the state sector is 38.5 hours versus 42.4 hours in the private sector.

$$u(c_{it}^j) = \frac{(c_{it}^j)^{1-\theta}}{1-\theta} \qquad \text{where } \theta \geq 0. \tag{1.3}$$

Utility from children is:

$$\varphi(n_{it}) = (\phi + \varepsilon_{it}^n)\frac{(n_{it})^{1-\tau}}{1-\tau} \tag{1.4}$$

where $\phi > 0, \tau \geq 0$ and ε_{it}^n denotes a random taste shock.

1.3.2 Budget constraints

In this subsection, I describe the child costs, the wage processes in each job sector and the budget constraints when women receive their wages and when they do not.

Child costs

If a woman decides to have a baby, she incurs a child cost of z_r^j which is paid every period t after the child's birth. The cost depends on the geographical region r and the woman's job sector j. I model this dependence as a function of the mother's working hours and the provision of fringe benefits (vacation leave and sick leave).[8] My measure for child costs is higher in the private sector ($z_r^p > z_t^s$), because of longer working hours and lower incidence of fringe benefits in that sector.[9] If

[8]Note that in this model there is no choice over the "quality" of children. This is not only a simplification but largely reflects the reality of transition economies where the overwhelming majority of children go to free public schools. In both the Russian and Ukrainian data, expenditure on children for school fees is minimal. Moreover, conditional on total household income, the overall reported expenditure on children does not differ across job sectors. Results available upon request.

[9]Under the labor code in both Russia and Ukraine, paid vacation leave and sick leave are benefits entitled to all employees regardless of their job sectors. In practice, however, both type of benefits are much less common in the private sector. In Russia, only 78% (72%) of women in the private sector report being given paid vacation leave (sick leave) by their employers versus 99% (99%) of women in the state sector. In Ukraine, the numbers are even lower - only 49% (46%) of women in the private sector report being given paid vacation leave (sick leave) versus 91% (89%) of women in the state sector. See Appendix A for more details on how I construct the measures of child costs from the data.

a woman is not working, the child cost is normalized to 0: $z_r^{nw} = 0$.

By including these differences in child costs by job sector, the model allows for testing competing reasons for why women may prefer to have children in one sector over another. One reason might be these differences in child costs, whereas another reason might be differences in risk of wage loss, which is described below.

Wage uncertainty

When a woman decides which job sector to work in, she faces uncertainty about her monthly wage payment. Wages in Russia and Ukraine are paid monthly and there is wage uncertainty over the payment of each monthly wage throughout the year. I model the annual wage W_{it}^j as a random variable with a discrete probability distribution. With probability π_{rt}^j, a woman working in sector j in region r at year t receives her monthly wage w_{it}^j throughout the year: $W_{it}^j = 12\ w_{it}^j$; with probability $1 - \pi_{rt}^j$, she fails to receive at least one of her monthly wages. In sector nw, the distribution is degenerate: $w_{it}^{nw} = 0$ and $\pi_{rt}^{nw} = 1$. I let ρ_m^j denote the probability that a woman working in sector j receives exactly m $(m = 0, 1, 2, 3, ..., 11)$ of her monthly wages by the end of the year (conditional on her failing to receive at least one monthly wage).[10]

Note that this type of wage uncertainty is not a mean-preserving spread. The probability of experiencing wage loss affects both the first and the second moment of the wage distribution. An increase in the risk of wage loss reduces the mean and increases the variance of the wage distribution, except for very extreme probabilities. In Appendix 1.B I derive the exact relationship between the probability π_{rt}^j and the mean and variance of the wage distribution.

[10] I assume that the probability ρ_m^j, for $m = 0, 1, 2, ..., 11$, i.e. the *distribution of duration* of wage uncertainty, does not vary across geographical regions or time, whereas the incidence of wage uncertainty π_{rt}^j does. This assumption is made for two reasons. The first reason is to simplify the estimation of the model. The second reason is that the sample size is not large enough to estimate each of these 12 parameters separately by geographical region and time period, in addition to job sector. Moreover, this assumption is not obviously violated in the data. In both Russia and Ukraine the average duration of wage non-payment stays roughly stable across years. In any given year, the average duration of wage arrears is in the range of 2-3 months, the average duration of compulsory unpaid leave is in the range of 1-2 months, and the average duration of unemployment is in the range of 5-6 months.

Wages

Conditional on receiving a monthly wage, the wage depends on individual characteristics (education and experience in each job sector), and on community characteristics (geographical region and urbanization level):

$$\begin{aligned} \ln w_{it}^j = \alpha_0^j + \alpha_1^j X_{it}^s + \alpha_2^j (X_{it}^s)^2 + \alpha_3^j X_{it}^p + \alpha_4^j (X_{it}^p)^2 \\ + \mathbf{E}_i^j \prime \alpha_e^j + \mathbf{R}_i^j \prime \alpha_r^j + \alpha_u^j U_i^j + \varepsilon_{it}^j \quad \text{for } j = p, s, \end{aligned} \tag{1.5}$$

where X_{it}^j is experience (in years) working in sector $j = \{p, s\}$, $\mathbf{E}_i^j$ is a vector of education categories - high school only and college or higher (with less than high school being the omitted category), $\mathbf{R}_i^j$ is a vector of geographical region indicators (8 regions in total in Russia and 4 in Ukraine), U_i^j is an indicator for an urban population center and ε_{it}^j is a random wage shock.

Experience evolves based on past job sector choices:

$$\begin{aligned} X_{it}^p &= X_{it-1}^p + \mathbf{1}[j_{it} = p] - \eta(1 - \mathbf{1}[j_{it} = p]) \\ X_{it}^s &= X_{it-1}^s + \mathbf{1}[j_{it} = s] - \eta(1 - \mathbf{1}[j_{it} = s]), \end{aligned} \tag{1.6}$$

where η is the rate at which sector experience depreciates if a woman leaves the given sector.

From a woman's perspective, her partner's annual earnings are not realized until after she makes her choices of job sector and fertility. Hence, a woman forms her expectations about her husband's potential earnings based on her own demographics:

$$\bar{y}_{it}^h = E_i \prime \alpha_e^h + R_i \prime \alpha_r^h + \alpha_u^h U_i + \lambda_t, \tag{1.7}$$

where h denotes husband or partner and λ_t is an age indicator.

Shocks

The vector of shocks is $\varepsilon_{it} = \{\varepsilon_{it}^p, \varepsilon_{it}^s, \varepsilon_{it}^n\}$,where the first two correspond to wage shocks in the private and state sectors, and the third is a taste shock for children. I assume no serial correlation and joint normal distribution $\varepsilon_{it} \sim N(0, \Sigma)$. I allow for covariance between the wage shocks in the two sectors ($\sigma_{\varepsilon^s \varepsilon^p}$) but the taste shock for children is independent of either wage shock.

Budget constraints

Assuming that there is no intertemporal borrowing or lending, household expenditures in year t are constrained by the following equation:

$$c_{it}^j e_{it} + z_r^j n_{it} = W_{it}^j + \bar{y}_{it}^h \tag{1.8}$$

where e_{it} is the adult-equivalent household size ($e_{it} = 1.7 + .5n_{it}$), using the OECD equivalent scale (OECD 1982).[11]

The assumption of no borrowing or lending is not unrealistic in the context of the 1990s when financial markets were undeveloped and prolonged periods of inflation made cash savings a useless insurance. Admittedly, the assumption is more problematic in the context of the 2000s.[12]

[11]Note that this specification assumes that child costs do not vary with a child's age. It would be more realistic to have a cost function that decreases with a child's age. However, this would significantly increase the dimensionality of the state space that needs to be computed. The size of the state space will no longer depend only on the number of children in every period t, but also on all possible sequences of children's ages. Age-dependent child costs have been used in the literature (e.g. Hotz and Miller 1988), but only in models which include other simplifying assumptions, such as no endogenous work experience. Given that endogenous work experience is a key aspect of the dynamic decisions over job sector, I have opted to simplify the child cost function instead.

[12]Even in the late 2000s a substantial majority of households report zero annual saving and borrowing. In 2007, the last ULMS round, only 18% of Ukrainian households report non-zero saving and only 14% of Ukrainian households report non-zero borrowing during the past 12 months. In 2010, the last RLMS round, only 17% of Russian households report non-zero saving and only 6% of Russian households report non-zero borrowing during the past month. (The RLMS questionnaire only asks about saving and borrowing for the past month.)

1.3.3 The woman's planning problem

The state space is $\Omega_t = \{n_{t-1}, X^s_{t-1}, X^p_{t-1}, \mathbf{E}, \mathbf{R}, U, \varepsilon_t, z^j_r, \pi^j_{rt}, \rho^j_m, f_t\}$. At time t, the problem becomes

$$V_t(\Omega_t) = \max_{j_t, b_t}\{U(j_t, b_t, \Omega_t) + \beta f_t E_t V_{t+1}(\Omega^f_{t+1}|\Omega_t, j_t, b_t) + \beta(1-f_t)E_t V_{t+1}(\Omega^{nf}_{t+1}|\Omega_t, j_t, b_t)\} \tag{1.9}$$

where β is the discount factor. Ω^f_{t+1}(or Ω^{nf}_{t+1}) is the state space in the next period if the woman is able (or not able) to conceive a child:

$$\begin{aligned}\Omega^f_{t+1} &= \{n_{t-1} + b_t, X^s_t, X^p_t, E, R, U, \varepsilon_{t+1}, z^j_r, \pi^j_{rt+1}, \rho^j_m, f_{t+1}\} \\ \Omega^{nf}_{t+1} &= \{n_{t-1}, X^s_t, X^p_t, E, R, U, \varepsilon_{t+1}, z^j_r, \pi^j_{rt+1}, \rho^j_m, f_{t+1}\}.\end{aligned} \tag{1.10}$$

1.3.4 Expectations about uncertainty

In solving the model, I assume that individuals optimize by expecting current period uncertainty π^j_{rt} to persist in the future, i.e. individuals set $\pi^j_{rt+1} = \pi^j_{rt+2} = \pi^j_{rt+3} = ...\pi^j_{rT} \equiv \pi^j_{rt}$. However, when the next period $t+1$ comes and if uncertainty turns out to be different, individuals re-optimize and set the latest observed level of uncertainty π^j_{rt+1} as their new expectations for the future.

The assumption of static expectations is meant to reflect the inability of employees to predict the future evolution of wage uncertainty in an environment with nonlinear trends in π^j_{rt}. Another reason for the unpredictability of future risk of wage non-payment is that this risk is coming from the employers. Employers are the ones who decide when, and if, to pay wages to their employees. Given this context, I consider static expectations more plausible than rational expectations.[13]

[13]Of course, the assumption of static expectations is not the only alternative to rational expectations. Another plausible form of expectations is adaptive expectations, where current expectations are based on previous expectations and the deviation of those expectations from past realizations of wage uncertainty. Such alternatives are beyond the present scope of this paper.

1.3.5 Problem solution

The problem does not have a closed form solution, and I use backwards induction to numerically solve the dynamic problem. The terminal condition is that the continuation value function after the last period of childbearing T is zero: $V_{T+1} = 0$. To evaluate the continuation value functions in every period t before the last, I use the sparse-grid methods for numerical integration of multivariate functions, which are standard methods in the literature.

1.3.6 Implications of the model

The model has two main observable implications for the effects of wage uncertainty on the job sector and fertility decisions. The first implication is about the timing of fertility by job sector. Note that when a private sector job has a lower likelihood of wage payment (e.g. $\pi^p_{rt} < \pi^s_{rt}$), deciding to have one more child ($b_{it} = 1$) in that sector is relatively more expensive. This arises because the child cost needs to be paid every period in the future regardless of whether one receives a wage or not and because individuals expect current period uncertainty to persist in the future. Because there is no intertemporal borrowing or saving, a woman cannot shift resources over time. As a result, childbearing in a job sector that has higher wage uncertainty is postponed to a later period, so that there would be fewer remaining periods over which the child costs are to be paid.

The second implication of the model is about completed fertility (n_{it}) by job sector. Whereas only current period uncertainty matters for the timing of fertility, the entire sequence of past wage uncertainty matters for completed fertility ($n_{it} = \sum_{k=1}^{k=t} b_{ik}$; completed fertility is the sum of births, and the decision for each birth depends on contemporaneous wage uncertainty.) Women who spend a greater portion of their childbearing years in jobs with higher wage uncertainty are likely to have fewer children in total than their counterparts in jobs with lower wage uncertainty.

1.4 Data and empirical analysis

1.4.1 The data

I use data from Phase II of the Russia Longitudinal Monitoring Survey (RLMS) and the Ukrainian Longitudinal Monitoring Survey (ULMS). These surveys contain panel household data from the early 1990s until the late 2000s. I exploit fifteen waves of the RLMS (between 1994 and 2010) and three waves of the ULMS (2003, 2004 and 2007) where the latter survey contains complete fertility history and retrospective information on labor market experience since 1997.

The RLMS is a panel of Russian households interviewed annually between 1994 and 2010, except for 1997 and 1999 when no interviews took place. An individual questionnaire collects detailed information on demographics and labor market participation from each adult in the household, and a household questionnaire collects information on total household income from one designated adult in the household. In addition, a community questionnaire collects information on prices and infrastructure in the population center. In the first year of the panel 8893 adults in 3975 households were interviewed to create a nationally representative sample. The same addresses are visited in each subsequent round and all household members are re-interviewed. Panel attrition is relatively modest - on average 79% of all adult household members are present and re-interviewed in the next survey round. This low level of attrition is facilitated by low levels of residential mobility and moderate non-response rates (Heeringa 1997).[14] Additionally, in each survey round, new households are added to maintain sample size and representation such that the survey can be used both as a panel and repeated cross-sections.

The ULMS is a panel data set of a nationally representative sample of individuals between the ages of 15 and 72 inclusive, who were interviewed in three waves in 2003, 2004 and 2007. In 2003, 8,537 individuals in 4,056 households were interviewed. 7,200 and 6,774 individuals from the original sample were re-interviewed in 2004 and 2007 respectively Like the RLMS, the ULMS panel is created by visiting the same household addresses. Similar to Russia, Ukraine also has low

[14]For more details on the survey design, see http://www.cpc.unc.edu/projects/rlms-hse/project.

residential mobility which has helped keep panel attrition modest. Out of the original respondents in the 2003 survey, the response rate in the 2004 follow-up is 81% and the response rate in the 2007 follow-up (from the original respondents in the 2003 survey) is 65%. These response rates also reflect the fact that individuals who reach age 73 by the time of the follow-up surveys are not re-interviewed. The survey's individual questionnaire collects detailed information on individual labor market activities and demographics such as education, marital status (and marital history) and fertility history. Similar to the Russian survey, a household questionnaire collects information on total household income. Even though the first Ukrainian survey took place in 2003, the ULMS is just as well suited as the RLMS for analyzing women's labor market activities since the fall of the Soviet Union. In particular, respondents are asked for their labor market activity at two pre-transition points in time as well as for their complete labor market history starting from 1997 until the interview date. To minimize possible recall error, the survey specifically asks about employment in 1986 and 1991 because these years coincide with momentous events, namely the Chernobyl nuclear explosion and the fall of the Soviet Union.

The main sample used in the regression analysis consists of women who are of childbearing age (defined as ages 15 to 49 inclusive) during the transition period 1994-2010 for Russia and 1997-2007 for Ukraine and who have non-missing information on education, employment and fertility history. The RLMS sample size is 11,551 women, who are observed on average for five survey rounds for a total of 60,257 observations. The ULMS sample size is 3,531 women for a total of 38,841 observations.

I define job sector as follows: an individual works in the private sector (p) if 1) the main economic activity is not working at an enterprise but instead is recorded as farming, entrepreneurship or other individual economic activity (e.g. sewing clothes, doing repairs, performing deliveries) or 2) the main economic activity is working at an enterprise in which the state is *not* an owner or co-owner. Alternatively, an individual works in the state sector (s) if the main economic activity is working at an enterprise in which the state is an owner or co-owner.[15] An individual is defined

[15]There is no clear consensus in the transition literature on whether privatized firms more closely resemble state firms or new private firms in terms of their treatment of workers. For example, in

as non-working (*nw*) if she reports that she is not working. This leaves less than 2% of employed individuals with unclassified sector.[16]

1.4.2 Sample characteristics

In Tables 1.1 and 1.2, I show selected characteristics for two cohorts of women who had reached childbearing age by 1992, the first year after the fall of the Soviet Union. The younger cohort consists of women who were 15-24 years of age in 1992. These are women whose childbearing started almost simultaneously with the start of the transition. The older cohort consists of women who were 25-40 years of age in 1992. These are women who spent some portion of their childbearing years while the Soviet Union still existed. I then subdivide these two cohorts by their labor market experience. In the odd-numbered columns are women who spent less than half of their labor market experience since 1992 in the state sector and in the even-numbered columns are women who spent more than half of their labor market experience since 1992 in the state sector. I focus on the following characteristics: indicators for college degree and high school diploma only, indicator for married, age, household income per capita, indicator for a pensioner living in the same household, total number of kids and age at first birth.

Consider the younger cohort given in columns (1) and (2) of Tables 1.1 and 1.2. In both Russia and Ukraine women who had more state sector experience during the transition have significantly higher completed fertility and significantly earlier start of childbearing. Note that in the case of Russia these women achieved higher fertility despite a significantly higher education. Also note that the higher fertility for women in column (1) cannot be attributed to a higher propensity to

terms of layoffs privatized firms have higher rates than state firms in Russia (e.g. Lehmann and Wadsworth 2000) whereas in terms of arrears, such firms have lower rates than state firms in Russia (e.g. Earle and Sabirianova 2002). Given this intermediate standing and the fact that the RLMS does not explicitly ask for original firm ownership, I classify jobs in privatized firms as private sector.

[16]This definition of firm ownership is subject to measurement error due to workers' imperfect knowledge. A factor that potentially mitigates this error is that the privatization process in both Russia and Ukraine involved citizens quite closely. Citizens were given vouchers (either free or for a small fee) against which they could acquire shares in state-owned companies.

live with extended family members who are retired and could in principle assist with childcare.[17] These differences are consistent with the implications of the model in that completed fertility is higher and delay of childbirth is shorter among women exposed to less wage uncertainty during the transition.

Among the older cohort (columns (3) and (4)) more state sector experience is also associated with higher completed fertility. This is despite the fact that state sector experience is also associated with higher education and, in the case of Russia, lower propensity to live with an extended family member who is a pensioner. Note, however, that for this older cohort, there is no difference by job-sector experience with respect to the start of childbearing in both Russia and Ukraine. This is a placebo test of the model. According to the model, only contemporaneous uncertainty should matter for the timing of fertility. Almost all women in the older cohort started childbearing during the Soviet Union when there were no private jobs and no difference in wage uncertainty across jobs.

These fertility differences by labor market experience are only suggestive. In the next sections I will try to isolate the effects of uncertainty more precisely using a regression framework and structural estimation.

1.4.3 Wage uncertainty measures

I use two measures of wage uncertainty. For the first measure, I estimate π_{rt}^{j} - the overall probability that an individual working in sector $j = p,s$ in region r receives her contracted wage for each month of the year t. This probability is a function of 1) the probability the she is not fired, q_{rt}^{j}, 2) the probability that she is not sent on compulsory unpaid leave, v_{rt}^{j}, (conditional on still being employed) and 3) the probability that she is paid on time, κ_{rt}^{j}, (conditional on not being sent on unpaid leave). Each of these three probabilities is estimated as the incidence rate reported in the survey among women working in sector j in region r at time t. Further details on the construction

[17] This finding is consistent with studies who find no relationship between informal childcare provided by grandparents and labor market participation of young women in the former Soviet Union (Aassve et al. 2012)

of the wage uncertainty measures are provided in Appendix 1.A.

Figure 1.7 and Figure 1.8 show the time trends of the different sources of wage uncertainty in Russia and Ukraine respectively. For both countries all three sources of wage uncertainty in both the state and the private sector reach their peak in the last years of the 1990s and then decrease throughout the 2000s. For both countries the incidence of job dismissal is unambiguously higher in the private sector throughout the period. The incidence of compulsory unpaid leave is also always higher in the private sector except for the late 1990s in the case of Ukraine. Wage arrears, on the other hand, differ across sectors mostly in the late 1990s and early 2000s with very little difference by sector in the second half of the 2000s. In addition, in that first half of the period wage arrears in Russia are higher in the state sector than the private sector, whereas in Ukraine it is the reverse. Nevertheless, overall wage uncertainty of the private sector is unambiguously higher than overall wage uncertainty in the state sector for both Russia and Ukraine. The gap in both countries is largest in the first half of the survey period with a gap of over 10 percentage points in Russia and a gap of close to 20 percentage points in Ukraine. In the second half of the survey period the gap closes considerably with wage uncertainty being almost the same across the two sectors, except for the time around the 2009 financial crisis when job dismissals and compulsory unpaid leave rise in the private sector.

Figures 1.9 through 1.11 show the large geographical variation in wage uncertainty within Russia and Ukraine. For example, compare the metropolitan region consisting of Moscow and St. Petersburg with Ural. In the former region, private jobs were very uncertain in the 1990s but private sector uncertainty decreased considerably during the 2000s. State jobs, on the other hand, exhibited low level of wage uncertainty throughout the entire period, and as a result, the relative gap in wage uncertainty across the two sectors closed over time. In Ural, however, state jobs were briefly more uncertain than private jobs during the end of the 1990s (due to high incidence of wage arrears). During the 2000s, however, wage uncertainty in state jobs decreased much more rapidly than wage uncertainty in private jobs. As a result, the relative gap in wage uncertainty across the two sectors in Ural increased over time. In general, the regional variation of wage uncertainty is

greater in the private than in the state sector, and private sector jobs are more uncertain in regions which are further away from the capital.

The geographical variation within Ukraine is somewhat smaller but still noticeable. In Western Ukraine the wage uncertainty of the private sector was considerably larger than the uncertainty of the state sector during the second half of the 1990s but the gap between the sectors decreased through the 2000s. In the other three regions, however, the gap remained roughly constant throughout the 1990s and 2000s with the private sector consistently exhibiting higher wage uncertainty. In general, similar to Russia, the regional variation of wage uncertainty is much greater in the private than in the state sector, and private sector jobs are more uncertain in Western and Eastern Ukraine than in North-Central Ukraine which contains the capital.

As a robustness check, I also employ an alternative measure of wage uncertainty. The RLMS community questionnaire collects annual information on closures of state enterprises in the local population center.[18] The information is coded as an indicator for whether any closures occurred during the past year. Table 1.3 shows the link between closures and the three probabilities of wage non-payment. The first row reports the difference in wage uncertainty between state and private jobs in places with no closures of state enterprises during the past year. In such places, having a state job is associated with a higher risk of wage arrears but lower risks of compulsory unpaid leave and job loss. Living in a community where there have been closures of state enterprises is not on its own associated with a higher absolute level of wage uncertainty except for a higher risk of job loss (as would be expected). The results in the third row of the table suggest that closures of state enterprises might be used as an alternative measure of relative uncertainty of state jobs. Having a state job in a place with recent closures of state enterprises exposes workers to a greater risk of wage non-payment for each of the three risk categories.

[18] Unfortunately, the ULMS does not contain such information.

1.4.4 Cross-sectional estimation

To test the empirical predictions of the model in Section 1.3, I exploit the regional and time variation in relative wage uncertainty:

$$Birth_{i,r,t+1} = \beta_1 Job^s_{i,r,t} + \beta_2(\pi^s_{r,t} - \pi^j_{r,t}) + \beta_3(Job^s_{i,r,t} * (\pi^s_{r,t} - \pi^j_{r,t})) + \gamma' Z_{i,r,t} + \delta_r + \theta_t + \varepsilon_{i,r,t} \quad (1.11)$$

where $Birth_{i,r,t+1}$ is an indicator for individual i living in region r having given birth at time $t+1$, $Job^s_{i,r,t}$ is an indicator for holding a state job (the reference category being either a private job or non-working), $(\pi^s - \pi^j)_{r,t}$ is the relative wage uncertainty in sector $j = p, nw$ relative to the state sector s, $Z_{i,r,t}$ is a vector of controls for socio-economic and demographic characteristics (linear and quadratic terms in age, educational degrees, ethnicity, marital status, annual household income per capita, number of existing children by time t, annual hours worked, indicator for receiving fringe benefits, living in a rural area), δ_r is a vector of region fixed effects and θ_t is a vector of time fixed effects. In this estimation, as well as all subsequent ones, I cluster the standard errors by survey site. For closure of state enterprises, I estimate Equation (1.11) except that $(\pi^s - \pi^j)_{r,t}$ is replaced with the indicator for closures of state enterprises in region r between time t and $t-1$.[19]

The variables included in $Z_{i,r,t}$ are purposefully selected so that they control for other work

[19] In Appendix 1.C, I provide results of a more flexible specification, where π^s_{rt} and π^p_{rt} enter the equation separately. As Table 1.13 shows, the coefficients on π^s_{rt} and π^p_{rt} are both statistically significant, opposite in sign and roughly equal in magnitude. Note that in this specification the estimated coefficients in row 1 are no longer economically meaningful. These coefficients capture the fertility difference between sectors when both sectors have 0 likelihood of wage payment ($\pi^s_{rt} = \pi^p_{rt} = 0$).

In Appendix 1.C, I also present results from a specification, in which I allow the fertility effect to differ based on the source of uncertainty: unemployment, compulsory unpaid leave and wage arrears (Table 1.14). In Russia, there is a strong positive association between fertility and a state job when state jobs have lower wage uncertainty than private jobs, regardless of the source of uncertainty. The fertility effect is somewhat stronger when the source of uncertainty is unemployment or unpaid leave than when the source is wage arrears. In Ukraine, the fertility effect is somewhat stronger when the source of uncertainty is unpaid leave or wage arrears than when the source is unemployment. Nevertheless, these differences by source of uncertainty are not large in either country.

aspects that differ across job sectors and that influence fertility decisions. These are the hours worked, the overall household income and the provision of fringe benefits on the job. Note that I do not include explicit measures for child costs z_{rt}^{j} but control for them implicitly by including the hours worked and an indicator for receiving fringe benefits. Similarly, I control for overall household income, which includes both the woman's wage and the husband's wage (if she is married).

The coefficient of main interest is the coefficient on the interaction term β_3. According to the model in Section 3, this coefficient is predicted to be positive. It is a type of "triple-difference" estimator (across sectors, across regions and across time) of the effect of uncertainty on fertility. I consider it a cleaner measure of the effect of uncertainty than the main effects β_1 and β_2. β_1 might plausibly reflect unobservable heterogeneity that makes some women prefer both having state jobs and having children. Similarly, β_2 might reflect other regional characteristics that are correlated with the inter-sector gap in uncertainty and that also encourage fertility.

Table 1.4 shows the results from the cross-sectional estimation where uncertainty is measured as the probability of receiving one's wage. As discussed above, the effect of uncertainty is captured by the coefficient on the interaction term between the indicator for a state job and the measure of uncertainty of a state job relative to a private job (columns (1) and (3)) or relative to not working (columns (2) and (4)). The first row in the table gives the fertility difference across sectors when there is no difference in relative uncertainty. When there is no gap in relative uncertainty, having a state job is associated with only slightly higher fertility than having a private job or having no job. When the state job is in a region and at a time period where state jobs are relatively less uncertain, there is a strong positive association between fertility and having a state job. In comparison to the private sector, a 1 percentage point increase in the relative likelihood of wage payment in the state sector is associated with 6 percentage points (5 percentage points) increase in the relative likelihood of having a baby in the state sector in Russia (Ukraine). The effect is only slightly lower when the state sector is compared with non-working.

This result is robust to the alternative measure of wage uncertainty, local closures of state enterprises (Table 1.5). Note that with this definition of relative uncertainty, a higher measure

implies a higher relative uncertainty of state jobs and so the interaction term in row (3) (as well as the main effect in row (2)) should have an opposite sign from the interaction term in Table 1.4. In line with the other proxy of uncertainty, this measure also shows a rather strong relationship with fertility. Holding a state job in a place where state firms recently closed down reduces the relative likelihood of having a baby in the state versus the private or non-working sector by 2 percentage points.

1.4.5 Panel estimation

Even though the coefficient β_3 in Equation (1.11) is likely a better estimate of the relationship between wage uncertainty and fertility than either β_1 or β_2, it is still only evidence of association. Indeed, in a cross-section, it is almost impossible to show that the relationship between wage uncertainty and fertility is not driven by unobserved heterogeneity. However, if the unobserved heterogeneity is time-invariant, it does not affect a time-series relationship between a woman's job sector choice and her fertility behavior.

To test this hypothesis, I examine the fertility behavior of a sub-sample of women, all of whom work initially in the private sector but some later switch to the state sector. According to the model in Section 1.3, women who switch to the state sector should be more likely to have children than women who remain in the private sector. In addition, these fertility differences should emerge only after the job sector switch. Any fertility differences prior to the switch would indicate that there is some unobserved heterogeneity between 'stayers' and 'switchers.' In such a case, the group of 'stayers' would not be a proper counterfactual for the group of 'switchers.'

Table 1.6 displays the results of this analysis. The first four columns show the fertility difference between "switchers" and "stayers" in the 3 years leading up to and including the year of the sector switch. In both Russia and Ukraine, the two groups of women have the same number of children prior to the sector switch. Yet, starting one year after the switch, the "switchers" have on average 0.1 more children than the "stayers."

Table 1.7 shows the same exercise but now I include an interaction between the gap in inter-

sector uncertainty and an indicator for the sector switch. The coefficient on the interaction term shows that, if the switch occurs in a region and at a time when private jobs are relatively more uncertain, then the fertility gap after switch is more pronounced. For every 1 percentage point increase in the relative wage uncertainty of private jobs, women who switch sectors have about 0.5 more children in Russia and about 0.6 more children in Ukraine in the years immediately after the switch. These results suggest that unobserved heterogeneity among women is an unlikely driver of the cross-sectional association between uncertainty and fertility. Of course, there still might be unobservable heterogeneity that varies with time based on the current job sector, but I consider this less plausible.

When interpreting these results, one should also keep in mind the role of the husband's job sector. Although not part of the main analysis, I briefly examine this issue in Section 6.

1.5 Structural estimation

The vector of structural parameters to be estimated is

$$\rho = \{a_0^j, a_1^j, a_2^j, a_3^j, a_4^j, \mathbf{a}_e^j, \mathbf{a}_r^j, a_u^j, \eta, \delta^j, \theta, \phi, \tau, \sigma_{\varepsilon^s,\varepsilon^p}, \sigma^2_{\varepsilon^j}, \sigma^2_{\varepsilon^n}\} \text{ for } j = \{p,s\}, \tag{1.12}$$

where $\{a_0^j, a_1^j, a_2^j, a_3^j, a_4^j, \mathbf{a}_e^j, \mathbf{a}_r^j, a_u^j\}$ are the wage determinants in each sector,[20] $\{\eta\}$ is the depreciation rate of experience, $\{\delta^j, \theta, \phi, \tau\}$ are the preference parameters, and $\{\sigma_{\varepsilon^s,\varepsilon^p}, \sigma^2_{\varepsilon^j}, \sigma^2_{\varepsilon^n}\}$ are the shocks. I assume that the parameters governing husband's income $\bar{y}_{it}$ and the measures of wage uncertainty π_{rt}^j, ρ_m^j and child costs z_r^j in each sector can be taken directly from the survey data. The remaining set of exogenous parameters is described in Table 1.8. Given the slightly different

[20]Note that $\mathbf{a}_e^j$ is a vector of two parameters – the sector-specific wage return to high school degree and the sector-specific wage return to college or higher degree. Similarly, $\mathbf{a}_r^j$ is a vector of the sector-specific wage returns to living in a geographical region r. In the case of Russia, I estimate 7 such parameters (8 regions in total). In the case of Ukraine, I estimate 3 such parameters (4 regions in total.)

timing of the RLMS and ULMS, I estimate the model on the 1973-1976 cohort of Russian women and on the 1976-1979 cohort of Ukrainian women. As the model does not address the process of family formation, I restrict the sample to women who are either officially married or cohabitating with partners throughout the period. The optimization process begins at initial age of 18 to 21 years for all individuals in the sample. In the Russian case, a period t corresponds to one of the 15 survey waves carried out between 1994 and 2010; as a reminder, surveys were annually conducted except in 1997 and 1999. In the Ukrainian case, a period t corresponds to 1 of the 11 years between 1997 and 2007, which is the period when the survey collected uninterrupted labor market history. The discount factor β is set to 0.95.[21] The values used for f_t, the probability of conception at age t, are given in the last four rows of Table 1.8. These are values taken from demography studies on the relationship between age and infertility (Menken et al. 1986).

1.5.1 Estimation method

I estimate the vector of parameters ρ by indirect inference (Gourieroux et al. 1993) in which I use the following set of auxiliary parameters. Note that ρ is identified by the *entire* set of auxiliary parameters, although there is a close correspondence between some of the parameters of the structural model and the auxiliary one. In particular, to identify the parameters governing the wage processes $a_0^j, a_1^j, a_2^j, a_3^j, a_4^j, \mathbf{a}_e^j, \mathbf{a}_r^j, a_u^j, \eta, \sigma_{\varepsilon^j}^2$ I use as auxiliary parameters the coefficients and residuals from OLS regressions of log wages by sector. To identify $\sigma_{\varepsilon^s,\varepsilon^p}$ specifically, I use the OLS residuals from the subsample of individuals who switch between the two sectors. To identify the parameters governing preferences for children ϕ, τ and $\sigma_{\varepsilon^n}^2$, I use as auxiliary parameters the coefficients and residuals from Poisson regressions of the number of children on age. Lastly, to identify the parameters for leisure and risk preferences δ^j and θ, I use the average proportion employed in each sector by age and the average number of children in each sector by age.

