The Unintended Consequences of Relaxing Birth Quotas: Theory and Evidence*

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Abstract

This study examines the impact of easing birth quotas on fertility transition, considering both extensive and intensive margins. Employing an extended Barro-Becker model, we anticipate asymmetric effects on birth rate based on household preferences. Utilizing a distinctive two-child policy in China that allowed eligible couples to have a second child, combined with a difference-in-differences framework, our analysis reveals a notable increase in second-child births following the relaxation of birth quotas. However, this positive outcome is counterbalanced by a discernible reduction in first-child births. The observed decline or postponement in first childbearing is linked, in part, to the rising costs associated with child rearing, aligning with our theoretical predictions. Consequently, our findings advocate for a nuanced policy approach. Instead of universally relaxing or abolishing birth quotas, policymakers should prioritize initiatives that address the financial burdens of child rearing faced by prospective parents. This targeted strategy is posited to more effectively enhance overall fertility rate.

JEL classification: J13; J18; H23
Keywords: birth quotas; childbearing decisions; child-rearing costs; China

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1 Introduction

China is currently undergoing a profound demographic shift marked by a significant decline in fertility rate and an increasingly aging population. In response to these demographic challenges, Chinese authorities have implemented a substantial relaxation of birth quotas over the past decade. This policy adjustment has granted eligible couples, previously constrained by the longstanding one-child policy (OCP), the opportunity to have a second or even a third child.[1] During this period, the country witnessed a notable shift with a surge in second and higher-order births, a trend anticipated due to the policy change. However, unexpectedly, there was a somewhat simultaneous decrease in the overall birth rate, indicating an even faster decline in the first-child birth rate.[2] This paradox prompts a particular curiosity: does the relaxation of birth quotas exert asymmetric effects on fertility transitions along the extensive margin (the decision of whether to have a child at all) and intensive margin (the decision of whether to have a second child)? If so, what are the underlining mechanisms driving these shifts?

In this paper, we present a theoretical framework to explore the impact of relaxing birth quotas on households' fertility decisions. This framework extends the canonical Barro-Becker framework (Becker and Barro (1988); Barro and Becker (1989)) by incorporating birth quotas and the heterogeneous fertility preferences of households. Analytically, the relaxation of birth quotas influences the total birth rate through two distinct channels. The first channel, which we term the "limits-raising channel", embodies a natural and intuitive expectation. By easing birth quotas, households with the desire and financial capability to support more children, previously restricted, are now legally enabled to expand their families. This channel is associated with the intensive margin of fertility decision-making.

Furthermore, merely easing birth quotas can exacerbate the overall expenses of raising children. This is particularly evident when households, deeply inclined towards larger families and potentially less aware of the associated financial burdens, choose to further enlarge their families. With the expansion of these households, a crowding-out effect ensue, wherein the sudden surge in newborns amplified demands on nursing, caregiving, food, housing and future education, collectively escalating the overall cost of child rearing. This expectation, corroborated below, introduces additional stress for newlyweds preparing for their first child. Consider a

[1] In 2011, China began relaxing its OCP, allowing 'eligible couples' where both partners were only children to have a second child, and in 2014, it expanded the policy to 'eligible couples' with at least one partner being an only child, and in 2016, it expanded the policy to all married couples. In 2021, the introduction of the third-child policy permits all married couples to have a third child.

[2] Between 2011 and 2022, the proportion of second and higher-order births among annual births rose substantially from 33.7% to 53.9%, illustrating the impact of the policy shift. Paradoxically, despite this increase in multiple-order births, the overall birth rate declined from 14.6‰ to a record low of 6.8‰ (National Bureau of Statistics of China (2023)).

scenario where couples aged 20-30, preparing for their first child, suddenly contend with an influx of newborns from 30-40-year-old couples, now permitted to have an "additional" child. This phenomenon encapsulates the "child-rearing cost channel", associated with the extensive margin and applicable within a dynamic general equilibrium context, where individuals make decisions while considering the behavior of others. This channel exerts a downward pressure on households' birth rate, particularly for prospective parents with lower preferences, thereby counteracting the effects of the "limits-raising channel".

Few countries encounter a scenario where the cost of child rearing becomes such a pivotal factor in influencing decisions regarding childbearing, as is evident in China. The average costs of child rearing in China were among the highest globally, as reported by the YuWa Population Research Institute in 2022.[3] According to this Beijing-based think-tank,[4] the average cost of raising a child to the age of 17 in China in 2019 stood at 485,000 yuan ($76,629) for a first child, which was 6.9 times China's per capita GDP that year. Notably, child rearing costs were even higher in major cities, reaching more than 1 million yuan in Shanghai and 969,000 yuan in Beijing.[5] Among the 13 countries included in the study, China ranked as the second-highest, trailing only behind South Korea, which has the lowest birth rate in the world.[6]

The prominence of the "child-rearing cost channel" is further accentuated by the absence of a comprehensive support system, raising concerns about the unintended consequences of the recent policy shift. Despite the relaxation of birth quotas, the implemented policy lacks a corresponding set of measures aimed at mitigating the financial burden associated with raising children. Crucial components such as cash and tax subsidies, along with housing subsidies, are notably absent from this policy framework, which is essential for addressing the economic burdens of child rearing. Furthermore, the existing infrastructure to support families, including childcare facilities and access to professional nannies, remains woefully inadequate.[7]

[3]According to YuWa Population Research Institute (2022), the average cost of raising a child to the age of 17 in China consists of the costs during pregnancy (2.06%), expenses related to childbirth and postpartum confinement (3.09%), the costs of raising infants aged 0-2 (13.33%), the cost of raising children aged 3-5 (20.75%), the cost of raising children aged 6-14 (44.65%), and the cost of raising children 15-17 (16.12%).

[4]Findings of this report has been cited widely by international media such as CNN, Reuters, ABC and Financial Times.

[5]It echoes with a 2019 Shanghai Academy of Social Sciences report, which revealed that an average family living in Shanghai's upscale Jingan District spent almost 840,000 yuan per child from birth through junior high school, which typically ends at age 15, including 510,000 on education alone. Findings of this report has been cited by international media such as Reuters, VOA, and The Diplomat.

[6]The U.S. figure, based on 2015 data, stood at 4.11 times per capita GDP, while Japan's figure, based on 2010 data, stood at 4.26. Other included countries were Australia (2.08), Singapore (2.1), Sweden (2.91), Switzerland (3.51), Ireland (3.57), Germany (3.64), Canada (4.34), New Zealand (4.55), United Kingdom (5.25), Italy (6.28).

[7]The existing infrastructure supporting child rearing, encompassing prenatal tests and deliveries, postpartum centres, childcare facilities, education, and health care, face significant inadequacies. Notably, in 2020, there were over 47 million children under 3 years old in China, with merely 5% of them benefiting from day care services. Compounding the issue, less than 20% of these services have government-guided pricing. This accessibility crisis,

To estimate the causal effects of relaxing birth quotas,[8] we take advantage of a distinctive two-child policy adopted by local governments in China[9] in different months in 2014, combined with a difference-in-differences (DID) framework. To check the identifying assumption that the decision about whether and when to relax birth quotas is uncorrelated with any prior trends in the outcomes of interest, we test whether any treatment effects exist in the pretreatment periods based on an estimator proposed by de Chaisemartin and D'Haultfoeuille (2020) that works well in the presence of heterogeneous treatment effects (e.g. by groups or over time). To check the identifying assumption that the timing of relaxing birth quotas does not coincide with any province-specific shocks or policies that might influence the outcomes of interest, we conduct additional placebo analyses focusing on individuals less affected by the relaxation of birth quotas and analysing alternative censuses prior to the relaxation of birth quotas.

Such an empirical framework, however, poses challenges in measuring childbearing outcomes and the costs associated with child-rearing. As for the childbearing outcomes, most studies use low-frequency annual births, which is more likely to be affected by unobservable social and economic changes, to investigate the impact of (relaxing) birth quotas. In this study, we take advantage of China's 2015 population census to construct monthly panel data on births. Importantly, we assume a minimum amount of time, about one year, for the policy to take effect based on the birth permit process (about a month), time to conceive (one month) and pregnancy (10 months). For more details, see **Appendices C.1 to C.4**.

In addressing the challenges of obtaining high-quality and high-frequency panel data on child-rearing costs in China, this study adopts a multi-pronged approach. We construct three complementary measures to delineate changes in the average cost of child rearing across specific regions over time. The first measure entails the collection of overall and expenditure-based Consumer Price Indices (CPIs). These indices aim to capture both the general shifts in the cost of final goods and services and the specific alterations in the cost of items more pertinent to raising a child. For the second measure, we gather data on overall and square-meter-based residential real estate sale price indices. This enables us to gauge variations in housing costs and discern changes in housing expenses for different demographic groups. The third measure involves the

coupled with the absence of regulated pricing, renders in-demand private facilities financially inaccessible for many families (Xinhua (2020)).

[8]Our empirical exercise focuses on short-term effects of relaxing birth quotas using real-world policies and recent fertility data. An underlying assumption is that people incorporate these changes in price and interest rate into their fertility decisions. This assumption is intuitive because raising a child is costly for most Chinese families, and thus couples' fertility decisions would be sensitive to any changes in child-rearing costs.

[9]China's constitution provides for three de jure levels of government. Currently, however, there are five practical (de facto) levels: (1) provincial (including province, autonomous region, municipality, and special administrative region), (2) prefecture, (3) county, (4) township, and (5) village. For the sake of simplicity, we treat 'local' and 'provincial' as interchangeable.

collection of data on nanny wages at the individual level, offering insights into fluctuations in the overall cost of hiring local nannies. Unlike the continuous increase in demand for final goods and services throughout a child's development (e.g., housing, education, health care), the demand for a nanny is likely to peak during the child's initial year of life and subsequently decline. For a more in-depth exploration of these measures, please refer to **Section 3.3**.

Our main empirical findings are as follows. First, we find that relaxing birth quotas substantially increases second-child births. The impact is most pronounced for couples covered by pensions, which supports the intergenerational altruism motive rather than the old-age security motive for fertility. Second, it significantly reduces or delays first-child births. Further analyses suggest that higher child-rearing costs induced by the relaxation of birth quotas are a likely reason for this trend. Third, the declining first-child births do not offset the increased second-child births. Overall, these findings align with predictions from an extended Barro-Becker model, as described in **Section 2**.

This study makes several contributions to the literature. First, it advances our understanding of heterogeneous fertility transitions along the extensive and intensive margins (Doepke, Hannusch, Kindermann and Tertilt (2023)). While prior research has predominantly focused on changes in the number of children per family (intensive margin), this paper sheds light on the often-overlooked dimensions of fertility transitions along both margins, drawing insights from China's relaxation of the OCP. Noteworthy a few studies in this realm have primarily concentrated on developed countries, such as Gobbi (2013), Aaronson, Lange and Mazumder (2014), Baudin, De La Croix and Gobbi (2015) and Hwang (2023).

Second, this study contributes to the theoretical analysis of OCP's impact. The existing literature has primarily tackled empirical issues, with only a few studies delving into the theoretically implications (Liao (2013); Choukhmane, Coeurdacier and Jin (2023)). Liao (2013) is one of the few studies to theoretically investigate the general equilibrium effect of relaxing the OCP on childbearing decisions. Using a calibrated general-equilibrium model, the author reveals that skilled parents may choose to have fewer children when the OCP is lifted due to wage rate effects. This paper specifically focuses on the child-rearing cost effects within the general equilibrium framework, providing a nuanced understanding of the dynamics at play.

Third, this study adds to the extensive body of empirical studies investigating the economic impacts of China's OCP. Previous studies have examined the effect of China's OCP on various family and labor market outcomes, such as sex ratio (Ebenstein (2010); Li, Yi and Zhang (2011)), twins conceived through assisted reproductive technology (Huang, Lei and Zhao (2016)), labor supply (Wang, Zhao and Zhao (2017); Zhang (2017)), human capital (Rosenzweig and Zhang (2009); Qian (2009); Liu (2014)), marriage market (Huang, Pan and Zhou (2023)), income and

savings (Huang, Lei and Sun (2021)). However, the effects of China's OCP on housing prices, cost of living and real incomes remain largely underexplored (Zhang (2017)). Our study addresses a notable empirical research gap on the effect of China's OCP on child-rearing costs.

Finally, this study engaged with ongoing debates surrounding the fundamental motives that drive fertility decisions. Existing discussions present divergent perspectives on the primary motives behind fertility decisions. Whereas some studies suggest that the main motive for fertility is intergenerational altruism, others provide evidence that favours the old-age security motive (Nugent (1985); Becker and Barro (1988); Cigno (1992); Rendall and Bahchieva (1998); Rossi and Godard (2022)). Our study involves the identification of couples with and without pensions. Through this lens, we seek to discern which group is more likely to have a second child with the relaxation of China's OCP.

2 Relaxation of China's OCP and its Economic Effects

In this section, we briefly introduce China's OCP and its reforms. We also extend the Barro-Becker framework to incorporate birth quotas and heterogeneous fertility preferences, by household, into a dynamic general equilibrium model. This theoretical analysis of the economic effects of relaxing birth quotas serves as a guide to our empirical analysis.

2.1 China's OCP and its reforms

The mandatory OCP in China was formally introduced in 1979 and enforced by the National Family Planning Commission. Prior to the OCP, authorities adopted a mostly voluntary 'later-longer-fewer' policy (Chen and Fang (2021)). After the OCP, most couples were permitted to have one child, and any violations incurred a fine and other punishments, including seizure of property and forced dismissal from government employment (Ebenstein (2010)).[10] Moreover, children without a proper birth permit were denied access to hukou, which determines a Chinese citizen's right to vote and stand for election, right to work, right to education, and right to social security. The OCP was implemented more strictly in urban areas than rural areas and beginning in 1984, couples in many rural areas were permitted to have a second child if their first child was female (i.e. the so-called 1.5-child policy). See Zhang (2017) for a review of the OCP.

In the early 2000s, the National Family Planning Commission began to relax the OCP to address declining fertility rates. In 2011, all couples in which both partners were single offspring

[10] Although the OCP was a de jure birth-quota system, it functioned as a de facto pricing system to some extent (Scharping (2003)). To simplify the (theoretical) analysis, this article defines the OCP as a birth-quota system that restricts the number of children each Chinese couple can have. That said, using an alternative definition of the OCP (i.e. a pricing system) is not likely to affect our main findings.

were allowed to have a second child, and in November 2013, this policy expanded to include couples with only one such partner (Ouyang (2013)). Eligible couples still had to apply for a birth permit in advance, and couples who had a second child without obtaining birth permits were fined 500 to 2,000 RMB and subject to other punishments, such as denial of reimbursement of childbirth-related medical costs and subsidies. In December 2015, the universal two-child policy was announced, allowing all couples to have a second child (Li, Xue, Hellerstein, Cai, Gao, Zhang, Qiao, Blustein and Liu (2019)) and importantly, dropping the requirement for a birth permit. In this article, we focus on the effects of the partial two-child policy owing to data limitations. Given that the universal two-child policy is essentially an extension of the partial two-child policy, analyzing its effects can help identify the effects of the universal two-child policy by providing a lower bound of estimates.

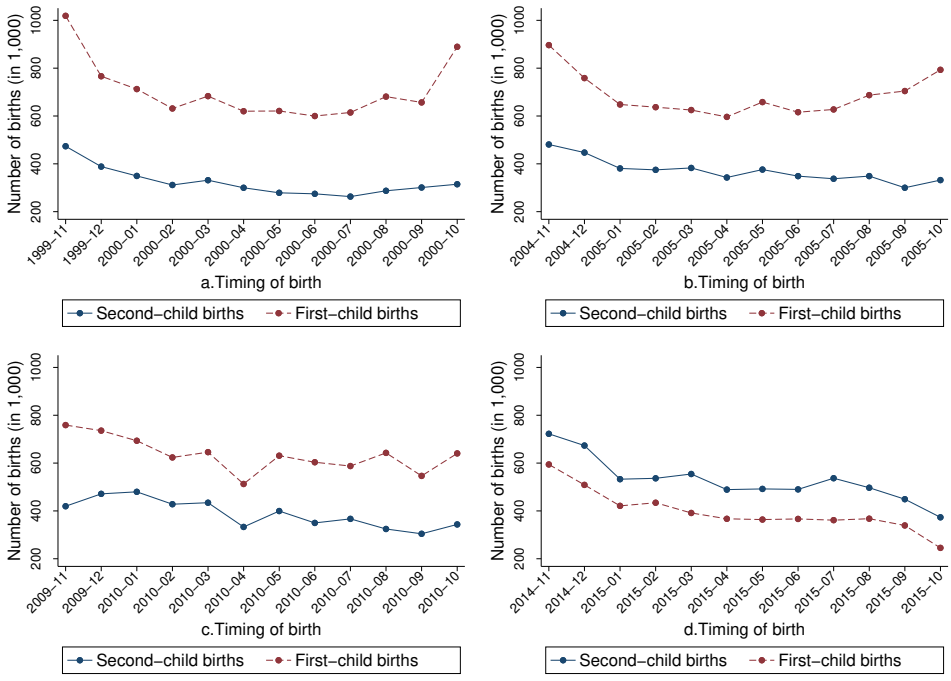


Figure 1: Number of first-child and second-child births over time (2000 – 2015).

Notes: This graph shows the number of monthly first-child and second-child births (a) from November 1999 to October 2000, (b) from November 2004 to October 2005, (c) from November 2009 to October 2010, and (d) from November 2014 to October 2015. Data on monthly first-child and second-child births in each panel are constructed by the authors using China's Population Censuses in 2000, 2005, 2010, and 2015, respectively.

Figure 1 depicts the monthly counts of first-child and second-child births spanning from 2000 to 2015. We observed a gradual decline in first-child births during this period, while second-child

births remained relatively steady until 2010. By 2015, there was a notable uptick in second-child births, yet first-child births saw a significant decline. **Figure 2** illustrates the proportion of second and higher-order births alongside the overall birth rate from 2011 to 2020. Intriguingly, even as the proportion of second and higher-order births rose following the relaxation of the OCP, the overall birth rate experienced a decline. Collectively, these findings indicate distinct fertility patterns in response to changes in birth quotas, both in terms of the extensive margin (the decision of whether to have a child at all) and intensive margin (the decision of whether to have a second child).

