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

Zhangfeng Jin Shiyuan Pan Zhijie Zheng

October 22, 2024

**Abstract**

This study investigates the effects of easing birth quotas on fertility transition, examining both extensive and intensive margins. Employing an extended Barro-Becker model, we anticipate asymmetric effects on birth rates based on household preferences. By analyzing the unique context of China's two-child policy, which permitted eligible couples to have a second child, and employing a difference-in-differences framework, our analysis reveals a significant increase in second-child births following the relaxation of birth quotas. However, this positive trend is offset by a noticeable reduction in first-child births. The observed decline or postponement in first childbearing is partly attributed to the escalating costs of child rearing, consistent 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 alleviate the financial burdens of child rearing for prospective parents. This targeted strategy is proposed to effectively enhance overall fertility rates.

JEL classification: J13; J18; H23

Keywords: birth quotas; childbearing decisions; child-rearing costs; China

*Jin: School of Economics, Zhejiang University of Technology, 288 Liuhe Road, Hangzhou, China, 310027 (email: zhangfeng1019@gmail.com); Pan: CRPE and School of Economics, Zhejiang University, 38 Zheda Road, Hangzhou, China, 310027 (email: shiyuanpan@zju.edu.cn); Zheng: Bay Area International Business School, Beijing Normal University (email: zhengzhijie@bnu.edu.cn). We appreciate valuable comments from Junsen Zhang, Magne Mogstad, Kevin Lang, Heather Royer, Lee Benham, Uta Schoenberg, Fabian Lange, Hanming Fang, Phillip Levine, David Weil, James Heckman, Xin Meng, and participants at the Jiangxi University of Finance and Economics (2020), Peikang Chang Research Foundation of Development Economics (2021), China Meeting of the Econometric Society (2021), Central University of Finance and Economics (2022), Nanjing University Adam Smith Forum (2023), Asian Meeting of the Econometric Society in East and Southeast Asia (2023), Sun Yat-sen University (2023), Sixth Annual Conference of Government and Economics (2024), NBER Conference on Fertility and Declining Population Growth in High-Income Countries (2024) and 2024 Asian Meeting of the Econometric Society in China. Jin is particularly grateful to mentors and fellow mentees of the Society of Labor Economist Mentoring Program (2023-2024) for providing insightful comments and suggestions on this fertility project. We are grateful to China's National Bureau of Statistics and Tsinghua China Data Center for providing access to the Population Census Data, and China's National Health Commission for providing access to China Migrants Dynamic Survey. Any opinions expressed herein are those of the authors.

# 1 Introduction

China is currently undergoing a significant demographic shift, marked by a pronounced decline in fertility rates and a rapidly aging population. To tackle these demographic challenges, Chinese authorities have implemented substantial relaxations of birth quotas in the past decade. This policy shift has allowed eligible couples, previously constrained by the longstanding One-Child Policy (OCP), the freedom to have a second or even a third child.[1] During this period, the country has witnessed a notable change, with a surge in second and higher-order births, a trend expected due to the policy change. However, unexpectedly, there has been an even sharper decline in first-child births.[2] This paradox prompts intriguing questions: does the relaxation of birth quotas have asymmetric effects on fertility transitions, affecting along the decision of whether to have a child at all (the extensive margin) and the decision of whether to have a second child (the intensive margin)? If so, what are the underlying mechanisms driving these shifts?

In our paper, we present a theoretical framework designed to elucidate how the relaxation of birth quotas influences households' decisions regarding fertility. This framework builds upon the well-established Barro-Becker model (Becker and Barro (1988); Barro and Becker (1989)), incorporating factors such as birth quotas and heterogeneous fertility preferences among households. When birth quotas are relaxed, they affect the overall birth rate through two primary pathways. The first pathway, which we term the "limits-raising channel", operates on a straightforward principle. By loosening birth quotas, families previously constrained by these regulations and desiring to support more children can now legally expand their households. This pathway mainly influences decisions regarding the number of children within families.

