

# CAN MATERNITY BENEFITS HAVE LONG-TERM EFFECTS ON CHILDBEARING?

## EVIDENCE FROM SOVIET RUSSIA

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This paper quantifies the effects of Russia's 1981 expansion in maternity benefits on completed childbearing. The program provided one year of partially paid parental leave and a small cash benefit upon a child's birth. I exploit the program's two-stage implementation in an event-study framework and find evidence that women had more children as a result of the program. Fertility rates rose immediately by approximately 8.2 percent over twelve months. The increase in fertility rates not only persisted for the ten-year duration of the program, but it reflected large increases in higher order births to older women who already had children before the program started. The program induced nearly 5 million births over its duration, where an extra birth cost the government about 1.4 times a year's average national earnings.

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Eighty-four percent of developed countries offer subsidies or parental leave benefits at an average cost of 2.6 percent of GDP (United Nations 2013). Some of these programs are tremendously expensive.<sup>2</sup> Countries implement these programs in part to increase childbearing, because they are worried that below replacement fertility levels accompanied by an increase in life expectancy may negatively affect their economies in several ways. First, this demographic shift threatens the ability of many countries to finance old-age benefits.<sup>3</sup> Second, a shrinking working-age population compared to a rising elderly population may result in lower economic growth because of a decline in workers per capita (Bloom et al. 2009). Although maternity benefits are costly, they may be ineffective if they result in only a short-run increase in childbearing due to a shift in timing of childbearing, instead of a long-run increase due to women having more children.

Whether these programs are effective in raising childbearing is an open question. The provision of more generous parental benefits is associated with a country's demand for children, which makes estimating the effects of programs themselves difficult. To address this problem, the literature uses a variety of natural experiments in different countries to provide evidence that parental leave (Lalive and Zweimüller 2009, [Austria]) and cash transfer programs (Cohen et al. 2013, [Israel]; Gonzalez 2012, [Spain]; Milligan 2005, [Canada]) have short-run, positive effects on childbearing in developed countries. But, the limited time-horizon of available data and empirical methods limit inferences about long-run effects. Consequently, much of the estimated effects may reflect changes in the timing rather than a permanent increase in childbearing.

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<sup>2</sup> Germany, for instance, spends nearly 100 billion dollars per year on family benefits.

<sup>3</sup> The total fertility rate is below the replacement level of 2.1 in 113 countries (CIA Factbook 2013). Population aging is a concern for 92 percent of developed countries, where 22 percent of the population is over 60 (United Nations 2013).

This paper leverages the two-stage introduction of Russia's 1981 expansion of maternity benefits to evaluate both its short-run and long-run effects on childbearing. Similar to the goals of programs in developed countries today, the program was intended to increase completed childbearing by providing a sizable expansion in partially paid parental leave until a child turned one, unpaid parental leave until a child turned a year and a half, and cash transfers at the birth of the first, second or third child. Eighty-five percent of women were eligible for benefits, because they met the provision stipulating that they be in the labor force.

My research design uses the Soviet government mandate that the benefits start in 32 oblasts in 1981 (similar to states; I call these oblasts "early beneficiaries"), and then in 50 oblasts ("late beneficiaries") one year later. The historical vantage point allows me to estimate long-run effects of the maternity benefits expansion. Another contribution of the project is the reconstruction of Russian population and characteristics of regions data from recovered vital statistics, censuses, yearbooks, and surveys, which I cull and translate from published sources.

My results show that the 1981 Russian maternity benefit program is associated with an immediate and sustained increase in childbearing. Fertility rates rose *immediately* after the program started by approximately 8.2 percent in the first twelve months. The elasticity of fertility rates with respect to a change in cost of a child is -3.7, which is in the range of short-run effects found in other studies.<sup>4</sup> Three empirical findings underscore that this increase reflects higher completed childbearing. First, period fertility rates remained on average 14.6 percent higher for the ten-year duration of the program. Second, children born after the expansion were more likely

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<sup>4</sup> I calculate the price elasticity of fertility rates with respect to changes in the cost of having a child for the first 18 years after birth. Most of the prior literature does not convert their estimates into price elasticities. I calculate the price elasticity in Lalive and Zweimüller (2009) as -4.4, the elasticity in Gonzalez (2012) as -3.8, the elasticity in Milligan (2005) as -4.1, while the reported elasticity in Cohen et al. (2003) is -0.54.

to have been higher order births. Third, children born after the expansion were born to mothers who were older and had a longer interval since their previous birth.

This study is the first to find a positive effect of maternity benefits on long-run fertility rates.<sup>5</sup> This is at odds with some papers that have argued that maternity benefits have small or no effects on raising short-run fertility rates and long-run childbearing (Demeny 1986; Gauthier et al. 2007).<sup>6</sup> Theoretically, maternity benefits could decrease childbearing over the longer term. To demonstrate this, I develop two theoretical extensions of the Becker and Lewis (1973) model that incorporate two types of maternity benefits—paid leave, which lowers the opportunity cost of childbearing, and a cash transfer, which increases income. The first extension allows parents to choose both the number of children and their level of investments in each child (“quality”), which has the well-known effect of making the income effect theoretically ambiguous. The second extension achieves the same result through non-standard channels: by allowing women to make endogenous choices about their time off from work.

Theoretically, the effect of maternity benefits on childbearing depends on the shadow price of a child (Becker 1965) and on income (Becker and Tomes 1976; Galor and Weil 1999). My empirical results suggest the opportunity cost of having a child as the mechanism for the effect of maternity benefits on childbearing, because less educated and more rural areas experienced a greater increase in fertility rates for the duration of the program. Because the benefits were the same amounts for everyone, women with lower opportunity costs experienced

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<sup>5</sup> Lalive and Zweimüller (2009) find suggestive evidence of long-run effects when women who already had one child were more likely to progress to a second child if they randomly received extra paid parental leave for the first child. The women who were eligible for extra leave were also eligible for an automatic renewal of benefits if they gave birth to the second child up to 27.5 months after the first birth, while women who were not eligible could only receive automatic renewal if they gave birth to the second child 15.5 months after. Thus, this result may largely be due to other types of benefits, and is not applicable to typical interventions affecting benefits for future births.

<sup>6</sup> Papers on the United States focus on low income women to study effects of welfare policies on fertility rates, and find inconclusive evidence due to a large variation in results (Hoyne 1997; Moffitt 1998).

larger relative decreases in the opportunity costs of having a child. Maternity benefits may have had long-run effects on childbearing in Russia, because Russian women had both low opportunity costs of taking time off and low costs of raising children. My best evidence on this comes from the fact that Russian women had lower earnings compared to men and a flatter age-earnings profile than women in other countries (Brainerd 2000, Gregory and Kohlhasse 1988). They also had access to wide-spread and affordable preschool care for children of *all* ages.

In summary, this paper shows that maternity benefits can have an effect on both short-run *and* long-run childbearing behavior. In the Russian case, the program induced nearly 5 million births over its duration, where an extra birth cost the government about 1.4 times a year's average national earnings. This paper also provides a cautionary tale for countries interested in designing their maternity benefit programs: behavioral responses to these incentives may vary tremendously across contexts and will reflect the interaction of maternity benefits with other social and public programs.

## **I. Russian Family Benefits**

Before 1981, the major beneficiaries of family subsidies were families with many children or low income families. From 1947, women received a one-time payment beginning with their third child and monthly supplements until a child's fifth birthday beginning with their fourth child (Presidium Verhovnogo Soveta 1947). In addition, after 1974, families with monthly per capita income below a threshold received monthly supplements for each child under the age of eight (Presidium Verhovnogo Soveta 1974). However, these benefits did not provide incentives to an "average" family to have a second or a third child.

The government also provided limited benefits to working mothers, but did not provide financial support for women who wanted to stay home with a child for a longer period of time. The most generous benefit was a fully-paid maternity leave of 56 days before and 56 days after a birth. In addition, women could take an unpaid job-protected parental leave until a child turned one (Goskomtrud 1970). Job protected leave was an important feature in Soviet Russia, where about half of the labor force consisted of women in 1980.

In the late 1970s the Soviet government formulated a pronatalist policy with the goal of encouraging all women to have second and third births. The government was interested in securing the replacement level of children “from each physically and morally healthy family, instead of a maximum of children from a minimum of families” (Desfosses 1981). One of the motivations was the desire to achieve more population growth in areas with labor shortages (DiMaio 1981).<sup>7</sup> In January 1978, Litvinova (1978), who was the Senior Research Associate with the Institute of State and Law of the USSR Academy of Sciences, wrote that the government needs to be concerned about the “quality” of the population. She added that “the state cannot be indifferent to what kind of population increase occurs, whether it is highly mobile or, owing to a variety of circumstances (including large families and language barriers), bound to one specific region”.

#### *A. Description of the Maternity Benefits Expansion*

The outcome of these discussions was the 1981 maternity benefit program that aimed to increase childbearing by providing “good conditions for population growth”, to “ease the status

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<sup>7</sup> Women in the Central Asia republics (e.g. Kazakhstan, Uzbekistan) were the majority of beneficiaries of income-tested and higher parity births benefits in the Soviet Union, because they had more children compared to women in Russia. Mobility from the Central Asia republics was low, thus it was difficult to move excess population to labor-shortage regions. One way to even out the spread of population was to provide incentives for childbearing in areas with lower fertility rates and labor shortages.

of working mothers”, and to “decrease the differences in standard of living depending on having children” (TSK KPSS 1981). The maternity benefit program provided three new benefits. First, women received partially paid parental leave until a child turned one, which represented a flat monthly payment equaling roughly 27 percent of the average national female monthly salary.<sup>8</sup> Second, women could keep their job while staying home until their child turned 18 months old. Third, women received a one-time per birth cash transfer which was about 38 percent of the average national monthly salary for first births, and 76 percent of the average national monthly salary for second and third births.<sup>9</sup>

Unlike previous programs for poor or large families, most families were eligible. Women who worked for at least a year as well as students regardless of work experience were eligible for the benefits.<sup>10</sup> Non-working women were also eligible for a small flat payment at the birth of the first, second, and third children which equaled to about 20 percent of the average national female monthly wage. Given that 85 percent of women were employed and 59 percent of college students were women, this program covered the vast majority of women. Notably, the program left previous means-tested benefits and benefits for large families unchanged.<sup>11</sup>

Even though the benefits lasted about ten years, women probably expected them to stay in place permanently. The government had never cancelled previous benefits and typically expanded them. In fact, in 1989 the government further expanded partially paid parental leave

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<sup>8</sup> Women received 50 rubles per month until the child turned one in Siberia, Far East and the Northern regions of Russia, and 35 rubles in the rest of Russia. Wages were higher in regions that received 50 rubles per month, so benefits represented the same share of average national monthly salary. The benefit at the birth of the first child was 50 rubles, while the benefit at the birth of the second and third child was 100 rubles.

<sup>9</sup> The shares reported are for women in late beneficiary regions. The shares were lower for women in the early beneficiary regions, because wages were higher there.

<sup>10</sup> Students from a wide variety of institutions could receive the benefit – universities, secondary special, professional-technical schools, clinical, and schools that provided courses to improve qualifications.

<sup>11</sup> Before 1980, women received a one-time cash transfer of 20 rubles after the birth of the third child. Under the new program, this benefit became 100 rubles.

until the child turned 18 months, and unpaid leave until the child turned three (Sovmin 1989). However, by 1992, hyperinflation reduced the benefit to almost zero and the subsequent collapse of the Soviet Union ended the program.

## II. Expected Effects of Russian Maternity Benefits on Childbearing

Introducing paid parental leave and birth transfers directly reduced the cost of having a child for working women. But, evaluating the effects of maternity benefits on childbearing decisions is more complicated than what the standard model of consumer behavior suggests. This section begins with that standard model and then extends it in two ways to discuss the theoretical effects of maternity benefits on childbearing.

### A. *Maternity Benefits in the Neo-Classical Consumer Model*

In the neoclassical consumer model parents maximize utility,  $U(n, z)$ , choosing the quantity of children,  $n$ , and another consumption good,  $z$ . Parents face a simple lifetime budget constraint,  $[t(w_f - a) - b]n + \pi_z z = T(w_f + w_m)$ , where both the husband and wife can work for units of time,  $T$ , while the wife receives,  $w_f$ , for her work and the husband receives,  $w_m$ , for his work. I assume that only women take time off,  $t$ , to take care of each child, for which they receive a benefit,  $a$ , thus making the benefit-adjusted opportunity cost of childrearing,  $t(w_f - a)$ . This model does not limit the length of time a mother can receive the benefit, but it is less than one year in the context I study. Finally, women receive a cash transfer,  $b$ , for each child.

This model predicts that both an increase in paid leave and cash transfers will increase the quantity of children, as long as children are normal goods. This reflects the fact that the effect of maternity benefits,  $\frac{dn}{da}$  and  $\frac{dn}{db}$ , is the sum of income and substitution effects: the income effect

is positive if children are normal goods, while the substitution effect is always positive. However, the quantity of children need not increase when the cost of children decreases, even if children are normal goods, once the model allows parents to choose both quantity *and* quality (Becker and Lewis 1973; Willis 1973) or endogenously choose mothers' *time off from work* (Becker 1965). In the next sections I provide two important extensions of these models that incorporate maternity benefits.

### *B. Maternity Benefits with Endogenous Choice of Child Quantity and Quality*

Maternity benefits may lead women to have fewer children due to the interaction of quantity and quality. It may become optimal for women to have fewer children, if they invest more into each child because of the benefits (and thus produce children of a higher quality). To demonstrate this, I incorporate the choice of quality of each child,  $q$ , in the neoclassical consumer model. Parents maximize utility,  $U(n, q, z)$ , and face a lifetime budget constraint,  $\pi qn + [t(w_f - a) - b]n + \pi_z z = T(w_f + w_m)$ . Parents pay,  $\pi$ , for a quality unit of a child which they must spend on each child individually, such as college tuition. This budget constraint differs from the one in the standard quantity-quality model, because it incorporates maternity benefits into the cost of a child.

Even if children are normal goods, the observed impact of an increase in income on the number of children may be negative. Children are defined as normal goods if the true income elasticity of quantity ( $\eta_n$ ) is positive:  $\eta_n$  uses a measure of income that is calculated using shadow prices (marginal costs) whose ratios in equilibrium are equal to the marginal rates of substitution in the utility function. To fix ideas, consider an increase in income where both child quantity and child quality are normal goods, but the true income elasticity is greater for quality

than for quantity,  $\eta_q > \eta_n$ . At first, both quantity and quality will rise, where quality will rise more than quantity. As shown by Becker and Lewis (1973), the fact that  $q$  and  $n$  enter the budget constraint multiplicatively leads to an increase in  $n$  to increase the shadow price of  $q$  and vice versa. Once quality increases by more than quantity, the shadow price of child quantity will rise by more than the shadow price of quality. As a result, an increase in income may reduce childbearing if the true income elasticity for quality is large enough relative to the true income elasticity for quantity. This intuition can be formalized in the following proposition and corollary.

**PROPOSITION 1:** The effect of parental leave ( $a$ ) on childbearing is positive if the difference of the true income elasticity for quantity and the true income elasticity for quality,  $\eta_n - \eta_q$ , is sufficiently large enough.

*Proof:* I differentiate the first order conditions of the utility maximization problem and the budget constraint (see the full details in appendix A). The resulting elasticity of childbearing with respect to paid leave may be decomposed into a combination of income and substitution elasticities,

$$\frac{d\ln(n)}{d\ln(a)} = \frac{atn}{I} \left( \underbrace{\frac{d\ln(n)}{d\ln(I)} \Big|_{\pi's \text{ const}}}_{\text{Income Elasticity}} \right) + \underbrace{\frac{d\ln(n)}{d\ln(a)} \Big|_{I \text{ const}}}_{\text{Substitution Elasticity}}$$

The income elasticity equals to  $\frac{d\ln(n)}{d\ln(I)} \Big|_{\pi's \text{ const}} = f(\eta_n) - g(\eta_q)$ , where  $f_{\eta_n} > 0$  and  $f_{\eta_q} > 0$ . Thus, the income elasticity of childbearing is positive if  $f(\eta_n) - g(\eta_q) > 0$ , which is true when  $\eta_n - \eta_q$  is large enough. The substitution elasticity is always positive.  $\square$

**COROLLARY 1:** The effect of cash transfers ( $b$ ) on childbearing is positive if the difference of the true income elasticity for quantity and the true income elasticity for quality,  $\eta_n - \eta_q$ , is sufficiently large enough.