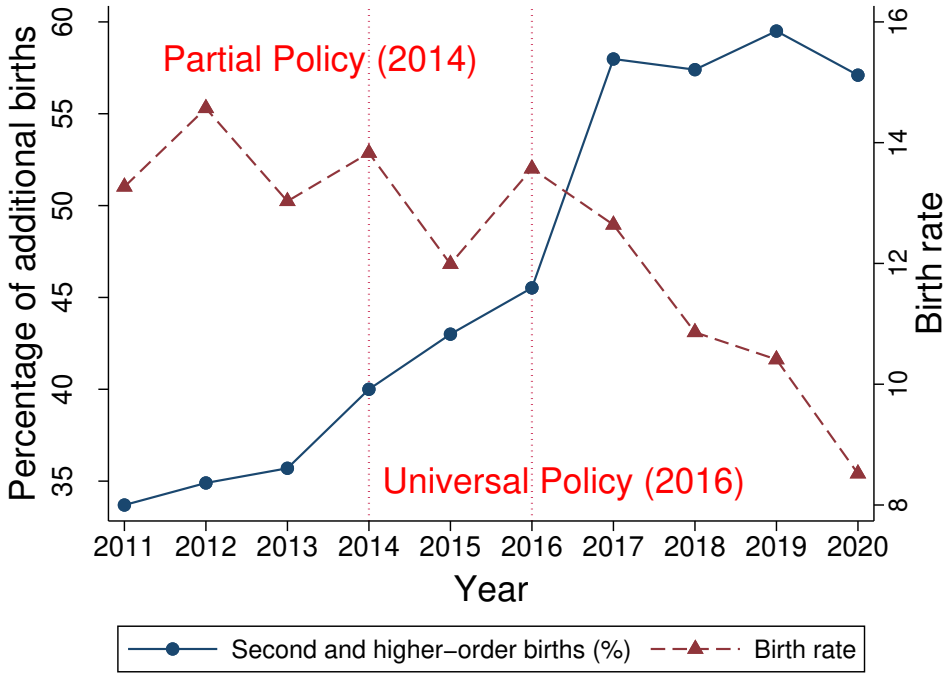


Figure 2: Percentage of second and higher-order births and birth rate (2011-2020).

Notes: The solid line denotes the trends in the percentages of second and higher-order births from 2011 to 2019. The dashed line denotes the trends of birth rates from 2011 to 2020. Partial Policy (2014) refers to the partial two-child policy that was announced at the national level in November 2013 and was then implemented by local governments at different time points in 2014. Universal Policy (2016) refers to the universal two-child policy that was announced at the national level in December 2015 and was then implemented by local governments at different points in time in 2016. Annual data on the percentage of second and higher-order births and birth rate are reported by the National Bureau of Statistics of China.

2.2 Economic effects of relaxing birth quotas

2.2.1 Households

The economy consists of two types of households who differ in their fertility preferences, with α^h and $\alpha^l < \alpha^h$ denoting the households with high and low fertility preferences, respectively. The respective population sizes of households α^h and α^l are given by q and $(1-q)$, where $q \in (0,1)$. To keep things as simple as possible, we assume that members in each household have the same fertility preference, and the preferences of descendants within the same household remain unchanged over time.

Each person lives for two periods, childhood and adulthood. In every period, each adult chooses the desired number of children and raises them in that period.[11] Specifically, at the beginning of $t = 0$, each adult in household α^h chooses a number $n(\alpha^h, 0)$ of children and raises them in that period. In generation $t = 1$, the $q \cdot n(\alpha^h, 0)$ children grow up to be adults, and each adult then decides to have and raise $n(\alpha^h, 1)$ children. Therefore, the number of adults and children in household α^h in generation $t = 1$ are $q \cdot n(\alpha,^h 0)$ and $q \cdot n(\alpha^h, 0) \cdot n(\alpha^h, 1)$, respectively, implying that the birth rate of household α^h is $n(\alpha^h, 1)$. By the same token, in generation t, the total numbers of adults and children in household α^h are $q \cdot n(\alpha^h, 0) \cdot n(\alpha^h, 1) \cdots n(\alpha^h, t-1) = q \cdot \prod_{t=0}^{t-1} n(\alpha^h, t) = qN(\alpha^h, t)$ and $qN(\alpha^h, t+1)$, respectively. Accordingly, the birth rate of household α^h is $n(\alpha^h, t)$. Similarly, the total numbers of adults and children, and the corresponding birth rate in household α^l in generation t are $(1-q) \cdot \prod_{t=0}^{t-1} n(\alpha^l, t) = (1-q)N(\alpha^l, t)$, $(1-q)N(\alpha^l, t+1)$, and $n(\alpha^l, t)$, respectively.

Following Becker and Barro (1988) and Barro and Becker (1989), we assume that the utility function of each adult in generation t in household $\alpha = \{\alpha^h, \alpha^l\}$ is

$$U(\alpha, t) = [c(\alpha, t)]^{\sigma}/\sigma + a(n(\alpha, t)) \cdot n(\alpha, t) \cdot U(\alpha, t+1), \quad \sigma < 1 \tag{1}$$

where $1/(1-\sigma)$ is the elasticity of intertemporal substitution in consumption, $c(\alpha, t)$ is the adult's consumption of final good, and $U(\alpha, t+1)$ is the utility attained by each child. Parents with different fertility preferences have unequal degrees of altruism; the higher the degree of altruism, the more children that an adult is likely to have.[12] The function $a(n(\alpha, t))$ measures the degree of altruism toward each child and is specified as

$$a(n(\alpha, t)) = \alpha \cdot [n(\alpha, t)]^{-\epsilon}, \quad 0 < \epsilon < 1. \tag{2}$$

Equation (2) suggests that given the utility function of each child $U(\alpha, t+1)$, the degree of altruism decreases with the number of children.

[11] Our model considers neither gender difference nor the fertility process for simplicity.

[12] Fertility preferences can be affected by factors such as individual characteristics and culture.

The labor market is competitive. Each adult owns one unit of time in each period and earns the same wage rates $w(t)$. Parents leave a bequest of capital $k(\alpha, t+1)$ in generation $t+1$ to each child born in generation t. Capital $k(\alpha, t)$ earns rent at the rate $r(t)$ in generation t. An adult of household α in generation t spends their earnings and inheritance $w(t) + [1 + r(t)]k(\alpha, t)$ on own consumption $P(t)c(\alpha, t)$, on bequests to children $n(\alpha, t)k(\alpha, t+1)$, and on child-rearing costs, where $P(t)$ are the prices of final good. Raising a child costs $\beta(\alpha, t)$, so $n(\alpha, t)\beta(\alpha, t)$ is total child-rearing costs. Therefore, the budget constraint for an adult in household α is $w(t) + [1 + r(t)]k(\alpha, t) = P(t)c(\alpha, t) + n(\alpha, t)[\beta(\alpha, t) + k(\alpha, t+1)]$.

Solving the optimization problem yields the equilibrium conditions summarized in the following lemma.

Lemma 1. *The optimal consumption and birth rate are given by, respectively,*

$$c(\alpha, t) = \left(\frac{\sigma}{1 - \epsilon - \sigma}\right) \frac{\beta(\alpha, t-1)[1 + r(t)] - w(t)}{P(t)}, \quad (t = 1, 2, \cdots) \tag{3}$$

$$[n(\alpha, t)]^{\epsilon} = \alpha[1 + r(t+1)] \left[\frac{P(t)}{P(t+1)}\right]^{\sigma} \left\{\frac{\beta(\alpha, t-1)[1 + r(t)] - w(t)}{\beta(\alpha, t)[1 + r(t+1)] - w(t+1)}\right\}^{1-\sigma}, \quad (t = 1, 2, \cdots) \tag{4}$$

Proof. See the Appendix A.1. □

Equation (4) reveals that the optimal number of children $n(\alpha, t)$, for all $t \geq 1$, born for each adult of household α depends on the degree of altruism, prices of final good, interest rates, wage rates, and child-rearing costs at different periods $\beta(\alpha, t-1)$ and $\beta(\alpha, t)$. The initial birth rate $n(\alpha, 0)$ is determined in (A.6). We further assume that child-rearing costs are identical across households in the same period,[13] and only final good is required such that[14]

$$\beta(\alpha, t) = \beta(t) = \mu P(t), \quad \mu > 0. \tag{5}$$

Given this assumption, Equation (4) then shows that the difference in birth rates is driven by the unequal fertility preferences in that the macroeconomic factors (i.e., $P(t)$, $r(t)$ and $w(t)$) affect households symmetrically. Therefore, the relation $n(\alpha^h, t) > n(\alpha^l, t)$ must hold for all $t \geq 1$ because of $\alpha^h > \alpha^l$.

[13] Assuming that families with different fertility preferences are related to unequal child-rearing costs largely complicates the model analysis. More importantly, we do not see direct evidence linking child-rearing costs with families' fertility preferences.

[14] First, the assumption that raising a child does not require an adult to spend time helps ensure the uniqueness of the steady-state equilibrium. See **?** for a detailed discussion. More importantly, this assumption is enough to capture the impact of relaxing fertility restrictions on child-rearing costs.

2.2.2 Production

The economy produces a unique final good and the production sector is competitive. A representative firm produces the unique final good using labor $L(t)$ and capital $K(t)$ according to the Cobb-Douglas technology $Y(t) = [L(t)]^{\eta}[K(t)]^{1-\eta}$. Solving the cost-minimization problem gives rise to the demand functions for labor and capital, $L(t) = \eta P(t)Y(t)/w(t)$ and $K(t) = (1-\eta)P(t)Y(t)/r(t)$. So far, we have not chosen the numeraire in the economy. Henceforth, we take wage rates as the numeraire so that $w(t) \equiv 1$.[15] The prices of final good and interest rates can thus be expressed as

$$P(t) = 1/\left\{\eta[k(t)]^{1-\eta}\right\}, \quad \text{and} \quad r(t) = (1-\eta)/\left[\eta k(t)\right], \tag{6}$$

where $k(t) \equiv K(t)/L(t)$ is the capital stock per capita/adult.

2.2.3 The Government

The government adopts different fertility restrictions, or birth quotas, to affect population growth. Suppose that the number of children that each household has in generation t satisfies the following condition $n(\alpha, t) \in [\underline{n}, \overline{n}]$, where $\underline{n}$ and $\overline{n}$ are the minimum and maximum number of children. In our model economy, the government strictly controls the maximum number of children that each household is permitted to have, namely $\underline{n} < \tilde{n} < \overline{n}$. Therefore, the number of children $n(\alpha, t)$ in generation t satisfies the following condition

$$n(\alpha, t) = \begin{cases} \tilde{n} & \text{if } n(\alpha, t) \geq \tilde{n} \\ n(\alpha, t) & \text{if } n(\alpha, t) < \tilde{n}. \end{cases} \tag{7}$$

2.2.4 Aggregation

The total adults at the beginning of $t = 0$ is defined as $\hat{N}(0) = q + (1-q) = 1$, and the total children in generation 0 or the total adults in generation 1 is $\hat{N}(1) = qn(\alpha^h, 0) + (1-q)n(\alpha^l, 0)$. Suppose that the government initiates a population policy at the beginning of generation $t = 1$ that restricts each household to have at most $\tilde{n} = n(\alpha^l, 1) > 0$ children, the number of children that each agent in household α^l is willing to have. It means that the restricted fertility constraint is not binding on the α^l-type household, but it is binding on the α^h-type household for $n(\alpha^h, 1) >$

[15] Selecting wage rates as the numeraire brings some convenience. In the current setting, changes in fertility restrictions can affect the prices of final good and thus child-rearing costs. If, instead, choosing the prices of final good as the numeraire, we need to assume that raising a child involves time devoted by adults in order to capture the child-rearing costs channel. The modification complicates our analysis.

$n(\alpha^l, 1)$ in a restriction-free environment. Given this fertility policy, the total number of adults at period of $t = 2$ and the total birth rate at period $t = 1$ are $\hat{N}(2) = q \cdot n(\alpha^h, 0) \cdot \tilde{n} + (1 - q) \cdot n(\alpha^l, 0) \cdot \tilde{n}$ and $B(t = 1) = \hat{N}(2)/\hat{N}(1) = \tilde{n}$. Holding constant the population policy, all members in households α^h and α^l share the same birth rate since $t = 1$, and the birth rates of both types of households no long change. Accordingly, the total birth rate remains unchanged over time such that

$$B(t > 1) = B(t = 1) = n(\alpha^h, t > 1) = n(\alpha^l, t > 1) = \tilde{n}. \tag{8}$$

Given this constant population growth rate, the economy is then on its balanced growth path, in which the growth rate of output per adult is zero.[16]

2.2.5 Short-term effects of relaxing birth quotas

In this subsection, we explore the short-term effects of relaxing birth quotas and leave the discussion on the long-term effects in the Appendix A.3. Suppose that at the very beginning of generation $t = s > 1$, the government eases the restricted population policy from $\tilde{n}$ to $n^* \in (\tilde{n}, n^h)$, where n^h is the birth rate of household α^h in the equilibrium without population control.[17] Then, upon the arrival of a new population policy, agents in both types of households may alter their fertility decisions in response to the potential change of macroeconomic factors.

We first analyze how the relaxation of birth quotas affects the α^l-type household's fertility decision. Using (5) and the fact the wage rate is normalized to one, we can reduce (4) to

$$(\tilde{n})^\epsilon = \alpha^l \cdot [1 + r(s)], \tag{9}$$

where we have also applied the balanced growth condition that prices of final good and interest rates remain stationary before any policy intervention, that is, $P(s-1) = P(s-2)$ and $r(s) = r(s-1)$. As the new population control policy acts, the prices of final good and interest rates may deviate from their previous balanced growth path equilibrium values. In this case, $n(\alpha^l, s)$ is determined according to

$$\left[n(\alpha^l, s)\right]^\epsilon = \alpha^l \cdot [1 + r(s+1)] \cdot \left[\frac{P(s)}{P(s+1)}\right]^\sigma \cdot \left\{\frac{\mu P(s-1)[1 + r(s)] - 1}{\mu P(s)[1 + r(s+1)] - 1}\right\}^{1-\sigma}. \tag{10}$$

[16] We define the decentralized equilibrium in Section A.2.

[17] We here take the case of $n^* < n^h$ as an example. Letting $n^* \geq n(\alpha^h, s-1)$ would not make any difference on the mechanisms illustrated below.

Combining (9) and (10) yields

$$\left[\frac{n(\alpha^l,s)}{n(\alpha^l,s-1)}\right]^{\epsilon}=\underbrace{\left[\frac{1+r(s+1)}{1+r(s)}\right]}_{\Phi_1>1}\cdot\underbrace{\left[\frac{P(s)}{P(s+1)}\right]^{\sigma}}_{\Phi_2<1}\cdot\underbrace{\left\{\frac{\mu P(s-1)[1+r(s)]-1}{\mu P(s)[1+r(s+1)]-1}\right\}^{1-\sigma}}_{\Phi_3<1}. \tag{11}$$

Intuitively, raising the permitted number of children from $\tilde{n}$ to n^* allows the α^h-type household members to give birth to more children. Children in generation s then suddenly increases, which means that the capital per adult in generation $s+1$ goes down. This is the partial equilibrium result and we label this as "limits-raising channel". Equation (11) then reveals that a relaxation of birth quotas generates additional general equilibrium influence by adjusting households' fertility decision via the following approaches.

From Equation (6), we can see that both the prices of final good and interest rates at period of $s+1$ rise in response. Higher interest rates $r(s+1)$ (relative to $r(s)$) produces two opposing effects on the α^l-type household's fertility decisions. On the one hand, it tends to raise the birth rate, captured by the aggregate term $\Phi_1>1$. Intuitively, increased interest rates $r(s+1)$ encourage households to have more children in that fertility is the rate of investment in these descendants; that is, higher interest rates imply increased return rates of saving in the form of the number of descendants. It is therefore optimal for households to consume less and spend more income on child-rearing, resulting into an increase in the number of descendants.[18]

On the other hand, higher interest rates can also generate a counteracting negative effect on the α^l-type household's birth rate. Because of higher interest rates $1+r(s+1)$, the household faces an increased opportunity cost of raising a child; higher interest rates in the future make current spending on child-rearing more expensive, captured by $\Phi_3<1$. According to Equation (3), larger child-rearing costs in generation s are associated with higher consumption per adult in generation $s+1$. Then, as shown in Equation (4), the birth rates in generation s need to decrease to ensure an increase in consumption per adult in generation $s+1$.

Moreover, higher prices of final good $P(s+1)$ (relative to $P(s)$), captured by $\Phi_2<1$, tend to decrease the α^l-type household's fertility rate. Similarly, Equation (3) implies that higher prices of final good reduces consumption per adult in generation $s+1$. As a result, households must limit the number of children in generation s to ensure consumption growth in generation $s+1$. We label the latter two negative effects, Φ_2 and Φ_3, together as "child-rearing cost channel". Therefore, if the "child-rearing cost channel" is sufficiently large, the α^l-type household's birth rate will decline, that is, $n(\alpha^l,s)<n(\alpha^l,s-1)=\tilde{n}$. Otherwise, the α^l-type household's birth rate

[18]This can also be verified in (A.8); an increased number of descendants is at the sacrifice of a lower level of consumption per adult. Since $\hat{N}(t)$ or $L(t)$, and $K(t)$ are given in generation t, the level of output would not be affected by the announcement of population control policy.

will increase.

As for the impact on the α^h-type household's fertility decision, since the government relaxes the birth quotas in a large magnitude (i.e., n^* is larger than $\tilde{n}$ in a sufficiently large scale), which supports the dominance of the "limits-raising channel" (a partial equilibrium outcome) over the other general equilibrium impacts. As a result, the α^h-type household's birth rate always increases. We now examine the influence on the total birth rate, which is given by

$$\begin{aligned} B(t=s) &= \frac{\hat{N}(s+1)}{\hat{N}(s)} = \frac{q \cdot n(\alpha^h,0) \cdot (\tilde{n})^{s-1} \cdot n^* + (1-q) \cdot n(\alpha^l,0) \cdot (\tilde{n})^{s-1} \cdot n(\alpha^l,s)}{q \cdot n(\alpha^h,0) \cdot (\tilde{n})^{s-1} + (1-q) \cdot n(\alpha^l,0) \cdot (\tilde{n})^{s-1}} \\ &= \frac{q \cdot n(\alpha^h,0) \cdot n^* + (1-q) \cdot n(\alpha^l,0) \cdot n(\alpha^l,s)}{q \cdot n(\alpha^h,0) + (1-q) \cdot n(\alpha^l,0)}. \end{aligned} \tag{12}$$

Since that the total birth rate in periods ahead of the new population policy is $\tilde{n}$, we have

$$B(t=s) \gtrless B(t=s-1) \Leftrightarrow q \cdot n(\alpha^h,0) \cdot (n^* - \tilde{n}) \gtrless (1-q) \cdot n \cdot (\alpha^l,0) \cdot \left(\tilde{n} - n(\alpha^l,s)\right). \tag{13}$$

Given a sufficiently large relaxation of birth quotas (i.e., $n^* > \tilde{n}$), the total birth rate is likely to increase if (1) the low preference household's birth rate increases (i.e., $n(\alpha^l,s) > \tilde{n}$); (2) the size of household α^h at period $t = 1$ is large relative to the one of household α^l, represented by the comparison of $qn(\alpha^h,0)$ and $(1-q)n(\alpha^l,0)$. The following proposition summarizes these results.