However, simply relaxing birth quotas can lead to higher costs associated with raising children. This becomes evident when families, strongly inclined toward larger households and perhaps less aware of the financial challenges involved, opt to expand further. As these families grow, a crowding-out effect may occur, where the sudden increase in newborns places greater demands on resources such as healthcare, caregiving, food, housing, and education, ultimately driving up child-rearing costs. These short-term costs are influenced not only by actual child-bearing outcomes but also by the expectation of a surge in second-child births. This expectation

[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 2019, the number of second and higher-order births among annual births rose substantially from 5.4 to 8.7 million, illustrating the impact of the policy shift. Paradoxically, the number of first births declined from 10.6 to a record low of 5.9 million (National Bureau of Statistics of China (2023)).

is especially pertinent among newlyweds preparing for their first child. Imagine couples in their 20s to 30s, gearing up for parenthood, suddenly confronted with an influx of newborns from couples aged 30 to 40, who are now permitted to have an "extra" child. This projection is further supported by the estimates from the Chinese authorities.[3] This scenario illustrates what we term the "child-rearing cost channel", which influences societal decisions within a dynamic general equilibrium framework, where individual choices are shaped by broader societal behavior. This channel exerts downward pressure on birth rates, especially for prospective parents with low fertility preferences, thus partially counterbalancing the effects of the "limits-raising channel".

In few countries does the cost of raising children play as significant a role in family planning decisions as it does in China. According to a 2022 report from the YuWa Population Research Institute, China stands among the highest globally in terms of child-rearing expenses.[4] The institute, headquartered in Beijing,[5] found that the average cost of raising a child to the age of 17 in China in 2019 was 485,000 yuan ($76,629) for a first child. This amount equated to 6.9 times China's per capita GDP that year. Notably, the costs were even higher in major cities, with Shanghai exceeding 1 million yuan and Beijing close behind at 969,000 yuan.[6] Among the 13 countries analyzed, China ranked second-highest in child-rearing costs, trailing only South Korea, which boasts the world's lowest birth rate.[7]

The importance of the "child-rearing cost channel" is highlighted by the absence of a robust support system, raising concerns about the unintended consequences of recent policy shifts. Despite the relaxation of birth quotas, the current policy lacks measures to alleviate the financial strain of raising children. Critical components such as cash and tax subsidies, along with hous-

[3] For instance, in early 2014, Ma Xu, a deputy of the National People's Congress and Director of the Scientific Research Institute of the National Health and Family Planning Commission, predicted that the partial implementation of the Two-Child Policy could result in about 2 million additional births. He also projected that full implementation of the Two-Child Policy could raise the annual births by approximately 10 million (https://www.guancha.cn/life/2014_03_04_210509.shtml). Similarly, in late 2015, Yang Wenzhuang, Director of the Grassroots Family Planning Guidance Department at the National Health and Family Planning Commission, estimated that by 2020, births could reach 17 million annually, with potential peaks exceeding 20 million and annual population growth of over 3 million due to the policy (https://news.cnr.cn/native/gd/20151111/t20151111_520467533.shtml). However, actual birth outcomes fell short of these projections.

[4] 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%).

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

[6] 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.

[7] 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).

ing assistance, are conspicuously absent from this framework, essential for easing the economic burdens of child rearing. Furthermore, the existing infrastructure supporting families, including childcare facilities and access to professional caregivers, remains inadequate.[8]

This paper employs a robust empirical framework to analyze the causal effects of relaxing birth quotas in China, leveraging the unique implementation of a two-child policy by local governments in different months of 2014.[9] The methodology utilizes a difference-in-differences (DID) approach, enabling the examination of short-term effects using real-world policies and high-frequency monthly data.[10] To ensure the validity of the analysis, an estimator proposed by de Chaisemartin and D'Haultfoeuille (2020) is utilized to confirm that the decision to relax birth quotas is not correlated with prior trends in the outcomes under investigation. This approach is particularly effective in handling heterogeneous treatment effects, such as those observed across different groups or over time. Additionally, this study evaluates whether the timing of relaxing birth quotas aligns with province-specific shocks or policies that could influence the outcomes of interest. This is accomplished through placebo analyses, focusing on individuals less affected by the relaxation of birth quotas, and considering alternative censuses conducted before the policy change.