*Proof:* See appendix A.  $\square$

While the theoretical income effect is ambiguous, a growing empirical literature suggests that husband's earnings (Lindo 2010, Black et al. 2013) and housing wealth (Lovenheim and Mumford 2013) have positive effects on completed childbearing.

### *C. Maternity Benefits with Endogenous Choice of Time Off*

Another way maternity benefits may lead women to have fewer children is through the interaction of quantity and time off from work. To see this, I extend the neoclassical consumer model by incorporating the choice of time off from work,  $t$ , for the mother. Households maximize utility,  $U(n, t, z)$ , and face a lifetime budget constraint,  $[t(w_f - a) - b]n + \pi_z z = T(w_f + w_m)$ . Unlike in the quantity-quality extension, the household's problem no longer involves quality, but the mother may now choose time off from work,  $t$ , while paying wages less the parental leave benefit,  $w_f - a$ , for each time unit spent out of the labor force for each child.

A notable feature of parental leave benefits is that they subsidize not only the quantity, but also potentially the quality of children because they reduce the opportunity cost of the mother's time with her child.<sup>12</sup> This feature renders the interpretation of time off from work,  $t$ , as similar to that of quality,  $q$ , in the previous extension. If women spend more time out of the labor force when they have access to partially paid leave, it may result in a greater cost of a child in a world with paid leave compared to the world with unpaid leave. In this case, it may become optimal for women to have fewer children.

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<sup>12</sup> Assuming public child care, the care of relatives or nannies are not perfect substitutes or better than a mothers' time with the child, a mother spending time with the child at a young age may improve child quality. A growing empirical literature finds a positive link between maternity benefits and children's outcomes (Rossin 2011, Carneiro et al. 2014).

Again, consider an increase in income where both child quantity and time off from work are normal goods. Time off will rise by more than quantity, if the true income elasticity is greater for time off than for quantity,  $\eta_t > \eta_n$ . Then, the shadow price of quantity will rise by more than the price of time off from work because  $t$  and  $n$  enter the budget constraint multiplicatively. Finally, an increase in income may reduce childbearing if the true income elasticity for time off is large enough relative to the true income elasticity of quantity. This intuition can be formalized in the following proposition and corollary.

**PROPOSITION 2:** The effect of parental leave ( $a$ ) on childbearing is positive if the difference of the true income elasticity of quantity and the true income elasticity of time off,  $\eta_n - \eta_t$ , is sufficiently large enough and the partial elasticity of substitution between  $n$  and  $t$ ,  $\sigma_{nt}$ , is positive.

*Proof:* I differentiate the first order conditions of the utility maximization problem and the budget constraint (see the full details in appendix A). The resulting elasticity of childbearing with respect to paid leave may be decomposed into a combination of income and substitution elasticities:

$$\frac{d\ln(n)}{d\ln(a)} = \frac{atn}{I} \left( \underbrace{\frac{d\ln(n)}{d\ln(I)} \Big|_{\pi's \text{ const}}}_{\text{Income Elasticity}} \right) + \underbrace{\frac{d\ln(n)}{d\ln(a)} \Big|_{I \text{ const}}}_{\text{Substitution Elasticity}}$$

The income elasticity equals to  $\frac{d\ln(n)}{d\ln(I)} \Big|_{\pi's \text{ const}} = f(\eta_n) - g(\eta_t)$ , where  $f_{\eta_n} > 0$  and  $f_{\eta_t} > 0$ . Thus, the income elasticity of childbearing is positive if  $f(\eta_n) - g(\eta_t) > 0$ , which is true if  $\eta_n - \eta_t$  is large enough. The substitution elasticity equals to,  $\frac{d\ln(n)}{d\ln(a)} \Big|_{\pi's \text{ const}} = h(\sigma_{nt})$ , where  $h_{\sigma_{nt}} > 0$ . Thus the substitution elasticity of childbearing is positive if  $\sigma_{nt} > 0$ .  $\square$

**COROLLARY 2:** The effect of cash transfers ( $b$ ) on childbearing is positive if  $\eta_n - \eta_t$  is sufficiently large enough.

*Proof:* See appendix A.  $\square$

#### *D. Heterogeneous Responses to Benefits Based on Opportunity Cost and Income*

The effect of prices on childbearing depends on the shadow price of a child (Becker 1965) and on income (Becker and Tomes 1976). Similarly, women's adjustment of childbearing in response to maternity benefits depends on their opportunity costs, costs of raising a child and their incomes. The neoclassical model provides different predictions on what types of women adjust their childbearing the most compared to models that allow for the choice of quality or time off from work.

The neo-classical consumer model predicts that women with lower wages and lower incomes will increase their childbearing the most. Due to the fact that the benefit is the same across all women, the benefit represents a greater relative decrease in the opportunity cost of a child and a greater relative increase in income for women with lower opportunity costs and incomes. Thus women with lower opportunity costs experience a larger relative shift of their budget constraint, which leads them to increase their childbearing by more.

However, in models that endogenize either quality or time off women with lower wages and lower incomes may increase their childbearing the least. This is because women for whom the benefits are relatively more important may use the extra income to have children of higher quality both in terms of monetary and time investments. As shown in previous extensions, the sign of the income effect may be negative when  $\eta_q$  and  $\eta_t$  are large relative to  $\eta_n$ . Women with lower opportunity costs may have a more negative income effect, because their  $\eta_q$  and  $\eta_t$  may be larger. This is because it may be less costly to increase child quality when an individual starts at a low initial level of quality.

### III. Roll-Out of Russian Maternity Benefits and Expected Timing of Responses

On January 22, 1981, the Soviet government passed a ruling about its intention to expand maternity benefits (TSK KPSS 1981). The ruling described the components of the program, the exact benefit amounts, and eligibility requirements. The ruling stated that the program was to be implemented in waves around the country. The early beneficiaries of the program were the Far East, Siberia and the Northern regions of Russia, while the late beneficiaries of the program were the rest of the regions in Russia. The announcement stated that the early beneficiaries would start receiving benefits in 1981, but did not provide information on the exact date of when benefits would start in late beneficiaries.

The Russian government may have decided the order of the benefits based on the fact that it prioritized population growth in areas with labor shortages that were important in terms of industrial production (Weber and Goodman 1981). Perevedentsev (1974), one of the most highly respected population economists, wrote “The task for the next 10 years is to balance the distribution of the country’s population and natural resources. It is the East that has the most abundant natural resources.”<sup>13</sup> Figure 1 shows a map of the roll-out of benefits across Russia, where the early beneficiary oblasts are shaded. The early beneficiary oblasts were less populous and only 25 percent of the Russian population resided there.<sup>14</sup>

Early and late beneficiary regions were similar in many respects. Table 1 shows that the share of employed women, educated individuals, women living in rural areas and children in preschool were similar in 1980. However, the population in early beneficiary oblasts is, on

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<sup>13</sup> When discussing the Soviet population distribution policy, Chinn (1977) wrote that “ideally, Siberia and the Far East should grow relative to the rest of the country.”

<sup>14</sup> Note that based on this evidence, the order of benefits was based on some fixed characteristics of regions, and not based on factors that changed over time. Including region fixed effects in regressions takes these fixed characteristics into account.

average, younger relative to the late beneficiary oblasts, which may explain higher fertility rates in early beneficiary oblasts.

Women were aware of the details of the maternity benefit program even before the ruling, because the introduction of this program was widely publicized in major Russian newspapers. Next, I present information that I translated from original sources which were in Russian. The first mention of the program was on December 2<sup>nd</sup>, 1980 in both major Russian newspapers (*Pravda* 1980; *Izvestija* 1980). Several other articles discussed these plans in more detail before and after the official announcement in January and mentioned that they would be introduced in waves starting in 1981 (*Pravda* 1981a; *Izvestija* 1981a).<sup>15</sup> The newspapers also mentioned that newspaper editors received a lot of mail with positive reviews of the program, and one woman wrote that every house in her town was talking about the new program (*Pravda* 1981b; *Izvestija* 1981b).

Later on September 2<sup>nd</sup> in 1981, a government ruling announced the exact timing of the start of benefits across regions (Sovmin 1981), and it appeared in a major newspaper shortly afterwards (*Izvestija* 1981d). This time it stated that the early beneficiaries would receive benefits starting on November 1, 1981. The late beneficiaries would receive benefits starting on November 1, 1982.<sup>16</sup> The timing of a woman's benefit eligibility depended on her location of permanent work or study and not on the place of birth of the child or residence.<sup>17</sup>

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<sup>15</sup> The most detailed description of the law appears on March 31<sup>st</sup> in 1981 as front page news in both major Russian newspapers (*Izvestija* 1981c; *Pravda* 1981c).

<sup>16</sup> Only women who gave birth after the policy's implementation were eligible to receive the one time birth transfer, while women could receive the monthly paid parental leave for the remaining months after implementation until the child turned one. For example, a woman who gave birth to her first child in November, 1981 in an early beneficiary region received 50 rubles as the birth transfer and ten 50 ruble payments until her child turned one. But, a woman who gave birth to her first child in June, 1981 in an early beneficiary region only received eight 50 ruble payments until her child turned one (Goskomtrud 1982).

<sup>17</sup> Housing shortages and the internal passport system limited geographic mobility (Brainerd 1998).

These policy changes lead to the following hypotheses. I expect an increase in childbearing in early beneficiary regions in 1981, and a larger increase in childbearing in 1982 when benefits were in place for a full year. It is likely that women adjusted their childbearing decisions after they found out about the program, because they knew that the program would start in 1981. Fertility rates could go up in early beneficiaries due to an increase in conceptions and in foregone abortions. Fertility rates may have taken longer to adjust if it was costly for women to switch methods of contraception. Given that the main method of contraception was abortion, it is reasonable to expect a quick adjustment of fertility rates (Popov 1991).

Women living in late beneficiary regions could respond in two different ways. First, they may have increased childbearing when they became eligible for benefits similarly to women in early beneficiaries. This would manifest as an increase in fertility rates in 1982 and a larger increase in 1983 when benefits were in place for a full year. Second, they may have decided to postpone childbearing after the announcement to be able to take advantage of benefits once they started in their region. This could have resulted in a reduction of fertility rates in these regions before November 1982. However, the incentive to postpone was not as strong because they would still receive some benefits even if they gave birth before their region became eligible.<sup>18</sup> Thus, the sign of the change in childbearing in late beneficiaries is ambiguous in 1982, because of the combination of a potential increase and delay in childbearing.

#### **IV. The Effect of Maternity Benefits on Childbearing in the Short-Run**

I analyze whether maternity benefits affected childbearing in the short-run by quantifying their effect within a year of the start of benefits. To do this, I construct the general fertility rate

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<sup>18</sup> If women gave birth before the start of the program in their region, they could receive benefits for the remaining months until the child turned one.

(GFR) which consists of the annual number of births per thousand women ages 15 to 44. I use non-public data on the number of births by region and year for my time period of interest which I obtained from the Russian Federal State Statistics Service (Rosstat). For the denominator, I use the 1989 Russian census to estimate the number of women of childbearing age in each year. The details for fertility rate estimation are in appendix B.

*A. Descriptive Evidence on Childbearing Responses to Maternity Benefits*

The evolution of fertility rates in early and late beneficiary oblasts provides preliminary evidence of a positive effect of the maternity benefit program on childbearing. Figure 2 plots the GFR separately for each area. Early and late beneficiaries had similar trends in fertility rates before the start of the program. Due to the fact that regions were eligible for maternity benefits for part of the year in 1981 and for the whole year in 1982, the GFR in the early beneficiaries jumped in 1981 and rose even further in 1982, which is consistent with the expected response of fertility rates. The GFR in the late beneficiaries did not decrease in 1981, which indicates that women did not delay childbearing after the announcement of the benefits. Also, GFR in late beneficiary regions increased in 1982, which likely resulted from a combination of an increase in childbearing among women induced to have children due to the benefits and a decrease in childbearing of women who decided to postpone childbearing until their regions became eligible for benefits. Similar to behavior in early beneficiary oblasts, GFR in late beneficiary oblasts jumped the most in 1983 when benefits were in place for the whole year. Finally, fertility rates in both early and late beneficiaries stayed higher after the start of the program, suggesting the program increased completed childbearing.

### B. Generalized Differences-in-Differences Framework

I next use this two-stage implementation in a generalized differences-in-differences framework to adjust these raw comparisons for other covariates and construct confidence intervals (Jacobson et al. 1993),

$$GFR_{o,y} = \alpha + \gamma_y + \delta_o + \sum_{t=1975}^{1979} \theta_t * D_o * 1(y=t) + \sum_{t=1981}^{1986} \pi_t * D_o * 1(y=t) + X_{o,y} + \varepsilon_{o,y} \quad (1)$$

where  $GFR_{o,y}$  is the general fertility rate in oblast  $o$  and year  $y$ ,  $\gamma$  is a set of year fixed effects that capture changes common to all oblasts,  $D_o$  equals one if an oblast was an early beneficiary and  $\delta$  is a set of oblast fixed effects that capture time-invariant oblast level differences. The dummy for the year before the start of the program,  $1(y = 1980)$ , is omitted which normalizes the estimates for  $\theta$  and  $\pi$  to zero in 1980.<sup>19</sup> The regression also includes the following limited set of covariates that vary at the year and oblast level and measure production, agricultural output and economic activity: amount of bricks, concrete, timber, meat, and canned goods produced, as well as the value of retail trade.<sup>20</sup> These covariates allow me to test for whether the change in fertility rates was due to some other coincidental economic shocks across regions. The point estimates of interest,  $\theta$  and  $\pi$ , directly test whether fertility rates were on parallel trends before the start of the benefits and whether estimates diverged after implementation.

The coefficient  $\pi_{1981}$  captures the treatment effect of the program on fertility rates in early beneficiary regions in the first two months of the program;  $\pi_{1982}$  captures the effect in the first full year of implementation. These coefficients may be biased upward if women in late

<sup>19</sup> For notational convenience,  $\theta$  is a vector that contains all the coefficients,  $\theta_t$  for years,  $t$  from 1975 to 1979, while  $\pi$  is a vector that contains all the coefficients,  $\pi_t$  for years,  $t$  from 1981 to 1986.

<sup>20</sup> These covariates represent a sample of the most relevant statistics available for that time period at the oblast and year level, which I manually entered using 1975-1992 “Narodnoe Hozyaystvo” yearbooks.

beneficiaries delayed childbearing in response to the benefit announcement. I test whether the fertility rate in late beneficiaries dropped discontinuously between November 1981 and October 1982 in response to the program announcement.<sup>21</sup> I find no evidence of this and instead estimate a statistically insignificant 0.63 percent increase in fertility rates in these regions in the year when they were not eligible for benefits.

The coefficients  $\pi_{1983}$  to  $\pi_{1986}$  capture the reversion of the mean difference to its pre-program level after the late beneficiary oblasts gained eligibility. If the late beneficiaries adjusted their fertility rates once they became eligible for the full year of benefits in 1983,  $\pi_{1983}$  to  $\pi_{1986}$  should be smaller in magnitude compared to  $\pi_{1982}$  when the two areas differed in eligibility.

### *C. Results: Effect of Maternity Benefits Using Annual Data*

Figure 3 displays estimates from specification (1), which represent the covariate-adjusted differences in fertility rates between the early and late beneficiaries in a year compared to the difference in 1980. The results are weighted by the population of women aged 15 to 44 in 1980 in each oblast.<sup>22</sup> The standard errors are clustered at the oblast-level to allow for an arbitrary correlation structure within an oblast.