Proposition 1. *Relaxing fertility restrictions (i) raises the short-term child-rearing costs; (ii) raises or reduces the the short-term low preference (α^l-type) household's birth rate and total birth rate; (iii) raises the high preference (α^h-type) household's birth rate.*

Proof. Proven in text. □

To help build up a more straightforward and closer connection of our theoretical predictions and empirical tests, we make a numerical example as follows. Suppose that the one-child policy is represented by $\tilde{n} = 1$. This policy was phased out in 2015 and replaced by the universal two-policy such that $n^* = 2$. According to the Proposition 1, our model predicts that the relaxation of birth quotas does release the birth restriction on the high preference household, which, of course, leads to an increased second-child birth rate. The other side of the coin is that the sudden increase number of children may create an equilibrium influence on the economy. This in turn can enhance or depress the fertility desire of the low preference households. In this case, the first-child birth rate may decrease or increase in response. Once the rise in the second-child birth rate is not able to cover the decline in the firstborn birth rate, the total birth rate could even decrease as a result of relaxing birth quotas. The best result, the government expects to see, is

that both the first-child and second-child birth rate increase. At least, the increased birth rate of the second-child is able to remedy the loss of the first-child birth rate.

3 Data, Variables, and Summary Statistics

3.1 Childbearing outcomes

To measure childbearing outcomes, we use a 10% random sample of the 2015 population census conducted by China's National Bureau of Statistics (NBS). The data include about 1.37 million individuals who collectively represent the total population in China. An obvious advantage is that the census data contain rich details on childbirth for women between 15 and 50 years of age, such as the timing of childbirth at the monthly level.[19] Based on the timing of childbirth, we define a birth in month t as a woman giving birth to a child in month t. We then aggregate all births at the province level, combined with sampling weights provided by the NBS, to calculate total births in province s in month t. The census data also include individual characteristics such as age, gender, education, hukou status (rural/urban), number of total births and social security participation, which allows us to calculate monthly births by cohort. In particular, information on the number of total births helps us distinguish between first children and second children. As a result, we construct panel data on monthly second-child (first-child) births for all 31 provinces from November 2014 to October 2015.

3.2 Relaxation of birth quotas

To quantify relaxation of birth quotas, we use data on the two-child policy announced by the Chinese government in November 2013 and gradually adopted by provinces on different dates in 2014.[20] For example, Zhejiang, Jiangxi and Anhui adopted in January 2014, and Hebei and Shandong adopted in May 2014, whereas Tibet and Xinjiang adopted in November and December 2014, respectively. We focus on variations in the timing of this relaxed birth quotas to identify its causal impacts. We further estimate the treatment effect by taking advantage of variations in exposure to the relaxed birth quotas due to differences in the numbers of eligible couples (for details, see **Section 5.6**). **Figure 3a** illustrates the timing distribution of relaxing birth quotas across provinces. Given that different provinces adopt the same policy on different dates, even within the same month (e.g. February 5^{th} versus February 25^{th}), we assume it takes effect in

[19] A possible concern is that inferring births using a retrospective history may lead to significant measurement error due to infanticide. However, childbirth outcomes in the census include children who were not alive at the time of the census survey, which largely mitigates this concern.

[20] See **Appendix C** for details. Also, as discussed in **Section C.5**, the anticipation effect is not likely to be a major concern in our context.

the same (next) month when adopted before (after) the mid-month. We also assume it takes a minimum of one year to plan for and have a second child, which includes one month to obtain a birth permit, one month to get pregnant (Olsen, Juul and Basso (1998); Gnoth, Godehardt, Godehardt, Frank-Herrmann and Freundl (2003); Eisenberg, Thoma, Li and McLain (2021)) and 10 months of pregnancy. See **Appendices C.1 to C.4** for more details on the recoding of the timing of relaxing birth quotas.

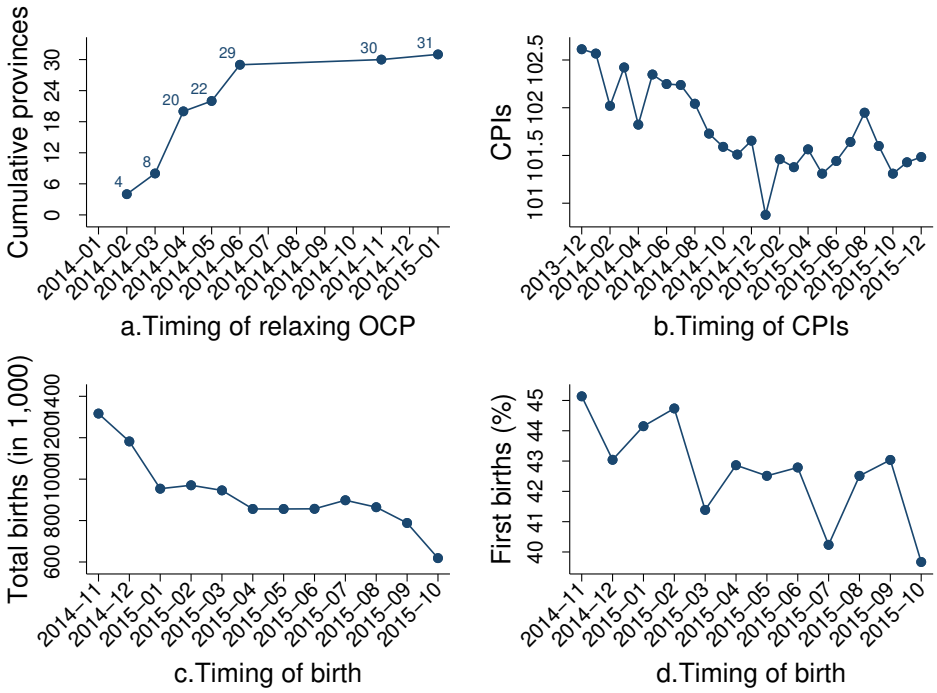


Figure 3: Descriptive summary for the relaxation of birth quotas, CPIs and births

Notes: This graph shows (a) the cumulative number of provinces that relaxed birth quotas over time, (b) the trends in consumer price indices over time, (c) the total number of births (in 1,000 people) over time, and (d) the percentage of first-child births based on total births over time. All variables are at the monthly level. Data on the timing of relaxing birth quotas are collected from official documents. Data on CPIs come from China's National Bureau of Statistics. Data on total births and percentage of first-births from November 2014 to October 2015 are constructed using the 2015 census.

3.3 Child-rearing costs

Measuring child-rearing costs is challenging in the context of China largely due to limited data. Because we cannot directly calculate overall child-rearing costs for each household, we use an alternative approach and measure changes in the price of the final good (and services) locally. Specifically, if raising a child costs a fixed unit of final good, then the change in the

price of final good locally can capture changes in cost of child rearing locally. To some extent, this approach reflects the child-rearing costs defined in the theoretical model. We construct three different measures for changes in the price of final good and services locally. The first refers to CPIs, which measure the average change in prices over time that consumers pay for a basket of goods and services. Data on CPIs in each province from August 2014 to July 2015 are obtained from China's National Bureau of Statistics. The overall CPIs, which aggregate the prices of a basket of goods and services, may overstate or understate the true costs of raising a child. To mitigate this concern, we look at CPIs by expenditure category to distinguish between expenditures that are more relevant to childrearing. **Figure 3b** shows the average CPIs over time. Overall CPIs decreased gradually beginning in December 2013 and reached their lowest point in January 2015.

It is noteworthy that housing costs, which account for a considerable part of the cost of raising a child in China, are not fully captured by CPIs. Previous studies have revealed the importance of home prices for fertility decisions (Yi and Zhang (2010); Lovenheim and Mumford (2013); Dettling and Kearney (2014); Liu, Xing and Zhang (2020)). Thus, to complement the CPIs findings, we also assess changes in the cost of child rearing associated with residential real estate sale prices, which we obtain from China's NBS from August 2014 to July 2015 in 70 large and medium-sized cities except in Tibet. In addition to overall residential real estate sale price indices, we also sort residential real estate sale price indices by type of building (i.e., newly constructed residential, newly constructed commercial and second-hand residential) and by floor area (i.e., 90 m^2 and smaller, 90 to 144 m^2 and larger than 144 m^2).

An advantage of residential real estate sale price indices by category is that they allow us to examine the heterogeneous impacts of relaxing birth quotas on the cost of housing across households with different sensitivity to housing costs. For example, if the area of residential buildings captures households' sensitivity to housing costs, disadvantaged couples should be more sensitive to prices of smaller residences (i.e., 90 m^2 and below) than to larger ones (i.e., above 144 m^2).

Both CPIs and residential real estate sale price indices variables measure changes in the cost of raising a child from the time when the relaxed birth quotas begins to affect local births to at least 12 months after the relaxation of birth quotas. As a child grows, demand for some goods and services (e.g. housing, education, health care) continuously increases, whereas demand for others (childcare) is likely to peak during the first year of a child's life and then decline. To further capture changes in the cost of child rearing with respect to hiring a nanny, we construct a third measure, nanny wages, at the individual level. Specifically, we use the China Migrants Dynamic Survey ($2012 - 2015$), a nationally representative dataset conducted by the National

Health Commission of the People's Republic of China with a total sample size of 764,288. This micro-level dataset contains wage information for migrant workers across occupations (including nannies), which enables us to estimate the causal impact of relaxing birth quotas on nanny wages in a DID setup.

Table 1: Summary statistics for births

Variables	N	Mean	Median	Std. Dev.	Min	Max
Monthly second-child births	372.00	17,062.00	13,656.85	13,731.51	0.00	75,567.24
Monthly second-child births of women from rural	372.00	10,317.26	7433.91	9905.70	0.00	63,096.05
Monthly second-child births of women from urban	372.00	6744.74	5242.15	5572.66	0.00	32,850.81
Monthly second-child births of women with college degree	372.00	2424.38	1687.54	2587.30	0.00	14,227.99
Monthly second-child births of women without college degree	372.00	14,637.62	11,503.95	12,645.93	0.00	72,435.29
Monthly second-child births of old women	372.00	10,929.86	8135.93	9523.78	0.00	59,647.27
Monthly second-child births of young women	372.00	6132.14	4490.26	5325.93	0.00	26,971.66
Monthly second-child births of women with pension	372.00	12,368.95	9794.39	11,064.35	0.00	65,058.43
Monthly second-child births of women without pension	372.00	4693.05	3287.09	4597.44	0.00	25,136.28
Monthly second-child births of Han women	372.00	15,021.90	9920.99	13,639.28	0.00	75,567.24
Monthly second-child births of minorities women	372.00	2040.10	501.49	3748.31	0.00	23,855.63
Monthly second-child births of native women	372.00	15,962.11	12,735.53	13,259.95	0.00	74,028.93
Monthly second-child births of migrant women	372.00	1099.89	0.00	1924.65	0.00	11,396.63
Monthly first-child births	372.00	12,799.36	10,386.18	10,001.90	0.00	48,436.91
Monthly first-child births of women with college degree	372.00	3733.19	2732.64	3771.27	0.00	23,112.90
Monthly first-child births of women without college degree	372.00	9066.17	7097.01	7709.10	0.00	41,661.64
Monthly first-child births of old women	372.00	4364.47	3297.94	3899.93	0.00	20,908.30
Monthly first-child births of young women	372.00	8434.88	6439.68	7168.76	0.00	36,508.29

Notes: We calculate monthly first-child and second-child births using China's Population Census in 2015. We define women from rural (urban) areas as those women who were (not) granted land-contracting rights. We define old (young) women as those women above (below) the median value of the childbearing age distribution. We define native women as those women registered in the same province, and define migrant women as those women registered in other provinces.

Tables 1 and 2 show the summary statistics of the main variables. **Figures 3c and 3d** show the number of total births at the monthly level and the percentage of first-child births among total births at the monthly level, respectively. **Figure 4** shows the monthly second-child births between treatment and control groups over time in an event-study setup.[21] We find that monthly second-child births rebounded significantly in the treatment group after the policy relaxation. By contrast, we do not find similar patterns for the control group.

4 Model Specification and Identification Strategy

For each birth outcome, the following regression specification evaluates the impact of relaxing birth quotas as a quasi-experiment. The treatment is the degree of exposure to the relaxation of

[21]The treatment group includes provinces that relaxed birth quotas between February and April 2014 and the control group includes provinces that did not relax birth quotas between February and May 2014. Using the number of monthly second births between November 2014 and October 2015 as the outcome variable, we define 'time since relaxation of birth quotas' as the time lag between the date when second births occurred and the date 12 months after the province relaxed birth quotas.

Table 2: Summary statistics for child-bearing costs

Variables	N	Mean	Median	Std. Dev.	Min	Max
Overall consumer price indices	372.00	101.52	101.50	0.64	99.90	103.60
Consumer price indices of food	372.00	102.22	102.30	1.33	97.60	105.60
Consumer price indices of tobacco and cigarette	372.00	100.38	99.90	1.98	95.70	106.00
Consumer price indices of clothing	372.00	102.93	102.65	2.07	97.20	115.20
Consumer price indices of household facility, article and maintenance service	372.00	101.07	101.00	0.86	99.20	104.00
Consumer price indices of healthcare	372.00	101.59	101.50	0.94	98.30	105.00
Consumer price indices of transportation and communications	372.00	98.94	99.00	1.27	94.90	103.20
Consumer price indices of education, culture and entertainment	372.00	101.68	101.50	1.60	96.60	108.90
Consumer price indices of housing	372.00	101.12	101.10	1.24	96.90	105.10
Sale price indices of residential buildings	840.00	95.83	95.40	3.10	89.20	123.60
Sale price indices of new constructed residential buildings	840.00	95.63	95.20	3.18	88.80	124.00
By floor area: a. 90 m2 and below	840.00	95.99	95.60	3.28	87.60	124.10
b. From 90 to 144 m2	840.00	95.68	95.30	3.18	89.10	124.80
c. Above 144 m2	840.00	95.00	94.50	3.47	86.60	123.50
Sale price indices of second-hand residential buildings	840.00	96.44	96.15	3.08	85.80	124.30
By floor area: a. 90 m2 and below	840.00	96.67	96.50	3.25	84.80	125.10
b. From 90 to 144 m2	840.00	96.53	96.20	2.97	88.60	124.20
c. Above 144 m2	840.00	95.67	95.50	3.12	86.90	121.80

Notes: We collect data on monthly consumer price indices from August 2014 to July 2015 in each province, and collect data on sale price indices of residential buildings from August 2014 to July 2015 in 70 large and medium-sized cities across all provinces except for Tibet. Both datasets are provided by National Bureau of Statistics of China.

birth quotas.[22] We estimate the treatment effect in a DID framework, following Chandra, Gruber and McKnight (2010) and Agersnap, Jensen and Kleven (2020). The model specification is as follows:

$$Y_{s,t_{birth}} = \boldsymbol{\alpha}' I_s + \boldsymbol{\beta}' I_{t_{birth}} + \gamma \left(I_{s,t_{birth} \geq t_{policy+r}} \right) + \varepsilon_{s,t_{birth}}, \tag{14}$$

where $Y_{s,t_{birth}}$ is the logarithm of monthly second (first, total) births in province s in month t_{birth},[23] I_s is a vector of province fixed effects, $I_{t_{birth}}$ is a vector of time fixed effects, $I_{s,t_{birth} \geq t_{policy+r}}$ is a dummy for an observation r months after birth quotas are relaxed in province s (specifically, an interaction of an indicator variable for being in province s when birth quotas are relaxed and an indicator for being in the r^{th} month after birth quotas are relaxed), or an observation after the de facto relaxation of birth quotas. r is the time interval between the date when a province relaxed birth quotas and the date when this policy de facto affected the number of local births. In the benchmark model, we assume that $r = 12$, meaning that it takes about 12 months, as previously

[22] In our benchmark model, we mainly quantify the degree of exposure to the relaxation of birth quotas based on the timing of policy adoption. In **Section 5.6**, we make further use of differences in the number of couples eligible to have a second child and estimate the treatment effect based on a triple difference strategy. We find similar results using these two different approaches.

[23] In our empirical analysis, we mainly use the number of births rather than birth rate (or fertility rate) as the outcome variable because the panel data analysed in the paper are at the monthly level within a calendar year, whereas we only have the data on total population at the yearly level. A likely concern is that ignoring differences in the number of women of childbearing age contaminates our results. Controlling for province fixed effects partially mitigates this concern. In fact, when using fertility rate (i.e. number of births divided by number of women of childbearing age) at the time of the 2015 Census survey as an alternative outcome variable, our main results remain robust (**Figure B.12**). As an additional robustness check, we use micro-level data to conduct the analysis and find similar results.