However, the empirical framework faces challenges in measuring childbearing outcomes and the expenses associated with child-rearing. Unlike earlier studies that relied on annual birth data, this study utilizes China's 2015 population census to create monthly panel data on births. It's important to note that a minimum timeframe of approximately one year is assumed for the policy to take effect, considering factors such as the birth permit procedure, conception time, and pregnancy duration. Detailed information and further explanation regarding these aspects can be found in **Appendices C.1 to C.4**.

To address the challenges of obtaining detailed monthly data on the cost of child-rearing in China, this study adopts a multifaceted approach. We develop three complementary measures to track changes in the average cost of raising children across specific regions over time. The

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

[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.

[10] An underlying assumption is that people incorporate 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.

first measure involves compiling both overall and expenditure-based Consumer Price Indices (CPIs) to capture general shifts in the cost of goods and services, as well as specific changes in expenses relevant to child-rearing. For the second measure, we collect data on overall and square-meter-based residential real estate sale price indices to assess variations in housing costs and understand how housing expenses vary among households with different sensitivities to such costs. The third measure entails gathering data on nanny wages at the individual level to provide insights into fluctuations in the overall cost of hiring a nanny. Unlike the consistent demand for goods and services throughout a child's upbringing (e.g., housing, education), the demand for nanny services is likely to peak during the child's first year and then decline. For a more detailed examination of these measures, please refer to **Section 3.3**.

Our main empirical findings reveal several significant insights. Firstly, the relaxation of birth quotas leads to a notable average increase of approximately 37.9% in monthly second-child births in the short term. This trend aligns with a broader annual surge of 27.6% in the total number of second-child births observed shortly after the policy adoption nationwide. The impact is particularly pronounced among younger women, those with lower levels of education, non-homeowners, and those whose first child is a girl. Additionally, couples covered by pensions experience the most significant impact, supporting the notion that the motive of intergenerational altruism, rather than old-age security, motivate fertility decisions. These results suggest that while child-rearing costs play a role, fertility preferences play a more crucial role in determining second-child births.

Secondly, there is an average decrease of approximately 23.5% in monthly first-child births in the short term following the relaxation of birth quotas. This decline mirrors the annual decrease of about 16.1% in the total number of first-child births observed shortly after the policy adoption nationwide. The decline is more pronounced among younger women, those with lower levels of education, and non-homeowners, indicating a higher sensitivity to child-rearing costs within these groups. This underscore that the impact is more significant for couples who are more attuned to the financial considerations associated with raising children.

Thirdly, the relaxation of birth quotas leads to a significant increase in the cost of child rearing, particularly in the prices of goods and services directly relevant to raising a child, such as residential real estate and nanny services. These findings suggest that the increased costs of child rearing following the relaxation of birth quotas contributes to the asymmetric effects on fertility transitions across both the extensive and intensive margins. Lastly, the relaxation of birth quotas witnesses a substantial short-term increase in the total number of monthly births, indicating that the decline in first-child births does not offset the heightened second-child births during this period. Overall, our findings align with the predictions from an extended Barro-Becker model, as elucidated in **Section 2**.

This study makes several significant contributions to the literature. Firstly, it enhances our comprehension of heterogeneous fertility transitions along both 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 illuminates the often-overlooked dimensions of fertility transitions along both margins, drawing insights from China's relaxation of the OCP. Notably, few studies in this field 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).

Furthermore, this study contributes to the theoretical analysis of OCP's impact. While existing literature has mainly tackled empirical issues, only a handful of studies have delved into theoretical implications (Liao (2013); Choukhmane, Coeurdacier and Jin (2023)). Liao (2013) stands out as one of the few studies that theoretically investigate the general equilibrium effect of relaxing the OCP on childbearing decisions. Through a calibrated general-equilibrium model, the author reveals that skilled parents may opt for fewer children when the policy is lifted due to wage rate effects. This paper specifically zooms in on the child-rearing cost effects within the general equilibrium framework, offering a nuanced understanding of the underlying dynamics.

Moreover, this study contributes to the extensive body of empirical research examining the economic impacts of China's OCP. Previous studies have explored the policy's effects on various family and labor market outcomes, such as the 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 the cost of living, housing prices, and real incomes remain largely underexplored (Zhang (2017)). Our study fills a notable empirical research gap by investigating the effect of China's OCP on child-rearing costs.