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<sup>21</sup> For the sample of late beneficiary oblasts, I estimate  $GFR_{o,y,m} = \alpha + \gamma_1 y(m) + \beta_3 \text{post} + \delta_o + \delta_m + \epsilon_{o,y,m}$ , where the unit of observation is at the oblast,  $o$ , year,  $y$ , and month,  $m$ , level. The specification includes a linear time trend,  $y(m)$ , which accounts for any smooth fertility trends, a set of month fixed effects,  $\delta$ , which account for seasonality in births, and a dummy,  $\text{post}$ , for the period from November 1981 to October 1982. The coefficient of interest is  $\beta_3$  which tests for a discontinuous change in GFR in late beneficiaries during the year when they were not eligible for benefits but the early beneficiaries were. My result is robust to the inclusion of flexible polynomials in time. The details for the construction of month-level fertility rates are described later in this section.

<sup>22</sup> The motivation for weighting is to correct for heteroskedasticity that is related to population size in the oblast by year error terms, when the estimates of the treatment effect with or without weighting may differ (Solon et al. 2013, DuMouchel and Duncan 1983). I fail to reject that the treatment effect coefficients in weighted and unweighted specifications are the same using a Hausman test (Deaton 1997). Thus, I presented the weighted results in the paper.

These results support the hypothesis that increasing maternity benefits increased fertility rates. First, there is no difference in fertility rate trends in the early and late beneficiaries five years before program implementation. The point estimates for years 1975 to 1979 are individually indistinguishable from zero and follow a flat trend. Second, these results adjusted for covariates support the findings from the unadjusted series from figure 2. Third, the difference between early and late beneficiaries rises from 1981 to 1982 and reverts toward the program mean from 1983 to 1986.

Fertility rates rose immediately once early beneficiaries became eligible for benefits in 1981. As expected, the increase in GFR was larger when early beneficiaries were eligible for the whole year (1982) than for part of the year (1981) of benefits. Estimates in table 2 imply that GFR jumped by 2.4 and 6.2 births per 1,000 women of childbearing age in the first partial (1981) and full year (1982) of benefits. This represents a 3.2 and 8.2 percent increase over a pre-treatment mean of 76.0 in early beneficiaries. Fertility rates rose immediately once late beneficiaries became eligible for the full year of benefits in 1983. This is evidenced by the fact that estimates for years 1983 to 1986 are smaller in magnitude than estimates for 1982 and are not statistically different from zero. Thus, the difference between the GFR of early and late beneficiaries shrank due to an increase in fertility rates in late beneficiaries once they became eligible for benefits.

A potential threat to internal validity of these estimates could occur if coincidental policies or economic factors affected outcomes. It is unlikely that the estimated effect of maternity benefits is due to other policies or factors, because other factors would need to change in a specific order in early and late beneficiaries. In particular, my estimates will not capture the effects of maternity benefits if other factors affecting childbearing changed discontinuously in

1981 in early beneficiaries, and in 1982 in late beneficiaries. In addition to controlling for covariates, I also gather more aggregated data on additional relevant measures of economic activity to look for broader patterns in the evolution of the growth of industrial product, production of oil, and natural gas in a subset of early beneficiaries and in all of Russia.<sup>23</sup> In contrast to hypotheses that the increase in fertility rates was due to other factors, figures in appendix D show little evidence in changes in economic activity at the start of the program, because all indicators visually stayed on the same trend during the relevant period.

#### *D. Testing for the Role of Abortion Using Monthly Data*

The immediate increase in fertility rates could be due to an increase in foregone abortions as well as an increase in conceptions. In the 1980s, abortion was the most widely used method of fertility regulation in Russia, in part because women had limited access to and education on other types of contraception (Popov 1991). Abortions were legal up to the 12<sup>th</sup> week of pregnancy, but illegal abortions were also prevalent. Thus, fertility rates may have gone up six months after women found out about the benefits if only legal abortions were present. Fertility rates may have gone up even earlier if women had illegal abortions. Further, fertility rates may have gone up nine months after women found out about the benefits due to an increase in conceptions. The adjustment of fertility rates would not be as immediate, if women were using contraceptive methods that required medical help to remove or a waiting period until full fecundity.

The analysis using annual data hides the month-to-month dynamics of fertility rate adjustment that would allow me to distinguish between births due to foregone abortions or

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<sup>23</sup> I was only able to collect aggregated data on economic indicators in 16 of the early beneficiary oblasts. Data on these indicators is not available for all oblasts at the year and oblast level.

conceptions. To examine these patterns, I construct month-level fertility rates using retrospective information from the 2002 census to estimate the monthly number of births.<sup>24</sup>

The census does not include individuals who were born in Russia, and who emigrated or died before the census. Consequently, my estimates of the number of births using the census are lower than the true number of births due to mortality and mobility. To correct these estimates, I use information on the true number of births by oblast and year from the vital statistics data presented in the previous section. Specifically, I calculate the proportion of births that are still present in the census

$$p_{o,y} = (Births_{o,y}^{Census}) / (Births_{o,y}^{Vital})$$

where  $Births_{o,y}^{Census}$  is the number of births in oblast,  $o$ , in year,  $y$ , recorded in the census, and  $Births_{o,y}^{Vital}$  are the number of births in oblast,  $o$ , in year,  $y$ , in the vital statistics data. To adjust census estimates of the number of births, I scale them by the proportion of births that are still present in the census:

$$Births_{o,y,m}^{adjusted} = \frac{Births_{o,y,m}^{Census}}{p_{o,y}} \quad (2)$$

This procedure scales birth rates in each year by a common factor and assumes no differential mortality/mobility by month. Thus, these estimates preserve some of the information on seasonal variation but are rescaled to be comparable to the annual estimates in the previous section.

I use month-level fertility rates to estimate the following extension of specification (1):

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<sup>24</sup> These data need to be constructed because month-level fertility rates at the region-level are not published for Russia in my time period of interest.

$$FR_{o,y,m} = \alpha + \gamma_y + \gamma_{m,o} + \delta_o + \sum_{t=1978}^{1983} \sum_{k=1}^{12} \pi_{t(k)} * D_o * 1(y = t, m = k) + X_{o,y} + \varepsilon_{o,y,m} \quad (3)$$

where  $FR_{o,y,m}$  is measured in oblast  $o$ , year  $y$  and month  $m$ ,  $\gamma$  is a set of month-by-oblast fixed effects that capture seasonality in fertility rates in each oblast,  $1()$  is a dummy for month of observation, while the rest of the variables are the same as those defined in equation (1). Dummies for months in 1977 are omitted, which normalizes the estimates for  $\pi$  to zero in 1977.<sup>25</sup> The coefficient,  $\pi_{t(k)}$ , represents the difference between fertility rates in early and late beneficiaries in year,  $t$ , and month,  $k$ , compared to the difference in 1977.

Estimates of  $\pi$  in figure 4 using month-level census data suggest that the increase in fertility rates was partly due to foregone abortions and partly due to an increase in conceptions. Fertility rates in early beneficiary oblasts rose about six months after women found out about the benefits compared to fertility rates in late beneficiaries, which may be attributable to an increase in foregone abortions. This is before benefits began and may reflect the fact that women planning abortions may not have known when they were due. Further, fertility rates rose even further in the period between November 1981 and October 1982 and stayed higher in every month during that period, which may be attributable to an increase in conceptions. Finally, the difference in fertility rates declined once women in late beneficiaries also received the benefits.

## V. The Long-Run Effect of Maternity Benefits on Childbearing

This short-run increase in fertility rates may have been the result of two channels – women gave birth sooner, or women had children they would not have otherwise had. Given that the goal of maternity benefits was to induce women to have more children, I test whether

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<sup>25</sup> For notational convenience,  $\pi$  is a vector that contains all the coefficients,  $\pi_{t(k)}$  for years,  $t$ , from 1978 to 1983 and months,  $k$ , from 1 to 12.

maternity benefits resulted in an increase in completed childbearing in three ways. First, I examine the response in fertility rates by parity. Second, I estimate the effect of the program on fertility rates over its duration. Third, I estimate the change in the demographic composition of mothers after the start of the program.

#### *A. The Effect of Maternity Benefits on Fertility Rates by Parity*

As a first test of whether completed childbearing increased due to the program, I examine the response of fertility rates by parity. An increase in first birth fertility rates likely reflected women having desired children sooner, because most women in Russia had at least one child. However, an increase in higher parity fertility rates suggests women had children they would not have otherwise had. Because, fertility rates by parity and oblast are not published in Russia for my time period of interest, I construct first birth and higher order (second and higher order births) fertility rates using the 2010 Census data. I then use equation (1) to test whether the increase in fertility rates was due to first or higher parity births.

The results suggest that the short-run increase in fertility rates was due to an increase in completed childbearing, because only higher parity fertility rates increased after the start of benefits. Panel A in table 3 shows that the higher parity fertility rate in early beneficiaries increased by 17.9 percent (6.7 divided by the pre-program mean of 37.4) in the first year of benefit receipt.<sup>26</sup> Higher parity fertility rates in early and late beneficiaries followed a parallel trend five years before the start of the program in panel A of figure 5, which supports the internal validity of these results. The late beneficiaries responded to the program after they became eligible for it; the difference between the fertility rates of early and late beneficiaries shrank

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<sup>26</sup> I can only obtain fertility rates by parity using census data, so the data used in these regressions is at the month level. The specification is the same as in (1), but also includes month by oblast fixed effects.

starting in 1983. Panel B of figure 5 shows that first parity fertility rates did not change after the program, because there is no trend-break in the difference between early and late beneficiaries after the program compared to before the program began.

*B. The Effect of Maternity Benefits on Long-Run Childbearing*

As a second test of whether completed childbearing increased due to the program, I examine the evolution of fertility rates for the ten-year duration of the program. Fertility rates would have risen temporarily, and then have fallen below previous levels, if women only shifted the timing of childbearing. If true, this would have resulted in a zero net increase in fertility rates over the duration of the program. However, fertility rates would have stayed consistently higher for the duration of the program, if women increased their completed childbearing.

To quantify the effects of maternity benefits on childbearing for the duration of the program, I employ an event-study framework. This allows me to study the dynamics of the response over a longer time period. My specification compares outcomes within an oblast before and after the start of the program compared to outcomes right before the start of the program (Bailey et al. 2014). The specification is the following:

$$GFR_{o,y} = \alpha + \gamma_y + \delta_o + \sum_{t=-5}^{-1} \theta_t * 1(y-y^*=t) + \sum_{t=1}^{11} \pi_t * 1(y-y^*=t) + X_{o,y} + \varepsilon_{o,y} \quad (4)$$

where  $y^*$  is the year before the start of the program which is different for early and late beneficiaries ( $y^* = 1980$  for early beneficiaries, while  $y^* = 1981$  for late beneficiaries), and  $1()$  is a dummy that represents years relative to the start of the program. In this specification,  $\pi_t$  measures the effect of maternity benefits on fertility rates  $t$  years after the start of the program. Note that  $t=0$  is omitted. The coefficient  $\pi_t$  when  $t=1$  should be the smallest in equation (4),

because in 1981 the early beneficiary regions while in 1982 the late beneficiary regions only had the benefit in place for two months. Estimates of  $\theta$  document whether pre-existing trends bias estimates of  $\pi$ , and whether “effects” preceded the program.

The event-study analysis has a causal interpretation if the timing of benefit receipt was conditionally random, which is supported by the fact that fertility rates were on parallel trends in early and late beneficiaries. Also, the fact that the government mandated the order of benefits provides evidence against specific types of regions choosing to receive the benefits first. There are no official published reasons for the choice of the order of benefits by the government. But, because one of the goals was to increase childbearing in less populated and labor shortage regions, it is reasonable that the government prioritized the early beneficiary regions that met these criteria.

I find evidence that completed childbearing increased as a result of the program, because the increase in childbearing was sustained for the entire duration of the program. In figure 6, I present event-study coefficients  $\theta$  and  $\pi$  from specification (4) that show changes in fertility rates relative to the year before the benefits started. Estimates to the left of the vertical axis capture the evolution of fertility rates before the start date, and estimates to the right of the vertical axis capture the effect of maternity benefits on fertility rates 1 to 10 years after the start of the program.

The maternity benefit program was associated with a sustained increase in childbearing, which was due to an increase in higher order fertility rates. Panel A of figure 6 shows that fertility rates were higher for ten years after the start of the program compared to before the program. Table 4 shows that fertility rates were on average 14.6 percent higher 1 to 10 years

after the start of the program. There is no evidence that differential pre-existing trends may bias this analysis or that effects preceded the program, because estimates of  $\theta$  display a flat trend in the five years before the start of the program. Consistent with the previous section, panel B of figure 6 exhibits a trend-break and shows that higher parity fertility rates were on average 27.8 percent higher for the ten-year duration of the program.

### *C. The Effect of Maternity Benefits on the Composition of Mothers*

As a third test of whether completed childbearing increased due to the program, I quantify changes in the composition of mothers who gave birth after the program. The increase in fertility rates was due to women having desired children sooner, if women who had children after the program were younger and waited less to have another child compared to women who had children before the program. No publicly available vital statistics data on the characteristics of mothers in Russia are available. Instead, I use the Generations and Gender Survey conducted in 2004 to obtain characteristics of mothers based on when their children were born.

I find evidence that completed childbearing increased as a result of maternity benefits, because women who had children after the program were older and waited longer to have a child. The panels of figure 7 present results from specification (4), where I use age of mother at birth (panel A), years since last birth (panel B) and number of previous children at the time of birth (panel C) as dependent variables.<sup>27</sup> Women who had children 1 to 10 years after the start of the program were consistently older, had more previous children, and waited more years to have a child compared to women who had children before the start of the program. These effects are summarized in table 5, and are statistically different from zero. Again, there is no evidence that

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<sup>27</sup> Time relative to treatment is 0 if birth year is 1981 for early beneficiaries and if birth year is 1982 for late beneficiaries. Also, the unit of observation in these regressions is at the birth-mother level: I treat each birth of the mother as a separate observation. The standard errors are clustered at the oblast level.

differential pre-existing trends may bias this analysis or that effects preceded the program, because estimates of  $\theta$  display a flat trend in women's characteristics for five years before the start of the program.

## **VI. Testing for Mechanisms in Responses to Maternity Benefits**

The increase in childbearing after the expansion of maternity benefits may be due to three channels. First, if the opportunity cost channel predominates then women with lower opportunity costs of having a child may increase childbearing by more, because the benefit is a higher fraction of their salary, and they have a more flexible work schedule to accommodate another child. Second, if the child cost channel predominates then women for whom the cost of raising a child is lower, for instance lower cost of childcare, may increase childbearing by more. Third, if the housing size channel predominates then women who do not have housing size constraints may increase childbearing by more.

Women in rural areas may have increased childbearing by more because of the opportunity cost and housing size channels, while less educated women may have increased childbearing by more because of the opportunity cost channel. Women in rural areas were mostly employed as manual laborers and were heavily underrepresented in the prestigious occupations of machine operators which required special training and skill (Bridger 1987). As a result, women's opportunity costs in rural areas were low due to low wages, and the seasonal nature of their work, which gave them greater flexibility in caring for a child compared to women in urban areas. Moreover, women in rural areas had larger housing which allowed them more room to have another child, whereas women in urban areas were more space constrained. Less educated

women had lower opportunity costs due to lower wages and due to less of a penalty on wages from taking longer leave because of slower skill depreciation.

To establish the importance of these different channels, I examine whether effects of maternity benefits differed across oblasts. In particular, I analyze whether fertility rates increased by more in more rural compared to more urban areas or less educated compared to more educated areas. For this purpose, I obtain characteristics of regions using the 1979 Russian census and follow the approach of Finkelstein (2007):

$$GFR_{o,y} = \alpha + \gamma_y + \sum_{t=-6}^{-1} \theta_t * Z_o * 1(y-y^*=t) + \sum_{t=1}^{10} \pi_t * Z_o * 1(y-y^*=t) + \delta_o + \varepsilon_{o,y} \quad (5)$$

where  $Z_o$  represents continuous variables at the oblast level measured in 1979 to be included in separate regressions: share of women age 15 to 44 in an oblast who were living in a rural area, and share of individuals age 10 and older who have not completed high school. All other covariates remain the same as in specification (4). This specification is the same as specification (4), except now the event-year dummies,  $1(y - y^* = t)$ , are interacted with  $Z_o$ .