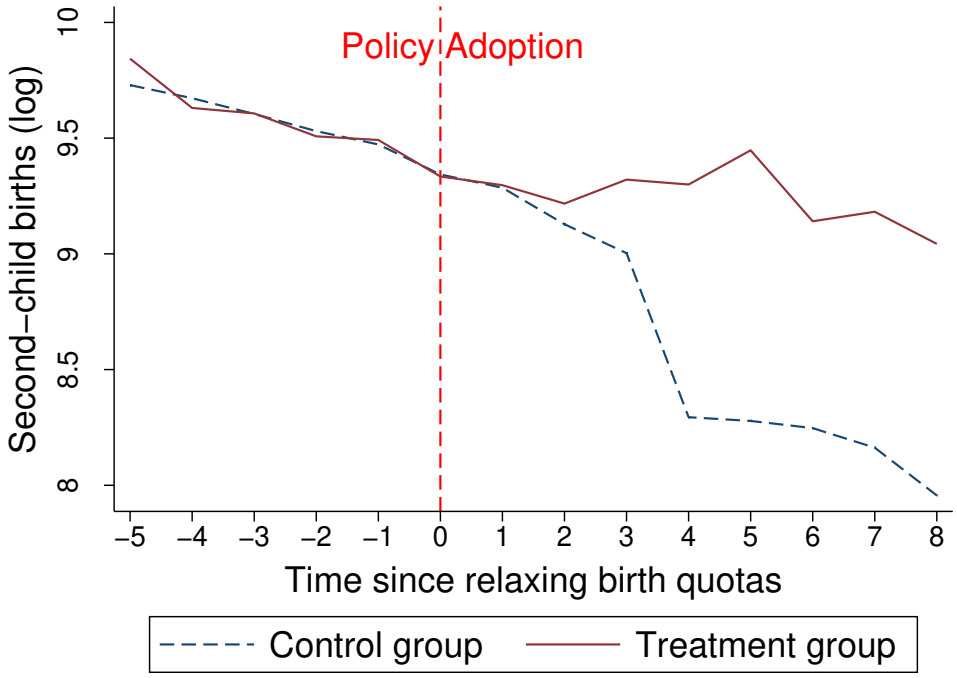


Figure 4: Number of second-child births (in log) over time

Notes: The treatment group consists of provinces that relaxed birth quotas between February 2014 and April 2014. The control group consists of provinces that relaxed the OCP after May 2014. We use the monthly second-child births (in log) between November 2014 and October 2015 as the outcome variable. We define "time since relaxing birth quotas" variable as the time lag between the date when second births occurred and the date 12 months after the province relaxed birth quotas.

described, for the policy to affect local second-child births. The error term is $\varepsilon_{s,t_{birth}}$, $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are vectors of coefficients to be estimated and γ is the coefficient of interest. We cluster standard errors at the provincial level.[24]

The assumptions underlying the identification of parameter γ are similar to all DID analyses, in that the decision about whether and when to relax birth quotas must be uncorrelated with any prior trends in birth outcomes and that the timing of relaxing birth quotas cannot coincide with any province-specific shocks or policies that might influence birth outcomes. Also, we note several potential challenges for the identification. Recent studies show that, conditional on a constant treatment effect across groups and over time, the two-way fixed effects regression above identifies the effect under the standard 'common-trends' assumption. However, it is often

[24]To mitigate the concern that clustering standard errors by province would not work well due to too few clusters, we use the wild bootstrap approach as a robustness check (Cameron, Gelbach and Miller (2008); Roodman, Nielsen, MacKinnon and Webb (2019)). Using the Stata code Boottest proposed by Roodman, Nielsen, MacKinnon and Webb (2019), we find our main results are robust.

implausible for the treatment effect to be constant. If the constant effect assumption is not supported, the two-way fixed effects regression identifies weighted sums of the average treatment effects (ATEs) in each group and period, with weights that may be negative (de Chaisemartin and D'Haultfoeuille (2020); Sun and Abraham (2020); Athey and Imbens (2022); Callaway and Sant'Anna (2021); Imai and Kim (2018); Goodman-Bacon (2018)).

The second potential identification problem regarding the correlation between unobserved, province-specific shocks and the relaxation of birth quotas is more difficult to investigate. We thus conduct a falsification test using birth outcomes for women migrating from other provinces (i.e. inter-provincial migrants). Under China's hukou system, the relaxation of birth quotas mainly affects fertility decisions of women registered in the same province (i.e. local natives). If the relaxation of birth quotas really works according to our hypothesis, we should not find that this policy has a similar and significant impact on birth outcomes for inter-provincial migrants who come from provinces that relaxed birth quotas at different points in time.

We go further and also use alternative population censuses (i.e., 2005 and 2010) to construct birth outcomes for all 31 provinces from November 2004 to October 2005 and birth outcomes for all 31 provinces from November 2009 and October 2010, respectively.[25] Given that these second-child births occurred prior to the relaxation of birth quotas, we use them to conduct placebo analyses. If our main estimates are not driven by other unobserved province-specific shocks (e.g., seasonal effects), then when using these counterfactual birth outcomes, we should not find that the relaxation of birth quotas has similar and significant impacts.

5 Empirical Results

This section reports the effects of relaxing birth quotas. First, we show its impact on second-child births and whether the impact is heterogeneous across groups. Second, we explore whether and how it affects first-child births. Third, we explore its effects on child-rearing costs, and finally, its impact on total births.

5.1 Pretreatment trends

A crucial assumption that justifies the DID approach is the presence of common parallel trends in the pretreatment periods. **Figure** 4 shows similar pretreatment trends. As a more formal test of this assumption, we use the estimator proposed by de Chaisemartin and D'Haultfoeuille

[25]This exercise partially mitigates the concern that seasonality effects (i.e. timing of fertility decisions within a calendar year) would confound the policy effect. If the seasonality effects dominate, we should find similar effects using the alternative Census data prior to the relaxation of birth quotas.

(2020).[26] If the common parallel assumption is valid, we should find that relaxation of birth quotas has no significant impacts in the pretreatment periods. **Figure 5a** shows the placebo effects in the pretreatment periods for the outcome variable.[27] The horizontal axis denotes the values corresponding to the time relative to the de facto relaxation of birth quotas, which are negative in the pretreatment periods, and the vertical axis represents the values corresponding to the changes in the outcome of interest (i.e. the logarithm of second-child births). The points denote the estimated treatment effects in the treatment periods and placebo effects in the pretreatment periods, and the error bars denote 95% confidence intervals of the estimated effects and placebos. The estimation results show that the placebo effects are close to zero and not statistically significant at the conventional level in the pre-treatment periods.

5.2 Impact of relaxing birth quotas on second-child births

Using equation (14), we estimate the impact of relaxing birth quotas on second-child births. Column 1 of **Table B.1** reports the main results. We find that the coefficient of interest is -0.007, which is not statistically significant at the conventional level. As mentioned earlier, we check the negative weight problem and compute the weights attached to each average treatment effect in group s and period t. We find that the negative weight problem exists in the data.[28] To reduce the concern that the two-way fixed effects model outlined above may not capture the average treatment effects in the presence of heterogeneous treatment effects, we use the estimator proposed by de Chaisemartin and D'Haultfoeuille (2020) to estimate the impact of relaxing birth quotas on second-child births. **Figure 5a** shows the main results.

The relaxation of birth quotas has had a pronounced effect on second-child births, with an average increase of approximately 37.9% in monthly figures in the immediate aftermath. This surge can be largely attributed to the pent-up demand among couples who were previously restricted from having more than one child. As a result of these prior limitations, a significant number of couples harboured the desire to expand their families with a second child. Supporting this observation, data from the National Health Commission of China reveals that the total number of second-child births escalated from 5.11 million in 2013 to 6.52 million in 2015. This represents a notable national uptick of about 27.6%.

[26] For each treatment period t, the estimator compares the change of the mean outcome between $t-1$ and t in two sets of groups: the reform adopters (i.e. those switching from untreated to treated) and those who remain untreated. For each pretreatment period t, the (placebo) estimator essentially compares the outcome's change from $t-2$ to $t-1$ in groups that relax birth quotas and those that do not relax birth quotas between $t-1$ and t. For details on the estimator, see de Chaisemartin and D'Haultfoeuille (2020).

[27] We use STATA code *did_multiplegt* to draw this figure.

[28] Under the common-trends assumption, $\hat{\gamma}$ estimates a weighted sum of 199 ATTs. A total of 159 ATTs receive a positive weight, and 40 receive a negative weight.

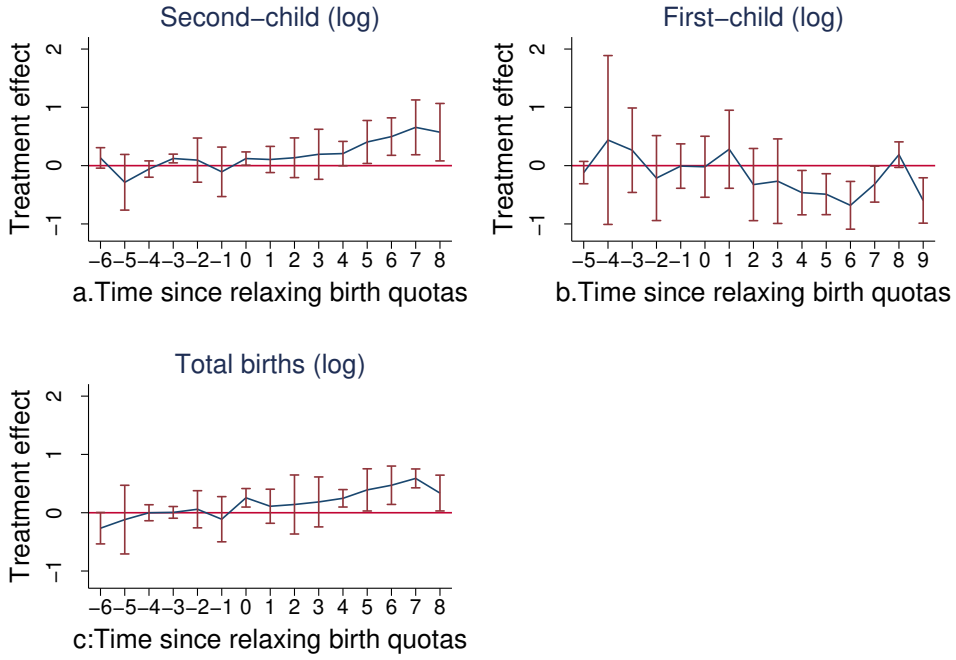


Figure 5: Effects of relaxing birth quotas on second-child, first-child and total births.

Notes: This graph shows the impact of relaxing birth quotas on (a) second-child births, (b) first-child births, and (c) total births. In each panel, the horizontal axis represents the values corresponding to the time relative to the (de facto) relaxation of birth quotas, which are negative in the pretreatment periods, and the vertical axis represents the values corresponding to the change in the outcome of interest. The points represent the estimated treatment effects in the treatment periods and placebo effects in the pretreatment periods, and the error bars represent 95% confidence intervals of the estimated effects and placebos. Both first-child,second-child and total birth variables are in the logarithm form. The panel data on first-child, second-child and total births from November 2014 to October 2015 are constructed using the 2015 census. To draw these figures, we followed the estimation strategy proposed by de Chaisemartin and D'Haultfoeuille (2020).

The impact is heterogeneous over time. Relaxing birth quotas increases second-child births by 12.7% in the month of de facto relaxing birth quotas, an impact that further increases to 92.7% in the seventh month of de facto policy adoption.

As discussed, relaxing birth quotas mainly applies to women registered in the same province. Using monthly births of inter-provincial migrants to conduct a falsification test, we do not find obvious evidence that relaxing birth quotas increases second-child births for inter-provincial migrants. By contrast, relaxing birth quotas significantly increases second-child births for local natives. The results are shown in **Figures B.1a and B.1b**.

Finally, we use the counterfactual monthly second-child births from the 2005 and 2010 population censuses to conduct placebo analyses. As shown in **Figures B.2a and B.2c**, we do not find

obvious evidence that relaxing birth quotas affects the counterfactual second-child births. These placebo analyses therefore further alleviate the concern that our estimation results are driven by unobserved province-specific shocks (e.g., seasonal effects). All of these findings suggest that relaxing birth quotas substantially increases second-child births, which is consistent with our theoretical model.

5.3 Which groups are more responsive to the relaxation of birth quotas?

We now further explore which groups are more responsive to the relaxation of birth quotas. Our main results are reported in Figures B.3 to B.7. As shown in **Figures B.3a and B.3b**, regarding childbearing age, the impact is more pronounced for women of relatively early rather than late childbearing age (below the median value of the childbearing age distribution). As shown in **Figures B.4a and B.4b**, the impact in relation to women's education is more pronounced for women without a college degree than women with a college degree or above. Regarding home-ownership, **Figures B.5a and B.5b** show that the impact of relaxing birth quotas on second-child births is more pronounced for women who do not own a home. In relation to the gender of children, couples in China are more likely to have a second child if their first child is a girl. **Figures B.6a to B.6d** corroborate this assumption and show that the impact is more pronounced for couples who have no male child or have a female child at the time of the survey.

Finally, we examine the impact of relaxing birth quotas in relation to pension accessibility. On the one hand, having access to pensions may not be associated with fewer children if parents derive utility from that of their children (i.e. the intergenerational altruism motive). On the other hand, having access to pensions may crowd out childbearing decisions if government-provided social security is substitutable for old age support from children (i.e. the old-age security motive for fertility). **Figures B.7a and B.7b** report the main results. We find that the impact of relaxing birth quotas on second-child births is more pronounced for couples covered by pensions than for couples without access to pensions, which favours the intergenerational altruism motive rather than the old-age security motive for fertility.

Overall, these findings suggest that fertility preferences, rather than sensitivity to child-rearing costs, matter more for second-child births and that couples with higher fertility preferences are more likely to have a second child after the relaxation of birth quotas, regardless of their economic conditions. These findings are again consistent with our theoretical model.

5.4 Does relaxing birth quotas affect first-child births?

As mentioned in Section 2, although relaxing birth quotas itself does not directly affect first-child births, it may indirectly increase or decrease (or delay) first-child births through general equilibrium effects. To empirically investigate the impact of relaxing birth quotas on first-child births, we assume that it takes at least one month to conceive a first child. Because the relaxed birth quotas remove the requirement for a birth permit to have a first child, we assume that it takes at least 11 months for the relaxation of birth quotas to affect local first-child births ($r = 11$).

Using the same empirical approach, we find that the coefficient of interest is -0.05, which is not statistically significant at the conventional level. We also find that the negative weight problem exists in the data.[29] To reduce the concern that the two-way fixed effects model outlined above may not capture the average treatment effects in the presence of heterogeneous treatment effects, we again use the estimator proposed by de Chaisemartin and D'Haultfoeuille (2020) to estimate the impact of relaxing birth quotas on first-child births. **Figure 5b** shows the main results.

A straightforward analysis combining both instantaneous and dynamic effects indicates that the relaxation of birth quotas results in an average reduction of approximately 23.5% in monthly first-child births in the short term. This marked decline can be attributed to the heightened responsiveness of second-child births to alterations in policy, as well as couples strategically postponing their first-child decisions in anticipation of potential increases in second-child births in the near future. corroborating this trend, data from the National Health Commission of China shows that the total number of first-child births decreased from 10.56 million in 2013 to 8.86 million in 2015, marking a significant national decrease of about 16.1%.

The impact is also heterogeneous over time. We do not find obvious evidence that relaxing birth quotas affects first-child births in the month of de facto relaxing birth quotas. However, the same policy leads to a decrease of 49.3% in first-child births in the sixth month of de facto policy adoption. Afterwards, we find short-lived recovery that is possibly induced by delayed first-child births.

Using monthly births of inter-provincial migrants to conduct a falsification test, we do not find obvious evidence that relaxing birth quotas reduces first-child births for inter-provincial migrants. By contrast, relaxing birth quotas significantly reduces first-child births for local natives (**Figures B.1c and B.1d**). Additionally, we do not find similar effects when using counterfactual first-child births from November 2004 to October 2005 and from November 2009 to October 2010 (**Figures B.2b and B.2d**), further mitigating the concern that our estimates are driven by

[29] Under the common-trends assumption, $\hat{\gamma}$ estimates a weighted sum of 169 ATTs. A total of 137 ATTs receive a positive weight, and 32 receive a negative weight.

unobservable factors such as seasonal effects. These findings suggest that relaxing birth quotas substantially decreases or delays first-child births.

When further exploring whether the impact of relaxing birth quotas on first-child births is heterogeneous across couples with different characteristics (e.g. childbearing age, education, home ownership, access to pension), our findings show that the negative impact of relaxing birth quotas on first-child births is more pronounced for women who are relatively younger, less educated and not homeowners. By contrast, we find no obvious negative effect for women who are relatively older, more educated and homeowners. These findings further suggest that, unlike second-child births, sensitivity to child-rearing costs is a likely underlying mechanism through which the relaxation of birth quotas crowds out or delays first-child births. Details can be found in Figures B.3 to B.6.

5.5 Does relaxing birth quotas affect child-rearing costs?

As mentioned before, higher child-rearing costs are likely an underlying mechanism through which relaxing birth quotas exerts asymmetric effects on first-child and second-child births. We now go a step further and examine whether and how child-rearing costs respond using three price measures associated with childrearing as proxies for changes in child-rearing costs as described in **Section 3.3**.

5.5.1 Impact of relaxing birth quotas on consumer price indices

We start by using local CPIs to capture changes in the cost of child rearing in terms of prices of final goods and services locally. To estimate the impact of relaxing birth quotas on local CPIs, we use the following model specification:

$$CPI_{s,t_{birth}} = \boldsymbol{\alpha}^{'} I_s + \boldsymbol{\beta}^{'} I_{t_{birth}} + \gamma_2 \left(I_{s,t_{birth} \geq t_{policy+r}} \right) + \varepsilon_{s,t_{birth}}, \tag{15}$$

where $CPI_{s,t_{birth}}$ is CPIs in province s in month t_{birth}, I_s is a vector of province fixed effects, $I_{t_{birth}}$ is a vector of time fixed effects and $I_{s,t_{birth} \geq t_{policy+r}}$ is an indicator for an observation r months after relaxation of birth quotas or an observation after the de facto policy adoption. In this model, we assume that $r = 12$, meaning that it takes at least 12 months for the relaxation of birth quotas to begin to affect the local costs of raising a child. As discussed, if it takes 12 months for the relaxation of birth quotas to affect local second-child births, then the impact of relaxing birth quotas on local CPIs is likely to be small in the short term and then gradually increase. The error term is $\varepsilon_{s,t_{birth}}$, $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are vectors of coefficients to be estimated and γ_2 is the coefficient of interest.

Using the same approach, we estimate the dynamic impacts of relaxing birth quotas on local CPIs.[30] As shown in **Figure 6a**, a simple average of the instantaneous and dynamic effects shows that relaxation of birth quotas leads to an increase of about 0.48 unit (or 32% higher compared with the median value of year-on-year growth in local CPIs in the sample) in local monthly CPIs in the short term.

The impact is also heterogeneous over time, increasing slightly in the beginning and then more rapidly over time, suggesting that it takes a certain number of months to prepare for pregnancy and to get pregnant after the relaxation of birth quotas.

To further capture shifts in the cost of items more pertinent to raising a child, we examine the impact of relaxing birth quotas on CPIs by expenditure categories. We find that local CPIs for food increase by 4 units 17 months, compared to a decline in CPIs for the tobacco, alcohol and clothing. Overall, the evidence suggests that relaxing birth quotas significantly affects the prices of final goods and services locally, particularly those more relevant to raising a child. See **Figures B.8 to B.9** for more details.