Lastly, this study contributes to ongoing debates regarding the fundamental motives driving fertility decisions, which currently present conflicting viewpoints. While some research suggests that intergenerational altruism is the primary driver, other studies provide evidence for the old-age security motive (Nugent (1985); Becker and Barro (1988); Cigno (1992); Rendall and Bahchieva (1998); Rossi and Godard (2022)). Our research further illuminates this issue by identifying which groups, such as those participating in social security, are more likely to have a second child following the relaxation of China's OCP.

The rest of this article is organized as follows. The subsequent section provides an overview of

China's OCP and its reforms, while introducing an extended Barro-Becker model incorporating birth quotas and diverse fertility preferences. Section 3 outlines essential variables and offers a brief description of the data. In Section 4, we detail the model specification and outline our identification strategy. Section 5 presents our empirical findings, and the final section concludes the study.

## 2 Relaxation of China's OCP and its Economic Effects

In this section, we provide a brief overview of China's OCP and its subsequent reforms. Additionally, we expand upon the Barro-Becker framework, integrating birth quotas and heterogeneous fertility preferences among households into a dynamic general equilibrium model. This theoretical examination of the economic implications of easing birth quotas lays the foundation for our empirical investigation.

### 2.1 China's OCP and its reforms

The OCP was formally introduced in China in 1979 and enforced by the National Family Planning Commission. Prior to the OCP, authorities had primarily adopted a voluntary "later-longer-fewer" policy, as discussed by Chen and Fang (2021). With the implementation of the OCP, most couples were restricted to having only one child, and violations incurred penalties, including fines and other punishments such as property seizure and forced dismissal from government employment (Ebenstein (2010)).[11] Moreover, the OCP had extensive consequences beyond birth restrictions. Children without a proper birth permit faced denials of access to hukou, a vital document determining a Chinese citizen's right, including the right to vote and stand for election, the right to work, the right to education, and the right to social security. Notably, the OCP was enforced more strictly in urban areas compared to rural areas. From 1984 onwards, there was a relaxation of policy in some rural areas where couples were permitted to have a second child if their first child was female, a policy known as the "1.5-child policy". See Zhang (2017) for a comprehensive review of the OCP.

In the early 2000*s*, in response to declining fertility rates, the National Family Planning Commission in China initiated a series of relaxations to the OCP. A significant change occurred in 2011, allowing couples where both partners were single offspring to have a second child. Later, in November 2013, this policy was expanded to include couples with only one single-offspring

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

partner (Ouyang (2013)). However, even with these relaxations, "eligible couples" were still required to apply for a birth permit in advance. Couples who had a second child without obtaining the necessary birth permits faced fines ranging from 500 to 2,000 RMB and other penalties, such as the denial of reimbursement of childbirth-related medical costs and subsidies. A crucial development occurred in December 2015 when 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)). Importantly, this policy change eliminated the requirement for a birth permit.

Even though the universal two-child policy is in place, this article concentrates on the impacts of the partial two-child policy due to data constraints. Examining the effects of the partial two-child policy is especially pertinent as it can offer insights into the effects of the universal two-child policy, acting as a conservative estimate.[12]