From the data I calculate the vector **m** of these auxiliary parameters. I simulate the model 30

[21] The discount factor is not well identified in such structural models and is usually exogenously set.

times to create a vector of simulated parameters $\widetilde{\mathbf{m}}_s(\rho)$ for each simulation $s = 1, \ldots, 30$.Then I find $\widehat{\rho}$ that solves

$$\widehat{\rho} = \arg\min_{\rho}(\mathbf{m} - \frac{1}{30}\sum_{s=1}^{30}\widetilde{\mathbf{m}}_s(\rho))\prime W^{-1}(\mathbf{m} - \frac{1}{30}\sum_{s=1}^{30}\widetilde{\mathbf{m}}_s(\rho)) \tag{1.13}$$

where the weighting matrix W is a diagonal matrix, whose elements are the variances of the auxiliary parameters $\mathbf{m}$ which I estimate from the data.[22]

1.5.2 Results

Table 1.9 presents the results from the structural estimation of the parameters. Consistent with the fact that the two countries share very similar patterns in fertility and wage uncertainty, the estimates based on Russian and Ukrainian data closely resemble one another. The difference in disutility of working in the private versus the state sector is larger in Ukraine than in Russia. This finding aligns with the fact that Ukraine has a larger difference in working hours and fringe benefit provision between the sectors.

With regards to the determinants of wages, the return to a college degree is higher in the state sector whereas the return to a high school degree is higher in the private sector in both countries. This is consistent with the educational patterns by job sector experience that are presented in Tables 1.1 and 1.2: women with more state sector experience are more likely to have a college degree whereas women with more private sector experience are more likely to have a high school degree only. With regards to previous experience in the same sector, the state sector has a slightly flatter experience-wage profile than the private sector. With regards to previous experience in the other sector, only the state sector gives significant returns to previous experience in the other sector. Previous private sector experience pays if one switches to the state sector, but previous state sector experience does not pay if one switches to the private sector. Lastly, the return to living in an urban population center is higher in the private sector than in the state sector. This difference in

[22]I do not use the full variance-covariance matrix of the auxiliary parameters because it can be biased in small sample sizes (Altonji and Segal 1996).

returns across the two sectors is especially large in Russia: the return to living in an urban center for private sector employees is almost double the return for state sector employees.

The model fits well the trends of employment by sector and the trends of fertility by sector (Figures $1.12 - 1.17$). The figures are drawn with calendar year on the horizontal axis, but the time trends displayed can alternatively be interpreted as age trends given that I estimate the model on one cohort of women. Note that the model matches the observed difference in the timing of fertility across sectors. The simulated gap in fertility between the two sectors is largest in the first couple of years because women in the state sector start childbearing earlier than their private-sector counterparts. Fertility growth in the non-work sector is approximately constant over time, consistent with the fact that there is no change in uncertainty over time in that sector.

Figures 1.18 and 1.19 illustrate the model fit by simulating the dynamics of the fertility response to switching sectors. The model predicts that women who switch to the state sector have on average 0.10 to 0.18 more children in the years immediately after the switch. These findings are in accord with the estimates taken from the actual data, presented in Table 1.6.

1.5.3 Counterfactual exercises

Having estimated the structural parameters of the model, I now analyze how job sector and fertility behavior adjust under two counterfactual scenarios for wage uncertainty and a counterfactual scenario for a fertility policy with more generous maternity leave.[23]

For the first counterfactual scenario, I equalize the wage uncertainty in the private sector to what is actually observed in the state sector for the same time period and geographical region ($\pi^p_{rt} \equiv \pi^p_{rt}$). The effects of this change are illustrated in Figures $1.20 - 1.23$. Eliminating the extra uncertainty in the private over the state sector affects both employment and fertility. It shifts some

[23]The model developed in this paper assumes that individuals have no means of borrowing or saving, an assumption which reflects the undeveloped credit markets in Russia and Ukraine. This assumption underlies the results of the counterfactual exercises presented here. Children would be less costly in the presence of wage uncertainty, if young adults had greater means of smoothing their consumption over time.

female workers away from the state into the private sector and increases fertility in the private sector to almost the same levels as fertility in the state sector. Moreover, notice that this counterfactual exercise produces larger effects in Ukraine than in Russia because the actual gap in wage uncertainty between sectors is larger in Ukraine. Columns 1 and 2 in Table 1.10 summarize the fertility effects of this exercise. Eliminating the difference in wage uncertainty between the two sectors closes the fertility gap between the sectors by about 51% in Russia, and by about 57% in Ukraine.

For the second counterfactual scenario, I change the wage uncertainty in each sector to equal the latest level observed in that sector ($\pi^p_{rt} \equiv \pi^p_{rT}, \pi^s_{rt} \equiv \pi^s_{rT}$). In effect, this hypothetical situation eliminates the high wage uncertainty levels of the late 1990s and replaces them with the much lower uncertainty levels of the late 2000s. The fertility results are illustrated in Figures 1.24 and 1.25. As this exercise reduces uncertainty in both sectors, counterfactual fertility is also higher than actual fertility in both sectors. Nevertheless, the effects are larger for the private sector which is consistent with that sector having experienced a greater decline in uncertainty from the late 1990s to the late 2000s. In addition, this counterfactual scenario produces a small fertility gap between the two sectors in Ukraine and reflects the small gap in uncertainty observed in Ukraine during the latest survey year. Columns 3 and 4 in Table 1.10 summarize the fertility effects of this counterfactual exercise. Had the high uncertainty levels of the 1990s been avoided in both sectors, overall fertility would have been 26% higher in Russia and 30% higher in Ukraine.

I implement a fertility policy that increases the duration of fully paid maternity leave in the third counterfactual scenario. According to current law, the duration of leave that is fully paid is 140 and 126 days in Russia and Ukraine respectively.[24] In particular, I consider a policy which increases the duration of fully paid leave by 20% in both countries. This amounts to an increase of 28 and 25.2 days in Russia and Ukraine respectively. Under this counterfactual, the probability that the policy is implemented in each sector equals the probability that maternity leave is implemented in

[24]Note that in both countries, women can take additional maternity leave that is only partially paid. In the model, I do not allow for this possibility.

each sector.[25] The fertility results are illustrated in Figures 1.26 and 1.27. As a result of this policy, fertility increases in both job sectors. However, the response is larger in the state sector, given the higher probability of actual policy implementation in that sector. As expected, this policy has smaller effects in Ukraine, where the probability of actual implementation is lower in both sectors, especially the private sector. Columns 5 and 6 in Table 1.10 summarize the fertility effects of this counterfactual exercise. A 20% increase in the duration of fully paid maternity leave, increases overall fertility by about 12% in Russia and 5% percent in Ukraine.

1.6 Discussion

In this section, I discuss other factors apart from wage uncertainty that potentially affect the fertility trends in Russia and Ukraine after the fall of the Soviet Union. Some factors are likely complementary to the wage uncertainty hypothesis; for example, wage uncertainty could influence marriage rates as well as fertility. In this section I also use information on the husband's job sector to help separate the two reasons which potentially make state jobs 'children-friendly,' e.g. lower wage uncertainty and higher fringe benefit provision.

1.6.1 Other potential factors affecting fertility

Even if the decline in fertility did not precede the economic transformation, it is still possible that it was driven by factors other than the labor market transformation. In particular, marriages also declined substantially during the course of the 1990s from 21 to 13 marriages per 1000 individuals in ages 15-44. To examine the effect of marriages on fertility, I construct a counterfactual fertility rate that holds marriages fixed at their 1989 level and use the observed rate of births per marriage during the transition period. Figures 1.28 and 1.29 display the trend of this counterfactual fertility

[25] I construct this probability in the same way as I do for the other fringe benefits, by using the observed incidence of this type of paid leave by job sector (as well as by geographical region and time period). In Russia, 96% of women in the state sector and 68% of women in the private sector report that their workplaces would allow them to take maternity leave. In Ukraine, the respective percentages are even lower: 78% in the state sector and 37% in the private.

next to the trend of actual fertility in Russia and Ukraine respectively. The fact that these two trends are so close together suggests that the change in marriage formation cannot single-handedly explain the change in fertility.

Another potential explanation for the fertility decline in the 1990s is the wider access to modern contraception that accompanied the opening up of the Russian economy. In the Soviet Union, birth control was practiced through traditional methods such as rhythm and withdrawal as well as abortions which were legal and free of charge (Popov et al. 1993). In the 1990s, modern birth control became widely available for the first time. By 1994, the first year of the RLMS survey, around 50% of women in childbearing age report using birth control, and around 49% of women who answered affirmatively cite IUD as their preferred method. This quick take-up of modern birth control suggests that modern birth control has cheaper emotional cost than abortions. Hence, it is plausible that reducing unwanted fertility became cheaper at the start of the 1990s. However, there are several problems with this explanation. First, since the mid-1990s, there have been no big changes in the availability of modern birth control. Second, the cost of an IUD (relative to the average wage) was highest in the late 1990s when fertility was lowest and the cost was declining for most of the 2000s when fertility was rising (Figure 1.30). Third and most importantly, contraception is a means of reducing unwanted fertility only. If it did cause the decrease in fertility in the 1990s, it must be that these women desired fewer children than before. A reduced desire for children is not inconsistent with the wage uncertainty hypothesis. Wider availability of modern contraception might very well be an explanation of *how* the fertility decline was actually accomplished. A more problematic story would be if women report wanting fewer children for non-economic reasons. Although the Russian survey does not contain information on fertility preferences independent from actual economic circumstances, the Ukrainian survey does ask for desired preferences under "ideal economic circumstances." Figure 1.31 shows the ideal number of children by women's birth cohort (aggregated into 5-year intervals). As the figure shows, there is no significant difference in desired number of children between women in older birth cohorts who completed their fertility in the Soviet Union and younger birth cohorts who spent at least a portion of their childbearing years

in the period after the Soviet Union fell.[26] In addition, note that the ideal number of children is greater than 2 for all birth cohorts, which is substantially higher than realized fertility in the period after 1991.

Figure 1.31 helps dismiss selective emigration from the Soviet Union as an alternative explanation for the fertility decline in the 1990s. In theory, emigration could have contributed to the fertility decline if non-emigrants had lower fertility intentions than emigrants.[27] However, desired fertility is unlikely to be correlated with migration status for older individuals. Hence, in Figure 1.31 desired fertility of older cohorts is unlikely to depend on these cohorts being non-emigrants. Desired fertility of younger cohorts might very well depend on these cohorts being non-emigrants. Nevertheless, it appears from the figure that younger non-emigrants do not have lower desired fertility than older cohorts.

1.6.2 The role of husband's job sector

In the model presented above, I focus on the role of a woman's job sector for her fertility. In this subsection, I briefly analyze the differences in fertility by job sectors of *both* partners. I restrict the analysis to couples in which both partners are employed in order to isolate the fertility effect of working in a state versus a private sector job. Such an analysis could be helpful in distinguishing between the role of higher fringe benefits and the role of lower wage uncertainty for higher fertility in the state sector.

Consider the following equation:

[26] This survey question is also asked of men. There is no statistically significant difference by gender (results available upon request).

[27] In Russia, there was substantial immigration from the other former Soviet republics, to the extent that net migration was positive throughout the 1990s. In Ukraine, there was some immigration too, but it was lower than emigration (primarily to Russia) throughout the 1990s.

$$Birth_{h,r,t+1} = \beta_0 + \beta_1 Both_State_{h,r,t} + \beta_2 Woman_State_{h,r,t} + \beta_3 Man_State_{h,r,t} + \gamma' Z_{h,r,t} + \delta_r + \theta_t + \varepsilon_{h,r,t} \quad (1.14)$$

where $Both_State_{h,r,t}$ is an indicator for a household h in which both partners have state jobs at time t, $Woman_State_{h,r,t}$ is an indicator for a couple in which only the woman works in a state job at time t (the man is in a private job) and $Man_State_{h,r,t}$ is an indicator for a couple in which only the man works in a state job at time t (the woman is in a private job). The outcome variable is an indicator for having a baby in the household at time $t+1$. The vector of controls $Z_{h,r,t}$ includes a linear and quadratic term for woman's age, indicators for her education degree, ethnicity, household income per capita, number of kids already born by time t, indicator for living in rural area and fixed effects for geographical region and survey year. Similarly to the main cross-section specification in Equation (1.11), $Z_{h,r,t}$ includes total household income per capita to control for the fact that different job sector choices might produce different income levels, in addition to different risks of wage non-payment, and that these income differences could affect fertility too.

There are two assumptions that are necessary in order to use this framework to disentangle the roles of fringe benefits and wage uncertainty. The first assumption is that only women use their jobs' fringe benefits for the purposes of child care. The second assumption is that child costs are borne out of total household resources so that the wage uncertainty of a man's job affects the ability to bear the child costs in the same way as the wage uncertainty of a woman's job. Given these assumptions, a comparison of β_2 and β_3 isolates the association between fertility and fringe benefits while holding constant total wage uncertainty in the household. Similarly, a comparison of β_1 and β_2 isolates the association between fertility and wage uncertainty of state jobs, while holding constant fringe benefits in the household.

Note that this specification is agnostic about the specific fringe benefits that might make it easier to have a child in a state job rather than in a private job. These might be observed characteristics

such as shorter hours, higher provision of paid leave, as well as unobserved characteristics such as less stressful or demanding work culture.

Table 1.11 shows the results of this estimation. All else equal, households in which both spouses work in the state sector have close to 2 percentage points higher likelihood of having a baby relative to households in which both spouses work in the private sector. Households in which only the woman works in the state sector are about 1 percentage point less likely to have a baby relative to households in which both spouses work in the state sector. Holding fringe benefits constant, having a husband in the private sector relative to a husband in the state sector increases household wage uncertainty and appears to have a significant negative effect on fertility (the first F-test shown in the table). However, holding household wage uncertainty constant, the fringe benefits of a woman's job do not have a statistically significant effect on fertility (the second F-test shown in the table).

These results are consistent with the results from the regression analysis and simulation exercises presented earlier: having a state job has a positive effect on fertility primarily because state jobs provide a lower risk of wage loss than private jobs.

1.7 Conclusion

In this chapter I show that wage uncertainty has had a significant effect on the fertility and job sector choices of Russian and Ukrainian women during the 1990s and 2000s.

I use individual- and household-level panel data to document large variation in wage uncertainty over time and across job sectors and geographical regions. The regression analysis suggests that women working in private jobs in a year in which private jobs in the region have a 1 percentage point higher likelihood of wage non-payment (relative to state jobs) have 5 to 6 percentage points lower propensity of having a baby (relative to women in state jobs) the following year. In addition, switching to a state job in which wage non-payment is 1 percentage point less likely is associated with about 0.4 to 0.6 more children in the year after the job switch.

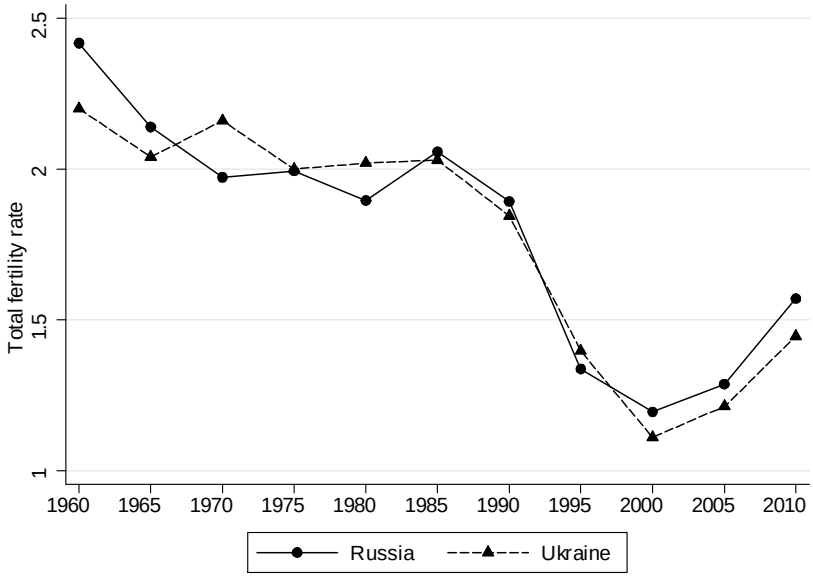
To interpret these findings, I build a stochastic dynamic model that incorporates the features of the Russian and Ukrainian labor markets. I estimate the key parameters of the model by indirect inference using detailed panel data. The model fits well the main fertility and employment trends as well as the dynamics of the fertility response to switching job sectors. I then use the model to perform counterfactual exercises with different levels of wage uncertainty. I find that the large uncertainty gap between the private and the state sector that persisted for most of the late 1990s and early 2000s has been responsible for over 51% of the fertility gap between the sectors. Additionally, had the high uncertainty levels during the late 1990s been avoided in both sectors, overall fertility would have been 26% to 30% higher.

Policymakers across Eastern Europe focus on stimulating fertility through more generous maternity leaves (Putin 2006). The findings in this chapter suggest that such policies are unlikely to be very effective for several reasons. As state jobs continue to constitute an ever smaller proportion of employment, the effects of more generous child care provisions depend primarily on the level of labor code enforcement in the private sector. Enterprise data show that private sector firms are 1.6 times less likely to have had a labor inspection in the past year, and if such inspections do occur, their duration is 40% shorter in the private than the state sector, controlling for the size of the firm (World Bank Enterprise Survey 2005). The persistently low level of labor code enforcement is reflected in the data. In the last survey round, only 74% of women in the private sector in Russia report being able to take maternity leave at their jobs. The respective number in Ukraine is much lower at 40%. Given these low levels of labor code enforcement, I estimate that a policy which raises the duration of fully paid maternity leave by 20% leads to only a 12% increase in fertility in Russia and a 5% increase in fertility in Ukraine. Moreover, stricter enforcement of maternity leave might not necessarily make such fertility policies more effective. Gimpelson et al. (2009) find that stricter enforcement targeted on protection of women suppresses female employment in Russia by making female employees relatively more costly for employers.

This chapter has documented a strong link between wage uncertainty and fertility behavior. The link exists both in the cross-section and in the time-series dimension. Women's choices of job sec-

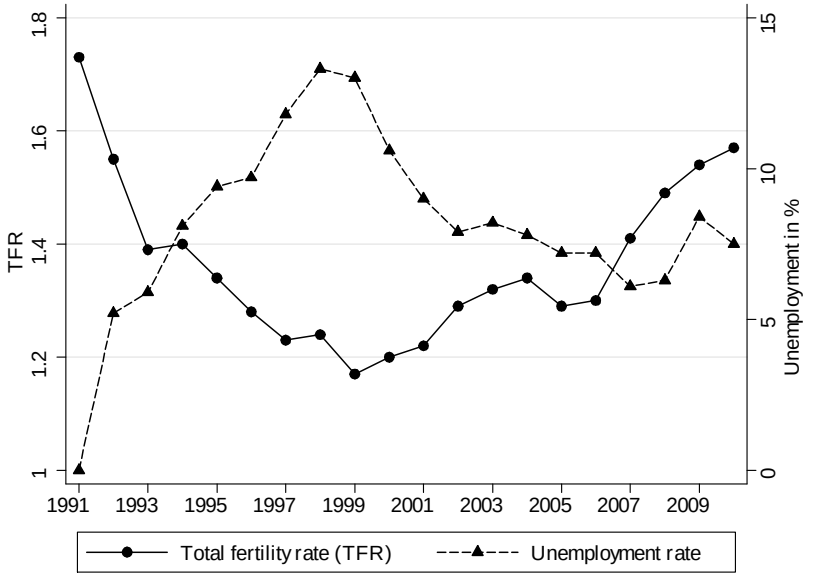
tor and fertility appear consistent with a framework in which children are thought of as irreversible investments made costlier by wage uncertainty. A considerable portion of the wage uncertainty in Russia and Ukraine is due to wage non-payment resulting from wage arrears and compulsory unpaid leave. Policies that target wage contract violations might be effective in increasing fertility rates in these transition economies.

Figure 1.1: Historical trends in fertility in Russia and Ukraine



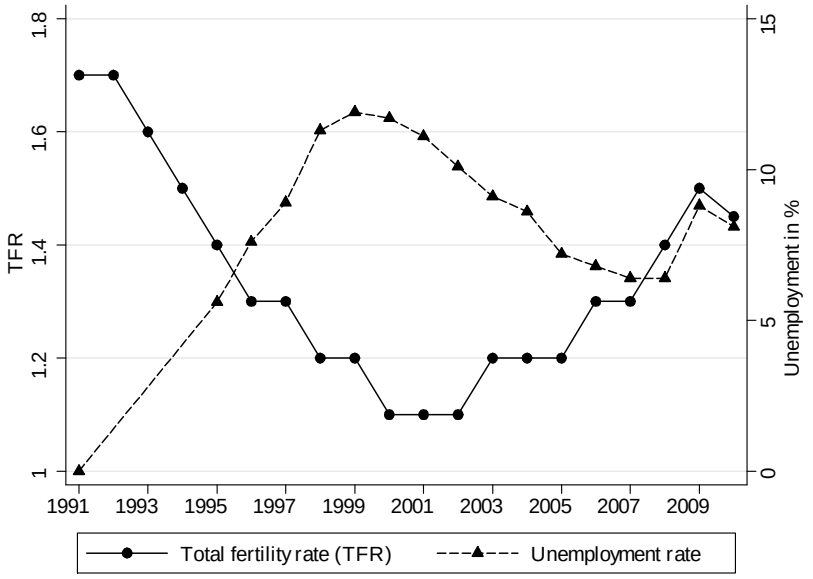
Notes: The data sources are the Russia Federal Statistics Committee and the Ukraine State Statistics Committee.

Figure 1.2: Fertility and unemployment trends in post-Soviet Union Russia



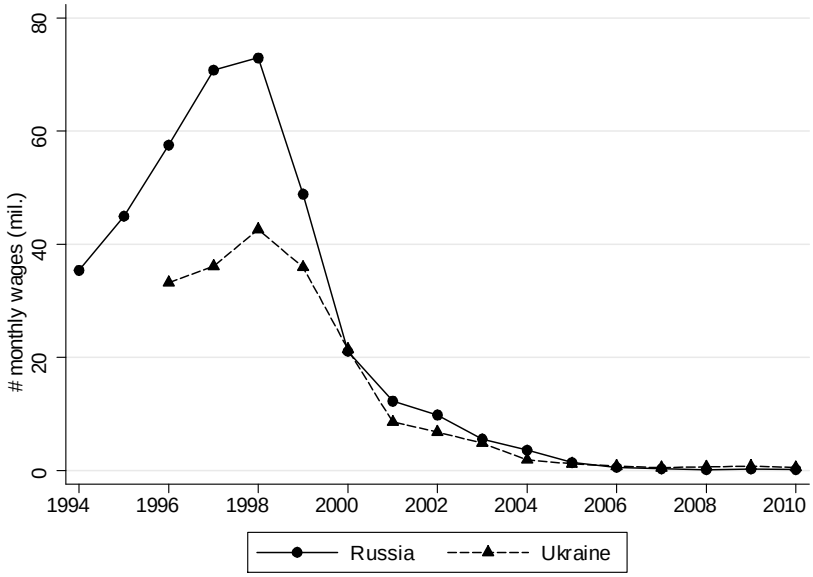
Notes: The data source is the Russia Federal Statistics Committee.

Figure 1.3: Fertility and unemployment trends in post-Soviet Union Ukraine



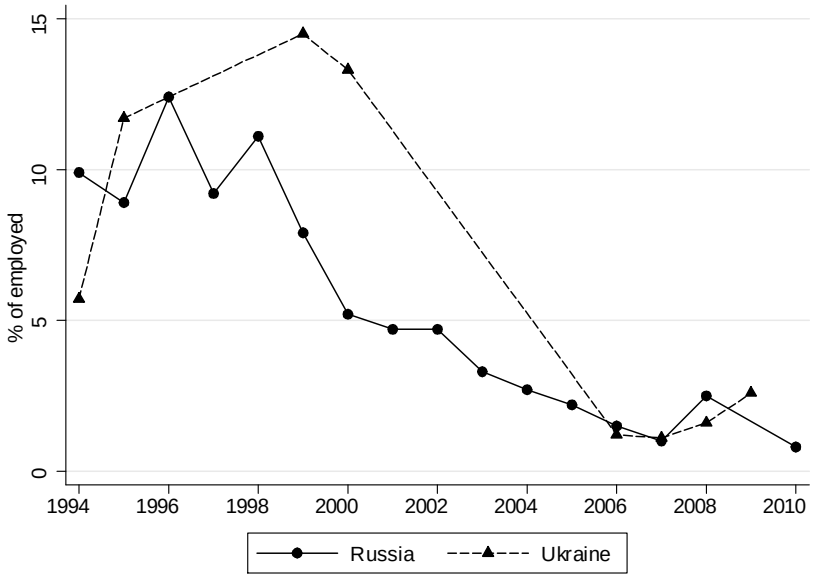
Notes: The data source is the Ukraine State Statistics Committee.

Figure 1.4: Stock of wage arrears in millions of average monthly wages



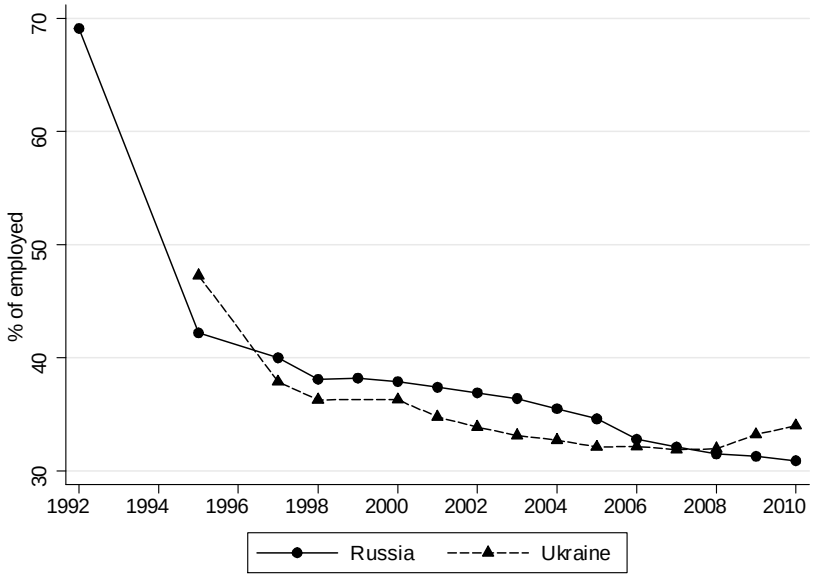
Notes: The data sources are the Russia Federal Statistics Committee and the Ukraine State Statistics Committee.

Figure 1.5: Incidence of unpaid or partially paid leave initiated by the employer



Notes: The data sources are the Russia Federal Statistics Committee and the Ukraine State Statistics Committee. Data exclude small enterprises.

Figure 1.6: Time trend in state sector employment



Notes: The data sources are the Russia Federal Statistics Committee and the Ukraine State Statistics Committee.

Figure 1.7: Time trends of sources of wage uncertainty in Russia

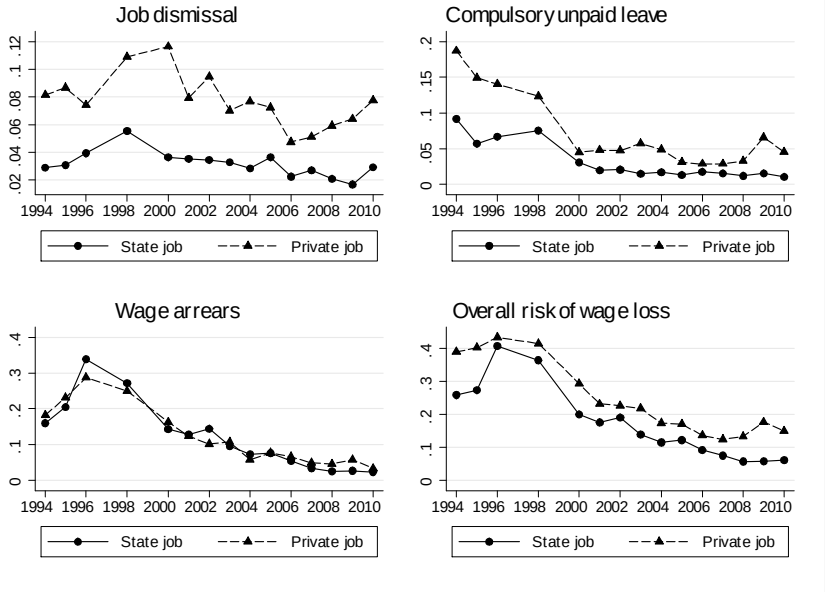
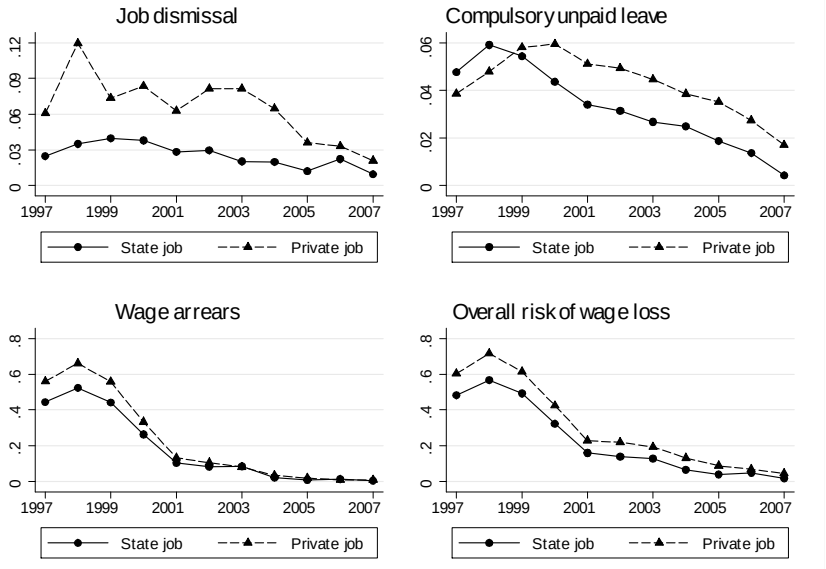


Figure 1.8: Time trends of sources of wage uncertainty in Ukraine



Notes: Displayed are the fractions of women in each sector who experience dismissal (among all employed), compulsory leave (conditional on no dismissal) or wage arrears (conditional on no unpaid leave) in a given year. Overall risk of wage loss is calculated as the likelihood of dismissal, unpaid leave or arrears.

Figure 1.9: Time trend of overall risk of wage loss by region in Russia

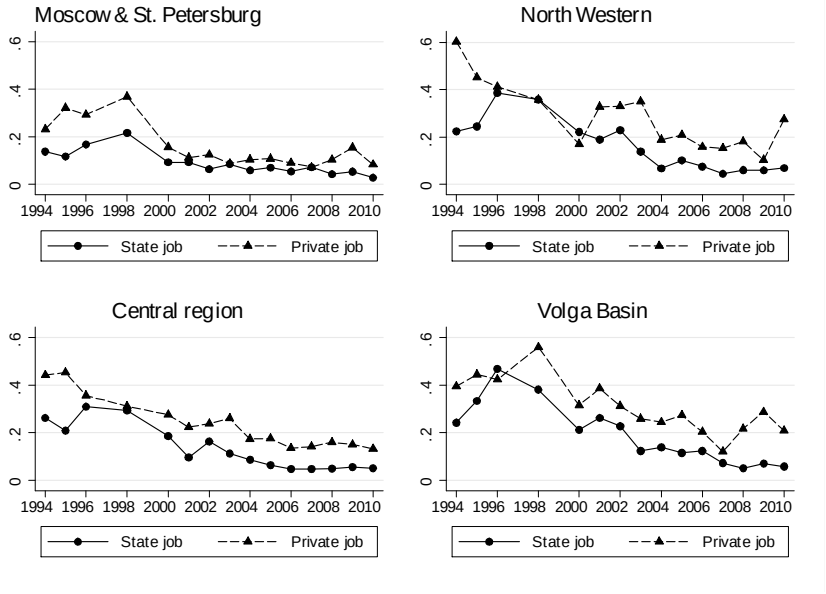


Figure 1.10: Time trend of overall risk of wage loss by region in Russia (cont.)