5.5.2 Impact of relaxing birth quotas on residential real estate sale price indices

To examine the impact of relaxing birth quotas on the cost of child rearing in terms of home prices, we use the following model specification:

$$HP_{s,t_{birth}} = \boldsymbol{\alpha}^{'} I_s + \boldsymbol{\beta}^{'} I_{t_{birth}} + \gamma_3 \left(I_{s,t_{birth} \geq t_{policy+r}} \right) + \varepsilon_{s,t_{birth}}, \tag{16}$$

where $HP_{s,t_{birth}}$ is the sale price indices of residential buildings in city s in month t_{birth}, I_s is a vector of city fixed effects, $I_{t_{birth}}$ is a vector of time fixed effects and $I_{s,t_{birth} \geq t_{policy+r}}$ is an indicator for an observation r months after the relaxation of birth quotas or an observation after the de facto policy adoption. As before, we assume that $r = 12$, meaning that it takes at least 12 months for the policy to begin to affect local costs of raising a child. The error term is $\varepsilon_{s,t_{birth}}$,, $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are vectors of coefficients to be estimated and γ_3 is the coefficient of interest. Standard errors are clustered at the city level.

As shown in **Figure 6b**, a simple average of the instantaneous and dynamic effects suggests that relaxing birth quotas increases the local monthly residential real estate sale price indices in the short term by about 1.29 units (or 28% higher compared with the median value of year-on-year growth in local residential real estate sale price indices in the sample).

The impact is also heterogeneous over time, increasing slightly in the beginning and then

[30] Using a two-way fixed effects model, we find that the coefficient of interest is 0.27. Further investigation shows an obvious negative weight problem in the data. Specifically, under the common-trends assumption, $\hat{\gamma}_2$ estimates a weighted sum of 112 ATTs. A total of 90 ATTs receive a positive weight, and 16 receive a negative weight.



Figure 6: Effects of relaxing birth quotas on consumer price indices and residential real estate sale price indices.

Notes: This graph shows the impact of relaxing birth quotas on (a) consumer price indices and (b) residential real estate sale price indices. In each panel, the horizontal axis represents the values corresponding to the time relative to the (de facto) relaxation of birth quotas, which are negative in the pretreatment periods, and the vertical axis represents the values corresponding to the change in the outcome of interest. The points represent the estimated treatment effects in the treatment periods and placebo effects in the pretreatment periods, and the error bars represent 95% confidence intervals of the estimated effects and placebos. The panel data on consumer price indices and residential real estate sale price indices from August 2014 to July 2015 come from China's National Bureau of Statistics. To draw these figures, we followed the estimation strategy proposed by de Chaisemartin and D'Haultfoeuille (2020).

more rapidly over time, suggesting that it takes a certain number of months to get pregnant after relaxation of birth quotas.

To further capture changes in housing expenses for different demographic group, we examine the impact of relaxing birth quotas on residential real estate sale prices by floor area (90 m^2 and below, from 90 to 144 m^2 and above 144 m^2). We find that the impact decreases as floor area increases, which is true for both newly constructed (shown in **Figure B.10**) and second-hand (shown in **Figure B.11**) residential buildings. In other words, disadvantaged couples who are more sensitive to the cost of housing are actually more likely to be exposed to higher housing costs after birth quotas are relaxed, compared with advantaged couples who are less sensitive to the cost of housing.

A concern is that higher residential real estate sale prices are associated with higher assets for homeowners such that changes in sale prices affects both their child-rearing costs and total wealth. If this is true, first-child births should decrease more for non-homeowners than for homeowners. To test this hypothesis, we investigate the impact of relaxing birth quotas on first-child births by homeownership status. Unsurprisingly, we find a negative impact for non-homeowners and a positive impact for homeowners, consistent with previous findings on the impact of home prices on fertility decisions (Yi and Zhang (2010); Lovenheim and Mumford (2013); Dettling and Kearney (2014); Liu, Xing and Zhang (2020)). However, we do not find similar patterns for the impact of relaxing birth quotas on second-child births, possibly because other factors (e.g. fertility preferences) matter more for having a second child. More details are provided in Figures B.5a to B.5d.

5.5.3 Impact of relaxing birth quotas on nanny wages

To examine the impact of relaxing birth quotas on the cost of child rearing with respect to hiring a nanny, we use the following model specification:

$$y_{i,s,t} = \boldsymbol{\alpha}^{'} Caregiver_{i,s,t} + \boldsymbol{\beta}^{'} I_{i,t} + \gamma_{2012} \left(Caregiver_{i,s,t} \cdot I_{i,2012}\right) + \gamma_{2014} \left(Caregiver_{i,s,t} \cdot I_{i,2014}\right) + \delta^{'} I_{i,s} + K_{i,s,t} + \varepsilon_{i,s,t} \tag{17}$$

where $y_{i,s,t}$ is the logarithm of the monthly wage in April in province s in year t for migrant i, $Caregiver_{i,s,t}$ is a dummy variable indicating whether the main occupation for migrant i in province s in year t is a nanny, $Policy_{i,t}$ is a dummy variable indicating whether migrant i is exposed to the relaxation of birth quotas in year t, $I_{i,s}$ is the province fixed effects for migrant i, $I_{i,t}$ is the year fixed effects for migrant i, $I_{i,2012}$ and $I_{i,2014}$ refer to year dummies in 2012 and 2014, respectively. $K_{i,s,t}$ refers to individual characteristics such as gender, age, age squared, education, marriage status, hukou type, duration of migration and duration of migration squared for migrant i in province s in year t. $\varepsilon_{i,s,t}$ is the error term. The parameter of interest γ_{2014} captures the impact of relaxing birth quotas on nanny wages. The parameter γ_{2012} captures the extent of pre-treatment trends differences.

Considering the survey data provide wage information only for April of each survey year, our main analysis restricts samples to provinces that relaxed birth quotas from January to April 2014 and to survey years from 2012 to 2014. Moreover, we conduct two placebo analyses by either restricting samples to provinces that relaxed birth quotas after April 2014 (i.e. in the absence of the policy when conducting the survey) or using survey years in 2012, 2013 and 2015. The rationale for the first analysis is that nanny wages should not be affected in provinces where the survey was conducted in the absence of relaxing birth quotas, and the rationale for the second

analysis is that higher nanny wages attract more nannies and eventually narrow wage differences over time.

Our main findings show that relaxing birth quotas increases monthly nanny wages by about 9.3% on average in the short term (Columns 1 to 4 of Table 3). We do not find any significant differences in pre-treatment periods, supporting the common parallel assumption. Also, we find that the impact is larger for migrant workers registered in the same city (i.e., local migrants) than for migrant workers registered in other provinces (i.e., non-local migrants). Specifically, relaxing birth quotas increases nanny wages of local migrants and non-local migrants by 22.4% and 7.6%, respectively (Columns 5 and 6 of Table 3). A likely explanation is that the labor supply of local nannies is much less responsive to higher wages than non-local nannies due to the population size. Our main findings are rejected when using two different placebo analyses, further supporting that higher nanny wages are driven by relaxation of birth quotas rather than other unobservable factors (Columns 7 and 8 of Table 3).

Table 3: Impact of relaxing the OCP on nanny wages

Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Pre-treatment trends	−0.0055	0.0213	0.0211	0.0214	0.0587	−0.0384	−0.0897	0.0189
	(0.0304)	(0.0284)	(0.0284)	(0.0284)	(0.0370)	(0.0660)	(0.0774)	(0.0283)
Treatment effect	0.0743***	0.0891***	0.0892***	0.0891***	0.0739**	0.2028***	−0.0523	0.0224
	(0.0282)	(0.0264)	(0.0264)	(0.0264)	(0.0368)	(0.0575)	(0.0613)	(0.0278)
Nanny (o/1)	−0.3180***	−0.1875***	−0.1857***	−0.1860***	−0.1750***	−0.1838***	−0.1350***	−0.1886***
	(0.0206)	(0.0193)	(0.0193)	(0.0193)	(0.0270)	(0.0399)	(0.0444)	(0.0193)
Observations	360664	360664	360284	360284	200106	63784	86901	352527
R-squared	0.0565	0.1753	0.1768	0.1769	0.1842	0.1596	0.1209	0.2061
F-stat	829.9535	2017.1035	1934.1461	1843.0295	1075.4069	310.4696	442.4326	2178.5219
Individual control variables	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: This table reports the impacts of relaxing the OCP on nanny wages based on equation (17). Wage variables are in the logarithm form. We use samples in provinces that relaxed the OCP from January to April 2014 for the main analysis. Column 1 reports the estimates without controlling for individual characteristics. Column 2 controls for male, age, age squared and marriage status. Column 3 further controls for hukou type. Column 4 further controls for migration duration and migration duration squared. Columns 5 and 6 estimate the impact by places of origin. Specifically, Column 5 uses samples of migrants registered in the same city. Column 6 uses samples of migrants registered in other provinces Column 7 estimates the placebo effect by using samples in provinces that relaxed the OCP after April 2014. Column 8 estimates another placebo effect by using samples in survey years in 2012, 2013 and 2015. Standard errors are in parentheses. $*p < 0.1$, $**p < 0.05$, $***p < 0.01$.

Overall, utilizing three distinct measures to assess the variations in the cost of child rearing locally, we consistently observe evidence indicating that the relaxation of birth quotas significantly amplifies the expenses associated with raising a child in the local context. This finding further strengthens the support for the mechanisms outlined in the theoretical model, which posits asymmetric effects of relaxing birth quotas on first-child and second-child births.

Notably, the discernible short-term impact of relaxing birth quotas on nanny wages provides

partial support for the idea that forward-looking couples encounter increased child-rearing costs in the current period due to anticipated growth in second-child births. Consequently, these couples adjust their current childbearing decisions in response to the relaxation of birth quotas.[31]

5.6 Do more eligible couples magnify the policy impact?

In the previous analyses, we use variations in the timing of birth quotas changes across provinces to estimate the treatment effect in a DID framework. In this subsection, we further take advantage of variations in the exposure to these birth quotas changes and resulting differences in the number of couples eligible to have a second child. We estimate the treatment effect based on a triple-difference strategy. Specifically, we use the ratio of eligible couples at the provincial level and examine whether the effect of the birth quotas change depends on the ratio of eligible couples across reform provinces. The model specification is as follows:

$$Y_{s,t_{birth}} = \boldsymbol{\alpha}^{'} I_s + \boldsymbol{\beta}^{'} I_{t_{birth}} + \gamma_4 \left(I_{s,t_{birth} \geq t_{policy+r}} \right) + \gamma_5 \left(I_{s,t_{birth} \geq t_{policy+r}} \right) \cdot Intensity_s + \varepsilon_{s,t_{birth}}, \tag{18}$$

where $Y_{s,t_{birth}}$ refers to second-child births, first-child births, CPIs, or housing price indices. $Intensity_s$ is the ratio of eligible couples in province s in which at least one partner is an only child. γ_5 is the parameter of interest, which shows the extent to which the effect of relaxing birth quotas depends on the proportion of eligible couples across provinces. Other model setups are similar as before.

Using the same approach, we estimate the extent to which the proportion of eligible couples influences the dynamic fertility effects of relaxing birth quotas. Our results show that the positive impact of relaxing birth quotas on second-child births increases with the proportion of eligible couples, and the negative impact of relaxing birth quotas on first-child births decreases with the proportion of eligible couples. These findings further support that relaxing birth quotas encourages some couples to have a second child but crowds out other couples' first-child births. Additional evidence suggests that the positive impact of relaxing birth quotas on child-bearing costs increases with the proportion of eligible couples. More details are provided in Figures 7 and 8.

[31] It is important to acknowledge a concern raised in our analyses thus far – they may not fully elucidate the substantial decline in fertility. Even if first births decrease, this decline cannot surpass the increase in second births; otherwise, the price effects that drive a reduction in first births would be absent. As discussed in Section 2, the price effects manifest in the current period due to the expected uptick in second-child births. This does not necessarily imply that the decline in first births cannot exceed the rise in second births during the current period.

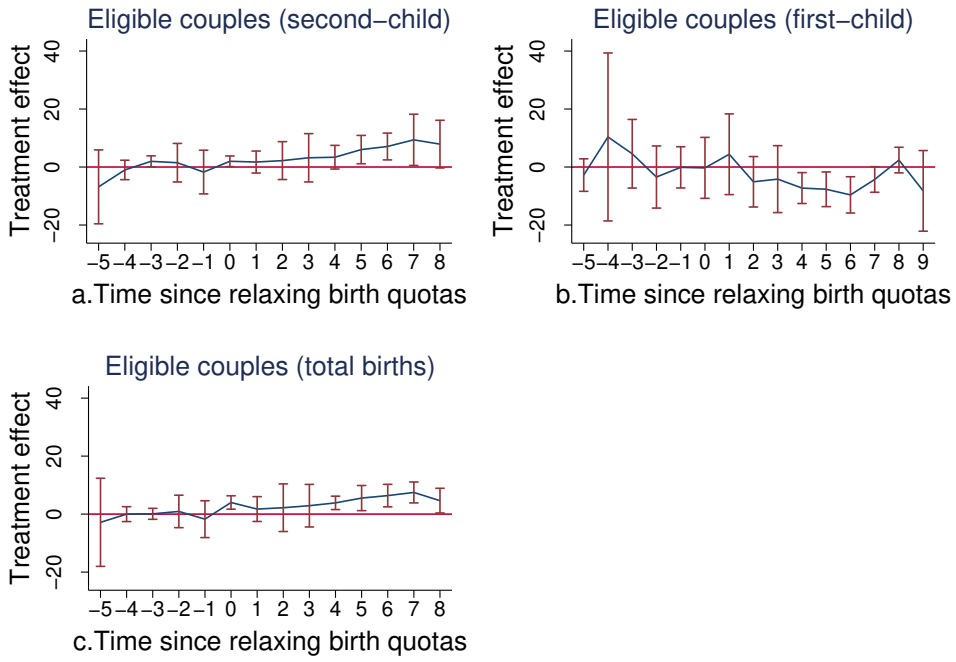


Figure 7: Effects of eligible couples on the effects of relaxing birth quotas.

Notes: This graph shows how the proportion of eligible couples affects the impact of relaxing birth quotas on (a) second-child births, (b) first-child births and (c) total births. In each panel, the horizontal axis represents the values corresponding to the time relative to the (de facto) relaxation of birth quotas, which are negative in the pretreatment periods, and the vertical axis represents the values corresponding to the change in the outcome of interest. The points represent the estimated treatment effects in the treatment periods and placebo effects in the pretreatment periods, and the error bars represent 95% confidence intervals of the estimated effects and placebos. Eligible couples are defined as couples in which at least one of the marital partners is an only child. Both first-child, second-child and total births variables are in the logarithm form. The panel data on first-child,second-child and total births from November 2014 to October 2015 are constructed using the 2015 census. To draw these figures, we followed the estimation strategy proposed by **?**.

5.7 Does relaxing birth quotas affect total births?

Thus far, we have shown that relaxing birth quotas increases second-child births but decreases or delays first-child births, the latter of which is probably caused by increased child-rearing costs in the short term. To determine whether decreased first-child births offset the increased second-child births, we further investigate the impact of relaxing birth quotas on total births. Using the same empirical approach, we find that relaxing birth quotas substantially increases total births, suggesting that the decrease in first-child births is partially offset by the increased second-child after relaxation of birth quotas. This finding provides further support that relaxing birth quotas leads to higher child-rearing costs in the short term. See **Figure 5c** for details.

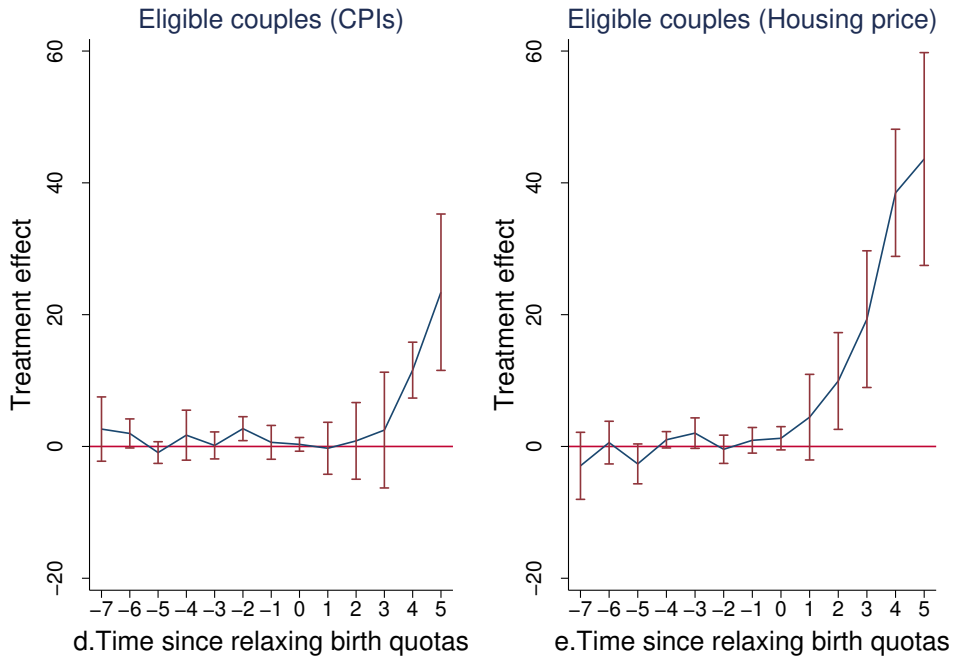


Figure 8: Effects of eligible couples on the effects of relaxing birth quotas (cont.).

Notes: This graph shows how the proportion of eligible couples affects the impact of relaxing birth quotas on (a) CPIs and (b) housing price indices. In each panel, the horizontal axis represents the values corresponding to the time relative to the (de facto) relaxation of birth quotas, which are negative in the pretreatment periods, and the vertical axis represents the values corresponding to the change in the outcome of interest. The points represent the estimated treatment effects in the treatment periods and placebo effects in the pretreatment periods, and the error bars represent 95% confidence intervals of the estimated effects and placebos. Eligible couples are defined as couples in which at least one of the marital partners is an only child. The panel data on CPIs and housing price indices from August 2014 to July 2015 come from China's National Bureau of Statistics. To draw these figures, we followed the estimation strategy proposed by **?**.