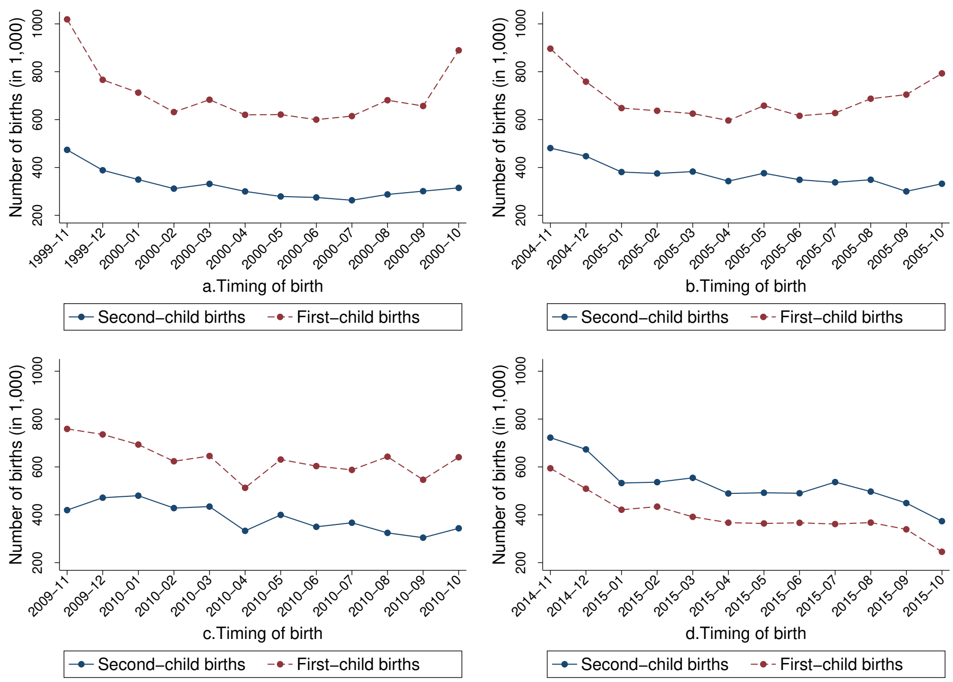


Figure 1: Number of first-child and second-child births over time (2000 – 2015).

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** displays the monthly number of first-child and second-child births based on China's

[12]In **Section 5.6**, we explore how the effect of relaxing birth quotas depends on the number of "eligible couples".

Population Censuses from 2000 to 2015. We observed a gradual decrease in first-child births during this period, while second-child births remained relatively steady until 2010. By 2015, there was a noticeable increase in second-child births, while first-child births saw a significant decline. This trend is not only observed at the national level but also consistently across all provinces. **Figure 2** illustrates the number of second and higher-order births alongside first births reported by the National Bureau of Statistics annually from 2011 to 2019. Once again, as the number of second and higher-order births rose following the relaxation of the OCP, the number of first births experienced a decline. These findings collectively 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 the intensive margin (the decision of whether to have an additional child).

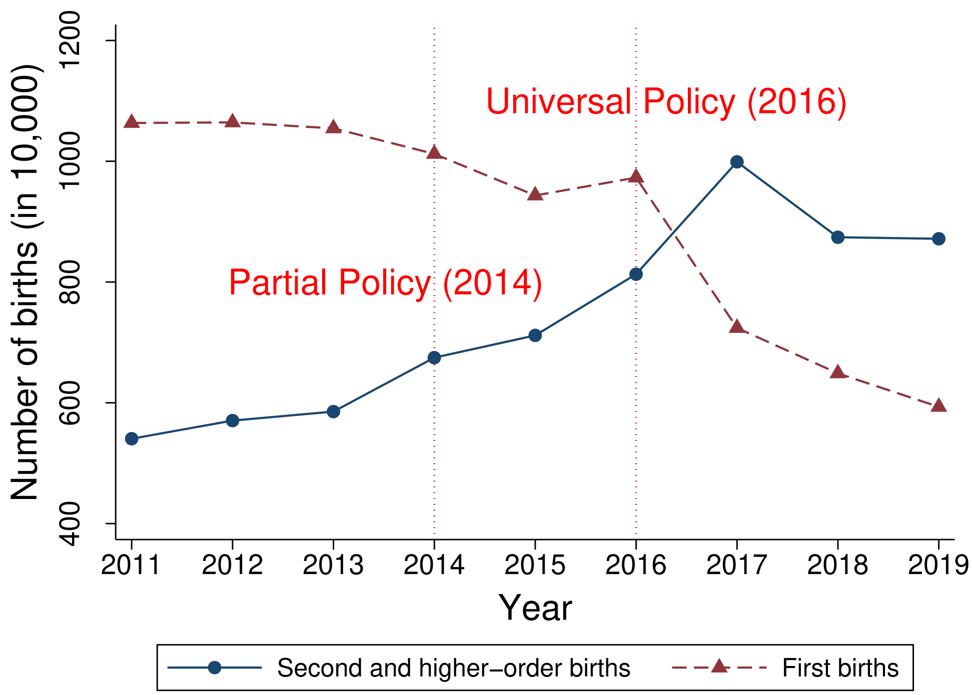


Figure 2: Number of second and higher-order births and first births (2011-2019).