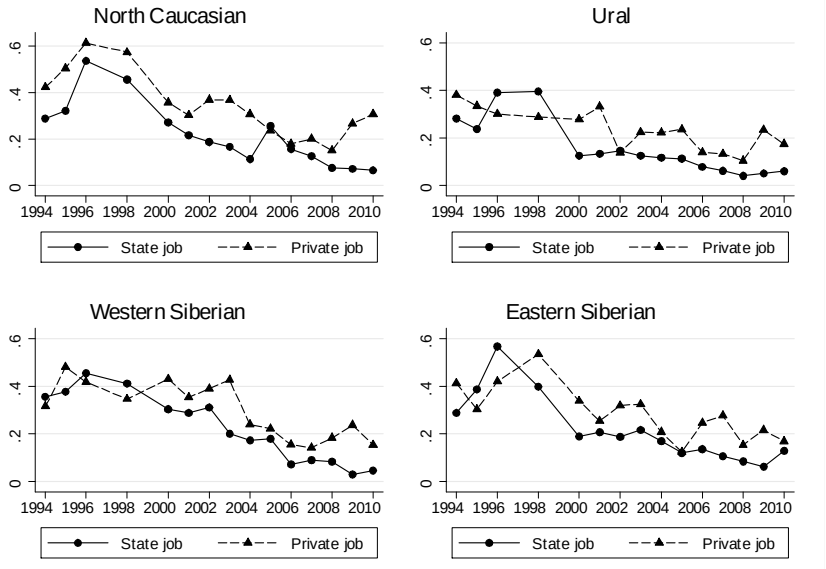


Figure 1.11: Time trend of overall risk of wage loss by region in Ukraine

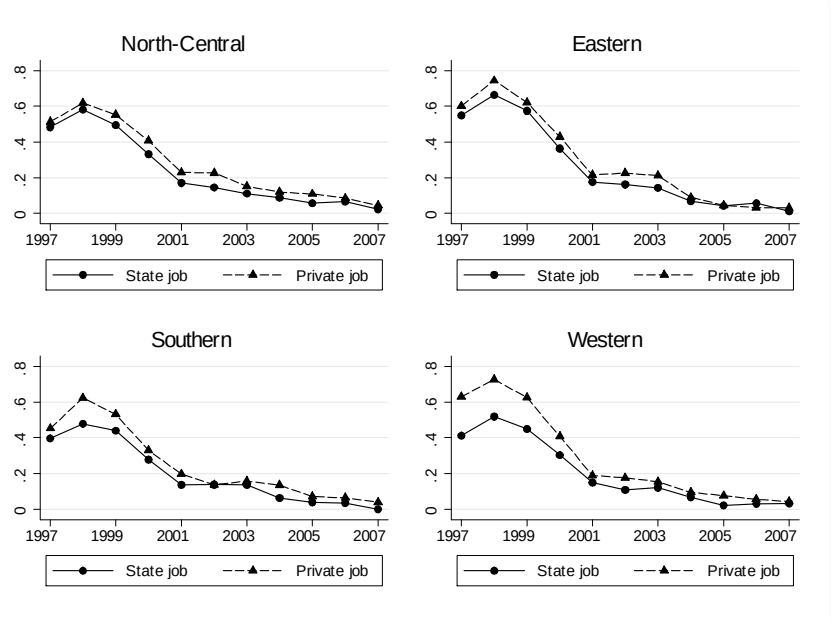


Figure 1.12: Model fit for the share employed in state jobs in Russia

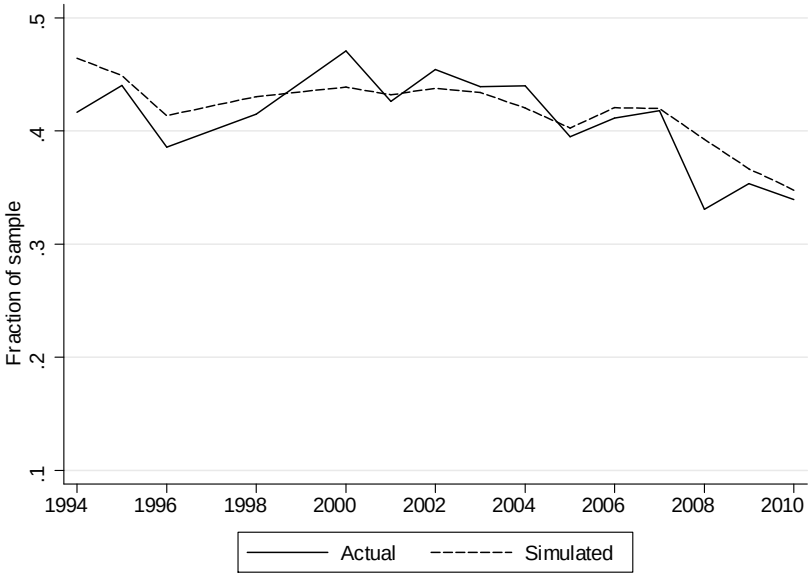


Figure 1.13: Model fit for the share employed in private jobs in Russia

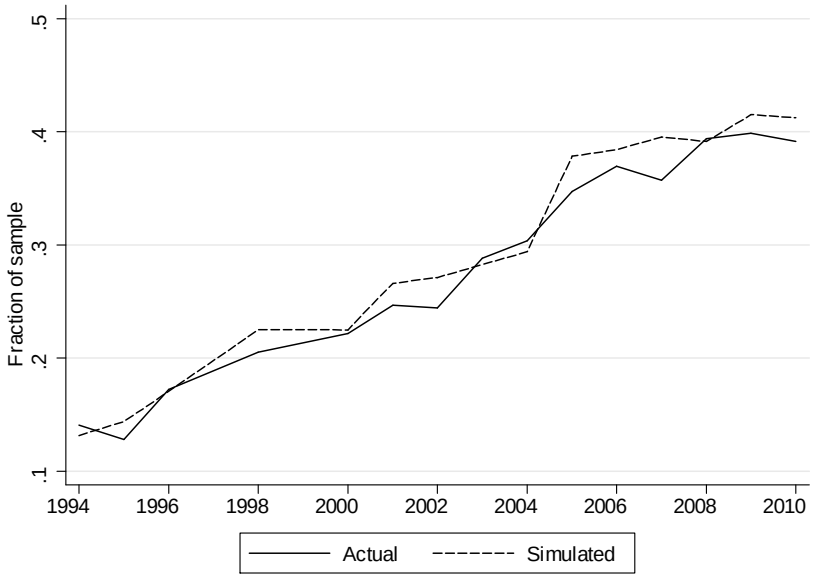


Figure 1.14: Model fit for share employed in state jobs in Ukraine

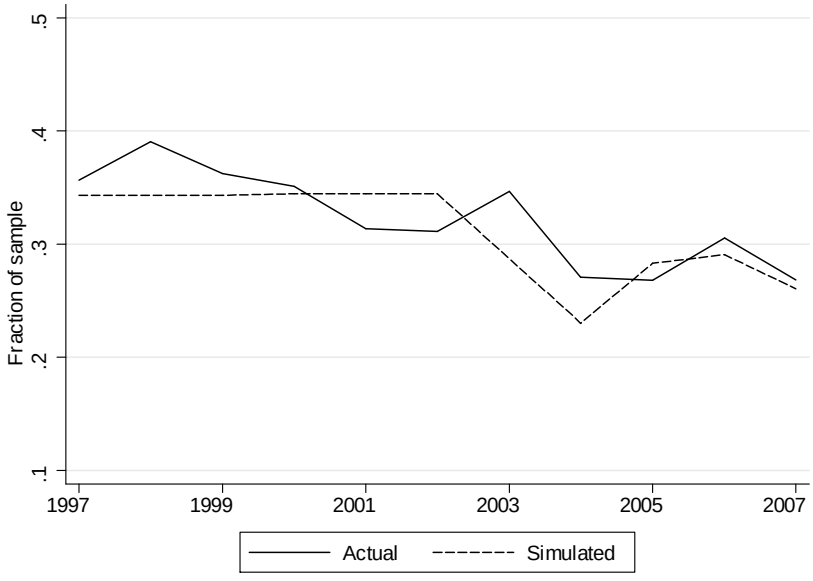


Figure 1.15: Model fit for share employed in private jobs in Ukraine

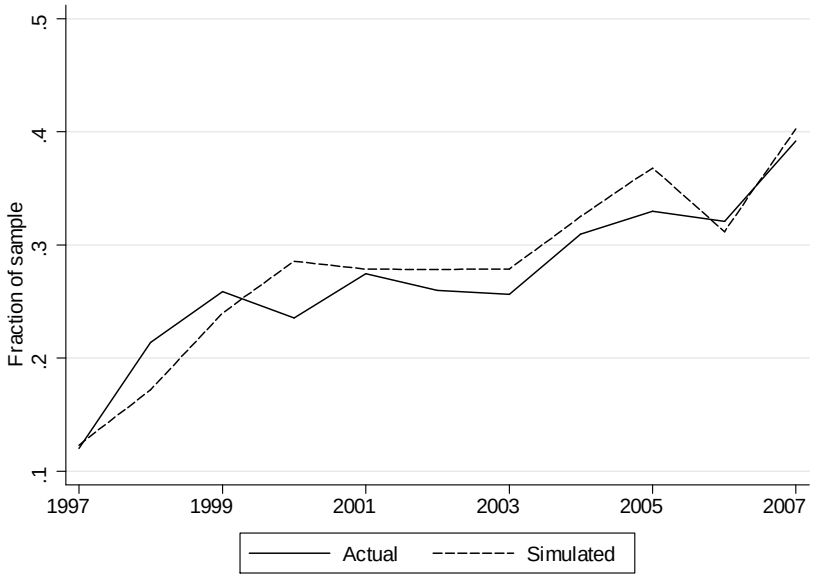


Figure 1.16: Model fit for fertility by job sector in Russia

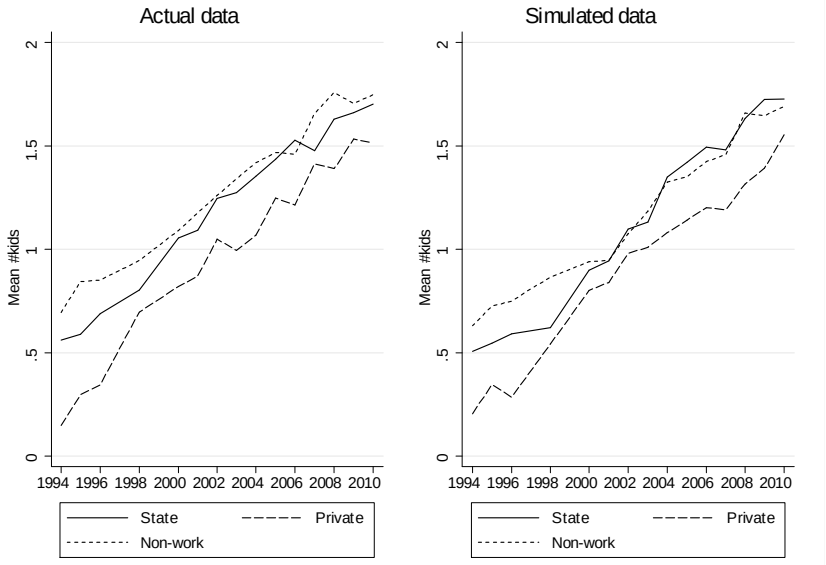


Figure 1.17: Model fit for fertility by job sector in Ukraine

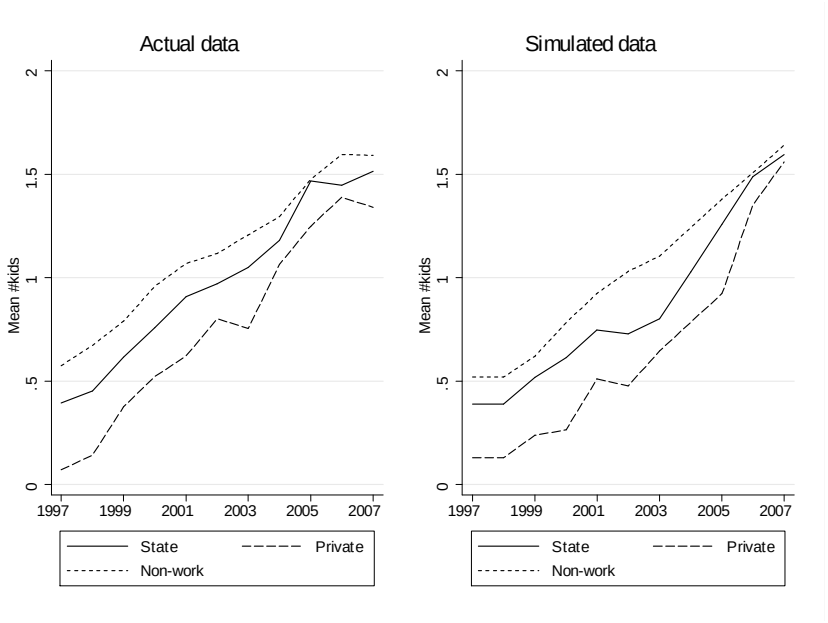
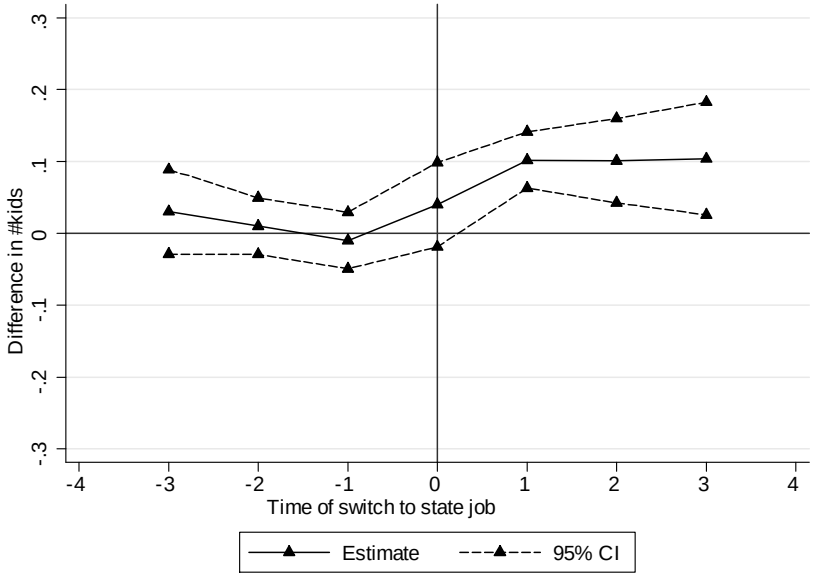
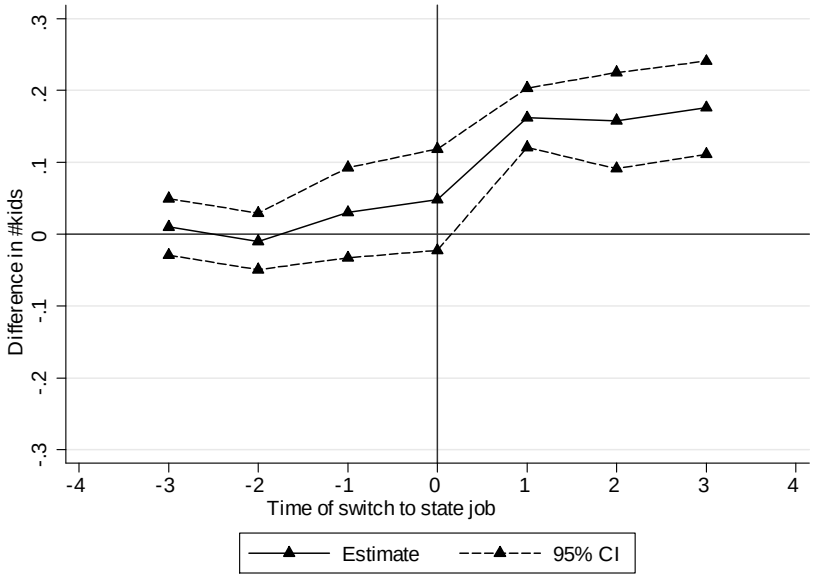


Figure 1.18: Simulated fertility response to switching sectors in Russia



Notes: Displayed are the regression estimates of the fertility difference between women who switch to state jobs and women who remain in private jobs.

Figure 1.19: Simulated fertility response to switching sectors in Ukraine



Notes: Displayed are the regression estimates of the fertility difference between women who switch to state jobs and women who remain in private jobs.

Figure 1.20: State employment shares under the counterfactual $\pi_{rt}^{p} = \pi_{rt}^{s}$

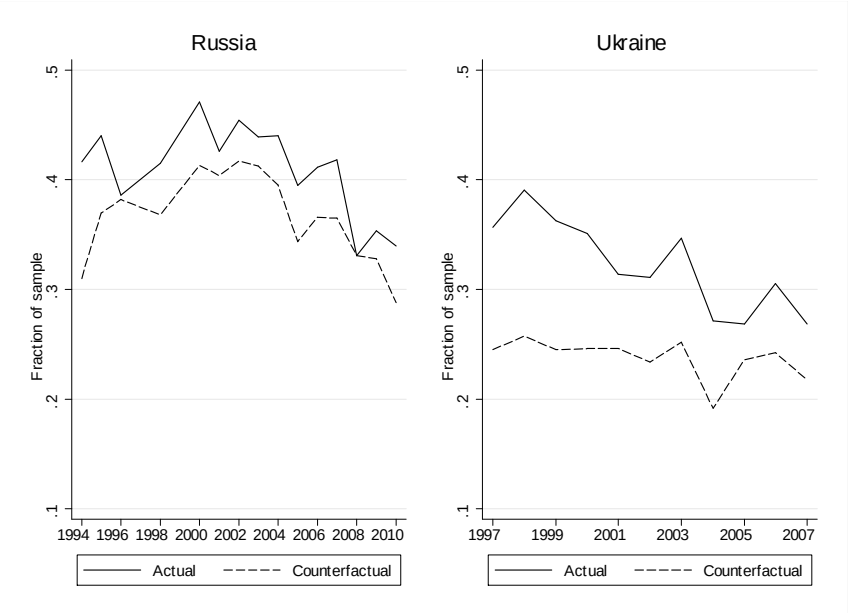


Figure 1.21: Private employment shares under the counterfactual $\pi_{rt}^{p} = \pi_{rt}^{s}$

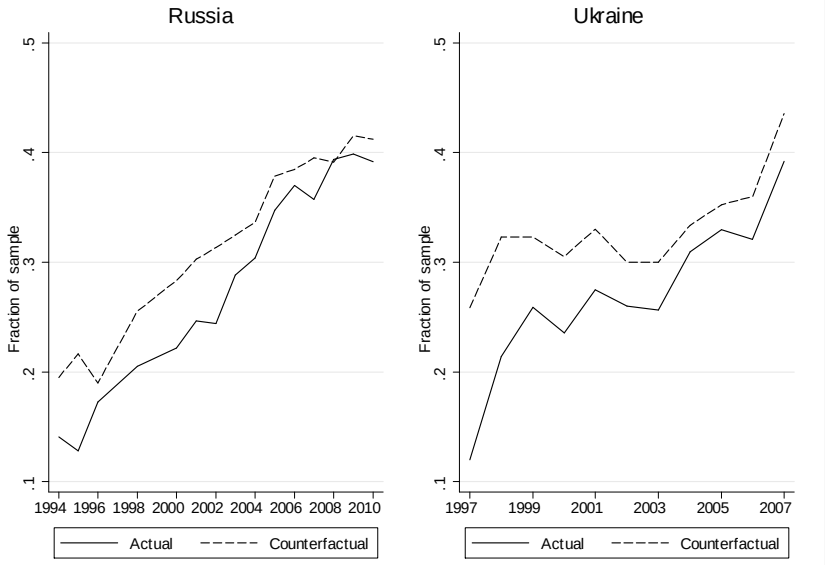


Figure 1.22: Fertility by sector in Russia under the counterfactual $\pi_{rt}^{p} = \pi_{rt}^{s}$

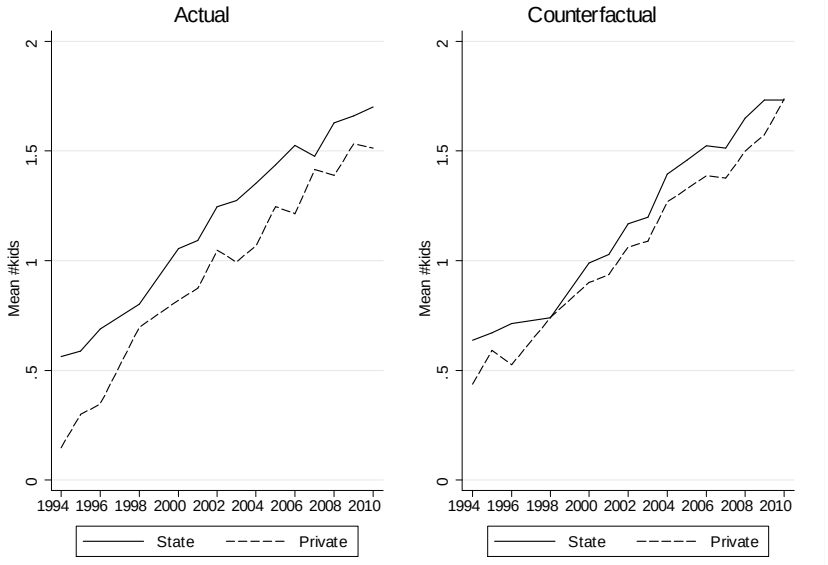


Figure 1.23: Fertility by sector in Ukraine under the counterfactual $\pi^p_{rt} = \pi^s_{rt}$

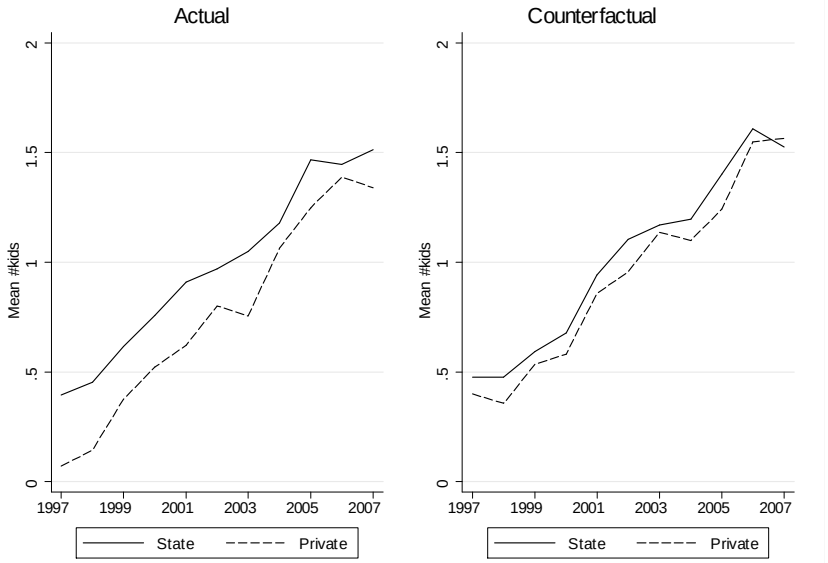


Figure 1.24: Fertility by sector in Russia under the counterfactual $\pi^p_{rt} = \pi^p_{r2010}$, $\pi^s_{rt} = \pi^s_{r2010}$

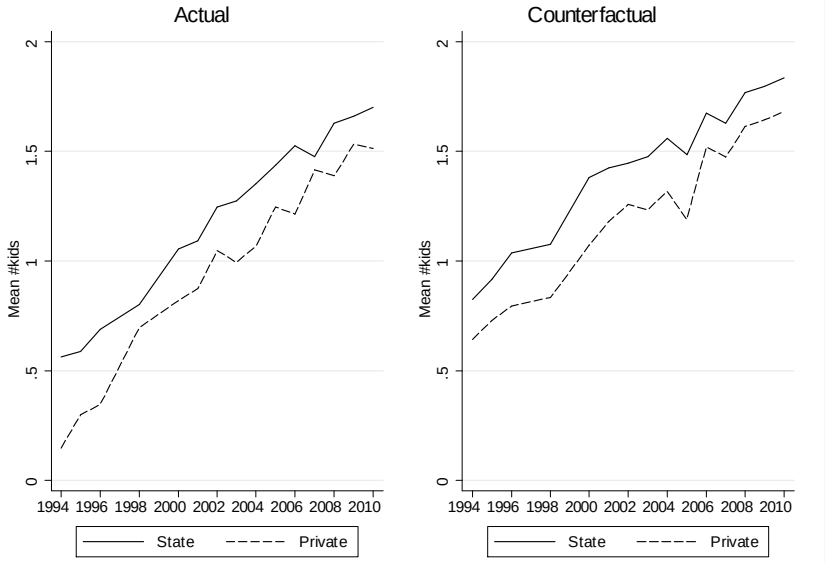


Figure 1.25: Fertility by sector in Ukraine under the counterfactual $\pi^p_{rt} = \pi^p_{r2007}$, $\pi^s_{rt} = \pi^s_{r2007}$

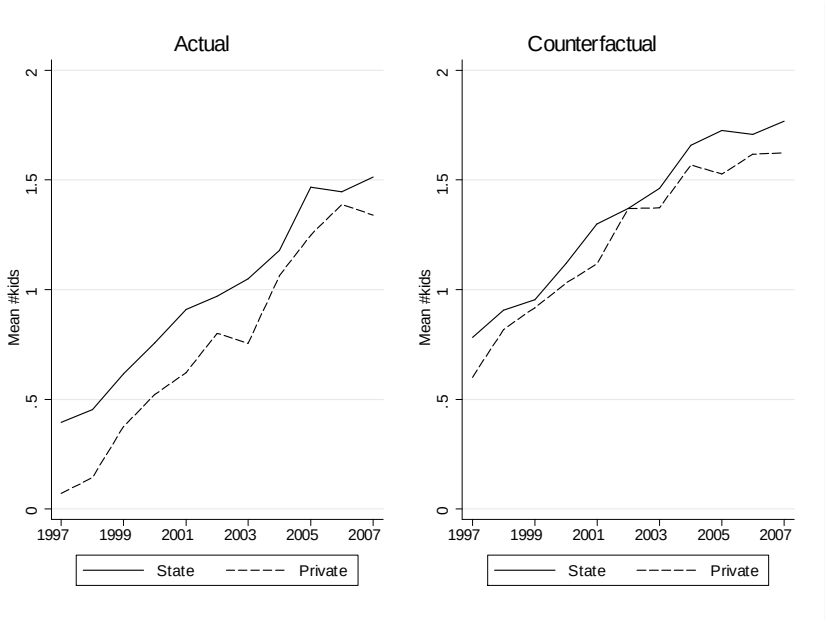


Figure 1.26: Fertility in Russia under a counterfactual policy that raises the duration of fully paid maternity leave

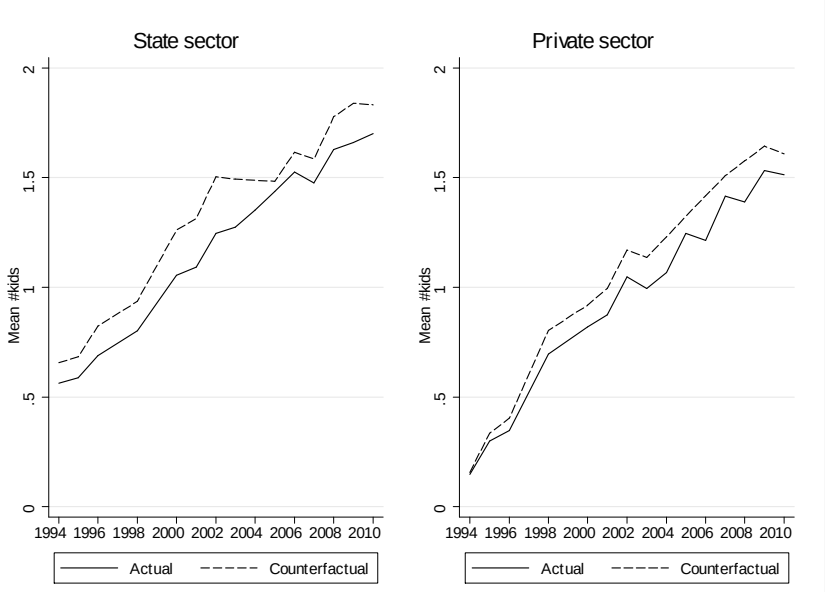


Figure 1.27: Fertility in Ukraine under a counterfactual policy that raises the duration of fully paid maternity leave

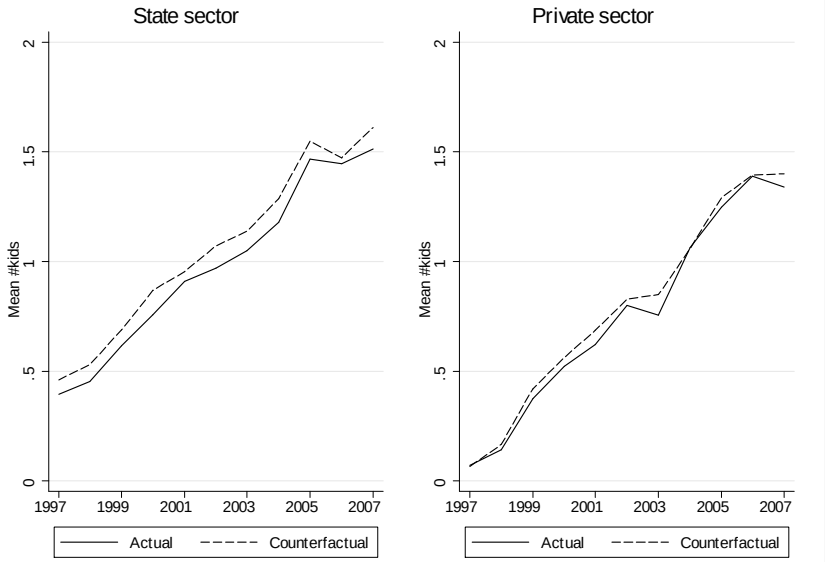
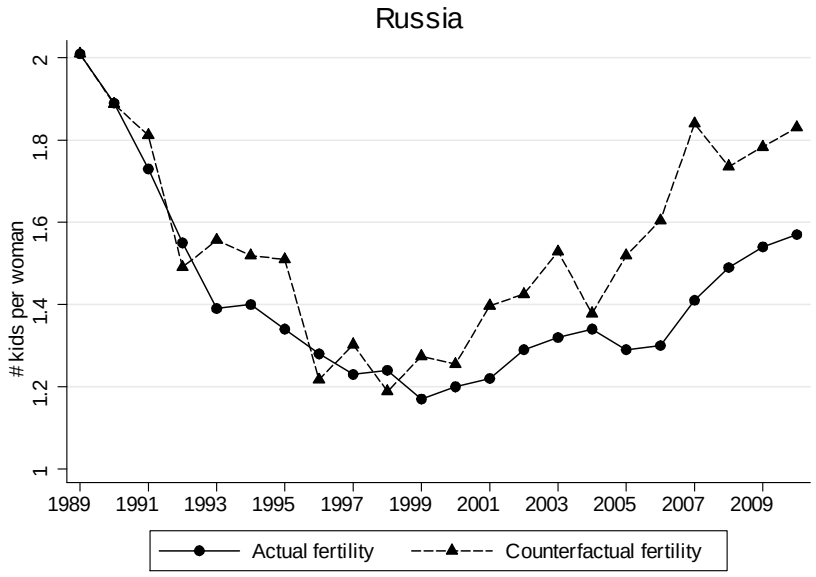


Figure 1.28: Counterfactual fertility in Russia with marriage rate fixed at 1989 level



Notes: Own calculations using marriage data from Russia Federal Statistics Committee.

Figure 1.29: Counterfactual fertility in Ukraine with marriage rate fixed at 1989 level



Notes: Own calculations using marriage data from Ukraine State Statistics Committee.

Figure 1.30: Availability and access to modern contraception in Russia

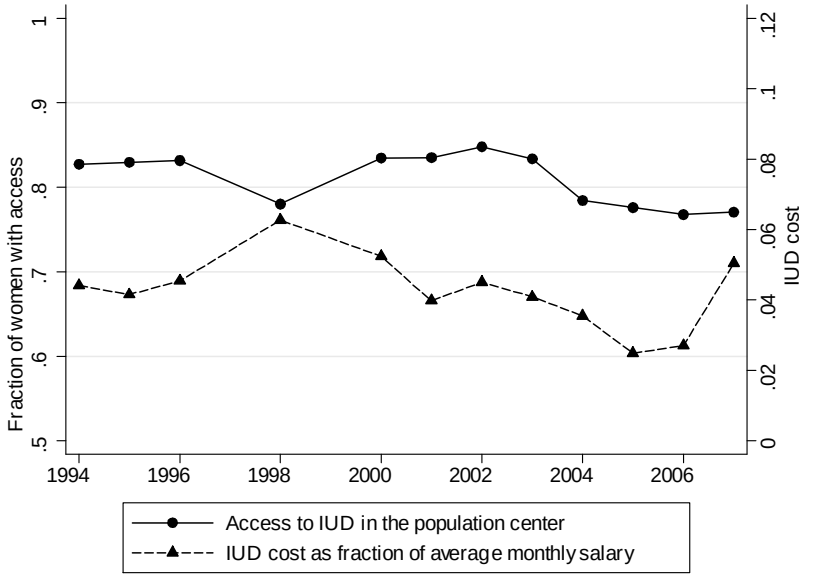
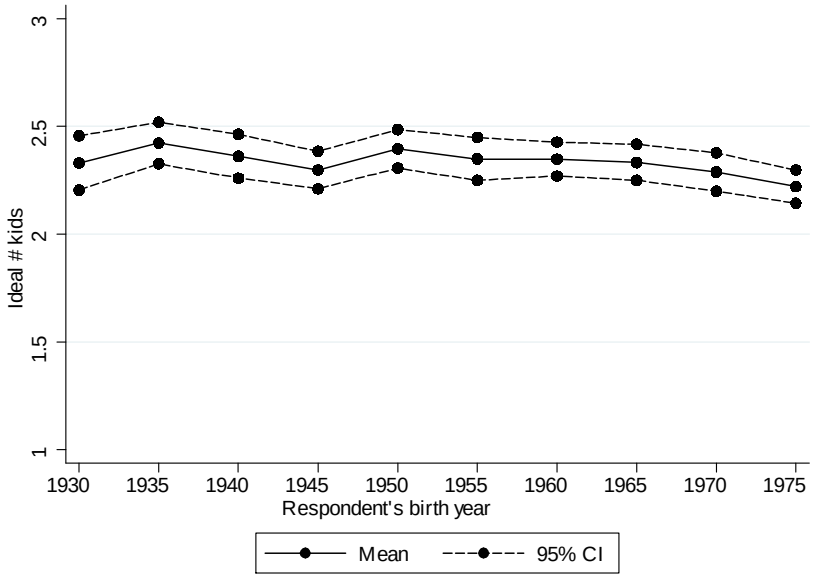


Figure 1.31: Ideal number of children by respondent's birth year in Ukraine



Notes: Displayed is the response to the survey question: 'If you could start your adult life" "over again under ideal economic conditions, how many children would you want to have?'"