5.8 A competing hypothesis: Marriage market

In this subsection, we investigate a competing mechanism, the marriage market, for asymmetric effects of relaxing birth quotas on first-child and second-child births. To estimate the impact on marriage outcomes, we construct monthly data on the number of marriages across provinces. The model specification is as follows:[32]

$$Marriage_{s,t_{birth}} = \boldsymbol{\alpha}^{'} I_s + \boldsymbol{\beta}^{'} I_{t_{birth}} + \gamma_6 \left(I_{s,t_{birth} \geq t_{policy+r}} \right) + \varepsilon_{s,t_{birth}}, \tag{19}$$

[32]China's Ministry of Civil Affairs provides the number of marriages at the quarterly level only in each province. Thus, we use the cubic spline interpolation method to calculate the number of marriages at the monthly level from October 2013 to September 2014.

where $Marriage_{s,t_{birth}}$ is the number of marriages in province s in month t_{birth}, I_s is a vector of province fixed effects, $I_{t_{birth}}$ is a vector of time fixed effects and $I_{s,t_{birth} \geq t_{policy+r}}$ is an indicator for an observation r months after the relaxation of the OCP or an observation after the de facto policy adoption. In this model, we assume that $r = 0$ (i.e., relaxation of birth quotas affects local marriage markets immediately) because conceptually, marrying someone is much less time-consuming than giving birth to a child and because forward-looking individuals would expect changes in the returns of getting married to occur immediately upon relaxing birth quotas. The error term is $\varepsilon_{s,t_{birth}}$, $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are vectors of coefficients to be estimated and γ_6 is the coefficient of interest.

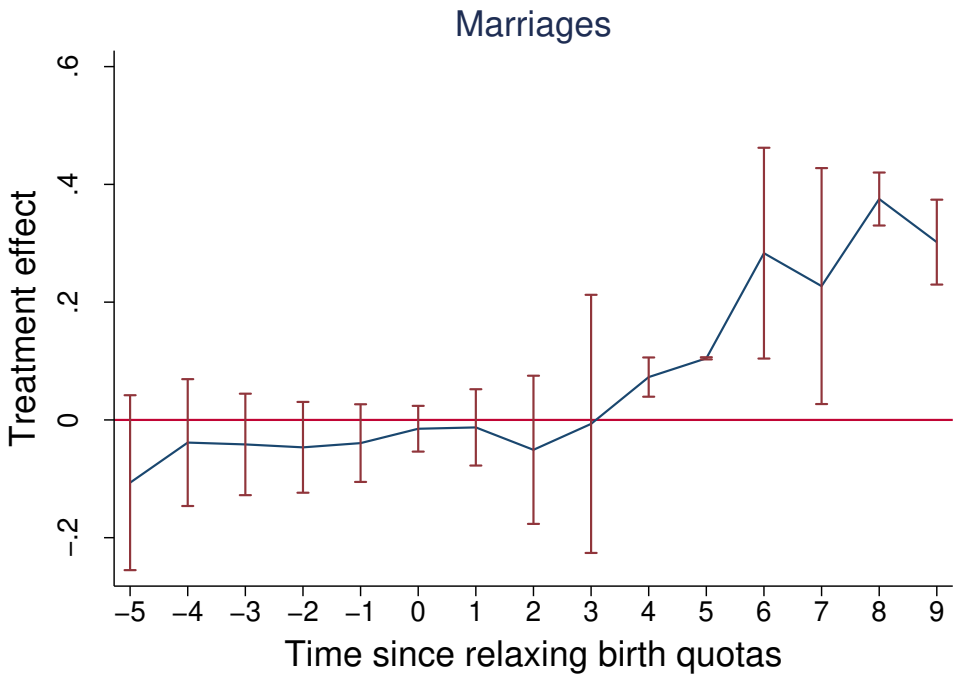


Figure 9: Impact of relaxing birth quotas on local marriages

Notes: This graph shows the impact of relaxing birth quotas on total marriages. In each panel, the horizontal axis represents the values corresponding to the time relative to the (de facto) relaxation of birth quotas, which are negative in the pretreatment periods, and the vertical axis represents the values corresponding to the change in the outcome of interest. The points represent the estimated treatment effects in the treatment periods and placebo effects in the pretreatment periods, and the error bars represent 95% confidence intervals of the estimated effects and placebos. Total marriages variables are in the logarithm form. Since China's Ministry of Civil Affairs only provides the number of marriages at the quarterly level in each province, we resort to the cubic spline interpolation method to calculate the number of marriages at the monthly level from October 2013 to September 2014. To draw these figures, we followed the estimation strategy proposed by de Chaisemartin and D'Haultfoeuille (2020).

We use the same approach to estimate the dynamic impacts of relaxing birth quotas on local

marriages.[33] Our main results are reported in **Figure 9**. We find that relaxing birth quotas significantly increases the number of marriages, echoing a previous study by Huang, Pan and Zhou (2023), who find that the OCP induced a significantly lower marriage rate. Specifically, we find that the number of marriages increases by 7.4% in the fourth month of relaxing birth quotas and increases to 35.1% in the ninth month. A simple average of the instantaneous and dynamic effects shows that relaxing birth quotas leads to an increase of about 13.4% in the number of local marriages. As a result, the positive impact of relaxing birth quotas on the marriage market strengthens our previous argument that higher child-rearing costs are likely an underlying mechanism for the declining first-child births after relaxing birth quotas.

6 Conclusion

This study investigates the economic effects of relaxing birth quotas. An extended Barro-Becker model predicts asymmetric effects of relaxing birth quotas on first-child and second-child births. Using a DID framework to examine a change from a one-child to a two-child policy implemented by different Chinese local governments in different months, we find that increasing the birth quota substantially increases second-child births. This impact is concentrated among couples covered by pensions rather than couples without access to pensions, which supports the intergenerational altruism motive rather than the old-age security motive for fertility. Moreover, we provide some of the first evidence that relaxing birth quotas leads to fewer first-child births, and further analyses suggest that higher costs of raising a child are a likely underlying mechanism in this phenomenon. Overall, the findings suggest that policies designed to boost fertility rates should prioritize reducing the child-rearing costs borne by prospective parents. More importantly, the observed spill-over effects indicate that influencing fertility through birth quotas is not a one-size-fits-all approach.

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[33] Using a two-way fixed effects model, we find that the coefficient of interest is -0.08. Futher investigation reveals an obvious negative weight problem in the data. Specifically, under the common-trends assumption, $\hat{\gamma}_6$ estimates a weighted sum of 229 ATTs. A total of 169 ATTs receive a positive weight, and 60 receive a negative weight.

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Appendix A

A.1 Proof of Lemma 1

To solve the optimization problem, we substitute out for $U(\alpha, t+1)$ (and $U(\alpha, t+2)$, etc.) in equation (1) and rewrite the utility function for household α in generation $t = 0$ as

$$U(\alpha, 0) = \sum_{t=0}^{\infty} \alpha^t \cdot [N(\alpha, t)]^{1-\epsilon} \cdot [c(\alpha, t)]^{\sigma}/\sigma. \tag{A.1}$$

Moreover, the dynastic budget constraint, which equates the present value of all resources to the present value of all expenditures, is expressed as

$$k(\alpha, 0) + \sum_{t=0}^{\infty} d(t) N(\alpha, t) w(t) = \sum_{t=0}^{\infty} d(t)[N(\alpha, t) P(t) c(t) + N(\alpha, t+1)\beta(\alpha, t)], \tag{A.2}$$

where $d(t) = \prod_{t'=0}^{t} [1 + r(t')]^{-1}$. Thus, the optimization problem can be expressed as choosing consumption $c(\alpha, t)$ and number of children $n(\alpha, t)$ to maximize utility $U(\alpha, 0)$ subject to the dynastic budget constraint. The corresponding Lagrange function is

$$\begin{aligned} L = &\sum_{t=0}^{\infty} \alpha^t [N(\alpha, t)]^{1-\epsilon} [c(\alpha, t)]^{\sigma}/\sigma + \lambda(\alpha, 0)[k(\alpha, 0) + \sum_{t=0}^{\infty} d(t) N(\alpha, t) w(t) \\ &- \sum_{t=0}^{\infty} d(t) N(\alpha, t) P(t) c(\alpha, t) - \sum_{t=0}^{\infty} d(t) N(\alpha, t+1)\beta(\alpha, t)], \end{aligned} \tag{A.3}$$

where $\lambda(\alpha, 0)$ is the Lagrange multiplier. The first-order conditions related to $c(\alpha, t)$ and $N(\alpha, t)$ are

$$\alpha^t [N(\alpha, t)]^{1-\epsilon} [c(\alpha, t)]^{\sigma-1} = \lambda(\alpha, 0) d(t) N(\alpha, t) P(t), \quad (t = 0, 1, \cdots) \tag{A.4}$$

and

$$\frac{(1-\epsilon)\alpha^t[c(\alpha,t)]^\sigma}{\sigma[N(\alpha,t)]^\epsilon} = \lambda(\alpha,0)[d(t)P(t)c(\alpha,t) + d(t-1)\beta(\alpha,t-1) - d(t)w(t)], \quad (t=1,2,\cdots) \tag{A.5}$$

respectively. By iterating (A.4) forward one period, we obtain an equation of $t+1$. Combing the resulting equation and (A.4) yields

$$\left[\frac{c(\alpha,t+1)}{c(\alpha,t)}\right]^{1-\sigma} = \frac{\alpha[1+r(t+1)]}{[n(\alpha,t)]^\epsilon} \cdot \frac{P(t)}{P(t+1)}, \quad (t=0,1,\cdots), \tag{A.6}$$

Further, using the form of $\lambda(\alpha,0)$ from (A.4), we can rewrite (A.5) as (3). It shows that at period of $t \geq 1$, consumption per member of household α is independent of initial capital stock. Moreover, as in **?** and **?**, this extended model also implies that children are a net financial burden to altruistic parents; that is, the costs of raising a child would exceed that child's lifetime earnings. Finally, combining (A.6) and (3) yields the optimal number of children born such that

$$[n(\alpha,t)]^\epsilon = \alpha[1+r(t+1)]\left[\frac{P(t)}{P(t+1)}\right]^\sigma \left\{\frac{\beta(\alpha,t-1)[1+r(t)]-w(t)}{\beta(\alpha,t)[1+r(t+1)]-w(t+1)}\right\}^{1-\sigma}, \quad (t=1,2,\cdots). \tag{A.7}$$

A.2 Decentralized Equilibrium

In equilibrium, households maximize utility, firms maximize profits, and all markets clear. In generation $t \geq 1$, the total number of adults and children are $\hat{N}(t) = qN(\alpha^h,t) + (1-q)N(\alpha^l,t)$ and $\hat{N}(t+1) = qN(\alpha^h,t+1) + (1-q)N(\alpha^l,t+1)$, respectively. Therefore, the final good market-clearing condition is

$$C(t) + \mu\hat{N}(t+1) = Y(t), \tag{A.8}$$

where $C(t) \equiv qN(\alpha^h,t)c(\alpha^h,t) + (1-q)N(\alpha^l,t)c(\alpha^l,t)$ is the total consumption of final good, the left-hand side (LHS) and the right-hand side (RHS) of (A.8) are the aggregate demand and supply of final good, respectively. The labor market-clearing condition is $L(t) = \hat{N}(t)$, where the LHS is the total labor demand and the RHS denotes the total labor supply, namely the total number of adults. The total capital stocks is defined as $K(t) = qN(\alpha^h,t)k(\alpha^h,t) + (1-q)N(\alpha^l,t)k(\alpha^l,t)$, and the capital market also clears.

A.3 Long-term effects of relaxing birth quotas

In this subsection, we examine the long-run impact of the population control policy on birth rates. We first show that a change in the population control policy in this model induces a long

transitional dynamic and there only exists an asymptotic steady state equilibrium. To see this, we need to characterize the change of the total birth rate. Total number of adults in generation $t+2$ is

$$\hat{N}(s+2) = q \cdot n(\alpha^h, 0) \cdot (\tilde{n})^{s-1} \cdot (n^*)^2 + (1-q) \cdot n(\alpha^l, 0) \cdot (\tilde{n})^{s-1} \cdot n(\alpha^l, s) \cdot n(\alpha^l, s+1). \tag{A.9}$$

$n(\alpha^l, s+1)$ could deviate from $n(\alpha^l, s)$ because the birth rates of both types of households are unequal since the period of s, which implying that the total birth rate is changing over time. As a result, the prices of final goods and interest rates are continuing to change, causing households to reconsider their fertility decisions. Recall the total number of adults in generation $s+1$ in (12). We then obtain the total birth rate in generation $s+1$ such that

$$B(s+1) = \frac{\hat{N}(s+2)}{\hat{N}(s+1)} = \frac{q \cdot n(\alpha^h, 0) \cdot (n^*)^2 + (1-q) \cdot n(\alpha^l, 0) \cdot n(\alpha^l, s) \cdot n(\alpha^l, s+1)}{q \cdot n(\alpha^h, 0) \cdot n^* + (1-q) \cdot n(\alpha^l, 0) \cdot n(\alpha^l, s)}. \tag{A.10}$$

Following the same logic, we can conclude that the total birth rate at period of $s+j$ is

$$B(s+j) = \frac{\hat{N}(s+j+1)}{\hat{N}(s+j)} = \frac{q \cdot n(\alpha^h, 0) \cdot n^* + (1-q) \cdot n(\alpha^l, 0) \cdot \frac{n(\alpha^l, s)}{n^*} \cdots \frac{n(\alpha^l, s+j)}{n^*} \cdot n(\alpha^l, s+j+1)}{q \cdot n(\alpha^h, 0) + (1-q) \cdot n(\alpha^l, 0) \cdot \frac{n(\alpha^l, s)}{n^*} \cdots \frac{n(\alpha^l, s+j)}{n^*}}, \quad j > 0, \tag{A.11}$$

where we implicitly assume that n^* is sufficiently large so that $n^* > \max\{n(\alpha^l, s), \cdots, n(\alpha^l, s+j+1)\}$ is satisfied. As time goes by, the total birth rate is growing and approaches n^* for $j \to +\infty$, by applying the L'Hôpital's rule. During this transitional path towards an asymptotic steady state equilibrium, the family structure is varying. Specifically, the fraction of the α^h-type household is gradually rising and reaches one for $j \to +\infty$, while the fraction of the α^l-type household is declining and becomes negligible, approaching zero, for $j \to +\infty$.

Proposition 2. *Starting from the next period after the government relaxes birth quotas, the economy starts its transitional dynamics. Along the transitional path, the total birth rate is rising, the size of α^h-type household is increasing while the size of the α^l-type household is decreasing.*

Proof. Proven in text. □

Appendix B

Table B.1: Impact of relaxing birth quotas on second-child births using the two-way fixed effects model

Variables	(1) Second birth	(2) Native	(3) Migrant	(4) Old	(5) Young	(6) College	(7) No college	(8) Pension	(9) No pension
Treatment effect	−0.007	0.016	0.209	−0.113	0.220	0.030	0.021	−0.001	0.030
	(0.102)	(0.110)	(0.225)	(0.095)	(0.176)	(0.129)	(0.103)	(0.112)	(0.121)
Observations	371	371	174	370	353	291	362	365	340
Adjusted R^2	0.860	0.879	0.653	0.804	0.736	0.455	0.875	0.848	0.632

Notes: This table reports the impact of relaxing birth quotas on second-child births using the two-way fixed effects model. Birth variables are in the logarithm form. Column 1 shows the impact of relaxing birth quotas on second-child births. Columns 2 and 3 show the impact of relaxing birth quotas by migration status. Columns 4 and 5 show the impact of relaxing birth quotas by childbearing age. Columns 6 and 7 show the impact of relaxing birth quotas by education degree. Columns 8 and 9 show the impact of relaxing birth quotas by access to pension. Standard errors are clustered at the provincial level and are in parentheses. * p < 0.1, ** p < 0.05, *** p < 0.01.

Appendix C

C.1 Recoding the timing of relaxing the OCP

As already mentioned, using monthly birth data for estimation implies that assumptions are required for the minimum time it takes for the relaxation of the OCP to noticeably affect local second-child (first-child) births. There are two main reasons. First, without recoding the timing of relaxing the OCP, we cannot estimate the effect of relaxing the OCP on childbearing decisions in a DID setup. This is because the census data provide monthly births from November 2014 to October 2015, while most provinces relaxed the OCP before November 2014. In other words, the timing of monthly births mostly does not overlap with the timing of relaxing the OCP. As a result, the untreated provinces will be too few to estimate the effect of relaxing the OCP on childbearing decisions in a DID setup. Second, even if we have early data on monthly births that overlap the timing of relaxing the OCP, estimating the effect of relaxing the OCP on childbearing decisions without recoding the timing of relaxing the OCP remains problematic because its effect on childbearing decisions would be largely contaminated by biological effects due to the fertility process. We further discuss how to recode the timing of relaxing the OCP as follows.

C.2 Time to obtain a birth permit for a second child (1 month)

Under this policy, prospective parents are required to apply for a birth permit in advance of having a second child. According to the official procedure and documents for granting birth permits, the average time for obtaining a birth permit from the local authorities ranges from 15 days to 45 days. Thus, we assume that it takes 1 month on average for prospective parents to

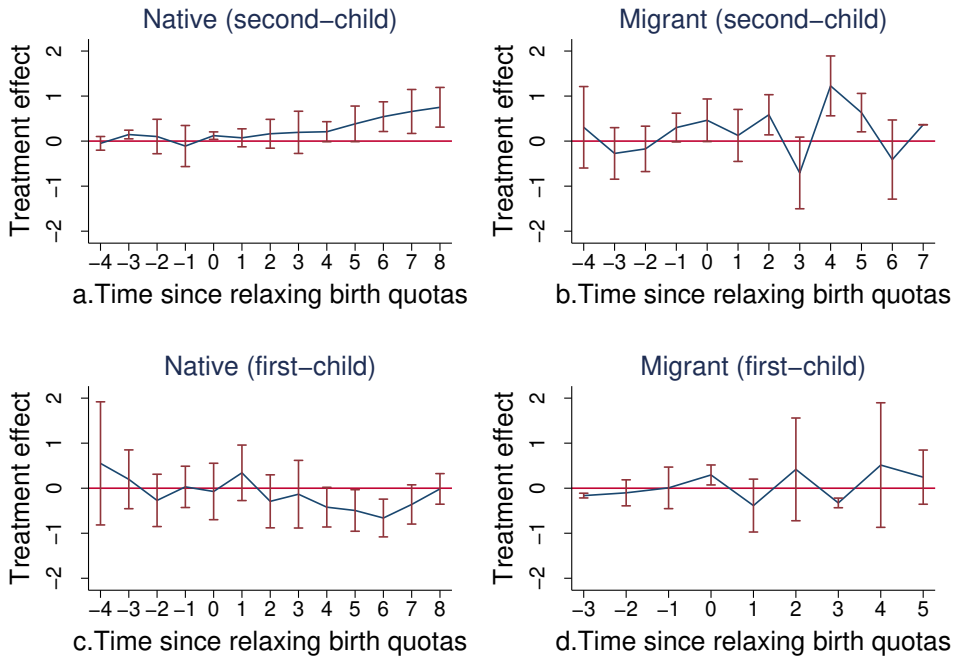


Figure B.1: Migration status and the effects of relaxing birth quotas on second-child and first-child births

Notes: This graph shows the impact of relaxing birth quotas on second-child births of household (a) registered in the same province and (b) registered in a different province, and the impact of relaxing birth quotas on first-child births of households (c) registered in the same province and (d) registered in other province. In each panel, the horizontal axis represents the values corresponding to the time relative to the (de facto) relaxation of birth quotas, which are negative in the pretreatment periods, and the vertical axis represents the values corresponding to the change in the outcome of interest. The points represent the estimated treatment effects in the treatment periods and placebo effects in the pretreatment periods, and the error bars represent 95% confidence intervals of the estimated effects and placebos. We define migrants as those individuals registered in a different province. Both first-child and second-child birth variables are in the logarithm form. The panel data on first-child and second-child births from November 2014 to October 2015 are constructed using the 2015 census. To draw these figures, we followed the estimation strategy proposed by de Chaisemartin and D'Haultfoeuille (2020).

obtain a birth permit from local authorities in order to be legally allowed to have a second child.