Notes: The solid line denotes the trends in the number of second and higher-order births from 2011 to 2019. The dashed line denotes the trends in the number of first births from 2011 to 2019. 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 number of second and higher-order births and first births 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 distinct types of households, each characterized by different fertility preferences: high-fertility households ($\alpha^h$) and low-fertility households ($\alpha^l$), where $\alpha^l < \alpha^h$. For example, households willing to have two children can be considered as high-fertility households, while those preferring only one child can be considered low-fertility households. The population sizes of high and low fertility households are represented by $q$ and $(1-q)$, respectively, where $q \in (0,1)$. For simplicity, we assume uniform fertility preferences within each household, and these preferences remain constant among descendants within the same household over time.

Each individual experiences two distinct life stages: childhood and adulthood. During each stage, adults decide on the desired number of children and raises them accordingly.[13] Specifically, at the onset of $t=0$, each adult in household $\alpha^h$ selects a number of children, denoted as $n(\alpha^h,0)$, to raise during that period. In generation $t=1$, the offspring, numbering $q \cdot n(\alpha^h,0)$, mature into adults. Each adult then make decisions regarding the desired number of children to raise during this period, denoted as $n(\alpha^h,1)$. Consequently, in generation $t=1$, household $\alpha^h$ consists of $q \cdot n(\alpha,^h 0)$ adults and $q \cdot n(\alpha^h,0) \cdot n(\alpha^h,1)$ children. The birth rate of household $\alpha^h$ is determined by $n(\alpha^h,1)$.

Extending this reasoning to generation $t$, the total numbers of adults and children in household $\alpha^h$ become $q \cdot n(\alpha^h,0) \cdot n(\alpha^h,1) \cdots n(\alpha^h,t-1) = q \cdot \prod_{t=0}^{t-1} n(\alpha^h,t) = qN(\alpha^h,t)$ and $qN(\alpha^h,t+1)$, respectively. Consequently, the birth rate of household $\alpha^h$ is given by $n(\alpha^h,t)$. A similar analysis applies to household $\alpha^l$, where the total numbers of adults and children, and the corresponding birth rate 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.

Building upon the frameworks established by Becker and Barro (1988) and Barro and Becker (1989), we adopt a utility function for each adult in generation $t$ within household $\alpha = \{\alpha^h, \alpha^l\}$:

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

where $1/(1-\sigma)$ represents the elasticity of intertemporal substitution in consumption, $c(\alpha,t)$ denotes the adult's consumption of the final good, and $U(\alpha,t+1)$ represents the utility attained by each child. Altruism levels vary among parents with different fertility preferences, with higher degrees of altruism corresponding to a greater likelihood of having more children.[14] The function

[13] To simplify our model, we do not consider gender difference and the intricacies of the fertility process.
[14] Fertility preferences can be affected by factors such as individual characteristics and culture.

$a(n(\alpha, t))$ quantifies altruism towards each child, defined as:

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

It's important to note that the degree of altruism decreases as the number of children increases, as indicated by Equation (2).

In a competitive labor market, each adult has one unit of time per period and earns uniform wage rates, denoted as $w(t)$. Parents passs on capital, represented by $k(\alpha, t+1)$, to each child born in generation $t+1$. This capital earns a rental yield at the rate $r(t)$ during generation $t$. For an adult of household $\alpha$ during generation $t$, their earnings and inheritance, i.e., $w(t) + [1 + r(t)]k(\alpha, t)$, are allocated among personal consumption $P(t)c(\alpha, t)$, bequests to children $n(\alpha, t)k(\alpha, t+1)$, and child-rearing costs. Here, $P(t)$ represents the prices of the final good.

Raising children incurs costs, denoted as $\beta(\alpha, t)$, which make the total expenses for raising $n(\alpha, t)$ children equal to $n(\alpha, t)\beta(\alpha, t)$. Consequently, the budget constraint for an adult in household $\alpha$ in generation $t$ is expressed as: $w(t) + [1 + r(t)]k(\alpha, t) = P(t)c(\alpha, t) + n(\alpha, t)[\beta(\alpha, t) + k(\alpha, t+1)]$.