As before, this empirical strategy tests for a break in any pre-existing differences in the level or trend of fertility rates across regions around the time of the start of the program that is correlated with  $Z_o$ . The identifying assumption is that without benefits the differences in fertility rates before the start of the program would have continued on the same trend. Thus, this strategy does not assume that areas that differ in their composition of residents had the same level or growth of fertility rates before the program.

The results show that increased maternity benefits are associated with a greater increase in fertility rates among more rural and less educated areas. Figure 8 plots estimates of  $\theta$  and  $\pi$

from specification (5) that correspond to interactions of the event-year dummies with independent variables measuring a region's rural (panel A) and education status (panel B) in separate regressions. The time pattern of  $\pi$  (estimates to the right of the vertical axis) presents changes in fertility rates after the start of the program in areas with larger expected increases in childbearing relative to areas with smaller expected increases. The dashed lines indicate a 95-percent confidence interval for each coefficient.

Both panels show a notable level-shift in coefficients after the start of the program in a region. The coefficients are individually statistically indistinguishable from zero in the years leading up to the start of the program. This indicates that before the start of the program fertility rates evolved similarly in areas with different shares of urban women and educated individuals. However, after the start of the program the coefficients,  $\pi$ , jump discontinuously in both panels, and are summarized in table 6. This indicates a larger increase in fertility rates in more rural compared to more urban areas, and a larger increase in fertility rates in less educated compared to more educated areas. After the initial jump, the GFR in more rural or less educated areas continued increasing by more for the first eight years after the start of the program. However, the difference across regions disappeared ten years after the start of the program, which is consistent with benefits losing their value due to inflation and the end of the program.

The fact that fertility rates increased by more in more rural and less educated regions indicates that the increase in fertility rates was largely due to the opportunity cost channel. This provides suggestive evidence that the magnitudes of effects of maternity benefits depend on the opportunity costs of women's work. Areas where women earn less have more flexible schedules and have flatter age-earnings profiles may experience larger increases in fertility rates due to maternity benefits.

## VII. Conclusion and Discussion

Low fertility rates are an important concern for many OECD countries, who have implemented various family friendly policies. I find an immediate response in fertility rates after the introduction of paid parental leave and cash transfers in Russia. Moreover, I find that the effects on fertility rates persist in the long run. These results indicate that maternity benefits affected both the timing and the number of children women had. This study provides new evidence on the topic of whether family policy can raise childbearing.

These results imply that childbearing is elastic with respect to the cost of a child. A back-of-the envelope calculation quantifies this price elasticity. I use estimates from Russian demographers to calculate the cost of a child as the sum of the costs over the first 18 years of a child's life (Valentei 1987). In Russia, maternity benefits decreased the cost of a child by 2.2 percent.<sup>28</sup> According to estimates in table 2, this is associated with an 8.2 percent increase in fertility rates (6.2/76) during the year after the start of the program, which means that the short-run price elasticity of childbearing equals to -3.7 (8.2/2.2).

How does the effect of maternity benefits on childbearing in Russia compare to effects of family benefits in other countries? To address this question, I compare the estimates of short-run elasticities from other studies, which figure 9 shows range from -4.4 to 0.54. The estimate of the elasticity in this study is in this range. I also construct corresponding confidence intervals for these elasticities using a parametric bootstrap method (Johnston and DiNardo 1997).<sup>29</sup> Figure 9 shows that in Austria the elasticity was equal to -4.4, in Spain it equaled -3.8, in Canada it was -

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<sup>28</sup> The details of this calculation are in appendix E.

<sup>29</sup> The procedure for calculating elasticities and their corresponding confidence intervals in other studies is described in appendix E.

4.1, while in Israel it was -0.54. The confidence intervals of all but the study on Israel overlap, which indicates that the estimate in the Russian context is in the range of other countries.<sup>30</sup>

This study is the first to find a positive long-run effect of maternity benefits on childbearing. My estimates imply that the program induced about 5 million births during its ten-year duration at a cost of roughly 2,830 rubles per birth induced.<sup>31</sup> In terms of what that meant at the time, the cost represented 1.4 times the average national yearly salary in 1980. How applicable is this finding in the context of other countries?

The long-run effect of maternity benefits may vary across countries because it may depend on the cost of a child or the opportunity cost of the woman's time. Maternity benefits may have had a long-run effect on childbearing in Russia, because women had lower opportunity costs and lower costs of raising children. In Russia, women earned substantially less than men, the age-wage profile was flatter than in other countries, and childcare was widespread and heavily subsidized. What are the policy implications? The results of this study suggest that maternity benefits may be effective in their goal of increasing long-run childbearing. Whether maternity benefits can increase long-run childbearing in other countries depends on opportunity costs of women, and the interaction of maternity benefits with other social or public programs.

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<sup>30</sup> This comparability may be less surprising because Russia in 1980 is comparable in many respects to OECD countries today in terms of key demographic and economic variables. Table 7 shows the characteristics of Russia in 1980 compared to characteristics of Austria, Spain, Italy, the United States, and Sweden in 2010-2013. Russia had a below replacement fertility rate, a high women's labor force participation rate where women earned less than men, and high rates of preschool enrollment. Also, the inequality in Russia is similar to European countries today, but lower than in the US.

<sup>31</sup> I obtain estimates of the number of induced births by multiplying the population of women age 15 to 44 in Russia in 1980 (32,696,400) by the event-study estimates from specification (4) and summing over event-years 1 to 11. I estimate that the program cost was about 14 billion rubles, where I assume that women took the entire leave (thus giving an upper bound of the cost). I calculate the cost from parental leave outlays by multiplying the fixed payments by 10 months (maximum time of leave) for each birth; I calculate the cost from the cash transfers by multiplying the fixed transfer by the number of births; I calculate the additional cost in terms of extra maternity benefits for births that were induced by multiplying the average national monthly salary by 4 months for the estimated induced 5 million births.

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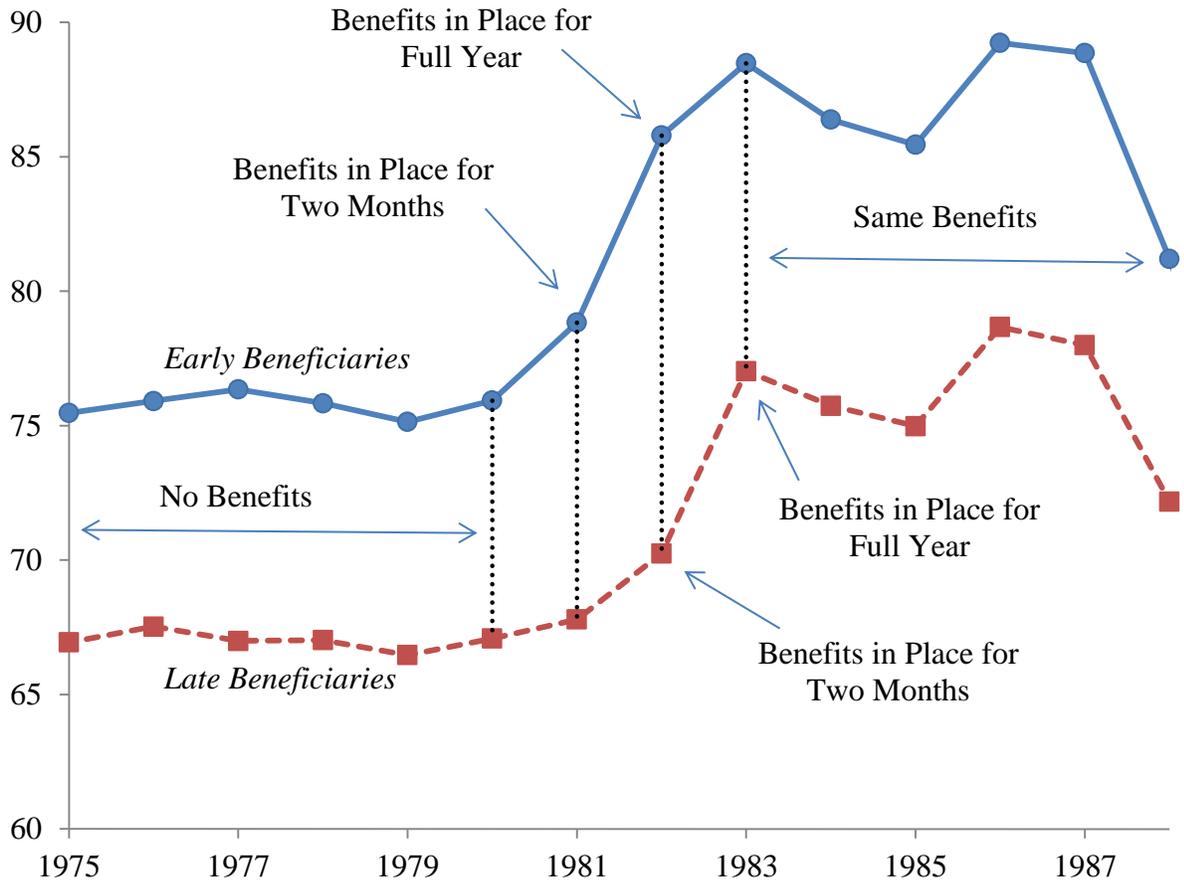
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**Figure 1. Map of Benefit Roll-Out Across Russia**



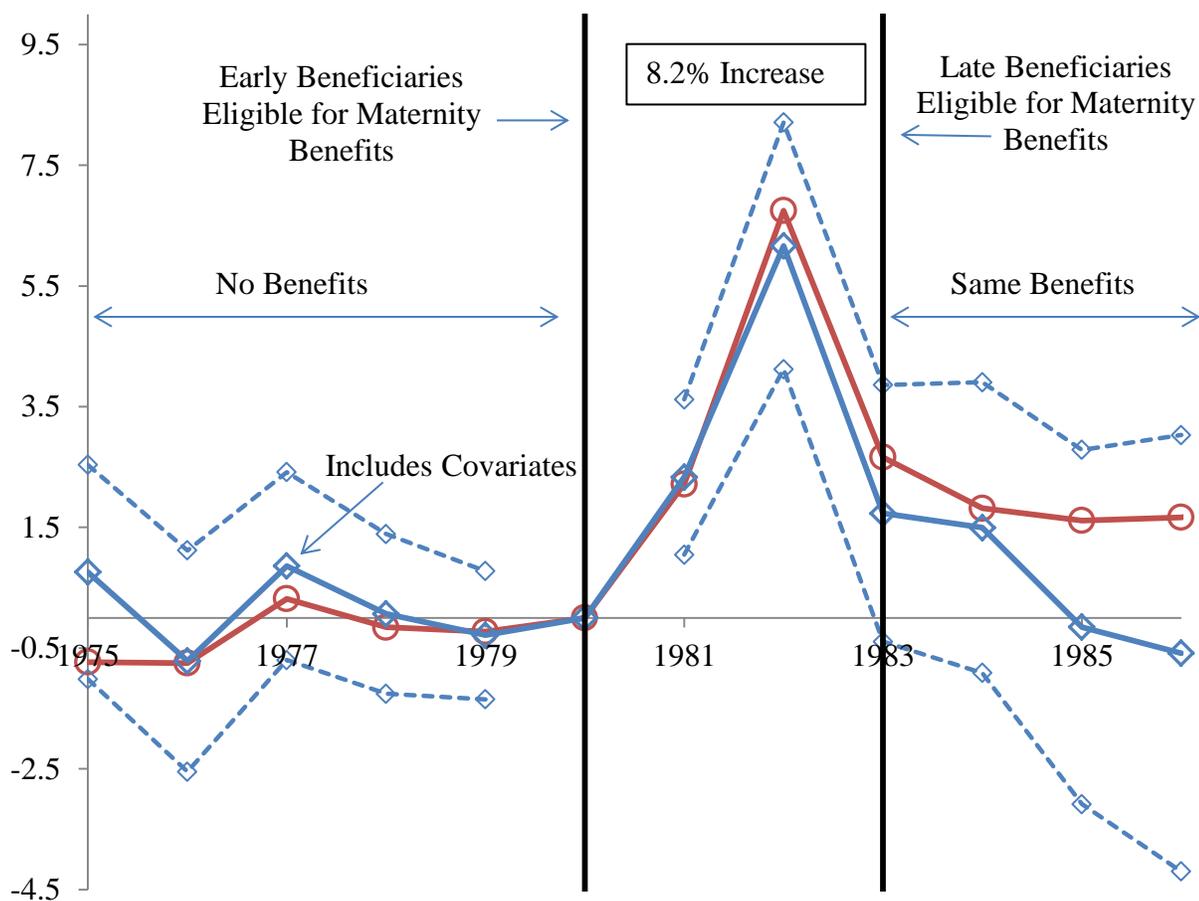
Notes: The early beneficiaries (Northern, Siberia and Far East regions) where maternity benefits started in November, 1981 are shaded. The late beneficiaries (rest of Russia) where maternity benefits started in November, 1982 are white.

**Figure 2. Descriptive Evidence of the Effect of Maternity Benefits on Fertility Rates**



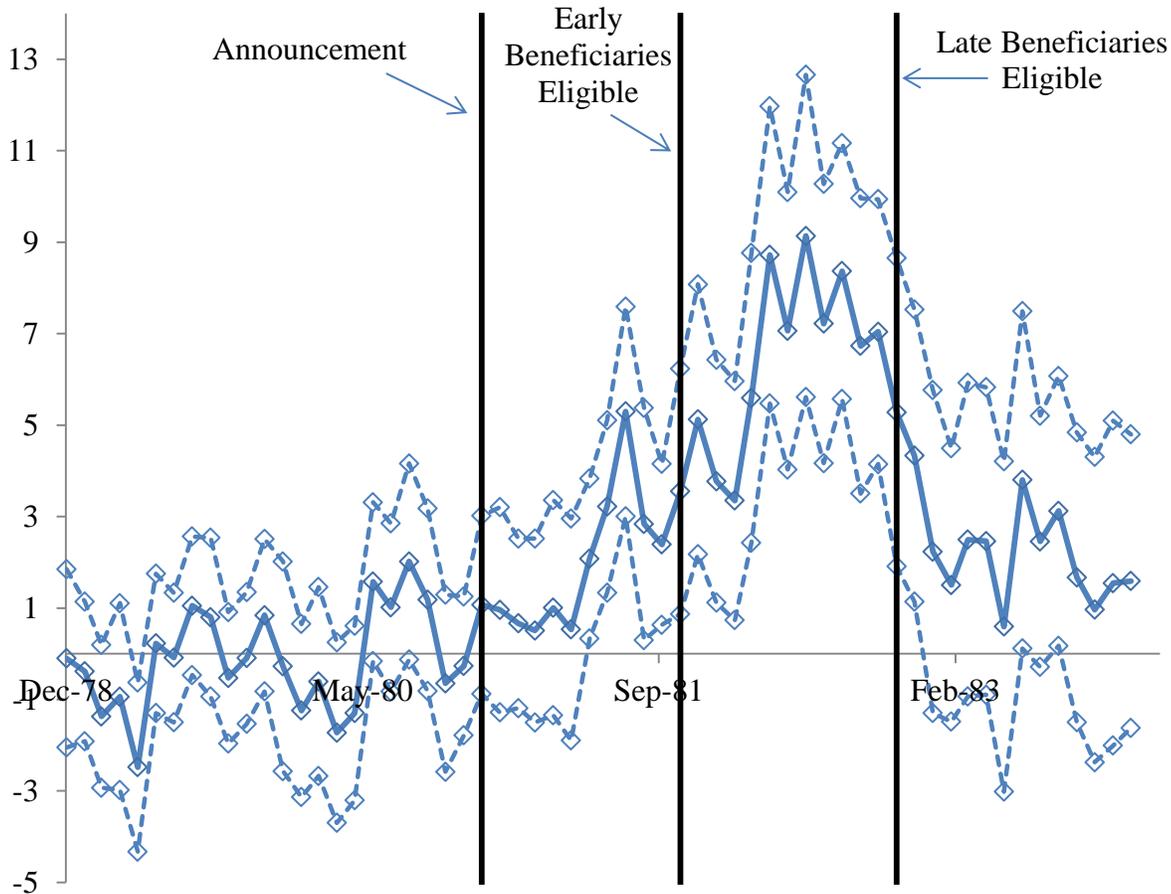
Notes: The figure plots the evolution of general fertility rates (GFR) in the early and late beneficiary regions. The GFR is the number of births per thousand women ages 15 to 44. Sources: Russian Federal State Statistics Service (Rosstat) and the 1989 Russian Census.