Table 1.1: Russia – selected sample characteristics

	Women aged 15-24 in 1992				Women aged 25-40 in 1992			
	(1)		(2)		(3)		(4)	
	<=6 years of state sector experience in 1992-2003		>6 years of state sector experience in 1992-2003		<= 6 years of state sector experience in 1992-2003		> 6 years of state sector experience in 1992-2003	
	Mean	St. error	Mean	St. error	Mean	St. error	Mean	St. error
College	**0.236**	**0.015**	**0.327**	**0.038**	**0.199**	**0.012**	**0.333**	**0.022**
High school	0.651	0.016	0.609	0.039	**0.689**	**0.014**	**0.593**	**0.022**
Married	0.764	0.015	0.724	0.036	0.750	0.013	0.735	0.020
Age in 2003	30.158	0.095	30.355	0.113	43.664	0.137	43.983	0.200
HH income per capita	3211.380	134.984	2886.869	185.577	3387.614	154.867	3291.060	158.435
Pensioner in HH	0.221	0.014	0.192	0.032	**0.233**	**0.013**	**0.148**	**0.016**
Total # kids by 2003	**1.171**	**0.028**	**1.306**	**0.060**	**1.747**	**0.033**	**1.861**	**0.044**
Age at first birth	**22.774**	**0.127**	**22.000**	**0.268**	24.051	0.158	24.375	0.238
N	1015				1564			

Notes: Those characteristics that differ between (1) and (2), and between (3) and (4) at the 5% significance level are in bold. Income is average monthly income in 2003, measured in current prices (rubles).

Table 1.2: Ukraine – selected sample characteristics

	Women aged 15-24 in 1992				Women aged 25-40 in 1992			
	(1)		(2)		(3)		(4)	
	<=6 years of state sector experience in 1992-2003		>6 years of state sector experience in 1992-2003		<= 6 years of state sector experience in 1992-2003		> 6 years of state sector experience in 1992-2003	
	Mean	St. error	Mean	St. error	Mean	St. error	Mean	St. error
College	0.171	0.017	0.196	0.026	**0.162**	**0.015**	**0.231**	**0.014**
High school	0.646	0.022	0.692	0.030	**0.701**	**0.018**	**0.628**	**0.016**
Married	0.735	0.020	0.721	0.029	0.774	0.017	0.785	0.014
Age in 2003	30.241	0.134	30.546	0.168	43.618	0.184	44.002	0.151
HH income per capita	162.254	5.753	161.310	6.876	164.201	5.574	170.520	4.462
Pensioner in HH	0.058	0.011	0.054	0.015	0.075	0.011	0.062	0.008
Total # kids by 2003	**1.198**	**0.047**	**1.321**	**0.037**	**1.761**	**0.036**	**1.858**	**0.029**
Age at first birth	**21.935**	**0.200**	**21.383**	**0.130**	22.551	0.159	22.705	0.131
N	726				1496			

Notes: Those characteristics that differ between (1) and (2), and between (3) and (4) at the 5% significance level are in bold. Income is average monthly income in 2003, measured in current prices (hryvnias).

Table 1.3: Relationship between closures of state firms and incidence of wage non-payment

Dependent variable:	(1) Wage arrears at time t	(2) Compulsory unpaid leave at time t	(3) Job loss at time t
Has a state job at time $t-1$	0.013***	-0.021***	-0.026***
	(0.004)	(0.003)	(0.003)
Closings of state firms in the population center between time $t-1$ and t (dummy)	0.002	-0.002	0.012*
	(0.006)	(0.005)	(0.007)
Interaction term (state job X state closures)	0.016**	0.011**	0.021***
	(0.007)	(0.005)	(0.007)
N	10,574	10,574	10,574
N*t	33,699	33,699	33,699
R^2	0.121	0.132	0.177

Notes: Reported are the coefficients from linear probability models with robust standard errors, clustered by survey site, in parentheses. Controls include linear and quadratic terms for age and tenure, dummies for educational degrees, geographical region, living in rural area and survey year. *** p<0.01, ** p<0.05, * p<0.1

Table 1.4: Cross-sectional results with uncertainty measured as the probability of wage payment

Dependent variable:	Gives birth at time $t+1$			
	(1)	(2)	(3)	(4)
	State or private	State or non-work	State or private	State or non-work
	Russia		Ukraine	
Has a state job at time t	0.006**	0.016**	0.008***	0.010*
	(0.003)	(0.007)	(0.003)	(0.006)
Relative wage uncertainty of sector j over sector s in region r at time t $=(\pi^s-\pi^j)_{rt}$	0.003	-0.007	-0.023	-0.029
	(0.027)	(0.031)	(0.021)	(0.038)
Interaction term state job X $(\pi^s-\pi^j)_{rt}$	0.064**	0.047***	0.046**	0.027***
	(0.031)	(0.020)	(0.021)	(0.010)
N	4,981	5,593	3,257	3,799
N*t	15,874	17,825	26,142	32,672
R^2	0.153	0.149	0.094	0.101

Notes: Reported are coefficients from a linear probability model with robust standard errors, clustered by survey site, in parentheses. Controls include age, age^2, monthly hours worked, indicators for educational degree, married, ethnicity, receiving fringe benefits, # kids by time t, per capita household income, living in rural area, survey year and region. *** p<0.01, ** p<0.05, * p<*0.*1.

Table 1.5: Cross-sectional results with uncertainty in state jobs measured as closure of state firms

Dependent variable:	Gives birth at time $t + 1$	
	(1)	(2)
	State or private	State or non-work
	Russia	
Has a state job at time t	0.006**	0.011*
	(0.003)	(0.006)
Closure of state firms in the population center between time $t - 1$ and t (dummy)	-0.006	0.007
	(0.004)	(0.005)
Interaction term (state job X state closures)	-0.018***	-0.017***
	(0.006)	(0.006)
N	4,981	5,593
N*t	15,874	17,825
R^2	0.158	0.147

Notes: Reported are coefficients from a linear probability model with robust standard errors, clustered by survey site, in parentheses. Controls include age, age^2, monthly hours worked, indicators for educational degree, married, ethnicity, receiving fringe benefits, # kids by time t, per capita household income, living in rural area, survey year and region. *** p<0.01, ** p<0.05, * p<0.1

Table 1.6: The fertility effect of switching from a private to a state job

Dependent variable: Total # children	(1) by time t-3	(2) by time t-2	(3) by time t-1	(4) by time t	(5) by time t+1	(6) by time t+2	(7) by time t+3
Panel A: Russia							
Switch from private to state job at time *t*	0.025 (0.048)	0.022 (0.048)	0.026 (0.048)	0.027 (0.048)	0.098** (0.048)	0.099** (0.048)	0.106** (0.049)
N	1,013	1,013	1,013	1,013	1,013	1,013	1,013
R^2	0.399	0.390	0.372	0.356	0.345	0.324	0.305
Panel B: Ukraine							
Switch from private to state job at time *t*	0.022 (0.053)	0.008 (0.052)	0.001 (0.052)	0.012 (0.052)	0.124** (0.054)	0.116** (0.052)	0.129*** (0.053)
N	961	961	961	961	961	961	961
R^2	0.228	0.219	0.208	0.201	0.348	0.305	0.274

Notes: Reported are the results from a linear regression with robust standard errors, clustered by survey site, in parentheses. Controls include linear and quadratic terms for age, monthly hours worked, indicator for educational degrees, marital status, ethnicity, receiving fringe benefits, per capita household income, living in rural area, survey year and region. *** p<0.01, ** p<0.05, * p<0.1

Table 1.7: The effect of switching from a private to a state job based on relative uncertainty of private job

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dependent variable: total # children	by time t-3	by time t-2	by time t-1	by time t	by time t+1	by time t+2	by time t+3
Panel A: Russia							
Switch from private to state job at time *t*	0.031	0.028	0.030	0.060	0.082	0.084	0.099*
	(0.051)	(0.050)	(0.050)	(0.051)	(0.051)	(0.051)	(0.052)
Relative regional uncertainty of p jobs $=(\pi^s-\pi^p)_{rt}$	0.237	0.170	0.078	0.124	0.154	0.122	0.077
	(0.161)	(0.163)	(0.162)	(0.165)	(0.162)	(0.164)	(0.165)
Interaction term = Switch X $(\pi^s-\pi^j)_{rt}$	0.232	0.265	0.218	0.264	0.434**	0.473**	0.490**
	(0.297)	(0.289)	(0.286)	(0.208)	(0.219)	(0.221)	(0.233)
N	1,013	1,013	1,013	1,013	1,013	1,013	1,013
R^2	0.400	0.391	0.373	0.357	0.345	0.325	0.305

Table 1.7: The effect of switching from a private to a state job based on relative uncertainty of private job (continued)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dependent variable: total # children	by time t-3	by time t-2	by time t-1	by time t	by time t+1	by time t+2	by time t+3
Panel B: Ukraine							
Switch from private to state job at time *t*	-0.063	-0.071	-0.077	-0.068	-0.048	-0.043	-0.038
	(0.068)	(0.069)	(0.069)	(0.070)	(0.069)	(0.070)	(0.072)
Relative regional uncertainty of p jobs $=(\pi^s-\pi^p)_{rt}$	-0.157	-0.163	-0.193	-0.182	-0.193	-0.174	-0.204
	(0.137)	(0.137)	(0.137)	(0.137)	(0.138)	(0.140)	(0.141)
Interaction term = Switch X $(\pi^s-\pi^j)_{rt}$	0.454	0.333	0.327	0.333	0.548**	0.633**	0.662**
	(0.366)	(0.271)	(0.272)	(0.274)	(0.269)	(0.283)	(0.285)
N	961	961	961	961	961	961	961
R^2	0.122	0.188	0.177	0.357	0.328	0.298	0.274

Notes: Reported are the results from linear regressions with robust standard errors, clustered by survey site, in parentheses. Controls include age, age^2, monthly hours worked, indicators for educational degree, married, ethnicity, receiving fringe benefits, per capita household income, living in rural area, survey year and region. *** p<0.01, ** p<0.05, * p<0.1

Table 1.8: Exogenous parameters of the structural model

Parameter	Value	
	Russia	Ukraine
Cohort of women	1973-1976	1976-1979
Marital status	Married/cohabitating	Married/cohabitating
Initial age	18-21	18-21
Years in each period	1-2	1
Number of periods	15	11
Discount factor	0.95	0.95
Probability of conceiving at ages 18-19	1	1
Probability of conceiving at ages 20-24	0.954	0.954
Probability of conceiving at ages 24-29	0.909	0.909
Probability of conceiving at ages 30-34	0.834	0.834
Probability of conceiving at ages 35-39	0.746	0.746

Table 1.9: Estimated structural parameters

Parameter (symbol)	Estimate (standard error)	
	Russia	Ukraine
Risk aversion (θ)	0.545 (0.062)	0.672 (0.071)
Scale of children preferences (ϕ)	94.013 (26.545)	110.072 (27.681)
Curvature of children preferences (τ)	0.504 (0.061)	0.982 (0.075)
Disutility from working in state job (δ^s)	-12.103 (3.312)	-14.001 (5.365)
Disutility from working in private job (δ^p)	-13.252 (5.729)	-19.127 (7.785)
Variance of wage shock in state job ($\sigma^2_{\varepsilon^s}$)	0.875 (0.107)	0.554 (0.060)
Variance of wage shock in private job ($\sigma^2_{\varepsilon^p}$)	0.706 (0.087)	0.592 (0.079)
Covariance of wage shocks ($\sigma_{\varepsilon^s,\varepsilon^p}$)	0.382 (0.049)	0.334 (0.038)
Variance of taste shock for children ($\sigma^2_{\varepsilon^n}$)	191.091 (14.699)	100.165 (7.248)
Return to state experience in state job (linear)	0.093 (0.007)	0.068 (0.009)
Return to private experience in private job (linear)	0.133 (0.011)	0.082 (0.009)
Return to state experience in state job (quadratic)	-0.001 (0.0003)	-0.008 (0.001)
Return to private experience in private job (quadratic)	-0.007 (0.0002)	-0.011 (0.001)
Return to private experience in state job (linear)	0.053 (0.021)	0.036 (0.007)
Return to state experience in private job (linear)	-0.004 (0.014)	0.006 (0.008)
Return to private experience in state job (quadratic)	-0.003 (0.001)	-0.008 (0.011)
Return to state experience in private job (quadratic)	0.001 (0.012)	0.005 (0.014)
Depreciation of experience (η)	0.056 (0.021)	0.063 (0.028)
Return to university degree in state job	0.972 (0.132)	0.983 (0.144)
Return to university degree in private job	0.781 (0.099)	0.698 (0.094)
Return to high school degree in state job	0.078 (0.015)	0.237 (0.056)
Return to high school degree in private job	0.287 (0.062)	0.393 (0.068)
Wage intercept in state job (α^s_0)	9.412 (0.119)	7.099 (0.086)
Wage intercept in private job (α^p_0)	10.138 (0.124)	7.602 (0.094)
Return to living in urban area in state job (α^s_u)	0.151 (0.052)	0.190 (0.088)
Return to living in urban area in private job (α^p_u)	0.289 (0.064)	0.215 (0.065)

Notes: I compute the standard errors by the indirect inference procedure, outlined in Gourieroux et al. (1993). The estimates of the sector-specific returns to living in each geographical region are available upon request. In both Russia and Ukraine, living in the capital region brings higher returns than living in any other region.)

Table 1.10: Analysis of fertility effects under different counterfactuals

Alternative regimes:	Private sector wage uncertainty reduced to state sector uncertainty $\pi_{rt}^{\,p} = \pi_{rt}^{\,s}$ for $\forall r,t$		Both sectors' wage uncertainty reduced to latest observed levels $\pi_{rt}^{\,p} = \pi_{rT}^{\,p}$ $\pi_{rt}^{\,s} = \pi_{rT}^{\,s}$		20% increase in the duration of fully paid maternity leave (approx. 1 extra month of full pay)	
	(1) Russia	(2) Ukraine	(3) Russia	(4) Ukraine	(5) Russia	(6) Ukraine
	Change in fertility gap between sectors		Change in completed fertility		Change in completed fertility	
% change	-51.3%	-57.4%	+26.2%	+30.4%	+12.1%	+5.2%
	(48.9, 53.7)	(54.9, 59.9)	(23.9, 28.5)	(28.0, 32.8)	(10.6, 13.6)	(4.8, 5.6)
unit change	-0.18	-0.25	+0.40	+0.40	+0.18	+0.07
baseline	0.35	0.43	1.51	1.32	1.51	1.32

Notes: The 95% confidence intervals of the point estimates are given in parentheses. The first row reports the fertility effect as percentage change, the second row reports the effect as unit change and the third row reports the baseline fertility before the counterfactual took place. The unit of measurement is the average number of children per woman observed in the last period *T*. For columns (1) and (2) the unit is the difference in the average number of children between the private and the state sector.

Table 1.11: Fertility by job sector of both spouses

Dependent variable:	Gives birth at year $t + 1$	
	(1)	(2)
	Russia	Ukraine
Job sectors within couple at year *t*:		
Wife state, husband state (s,s)	0.018***	0.017***
	(0.004)	(0.004)
Wife state, husband private (s,p)	0.010**	0.008**
	(0.005)	(0.004)
Wife private, husband state (p,s)	0.007	0.005
	(0.006)	(0.005)
F-test for equality of coefficients on (s,p) and (s,s)	4.42**	4.80**
	[0.036]	[0.028]
F-test for equality of coefficients on (s,p) and (p,s)	1.26	1.98
	[0.262]	[0.159]
N of households	3,811	1,619
N*t	13,622	11,423
R^2	0.071	0.087

Notes: Reported are the regression results from linear probability models. Controls include mother's age and age^2, indicators for mother's educational degree, ethnicity, # kids by time *t*, per capita household income, living in rural area, survey year and region. Robust standard errors, clustered by survey site, are given in parentheses and the p-values from two-sided F-tests are given in brackets. *** p<0.01, ** p<0.05, * p<0.1

1.A Construction of uncertainty and child cost measures

In this section, I describe in more detail how the wage uncertainty and the child cost measures are constructed from the survey data.

1.A.1 Wage uncertainty measures

The overall probability of receiving in full one's annual wage in job sector j in region r in year t is π_{rt}^{j}.This probability is calculated as

$$\pi_{rt}^{j} = q_{rt}^{j} v_{rt}^{j} \kappa_{rt}^{j}, \tag{1.15}$$

where q_{rt}^{j} is the probability that an individual is not fired from her job in job sector j in region r at time t, v_{rt}^{j} is the probability that an individual is not being sent on compulsory unpaid leave, conditional on being employed in job sector j in region r at time t, and κ_{rt}^{j} is the probability that an individual is not experiencing wage arrears, conditional on being employed and not being sent on compulsory unpaid leave in job sector j in region r at time t.

In the RLMS data, I define $1-q_{rt}^{j}$ as the fraction of individuals who are employed in job sector j in region r at time $t-1$ and who report their main status as "unemployed and looking for a job" at time t. I define $1-v_{rt}^{j}$ as the fraction of individuals employed in job sector j in region r at time t who report that their employer has sent them on compulsory unpaid leave at some point during the current year. I define $1-\kappa_{rt}^{j}$ as the fraction of individuals employed in job sector j in region r at time t who report that they are owed at least 1 month worth of wages during the current year.

In the ULMS data, I define $1-q_{rt}^{j}$ as the fraction of individuals who report leaving their job in job sector j in region r at time t as a result of one of the following scenarios: "closing down of enterprise/organization," "reorganization of enterprise/organization," "bankruptcy of enterprise/organization," "dismissal initiated by the employer," "personnel reduction," "expiring of employment contract," or "expiring of probation time." I define $1-v_{rt}^{j}$ as the fraction of individuals employed in job sector j in region r at time t who report that the administration has sent them

on compulsory leave at some point during time t. I define $1-\kappa_{rt}^{j}$ as the fraction of individuals employed in job sector j in region r at time t who report that they are owed at least 1 month worth of wages during the current year.

Note that these definitions of wage uncertainty have some limitations. First, unlike the ULMS data, in the RLMS data, there is no information on the reason for losing one's job. It is possible that some of the job separations in the RLMS data are initiated by the employee. To mitigate this concern, I construct the measure using out-of-work individuals who are looking for jobs. Given that unemployment benefits are extremely meager in Russia, it is not very likely that individuals who are currently looking for jobs would have left their old jobs of their own volition. Second, in both the ULMS data and the RLMS data wage arrears that have been restituted with only a short time lag are not counted.

The duration of wage non-payment is captured by the probabilities ρ_{m}^{j} that a woman working in sector j receives exactly m $(m=0,1,2,3,...,11)$ monthly wages within a year, conditional on at least one month of wage non-payment. Each probability ρ_{m}^{j} is calculated using the observed distributions for the duration of unemployment, compulsory unpaid leave, and wage arrears. More precisely, the probability ρ_{m}^{j} (the probability that she receives exactly m wages throughout the year) equals the probability that the combined duration of unemployment, compulsory unpaid leave, and wage arrears is $12-m$, for $m=0,1,...,11$. Figures 1.32-1.33 below display the distribution of ρ_{m}^{j} for $j=\{p,s\}$ in Russia and Ukraine.

Figure 1.32: Probability distribution of the number of monthly wages received (conditional on at least one wage non-payment) in Russia

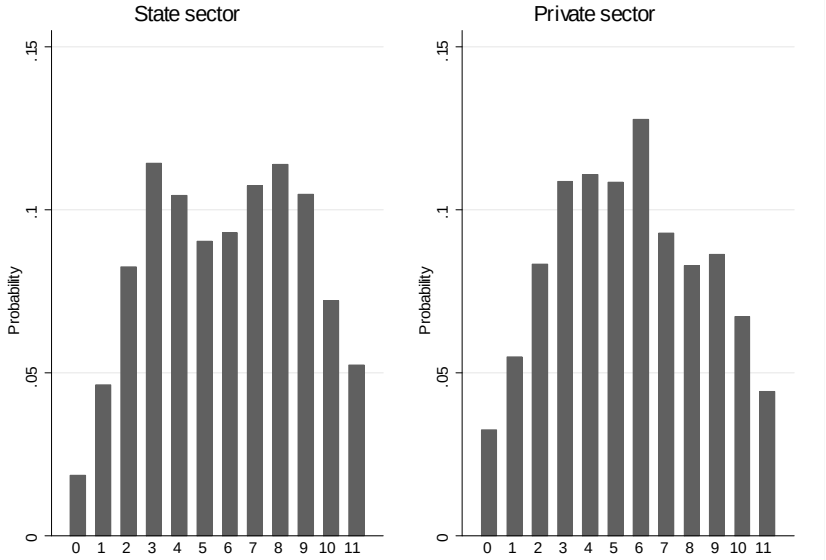
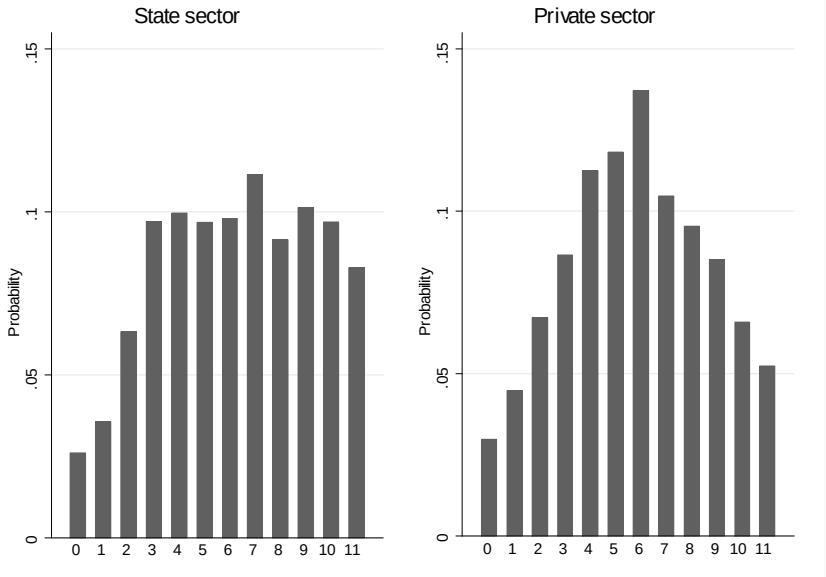


Figure 1.33: Probability distribution of the number of monthly wages received (conditional on at least one wage non-payment) in Ukraine



1.A.2 Child cost measures

I measure child costs z_r^j as the annual expenses for child care of a mother who lives in region r and works in job sector j. These are calculated as follows:

$$z_r^j = [h_{work}^j - h_{vacationleave}^j * \Pr(vacationleave)^j - h_{sickleave}^j * \Pr(sickleave)^j] * pcare_r \quad (1.16)$$

where h_{work}^j is the average annual hours worked by women in sector j, $h_{vacationleave}^j$ is the maximum annual hours of vacation leave, calculated as the product of the legally prescribed maximum of annual paid vacation days and the average daily hours worked by women in sector j, $\Pr(vacationleave)^j$ is the probability that a woman in sector j can actually take paid vacation leave (calculated as the fraction of women in sector j who report being able to take such leave from their employer), $h_{sickleave}^j$ is the maximum annual hours of sick leave, calculated as the product of the legally prescribed maximum of annual paid sick days and the average daily hours worked by women in sector j, and $\Pr(sickleave)^j$ is the probability that a woman in sector j can actually take paid sick leave (calculated as the fraction of women in sector j who report being able to take such leave from their employer). Lastly, $pcare_r$ is the average hourly price of childcare in region r, calculated from the weekly amount spent on childcare divided by the weekly number of hours of childcare, reported by households in region r (Lokshin 2004).

1.B Mean and variance of the wage distribution

In this section I derive the exact relationship between my measure of uncertainty $\pi_{r,t}^j$, the probability of receiving one's wage throughout the year t in region r in sector j, and the first two moments of the distribution of annual wages $W_{r,t}^j$.To simplify notation, I drop all region, time and sector subscripts and superscripts.

Consider first the formula for the mean of the distribution:

$$E[W] = \pi(12w) + (1-\pi)\sum_{m=0}^{11}\rho_m mw \tag{1.17}$$

where w is the monthly wage, and ρ_m is the probability of receiving exactly m $(m = 0, 1, 2, ..., 11)$ monthly wages, conditional on failing to receive at least one monthly wage. As discussed earlier, these probabilities are fixed parameters, which are estimated directly from the survey data.

Then, the derivative of the mean with respect to π is

$$\frac{\partial E[W]}{\partial \pi} = 12w - \sum_{m=0}^{11}\rho_m mw = w\left(12 - \sum_{m=0}^{11}\rho_m m\right) \geq w(12-11) > 0, \tag{1.18}$$

where I have used the fact that by definition the set of probabilities ρ_m sum to 1 ($\sum_{m=0}^{11}\rho_m = 1$).

Hence, an increase in the risk of wage non-payment (a decrease in π) decreases the mean of the wage distribution for any initial value of π.

Consider now the formula for the variance:

$$\begin{aligned}
Var[W] = \pi\Bigg(&(1-\pi)12w - (1-\pi)w\sum_{m=0}^{11}\rho_m m\Bigg)^2 \\
&+ (1-\pi)w^2\sum_{m=0}^{11}\rho_m m^2 - 2wmE[W] + E[W]^2) \\
&= \pi(1-\pi)^2 w^2\left(12 - \sum_{m=0}^{11}\rho_m m\right)^2 + (1-\pi)w^2\sum_{m=0}^{11}\rho_m m^2 \\
&\quad + (1-\pi)w^2\left(12\pi + (1-\pi)\sum_{m=0}^{11}\rho_m m\right)^2 \\
&\quad - 2(1-\pi)w^2\left(\sum_{m=0}^{11}\rho_m m\right)\left(12\pi + (1-\pi)\sum_{m=0}^{11}\rho_m m\right) \\
&= (1-\pi)w^2\left[\pi\left(12 - \sum_{m=0}^{11}\rho_m m\right)^2 + \sum_{m=0}^{11}\rho_m m^2 - \left(\sum_{m=0}^{11}\rho_m m\right)^2\right].
\end{aligned} \tag{1.19}$$

Then, the derivative of the variance with respect to π is

$$\frac{\partial Var[W]}{\partial \pi} = w^2 \left[\begin{array}{c} -2\pi \left(12 - \sum_{m=0}^{11} \rho_m m\right)^2 + \left(12 - \sum_{m=0}^{11} \rho_m m\right)^2 \\ -\sum_{m=0}^{11} \rho_m m^2 + \left(\sum_{m=0}^{11} \rho_m m\right)^2 \end{array} \right]. \tag{1.20}$$

Note that the derivative of the variance is a linear decreasing function of π. Hence, I can solve for the value of π at which the derivative becomes zero:

$$\pi_0 = \frac{1}{2} \left[1 - \frac{\sum\limits_{m=0}^{11} \rho_m m^2 - \left(\sum\limits_{m=0}^{11} \rho_m m \right)^2}{\left(12 - \sum\limits_{m=0}^{11} \rho_m m \right)^2} \right]. \tag{1.21}$$

For $\pi \in [\pi_0, 1]$, an increase in the risk of wage non-payment (a decrease in π) increases the variance of the wage distribution. The expression for this threshold value π_0 can be rewritten as

$$\pi_0 = \frac{1}{2} \left[1 - \frac{Var[M]}{(12 - E[M])^2} \right] \tag{1.22}$$

where M is the random variable denoting the number of monthly wages paid (conditional on failing to receive at least one monthly wage).

Note that this threshold π_0 has an upper bound of $\frac{1}{2}$ for any possible distribution of M ($Var[M] \geq 0$). In my sample, I can use the observed distribution of M to calculate the exact values of the threshold by job sector and by country.

Table 1.12 below presents the estimates for the first and the second moment of the distribution of M by job sector and by country, and the respective threshold values π_0. These threshold values are below the estimates of π for any year or geographical region in Russia (Figures 1.9 and 1.10). They are also below the estimates of π in Ukraine, except for year 1998 in the Eastern and Western regions (Figure 1.11). Hence, for the predominant majority of the Ukrainian sample and for the entire Russian sample, an increase in the risk of wage non-payment leads to a decrease in the mean and an increase in the variance of the wage distribution.

Table 1.12: Estimates for the threshold value π_0 by job sector and by country

	Russia		Ukraine	
	State sector (1)	Private sector (2)	State sector (3)	Private sector (4)
Mean of the number of monthly wages received $E[M]$	5.88	5.55	6.23	5.80
Variance of the number of monthly wages received $Var[M]$	8.64	8.59	9.17	8.20
Threshold value π_0	0.38	0.40	0.36	0.39

Notes: Reported is the threshold value π_0 for the probability of no wage loss, π. For values of π in the interval $[\pi_0, 1]$ a decrease in π leads to an increase in the variance of the wage distribution.

1.C Additional Results

Table 1.13: Cross-sectional results with uncertainty measured separately as the probability of wage payment in the state and private sector

Dependent variable:	Gives birth at year $t + 1$	
	(1)	(2)
	Russia	Ukraine
Has a state job at time t	0.006	-0.004
	(0.020)	(0.013)
Regional probability of receiving a wage in a state job π^{s}_{rt}	0.047	-0.012
	(0.032)	(0.024)
Regional probability of receiving a wage in a private job π^{p}_{rt}	0.028	0.014
	(0.031)	(0.021)
State job X π^{s}_{rt}	0.071**	0.052**
	(0.034)	(0.025)
State job X π^{p}_{rt}	-0.065**	-0.048**
	(0.031)	(0.021)
N	4,981	3,257
N*t	15,874	26,142
R^2	0.153	0.165

Notes: Reported are coefficients from a linear probability model with robust standard errors, clustered by survey site, in parentheses. Controls include age, age^2, monthly hours worked, indicators for educational degree, married, ethnicity, receiving fringe benefits, # kids by time t, per capita household income, living in rural area, survey year and region. *** p<0.01, ** p<0.05, * p<0.1

Table 1.14: Cross-sectional results with uncertainty measured separately by wage arrears, compulsory unpaid leave and unemployment

Dependent variable:	Gives birth at time $t+1$			
	(1) State or private	(2) State or non-work	(3) State or private	(4) State or non-work
	Russia		Ukraine	
Has a state job at time t	0.007*	0.009	0.009***	0.019***
	(0.004)	(0.006)	(0.003)	(0.008)
Relative uncertainty in sector j over sector s due to chance of unemployment $(q^s - q^j)_{rt}$	-0.012	0.025	0.028	0.044
	(0.063)	(0.022)	(0.037)	(0.031)
Interaction term state job X $(q^s - q^j)_{rt}$	0.073***	0.084***	0.038	0.027
	(0.021)	(0.035)	(0.039)	(0.028)
Relative uncertainty in sector j over sector s due to chance of unpaid leave $(\upsilon^s - \upsilon^j)_{rt}$	0.023	-0.061	-0.034	-0.050
	(0.061)	(0.071)	(0.033)	(0.039)
Interaction term state job X $(\upsilon^s - \upsilon^j)_{rt}$	0.087***	0.081**	0.066**	0.034*
	(0.035)	(0.037)	(0.031)	(0.020)
Relative uncertainty in sector j over sector s due to chance of wage arrears $(\kappa^s - \kappa^j)_{rt}$	-0.001	-0.016	-0.009	-0.044
	(0.041)	(0.034)	(0.021)	(0.029)
Interaction term state job X $(\kappa^s - \kappa^j)_{rt}$	0.042**	0.053*	0.055***	0.038***
	(0.021)	(0.031)	(0.022)	(0.016)
N	4,981	5,593	3,257	3,799
N*t	15,874	17,825	26,142	32,672
R^2	0.153	0.139	0.165	0.150

Notes: Reported are coefficients from a linear probability model with robust standard errors, clustered by survey site, in parentheses. Controls include age, age^2, monthly hours worked, indicators for educational degree, married, ethnicity, receiving fringe benefits, # kids by time t, per capita household income, living in rural area, survey year and region. *** p<0.01, ** p<0.05, * p<0.1.

Chapter 2

The Mental Cost of Pension Loss: The Experience of Russia's Pensioners during Transition[1]

2.1 Introduction

There is a growing literature on the correlates and determinants of subjective well-being (SWB) (Kahneman and Krueger 2006). In particular, the effects of income on SWB have received a lot of attention, especially in the context of developed countries.[2] Fewer studies have analyzed this relationship in the context of transition economies, and most of the existing studies have not identified an exogenous source of variation in income.[3] In transition economies, the public redistribution

[1] A slightly modified version of this chapter was published in *Research in Labor Economics*, 2012, Vol. 36: 191-240 (co-authored with Xiaotong Niu).

[2] For recent reviews, see Frey and Stutzer (2002), Di Tella and MacCulloch (2006) and Stutzer and Frey (2010).

[3] An exception is a study on lottery winnings and measures of SWB (Gardner and Oswald 2007). Whereas that paper examines the effects of an exogenous positive income shock, our study focuses on the effects of an exogenous negative income shock. Our analysis can reveal potentially different insights given experimental evidence that losses are felt more keenly than gains (Kahneman et al. 1991).

system can be a source of idiosyncratic income shocks for the household.[4] An important policy question is how the instability of social programs affects welfare as measured by SWB. Our results show that an adverse shock to the Russian pension system leads to a significant decline in measures of SWB and that the non-pecuniary cost of the shock is substantial.

This chapter exploits an episode of instability in the Russian pension system. In Russia, the elderly are a particular vulnerable group financially because most of them have no other personal source of income and live with few other employed adults. We investigate the impact of the 1996 pension crisis in Russia on several measures of SWB of pensioners and non-pensioner members of their households. In 1996 an uncertain political climate prior to the presidential elections of that year led to declines in economic output and in tax collection (the latter due to widespread tax exemptions granted before the elections). This resulted in a large decline in tax revenues relative to pension entitlements in most regions. The pension system experienced a significant funding crisis and many adults eligible for pensions were not paid in 1996. Nevertheless, the crisis was a one-time shock – the economy stabilized soon after the election and the pension system was restored in 1997. Still, the short-term impact of the crisis was severe: poverty rates doubled among affected pensioners (Jensen and Richter 2003).[5]

[4]The literature has also studied wage arrears in transition economies, another source of income loss that is rare in developed countries but widespread in transition economies. Earle and Sabirianova (2002, 2009) examine wage arrears in Russia using a framework that combines information on firms, workers, and local labor markets. According to these studies, wage arrears persisted for long periods throughout the 1990s and were the result of endogenous decisions by firms and workers. Hence, unlike pension arrears during the 1996 crisis, these episodes of wage arrears do not provide an exogenous source of income variation.