The equilibrium conditions, derived from solving the optimization problem, are 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 (3) shows that an increase in the net cost of creating a descendant in generation $t$ leads to a rise in consumption per person for that generation. Intuitively, when people are more costly to produce, it is optimal to endow each person produced with a higher level of consumption. Equation (4) indicates that the optimal number of children $n(\alpha, t)$, for all $t \geq 1$, born to each adult of household $\alpha$, depends on several factors including the degree of altruism, prices of the 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,[15] and only the final good is required such that[16]

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

Given this assumption, Equation (4) then shows that the difference in birth rates is driven by 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 $t \geq 1$ due to $\alpha^h > \alpha^l$.

### 2.2.2 Production

In this economy, there's a single final good produced within a competitive production sector. A typical firm employs labor $L(t)$ and capital $K(t)$ to produce the final good, following the Cobb-Douglas technology $Y(t) = [L(t)]^\eta [K(t)]^{1-\eta}$. The demand for labor and capital, derived from solving the cost-minimization problem, are given by $L(t) = \eta P(t)Y(t)/w(t)$ and $K(t) = (1-\eta)P(t)Y(t)/r(t)$.

At this point, the numeraire in the economy has not been specified. Moving forward, we choose wage rates as the numeraire, denoting $w(t) \equiv 1$.[17] The prices of the final good and interest rates are 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 $k(t) \equiv K(t)/L(t)$ represents the capital stock per capita or per adult.

### 2.2.3 The Government

The government employs distinct fertility restrictions, known as birth quotas, to shape population growth. Let's denote the number of children each household has in generation $t$ as $n(\alpha, t)$, subject to the condition $n(\alpha, t) \in [\underline{n}, \overline{n}]$, where $\underline{n}$ and $\overline{n}$ represent the minimum and maximum permissible number of children.

[15] This assumption simplifies the model analysis and does not seem to conflict with any direct evidence linking child-rearing costs with families' fertility preferences.

[16] The assumption, that raising a child does not necessitate adults to spend time, is vital for guaranteeing the uniqueness of the steady-state equilibrium. For an in-depth discussion, refer to Barro and Becker (1989). Moreover, this assumption proves sufficient in capturing the effect of relaxing birth quotas on child-rearing costs.

[17] This choice simplifies our analysis, particularly in capturing the effects of changes in fertility restrictions on prices of the final good and child-rearing costs. If, alternatively, the prices of the final good were chosen as the numeraire, capture the child-rearing costs would require incorporating the time spent by adults, adding complexity to our analysis.

In our model, the government imposes strict control over the maximum number of children allowed for each household, denoted as $\tilde{n}$, where $\underline{n} < \tilde{n} < \overline{n}$. Consequently, the number of children $n(\alpha, t)$ in generation $t$ adheres to 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 number of adults at the beginning of $t = 0$ is defined as $\hat{N}(0) = q + (1 - q) = 1$, and the total number of children in generation 0 or the total number of adults in generation 1 is $\hat{N}(1) = qn(\alpha^h, 0) + (1 - q)n(\alpha^l, 0)$.

Suppose, at the beginning of generation $t = 1$, the government introduces a fertility policy restricting each household to a maximum of $\tilde{n} = n(\alpha^l, 1) > 0$ children——the number each agent in household $\alpha^l$ is willing to have. This restriction is not binding for $\alpha^l$ households but constraining for $\alpha^h$-households when $n(\alpha^h, 1) > n(\alpha^l, 1)$ in a policy-free environment.

Given this policy, the total number of adults at $t = 2$ and the total birth rate at $t = 1$ are $\hat{N}(2) = q \cdot n(\alpha^h, 0) \cdot \tilde{n} + (1 - q) \cdot n(\alpha^l, 0) \cdot \tilde{n}$ and $B(t = 1) = \hat{N}(2)/\hat{N}(1) = \tilde{n}$. With this fertility policy in place, all members in households $\alpha^h$ and $\alpha^l$ share the same birth rate starting from $t = 1$, and the birth rates of both types of households no long change. Accordingly, the total birth rate remains constant thereafter. Thus

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

Given this consistent population growth rate, the economy is on a balanced growth path, signifying that the growth rate of output per adult is zero.[18]

### 2.2.5 Short-term effects of relaxing birth quotas

In this subsection, we explore the short-term effects of relaxing birth quotas, reserving the discussion on long-term effects for Appendix A.3. Suppose that at the outset of generation $t = s > 1$, the government eases the restricted fertility policy from $\tilde{n}$ to $n^* \in (\tilde{n}, n^h)$, where $n^h$ is the birth rate of household $\alpha^h$ in the equilibrium without fertility control.[19]

Upon the introduction of the new fertility policy, agents in both types of households may adjust their fertility decisions in response to the potential change of macroeconomic factors.