**Figure 3. The Short-Run Effect of Maternity Benefits on Fertility Rates**



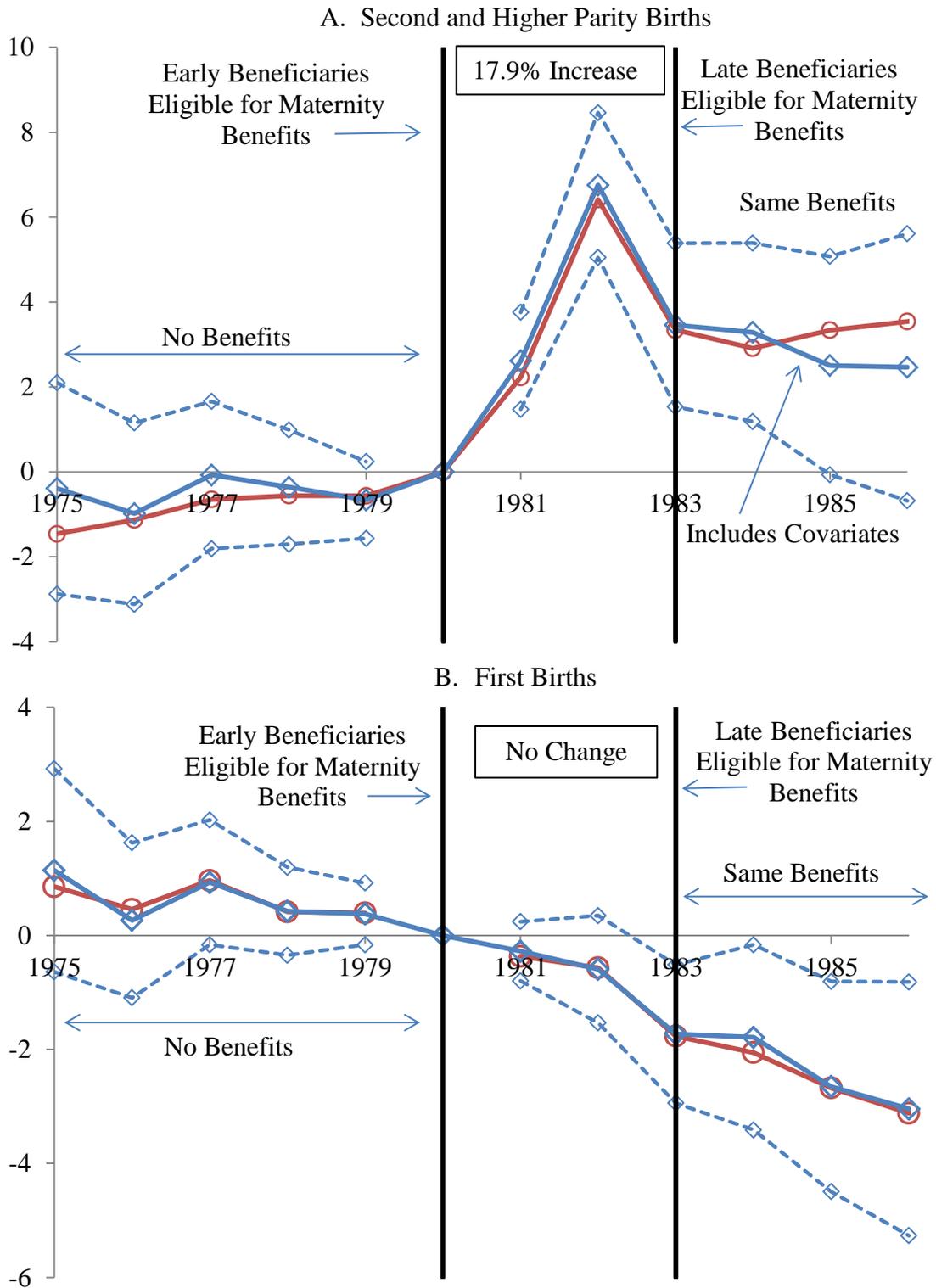
Notes: These coefficients represent the difference in GFR (general fertility rates) between the early and late beneficiary regions in each year relative to the difference in 1980. I present  $\theta$  and  $\pi$  from equation (1) using GFR as a dependent variable. The coefficient on year 1981 presents the effect of the program when maternity benefits were in place in early beneficiary regions for two months. The coefficient on year 1982 presents the effect of the program when maternity benefits were in place in early beneficiary regions for the full year. For the coefficients shown as circles the model includes year and oblast fixed effects; for the coefficients shown as diamonds the model also adds oblast by year covariates. Heteroskedasticity-robust standard errors clustered by oblast construct 95-percent, point-wise confidence intervals (dashed lines) for the coefficients that include oblast by year covariates. Regressions are weighted by the number of women who are ages 15 to 44 living in an oblast in 1980. See appendix B for details on the fertility rate construction, and appendix C on data descriptions. Sources: Rosstat, 1989 Russian Census, 1975-1986 “Narodnoe Hozyaystvo” yearbooks.

**Figure 4. The Short-Run Effect of Maternity Benefits on Monthly Fertility Rates**



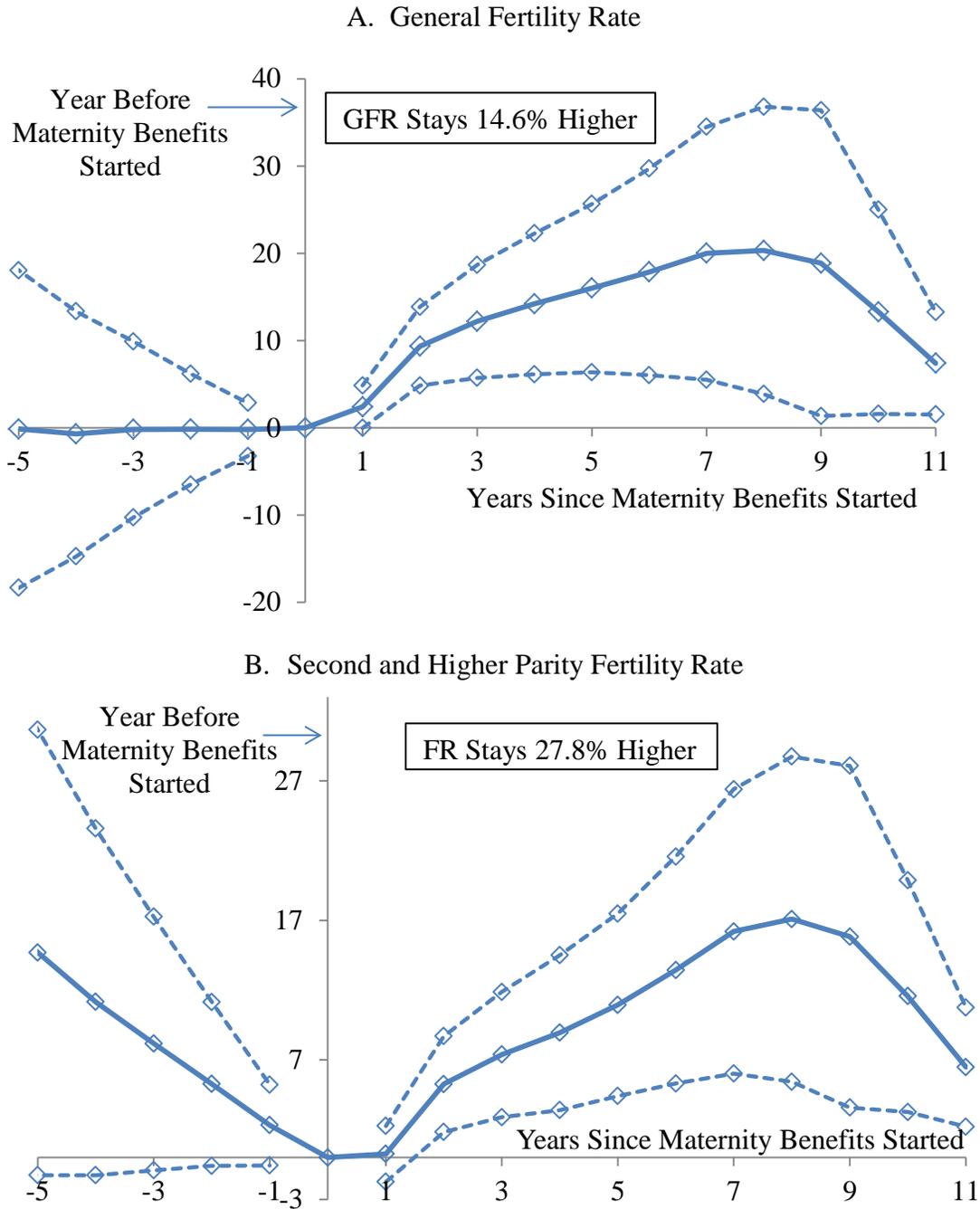
Notes: These coefficients represent the monthly difference in fertility rates between early and late beneficiaries from January 1978 to December 1983 relative to the difference in 1977, which is 0 by construction. I present  $\theta$  and  $\pi$  from equation (3). I use monthly fertility rates constructed using 2002 Census data and adjusted for mortality and mobility using equation (2). Monthly fertility rates represent number of births in a month per 1,000 women ages 15 to 44 in a year, and they have been multiplied by 12 to match the scale of the general fertility rate estimates in figure 3. Details on the fertility rate construction are in appendix B. The model includes year, oblast, month by oblast fixed effects, and annual covariates for each oblast. Weights are the number of women who are ages 15 to 44 living in an oblast in 1980. Heteroskedasticity-robust standard errors clustered by oblast construct 95-percent point-wise confidence intervals (dashed lines). Sources: Rosstat, 1989 and 2002 Russian Censuses, and 1975-1986 “Narodnoe Hozyaystvo” yearbooks.

**Figure 5. The Short-Run Effect of Maternity Benefits on Fertility Rates by Parity**



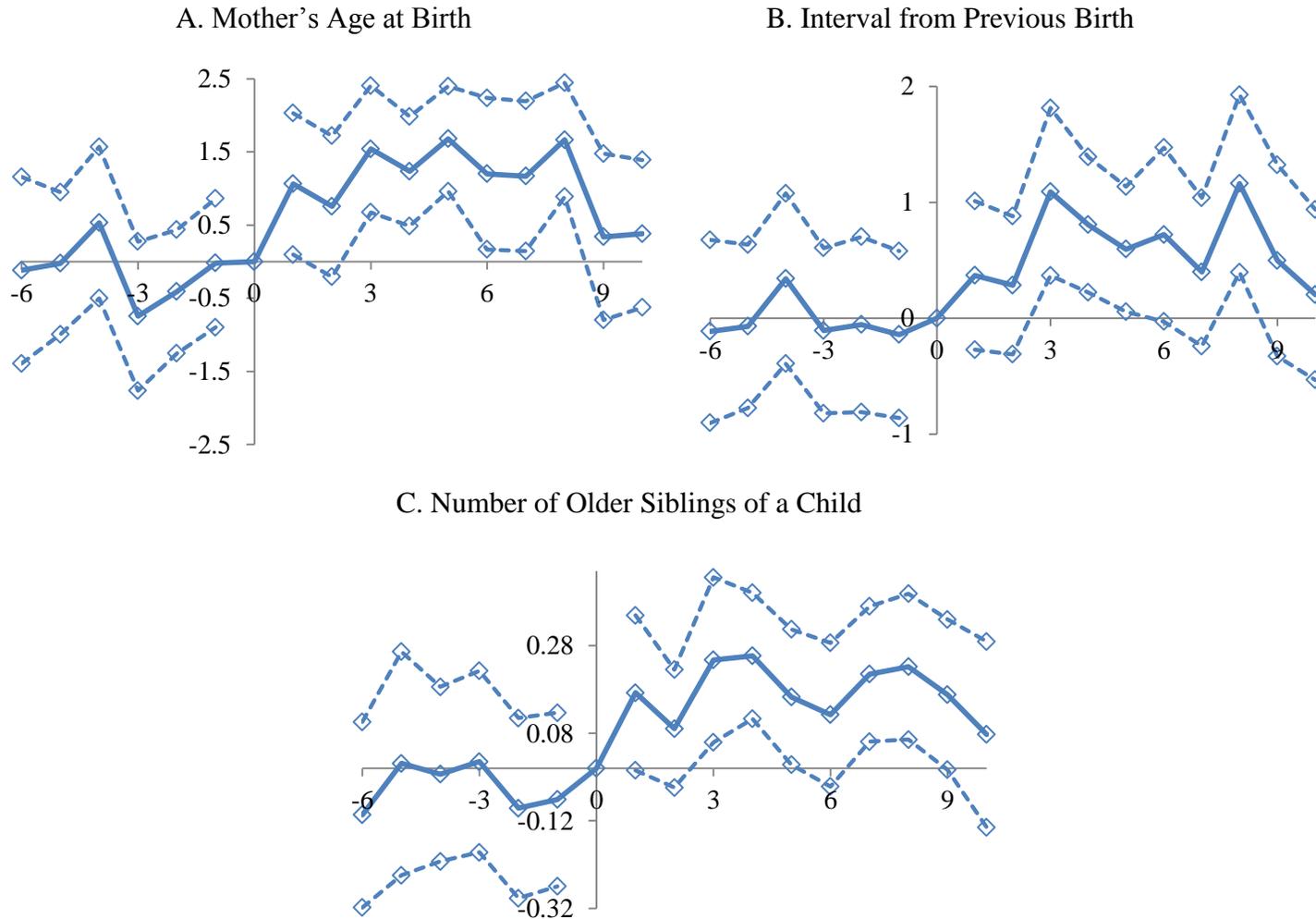
Notes: See notes for figure 3. The unit of observation is at the year-month-oblast level, and regressions include month by oblast fixed effects. Sources: Rosstat, 1989 and 2010 Russian Censuses, and 1975-1986 “Narodnoe Hozaystvo” yearbooks.