[5]A study closely related to ours is Jensen and Richter (2003). That study examines the same pension crisis as we do but it has a different empirical approach and different outcomes of interest. In particular, Jensen and Richter (2003) treat all individuals living in a pensioner household equally whereas we allow for a differential effect of the crisis on pensioners who are directly affected and non-pensioners who are only indirectly affected by living with pensioners. In addition, whereas our focus is on the effects on various aspects of subjective-well being, their focus is on health-related outcomes – calories, use of health services and medications. Jensen and Richter (2003) find a significant crisis effect on calorie intake but a more muted and varied effect on health reports and medication use. More details on the comparison between their results on objective health and our results on subjective well-being can be found in Section 2.6.

Our results show that the crisis significantly reduces life satisfaction of the affected pensioners. The availability of other measures of life evaluation in the data allows us to study which aspects of life are affected by the crisis. In particular, while self-assessed health is not affected by the experience of pension arrears, pensioner's expectations for future life satisfaction and future economic welfare and self perceptions of relative societal power are significantly lowered. There are also spillover effects to non-pensioners living in the same households with affected pensioners. Their reported life satisfaction declines just as strongly as the pensioners who are directly affected.

In addition, we find that the effect of the crisis is heterogeneous with respect to gender and expectations prior to the crisis. Male pensioners experience a greater decline in almost all measures of SWB. The crisis also has a stronger negative impact on those individuals – both pensioners and non-pensioners – who had more positive expectations about their future prior to the crisis.

Our study also looks at potential spill-over effects from the crisis beyond households with affected pensioners. We find no evidence of inter-household coping mechanisms to the pension crisis: the affected households neither give less nor receive significantly more money from extended family and friends during the crisis.

Lastly, we find some suggestive evidence that an institutional disruption in the pension system has non-pecuniary costs that extend beyond the income shock from the loss of pension income. We base this claim on several findings. First, our estimates of the effect of the crisis are 10 times larger than the predicted effect of a one-off income decline equal to the average pension size. Second, we find that non-pensioner household members report loss in life satisfaction equal to that of the pensioners directly affected by the crisis even when we control for the former's personal income. Third, we find no evidence that the magnitude of the crisis effect rises significantly with the number of pensioners in arrears in the household, which is what we would expect if the cost of the crisis were purely monetary. Finally, we find that the crisis has negative effects even on objective measures of well-being – immediate memory recall – for both pensioners and non-pensioners.

We contribute to the literature in several ways. First, our study establishes the causal effect of income on SWB using an exogenous shock to the redistribution system. Transition economies

provide a fruitful context for studying such income shocks. During the 1990s these economies experienced large and exogenous variations in incomes, which offer insights to many general economic questions. For example, Frijters et al. (2004a, 2004b, 2005) use the exogenous income variation between East and West Germany following reunification to study the effects of household income on life satisfaction (Frijters et al. 2004a, 2004b) and on health outcomes (Frijters et al. 2005). Similar to our findings, these studies find that rising household income is important for explaining rising life satisfaction in East Germany, and it is less important for explaining variation in health outcomes. Yet, the income shock in these studies is different in nature from ours. German reunification led to a steady increase in household income in East Germany, whereas the Russian pension crisis caused a one-time large decline in income. Another study which also focuses on Russia is Frijters et al. (2006). That paper examines how life satisfaction was affected by large swings in real household income that occurred in the late 1990s. Given the very large size of these income swings, the authors argue that they can be treated as exogenous shocks. Our contribution to this literature is that we analyze the effects of income on a level more disaggregate than the household – we are able to separately identify the effects of income loss when it is personally incurred and the effects of income loss when it is incurred only indirectly through shared household resources.

Second, unlike the studies cited above, our study explicitly exploits the fact that the pension income shock is due to an institutional failure. As a result, we can investigate whether there is any non-pecuniary cost to SWB apart from the monetary amount of pension loss. Such shocks to the redistribution systems might account for why cross-country studies find that individuals living in transition economies are much unhappier than would be predicted by their income levels (Deaton 2008). Our study on the impact of a pension crisis on subjective well-being provides some understanding into the causes of dissatisfaction with life in Russia, an economy in transition.

Third, our study provides an analysis on a richer set of subjective well-being measures in addition to the measure of life satisfaction, which is the most frequently used in the literature. These additional measures of SWB provide information on different dimensions of welfare and

can provide a deeper understanding of the effects of the pension crisis on overall life evaluation. In particular, the literature has emphasized the importance of relative income in determining SWB (Clark et al. 2008). Our data include several direct measures of self-assessed position in society: economic, respect and power rankings. These measures allow us to assess the effect of an income shock on one's perceived position relative to others in society.[6] Our data also include measures of expectations. According to the traditional lifecycle theory of consumption, one-time income shocks should have a smaller effect on utility than permanent shocks because of the opportunity for inter-temporal consumption smoothing. In hindsight, we know that the pension crisis was a one-time shock. However, if the pension eligible adults thought that the crisis would persist into the future, the pension crisis could have a large effect on expected permanent income, which might translate into a large income effect of the pension crisis on life satisfaction.

Fourth, our study builds on the literature of household resource allocation (Behrman 1995). Previous empirical studies have shown that household members share resources, but it is unclear whether such monetary spillovers translate into spillovers of life satisfaction.[7] If there is an intra-household response to the public crisis and we focus only on affected pensioners, we would underestimate the societal impact. We address this issue by looking at the effect of the pension crisis on the SWB of other individuals who are only indirectly affected by living together with pensioners who are directly affected. In addition, the crisis might affect other households linked to pensioner households through extended family networks. In Russia, such extended family networks are important source of informal monetary transfers, especially during the transition period. In particular, Kuhn and Stillman (2004) find that pensioners in Russia give substantial monetary transfers to their adult children. Unfortunately, we cannot identify adult children who are not living with their pensioner parents in our data. However, we have information on the amount of monetary transfers

[6]Lokshin and Ravallion (2005) find a significant cross-sectional relationship between household and personal income and perceived power in the context of Russia.

[7]In particular, in developing countries, pension income has been found to have a significant effect on the entire household and its members: pension receipt in South Africa changes household composition by increasing labor migration of prime-age household members (Ardington et al. 2009) and improves the health of children in the household (Duflo 2003).

sent and received by each household in our sample and we analyze the impact of pension arrears on these transfers.

Fifth, our study provides evidence supporting the theory of expectations being the reference point which determines preferences. In particular, we look at whether the crisis has a differential effect on SWB for groups with different pre-crisis expectations. Reference-dependent preferences in consumer theory suggest that utility is affected not only by the actual amount of resources but also by the amount of resources relative to a reference point. The reference point has been traditionally modeled as the status quo (Samuelson and Zeckhauser 1998). More recently, Köszegi and Rabin (2006) develop a model of consumer behavior in which the reference point is the rational expectations formed in the recent past about future outcomes. Laboratory experiments provide evidence in support of rational expectations as the reference point (Ericson and Fuster 2011). If it is expectations rather than the status quo that matter, we would expect those with higher pre-crisis expectations to experience a greater decline in SWB after controlling for household income. Indeed, this is what we find in our analysis.

The remainder of the chapter is organized as follows. Section 2.2 discusses the data and the measures of SWB used in the study. Section 2.3 describes the pension crisis of 1996. Section 2.4 presents the estimation strategy and addresses some identification issues. Section 2.5 reports the main estimates of the effect of pension arrears on the SWB of affected pensioners and non-pensioners living in households with affected pensioners. Section 2.6 discusses additional considerations and results from related objective measures of well-being. Lastly, Section 2.7 concludes with a discussion of policy implications.

2.2 Data

The sample we use comes from Phase II of the Russia Longitudinal Monitoring Survey (RLMS), a panel data of Russian households that started in 1994. Phase II data are collected in 1994, 1995, 1996, 1998, and annually starting in 2000. About 4,000 households were first interviewed in 1994.

Each household is given a household questionnaire. In addition, individual questionnaires are given to all members of the household except the very young and the very old.[8] Households that move out of their original dwellings are not followed in subsequent rounds and new households are added to the survey to maintain sample size and representation. However, given low migration and moderate non-response rates, sample attrition has been modest and especially so for older demographic groups which are the focus of our study.[9]

In our main analysis we focus on the first three survey rounds: 1994, 1995 and 1996.[10] Each round is administered in the last three months of the calendar year – October through December. The 1996 survey is the first survey collected after the pension crisis started earlier that year. In our study we exploit the panel nature of the data, which allows us to track affected individuals before and after the crisis. Pension eligibility in Russia is defined by age: all women over the age of 55 and all men over the age of 60 are eligible for pension receipts. Our sample consists of all pension-eligible adults in 1996 and other members of their households. The individual surveys include information on income and pension receipts in the past 30 days. We define an age eligible pensioner as being in pension arrears if he or she did not receive any pension income in the last 30 days.

We analyze several measures of SWB. In the survey respondents are asked to assess their overall life satisfaction and health status on a 5-point scale in which 1 is "very bad" and 5 is "very good." We refer to these variables as measures of life satisfaction (LS) and self-assessed health

[8]The survey also includes a community questionnaire that collects extensive data for each survey site. For more details on the survey design, see http://www.cpc.unc.edu/projects/rlms-hse/project.

[9]Our main analysis focuses on the sample of 1994, 1995, and 1996. In those survey rounds, among households in which the head is eligible for pension, the household attrition rate is between 11.6% and 13.4% (Heeringa 1997).

[10]In Section 2.6 we present additional analysis on the long-term effect of the pension crisis using data from the fourth survey round which took place in October 1998 – January 1999. We exclude data from 1998 (or later rounds) from our main analysis due to the difficulty in separating the effect of the 1996 pension crisis from the effect of the 1998 financial crisis which started before data was collected in 1998. Although we have no particular reasons to believe that the financial crisis – a country-wide shock – should affect differently our treatment and control groups, we also have no means of testing this assumption.

(SAH). We also look at the effect of the crisis on other aspects of life evaluation. The survey asks respondents to place themselves on a 9-step Cantril ladder for perceived economic, power and respect rankings in society (Cantril 1965). The English-translated question on perceived economic ranking is as follows: "And now, please, imagine a 9-step ladder where on the bottom, the first step, stand the poorest people, and on the highest step, the ninth, stand the rich. On which step are you today?" This measure is referred to as the Welfare Ladder Question (WLQ) in the literature. The English-translated question on perceived power ranking is as follows: "And now, please, imagine a 9-step ladder where on the bottom, the first step, stand people who are completely without rights, and on the highest step, the ninth, stand those who have a lot of power. On which step are you?" This measure is referred to as the Power Ladder Question (PLQ) in the literature (Lokshin and Ravallion 2005). Lastly, the question on perceived respect ranking is: "And now, another 9-step ladder where on the lowest step, are people who are absolutely not respected, and on the highest step stand those who are very respected. On which step of this ladder are you?" This measure is referred to as the Respect Ladder Question (RLQ).

In the data there is also a set of measures reflecting one's outlook on future economic welfare and future life satisfaction (both measured on a 5-point ordinal scale). The question about future economic welfare asks "How concerned are you about the possibility that you might not be able to provide yourself with the bare essentials in the next 12 months?" where an answer of 1 is "very concerned" and an answer of 5 is "not at all concerned." The question about future life satisfaction asks "Do you think that in the next 12 months you and your family will live better than today or worse?" where an answer of 1 is "much worse" and an answer of 5 is "much better."

Table 2.1 presents characteristics of our sample of pensioners and non-pensioners in 1994 and 1995, the two pre-crisis years. Over 70% of our pensioner sample is female. This is probably due to the much lower life expectancy of Russian men than Russian women and the lower age cut-off for pension eligibility for women. Pensioners who experience pension arrears during the 1996 crisis are similar to those in non-arrears in the years prior to the crisis in terms of age, gender, rates of marriage, unemployment, disability and household composition. Both groups also transfer

similar (positive) amounts of money to their extended families. The biggest difference between the groups is that the former are almost twice as likely to live in a rural area – close to 44% versus 23%. This difference likely accounts at least partially for the former's lower education and income. We observe the same patterns by arrears status among the non-pensioners living in pensioner households. In addition to the above differences, those in arrears are slightly more likely to live in households with more young children and pensioners. In the next section we describe the pension crisis and analyze in more detail the determinants of arrears status.

2.3 The pension crisis of 1996[11]

The old-age pension system in Russia is administered by the state pension fund, the Pension Fund of Russia (PFR). Pension receipt is based on age eligibility and not on current employment status. There are also provisions for early retirement and early pension receipt for serious medical conditions or work in hazardous conditions.[12] Until the pension reform in 2002, the PFR was financed mainly from payroll taxes from employers and, to a very small extent, from the government budget. In 1996 the uncertain political climate before the results of the July presidential elections triggered a decline in economic output. That decline was also accompanied by a decline in tax collection due to weaker tax enforcement and various tax exemptions given out in connection to the elections. As a result, the pay-as-you-go pension system experienced a serious shortage of revenues relative to entitlements. This funding crisis caused a sudden increase in the percentage of pensioners not receiving their pensions. Figure 2.1 shows the trend in pension arrears over the years. The rate of arrears was under 10% in 1994 and 1995. It jumped to more than 30% during the crisis in 1996 and went back down afterwards.

[11]The description of the pension crisis in this section relies heavily on Jensen and Richter (2003). Interested readers can find more details in that study.

[12]We exclude the group of early retirees from our main analysis because of sample selection issues. The effect of the crisis might be heterogeneous along the dimensions of personal or household characteristics that also determine early retirement. Since we do not observe the reasons for early retirement, we cannot adequately control for this potential heterogeneity in effect. Nevertheless, as discussed in Section 2.5.3, our results are robust to including early retirees in the sample.

The pension system in Russia is largely decentralized. Regional pension funds collect local payroll taxes and local pension offices disburse pensions out of these revenues. Any revenue surplus is sent to the central pension fund which distributes it to those regional funds which have more regional entitlements than regional tax revenues. Such decentralization can potentially lead to regional differences in the severity of the crisis. If this were the case, we would worry that the estimated effect is capturing the differences between administrative regions. Table 2.2 presents the fraction of age eligible pensioners in arrears by 8 main geographical regions.[13] The incidence of pension arrears was widespread during the crisis. Except for the Metropolitan region, which includes Moscow and St. Petersburg, all other regions experienced sizable increases in pension arrears during the crisis. Not only was the crisis widespread, but its impact on individuals was potentially considerable. Pension income is very important for pensioner households. In 1995 pension income represented 74.6% of the total household income for pensioner households in our RLMS sample.

The reason for the pension crisis is obvious: there was not enough in the pension fund to pay off all obligations. Less obvious is how administrators dealt with the fund shortage. Instead of lowering the pension amount for every eligible pensioner, resources were used to meet obligations for a subset of eligible adults. Figure 2.2 shows the average amount of pension receipt in the sample by survey round. All pensions are converted in June 1992 rubles. We see that the average pension amount among the non-arrears is comparable in 1995 and 1996, whereas the average amount among all pensioners drops sharply in this period due to increase in the number of arrears.

Since only some pensioners were paid, what were the selection criteria? We estimate a linear

[13]The administration of the pension system is done at a more disaggregated regional level – in 1996 there were 89 administrative districts in Russia that collected taxes and disbursed pensions independently of each other. 32 of these districts are included in the RLMS sample but the small number of observations does not allow us to present separate estimates at the district level. In Table 2.2 we show how the 8 geographical regions map to these administrative districts based on their debtor or donor status in 1996, as defined by the Pension Fund of Russia. Among the 32 districts included in the RLMS only 5 were donor districts in 1996. 2 of those districts map directly into the Metropolitan region, and the other three are each subsumed into the following three regions: Northern and North Western, Western Siberian and Eastern Siberian and Far Eastern. The remaining 4 geographical regions consist only of debtor districts.

probability model where the outcome variable is whether a pensioner is in arrears in 1996. We want to see which pre-crisis characteristics can predict arrears status in 1996. There were official guidelines suggesting that priority is given to those with low income, the unemployed and the single. Some regions also gave priority to the old and the disabled. Nevertheless, these guidelines were neither well, nor universally followed. As Table 2.3 shows, the probability of arrears during the crisis is not associated with pre-crisis household income in any region. Low pension income and being unemployed prior to the crisis are also not universal determinants of arrears status, with only a couple of regions showing a statistically significant negative association for at most one of these characteristics. Moreover, in several regions these variables are associated with arrears status in the direction opposite of the one predicted by the priority criteria. Neither are any of the other pre-crisis individual or household characteristics systematically related to arrears status.[14]

2.4 Empirical strategy

2.4.1 Main estimates

We employ a difference-in-difference (DD) approach to study the effect of pension arrears on SWB. In particular, we compare the change in the measures of SWB before and after the pension crisis for our treatment and control groups. Our pensioner sample consists of all individuals who are eligible for pensions in 1996 and who did not experience pension arrears in any year before the crisis. Our non-pensioner sample consists of all adult individuals who are living in the same household with at least one member of the pensioner sample. Within the pensioner sample we define the treatment group as those individuals who do not receive their pensions during the 1996 crisis and the control group as those individuals who do receive their pensions during the crisis. Within the non-pensioner sample we define the treatment group as adult individuals living with pensioners who are in arrears during the crisis and the control group as adult individuals living

[14]When we estimate this model for all regions combined (with regional dummies to allow for regional level differences) we find that the only statistically significant determinant of arrears status is living in a rural area.

with pensioners who are not in arrears. The identifying assumption is that in the absence of the crisis, there would be no differential time trends in the measures of SWB between treatment and control groups.

Figure 2.3 plots the time trend of all seven measures of SWB for the pensioner sample. These figures can be viewed as presentations of simple DD models without any controls. In the pre-crisis period from 1994 and 1995, pensioners who would go in arrears during the 1996 crisis report almost identical levels and trends in life satisfaction as those pensioners who would not go in arrears during the crisis (Figure 2.3a). The two groups also report very similar expectations for future life satisfaction and future economic welfare in the pre-crisis period (Figures 2.3c and 2.3d). Immediately after the 1996 crisis, however, the former group experiences sizable declines in all three of these measures while the latter group experiences no visible declines (or even experiences increases in expectations). The crisis effects on the three measures of self-perceived position in society seem more muted, but it still appears that the group affected by arrears during the crisis reports larger post-crisis declines in all three measures (Figures 2.3e, 2.3f and 2.3g). At the same time, the two groups' trends in self-assessed health do not appear to diverge in the immediate aftermath of the 1996 crisis (Figure 2.3b). Note that by 1998 pensioners affected by the 1996 crisis report levels of well-being that are close to (or even higher than) those of pensioners who were unaffected by the 1996 crisis. However, as mentioned before, the 1998 reports of well-being are after the severe financial crisis of that year and so this apparent convergence might be a product of the larger financial crisis, and not necessarily evidence that the pension crisis effects were short-lived.

The decentralized nature of the pension system implies the possibility of differential effects of the crisis across regions. In addition, the official guidelines that determine priority of pension receipts could also induce differences between the treatment and control groups. To guard against such potential problems, we control for pre-crisis characteristics and differential trends across regions. Even if the treatment, or arrears status, is randomly assigned, the inclusion of these

additional controls improves the precision of our estimates. The DD regression model is given by:

$$y_{it} = \beta_0 + \beta_1 Arrears_i + \beta_2 Arrears * Year1996_t + \beta_3 Year1995_t + \beta_4 Year1996_t + \gamma' Z_{it} + \varepsilon_{it} \quad (2.1)$$

where y_{it} is the SWB measure of individual i at time t (t = 1994, 1995, 1996), $Arrears_{it}$ is an indicator variable for the treatment group that equals 1 if the individual does not receive his or her pension during the pension crisis of 1996, and $Year1995_{it}$ and $Year1996_{it}$ are indicator variables for 1995 and 1996, and Z_{it} is a vector of exogenous demographic and pre-crisis socio-economic characteristics, which are likely to be correlated with both arrears status and SWB. These controls are as follows: age and age squared, indicators for female, married, high school graduate, college graduate, residing in rural area; indicators for pre-crisis disability status, pre-crisis unemployment status, low pre-crisis pension, pre-crisis total household income excluding pensions and pre-crisis number of household members in ages 0-6, 7-17, pre-pension age adults and pensioners.[15] We also include a set of indicators for the 8 main geographical regions (see Table 2.2 for classification) as well as their interactions with and to control for both permanent and time-variant regional differences.

We estimate Equation (2.1) by OLS and we cluster the standard errors at the level of administrative pension district-year.[16] The coefficient of interest, β_2, gives the effect of the pension crisis on the measures of SWB. Namely, it is the difference in the post-crisis change of SWB between treatment and control groups. This identification is valid if there are no other factors besides those

[15] All income measures in subsequent analysis are deflated to June 1992 rubles, using the CPI index provided by the RLMS survey team.

[16] In section 2.5.3 we demonstrate that our results are robust to alternative specifications such as an ordered probit which, unlike the OLS model, does not assume that the measures of SWB are cardinal. We use OLS as our preferred specification because the magnitudes of the effects estimated from this model are easier to interpret and because adding individual fixed effects in ordered probit models poses additional technical difficulties. In addition, Ferrer-i-Carbonel and Frijters (2004) show that OLS models perform very similarly to ordered probit models for measures of SWB.

accounted for in Z_{it} that differentially affect the trends in SWB of treatment and control groups.

We then proceed by adding individual fixed effects to the DD specification. The regression equation now becomes:

$$y_{it} = \beta_0 + \beta_1 Arrears_i * Year1996_t + \beta_2 Year1995_t + \beta_3 Year1996_t + I_i + \gamma' Z_{it} + \varepsilon_{it} \quad (2.2)$$

where I_i are the individual-specific effects. The rest of the variables are defined in the same way as in Equation (2.1), except that we have removed all time-invariant characteristics and Z_{it} now includes: a linear and quadratic term in age and interactions between the indicators for geographical region and survey year. Standard errors are clustered at the administrative pension district level. In this specification the coefficient of interest is β_1 which gives the within-person effect of the pension crisis. This specification is more conservative than the DD specification because it allows us to additionally control for any unobservable (permanent) individual characteristics that might be correlated with both arrears status and measures of SWB. This is our preferred specification for all subsequent analysis.

2.4.2 Channels of the pension crisis

Equations (2.1) and (2.2) give us the estimated effect of the pension crisis on SWB. We now proceed to explore the channels through which the pension crisis affects SWB. By definition, being in pension arrears during the crisis has a negative income effect on household resources. Yet, post-crisis household income excluding pensions is possibly endogenous with respect to arrears status: pensioners in arrears might go back to the labor force, or other non-pensioner household members might increase their labor supply to compensate for the pension loss. Similar endogeneity concerns apply to the post-crisis net monetary transfers from extended family and friends and to the post-crisis household composition. To separate the income effect of being in arrears from these possible household responses to the crisis, we include the following additional controls, measured at their

post-crisis levels: total household income excluding pensions (as a linear and quadratic term), net monetary transfers from extended family and friends and number of household members by age group (defined above).

Even though the pension crisis manifested itself as an income shock, it might also have had a non-pecuniary cost on SWB. In particular, we hypothesize that such an income shock which originates from an institutional failure reduces well-being in ways that are not captured fully by the size of the pension loss. To test this, we estimate Equation (2.2) by splitting the treatment group in two subgroups based on the intensity of treatment. We separately measure the effect of arrears for households who lost 2 or more pensions in the crisis and for households who lost exactly 1 pension in the crisis. If the crisis cost is experienced only in monetary terms, then the former group must experience a significantly greater decline in SWB.

Exploiting the rich set of SWB measures in our dataset, we also examine the possibility that the pension crisis affects other aspects of SWB which capture relative life evaluation, where the relation can be with respect to others in society or with respect to one's future self. To test this, we estimate the effect of pension arrears on the three measures of self-perceived relative standing in society (WLQ, PLQ, and RLQ) and the two measures of future expectations (for life satisfaction and economic welfare) using the econometric model as specified in Equations (2.1) and (2.2).

To further study the heterogeneity of the crisis effect, we estimate the effects separately for female and male pensioners. We also test the hypothesis that the pre-crisis expectation for the future is the relevant reference point which determines how the crisis impacts SWB. To do this we perform separate estimation for two groups that differ in their pre-crisis expectations. In our data we have two measures of expectations, expectations about future life satisfaction and future economic welfare, where each is defined on a 1 to 5 scale. We sum the two measures of expectations to obtain one measure of overall expectation measured on a 2 to 10 scale. We then take the average of this overall expectation in the two pre-crisis years: 1994 and 1995. Those individuals with an overall pre-crisis expectation measure of 4 or less are defined as having a low pre-crisis outlook and those with an expectation measure greater than 4 are defined as having a high pre-crisis outlook. The

outcomes of interests are life satisfaction, self-assessed health, and self-perceived rankings in the society (WLQ, PLQ, and RLQ). We estimate (2.2) separately for the group of low and the group of high pre-crisis outlook for the future.

2.4.3 Intra- and inter-household externalities

To see whether the adverse effects of the pension crisis extend beyond the pensioners directly affected, we study the effects of the crisis on SWB for the sample of non-pensioners who live in households with pensioners. We use the same strategy as in our analysis of the sample of pensioners. For the main results we estimate Equations (2.1) and (2.2), where in the DD estimates without individual fixed effects we add pre-crisis personal income as an additional control. This allows us to estimate the effect of the crisis separately from any other potential effects due to differences in personal pre-crisis resources between our treatment and control groups.

We then include as additional controls post-crisis measures such as personal income and household income excluding pensions (as linear and quadratic terms), net monetary transfers from family and friends and household composition. In this way we separate the income effect of pension loss from possible compensating responses by other household members.[17] In particular, we test whether living with a pensioner in arrears affects one's well-being separately from the current level of personal resources and whether the size of the effect rises with the number of pensioners in arrears. We also explore two additional dimensions of heterogeneity: we examine whether the effect of arrears differs by gender and by the level of pre-crisis expectations for the future. All these analyses are performed in same way as described above in the case of pensioners.

In addition, we attempt a test of whether the crisis impact extends even further into society beyond pensioner households to other households of extended family and friends. Because those latter households cannot be identified in the survey, we cannot directly test the effect of the crisis on measures of SWB for those households. Instead, we test whether households linked by familial

[17] As discussed before, by including these post-crisis measures we face the possibility that they are endogenous with respect to arrears status.

or social ties to pensioners in arrears are affected by the pension crisis through changes in the monetary transfers being sent or received by them. We estimate the household-level equivalent of the individual fixed effects specification in Equation (2.2):

$$y_{ht} = \beta_0 + \beta_1 Arrears_h * Year1996_t + \beta_2 Year1995_t + \beta_3 Year1996_t + I_h + \gamma' Z_{ht} + \varepsilon_{ht} \quad (2.3)$$

where we use the following set of outcomes: indicators for whether pensioner household h gives or receives monetary transfers to extended family and friends (the extensive margin), and the amounts of money given or received by pensioner household h (the intensive margin). In the vector of controls Z_{ht} we include interactions between geographical region and year, a linear and quadratic term for household income (excluding pensions) and the previously defined measures of household composition. I_h represents the household fixed effects. The coefficient of interest is β_1 which captures the effect of the crisis on the frequency and level of monetary transfers being sent or received by pensioner households.

2.4.4 Specification checks

In this section we present some specification checks on our identification strategy. The identification of the causal effect relies on the exogeneity of the crisis. First, we want to test whether the pension crisis was indeed unexpected. If the affected individuals had prior knowledge about the crisis, they might have adjusted their behavior to alleviate the effect of pension arrears prior to 1996. In such a case the post-crisis change in SWB would be underestimating the crisis impact. To find out if this is a cause of concern, we examine the difference between treatment and control groups in their last pre-crisis reports of expectations about future economic welfare and future life satisfaction.[18] We estimate the following model using data from 1995 – the last year before the

[18] As described in Section 2.2, the questions about future expectations only ask about the future 12 months, which is why we use the 1995 reports to see if the 1996 crisis was anticipated.

crisis:

$$y_{i1995} = \beta_0 + \beta_1 Arrears_i + \gamma' Z_{i1995} + \varepsilon_{i1995} \tag{2.4}$$

where y_{i1995} is one of the two measures of expectations, future life satisfaction and future economic welfare, $Arrears_i$ is an indicator equal to 1 if the individual is affected by the crisis in the following year 1996 and Z_{i1995} is the same set of controls defined further above. Given this specification, β_1 captures the difference in future expectations between our treatment and control groups one year before the crisis. A coefficient of zero would suggest that the 1996 pension crisis is not anticipated by those who are subsequently affected by it. Our results confirm that pensioners in arrears had no different pre-crisis expectations than pensioners not in arrears (if anything, the former had higher expectations). Non-pensioners in arrears had slightly lower pre-crisis expectations, but the difference is only marginally significant and only in the case of future life satisfaction (Table 2.4a). We interpret this as evidence that the crisis was unexpected for the pensioners who are affected by it personally. We cannot claim this with the same level of certainty for the non-pensioners who lived with these pensioners. However, any prior adjustment in expectations would mean that the results reported in Section 2.5.2 underestimate the crisis effects for the non-pensioner sample.

The validity of our strategy also depends on the assumption that there is no pre-existing difference in time trends between treatment and control groups. We examine this assumption by adding a separate pre-crisis time trend for the group in arrears and for the group in non-arrears:

$$\begin{aligned} y_{it} = {} & \beta_0 + \beta_1 Arrears_i * Year1995_t + \beta_2 Arrears_i * Year1996_t \\ & + \beta_3 Year1995_t + \beta_4 Year1996_t + I_i + \gamma' Z_{it} + \varepsilon_{it} \end{aligned} \tag{2.5}$$

where all variables are defined in the same way as in Equation (2.2). In this specification, the coefficient β_1 measures potential pre-crisis trend difference between treatment and control groups. We would reject the DD assumption if this coefficient were non-zero. The estimates are given in Tables 2.4b and 2.4c. All pre-crisis trend differences are economically and statistically insignificant for pensioners and, except for one marginally significant coefficient, for non-pensioners.

2.5 Results

2.5.1 The effect of the crisis on SWB of pensioners

Table 2.5 presents the estimates from Equation (2.1) for the pensioner sample. The odd-numbered columns contain the baseline model – the simple DD estimator without any controls – and the even-numbered columns contain the richer model that includes the vector of controls Z_{it}. The top row in the table gives the estimates of the crisis effects – the estimated coefficients on the interaction between arrears status and the crisis year 1996. Comparing the baseline and the richer model, we see that the inclusion of demographic and socio-economic characteristics makes little difference to the estimated effect of the crisis. The crisis has a significant effect on pensioners' life satisfaction: on the 5-point scale of life satisfaction, the crisis causes a decline of about .16 points. The crisis has even larger effects on expectations for the future: we estimate that both expectations for future life satisfaction and future economic welfare decline by about .22-.23 points (on a 5-point scale).

At the same time we find no significant effect on pensioners' subjective measure of health and small negative effects on pensioners' self-perceived relative rankings in society. One exception is the self-perceived power ranking which shows a marginally significant decline. Other determinants of SWB are as expected: married individuals, more educated individuals and those with higher pre-crisis household income tend to have higher levels of SWB, whereas those with low pre-crisis pensions, the unemployed and the disabled tend to have lower levels of SWB.

To put the magnitude of these estimates in a context, we compare the estimated effects of the crisis with the effects of some of the demographics. In particular, the crisis effect on life satisfaction is comparable to the effect of being unemployed or (in absolute value) the effects of being married or having a college degree. Note in particular that for all SWB measures the coefficient on arrears status during the 1996 crisis is close to 10 times as large in absolute value as the coefficient on pre-crisis household income. Household income is measured in 1000s of June 1992 rubles whereas an average pension in 1995 is around 1300 June 1992 rubles (Figure 2.2). A rough estimate shows that the loss of around 1300 rubles in the pension crisis appears to exert 10 times as large (in absolute

value) an effect on well-being as having extra 1000 rubles in household income. We consider this suggestive evidence that the pension crisis imposes a very large non-pecuniary cost for affected pensioners.[19]

The finding of substantial effects on LS and future expectations and no effect on SAH goes in line with our understanding of the former being flow quantities influenced by current levels of income and the latter being a stock quantity mostly unaffected by current levels of income. More surprising is our finding of a statistically insignificant effect on the pensioners' self-perceived economic ranking in society. One possible explanation might be that there was widespread decline in economic welfare during 1996 (as we describe earlier) and hence pensioners' relative economic standing in society was uninfluenced.

Results from our preferred specification which includes individual fixed effects are presented in Table 2.6. They confirm all our findings from the specification without fixed effects, and most of the crisis effects are even larger and more precisely estimated. In particular, the crisis effect on pensioners' life satisfaction is slightly larger (a decline of more than .18 points) and is now statistically significant at the 5% level. The self-perceived power ranking in society is also significant at the 5% level – the crisis causes a decline of about .19 points on the 9-point power scale. Moreover, the estimates change only imperceptibly when we include controls for post-crisis measures such as total household income other than pensions, net monetary transfers from extended family and friends and household composition. This suggests that allowing for a response to the crisis through other income sources or other household members has little to no effect on pensioners' experience of the crisis. In fact, we find little evidence of such response as discussed further below.

Table 2.7 shows how the size of the crisis effect depends on the total size of pension income lost,

[19]One reason might be that pensioners experience any loss of pension income more strongly that gains in household income – note that the coefficient on low pre-crisis pension is also much larger than the coefficient on pre-crisis household income (excluding pensions). Another reason is indicated by the significant crisis effects on future expectations. Namely, it might be that pensioners expected the crisis to last and so the crisis effects are large because they incorporate expected future pension losses. However, further investigation of the source of this non-pecuniary cost is beyond the scope of this paper.

measured by the number of pensioners in arrears. The odd-numbered columns show the estimates of the crisis effect when the crisis affects only 1 pensioner in the household and the even-numbered columns show the estimates when the crisis affects 2 or more pensioners in the household. These results demonstrate that the crisis effect does not rise substantially with the number of household pensions lost in the crisis. In fact, life satisfaction, the self-perceived economic and the self-perceived respect rankings all decline slightly less among pensioners in arrears when they are not the only ones in the household with lost pensions. For any of the measures, we cannot reject the hypothesis that losing just 1 pension in the household is as costly to SWB as losing more. We consider this another piece of suggestive evidence that the cost of the crisis cannot be fully measured with the loss of pension income.