C.3 Time to pregnancy for a second (first) child (1 month)

Based on previous studies in human reproduction (Olsen, Juul and Basso (1998); Gnoth, Godehardt, Godehardt, Frank-Herrmann and Freundl (2003); Eisenberg, Thoma, Li and McLain (2021)), we assume the minimum time to pregnancy for a second (first) child is 1 month.

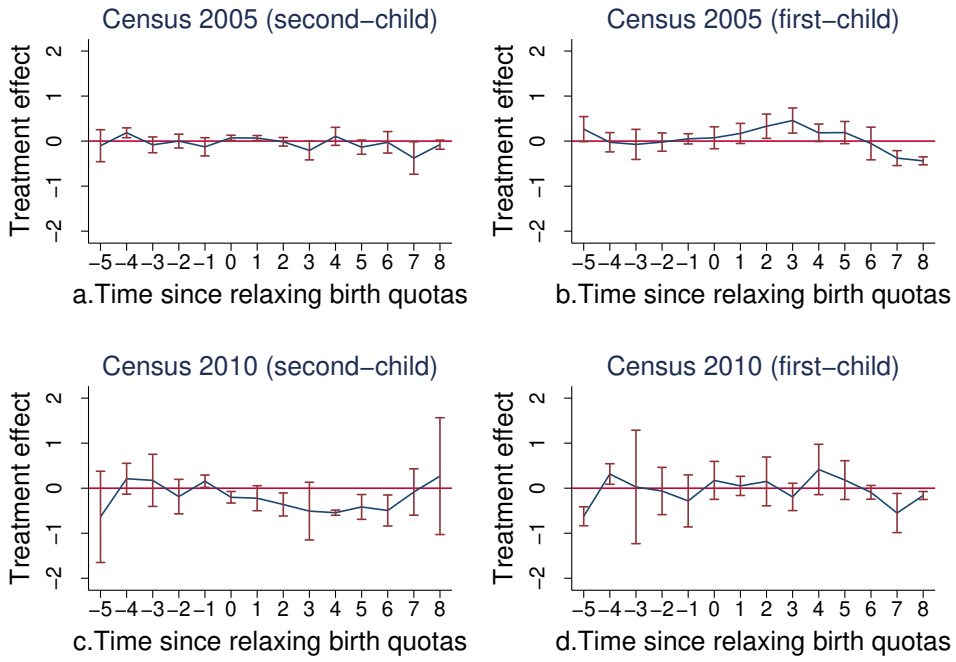


Figure B.2: Placebo analysis on the impact of relaxing birth quotas using censuses 2005 and 2010

Notes: This graph shows the impact of relaxing birth quotas on (a) counterfactual second-child births using census 2005, (b) counterfactual first-child births using census 2005, (c) counterfactual second-child births using census 2010, and (d) counterfactual first-child births using census 2010. In each panel, the horizontal axis represents the values corresponding to the time relative to the (de facto) relaxation of birth quotas, which are negative in the pretreatment periods, and the vertical axis represents the values corresponding to the change in the outcome of interest. The points represent the estimated treatment effects in the treatment periods and placebo effects in the pretreatment periods, and the error bars represent 95% confidence intervals of the estimated effects and placebos. Both first-child and second-child birth variables are in the logarithm form. The panel data on first-child and second-child births from November 2004 to October 2005 are constructed using the 2005 census, and the panel data on first-child and second-child births from November 2009 to October 2010 are constructed using the 2010 census. To draw these figures, we followed the estimation strategy proposed by de Chaisemartin and D'Haultfoeuille (2020).

C.4 Duration of pregnancy (10 months)

A full-term pregnancy generally lasts between 9 and 10 months. Moreover, our birth data are at the monthly level. For example, birth data in December consists of all births that occurred from December 1 to December 31. Thus, we assume that a pregnancy lasts 10 months on average to capture the impact of relaxing the OCP on local births.

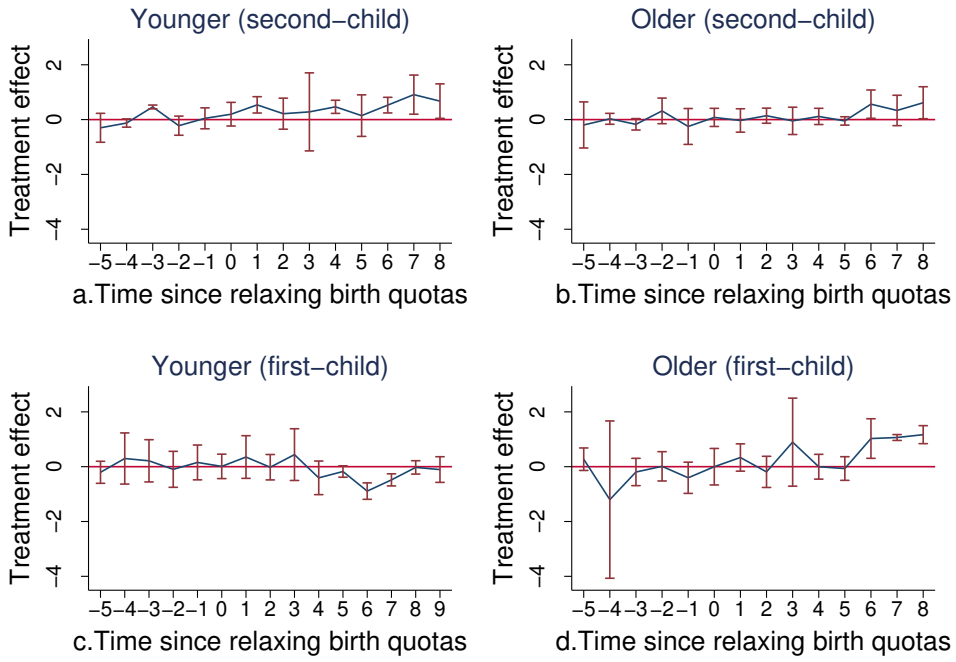


Figure B.3: Childbearing age and effects of relaxing the OCP on second-child and first-child births.

Notes: This graph shows the impact of relaxing birth quotas on second-child births of (a) couples of relatively early childbearing age (below the median value of the childbearing age distribution), (b) couples of relatively late childbearing age (above the median value of the childbearing age distribution) and first child births of (c) couples of relatively early childbearing age and (d) couples of relatively late childbearing age. In each panel, the horizontal axis represents the values corresponding to the time relative to the (de facto) relaxation of OCP, which are negative in the pretreatment periods, and the vertical axis represents the values corresponding to the change in the outcome of interest. The points represent the estimated treatment effects in the treatment periods and placebo effects in the pretreatment periods, and the error bars represent 95% confidence intervals of the estimated effects and placebos. Relatively early childbearing age are defined as women below the median value of the childbearing age distribution. Second-child and first-child births variables are in the logarithm form. The panel data on second-child and first-child births from November 2014 to October 2015 are constructed using the 2015 census. To draw these figures, we followed the estimation strategy proposed by de Chaisemartin and D'Haultfoeuille (2020).

C.5 Anticipation effect

Some may worry about an anticipation effect (i.e. couples in the late-reform area delay their fertility decisions until the reform date). However, this is not likely to be a major concern in our context. First, it remained illegal at that time for ineligible couples to have a second child, and couples violating the OCP were forced to pay a fine and were subject to a variety of other punishments, including the seizure of property and forced dismissal from government employment.

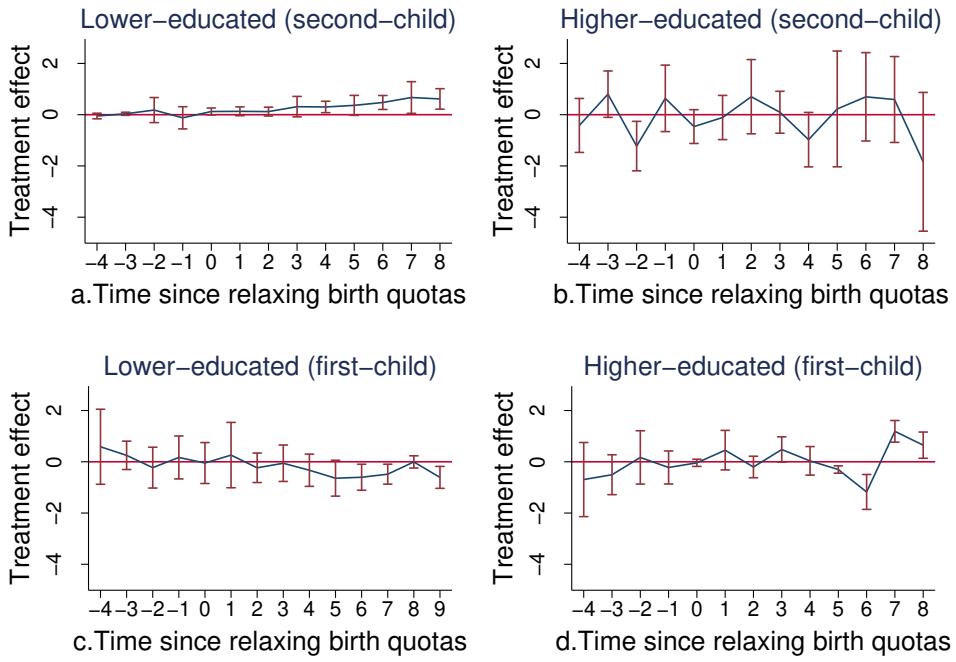


Figure B.4: Education and effects of relaxing birth quotas on second-child and first-child births.

Notes: This graph shows the impact of relaxing birth quotas on second-child births of (a) relatively lower-educated women, (b) relatively higher-educated women and first child births of (c) relatively lower-educated women and (d) relatively higher-educated women. In each panel, the horizontal axis represents the values corresponding to the time relative to the (de facto) relaxation of birth quotas, which are negative in the pretreatment periods, and the vertical axis represents the values corresponding to the change in the outcome of interest. The points represent the estimated treatment effects in the treatment periods and placebo effects in the pretreatment periods, and the error bars represent 95% confidence intervals of the estimated effects and placebos. Relatively higher-educated women are defined as women with a college degree or more. Second-child and first-child births variables are in the logarithm form. The panel data on second-child and first-child births from November 2014 to October 2015 are constructed using the 2015 census. To draw these figures, we followed the estimation strategy proposed by de Chaisemartin and D'Haultfoeuille (2020).

Second, there was no clear guideline on how quickly the reform should be implemented across the country. In fact, on November 17, 2013, Peian Wang, vice Minister of the National Health and Family Planning Commission, said in an official interview that there would be no unified timetable for implementation of the policy in the country. All provinces (autonomous regions and municipalities) would decide on their own time arrangements according their specific condition (http://www.nhc.gov.cn/xcs/s3574/201311/24fb42100d9945a99de21712f3f38f55.shtml). On December 7, 2013, Wenzhuang Yang, director of the family planning instruction department of China's National Health and Family Planning Commission said during an interview with China

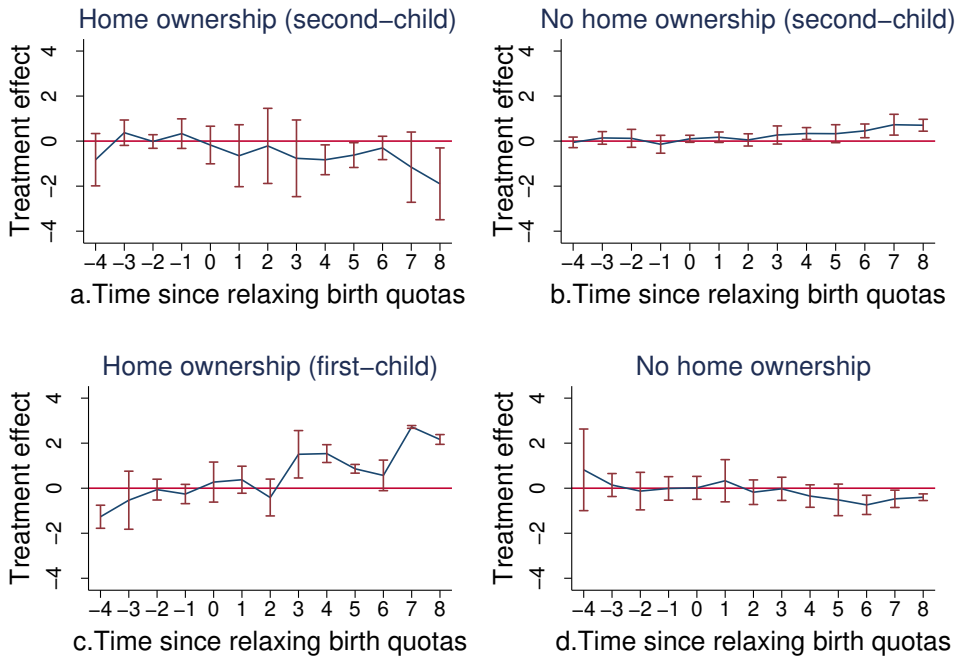


Figure B.5: Homeownership and the effects of relaxing birth quotas on second-child and first-child births.

Notes: This graph shows the impact of relaxing birth quotas on second-child births of household (a) with homeownership and (b) without homeownership and the impact of relaxing the OCP on first-child births of households (c) with homeownership and (d) without homeownership. In each panel, the horizontal axis represents the values corresponding to the time relative to the (de facto) relaxation of birth quotas, which are negative in the pretreatment periods, and the vertical axis represents the values corresponding to the change in the outcome of interest. The points represent the estimated treatment effects in the treatment periods and placebo effects in the pretreatment periods, and the error bars represent 95% confidence intervals of the estimated effects and placebos. Households with homeownership are defined as households that have purchased the house. Both first-child and second-child birth variables are in the logarithm form. The panel data on first-child and second-child births from November 2014 to October 2015 are constructed using the 2015 census. To draw these figures, we followed the estimation strategy proposed by de Chaisemartin and D'Haultfoeuille (2020).

Central Television (CCTV) that the policy would probably be implemented early in the next year after local administrations finished preparations and local legislatures amended the regulation. According to Yang, it was up to local administrations to arrange the exact timetable for implementing the policy, according to specific circumstances, but there should not be a major time gap in introducing the new rule from region to region. Yang also asked willing couples to carefully plan the timing of their second baby, stating that "there is no need to rush as the policy will be a long-standing one on the mainland".

Third, even if some sort of anticipation effect exists and it induces some couples to prepare

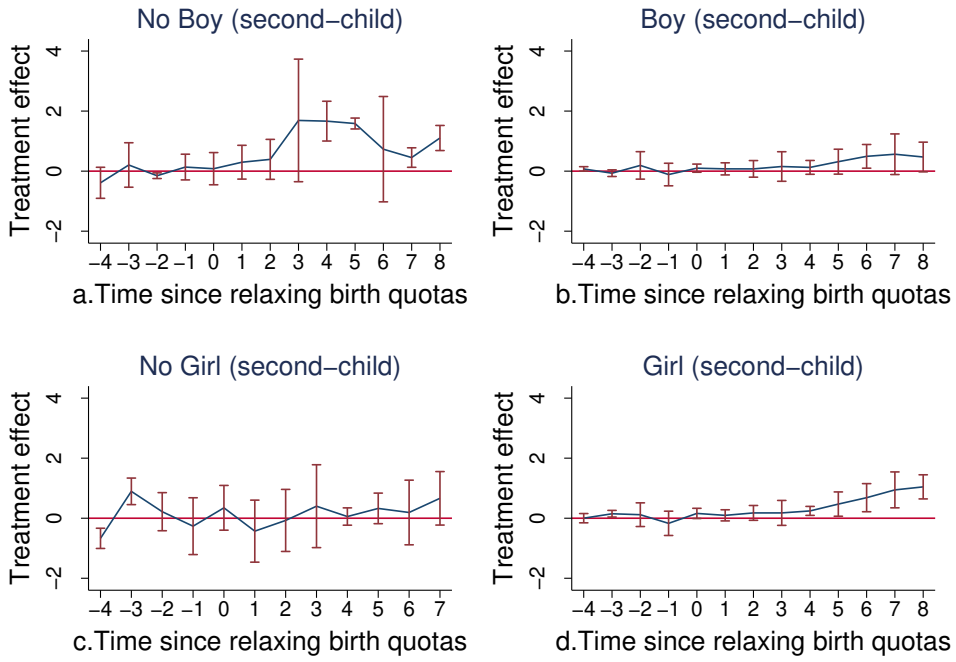


Figure B.6: Gender of children and the effects of relaxing birth quotas on second-child.

Notes: This graph shows the impact of relaxing birth quotas on second-child births of household (a) with no boy, (b) with at least a boy, (c) with no girl and (d) with at least a girl at the time of the survey. In each panel, the horizontal axis represents the values corresponding to the time relative to the (de facto) relaxation of birth quotas, which are negative in the pretreatment periods, and the vertical axis represents the values corresponding to the change in the outcome of interest. The points represent the estimated treatment effects in the treatment periods and placebo effects in the pretreatment periods, and the error bars represent 95% confidence intervals of the estimated effects and placebos. Second-child birth variables are in the logarithm form. The panel data on second-child births from November 2014 to October 2015 are constructed using the 2015 census. To draw these figures, we followed the estimation strategy proposed by de Chaisemartin and D'Haultfoeuille (2020).

for a second child (or delay fertility decisions of a first child) before local government's relaxation of the OCP, then our estimate would provide a lower bound of the effect of relaxing the OCP on second-child (or first-child) births (Malani and Reif (2015)).