**Figure 6. The Long-Run Effect of Maternity Benefits on Fertility Rates**



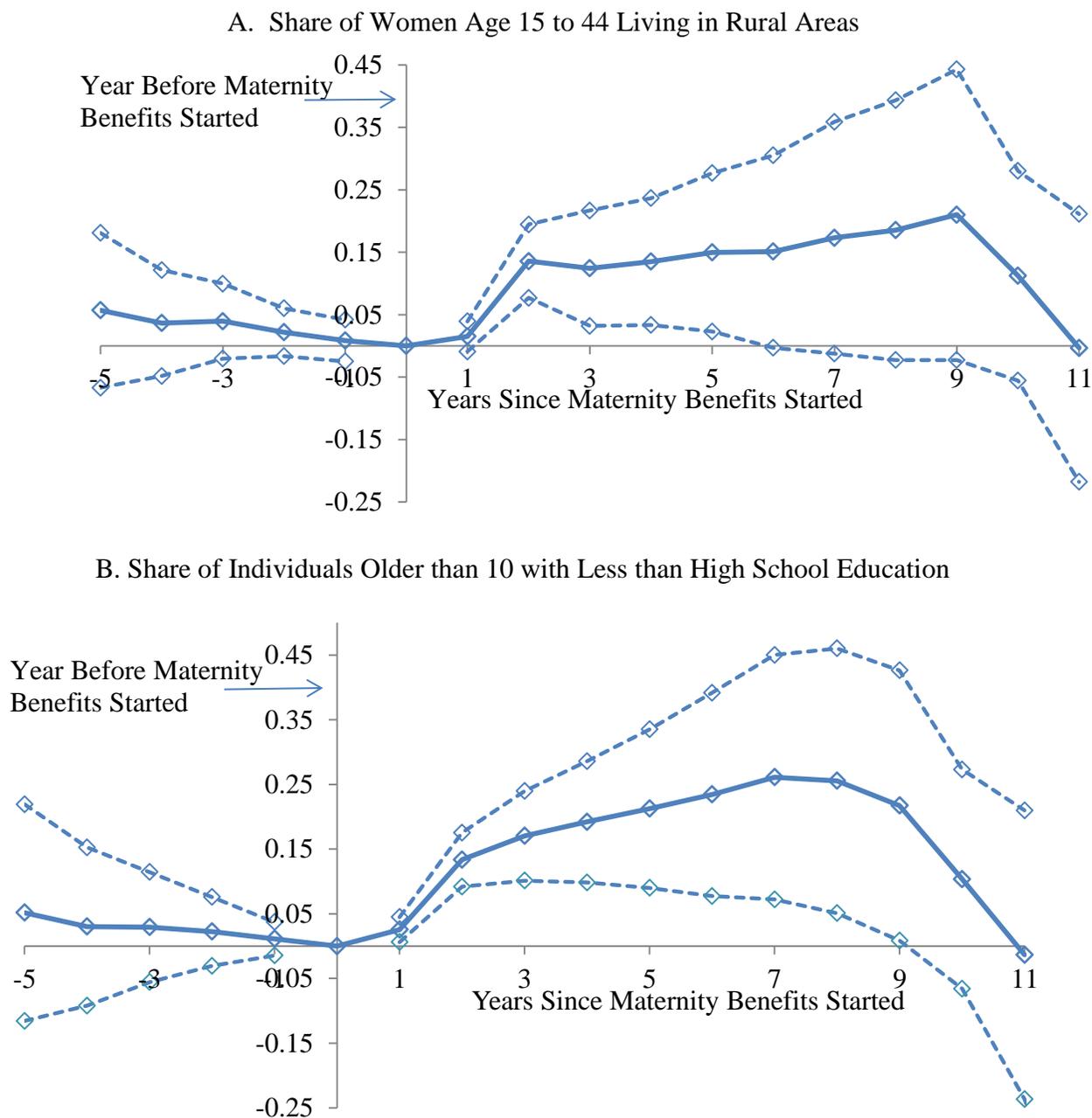
Notes: These coefficients show the evolution of fertility rates conditional on covariates 5 years before and 11 years after maternity benefits started. I present  $\theta$  and  $\pi$  from equation (4) using GFR as a dependent variable in panel A, and using second and higher parity fertility rate as a dependent variable in panel B. Years since maternity benefits started equal to zero if birth year equals 1980 in early beneficiary oblasts, and equal to zero if birth year equals 1981 for late beneficiary oblasts. Regressions are weighted by the number of women who are ages 15 to 44 living in an oblast in 1980. Heteroskedasticity-robust standard errors clustered by oblast construct 95-percent, point-wise confidence intervals (dashed lines). Sources: See sources for figure 5.

**Figure 7. The Effect of Maternity Benefits on the Demographic Composition of Mothers**



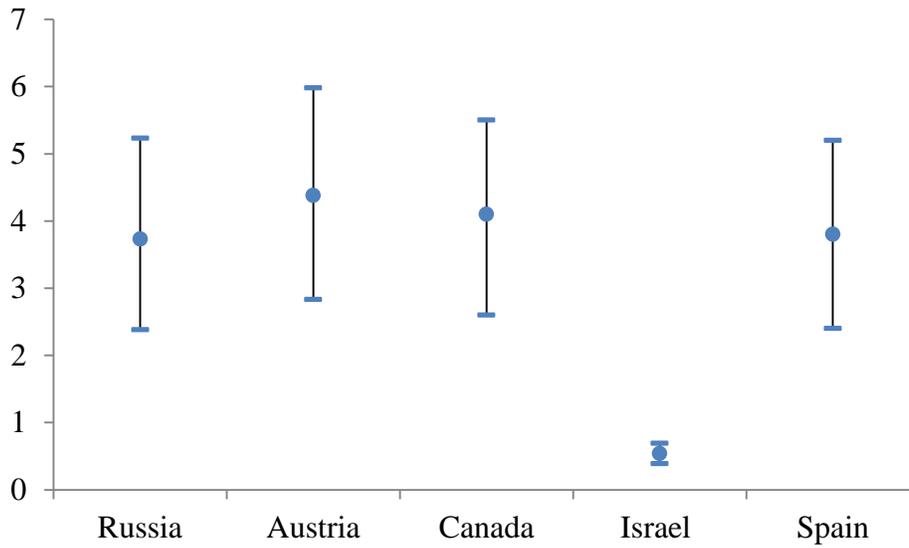
Notes: These coefficients show the evolution of the composition of mothers conditional on covariates before and after maternity benefits started. I present  $\theta$  and  $\pi$  from equation (4) using mother's age at birth, interval from previous birth and number of older siblings of a child as dependent variables. The unit of analysis is a mother-birth observation. The x-axis presents years since benefits started. Years since benefits started equal to zero if birth year equals 1981 for early beneficiaries, and equal to zero if birth year equals 1982 for late beneficiaries. The coefficient on years since treatment=0 is normalized to zero. Results are weighted using survey sampling weights. Heteroskedasticity-robust standard errors clustered by oblast construct 95-percent, point-wise confidence intervals (dashed lines). Source: Generations and Gender Survey, 2004.

**Figure 8. Heterogeneous Responses of Fertility Rates to Maternity Benefits Across Regions**



Notes: The coefficients show the pattern over time in GFR in regions with more rural compared to regions with more urban resident women (panel A), and in regions with less educated compared to regions with more educated individuals (panel B). I present  $\theta$  and  $\pi$  from equation (5) using GFR as a dependent variable. These coefficients present interactions of event-year dummies with a continuous oblast level characteristic measured in 1979 (share of women age 15 to 44 in rural areas {panel A}; share of individuals with less than high school education {panel B}). Years since maternity benefits started equal 0 if birth year equals 1980 in early beneficiaries, and if birth year equals 1981 for late beneficiaries. Regressions are weighted by the number of women who are ages 15 to 44 living in an oblast in 1980. Heteroskedasticity-robust standard errors clustered by oblast construct 95-percent, point-wise confidence intervals (dashed lines). Sources: 1979 and 1989 censuses, Rosstat, 1975-1992 “Narodnoe Hozyaystvo”.

**Figure 9. Comparison of Short-Run Childbearing Elasticities across Countries**



Notes: This plot presents estimates of short-run elasticities of childbearing with respect to the cost of a child until the child is 18 years old in Russia (calculated in this paper) and in other countries based on estimates from other studies. The plot shows the estimate of the elasticity (circle) which is surrounded by its 95-percent confidence interval. The figure presents estimates for the following countries: Austria (Lalive and Zweimüller 2009), Canada (Milligan 2005), Israel (Cohen et al. 2013), and Spain (Gonzalez 2012). Details of elasticity and confidence interval constructions are in appendix E.

**Table 1. Characteristics of Early and Late Beneficiary Oblasts**

	Early Beneficiaries	Late Beneficiaries
<i>Proportion of Population in 1979</i>		
with college degree among employed	9.0	9.1
with less than high school among employed	46.2	48.5
employed among women ages 15 to 54	84.1	85.3
in rural areas among women ages 15 to 44	25.7	24.9
in preschool among children younger than 7	49.0	50.2
65 years and older among women	9.6	14.7
Number of Oblasts	37	51
Share of Total Population Living in Oblasts	25.5	74.5

Notes: These statistics are based on the entire population of Russia. Sources: 1979 Census, 1979 “Narodnoe Hozyaystvo” Yearbook.

**Table 2. The Short-Run Effect of Maternity Benefits on Fertility Rates**

Dependent Variable	(1)	(2)	(3)
	General Fertility Rate		
	Before Program Mean in 1980: 75.95		
<i>Before Program</i>			
(Years 1975 to 1979)*Early Beneficiary	-0.309 [0.747]	-0.309 [0.779]	0.142 [0.632]
<i>After Early Beneficiaries Eligible</i>			
(Year 1981)*Early Beneficiary	2.218 [0.618]	2.218 [0.645]	2.354 [0.634]
(Year 1982)*Early Beneficiary	6.748 [1.291]	6.748 [1.347]	6.192 [1.031]
<i>After Everyone Eligible</i>			
(Years 1983 to 1986)*Early Beneficiary	1.940 [2.139]	1.940 [2.233]	0.502 [1.292]
R-squared	0.152	0.945	0.962
<b>Covariates</b>			
	Year FE	Year FE, Oblast FE	Year FE, Oblast FE, X <sub>o,y</sub>
Oblast-year cells	984	984	984
Oblasts	82	82	82

Notes: These coefficients represent the difference in GFR between the early beneficiaries compared to the late beneficiaries in grouped years relative to the difference in 1980. Standard errors clustered at the oblast-level are in brackets. These results are comparable to the coefficients in figure 3. These coefficients represent interactions of the grouped year dummies and the dummy for the early beneficiary regions using the regression model described in equation (1). The grouped years represent: the period before the start of the program (years 1975 to 1979), the year when benefits were in place for only two months (1981) in early beneficiaries, the year when the benefits were in place for the whole year (1982) in early beneficiaries, and the period after all regions became eligible for benefits (years 1983 to 1986). I omit the year before the start of the program (1980), so estimates for that year are normalized to zero. The unit of observation is oblast by year. Coefficients in column 1 include year fixed effects, coefficients in column 2 add oblast fixed effects, while coefficients in column 3 add annual oblast-level covariates where the covariates are the same as included in equation (1). Regressions are weighted by the number of women who are ages 15 to 44 living in an oblast in 1980. Sources: See notes for figure 3.

**Table 3. The Short-Run Effect of Maternity Benefits on Fertility Rates by Parity**

Dependent Variable	(1)	(2)	(3)
	A. Second and Higher Order Birth		
Before Program Mean in 1980: 37.44			
<i>Before Program</i>			
(Years 1975 to 1979)*Early Beneficiary	-0.932 [0.654]	-0.932 [0.683]	-0.385 [0.766]
<i>After Early Beneficiaries Eligible</i>			
(Year 1981)*Early Beneficiary	2.589 [0.579]	2.589 [0.604]	2.616 [0.570]
(Year 1982)*Early Beneficiary	7.325 [1.137]	7.325 [1.187]	6.717 [0.865]
<i>After Everyone Eligible</i>			
(Years 1983 to 1986)*Early Beneficiary	4.371 [1.885]	4.371 [1.968]	2.856 [1.164]
R-squared	0.212	0.914	0.932
Dependent Variable	B. First Birth		
	Before Program Mean in 1980: 38.50		
<i>Before Program</i>			
(Years 1975 to 1979)*Early Beneficiary	0.617 [0.449]	0.617 [0.469]	0.523 [0.509]
<i>After Early Beneficiaries Eligible</i>			
(Year 1981)*Early Beneficiary	-0.367 [0.256]	-0.367 [0.267]	-0.259 [0.260]
(Year 1982)*Early Beneficiary	-0.567 [0.429]	-0.567 [0.448]	-0.524 [0.465]
<i>After Everyone Eligible</i>			
(Years 1983 to 1986)*Early Beneficiary	-2.403 [0.864]	-2.403 [0.902]	-2.350 [0.830]
R-squared	0.862	0.951	0.954
Covariates	Year FE, Month FE	Year FE, Month-Oblast FE, Oblast FE	Year FE, Month-Oblast FE, Oblast FE, X <sub>o,y</sub>
Oblast-year-month cells	11,808	11,808	11,808
Oblasts	82	82	82

Notes: See notes for table 2. The unit of observation is at the year-month-oblast level. These results are comparable to the coefficients displayed in figure 5. Sources: See sources for figure 5.

**Table 4. The Long-Run Effect of Maternity Benefits on Fertility Rates**

Dependent Variable	(1)	(2)	(3)	(4)	(5)	(6)
	General Fertility Rate Before Program Mean: 69.98			Second and Higher Order Fertility Rate Before Program Mean: 34.60		
<i>Before Program</i>						
(Event Years -5 to -1)	-0.0137 [1.020]	-0.0137 [1.050]	-0.0610 [1.018]	1.434 [1.000]	1.434 [1.029]	1.359 [0.985]
<i>After Program</i>						
(Event Years 1 to 3)	8.994 [1.422]	8.994 [1.464]	9.063 [1.497]	7.047 [0.979]	7.047 [1.008]	7.004 [0.999]
(Event Years 4 to 6)	10.62 [2.178]	10.62 [2.242]	10.71 [2.277]	9.933 [1.590]	9.933 [1.637]	9.985 [1.647]
(Event Years 7 to 10)	10.77 [3.345]	10.77 [3.444]	10.84 [3.463]	11.71 [2.589]	11.71 [2.664]	11.82 [2.636]
R-squared	0.279	0.914	0.914	0.286	0.885	0.887
Covariates	Year FE	Year FE, Oblast FE	Year FE, Oblast FE, $X_{o,y}$	Year FE, Month FE	Year FE, Month-Oblast FE, Oblast FE	Year FE, Month-Oblast FE, Oblast FE, $X_{o,y}$
Observations	1,444	1,444	1,444	17,328	17,328	17,328
Oblasts	82	82	82	82	82	82

Notes: The coefficients represent dummies of grouped years to maternity benefits (event years) from the regression equation (4) where GFR (columns 1 to 3) and higher order fertility rates (columns 4 to 6) are dependent variables. Standard errors clustered at the oblast-level are in brackets. These estimates are directly comparable to coefficients presented in figure 6. The grouped years represent: the years before maternity benefits started (from 5 to 1 years before the program), 1 to 3, 4 to 6, and 7 to 10 years after the program started. The year before maternity benefits started is omitted, so estimates are normalized to zero for that event year. The unit of observation is oblast by year for columns 1 to 3, while the unit of observation is oblast by year by month for columns 4 to 6. Regressions are weighted by the number of women who are ages 15 to 44 living in an oblast in 1980. Sources: see sources for figure 6.

**Table 5. The Long-Run Effect of Maternity Benefits on the Composition of Mothers**

Dependent Variable	(1) <b>Mother's Age at Birth</b>	(2) <b>Number of Older Siblings of Child</b>	(3) <b>Interval From Previous Birth</b>
Before Program Mean	24.85	0.782	2.723
<i>After Program</i>			
(Event Years 1 to 3)	1.585 [0.322]	0.231 [0.0615]	0.899 [0.207]
(Event Years 4 to 6)	2.089 [0.432]	0.219 [0.101]	1.191 [0.272]
(Event Years 7 to 10)	1.799 [0.578]	0.23 [0.146]	1.368 [0.438]
R-squared	0.041	0.04	0.036
Oblasts	69	69	69
Observations (individuals)	4,457	4,457	3,327

Notes: This table shows the change in average age at birth, number of children before birth and interval from previous birth for women who had children 1 to 3, 4 to 6, and 7 to 10 years after the start of the program. Standard errors clustered at the oblast level are in brackets. These coefficients are directly comparable to the coefficients presented in figure 7. The unit of observation is at the woman-birth level. The grouped event years represent: dummies for 1 to 3, 4 to 6, and 7 to 10 years after maternity benefits started. The omitted category is 7 to 0 years before the program started and is normalized to zero. Regressions are weighted by the survey sampling weights. Source: Generations and Gender Survey, 2004.