Tables 2.8 and 2.9 examine other potential sources of heterogeneity in the effects of the crisis. In particular, we find that males and females report different experiences of the crisis along different aspects of life evaluation. The pension crisis causes somewhat greater declines in life satisfaction and self-perceived power ranking in society among male than among female pensioners. The two measures of future expectations, on the other hand, show little difference between the genders. We find that pre-crisis expectations also influence how pensioners experience the crisis. In particular, the crisis causes a greater decline in all measures of subjective well-being for those pensioners who had higher pre-crisis expectations. We believe this finding lends some support to the theory that prior expectations, and not just the current status quo, matter for experienced utility.

2.5.2 The effect of the crisis on SWB of non-pensioners

The crisis effect is not limited to the pensioners personally affected. We find that even non-pensioners experience worsening in SWB as a result of living in the same households with pensioners in arrears (Tables 2.10 and 2.11). For non-pensioners we find that living with a pensioner in arrears during the crisis causes a large and statistically significant decline in life satisfaction. In fact, we estimate a decline of over .20 points on the 5-point life satisfaction scale, which is slightly larger (though statistically indistinguishable) than the decline among pensioners. The strong crisis

effect on the life satisfaction of non-pensioners is quite robust to the inclusion not only of post-crisis household income and household composition, but also of post-crisis personal income. That the decline in life satisfaction is unaffected by the level of personal resources is again suggestive of some non-pecuniary cost of the pension crisis. The other measures of SWB appear to be affected less for non-pensioners, except for the three self-perceived rankings in society. These SWB measures exhibit large declines for non-pensioners but are measured much more imprecisely than in the larger pensioner sample. The self-perceived respect ranking shows the largest decline among the three measures and is also statistically significant at the 10% level (in the specification without individual fixed effects).

Similar to pensioners, we also find that the decline in SWB for non-pensioners does not rise substantively with the number of pensioners in arrears in the household (Table 2.12). Although the decline in life satisfaction is somewhat larger for non-pensioners who lost more than 1 pension in the household, the decline in self-perceived respect ranking is somewhat larger for non-pensioners who lost exactly one pension in the household. In fact, consistent with our findings for pensioners, the decline in each SWB measure is statistically indistinguishable between the group who lost several pensions in the household and the group who lost only one pension in the household.

Non-pensioners also show heterogeneity in their experiences of the crisis based on gender and pre-crisis expectations. Table 2.13 presents estimated effects of the pension crisis by gender. Unlike pensioners, the significant decline in life satisfaction among non-pensioners appears to be largely driven by women. The effect of pre-crisis expectations, however, is consistent across the pensioner and non-pensioner sample: non-pensioners who had higher pre-crisis expectations for the future also report greater decline in life satisfaction after the crisis (Table 2.14).

Our results indicate that the pensioner household cannot be analyzed as a homogenous unit when we examine the effect of the pension crisis. Pensioners and non-pensioners show similarities in the effect of the crisis on some aspects of subjective well-being (life satisfaction) but differences in others (future expectations). And even in those aspects of well-being where the crisis affects the two groups similarly, we find that women and men experience the crisis differently.

2.5.3 Robustness checks

In this section we describe a series of robustness checks to establish that our results are not sensitive to changes in the baseline specification we pick. All results are available upon request.

The first robustness check we perform is to test whether our results depend on the assumption that SWB is measured on a cardinal scale – an underlying assumption in our linear regression models. All of our results are qualitatively robust to estimating ordered probit models, which do not assume a cardinal scale.[20]

The second check is to recode each measure of SWB by collapsing them into binary 0-1 variables. For LS, SAH and the two measures of future expectations, we recode as 1 answers that are above 3 on the original 1-5 scales. As described in Section 2.2, answers above 3 correspond to strictly positive reports of current well-being or future expectations. For the three measures of societal rankings, we recode as 1 all answers that are above 4 on the original 1-9 scales. As described in Section 2.2, these answers correspond to reports of higher than average societal ranking. These alternative specifications produce the same qualitative results as the main specifications reported in the previous sections.

The third type of robustness checks we run is to test our sample selection. In our main analysis we exclude pensioners (and their non-pensioner household members) who report arrears in any year other than 1996 when the pension crisis occurred. The first check we do is to add these individuals to our sample and consider them part of our control group. We treat these individuals as controls because we are interested in the effect of the 1996 pension crisis and not in the general effect of failing to receive a pension. The second sample check we do is to expand our definition of pensioners to include people who receive early pensions due to early retirement because of serious illnesses or work in dangerous occupations. Once again, the main qualitative conclusions are robust to all of these robustness checks.

[20] As discussed previously, the ordered probit model does not allow for a straightforward inclusion of individual fixed effects. Hence, we test the cardinality assumption by estimating ordered probit models without fixed effects and compare the results with the results from OLS models without fixed effects (the latter are given in Tables 2.5 and 2.10).

2.6 Additional considerations

In this section we consider in turn the effects of the crisis that 1) go beyond measures of SWB, 2) go beyond pensioner households and 3) go beyond the immediate aftermath of the crisis.

In this paper we find that the pension crisis has a significant negative effect on several aspects of life evaluation both for the directly affected pensioners and for the other non-pensioner members of their households. Is this effect also mirrored in changes in the objective conditions of their lives? Jensen and Richter (2003) focus on the implications of the crisis on health-related measures of well-being. They find that pension crisis leads to a significant reduction in caloric intake, but has a more muted and varied response in terms of health outcomes. For example, medication use and doctor visits decline significantly for men but not for women, and physical limitations rise only marginally for men and not at all for women. These marginally significant effects on objective health reports are consistent with the insignificant effects we find on subjective health reports.

A possible explanation for the overall insignificant effect on health might be found in the changing composition of household expenditures. Jensen and Richter (2003) show that the crisis reduces calorie intakes. We explore this further. In particular, we find that households affected by the crisis spend a smaller share on meat and sugar and a larger share on home-grown produce, with the changes in the latter two shares being both economically and statistically significant (Table 2.15a).[21] It is possible that these changes in consumption patterns exert some positive effect on health that partially counters the negative effect from the loss of pension income.

Another health-related aspect of well-being that is relevant for the pension-aged is cognitive ability and particularly memory. Our interest in analyzing this aspect of well-being also lies in the suggestive evidence we reported earlier that the cost of the crisis is not entirely explained by the income loss from pension arrears. If some of the non-pecuniary cost is stress-related, it might be reflected in measures of memory. The RLMS survey includes tests of immediate memory recall and delayed memory recall. In the first set of tests, after reading ten words, the interviewer asks

[21] We find no significant change (both economically and statistically) in any other category of expenditure shares. All results are available upon request.

each respondent to recall as many words as possible. The immediate memory recall is measured by the number of words the respondent could recall immediately. After asking several other questions, the interviewer again asks the respondent to recall those ten words. The delayed memory recall is measured by the number of words the respondent could recall after responding to several other questions. We code those measures as 0-1 binary variables where 1 stands for having correctly memorized more than half of the words on the list. We find that the crisis has negative effects on both immediate and delayed memory recall for pensioners and an even larger negative effect on immediate memory recall for non-pensioners (Table 2.15b). While the effects on delayed memory are statistically insignificant, the negative effects on immediate memory recall are significant at the 10% level for both pensioners and non-pensioners.[22] These results are in line with the previously discussed evidence in favor of non-pecuniary costs from the pension crisis.

Overall, it appears that the crisis affected both objective and subjective measures of well-being, although the effects on the latter seem stronger. It is plausible that the measures of SWB capture elements of experienced utility which are not measured by objective circumstances. We believe that the inclusion of SWB measures in the analysis of the pension crisis paints a more complete picture of the consequences of such adverse events on the individuals personally affected and on the other members of their households.

The second theme we explore in this section is whether the monetary impact of the crisis extends beyond the pensioner households to other family members and friends linked through an informal network of inter-household monetary transfers. Our estimates of Equation (2.3) show that the pension crisis does not have a statistically significant effect on neither the frequency nor the size of monetary transfers in the network of extended family and friends; the income shock from the crisis appears to be borne entirely by pensioner households (Table 2.15c). These results are surprising because pensioners experience a significant monetary loss due to the pension crisis. One possible explanation might be that the pensioners made prior commitments, so that we do not

[22]Unfortunately these cognitive measures were asked only in 1995 and 1996 (and not in 1994) and were asked only for individuals aged 55 and above. It is possible that the power of the statistical tests is affected by the smaller sample size.

observe immediately an inter-household response to this crisis. Alternatively, if the pensioners' social network consists primarily of people who are also affected by the pension crisis, that would diminish the scope for inter-household insurance.[23]

The last theme we explore is the impact of the crisis beyond the immediate aftermath. In monetary terms the pension crisis was a one-time shock. Most pensioners went back to receiving their pensions in 1998, the first survey done after 1996, and the average pension adjusted for inflation is constant across survey rounds. Yet, does this necessarily imply that subjective well-being would return to its pre-crisis level once individual income recovers? Unfortunately, establishing the long-term effects of the pension crisis is problematic given that the first survey after 1996 is collected after the 1998 financial crisis. Hence, the interpretation of results which include the 1998 survey year relies on the assumption that the financial crisis does not differentially impact our treatment and control groups. We present these results with this caveat in mind.

We estimate the effect of the crisis after 1996 as follows:

$$y_{it} = \beta_0 + \beta_1 Arrears_i + \beta_2 Arrears_i * Year1996_t + \beta_3 Arrears_i * Year1998_t + \beta_5 Year1996_t + \beta_6 Year1998_t + I_i + \gamma' Z_{it} + \varepsilon_{it} \quad (2.6)$$

where we add an indicator for 1998 and its interaction with arrears status; all remaining variables are defined in the same way as in Equation (2.2). The coefficient of interest is β_3 which captures the difference in time trend between arrears and non-arrears groups in 1998, two years after the pension crisis. The estimates of this coefficient are given in Table 2.16. None of them are statistically significant for the pensioner sample. For the non-pensioners, the only two estimates that are statistically significant – the effects on future expectations – are in fact positive (Table 2.16). In other words, the trend in SWB over 1994-1998 is no worse for individuals affected by the 1996 pension crisis – either directly as pensioners in arrears or indirectly as living with pensioners in

[23] Also consistent with the hypothesis of limited scope for insurance is the lack of significant increase in the amount formally borrowed through loans or credits (results available upon request). It appears that the only insurance channel available during crisis was a decrease in the share of household resources formally lent to others (Table 2.15a).

arrears. As mentioned above, this short-lived effect could well be a result of the 1996 pension crisis being swamped by the larger 1998 financial crisis. This particular timing casts doubt on whether the short-lived effects we find here can be generalized to other shocks to the redistribution system. We cannot dismiss the possibility that had the financial crisis not occurred, the pension crisis effect would have persisted for longer.[24]

2.7 Discussion and conclusions

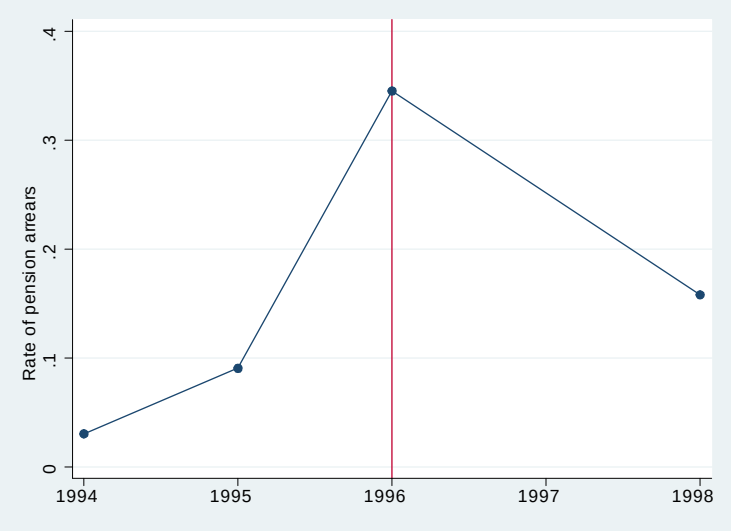
In this study we find that an exogenous shock to the redistribution system has a significant negative effect on the SWB of pensioners who were eligible for pensions but failed to receive them. The effect on well-being is stronger than predicted if the cost of the crisis were just the one-off monetary cost of pension loss. In particular, we find that pensioners' expectations for the future decline significantly and that the overall decline in SWB is just as strong for those who lose exactly 1 pension in the household as for those who lose more. The societal impact extends to non-pensioner members of these households. The reported decrease in well-being for non-pensioners remains unchanged even after accounting for their post-crisis personal income. In addition, mental well-being – measured by objective memory tests – also appears to decline by greater extent for individuals affected by the crisis – both pensioners and non-pensioners. One interpretation of all these results is that the pension crisis has a broader non-monetary effect on well-being. Such an effect has been documented for other life events such as unemployment (Luechinger et al. 2010). Indeed, our estimates show that being in pension arrears (regardless of whether experienced by oneself directly or by another household member) causes a decline in well-being of a similar size as that of being unemployed.

In conclusion, pensioner households appear to have borne the full monetary burden of pension loss caused by the 1996 crisis. We find no evidence of significant increase of monetary flows

[24] Jensen and Richter (2003) find that pensioners in arrears in 1996 have a higher rate of mortality in 1998 but it is uncertain whether even this effect can be attributed to the pension crisis alone – it might well be that it was the double shocks from the 1996 pension crisis and the 1998 financial crisis that led to higher mortality in 1998.

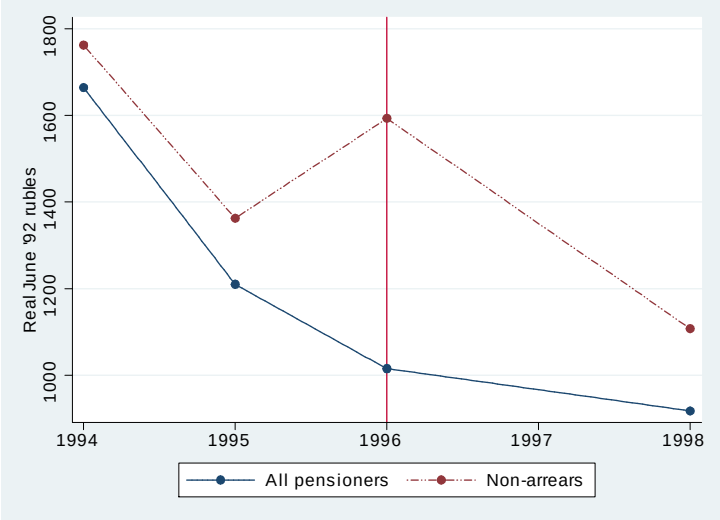
from extended family and friends. Within these pensioner households, we document effects on well-being that are consistent with significant additional non-pecuniary costs from the instability of the redistribution system. The crisis impact extends beyond the pensioners directly affected to the non-pensioners who live with them. Any public policies aimed to fully compensate for such disruptions in the redistribution system would need to be devised in ways that take these externalities into account.

Figure 2.1: Time trend in the rate of pension arrears



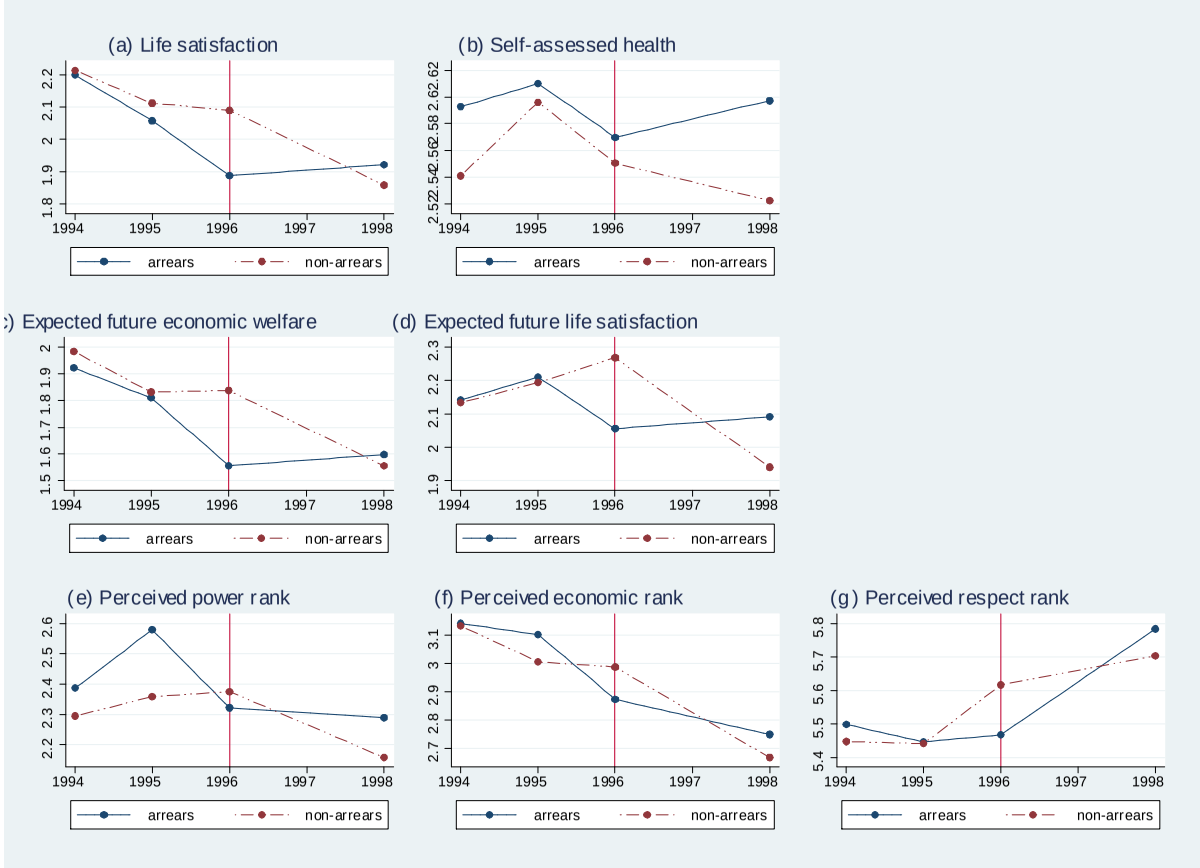
Notes: The figure plots the fraction of age-eligible pensioners in arrears by survey round.

Figure 2.2: Time trend in the amount of pension receipt



Notes: The solid line plots the average real pension income among all age-eligible pensioners by survey round. The dotted line plots the average real pension income among those who actually receive their pensions by survey round.

Figure 2.3: Time trends in measures of subjective well-being for pensioners



Notes: The figures plot the trends in all SWB measures for the pensioner sample. The solid lines represent individuals in pension arrears during the 1996 crisis, and the dotted lines represent individuals not in arrears during the crisis. For measures plotted in (a)-(d) the scale of responses is 1-5 and for measures plotted in (e)-(f) the scale is 1-9.

Table 2.1: Pre-crisis demographic characteristics by pension arrears status during the 1996 crisis

	Pensioners				Non-pensioners			
	Non-arrears		Arrears		Non-arrears		Arrears	
	Mean	SE	Mean	SE	Mean	SE	Mean	SE
Female	0.709	0.014	0.737	0.020	0.470	0.021	0.403	0.029
Married	0.590	0.015	0.588	0.022	0.594	0.020	0.604	0.029
College degree*	0.139	0.010	0.082	0.012	0.207	0.017	0.119	0.019
High school diploma	0.427	0.015	0.366	0.022	0.787	0.017	0.734	0.027
Years of age (in 1995)	65.293	0.243	64.831	0.363	35.133	0.526	34.874	0.774
Disabled	0.029	0.005	0.018	0.006	0.029	0.007	0.050	0.013
Unemployed	0.815	0.012	0.847	0.016	0.380	0.020	0.388	0.029
Personal income (000's)*					2.230	0.136	1.455	0.129
Pension income in bottom decile	0.130	0.010	0.218	0.019				
Hhld income w/o pensions (000's)*	2.059	0.120	1.495	0.123	5.143	0.248	3.498	0.231
Net monetary transfers from extended family (000's)	-0.279	0.074	-0.239	0.159	-0.117	0.047	-0.273	0.368
# of kids $<$ 7 years-old in hhld*	0.115	0.011	0.122	0.018	0.272	0.022	0.388	0.040
# of youths 7-18 years-old in hhld	0.164	0.014	0.161	0.021	0.535	0.029	0.496	0.043
# of adults in hhld	0.744	0.032	0.737	0.048	2.122	0.041	2.133	0.056
# of pensioners in hhld*	1.502	0.019	1.492	0.026	1.052	0.024	1.191	0.034
Rural area*	0.230	0.013	0.437	0.022	0.176	0.016	0.439	0.030
N	1089		490		581		278	

Notes: The pre-crisis demographic characteristics in bold are those that are statistically different at the 5% level between the pensioners in arrears and the pesioners not in arrears; and the asterisk indicate characteristics that are statistically different at 5% level between the non-pensioners living in households with at least one pensioner in arrears and those who are not. All income and monetary transfers are measured in 1000s of June 1992 rubles and calculated as the average of the two pre-crisis years, 1994 and 1995.

Table 2.2: Rates of pension arrears by region before and after the 1996 crisis

	1995		1996	
	Mean	SE	Mean	SE
Metropolitan: Moscow and St. Petersburg (both donor districts)	0.027	0.163	0.031	0.015
Northern and North Western (contains 1 donor district: Komi Republic)	0.057	0.232	0.327	0.048
Central and Central Black-Earth (only debtor districts)	0.091	0.288	0.285	0.023
Volga-Vaytski and Volga Basin (only debtor districts)	0.078	0.269	0.365	0.027
North Caucasian (only debtor districts)	0.142	0.349	0.526	0.032
Ural (only debtor districts)	0.080	0.271	0.302	0.033
Western Siberian (contains 1 donor district: Khanty-Mansi)	0.232	0.423	0.515	0.039
Eastern Siberian and Far Eastern (contains 1 donor district: Primorsky Krai)	0.000	0.000	0.309	0.038
Overall	0.090	0.006	0.345	0.011

Notes: Based on authors' own calculation using the RLMS. In parentheses we indicate whether the given geographical region contains administrative pension districts which were net donors of pension revenues during the 1996 pension crisis. In bold are those regions that do not experience a statistically significant increase at the 5% level in the rate of arrears between 1995 and 1996.

Table 2.3: Pre-crisis determinants of pension arrears during the 1996 crisis by region

Pre-crisis level	Region							
	1	2	3	4	5	6	7	8
Female	0.037	0.033	0.073	0.019	-0.058	0.044	-0.081**	0.059
	(0.034)	(0.103)	(0.045)	(0.045)	(0.072)	(0.026)	(0.017)	(0.025)
Married	-0.077	0.028	-0.015	-0.027	-0.025	-0.028	0.052	-0.068
	(0.048)	(0.193)	(0.066)	(0.082)	(0.203)	(0.126)	(0.074)	(0.100)
College	0.057	0.043	0.124	-0.001	-0.238*	-0.169	-0.359*	0.013
	(0.012)	(0.051)	(0.086)	(0.056)	(0.085)	(0.110)	(0.106)	(0.212)
High School	0.044*	0.015	0.091	-0.097	-0.153*	0.070	-0.035	-0.221
	(0.005)	(0.023)	(0.085)	(0.091)	(0.060)	(0.049)	(0.106)	(0.130)
Age	0.015	-0.202	-0.026	-0.037*	-0.000	-0.018	-0.004	-0.080
	(0.012)	(0.042)	(0.042)	(0.017)	(0.031)	(0.092)	(0.042)	(0.121)
Age squared (/100)	-1.137	16.591	1.863	2.447*	-0.186	0.332	0.131	5.251
	(0.916)	(2.919)	(2.933)	(1.073)	(2.300)	(6.536)	(3.382)	(8.781)
Disabled	-0.125	0.207	-0.010	-0.410**	0.117	-0.296	-0.230	-0.078
	(0.024)	(0.212)	(0.128)	(0.136)	(0.138)	(0.175)	(0.352)	(0.066)

Notes: Reported are coefficients from linear probability models of pension arrears status in 1996 on pre-crisis characteristics. Standard errors are clustered at the district-year level. All models also include a constant. Income is measured in 1000s of June 1992 rubles. *** p<0.01, ** p<0.05, * p<0.1

Table 2.3: Pre-crisis determinants of pension arrears during the 1996 crisis by region (continued)

Pre-crisis	Region							
level	1	2	3	4	5	6	7	8
Unemployed	0.041**	0.006	0.108	-0.018	0.074	0.053	-0.103*	-0.270
	(0.002)	(0.126)	(0.072)	(0.057)	(0.071)	(0.082)	(0.033)	(0.181)
Pension in bottom decile	0.274	-0.223	0.236**	0.013	0.046	0.060	-0.131**	-0.115**
	(0.101)	(0.122)	(0.083)	(0.059)	(0.099)	(0.137)	(0.029)	(0.022)
Other hhld income	0.003	0.009	-0.004	-0.011	-0.006	0.007	-0.008	-0.003
	(0.001)	(0.011)	(0.005)	(0.014)	(0.013)	(0.005)	(0.010)	(0.002)
# of 0-7 y/o in hhld	0.023	-0.040	0.114	-0.133	-0.052	-0.051	0.266***	-0.032
	(0.050)	(0.138)	(0.079)	(0.097)	(0.025)	(0.138)	(0.019)	(0.096)
# of 7-18 y/o in hhld	-0.007	-0.044	0.000	-0.050	0.037	-0.051	-0.033	0.118
	(0.015)	(0.015)	(0.047)	(0.056)	(0.079)	(0.091)	(0.031)	(0.089)
# of adults in hhld	-0.008	0.035	-0.018	-0.012	0.061*	0.002	-0.033	-0.036
	(0.014)	(0.027)	(0.019)	(0.055)	(0.023)	(0.024)	(0.023)	(0.049)
# of pensioners in hhld	0.034	0.062	-0.059	-0.045	-0.018	0.085	-0.009	0.095
	(0.012)	(0.053)	(0.046)	(0.064)	(0.087)	(0.084)	(0.027)	(0.070)
Rural area	-	0.503*	0.082	0.139	-0.107	0.107	0.354*	0.215
		(0.050)	(0.089)	(0.147)	(0.086)	(0.157)	(0.112)	(0.264)
N	125	88	369	337	205	198	123	134
R^2	0.154	0.403	0.082	0.060	0.057	0.087	0.254	0.126

Notes: Reported are coefficients from linear probability models of pension arrears status in 1996 on pre-crisis characteristics. Standard errors are clustered at the district-year level. All models also include a constant. Income is measured in 1000s of June 1992 rubles. *** p<0.01, ** p<0.05, * p<0.1

Table 2.4: Specification checks

(a) Pre-crisis levels of future expectations

	Future life satisfaction		Future economic welfare	
	(1) Pensioners	(2) Non-pensioners	(3) Pensioners	(4) Non-pensioners
Arrears	0.018	-0.187*	0.050	-0.046
	(0.054)	(0.097)	(0.063)	(0.093)
N	1,579	859	1,579	859
R^2	0.011	0.124	0.060	0.117

Notes: Reported are selected coefficients from OLS models specified in Equation (2.4) with standard errors clustered at district-year. *** p<0.01, ** p<0.05, * p<0.1

(b) Pre-crisis trends for pensioners

	Life satisfaction	Self-assessed health	Future life satisfaction	Future economic welfare
	(1)	(2)	(3)	(4)
Arrears* Year1995	-0.017	-0.029	0.056	0.053
	(0.070)	(0.027)	(0.079)	(0.091)
N	4,687	4,687	4,687	4,687
R^2	0.032	0.018	0.024	0.031

(b) Pre-crisis trends for pensioners (continued)

	Economic rank	Power rank	Respect rank
	(5)	(6)	(7)
Arrears* Year1995	0.115	0.066	-0.171
	(0.103)	(0.126)	(0.167)
N	4,687	4,687	4,687
R^2	0.027	0.022	0.022

Notes: Reported are selected coefficients from OLS models with individual fixed effects specified in Equation (2.5). *** p<0.01, ** p<0.05, * p<0.1

Table 2.4: Specification checks (continued)

(c) Pre-crisis trend for non-pensioners

	Life satisfaction	Self-assessed health	Future life satisfaction	Future economic welfare
	(1)	(2)	(3)	(4)
Arrears*	0.036	-0.031	-0.196*	0.031
	(0.140)	(0.076)	(0.112)	(0.110)
Year1995				
N	1,693	1,693	1,693	1,693
R^2	0.106	0.216	0.124	0.108

	Economic rank	Power rank	Respect rank
	(5)	(6)	(7)
Arrears*	0.081	0.250	0.051
Year1995	(0.158)	(0.188)	(0.227)
N	1,693	1,693	1,693
R^2	0.109	0.076	0.055

Notes: Reported are selected coefficients from OLS models with individual fixed effects specified in Equation (5). *** p<0.01, ** p<0.05, * p<0.1

Table 2.5: The effect of the 1996 crisis on SWB of pensioners (DD specifications)

	Life satisfaction		Self-assessed health	
	(1)	(2)	(3)	(4)
Arrears*	-0.167*	-0.159*	-0.014	-0.023
Year1996	(0.087)	(0.086)	(0.050)	(0.048)
Year1995	-0.115**	-0.168	0.043	0.069***
	(0.057)	(0.182)	(0.029)	(0.022)
Year1996	-0.131*	-0.183	0.004	-0.010
	(0.069)	(0.121)	(0.040)	(0.027)
Arrears	-0.034	0.003	0.033	0.041*
	(0.050)	(0.049)	(0.020)	(0.022)
Female		-0.046		-0.164***
		(0.033)		(0.025)
Married		0.114**		-0.078**
		(0.049)		(0.032)
College		0.126**		0.142***
		(0.054)		(0.029)
High school		0.056		0.083***
		(0.040)		(0.025)
Age		0.027		-0.004
		(0.031)		(0.018)
Age squared		-1.310		-1.386
(/100)		(2.269)		(1.306)
Disabled		-0.043		-0.406***
pre-crisis		(0.096)		(0.058)
Unemployed		-0.210***		-0.166***
pre-crisis		(0.045)		(0.027)
Low pension		-0.138***		-0.092***
pre-crisis		(0.045)		(0.034)
Hhold income		0.019***		0.003
pre-crisis		(0.006)		(0.002)
Rural		0.068		-0.010
		(0.044)		(0.029)
N	4,687	4,687	4,687	4,687
R^2	0.008	0.052	0.001	0.133

Table 2.5: The effect of the 1996 crisis on SWB of pensioners (continued)

	Future life satisfaction		Future economic welfare	
	(5)	(6)	(7)	(8)
Arrears*	-0.224**	-0.219**	-0.241***	-0.229***
Year1996	(0.100)	(0.093)	(0.085)	(0.076)
Year1995	0.063	-0.125	-0.139**	-0.113
	(0.056)	(0.191)	(0.065)	(0.190)
Year1996	0.136*	0.038	-0.140**	-0.129
	(0.071)	(0.162)	(0.066)	(0.120)
Arrears	0.012	0.004	-0.042	0.025
	(0.045)	(0.042)	(0.057)	(0.046)
Female		0.017		-0.244***
		(0.031)		(0.037)
Married		0.122**		0.128**
		(0.048)		(0.059)
College		0.013		0.331***
		(0.055)		(0.062)
High school		-0.013		0.111**
		(0.043)		(0.055)
Age		0.064***		-0.045
		(0.023)		(0.031)
Age squared		-4.072**		4.422**
(/100)		(1.622)		(2.204)
Disabled		-0.062		-0.282***
pre-crisis		(0.086)		(0.076)
Unemployed		-0.037		-0.100**
pre-crisis		(0.042)		(0.050)
Low pension		-0.086**		-0.126***
pre-crisis		(0.043)		(0.047)
Hhold income		0.007		0.021***
pre-crisis		(0.006)		(0.006)
Rural		0.117***		0.056
		(0.044)		(0.054)
N	4,687	4,687	4,687	4,687
R^2	0.004	0.019	0.010	0.065

Table 2.5: The effect of the 1996 crisis on SWB of pensioners (continued)

	Economic rank		Power rank		Respect rank	
	(9)	(10)	(11)	(12)	(13)	(14)
Arrears*	-0.166	-0.148	-0.209	-0.208*	-0.177	-0.131
Year1996	(0.118)	(0.111)	(0.130)	(0.117)	(0.183)	(0.149)
Year1995	-0.100	-0.091	0.104	-0.164**	-0.021	-0.050
	(0.084)	(0.095)	(0.089)	(0.081)	(0.170)	(0.185)
Year1996	-0.132	-0.072	0.101	-0.081	0.162	0.548***
	(0.090)	(0.100)	(0.093)	(0.067)	(0.169)	(0.185)
Arrears	0.052	0.075	0.156**	0.174***	0.028	0.061
	(0.068)	(0.058)	(0.071)	(0.064)	(0.106)	(0.086)
Female		-0.048		-0.043		0.087
		(0.045)		(0.047)		(0.059)
Married		0.274***		0.048		0.241***
		(0.072)		(0.058)		(0.084)
College		0.487***		0.380***		0.677***
		(0.077)		(0.091)		(0.128)
High school		0.394***		0.305***		0.384***
		(0.067)		(0.071)		(0.085)
Age		0.001		0.000		0.099*
		(0.036)		(0.035)		(0.059)
Age squared		0.592		0.376		-7.324*
(/100)		(2.547)		(2.485)		(4.288)
Disabled		0.078		0.061		0.012
pre-crisis		(0.130)		(0.120)		(0.200)
Unemployed		-0.259***		-0.263***		-0.277***
pre-crisis		(0.054)		(0.065)		(0.084)
Low pension		-0.287***		-0.127**		-0.356***
pre-crisis		(0.068)		(0.057)		(0.087)
Hhold income		0.028***		0.010*		0.004
pre-crisis		(0.010)		(0.006)		(0.011)
Rural		0.182***		0.193**		-0.074
		(0.062)		(0.076)		(0.100)
N	4,687	4,687	4,687	4,687	4,687	4,687
R^2	0.003	0.085	0.003	0.047	0.001	0.062

Reported are selected coefficients from OLS models specified in Equation (2.1) with standard errors clustered at district-year. Income is measured in 1000s of real rubles. Low pension pre-crisis is defined as having a pension income in the bottom decile of the pre-crisis pension income distribution.