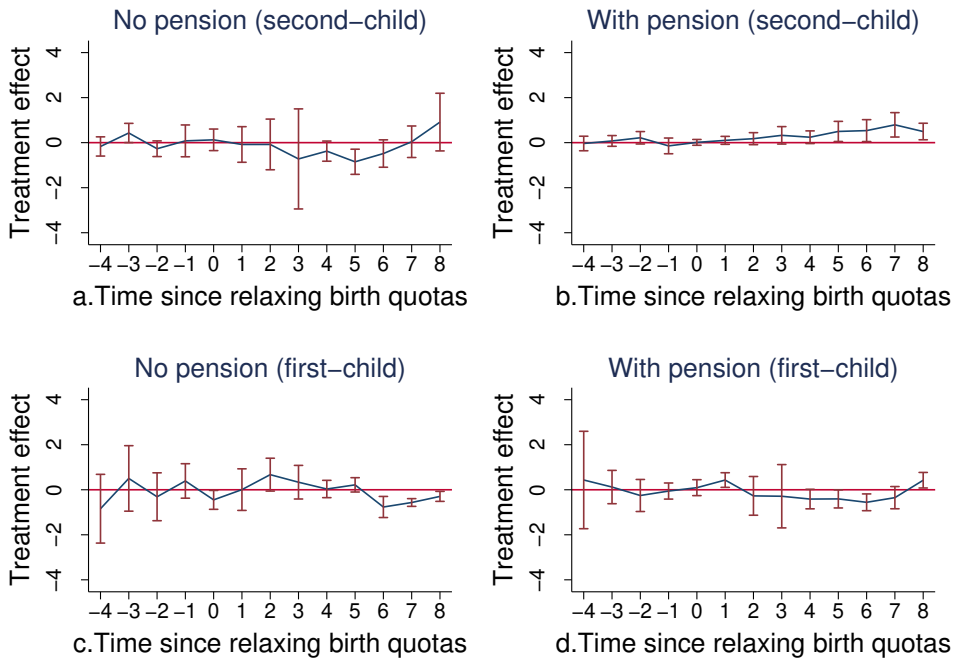


Figure B.7: Access to pension and the effects of relaxing birth quotas on second-child and first-child births.

Notes: This graph shows the impact of relaxing birth quotas on second-child births of household (a) with access to pension and (b) with no access to pension and the impact of relaxing the birth quotas on first-child births of households (c) with access to pension and (d) with no access to pension. In each panel, the horizontal axis represents the values corresponding to the time relative to the (de facto) relaxation of birth quotas, which are negative in the pretreatment periods, and the vertical axis represents the values corresponding to the change in the outcome of interest. The points represent the estimated treatment effects in the treatment periods and placebo effects in the pretreatment periods, and the error bars represent 95% confidence intervals of the estimated effects and placebos. Both first-child and second-child birth variables are in the logarithm form. The panel data on first-child and second-child births from November 2014 to October 2015 are constructed using the 2015 census. To draw these figures, we followed the estimation strategy proposed by de Chaisemartin and D'Haultfoeuille (2020).

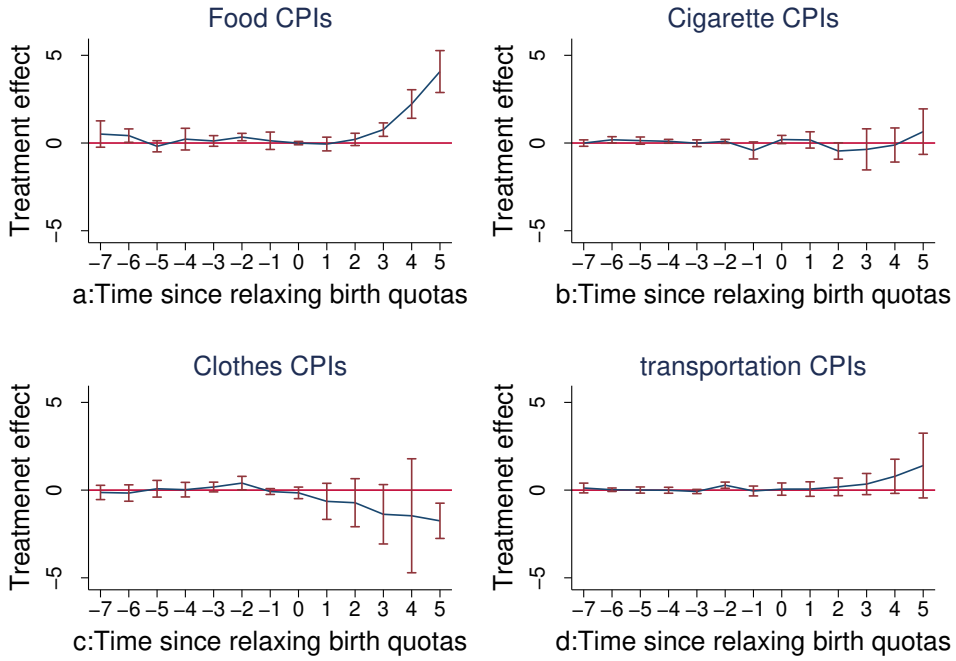


Figure B.8: Effects of relaxing birth quotas on CPIs by expenditure categories

Notes: This graph shows the impact of relaxing birth quotas on CPIs of (a) the food category, (b) the tobacco and alcohol category, (c) the clothing category, and (d) transportation and communications category. In each panel, the horizontal axis represents the values corresponding to the time relative to the (de facto) relaxation of birth quotas, which are negative in the pretreatment periods, and the vertical axis represents the values corresponding to the change in the outcome of interest. The points represent the estimated treatment effects in the treatment periods and placebo effects in the pretreatment periods, and the error bars represent 95% confidence intervals of the estimated effects and placebos. Both first-child and second-child birth variables are in the logarithm form. The panel data on CPIs by expenditure categories from August 2014 to July 2015 come from China's National Bureau of Statistics. To draw these figures, we followed the estimation strategy proposed by de Chaisemartin and D'Haultfoeuille (2020).

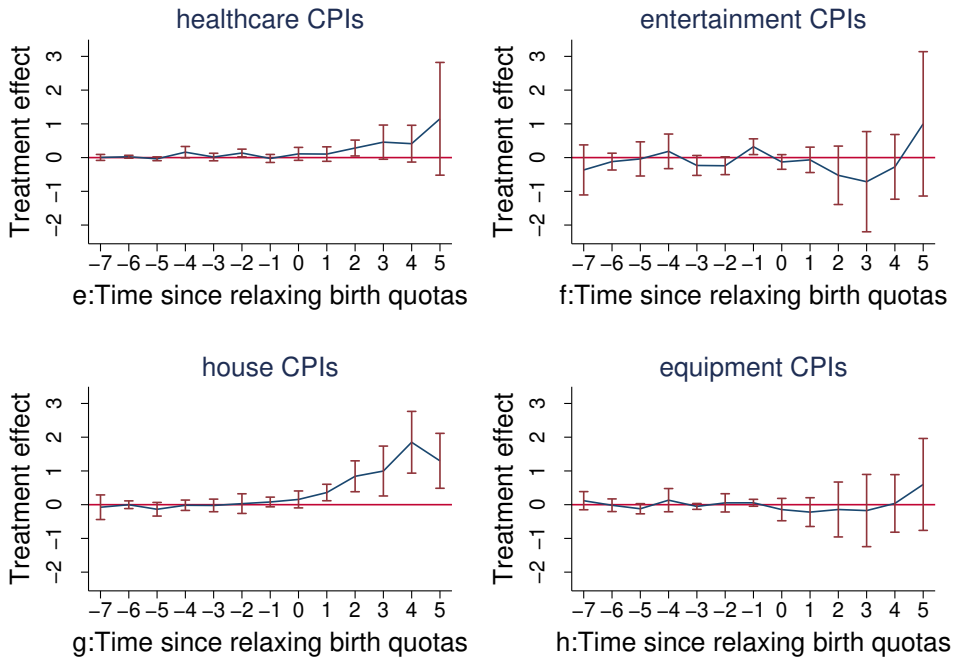


Figure B.9: Effects of relaxing birth quotas on CPIs by expenditure categories (cont.)

Notes: This graph shows the impact of relaxing birth quotas on the CPIs of (e) the health care category, (f) the education, culture, and entertainment category, (g) the housing category, and (h) the household facility, article, and maintenance service category. In each panel, the horizontal axis represents the values corresponding to the time relative to the (de facto) relaxation of birth quotas, which are negative in the pretreatment periods, and the vertical axis represents the values corresponding to the change in the outcome of interest. The points represent the estimated treatment effects in the treatment periods and placebo effects in the pretreatment periods, and the error bars represent 95% confidence intervals of the estimated effects and placebos. Both first-child and second-child birth variables are in the logarithm form. The panel data on CPIs by expenditure categories from August 2014 to July 2015 come from China's National Bureau of Statistics. To draw these figures, we followed the estimation strategy proposed by de Chaisemartin and D'Haultfoeuille (2020).

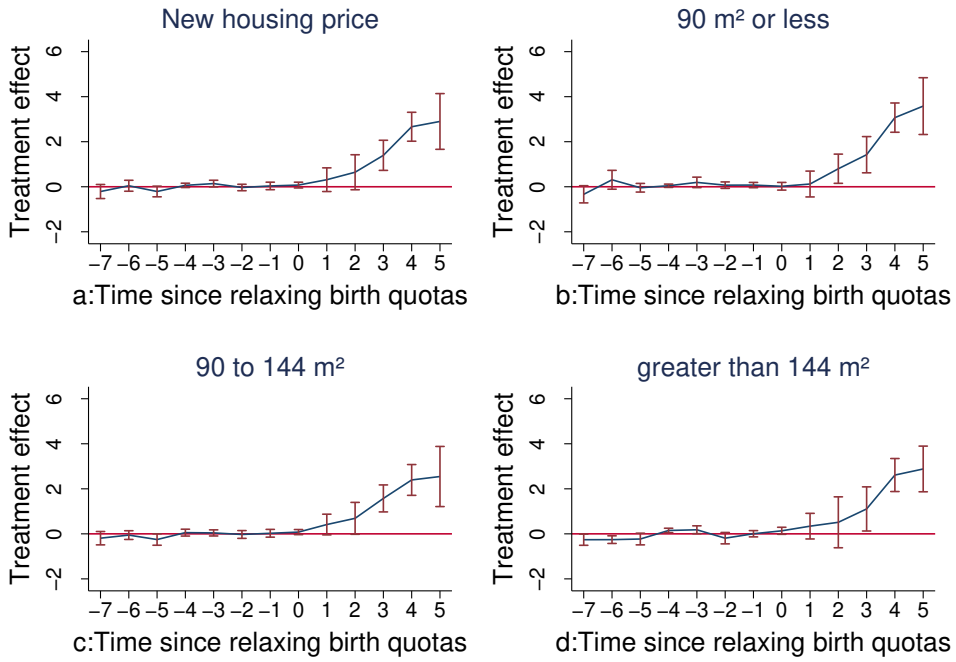


Figure B.10: Heterogeneous effects of relaxing birth quotas on sale prices of newly constructed residential buildings

Notes: This graph shows the impact of relaxing birth quotas on sale prices of (a) newly constructed residential buildings, (b) newly constructed residential buildings with floor area of 90 m^2 or less, (c) newly constructed residential buildings with floor area from 90 to 144 m^2, and (d) newly constructed residential buildings with floor area greater than 144 m^2. In each panel, the horizontal axis represents the values corresponding to the time relative to the (de facto) relaxation of birth quotas, which are negative in the pretreatment periods, and the vertical axis represents the values corresponding to the change in the outcome of interest. The points represent the estimated treatment effects in the treatment periods and placebo effects in the pretreatment periods, and the error bars represent 95% confidence intervals of the estimated effects and placebos. We collect data on sale price indices of residential buildings from August 2014 to July 2015 that are available in 70 large and medium-sized cities across all Chinese provinces except for Tibet, provided by the National Bureau of Statistics of China. To draw these figures, we followed the estimation strategy proposed by de Chaisemartin and D'Haultfoeuille (2020).

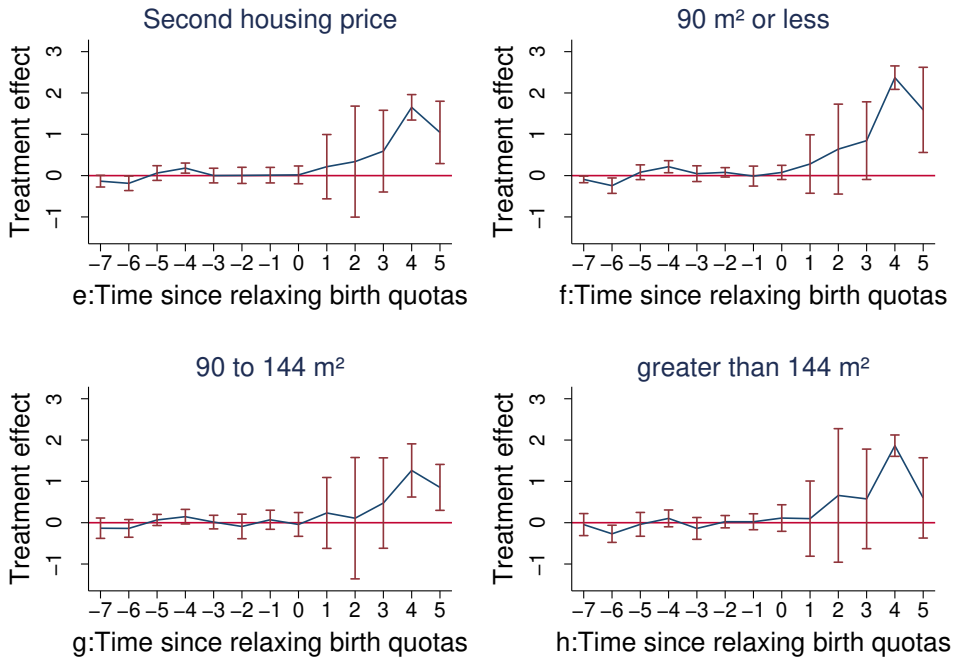


Figure B.11: Heterogeneous effects of relaxing birth quotas on sale prices of second-hand residential buildings.

Notes: This graph shows the impact of relaxing birth quotas on sale prices of (e) second-hand residential buildings, (f) second-hand residential buildings with floor area of 90 m^2 and below, (g) second-hand residential buildings with floor space from 90 to 144 m^2, and (h) second-hand residential buildings with floor space of above 144 m^2. In each panel, the horizontal axis represents the values corresponding to the time relative to the (de facto) relaxation of birth quotas, which are negative in the pretreatment periods, and the vertical axis represents the values corresponding to the change in the outcome of interest. The points represent the estimated treatment effects in the treatment periods and placebo effects in the pretreatment periods, and the error bars represent 95% confidence intervals of the estimated effects and placebos. We collect data on sale price indices of residential buildings from August 2014 to July 2015 that are available in 70 large and medium-sized cities across all Chinese provinces except for Tibet, provided by the National Bureau of Statistics of China. To draw these figures, we followed the estimation strategy proposed by de Chaisemartin and D'Haultfoeuille (2020).

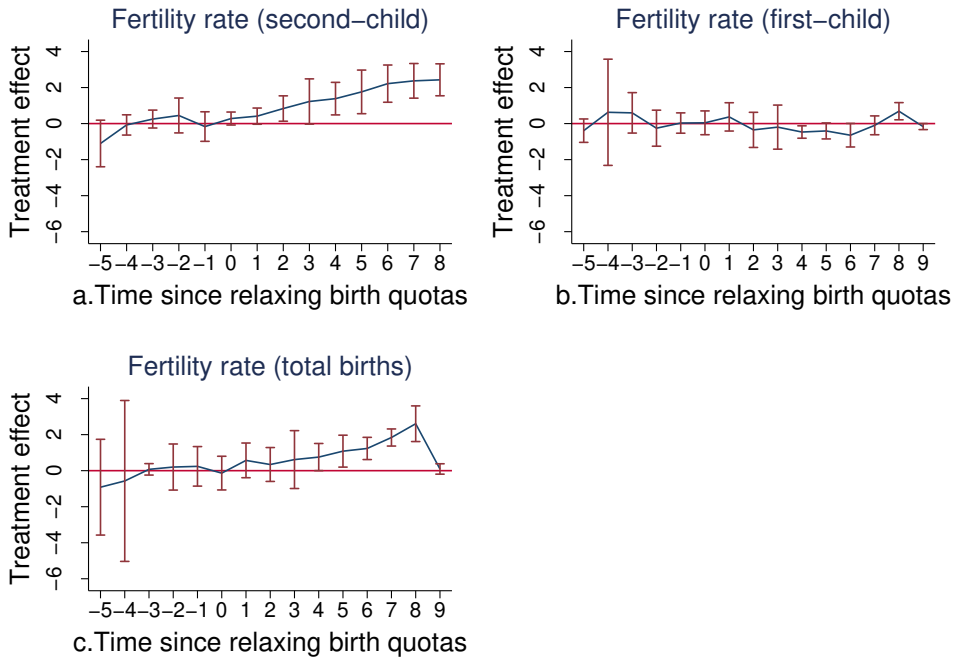


Figure B.12: Impact of relaxing birth quotas on second-child, first-child and total births fertility rates

Notes: This graph shows the impact of relaxing birth quotas on (a) second-child fertility rate (i.e. the number of births divided by the number of women of aged between 15 and 50 years old), (b) first-child and (c) total births fertility rate. In each panel, the horizontal axis represents the values corresponding to the time relative to the (de facto) relaxation of birth quotas, which are negative in the pretreatment periods, and the vertical axis represents the values corresponding to the change in the outcome of interest. The points represent the estimated treatment effects in the treatment periods and placebo effects in the pretreatment periods, and the error bars represent 95% confidence intervals of the estimated effects and placebos.The panel data on first-child, second-child and total births from November 2014 to October 2015 are constructed using the 2015 census. To draw these figures, we followed the estimation strategy proposed by de Chaisemartin and D'Haultfoeuille (2020).