**Table 6. Heterogeneous Fertility Rate Increases to Maternity Benefits across Regions**

	(1)	(2)	(3)
Dependent variable	A. General Fertility Rate		
<i>Before Program</i>			
(Event Years -5 to -1) * % Rural	0.026 [0.0203]	0.026 [0.0207]	0.023 [0.0250]
<i>After Program</i>			
(Event Years 1 to 3) * % Rural	0.116 [0.0406]	0.116 [0.0413]	0.127 [0.0424]
(Event Years 4 to 6) * % Rural	0.129 [0.0673]	0.129 [0.0687]	0.148 [0.0713]
(Event Years 7 to 10) * % Rural	0.140 [0.0899]	0.140 [0.0921]	0.157 [0.0885]
R-squared	0.498	0.913	0.914
Dependent variable	B. General Fertility Rate		
<i>Before Program</i>			
(Event Years -5 to -1) * % Less than HS	0.006 [0.0150]	0.007 [0.0159]	0.007 [0.0158]
<i>After Program</i>			
(Event Years 1 to 3) * % Less than HS	0.150 [0.0298]	0.149 [0.0287]	0.149 [0.0295]
(Event Years 4 to 6) * % Less than HS	0.183 [0.0515]	0.181 [0.0498]	0.181 [0.0507]
(Event Years 7 to 10) * % Less than HS	0.183 [0.0664]	0.182 [0.0640]	0.179 [0.0646]
R-squared	0.285	0.915	0.915
Covariates	Year FE	Year FE, Oblast FE	Year FE, Oblast FE, X <sub>o,y</sub>
Oblast-year cells	1,444	1,444	1,444
Oblasts	82	82	82

Notes: These coefficients present interactions of grouped event-year dummies with a continuous oblast level characteristic measured in 1979 (share of women age 15 to 44 living in rural areas {panel A}; share of individuals with less than high school education {panel B}). See the description of event-year dummies in table 4. Standard errors clustered at the oblast level are in brackets. These coefficients are directly comparable to the coefficients presented in figure 7. Weights are the number of women who are ages 15 to 44 living in an oblast in 1980. Sources: see sources in figure 7.

**Table 7. Similarity of Russia in 1980 to Other Countries Today**

	1980	2010-2013				
	<b>Russia</b>	<b>USA</b>	<b>Austria</b>	<b>Italy</b>	<b>Spain</b>	<b>Sweden</b>
TFR	1.9	1.9	1.4	1.4	1.3	1.9
Childless Women	9.5	14	21	24	13	13
In Formal Care (age <3)	35	43.2	13.9	24.2	39.3	46.7
In Formal Care (age 3-6)	65	66.5	75	90	91	92
Female LFP	85	67	70	54.4	70	79
Female/Male Wage Ratio	67	82.2	80.8	72.6	93.9	85.7
Gini	0.29	0.39	0.28	0.32	0.34	0.27

Sources: 1979 Russian Census, 1979 “Narodnoe Hozyaystvo” Yearbook, OECD Statistics.

## APPENDIX A

### Expected Effects of Maternity Benefits on Childbearing

#### Proof of Proposition 1

I derive the effect of family benefits on the number of children by maximizing utility,  $U(n, q, z)$ , subject to the lifetime budget constraint,  $\pi qn + [t(w_f - a) - b]n + \pi_z z = T(w_f + w_m)$ . The budget constraint is nonlinear in  $n$  and  $q$ , which leads to the ambiguous effect of the benefits on the number of children discussed in the paper. This problem is associated with the following first order conditions:  $\frac{U_n}{U_q} = \frac{\pi q + t(w_f - a) - b}{n\pi}$  (i),  $\frac{U_q}{U_z} = \frac{\pi n}{1}$  (ii),  $\frac{U_n}{U_z} = \frac{\pi q + t(w_f - a) - b}{1}$  (iii). Notably, the shadow price of quantity,  $p_n$ , depends on quality,  $q$ , while the shadow price of quality,  $p_q$ , depends on quantity,  $n$ .

I take the total derivative of the budget constraint and the first order conditions. I impose standard assumptions on the utility function  $U_n > 0, U_q > 0, U_z > 0, U_{nn} < 0, U_{qq} < 0, U_{zz} < 0$ , and I assume that the utility function is additively separable such that  $U_{nz} = U_{zn} = U_{nq} = U_{qn} = U_{qz} = U_{zq} = 0$ . The consumption good,  $z$ , serves as the numéraire and its price is set to one. The total derivatives of the first order condition (i) with respect to  $a$  is  $\pi n \left(\frac{dq}{da}\right) + \pi q \left(\frac{dn}{da}\right) + \left(\frac{dn}{da}\right)(t(w_f - a) - b) - nt + \left(\frac{dz}{da}\right) = 0$ , of condition (ii) is  $U_{nn}\pi n \left(\frac{dn}{da}\right) + U_n \pi \left(\frac{dn}{da}\right) = U_{qq} \left(\frac{dq}{da}\right)(\pi q + t(w_f - a) - b) + U_q \left(\pi \left(\frac{dq}{da}\right) - t\right)$ , and of condition (iii) is  $U_{qq} \left(\frac{dq}{da}\right) = U_{zz} \left(\frac{dz}{da}\right)\pi n + U_z \pi \left(\frac{dn}{da}\right)$ ,  $U_{nn} \left(\frac{dn}{da}\right) = U_{zz} \left(\frac{dz}{da}\right)(\pi q + t(w_f - a) - b) + U_z \left(\pi \left(\frac{dq}{da}\right) - t\right)$ .

After some algebraic manipulation of the total derivatives, I derive the following effect of parental leave on childbearing,

$$\frac{dn}{da} = \frac{\pi n^2 t U_{zz} (U_{qq} (\pi q + t(w_f - a) - b) + U_q \pi) - U_q t (U_{qq} + \pi n U_{zz})}{(U_{nn} \pi n + U_n \pi) (U_{qq} + \pi n U_{zz}) - (U_{zz} (-\pi^2 q n - (t(w_f - a) - b) n \pi + U_z \pi) (U_{zz} (\pi q + t(w_f - a) - b) + U_q \pi))}$$

whose sign is ambiguous. I decompose the above formula,

$$\frac{d \ln(n)}{d \ln(a)} = \frac{atn}{I} \left( \underbrace{\frac{d \ln(n)}{d \ln(I)} \Big|_{\pi s \text{ constant}}}_{\text{Income Elasticity}} \right) + \underbrace{\frac{d \ln(n)}{d \ln(a)} \Big|_{I \text{ constant}}}_{\text{Substitution Elasticity}}$$

where the elasticity of childbearing with respect to paid leave is the sum of income and substitution elasticities. This general formula does not make clear what generates the ambiguity of the sign of the overall effect.

To obtain more intuitive expressions for the income and substitution elasticities, I solve the maximization problem by replacing the non-linear budget constraint with a linear one by making it a function of shadow prices:  $p_n n + p_q q + p_z z = I - \pi q n = R$ . The shadow price of  $n$  is  $p_n = \pi q + t(w_f - a) - b$ , the shadow price of  $q$  is  $p_q = \pi$  and of  $z$  is  $p_z = \pi_z$ , which means that:

$$d\ln(p_n) = \left(\frac{1}{k_n}\right) \left[ (d\ln(\pi) + d\ln(q))k + d\ln(t) \left(\frac{(w_f - a)tn}{R}\right) + d\ln(w_f) \frac{tnw_f}{R} - d\ln(a) \frac{ant}{R} - d\ln(b) \frac{bn}{R} \right] \quad \text{and}$$

$$d\ln(p_q) = \left(\frac{k}{k_q}\right) (d\ln(\pi) + d\ln(n)). \text{ In these equations, } k = \frac{nq\pi}{R}; k_n = \frac{p_n n}{R}; k_q = \frac{pqq}{R}.$$

To derive the income and substitution elasticities I use the following propositions:

$$d\ln(n) = \eta_n d\ln\left(\frac{R}{p}\right) + k_z \sigma_{nz} d\ln(p_z) - (1 - k_n) \bar{\sigma}_n d\ln(p_n) + k_t \sigma_{nt} d\ln(p_t)$$

$$d\ln(t) = \eta_t d\ln\left(\frac{R}{p}\right) + k_z \sigma_{tz} d\ln(p_z) + k_n \sigma_{nt} d\ln(p_n) - (1 - k_t) \bar{\sigma}_t d\ln(p_t)$$

where the  $\sigma$ 's are the Allen partial elasticities of substitution in the utility function, and  $\bar{\sigma}_n$  is the average elasticity of substitution of  $n$  against  $z$  and  $q$ , and  $\bar{\sigma}_t$  is the same elasticity for  $t$  against  $z$  and  $n$ . The average elasticities are always positive. Also define:  $p = k_z d\ln(p_z) + k_n d\ln(p_n) + k_t d\ln(p_t)$ .

To derive the income elasticity, let income,  $I$ , change while holding the  $\pi$ 's constant, which results in:  $d\ln(p_n) = \left(\frac{k}{k_n}\right) d\ln(q)$ , and  $d\ln(p_q) = \left(\frac{k}{k_q}\right) d\ln(n)$ , and  $d\ln\left(\frac{R}{p}\right) = (1 - k) d\ln(I)$ . Plugging these into the above propositions, I obtain the following expression:

$$\underbrace{\frac{D}{1 - k} \left( \frac{d\ln(n)}{d\ln(I)} \right)}_{\text{Income Elasticity}} \Big|_{\pi \text{ s constant}} = (1 - k \sigma_{nq}) \eta_n - \frac{k(1 - k_n) \bar{\sigma}_n}{k_n} \eta_q$$

where  $D = (1 - k \sigma_{nq})^2 - \frac{k^2(1 - k_n)(1 - k_q) \bar{\sigma}_n \bar{\sigma}_q}{k_n k_q}$  and  $(1 - k \sigma_{nq})$  must be positive by the second order conditions. In the equation,  $\eta_q$  is the true elasticity of quality with respect to income, where it uses a measure of income that is calculated using shadow prices (marginal costs) whose ratios in equilibrium are equal to the marginal rates of substitution in the utility function, while  $\eta_n$  is the true elasticity of quantity with respect to income. Thus, if  $\eta_q > \eta_n$  by a sufficiently large enough magnitude, then the income elasticity could be negative. The sufficient condition for the income effect to be positive is if:  $(1 - k \sigma_{nq}) \eta_n > \frac{k(1 - k_n) \bar{\sigma}_n}{k_n} \eta_q$ .

In a similar manner I also derive the elasticity of substitution:

$$D \frac{d\ln(n)}{d\ln(a)} \Big|_{l \text{ const}} = \frac{(1 - k_n) \bar{\sigma}_n ant}{k_n R}$$

and the sign is unambiguously positive.

### Proof of Corollary 1

Following the same procedure as in the proof of proposition 1, I derive the following effect of the cash transfer,  $b$ , on childbearing:

$$\frac{dn}{db} = \frac{\pi n^2 U_{zz} (U_{qq} (\pi q + t(w_f - a) - b) + U_q \pi) - U_q (U_{qq} + \pi n U_{zz})}{(U_{nn} \pi n + U_n \pi) (U_{qq} + \pi n U_{zz}) - (U_{zz} (-\pi^2 q n - (t(w_f - a) - b) n \pi + U_z \pi) (U_{zz} (\pi q + t(w_f - a) - b) + U_q \pi))}$$

where the sign of the expression is ambiguous. The elasticity of childbearing with respect to cash transfers is:

$$\frac{d \ln(n)}{d \ln(b)} = \frac{bn}{I} \left( \underbrace{\frac{d \ln(n)}{d \ln(I)} \Big|_{\pi's \text{ constant}}}_{\text{Income Effect}} \right) + \underbrace{\frac{d \ln(n)}{d \ln(b)} \Big|_{I \text{ constant}}}_{\text{Substitution Effect}}$$

where it is the sum of income and substitution elasticities. This general formula does not make clear what generates the ambiguity of the sign of the overall effect.

To obtain more intuitive expressions for the income and substitution elasticities, I follow the procedure outlined in proposition 1. The expression for the income elasticity is the same as in proposition 1, whose sign is positive under the same conditions. The elasticity of substitution can be expressed as:  $\frac{d \ln(n)}{d \ln(b)} = \frac{(1-k_n) \bar{\sigma}_n b n}{k_n R}$ , and is always positive.

### Proof of Proposition 2

I derive the effect of family benefits on the number of children by maximizing,  $U(n, t, z)$ , subject to the lifetime budget constraint,  $[t(w_f - a) - b]n + \pi_z z = T(w_f + w_m)$ . The budget constraint is nonlinear in  $n$  and  $t$ , which leads to the ambiguous effect of the benefits on the number of children discussed in the paper. This problem is associated with the following first order conditions:  $\frac{U_n}{U_z} = \frac{t(w_f - a) - b}{1} = \frac{p_n}{p_z}$ ,  $\frac{U_n}{U_t} = \frac{t(w_f - a) - b}{(w_f - a)n} = \frac{p_n}{p_t}$ ,  $\frac{U_t}{U_z} = (w_f - a)n = \frac{p_t}{p_z}$ . Note that the shadow price of quantity,  $p_n$ , depends on time off from work,  $t$ , while the shadow price of time off from work,  $p_t$ , depends on quantity,  $n$ .

I take the total derivative of the budget constraint and the first order conditions. I impose standard assumptions on the utility function  $U_n > 0, U_t > 0, U_z > 0, U_{nn} < 0, U_{tt} < 0, U_{zz} < 0$ , and I assume that the utility function is additively separable such that  $U_{nz} = U_{zn} = U_{nt} = U_{tn} = U_{tz} = U_{zt} = 0$ . The consumption good  $z$  serves as the numeraire and its price is set to one. After substituting the total derivative formulas into each other I obtain the following expression,

$$\frac{dn}{da} = \frac{(-U_{zz} n (w_f - a) \pi_n + U_z (w_f - a)) n (U_{zz} n t (w_f - a) - U_z) + (t (U_{zz} \pi_n n - U_z)) (U_{tt} + U_{zz} n^2 (w_f - a)^2)}{(U_{nn} + U_{zz} \pi_n^2) (U_{tt} + U_{zz} n^2 (w_f - a)^2) - (-U_{zz} n (w_f - a) \pi_n + U_z (w_f - a)) (-U_{zz} \pi_n (w_f - a) n + U_z (w_f - a))}$$

where  $\pi_n = t(w_f - a) - b$ . The sign of this effect is ambiguous. I derive the elasticity of childbearing with respect to paid leave as:

$$\frac{d \ln(n)}{d \ln(a)} = \frac{ant}{I} \left( \underbrace{\frac{d \ln(n)}{d \ln(I)} \Big|_{\pi's \text{ const}}}_{\text{Income}} \right) + \underbrace{\frac{d \ln(n)}{d \ln(a)} \Big|_{I \text{ const}}}_{\text{Substitution}}$$

where it is a combination of income and substitution elasticities. This general formula does not make clear what generates the ambiguity of the sign.

To derive more intuitive expressions of income and substitution effects, I linearize the budget constraint, such that it equals to:  $p_n n + p_t t + p_z z = R = I + tn(w_f - a)$ , where  $p_n = t(w_f - a) - b$  and  $p_t = (w_f - a)n$ . Thus,  $dln(p_t) = dln(w_f) \left(\frac{w_f n t}{R}\right) \left(\frac{1}{k_t}\right) + dln(a) \left(\frac{a n t}{R}\right) \left(\frac{1}{k_t}\right) + dln(n)$ , and  $dln(p_n) = \left(\frac{1}{k_n}\right) [-dln(b) \left(\frac{b n}{R}\right) + dln(t) k_t + dln(w_f) \left(\frac{t w_f n}{R}\right) - dln(a) \left(\frac{a n t}{R}\right)]$ , where  $k = \frac{tn(w_f - a)}{R}$ ,  $k_n = \frac{p_n n}{n}$  and  $k_t = \frac{p_t t}{R} = k$ .

To calculate the income elasticity, hold prices constant which leads to:  $dln(p_t) = dln(n)$ ;  $dln(p_n) = dln(t) \left(\frac{k_t}{k_n}\right)$ ;  $dln\left(\frac{R}{p}\right) = (1 - k)dln(I)$ . This leads to the following expression:

$$\frac{dln(n)}{dln(I)} \Big|_{\pi s \text{ const}} = \frac{1}{1 - k_t \sigma_{nt}} (\eta_n (1 - k_t \sigma_{nt}) - \eta_t \frac{k_t^2 (1 - k_n) (1 - k_t) \bar{\sigma}_n}{k_n k_t})$$

where the income elasticity is positive if,  $\eta_n (1 - k_t \sigma_{nt}) > \eta_t \frac{k_t^2 (1 - k_n) (1 - k_t) \bar{\sigma}_n}{k_n k_t}$ .