Table 2.6: The effect of the 1996 crisis on SWB for pensioners (individual fixed-effect models)

	Life satisfaction		Self-assessed health	
	(1)	(2)	(3)	(4)
Arrears*Year1996	-0.187**	-0.183**	-0.024	-0.022
	(0.078)	(0.077)	(0.054)	(0.055)
Controls	No	Yes	No	Yes
N	4,687	4,687	4,687	4,687
R^2	0.028	0.033	0.013	0.015

	Future economic welfare		Future life satisfaction	
	(5)	(6)	(7)	(8)
Arrears*Year1996	-0.262***	-0.256***	-0.236***	-0.236***
	(0.075)	(0.076)	(0.058)	(0.057)
Controls	No	Yes	No	Yes
N	4,687	4,687	4,687	4,687
R^2	0.015	0.024	0.026	0.031

	Economic rank		Power rank		Respect rank	
	(9)	(10)	(11)	(12)	(13)	(14)
Arrears*Year1996	-0.128	-0.128	-0.188**	-0.185**	-0.176*	-0.170
	(0.088)	(0.088)	(0.088)	(0.088)	(0.101)	(0.102)
Controls	No	Yes	No	Yes	No	Yes
N	4,687	4,687	4,687	4,687	4,687	4,687
R^2	0.022	0.026	0.018	0.021	0.020	0.021

Notes: Reported are selected coefficients from OLS models with individual fixed effects specified in Equation (2.2). The even-numbered columns contain controls for current household income (excluding pensions), net transfers from extended family and friends and household composition. *** p<0.01, ** p<0.05, * p<0.1

Table 2.7: The effect of the 1996 crisis on SWB of pensioners by number of household arrears

	Life satisfaction		Self-assessed health	
	(1) Only 1 in arrears	(2) 2 or more in arrears	(3) Only 1 in arrears	(4) 2 or more in arrears
Arrears* Year1996	-0.204**	-0.167*	-0.043	0.006
	(0.083)	(0.096)	(0.057)	(0.062)
N	3,643	3,773	3,643	3,773
R^2	0.028	0.024	0.012	0.017

	Future life satisfaction		Future economic welfare	
	(5) Only 1 in arrears	(6) 2 or more in arrears	(7) Only 1 in arrears	(8) 2 or more in arrears
Arrears* Year1996	-0.222**	-0.323***	-0.172**	-0.345***
	(0.084)	(0.098)	(0.076)	(0.106)
N	3,643	3,773	3,643	3,773
R^2	0.016	0.017	0.024	0.022

	Economic rank		Power rank		Respect rank	
	(9) Only 1 in arrears	(10) 2 or more in arrears	(11) Only 1 in arrears	(12) 2 or more in arrears	(13) Only 1 in arrears	(14) 2 or more in arrears
Arrears* Year1996	-0.175*	-0.041	-0.130	-0.307**	-0.256*	-0.030
	(0.097)	(0.122)	(0.091)	(0.139)	(0.145)	(0.173)
N	3,643	3,773	3,643	3,773	3,643	3,773
R^2	0.021	0.021	0.012	0.019	0.020	0.023

Notes: Reported are selected coefficients from OLS models with individual fixed effects specified in Equation (2.2). In the odd-numbered columns the treatment group is restricted to pensioners who are the only ones in arrears in the household. In the even-numbered columns the treatment group is restricted to pensioners who have at least one other pensioner in arrears in the household besides themselves. In all cases the control group is pensioners who are not in arrears. *** p<0.01, ** p<0.05, * p<0.1

Table 2.8: The effect of the 1996 crisis on SWB of pensioners by gender

	Life satisfaction		Self-assessed health	
	(1) Men	(2) Women	(3) Men	(4) Women
Arrears* Year1996	-0.281**	-0.147	-0.012	-0.024
	(0.104)	(0.091)	(0.086)	(0.056)
N	1,333	3,354	1,333	3,354
R^2	0.040	0.039	0.060	0.020

	Future life satisfaction		Future economic welfare	
	(5) Men	(6) Women	(7) Men	(8) Women
Arrears* Year1996	-0.328***	-0.238***	-0.208*	-0.246***
	(0.115)	(0.082)	(0.109)	(0.059)
N	1,333	3,354	1,333	3,354
R^2	0.040	0.023	0.041	0.032

	Economic rank		Power rank		Respect rank	
	(9) Men	(10) Women	(11) Men	(12) Women	(13) Men	(14) Women
Arrears* Year1996	-0.203	-0.103	-0.396**	-0.107	-0.039	-0.210
	(0.190)	(0.094)	(0.176)	(0.102)	(0.188)	(0.138)
N	1,333	3,354	1,333	3,354	1,333	3,354
R^2	0.049	0.032	0.029	0.027	0.043	0.023

Notes: Reported are selected coefficients from OLS models with individual fixed effects specified in Equation (2.2). *** p<0.01, ** p<0.05, * p<0.1

Table 2.9: The effect of the 1996 crisis on SWB of pensioners by pre-crisis expectations

	Life satisfaction		Self-assessed health	
	(1)	(2)	(3)	(4)
	Low outlook	High outlook	Low outlook	High outlook
Arrears* Year1996	-0.156	-0.225*	0.009	-0.080
	(0.095)	(0.113)	(0.060)	(0.069)
N	2,832	1,855	2,832	1,855
R^2	0.030	0.064	0.019	0.033

	Economic rank		Power rank		Respect rank	
	(5)	(6)	(7)	(8)	(9)	(10)
	Low outlook	High outlook	Low outlook	High outlook	Low outlook	High outlook
Arrears* Year1996	-0.059	-0.228	-0.067	-0.352**	-0.062	-0.363**
	(0.132)	(0.200)	(0.123)	(0.143)	(0.147)	(0.173)
N	2,832	1,855	2,832	1,855	2,832	1,855
R^2	0.024	0.053	0.032	0.038	0.031	0.026

Notes: Reported are selected coefficients from OLS models with individual fixed effects specified in Equation (2.2). *** p<0.01, ** p<0.05, * p<0.1

Table 2.10: The effect of the 1996 crisis on SWB of non-pensioners (DD specifications)

	Life satisfaction		Self-assessed health	
	(1)	(2)	(3)	(4)
Arrears*	-0.249**	-0.224**	0.063	0.054
Year1996	(0.111)	(0.105)	(0.110)	(0.073)
Year1995	-0.073	0.127	0.051	0.009
	(0.077)	(0.091)	(0.055)	(0.144)
Year1996	-0.085	0.291***	0.037	-0.021
	(0.073)	(0.090)	(0.047)	(0.098)
Arrears	0.011	0.044	0.004	-0.027
	(0.072)	(0.068)	(0.055)	(0.039)
Female		-0.072		-0.248***
		(0.044)		(0.028)
Married		0.158***		0.041
		(0.045)		(0.032)
College		0.147***		0.111***
		(0.056)		(0.035)
High School		0.051		0.028
		(0.039)		(0.032)
Age		-0.105***		0.008
		(0.013)		(0.007)
Age squared		11.602***		-3.473***
(/100)		(1.683)		(0.917)
Disabled		-0.029		-0.972***
pre-crisis		(0.101)		(0.084)
Unemployed		-0.224***		0.019
pre-crisis		(0.050)		(0.031)
Personal income		-0.003		-0.015***
pre-crisis		(0.010)		(0.005)
Hhld income		0.025***		0.005
pre-crisis		(0.007)		(0.003)
Rural		-0.036		0.124***
		(0.064)		(0.043)
N	2,524	2,524	2,524	2,524
R^2	0.007	0.104	0.002	0.210

Notes: Reported are selected coefficients from OLS models specified in Equation (2.1) with standard errors clustered at district-year. Income is measured in 1000s of real rubles. *** p<0.01, ** p<0.05, * p<0.1

Table 2.10: The effect of the 1996 crisis on SWB of non-pensioners (continued)

	Future life satisfaction		Future economic welfare	
	(5)	(6)	(7)	(8)
Arrears*	0.129	0.090	-0.141	-0.137
Year1996	(0.177)	(0.132)	(0.112)	(0.103)
Year1995	0.054	-0.024	-0.049	-0.062
	(0.079)	(0.058)	(0.083)	(0.053)
Year1996	0.020	0.132	-0.054	-0.053*
	(0.072)	(0.167)	(0.084)	(0.030)
Arrears	-0.110	-0.085	-0.106*	-0.087
	(0.069)	(0.061)	(0.061)	(0.057)
Female		-0.135***		-0.150***
		(0.037)		(0.042)
Married		0.061		-0.010
		(0.046)		(0.054)
College		0.094		0.013
		(0.060)		(0.077)
High School		0.057		-0.006
		(0.046)		(0.051)
Age		-0.028**		-0.095***
		(0.012)		(0.012)
Age squared		0.322		10.235***
(/100)		(1.612)		(1.560)
Disabled		-0.191*		0.106
pre-crisis		(0.107)		(0.117)
Unemployed		0.064		-0.116**
pre-crisis		(0.043)		(0.048)
Personal income		0.008		0.020**
pre-crisis		(0.006)		(0.009)
Hhld income		0.012**		0.021***
pre-crisis		(0.005)		(0.006)
Rural		-0.102		0.062
		(0.067)		(0.061)
N	2,524	2,524	2,524	2,524
R^2	0.002	0.125	0.006	0.098

Notes: Reported are selected coefficients from OLS models specified in Equation (2.1) with standard errors clustered at district-year. Income is measured in 1000s of real rubles. *** p<0.01, ** p<0.05, * p<0

Table 2.10: The effect of the 1996 crisis on SWB of non-pensioners (continued)

	Economic rank		Power rank		Respect rank	
	(9)	(10)	(11)	(12)	(13)	(14)
Arrears*Year1996	-0.207	-0.236	-0.250	-0.215	-0.413*	-0.348*
	(0.183)	(0.150)	(0.243)	(0.174)	(0.244)	(0.205)
Year1995	-0.036	-0.004	0.134	0.184	-0.149	-0.101
	(0.084)	(0.088)	(0.118)	(0.279)	(0.123)	(0.257)
Year1996	0.091	0.561***	0.296**	0.712**	0.023	0.798***
	(0.109)	(0.163)	(0.122)	(0.287)	(0.164)	(0.197)
Arrears	0.043	0.032	0.143	0.083	0.020	-0.038
	(0.086)	(0.082)	(0.120)	(0.098)	(0.126)	(0.117)
Female		0.012		-0.106*		0.161**
		(0.058)		(0.061)		(0.063)
Married		0.283***		0.091		0.183**
		(0.073)		(0.077)		(0.087)
College		0.397***		0.554***		0.495***
		(0.100)		(0.116)		(0.113)
High school		0.235***		0.160**		0.229***
		(0.071)		(0.075)		(0.079)
Age		-0.120***		-0.092***		-0.070***
		(0.016)		(0.018)		(0.021)
Age squared (/100)		13.001***		8.929***		9.289***
		(2.126)		(2.428)		(2.631)
Disabled pre-crisis		-0.203		-0.347**		0.098
		(0.154)		(0.151)		(0.234)
Unemployed		-0.240***		-0.250***		-0.424***
pre-crisis		(0.059)		(0.073)		(0.102)
Personal income		0.013		0.013		0.007
pre-crisis		(0.011)		(0.012)		(0.010)
Hhld income		0.034***		0.003		0.021*
pre-crisis		(0.010)		(0.010)		(0.011)
Rural		0.073		0.038		-0.026
		(0.079)		(0.096)		(0.103)
N	2,524	2,524	2,524	2,524	2,524	2,524
R^2	0.001	0.107	0.004	0.081	0.004	0.074

Notes: Reported are selected coefficients from OLS models specified in Equation (2.1) with standard errors clustered at district-year. Income is measured in 1000s of real rubles. *** p<0.01, ** p<0.05, * p<0.

Table 2.11: The effect of the 1996 crisis on SWB for non-pensioners (individual fixed-effect models)

	Life satisfaction		Self-assessed health	
	(1)	(2)	(3)	(4)
Arrears*Year1996	-0.224***	-0.204**	0.055	0.060
	(0.078)	(0.084)	(0.050)	(0.051)
Controls	No	Yes	No	Yes
N	2,524	2,524	2,524	2,524
R^2	0.030	0.049	0.020	0.028

	Future economic welfare		Future life satisfaction	
	(5)	(6)	(7)	(8)
Arrears*Year1996	0.082	0.086	-0.109	-0.101
	(0.101)	(0.097)	(0.075)	(0.072)
Controls	No	Yes	No	Yes
N	2,524	2,524	2,524	2,524
R^2	0.016	0.026	0.012	0.030

	Economic rank		Power rank		Respect rank	
	(9)	(10)	(11)	(12)	(13)	(14)
Arrears*Year1996	-0.157	-0.134	-0.170	-0.161	-0.242	-0.226
	(0.098)	(0.102)	(0.140)	(0.136)	(0.181)	(0.193)
Controls	No	Yes	No	Yes	No	Yes
N	2,524	2,524	2,524	2,524	2,524	2,524
R^2	0.033	0.045	0.023	0.034	0.037	0.044

Notes: Reported are selected coefficients from OLS models with individual fixed effects specified in Equation (2.2). The even-numbered columns contain controls for current household income (excluding pensions), net transfers from extended family and friends and household composition. *** p<0.01, ** p<0.05, * p<0.1

Table 2.12: The effect of the 1996 crisis on SWB of non-pensioners by number of household arrears

	Life satisfaction		Self-assessed health	
	(1) Only 1 in arrears	(2) 2 or more in arrears	(3) Only 1 in arrears	(4) 2 or more in arrears
Arrears* Year1996	-0.183*	-0.264*	0.056	0.090
	(0.095)	(0.134)	(0.056)	(0.077)
N	1,667	2,008	1,667	2,008
R^2	0.031	0.035	0.024	0.027

	Future life satisfaction		Future economic welfare	
	(5) Only 1 in arrears	(6) 2 or more in arrears	(7) Only 1 in arrears	(8) 2 or more in arrears
Arrears* Year1996	-0.032	0.206	-0.130	-0.149
	(0.118)	(0.170)	(0.101)	(0.169)
N	1,667	2,008	1,667	2,008
R^2	0.023	0.024	0.018	0.016

	Economic rank		Power rank		Respect rank	
	(9) Only 1 in arrears	(10) 2 or more in arrears	(11) Only 1 in arrears	(12) 2 or more in arrears	(13) Only 1 in arrears	(14) 2 or more in arrears
Arrears* Year1996	-0.004	-0.489**	-0.199	-0.219	-0.219	-0.055
	(0.113)	(0.188)	(0.160)	(0.240)	(0.214)	(0.278)
N	1,667	2,008	1,667	2,008	1,667	2,008
R^2	0.038	0.035	0.033	0.035	0.043	0.050

Notes: Reported are selected coefficients from OLS models with individual fixed effects specified in Equation (2.2). In the odd-numbered columns the treatment group is restricted to pensioners who are the only ones in arrears in the household. In the even-numbered columns the treatment group is restricted to pensioners who have at least one other pensioner in arrears in the household besides themselves. In all cases the control group is pensioners who are not in arrears. *** p<0.01, ** p<0.05, * p<0.1

Table 2.13: The effect of the 1996 crisis on SWB of non-pensioners by gender

	Life satisfaction		Self-assessed health	
	(1) Men	(2) Women	(3) Men	(4) Women
Arrears* Year1996	-0.165	-0.238**	0.111*	0.008
	(0.127)	(0.111)	(0.062)	(0.072)
N	1,399	1,125	1,399	1,125
R^2	0.058	0.077	0.026	0.058

	Future life satisfaction		Future economic welfare	
	(5) Men	(6) Women	(7) Men	(8) Women
Arrears* Year1996	0.057	0.142	-0.198*	0.030
	(0.116)	(0.130)	(0.101)	(0.131)
N	1,399	1,125	1,399	1,125
R^2	0.030	0.054	0.048	0.053

	Economic rank		Power rank		Respect rank	
	(9) Men	(10) Women	(11) Men	(12) Women	(13) Men	(14) Women
Arrears* Year1996	-0.123	-0.144	-0.248	-0.058	-0.148	-0.303
	(0.113)	(0.151)	(0.187)	(0.154)	(0.207)	(0.249)
N	1,399	1,125	1,399	1,125	1,399	1,125
R^2	0.072	0.052	0.048	0.057	0.046	0.079

Notes: Reported are selected coefficients from OLS models with individual fixed effects specified in Equation (2.2). *** p<0.01, ** p<0.05, * p<0.1

Table 2.14: The effect of the 1996 crisis on SWB of non-pensioners by pre-crisis expectations

	Life satisfaction		Self-assessed health	
	(1) Low outlook	(2) High outlook	(3) Low outlook	(4) High outlook
Arrears* Year1996	-0.148	-0.255**	0.061	0.061
	(0.122)	(0.097)	(0.053)	(0.081)
N	1,335	1,189	1,335	1,189
R^2	0.069	0.085	0.053	0.038

	Economic rank		Power rank		Respect rank	
	(5) Low outlook	(6) High outlook	(7) Low outlook	(8) High outlook	(9) Low outlook	(10) High outlook
Arrears* Year1996	0.008	-0.298*	0.042	-0.416*	-0.228	-0.174
	(0.131)	(0.169)	(0.205)	(0.238)	(0.204)	(0.270)
N	1,335	1,189	1,335	1,189	1,335	1,189
R^2	0.071	0.080	0.046	0.076	0.056	0.067

Notes: Reported are selected coefficients from OLS models with individual fixed effects specified in Equation (2.2). *** p<0.01, ** p<0.05, * p<0.1

Table 2.15: The effect of the 1996 crisis on objective measures of well-being

(a) Household expenditure ratios

	Home production	Meat	Sugar	Lending
	(1)	(2)	(3)	(4)
Arrears* Year1996	0.018*	-0.011	-0.017**	-0.023**
	(0.010)	(0.012)	(0.008)	(0.009)
N	3,000	3,000	3,000	3,000
R^2	0.054	0.023	0.029	0.011

Notes: Reported are selected coefficients from OLS models estimated at the household level with household fixed effects as specified in Equation (2.3). *** p<0.01, ** p<0.05, * p<0.1

(b) Cognitive measures

	Immediate memory recall		Delayed memory recall	
	(1) Pensioners	(2) Non-pensioners	(3) Pensioners	(4) Non-pensioners
Arrears* Year1996	-0.043*	-0.208*	-0.043	0.043
	(0.024)	(0.118)	(0.031)	(0.101)
N	2,874	931	2,662	913
R^2	0.027	0.164	0.019	0.051

Notes: Reported are selected coefficients from OLS models with individual fixed effects specified in Equation (2.2). Memory measures are coded as 0-1 binary variables. 1 corresponds to having remembered more than half of the words in the list that the respondent is asked to memorize. These questions are asked in 1995 and 1996 (not in 1994) of respondents above the age of 54. *** p<0.01, ** p<0.05, * p<0.1

(c) Household transfers from extended family and friends

	Received any	Given any	Amount received	Amount given
	(1)	(2)	(3)	(4)
Arrears* Year1996	-0.018	-0.028	-0.050	-0.043
	(0.025)	(0.036)	(0.108)	(0.081)
N	3,636	3,636	3,636	3,636
R^2	0.023	0.038	0.021	0.064

Notes: Reported are selected coefficients from OLS models estimated at the household level with household fixed effects as specified in Equation (2.3). *** p<0.01, ** p<0.05, * p<0.

Table 2.16: Duration of the effect of the 1996 crisis

(a) Pensioners

	Life satisfaction	Self-assessed health	Future life satisfaction	Future economic welfare
	(1)	(2)	(3)	(4)
Arrears*	0.054	0.023	0.095	0.039
Year1998	(0.065)	(0.041)	(0.081)	(0.061)
N	5,855	5,855	5,855	5,855
R^2	0.038	0.022	0.027	0.043

	Economic rank	Power rank	Respect rank
	(5)	(6)	(7)
Arrears*	0.016	-0.135	-0.020
Year1998	(0.092)	(0.082)	(0.135)
N	5,855	5,855	5,855
R^2	0.039	0.022	0.017

Notes: Reported are selected coefficients from OLS models with individual fixed effects specified in Equation (2.6). *** p<0.01, ** p<0.05, * p<0.1

Table 2.16: Duration of the effect of the 1996 crisis (continued)

(b) Non-pensioners

	Life satisfaction	Self-assessed health	Future life satisfaction	Future economic welfare
	(1)	(2)	(3)	(4)
Arrears*	0.005	0.033	0.377***	0.246***
Year1998	(0.085)	(0.082)	(0.101)	(0.077)
N	3,183	3,183	3,183	3,183
R^2	0.043	0.028	0.038	0.028

	Economic rank	Power rank	Respect rank
	(5)	(6)	(7)
Arrears*	0.122	0.114	0.045
Year1998	(0.163)	(0.134)	(0.207)
N	3,183	3,183	3,183
R^2	0.046	0.028	0.032

Notes: Reported are selected coefficients from OLS models with individual fixed effects specified in Equation (2.6). *** p<0.01, ** p<0.05, * p<0.1

Chapter 3

Human Capital and Wage Inequality during Transition: Evidence from Bulgaria[1]

3.1 Introduction

Bulgaria's political transition from communism to democracy began in November 1989 and ended promptly and peacefully with democratic elections in June 1990. By contrast, Bulgaria's economic transition from a planned to a market economy has been prolonged and difficult. Per capita GDP returned to its pre-transition 1989 level only in 2007. The unemployment rate, virtually zero under communism, reached double digits in 1991 and remained there until 2006.

Under communism, Bulgaria's education system copied the Soviet model by emphasizing technical training adapted to the industrial and agricultural requirements of the Soviet bloc market. When that market collapsed, the human capital acquired under the old system did not fit the new market economy in which Bulgarian industry had to face Western competition. The problem of obsolete human capital has made education reform a priority for every Bulgarian government since 1990.

In transitioning from a planned to a market economy, the Bulgarian government has exercised

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less and less control over the wage distribution. The government no longer mandates parity between men's and women's wages, restricts labor market access for minorities, or stipulates earnings profiles across sectors. Wage distribution since 1990 has been driven by the internal demand of new private firms and by the external demand of increased trade and capital flows from Western Europe.

While there have been many studies on returns to human capital in developed and developing countries,[2] there have been relatively few studies on this topic specifically in the context of Eastern Europe. One problem is the lack of reliable data, or sometimes any data, for the region before 1990. Studies on the region have focused on the more developed transition economies because more data are available for these countries. Poorer countries, such as Bulgaria and Romania, have been largely ignored. Yet comparative studies (e.g. Flabbi, Paternostro and Tiongson 2007) show that there are significant differences in returns to human capital among Eastern European countries. Likewise, studies on wage inequality across the region (e.g. Rutkowski 2001) show that, in this respect also, the economic transition has affected former planned economies quite differently. Research on Bulgaria in particular has focused on returns to human capital only in a given year, and little attention has been paid to the relationship between these changing returns and the growth in wage inequality during Bulgaria's transition.

It is this gap in the literature that this chapter helps fill. I focus on Bulgaria for two reasons. First, studying Bulgaria is useful for understanding the transitions of the other poorer and less researched countries in the region as well. Moreover, among these poorer countries, Bulgaria stands out because its economy and education system had been most controlled by Moscow. As a result, Bulgaria experienced the greatest transformation of its human capital and wage structure after the collapse of the Soviet bloc market. Bulgaria, therefore, provides a fertile context in which to study returns to human capital and wage inequality during transition.

In this chapter, I use data from household surveys and study two early years of transition, 1995 and 1997, and two later years, 2001 and 2003. I find that returns to education increase but returns

[2]For a survey of such studies, see for example Psacharopoulos (1994).

to work experience stagnate. Returns to education are not significantly affected by an individual's gender, ethnicity, or sector of employment, but are significantly affected by an individual's age – in particular by whether an individual completed schooling during communism rather than after. Returns to education also differ between age groups, younger cohorts receiving higher returns than older ones. Finally, growing returns to education explain most of the growth in wage inequality during Bulgaria's transition.

3.2 Data

I use four Bulgarian Integrated Household Surveys conducted in 1995, 1997, 2001, and 2003, as part of the Living Standards Measurement Study of the World Bank. The first three surveys were conducted by Gallup International and the last one by the National Statistical Institute of Bulgaria.

Table 3.1 presents the main variables I use. I restrict my sample to hired workers between sixteen and sixty-five years old.[3] I use gross monthly wages since almost all Bulgarian wages are paid monthly. As each survey was conducted over several months, I deflate gross monthly wages by the monthly price indices calculated by the National Statistical Institute of Bulgaria.[4] I transform all wages to constant 2000 Bulgarian Leva.

As Table 3.1 shows, real monthly wages did not rise between 1995 and 2003. Real wages fell sharply between 1995 and 1997, a result of the 1996-1997 Bulgarian financial and currency crisis during which inflation climbed to 580%. This wage trend is consistent with the total wage bill from the Bulgarian national accounts data. Moreover, although 1995 was six years after the end of the planned economy, Table 3.1 shows that in 1995 only 17% of the sample worked in the private sector. By 2003, the private sector employed a majority – 61%.

Table 3.1 also shows that years of schooling in Bulgaria did not rise appreciably between 1995

[3]Sixteen is the earliest age at which Bulgarian children can leave school. Sixty-five is the age by which most Bulgarians retire. None of the results in Section 3.4 are sensitive to these age limits. In particular, raising the lower limit to eighteen or reducing the upper limit to sixty-three (the legal minimum age for retirement) leaves the results unchanged.

[4]Deflating by daily price indices is unnecessary because wages are paid at the end of the month.

and 2003. Under communism, Bulgaria already had high completion rates for secondary schools. As the dummy variable "Pre-1990 education" indicates, however, the kind of education received by Bulgarians has changed. This variable indicates whether individuals reported having completed schooling by 1990. In 1995, 90% of the sample had been educated under communism; in 2003, 69%.

Finally, Table 3.1 shows that the other demographic characteristics – mean age, work experience, proportion of women, proportion of ethnic minorities – were fairly constant across the survey years.

3.3 Empirical methods

3.3.1 Estimation of returns to human capital

I estimate returns to human capital using Mincer's (1974) earnings equation with additional covariates:

$$\begin{aligned}\ln Wage_{it} = {} & \beta_0 + \beta_1 Schooling_{it} + \beta_2 Experience_{it} + \beta_3 Experience\ squared_{it} \\ & + \beta_4 \mathrm{Pr}ivate_{it} + \beta_5 Female_{it} + \beta_6 Minority_{it} + \sum_{j=7}^{15} \beta_j \mathrm{Re}gion_{it\,j} + \varepsilon_{it}\end{aligned} \tag{3.1}$$

where $Wage_{it}$ is the gross monthly wages of individual i at time t; $Schooling_{it}$ is years of schooling; $Experience_{it}$ is years of potential work experience, defined as age minus years of schooling minus seven;[5] $\mathrm{Pr}ivate_{it}$ is a dummy variable for working in the private sector; $Female_{it}$ is a dummy variable for being female; $Minority_{it}$ is a dummy variable for being an ethnic minority;[6] and $\mathrm{Re}gion_{it\,j}$ is a dummy variable for region j, where j is one of the nine administrative regions in Bulgaria. Thus the dummy variables for gender and ethnicity account for personal character-

[5] In Bulgaria, children start school at age seven.

[6] Mostly ethnic Turks and Roma.

istics, while the dummy variables for private sector employment and region account for external economic characteristics.[7] I control for these characteristics because they likely affect both wages and the quantity and quality of human capital.[8] I estimate Equation (3.1) by ordinary least squares (OLS) for each sample year t and cluster the standard errors by household.[9]

One concern with estimating Equation (3.1) by OLS is that β_1 – the coefficient on schooling – is biased because the equation ignores individuals' abilities, and that the extent of the bias depends on the correlation between schooling and ability. To deal with this endogeneity bias, many instrumental variable approaches have been taken.[10] Given the limitations of my data, however, only two such approaches are possible here – using a change in laws that exogenously affects how many years of schooling some individuals receive (e.g. Angrist and Krueger 1991), or using parents' education as an instrument for schooling (e.g. Trostel, Walker and Woolley 2002). As I note in Section 3.2, however, the level of education in Bulgaria did not rise appreciably during the survey years. The Bulgarian education reforms of the 1990s affected education not in terms of years of schooling, but less measurably through changes in the curriculum. Moreover, although my data include parents' education, constructing such an instrumental variable would not eliminate the endogeneity bias. Studies on social capital in Bulgaria (e.g. Mitev 2005) have found that most Bulgarians find jobs through the network of extended family: in particular, parents often arrange for children to find jobs with relatives or colleagues of relatives. Hence parents' education affects their

[7] There could be a concern that by including the private sector dummy in Equation (3.1), β_1 no longer captures a potentially interesting component of returns to education. More schooling could be valuable if it helps in finding a job in the public sector, which on average pays higher salaries than the private sector (see Table 3.2). Omitting the private sector dummy, however, does not change the coefficient on schooling (or any other coefficient). The other alternative – examining the public and private sectors separately – is done in Section 3.4.1.

[8] I omit hours worked as a covariate because data were missing for many individuals and because the question was not consistently phrased across the survey years. Nevertheless, restricting the sample to those who reported hours worked and including hours worked as a covariate leave the results unchanged. This is because only a small fraction reported not working a forty-hour week, a fact that reflects the inflexibility of a labor market in Bulgaria with few part-time jobs.

[9] Clustering at a higher level, such as a district to account for the survey sample design, leaves the results unchanged.

[10] For a survey of such studies, see for example Card (1999).

children's earnings not only by its correlation with the children's education. Therefore, given the limitations of my data and given the biases that imperfect instrumental variables would introduce, I estimate Equation (3.1) by OLS.[11]

A second concern with estimating Equation (3.1) by OLS is that there may be selection into employment. The main approach taken to deal with this selection bias is the Heckman correction (1979). Indeed, studies on returns to human capital in Eastern Europe have found selection bias in some countries (Arandarenko, Kotzeva and Pauna 2006). In one study on Bulgaria (Dimova and Gang 2004), however, controlling for selection bias yields estimates of returns to human capital similar to OLS estimates. In these studies, the Heckman selection equation uses as exogenous variables the number of children, the number of disabled household members, and unemployment benefits – variables considered determinants of employment and not of wages. Yet even if estimates of returns to human capital are not biased from such general selection issues, there may still be selection bias due to the period I focus on. In particular, there may be individuals selected into unemployment after 1990 because they had human capital specific to the planned economy and difficult to transfer to the new market economy. To account for this selection bias, I use the 1995 survey, which includes data on what job an individual had in 1989. I construct a dummy variable for the selection equation indicating whether an individual worked either on a cooperative farm or at an industrial job in 1989. Bulgaria's agricultural sector experienced an especially hard transition because of political stalemate over land, while Bulgaria's industrial output fell by about 40% in the early years of transition. Thus among individuals working on cooperative farms or in the state-owned industries in 1989, many were likely selected into unemployment during transition. Had they been employed, their returns to human capital might have been lower than those of individuals who were actually employed. I account for this bias in a two-step Heckman regression.[12]

[11]Meta-analyses have found only slight differences in results when using different estimation methods (Ashenfelter, Harmon and Oosterbeek 1999).

[12]However, this approach to control for selection is limited as I have data on employment in 1989 only for individuals in 1995 and not for individuals in the subsequent samples. Moreover, this control by design can help with selection only among the older individuals – those that were already in the labor force by 1989 and who I consider most likely to have been affected by selection

Finally, I examine whether the estimate of the earnings equation (3.1) differs among demographic groups defined by gender, ethnicity, sector of employment, or age. Age in particular may affect returns to education in several ways. I examine differences in returns to education between individuals who completed schooling during communism and those who completed it after. I also examine differences in returns to education between cohorts in a survey year and how returns to education change for each cohort as it ages over the survey years.

3.3.2 Estimation of wage inequality

Next, I examine the relationship between returns to human capital and wage inequality during Bulgaria's transition. Admittedly, wage inequality captures only one part of the overall household income inequality. Transition economies experienced great changes from the communist to the post-communist period in the manner and level of public goods provision such as health and social services and thus studying measures that capture the evolution of the overall household income inequality during transition is a very vital issue. In this chapter I focus on wage inequality because wage data present much smaller measurement error problems than the valuation of public services' provision. Moreover, household composition in Bulgaria has remained virtually the same in the communist period and during the transition (e.g. Stoyanova and Kirova 2008). Thus at least in this respect I do not miss any important inequality dynamics due to changes in household structure.

I first calculate three measures of wage inequality: the variance of the logarithm of wages, the Gini coefficient, and the Theil-T index. To calculate the latter two, I use the formulas defined in Sen (1997):

$$G = 1 + \frac{1}{n} - \frac{2}{n^2 \mu} \sum_{i=1}^{n} i w_i \text{ for } w_1 \geq w_2 \geq ... \geq w_n \tag{3.2}$$

during the transition. Nevertheless, my main focus in this paper is to investigate the transition experience of Bulgarian workers who were actually employed in the post-1990 period. I use these selection-controlled estimates only for comparison with the OLS estimates.

$$T = \frac{1}{n}\sum_{i=1}^{n}\frac{w_i}{\mu}\ln(\frac{w_i}{\mu}) \tag{3.3}$$

where i indexes individual i out of sample size n, w_i is the individual's wage, and μ is the mean wage in the sample.

I calculate these three measures of wage inequality as they are common in the literature; thus I can compare my results to existing results on wage inequality in transition economies. However, as all three measures show the same trend of growing wage inequality during Bulgaria's transition, for decompositions of wage inequality I use the variance of log wages, which also has the additional advantage of being an additively decomposable measure.

I first decompose overall wage inequality in each survey year into inequality within groups and inequality between groups – the groups defined, as in the analysis of returns to human capital, by schooling, age, gender, ethnicity, and sector of employment. I use the formula

$$\sigma^2_{overall} = \sigma^2_{between\ groups} + \sum_{k=1}^{K}\rho_k\sigma^2_{group\ k} \tag{3.4}$$

where $\sigma^2_{overall}$ is wage inequality overall; $\sigma^2_{between\ groups}$ is wage inequality between groups, calculated by assigning every individual within a group the mean wage[13] of the group; and $\sum_{k=1}^{K}\rho_k\sigma^2_{group\ k}$ is wage inequality within groups, calculated by weighting wage inequality within each group k by ρ_k, the population share of group k, and summing for all groups.

Next, I decompose the change in wage inequality between the first and last survey years. I perform a semi-parametric decomposition (DiNardo, Fortin and Lemieux 1996) to determine how much of the growth in wage inequality between 1995 and 2003 is due to changes in sample characteristics. As I find that changes in sample characteristics contribute marginally to the growth in wage inequality, I move on to a parametric decomposition using my results on returns to human capital. I decompose the growth in wage inequality using the formula given by Yun (2006), which

[13]For the variance of log wages to be additively decomposable the geometric mean needs to be used in the between-group component.

combines the Juhn, Murphy and Pierce (1993) method of decomposition into aggregate characteristics, coefficients and residuals with the Fields (2003) method of decomposition into individual factors:

$$\sigma^2_{y^{2003}} - \sigma^2_{y^{1995}} = \sum_{k=1}^{K} (s_{ky^{2003*}} \sigma^2_{y^{2003*}} - s_{ky^{1995}} \sigma^2_{y^{1995}}) + \sum_{k=1}^{K} (s_{ky^{2003}} \sigma^2_{y^{2003}} - s_{ky^{2003*}} \sigma^2_{y^{2003*}}) + (\sigma^2_{e^{2003}} - \sigma^2_{e^{1995}}) \quad (3.5)$$

where $s_{ky} = \frac{\beta_k \sigma_{x_k} \rho_{x_k,y}}{\sigma y}$; $y^{2003*} = \beta_0^{1995} + \sum_{k=1}^{K} \beta_k^{1995} x_k^{2003} + e^{2003}$; y is the log wages in the given year; σ_x is the standard deviation of regressor x (e.g. schooling or work experience) used in the earnings equation (3.1); $\rho_{x,y}$ is the correlation between x and y; β is the coefficient (e.g. returns to work experience) I estimate by OLS; and e is the OLS residual. Fields (2003) refers to s_{ky} as the relative factor inequality weight for factor k (e.g. schooling). It measures the percentage of wage inequality in the given year accounted for by that factor. Juhn et al. (1993) refer to y^{2003*} as an auxiliary earnings measure. It computes counterfactual earnings in 2003: what earnings in 2003 would have been if returns to individual characteristics remained at their 1995 levels but the characteristics themselves and the unobservables were at their actual 2003 levels. Therefore, the difference between earnings in counterfactual 2003 and in 1995 captures the effect of changes in sample characteristics and unobservables whereas the difference between earnings in actual 2003 and in counterfactual 2003 captures the effect of changes in returns to characteristics.

Thus the first term on the right-hand side of Equation (3.5) is the growth in wage inequality between 1995 and 2003 that can be attributed to the sum of changes in each sample characteristic k (one of the wage determinants in the earnings equation (3.1)); the second term is the growth in wage inequality that can be attributed to the sum of changes in returns to each characteristic k; and the third term is the growth in wage inequality that can be attributed to changes in unobservables or residuals variance.

3.4 Results

3.4.1 Returns to human capital

Table 3.2 presents the estimate of the earnings equation (3.1) for each of the four survey years. In each successive survey year, returns to education increase: an additional year of schooling increases wages by 4.6% in 1995 and by 7.1% in 2003.[14] These estimates, moreover, are consistent with previous studies showing that returns to education in Bulgaria are lower than in the more developed transition economies in Central Europe. Flabbi et al. (2007) find that, between 1995 and 2003, returns to education in Hungary grew from 8.8% to 11.1%; in Poland, from 7.9% to 10.6%.

Table 3.2 also shows that returns to work experience stagnate around 2.5% to 2.9%. This stagnation appears specific to Bulgaria. Orazem and Vodopivec (1994) find that returns to work experience increased in Slovenia after communism, while Krueger and Pischke (1992) find that returns to work experience decreased in East Germany. Thus among transition economies, the trend in returns to work experience varies much more than the trend in returns to education. This is perhaps due to the greater variety across Eastern Europe in the pattern of economic restructuring than in that of educational reforms. Work experience acquired under communism is valued differently over the course of transition depending on the specifics of a particular country's economic transformation.

Table 3.2 shows that gender and ethnicity play large roles in determining wages. In each survey year, women and ethnic minorities earn less at all levels of schooling and work experience. While the wage gap for women shrinks over the later survey years, in 2003 the gap is still 28%. The wage gap for ethnic minorities shows no clear trend.

Table 3.2 shows that sector of employment also plays a role in determining wages. In each

[14] Only the 2003 estimate, however, is different from the 1995 and 1997 estimates at the 5% significance level. Note that the 1995 and 1997 samples are not independent because some of the 1995 individuals were re-interviewed and included in the 1997 sample. Given this fact I calculate the standard error of the difference between the 1995 and 1997 estimates and find this difference to be statistically insignificant.

survey year, individuals working in the private rather than public sector earn less at all levels of schooling and work experience. The wage gap between the private and public sectors is especially large in 1997, when the Bulgarian financial crisis destroyed the banking sector. It appears that the resulting credit crunch allowed private sector employers to let wages fall further behind those of the public sector.

Although the estimate of Equation (3.1) shows that gender, ethnicity, and sector of employment have statistically significant effects on wages, their effects on returns to human capital are for the most part not statistically significant. Thus I do not present estimates of Equation (3.1) separately based on gender, ethnicity, or sector of employment. The effect of gender on returns to human capital is small and statistically insignificant in all survey years; likewise for the effect of ethnicity, except in one year.[15] The effect of sector of employment on returns to human capital is statistically significant only in 2003. Table 3.3 shows that, in 2003, returns to education are a statistically significant 1.8% lower for individuals working in the private sector. This situation in Bulgaria contrasts with that of Eastern Europe overall. Flabbi et al. (2007) find that returns to human capital are much higher in private sectors across the transition economies, especially in the more developed Central European economies, and that Russia is the only other transition economy in which returns to human capital are not higher in the private sector.

Although in 2003 the private sector has significantly lower returns to education than the public sector, this does not change the basic conclusion for the overall trend of growing returns to education over time; in the last survey year returns to education in each sector are significantly higher than in the 1990s early transition years. There is one more notable point about the effect of sector of employment on wage determination in 2003: the gender gap in the private sector is significantly

[15]In 2003, returns to education are a statistically significant 1.9% lower for ethnic minorities. This result is consistent with other studies (e.g. O'Higgins 2009) showing that Roma minorities in Eastern Europe have lower returns to education. Despite this, the surveys I use were not designed specifically to be representative of the minority population in Bulgaria and thus are unlikely to reflect their situation fully. I do not focus on this result because I find no statistically significant difference in the other survey years nor do I find that ethnicity affects any of the other results in any survey year.

smaller than in the public sector in this year. Thus throughout the transition period women in the public sector are not gaining relative to their male counterparts. The decrease in the economy-wide gender gap in 2003 relative to 1995 which I note in Table 3.2 is due only to the decrease of the gender gap in the private sector. This finding that being in the private sector reduces the gender gap is surprising and opposite to the experience of not only the more developed transition economies such as Czech Republic and Slovakia (Jurajda 2003), but also those at a similarly lower level of development such as Romania (Skoufias 2003).

Not only are the effects of gender, ethnicity, and sector of employment on returns to human capital for the most part statistically insignificant, I find evidence that these OLS estimates are not biased due to individuals being selected into unemployment during the transition as a result of their possessing human capital less relevant in the new market economy. Table 3.4 shows that although the selection-corrected estimates of returns to human capital are lower than the OLS estimates as I predict in Section 3.3.1, the two sets of coefficients are virtually the same. I cannot reject the hypothesis that the correlation between the error term of the earnings equation (3.1) and that of the selection equation is zero.[16]

Nevertheless, I do find that individuals who moved from one industry during communism to another in the post-communist period receive much smaller returns to experience than those who stayed in the same industry. Table 3.5 shows that only 5.5% of the individuals who were already in the labor force by 1989 remain in the same industry in 1995 – 6 years into the transition period. These few individuals have very high returns to human capital: 7.8% return to schooling and 8.6% return to experience relative to 4.6% and 1.8% respectively for those individuals who changed industries by 1995. However, only the difference in returns to work experience is statistically significant at the 5% level. Thus it appears that low returns to experience in Bulgaria during the transition period were due to the very drastic transformation of the economy. It forced the predominant majority of the labor force to change industries and devalued their work experience

[16]Nevertheless, a strong predictor in the selection equation is the dummy variable indicating whether an individual worked either on a cooperative farm or at an industrial job in 1989. Such individuals are 24% more likely to be unemployed in 1995.

under communism.

Since returns to education are the only wage determinant markedly changing during Bulgaria's transition, I next examine whether having completed schooling during communism rather than after affects returns to human capital or the other wage determinants. I find that having completed schooling during communism affects only returns to education. The exact effect, however, depends on the level of schooling. Table 3.6 shows that, in each survey year after 1995, returns to secondary education are much lower for individuals having completed schooling under communism rather than after. This effect is even larger for returns to tertiary education. On the other hand, returns to primary education – small and in general statistically insignificant to begin with – are least affected by having completed schooling under communism. Thus the higher the level of schooling completed under communism, the less it is valued in the new market economy compared to the same level of schooling completed after.

Next, I examine differences in returns to education between cohorts. Table 3.7 shows, for each survey year, returns to education for four cohorts: 20- to 29-year-olds, 30- to 39-year-olds, 40- to 49-year-olds, and 50- to 59-year-olds. Returns to education are higher for the younger than the older cohorts: the oldest cohort – individuals in their 50s – receives the lowest returns to education in every survey year, except for the last year when it receives the second-lowest returns. However, whereas in the two early transition years it is the youngest individuals – those in their 20s and 30s who receive the highest returns, by the later transition period it is the middle-aged individuals in their 40s who receive the highest returns among all cohorts. Moreover, the differences in returns to education between cohorts become smaller. Whereas in 1995 and 1997 some of the younger cohorts have returns to education twice that of the older cohorts, by 2003 the differences are slight in comparison.

Table 3.8 shows how returns to education change for each cohort as it ages over the survey years.[17] Individuals in their 20s or 30s in 1995 receive increasing returns to education as they age over survey years. The largest increase, however, is received by individuals in their 40s in 1995.

[17]The oldest cohort – individuals who are in their 50s in 1995 – is omitted because aging this cohort through the later survey years leaves very few observations as individuals retire.

Although these individuals in 1995 receive lower returns to education than the younger cohorts, by 2003 their returns to education and those of the younger cohorts are virtually the same. This result explains most of the overall increase in returns to education that I note at the beginning of this section: returns to education grow over the survey years for two reasons: 1) as individuals age through the transition they receive greater schooling returns and 2) while in the early transition years only the young receive substantial schooling returns, by the later transition years the middle-aged catch up too.

3.4.2 Wage inequality

Given these results on returns to human capital, I examine the impact of these returns on wage inequality during Bulgaria's transition. Table 3.9 shows three measures of wage inequality – the variance of log wages, the Gini coefficient, and the Theil-T index – for each survey year. The three measures all show that wage inequality in 2003 is much greater than in 1995. Wage inequality is greatest in 1997 during the Bulgarian financial and currency crisis. This result is consistent with the results on returns to human capital in the previous section: the low R-squared for 1997 in Table 3.2 possibly indicates that wages in 1997 are poorly explained by human capital and other individual characteristics.[18] Table 3.9 shows that variance of log wages grows from 0.274 in 1995 to 0.314 in 2003, the Gini coefficient from 0.291 to 0.315. By comparison, in East Germany variance of log wages was 0.102 in 1991 (Krueger 1992) and grew only to 0.135 by 2000 (Gang and Yun 2002). Similarly, in the Czech Republic the Gini coefficient is fairly constant around 0.230 from 1997 to 2003; in Hungary, around 0.250 (Kaasa 2003). Only the other poorer transition economies, for example Romania, show the same pattern as Bulgaria: large wage inequality at the outset that grows further over time.

[18]The very high wage inequality recorded in 1997 relative to the period before or after could also be due to measurement error and the difficulty of calculating real wages during hyperinflation. The 1997 survey was conducted over a very big time interval: March through September. Nevertheless, the high variance of wages could also be due to real divergence during the crisis year: while some wages kept more at pace with inflation, others fell behind.

Next, I decompose overall wage inequality in each survey year into inequality within groups and between groups, with the groups defined in turn by gender, ethnicity, schooling, age and sector of employment. Table 3.10 shows that wage inequality between men and women is very small and continuously shrinks over the survey years. In each survey year, however, wage inequality among men is higher than among women – higher in fact than overall wage inequality. Growing wage inequality among men explains most of the growth in overall wage inequality; wage inequality among women grows more modestly.

Table 3.10 also shows the alternative decomposition based on ethnicity. Similarly to the decomposition by gender, most of the growth in overall wage inequality can be attributed to growing wage inequality within ethnic groups rather than between ethnic groups. Moreover, whereas wage inequality among ethnic minorities at the end of the period is only slightly higher than at the beginning, wage inequality among the Bulgarian majority grows considerably.

The next decomposition in Table 3.10 is based on educational attainment. The table shows that, as with the other decompositions, the growth in overall wage inequality is mostly due to growing wage inequality within groups rather than between. Yet this pattern might not persist into the future; in the last survey year wage inequality between educational groups grows sharply from its modest level in the early transition years. As Table 3.10 shows, wage inequality is stablest and in general least among individuals with secondary degrees. On the other hand, wage inequality is high and grows among those with tertiary degrees. In recent years, high-tech firms from Western Europe have opened more and more offices in Bulgaria. These firms hire among college-educated Bulgarians those with foreign language and computer skills, paying wages closer to the standard of living in Western Europe and thus increasing wage inequality among this educational group.

The next decomposition is based on age. Table 3.10 shows that, between 1995 and 2003, wage inequality grows least for the youngest cohort – individuals in their 20s. The effect of Bulgaria's transition on wage inequality is greatest for individuals in middle age. For example, in 1995 wage inequality is least among individuals in their 40s than among any other age group; in 2003, it is 28% greater in this age group. In comparison, overall wage inequality between 1995 and 2003

grows half as fast – by 15%. These results suggest that young Bulgarians, who have hardly any work experience under communism, adapt with equal success to the new market economy, while middle-aged Bulgarians, who have half of their lifetime careers under communism, adapt with much more varying success.

The final group decomposition of overall wage inequality is within and between sectors of employment. Table 3.10 shows that, between 1995 and 2003, wage inequality in the private sector shrinks and wage inequality in the public sector grows. Whereas in the first survey year wage inequality in the private sector is higher and that in the public sector is lower than overall wage inequality, by the last survey year the roles are reversed.

These group decompositions of overall wage inequality are only partly consistent with such decompositions in the other transition economies. Wage inequality between individuals with different levels of schooling is similar across Eastern Europe: as in Bulgaria, wage inequality is greatest among individuals with tertiary degrees and least among those with secondary degrees (Mitra and Yemtsov 2006). On the other hand, results vary across Eastern Europe when wage inequality is decomposed by gender, ethnicity, sector of employment, or age. These results across Eastern Europe are consistent with the explanation that, whereas wage inequality between individuals with different levels of schooling is largely due to common external factors like foreign direct investment, differences in wage inequality in the other group decompositions are largely due to different domestic policies.

From decomposing wage inequality by sample groups within a survey year, I now move on to analyzing the growth in wage inequality over the entire period. I first perform a semi-parametric decomposition (DiNardo et al. 1996) to determine how much of the growth in wage inequality in Bulgaria between 1995 and 2003 is due to changes in sample characteristics. Figure 3.1 shows the densities of log wages in 1995 and in 2003: again, wage inequality in 2003 is much greater than in 1995. Figure 3.2, on the other hand, shows the density of log wages in 1995 and what the density of log wages would have been in 2003 had sample characteristics been the same as in 1995. Wage inequality in this counterfactual 2003 is still much greater than in 1995. Thus most of the growth

in wage inequality between 1995 and 2003 is not due to changes in sample characteristics.

Given this result, I next determine how much of the growth in wage inequality is due to the estimated growth in returns to human capital. Table 3.11 shows the results from a parametric decomposition of the growth in wage inequality between 1995 and 2003 into changes in sample characteristics, changes in estimated coefficients, and changes in residuals variance. Changes in sample characteristics play no role in the growth in wage inequality (the contribution of these changes is in fact a negative 1.2%). Moreover, although the residuals variance increases from 1995 to 2003, this change explains less than half of the growth in wage inequality – 44.9%. Thus most of the growth in wage inequality is due to changes in the coefficients of earnings equation (3.1). Table 3.11 shows that, among these changes, growing returns to human capital, specifically education, explain most of the growth in wage inequality. While changes in the educational composition of the sample tend to reduce wage inequality, growing returns to education explain almost 78% of the growth in wage inequality.[19]

This result is comparable to the result found by Gang and Yun (2002) for East Germany between 1990 and 2000.[20] In both Bulgaria and East Germany, growing returns to human capital – in particular, growing returns to education – explain most of the growth in wage inequality during transition. Gang and Yun (2002) attribute this result for East Germany to skill-biased technological change. For Bulgaria, my results suggest one additional explanation. As I note in Section 3.4.1, returns to secondary and tertiary education are much lower for individuals having completed schooling under communism rather than after. Returns to education grew during Bulgaria's transition partly because over time successive cohorts were exposed to education more useful in the new market economy. At the same time, educational attainment did not increase appreciably. Thus the

[19]Changes in the sample composition with respect to gender, ethnicity, and sector of employment have smaller, mixed effects, as do changes in their coefficients. Changes in the gender composition and in the coefficient on gender tend to reduce inequality. Changes in the ethnic composition and in the coefficient on ethnicity tend to increase inequality slightly. Changes in the sample composition with respect to sector of employment and in its coefficient also tend to increase inequality slightly.

[20]Since their analysis was on male workers of German ethnicity, I compare only returns to education.

growing returns to education resulted in growing wage inequality.

3.5 Conclusions

Bulgaria, one of the poorest Eastern European countries, has experienced a difficult transition to a market economy. Returns to education in Bulgaria have been growing but modestly. Bulgaria still lags far behind not only the countries of Western Europe, but also the more developed transition economies, such as Hungary and Poland. Returns to human capital in these former communist countries were higher even during communism, and afterwards grew at least as fast as the returns to human capital in Bulgaria, often faster.

Moreover, unlike in other former communist countries such as Slovenia, returns to work experience in Bulgaria have stagnated. However, there is some indication that this might be only a temporary phenomenon. The low returns to experience appear to be at least partly due to the nonexistent value that the market economy places on work experience before 1990. Thus as more Bulgarians gain work experience under the market economy, the returns to work experience might start to increase.

The wage gap between men and women, and the wage gap between ethnic Bulgarians and ethnic minorities have remained substantial. Finding no explanation for these gaps in different returns to human capital suggests the need for more research into other causes, including discrimination (Jolliffe 2002).

More recent data are also needed to give a true picture of how the private sector in Bulgaria, still very small relative to that in more advanced transition economies, is growing. This growth, driven by greater integration with the economies of Western Europe and by the revival of privately-owned farms in Bulgaria, might raise returns to human capital in the private sector.

Looking at the evolution of overall wage inequality over the transition period 1995-2003, I identify growing returns to education to be the single most important factor pushing the growth in wage inequality. There is indication that in the early years of transition returns grew mostly as the

young benefitted from a higher education which was more relevant to the new market economy. However, over the course of the transition returns began to grow for the middle-aged as well. Thus unless these growing returns motivate an increase in educational attainment that is much more substantial than has occurred so far, wage inequality is likely to continue growing.

What this chapter leaves unexamined is the experience of the self-employed during transition – a transition experience that potentially would be very different from that of the hired employees examined here. In my data set the self-employed are a rapidly growing yet an extremely small portion of the sample each year up to 2003 included, so their analysis would require more recent surveys.

Lastly, this chapter finds that the transition after the end of communism proved a defining moment for returns to human capital in Bulgaria. Other studies show that countries across Eastern Europe likewise underwent a crucial period of transition. A natural extension of this research on human capital and inequality during transition would be to compare the period of transition after communism with the more recent period of transition after EU membership, a defining moment itself.

Figure 3.1: Growth in wage inequality between 1995 and 2003



Figure 3.2: Effect of characteristics on the growth of wage inequality

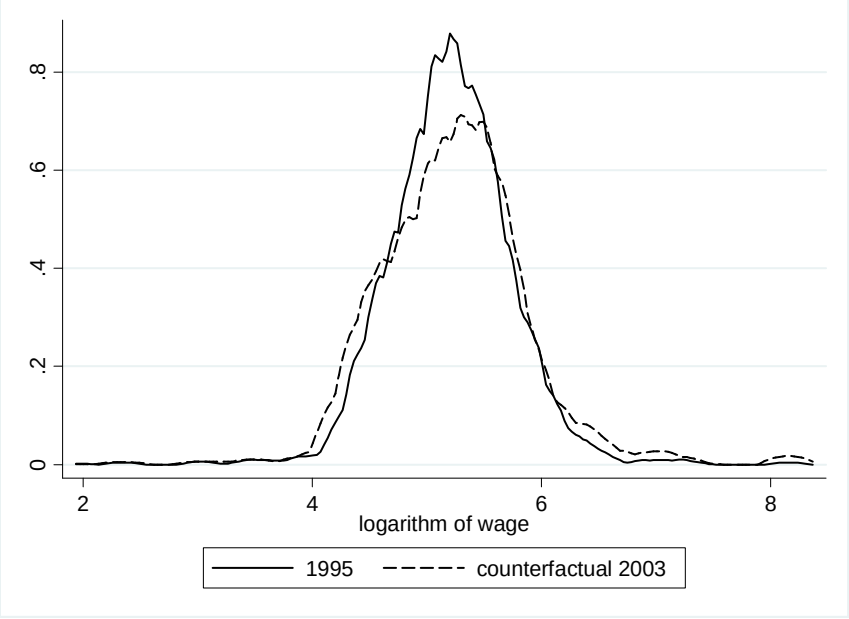


Table 3.1: Descriptive statistics

	1995	1997	2001	2003
Total wage bill (official data)	12816.41	7903.18	9668.42	10338.70
Gross monthly wage	211.40 (163.45)	111.02 (137.79)	206.62 (129.22)	203.81 (160.08)
Private sector dummy	0.172	0.231	0.519	0.612
Years of schooling	11.51 (3.05)	11.86 (2.92)	11.97 (2.95)	12.18 (3.18)
Pre-1990 education dummy	0.902	0.850	0.748	0.687
Years of work experience	21.15 (10.63)	21.62 (10.71)	20.85 (11.07)	20.50 (11.37)
Age	39.65 (9.99)	40.48 (10.13)	39.83 (10.65)	39.68 (10.85)
Female dummy	0.491	0.500	0.488	0.477
Ethnic minority dummy	0.090	0.089	0.095	0.102
Observations	1985	1601	1658	2528

Notes: The table reports the means of the variables above with standard deviations given in parentheses. The wage data are given in constant 2000 Bulgarian Leva (1BGN = 0.72 USD PPP in 2009). The total wage bill is from the national income accounts provided by the National Statistical Institute of Bulgaria.

Table 3.2: Returns to human capital

	Dependent variable: log monthly wages			
	1995	1997	2001	2003
Years of schooling	0.046 (0.005)	0.048 (0.009)	0.058 (0.006)	0.071 (0.004)
Experience	0.028 (0.004)	0.025 (0.007)	0.030 (0.004)	0.026 (0.003)
Experience squared (/100)	-0.055 (0.009)	-0.052 (0.017)	-0.061 (0.010)	-0.051 (0.008)
Female dummy	-0.351 (0.019)	-0.381 (0.035)	-0.330 (0.026)	-0.280 (0.018)
Ethnic minority dummy	-0.108 (0.040)	-0.139 (0.103)	-0.219 (0.064)	-0.155 (0.038)
Private sector dummy	-0.026 (0.037)	-0.140 (0.051)	-0.064 (0.030)	-0.057 (0.022)
R^2	0.253	0.137	0.216	0.291
Observations	1985	1601	1658	2528

Notes: The estimates are from OLS models with standard errors clustered by household in parentheses. Not reported but included in all regressions are a constant and 9 regional dummies.

Table 3.3: Public versus private sector wage determination

	Dependent variable: log monthly wages in 2003	
	Private sector	Public sector
Years of schooling	0.063	0.081
	(0.006)	(0.005)
Experience	0.026	0.030
	(0.004)	(0.006)
Experience squared (/100)	-0.050	-0.056
	(0.009)	(0.013)
Female dummy	-0.237	-0.334
	(0.024)	(0.029)
Ethnic minority dummy	-0.131	-0.197
	(0.050)	(0.055)
R^2	0.250	0.368
Observations	1548	980

Notes: The estimates are from OLS models with standard errors clustered by household in parentheses. Not reported but included in all regressions are a constant and 9 regional dummies.

Table 3.4: Selection into hired employment after transition

	Dependent variable: log monthly wages in 1995	
	OLS	Two-Step Heckman
Years of schooling	0.046	0.044
	(0.005)	(0.005)
Experience	0.028	0.026
	(0.004)	(0.005)
Experience squared (/100)	-0.055	-0.052
	(0.009)	(0.010)
Female dummy	-0.351	-0.352
	(0.019)	(0.021)
Ethnic minority dummy	-0.108	-0.097
	(0.040)	(0.042)
Private sector dummy	-0.026	-0.044
	(0.037)	(0.042)
R^2	0.253	
Uncensored observations	1985	1985
Censored observations		619
Inverse Mills ratio		-0.044
(lambda)		(0.074)
Rho		-0.097

Notes: The estimates are from an OLS regression and a two-step Heckman regression. The selection equation in the Heckman correction model includes the same covariates as in the OLS model plus an indicator for employment at a cooperative farm or at an industry job in 1989. Household clustered standard errors are in parentheses.

Table 3.5: Returns to experience after industry change in transition

	Dependent variable: log monthly wages in 1995	
	Same industry as in 1989	Different industry from 1989
Years of schooling	0.078	0.046
	(0.023)	(0.005)
Experience	0.086	0.018
	(0.033)	(0.005)
Experience squared (/100)	-0.173	-0.037
	(0.076)	(0.011)
Female	-0.297	-0.346
	(0.094)	(0.020)
Ethnic minority	0.150	-0.116
	(0.417)	(0.038)
Private sector	0.048	-0.037
	(0.097)	(0.045)
R^2	0.391	0.255
Observations	99	1691

Notes: The estimates are from OLS models with the same covariates as in equation (3.1) with the sample restricted to individuals who had completed their education by 1989. Not reported but included in all regressions are a constant and 9 regional dummies. Household clustered standard errors are in parentheses.

Table 3.6: Difference in returns to education acquired pre-1990 and post-1990

	Dependent variable: log monthly wages					
	1997		2001		2003	
Education:	Pre-1990	Post-1990	Pre-1990	Post-1990	Pre-1990	Post-1990
Primary	-0.008	0.380	0.309	1.240	0.052	0.125
	(0.149)	(0.297)	(0.121)	(0.212)	(0.066)	(0.107)
Secondary	0.248	0.832	0.535	1.344	0.295	0.443
	(0.151)	(0.230)	(0.126)	(0.206)	(0.068)	(0.104)
Tertiary	0.367	1.118	0.779	1.655	0.669	0.775
	(0.158)	(0.241)	(0.129)	(0.222)	(0.071)	(0.112)
R^2	0.126	0.133	0.211	0.246	0.267	0.307
Observations	1361	240	1240	418	1736	792

Notes: The estimates are from OLS models with the same covariates as in equation (3.1) with the exception of the linear term of years of schooling which has been substituted with dummies for completed primary, secondary and tertiary degreee. Household clustered standard errors are in parentheses.

Table 3.7: Differences in returns to education between cohorts

	Return to an additional year of schooling			
	1995	1997	2001	2003
Whole sample	0.046	0.048	0.058	0.071
20- to 29-year-olds	0.044 (0.014)	0.100 (0.028)	0.068 (0.017)	0.068 (0.011)
30- to 39-year-olds	0.059 (0.012)	0.047 (0.022)	0.067 (0.013)	0.080 (0.010)
40- to 49-year-olds	0.031 (0.009)	0.039 (0.019)	0.073 (0.013)	0.085 (0.009)
50- to 59-year-olds	0.027 (0.015)	0.036 (0.021)	0.039 (0.015)	0.071 (0.011)

Notes: The estimates are from OLS models with the same covariates as in equation (3.1). Household clustered standard errors are in parentheses.

Table 3.8: Differences in returns to education for three cohorts over time

	Return to an additional year of schooling			
	1995	1997	2001	2003
Cohort age:	20-29	22-31	26-35	28-37
	0.044	0.089	0.070	0.070
	(0.014)	(0.024)	(0.014)	(0.009)
Cohort age:	30-39	32-41	36-45	38-47
	0.059	0.054	0.040	0.078
	(0.012)	(0.020)	(0.017)	(0.009)
Cohort age:	40-49	42-51	46-55	48-57
	0.031	0.046	0.073	0.076
	(0.009)	(0.021)	(0.013)	(0.010)

Notes: The estimates are from OLS models with the same covariates as in equation (3.1). Household clustered standard errors are in parentheses.

Table 3.9: Overall wage inequality

	Inequality measures			
	1995	1997	2001	2003
Variance of log	0.274	0.731	0.356	0.314
Gini	0.291	0.443	0.296	0.315
Theil-T	0.169	0.383	0.154	0.186
Observations	1985	1601	1658	2528

Notes: For each sample year the income variable used to calculate the inequality measures is gross monthly wages.

Table 3.10: Decompositions of wage inequality

	1995	1997	2001	2003
Total inequality	0.274	0.731	0.356	0.314
Among primary education	0.245 (18.1%)	0.919 (21.4%)	0.326 (14.0%)	0.247 (10.6%)
Among secondary	0.242 (47.7%)	0.668 (50.6%)	0.332 (53.1%)	0.244 (43.7%)
Among tertary	0.279 (23.1%)	0.672 (23.3%)	0.292 (20.9%)	0.334 (29.0%)
Between degrees	0.020 (7.3%)	0.019 (2.6%)	0.032 (9.0%)	0.046 (14.6%)
Among women	0.236 (42.3%)	0.638 (43.6%)	0.294 (40.3%)	0.247 (37.6%)
Among men	0.258 (47.9%)	0.772 (52.8%)	0.380 (54.7%)	0.353 (58.8%)
Between sexes	0.027 (9.8%)	0.026 (3.6%)	0.018 (5.0%)	0.011 (3.6%)
Among ethnic Bulgarians	0.270 (89.8%)	0.716 (89.2%)	0.324 (82.4%)	0.304 (87.0%)
Among ethnic minorities	0.241 (7.9%)	0.797 (9.7%)	0.532 (14.2%)	0.248 (8.1%)
Between ethnicities	0.006 (2.3%)	0.008 (1.1%)	0.012 (3.4%)	0.016 (4.9%)
Within private sector	0.471 (29.6%)	0.695 (22.0%)	0.423 (37.1%)	0.304 (39.1%)
Within public	0.233 (70.3%)	0.736 (77.4%)	0.274 (61.7%)	0.316 (59.4%)
Between sectors	0.000 (0.1%)	0.005 (0.6%)	0.004 (1.2%)	0.005 (1.5%)

Table 3.10: Decompositions of wage inequality (cont.)

	1995	1997	2001	2003
Total inequality	0.274	0.731	0.356	0.314
Among 20- to 29-year-olds	0.283 (18.1%)	0.590 (13.8%)	0.363 (20.5%)	0.287 (18.7%)
Among 30- to 39-year-olds	0.287 (30.1%)	0.581 (19.7%)	0.318 (23.3%)	0.334 (30.3%)
Among 40- to 49-year-olds	0.226 (29.2%)	0.953 (46.5%)	0.441 (37.9%)	0.289 (24.7%)
Among 50- to 59-year-olds	0.294 (17.9%)	0.640 (18.2%)	0.243 (14.0%)	0.312 (21.1%)
Between age groups	0.008 (2.8%)	0.005 (0.6%)	0.009 (2.6%)	0.008 (2.6%)

Notes: Reported are the variance of log gross monthly wages for each group and between groups with the percentage contribution to total inequality in parentheses. Not reported but included in the within-group component in the decomposition by education are those with less than primary education and in the age decomposition – those aged 16-19 and those aged 60-65.

Table 3.11: Decomposition of the growth in wage inequality between 1995 and 2003

	Characteristics effect		Coefficients effect	
Total effect	-0.0005	(-1.17%)	0.0225	(56.30%)
Human capital	-0.0006	(-1.58%)	0.0303	(75.75%)
Years of schooling	-0.0009	(-2.15%)	0.0310	(77.69%)
Experience	0.0003	(0.65%)	-0.0044	(-10.98%)
Experience squared	-0.0000	(-0.08%)	0.0036	(9.04%)
Female	-0.0016	(-4.03%)	-0.0122	(-30.45%)
Ethnic minority	0.0002	(0.61%)	0.0032	(7.97%)
Private sector	0.0003	(0.74%)	0.0015	(7.97%)
Residuals effect		0.0179	(44.87%)	

Notes: The inequality measure is variance of log gross monthly wages. The coefficients and residuals used in the decomposition are from the same OLS model as in equation (3.1). The percentage contributions are in parentheses.

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