In a similar manner obtain the elasticity of substitution:

$$D \frac{dln(n)}{dln(a)} \Big|_{I \text{ const}} = \frac{k_a ((1 - k_n) \bar{\sigma}_n + \sigma_{nt} k_n) (1 - \sigma_{nt} k_t) + k_a (\bar{\sigma}_n + \frac{(1 - k_t) \bar{\sigma}_q}{k_t})}{(1 - k_n) \bar{\sigma}_n k_t k_n}$$

where  $D = (1 - k_t \sigma_{nt})^2 - \frac{(1 - k_t) \bar{\sigma}_q (1 - k_n) \bar{\sigma}_n k_t^2}{k_n k_t}$  and  $(1 - k \sigma_{nt})$  are positive by the second order conditions. The sign of the substitution elasticity is positive if  $\sigma_{nt} > 0$ .

### Proof of Corollary 2

Following the same procedure as in the proof of proposition 1, I derive the following effect of paid leave on childbearing:

$$\frac{dn}{db} = \frac{(-U_{zz} n (w_f - a) \pi_n + U_z (w_f - a)) (n U_{zz} n (w_f - a)) + (U_{zz} \pi_n n - U_z) (U_{tt} + U_{zz} n^2 (w_f - a)^2)}{(U_{nn} + U_{zz} \pi_n^2) (U_{tt} + U_{zz} n^2 (w_f - a)^2) - (-U_{zz} n (w_f - a) \pi_n + U_z (w_f - a)) (-U_{zz} \pi_n (w_f - a) n + U_z (w_f - a))}$$

where the sign is ambiguous. The elasticity of childbearing with respect to cash transfers is:

$$\frac{dln(n)}{dln(b)} = \frac{bn}{I} \left( \underbrace{\frac{dln(n)}{dln(I)} \Big|_{\pi s \text{ const}}}_{\text{Income}} \right) + \left( \underbrace{\frac{dln(n)}{dln(a)} \Big|_{I \text{ const}}}_{\text{Substitution}} \right),$$

which is a combination of income and substitution effects. The formula for the income effect as well as the condition for it to be positive are the same as in proposition 2. The substitution elasticity is:  $D \left(\frac{b}{n}\right) \frac{dln(n)}{dln(b)} \Big|_{I \text{ const}} = \frac{(1 - k_n) \bar{\sigma}_n b n}{k_n R}$  and is always positive.

## APPENDIX B

### Estimation of Fertility Rates Using Census Data

To estimate fertility rates, it is important to have information on the number of women of childbearing age. Information on the age structure of the population by region is only published in the decennial censuses. The 1989 census data is the closest to the year of the policy's start, compared to 2002 and 2010 census data, and should be the least affected by misclassification error due to mortality and mobility. Only individuals who have not died or moved out of the country appear in the census, thus the number of women of childbearing age calculated using 2002 census data will underestimate the true number of such women. The 1989 census data provide counts of men and women in one year age groups by region of residence as of January, 1989. I estimate birth year of the woman using her *age* in 1989 as  $1989 - \text{age} - 1$ , because the census took place between January 12 and January 19 in 1989. This calculation will only understate the birth year of women born between January 1<sup>st</sup> and January 11<sup>th</sup>. I use these data to backward-estimate the number of women each year (from 1975 to 1989) who are of childbearing age – ages 15 to 44. For instance, the number of women who are age 15 to 44 in 1979 is the same as the number of women who are age 25 to 54 in 1989. For years 1990 to 1992, I use published statistics in Rosstat on the number of women by age in each oblast.

My main outcome of interest is the General Fertility Rate (GFR) which is the number of births per thousand women of childbearing age. I estimate GFR in year  $y$  and region  $o$  as the number of children born in year  $y$  (except those who were still born), in oblast  $o$  as recorded in Rosstat data per thousand women aged from 15 to 44 in year  $y$ , and living in oblast  $o$  in 1989 as recorded in the 1989 Census.

$$(a) GFR_{y,o}^{1989} = \frac{\text{Number of Births}_{y,o} * 1000}{\text{Number of Women Age 15 to 44}_{y,o} \text{ from 1989 Census}}$$

I estimate the GFR in month  $m$ , year  $y$  and region  $o$  as the number of children born in month  $m$ , year  $y$  and oblast  $o$  and present in Russia in the 2002 Census per thousand women aged from 15 to 44 in year  $y$ , and living in oblast  $o$  in 1989 as recorded in the 1989 Census.

$$(b) GFR_{m,y,o}^{2002} = \frac{\text{Number of Births}_{m,y,o} \text{ from 2002 Census} * 1000}{\text{Number of Women Age 15 to 44}_{y,o} \text{ from 1989 Census}}$$

In my analysis comparing the effect of the program on first and higher parity births I will construct fertility rates (FR) by parity. I estimate fertility rates for first births in month,  $m$ , year,  $y$ , and oblast,  $o$ , as the number mothers born in oblast  $o$  and present in Russia during the 2010 Census who report that their first child was born in month,  $m$ , and year,  $y$ , per thousand women aged from 15 to 44 in year,  $y$ , and living in oblast,  $o$ , in 1989 as recorded in the 1989 Census.

$$(c) FR(1^{st} \text{ Birth})_{m,y,o}^{2010} = \frac{\text{Number of First Births}_{m,y,o} \text{ from 2010 Census} * 1000}{\text{Number of Women Age 15 to 44}_{y,o} \text{ from 1989 Census}}$$

Second, I estimate fertility rates for all higher parity births in month,  $m$ , year,  $y$ , and oblast,  $o$ , as the total number of births in month,  $m$ , year,  $y$ , and oblast,  $o$ , minus the number of first births estimate used in (c) per thousand women aged from 15 to 44 in year  $y$ , and living in oblast  $o$  in 1989 as recorded in the 1989 Census.

$$(d) FR(Higher Parity Birth)_{m,y,o}^{2010} = \frac{(Number\ of\ Births - Number\ of\ First\ Births)_{m,y,o}\ from\ 2010\ Census * 1000}{Number\ of\ Women\ Age\ 15\ to\ 44_{y,o}\ from\ 1989\ Census}$$

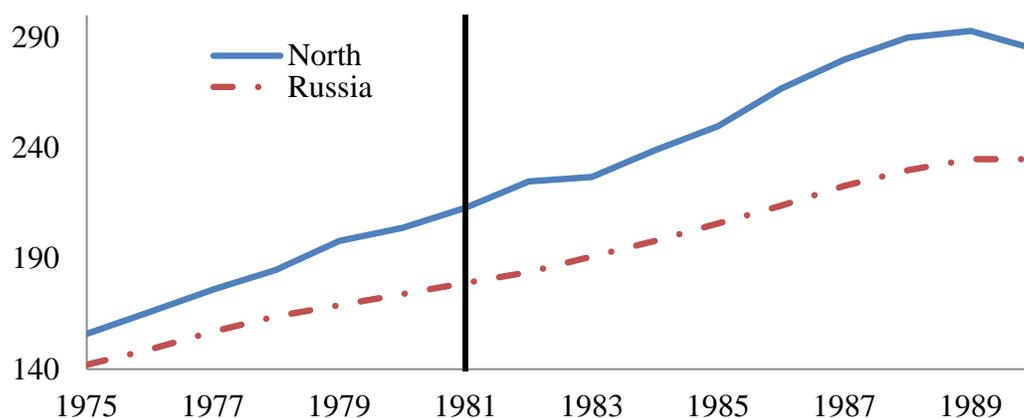
## APPENDIX C

### Data Description

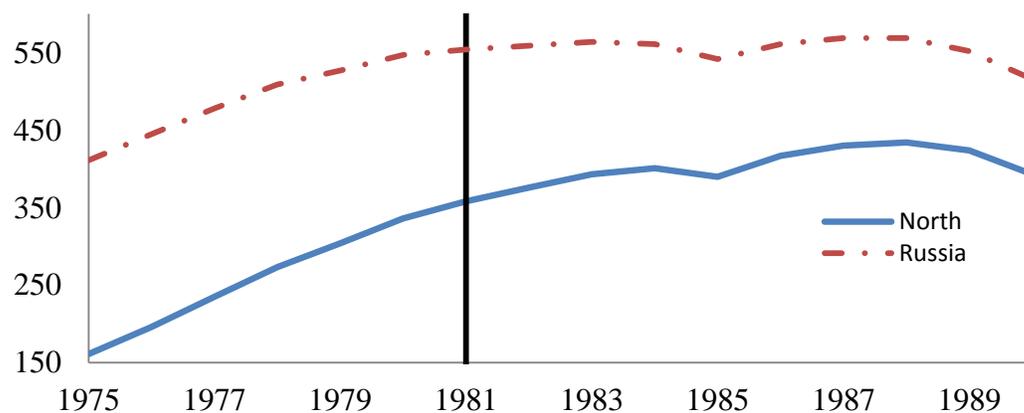
Data Source	Type	Description
1989 Russian Census	count data	counts of persons living in Russia in 1989 by age, sex, and oblast of residence
2002 Russian Census	count data	counts of persons living in Russia in 2002 by birth year, birth month and oblast at birth
2010 Russian Census	count data	counts of persons living in Russia in 2010 by birth year, birth month and oblast at birth; counts of women present in Russia in 2010 by birth year and birth month of their first child, and by oblast at birth
1979 Russian Census	count data	counts of persons living in Russia in January, 1979 by age, sex, oblast of residence, and by urban/rural status of the oblast; counts of persons living in Russia in January, 1979 who are older than 10 by education categories (only elementary, incomplete high school, complete high school, incomplete college, complete college)
2004 Generations and Gender Survey	micro-level data	sample of roughly 11,000 individuals aged 18 to 79 which is representative of Russia, where this analysis uses data on the birth year of the mother, and the birth year of her every child
1975-1992 "Narodnoe Hozyaystvo" Yearbooks	counts and averages data	characteristics of each oblast and in each year, where this analysis uses the production of bricks, concrete, timber, canned goods, meat, and the value of trade

## APPENDIX D

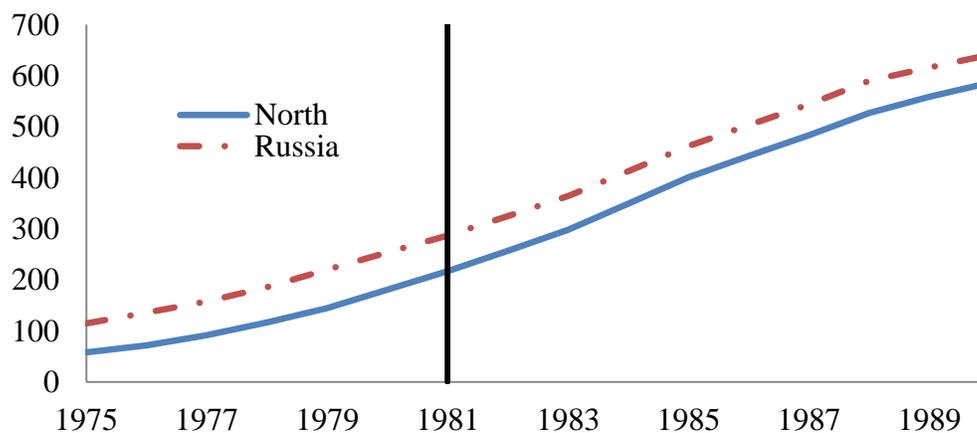
### Growth of Industrial Product (1970=100)



### Oil Production (million tons)



### Gas Production (milliard cubic meters)



Notes: The North is composed of 16 early beneficiary oblasts: Jakutsk ASSR, Murmansk, Kamchatsk, Magadan, Sahalin, Komi and Burjat ASSR, Krasnojarsk kraj, Primorsk, Habarovsk, Archangelsk, Tomsk, Tyumen, Irkutsk, Chitinsk, Amur oblasts. Sources: 1975-1990 "Narodnoe Hozyaystvo" yearbooks.

## APPENDIX E

### Creation of Childbearing Elasticities with Respect to Cost of a Child across Studies

I use a parametric bootstrap procedure to generate confidence intervals for the elasticity estimates in all studies (Johnston and DiNardo 1997). I generate 10,000 bootstrap draws of the reduced-form coefficients from normal distributions with means and standard errors equal to the point estimates reported in the paper. I calculate the percent change in childbearing,  $dln(n)$ , by dividing my bootstrap draws by the appropriate pre-treatment mean. I obtain an estimate of the cost of having a child until the child is 18 years old for a country and period relevant to the study, and generate 10,000 draws of average costs from normal distributions with means equal to the average cost estimate and standard errors equal to 10 percent of the cost for Russia, and 5 percent of the cost for all other studies. I calculate the percent change in costs,  $dln(c)$ , by dividing the change in monetary cost of a child induced by a policy by the bootstrap draws of average costs. Finally, I obtain 10,000 realizations of elasticities as:  $\frac{dln(n)}{dln(c)}$ . The values of the 2.5<sup>th</sup> and the 97.5<sup>th</sup> percentiles of the distribution of my generated elasticities constitute the 95-percent confidence interval for my estimated elasticity.

#### Russia

The percent change in the cost of a child is calculated as follows. The per-month cost estimate of a child is 90 rubles, which is provided by a Soviet demographer (Valentei 1987). I calculate the cost of a child over 18 years, which equals to  $90 \cdot 12 \cdot 18$ . The cost after the program subtracts parental leave payments (assuming full take-up),  $35 \cdot 10$ , and a cash transfer at the birth of a child, 75 (this is the average of the payment of 50 for a first birth and a payment of 100 for a second and third birth). Thus, the cost changed from  $90 \cdot 12 \cdot 18$  to  $90 \cdot 12 \cdot 18 - 35 \cdot 10 - 75$ , which represents a 2.2 percent decrease.

#### Austria

Guger (2003) uses the Austrian Consumer Survey from 2001 to estimate the monthly cost of a child in Austria as 500 euros, which is 389 euros in 1990 euros. The cost of raising a child before was:  $388 \cdot 12 \cdot 18$ , while after the expansion of paid leave the cost became:  $388 \cdot 12 \cdot 18 - 340 \cdot 12$  (the monthly paid leave was a fixed amount of 340 euros; I am assuming full-take up in this calculation). The cost of raising a child decreased by 4.8 percent. Childbearing went up by 21 percent (the coefficient scaled by the pre-treatment mean is,  $0.068/0.32$ , and the standard error of the coefficient equals to 0.12; these estimates come from table 7 in Lalive and Zweimüller (2009)). Thus, the elasticity equals to: -4.4.

#### Spain

The average total cost of a child is estimated to be 150,000 euros. This estimate comes from surveys conducted by the consumer organization CEACCU in 42 Spanish provinces (as of 2008). The cost of a child after the benefit became:  $150,000 - 2,500$ , where parents received a cash transfer of 2,500 euros at the birth of a child, which is a reduction in the cost of raising a child of

1.7 percent. The effect of the program on fertility rates is estimated at 0.063 percent (estimate in table A1 (Gonzalez 2012), with a standard error of 0.0115). This results in an elasticity of -3.8.

### Canada

I calculate the change in cost of a child by taking a weighted average of changes in costs for first, second and third and higher parity births based on numbers reported in the paper. The benefit was 500 dollars for the first birth, 1000 dollars for second births, and 8000 dollars for third and higher parity births. The annual costs of a child as reported in the paper were: 7,935, 6,348, and 5,324 for first, second, and third births respectively. I weight these costs of a child by the share of births in each category. The resulting change in costs equals to:  $(500/(7935*18))*0.45+(1000/(6348*18))*0.35+(8000/(5324*18))*0.2$ . Thus, costs decreased by 2.1 percent. The fertility rate increased by 0.087 percent (table 6 (Milligan 2005); model c). Thus, the elasticity equals to: -4.1.

### Israel

The elasticity, 0.54, and the standard error, 0.077, presented in this paper were derived in Cohen et al. (2013).