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

[19] We consider the case of $n^* < n^h$ as an example; the case $n^* \geq n(\alpha^h, s - 1)$ would lead to similar mechanisms.

We begin by analyzing the impact of the relaxation of birth quotas on the fertility decision of $\alpha^l$-type households. Utilizing (5) and normalizing wage rate to one, we can simplify (4) to:

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

As the new fertility policy is implemented, the prices of the final good and interest rates may deviate from their previous balanced growth path equilibrium values. In this scenario, $n(\alpha^l, s)$ is determined by:

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

Combining (9) and (10), we obtain:

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

Intuitively, raising the permitted number of children from $\tilde{n}$ to $n^*$ allows the $\alpha^h$-type households to have more children. Consequently, the number of children in generation $s$ increases, leading to a decrease in capital per adult in generation $s + 1$. This is the partial equilibrium result, denoted as the "limits-raising channel". Equation (11) then reveals that a relaxation of birth quotas generates additional general equilibrium influence by adjusting households' fertility decisions through the following channels.

From Equation (6), it is evident that both the prices of the final good and interest rates in period $s + 1$ rise in response. Higher interest rates $r(s + 1)$ (relative to $r(s)$) produces two opposing effects on the fertility decisions of $\alpha^l$-type households.

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 because fertility represents the rate of investment in these descendants. Higher interest rates imply increased return rates on saving in the form of the number of descendants. Therefore, households may choose to consume less and allocate more income to child-rearing, resulting in an increase in the number of descendants.[20]

On the other hand, higher interest rates can generate a counteracting negative effect on the birth rate of $\alpha^l$-type households. Due to higher interest rates $1 + r(s + 1)$, the household faces

[20] This effect can also be verified in (A.8); the increased number of descendants comes at the cost 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 the new fertility policy.

an increased opportunity cost of raising a child; higher interest rates in the future make current spending on child-rearing more expensive, which is 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 the final good $P(s+1)$ (relative to $P(s)$), captured by $\Phi_2 < 1$, tend to decrease the birth rate of $\alpha^l$-type households. Similarly, Equation (3) implies that higher prices of the 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, $\Phi_2$ and $\Phi_3$, together as "child-rearing cost channel". Therefore, if the "child-rearing cost channel" is sufficiently large, the birth rate of $\alpha^l$-type households will decline, that is, $n(\alpha^l, s) < n(\alpha^l, s-1) = \tilde{n}$. Otherwise, the birth rate of $\alpha^l$-type households will increase.

As for the impact on the fertility decision of $\alpha^h$-type households, since the government relaxes the birth quotas on a large scale (i.e., $n^*$ is significantly larger than $\tilde{n}$), supporting the dominance of the "limits-raising channel" (a partial equilibrium outcome) over the other general equilibrium impacts. Consequently, the birth rate of $\alpha^h$-type households always increases. Now, let's examine the influence on the total birth rate, 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 the total birth rate in periods ahead of the new fertility policy is $\tilde{n}$, we have

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

Given a sufficiently large relaxation of birth quotas (i.e., $n^* > \tilde{n}$), the total birth rate is likely to increase if (1) the birth rate of low preference households increases (i.e., $n(\alpha^l, s) > \tilde{n}$); (2) the size of household $\alpha^h$ at period $t=1$ is large relative to that of household $\alpha^l$, 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 birth quotas (i) increases short-term child-rearing costs; (ii) may increase or decrease the short-term birth rate of low preference households ($\alpha^l$-type) and the total birth rate; (iii) increases the short-term birth rate of high preference households ($\alpha^h$-type).*

*Proof.* Proven in text. □

### 2.2.6 Bridging theoretical predictions and empirical testing

To bridge theoretical predictions and empirical testing, we make several simplifications due to data limitations. First, we define fertility preferences based on the number of children desired. Households willing to have two children are categorized as high-fertility households, while those preferring only one child are classified as low-fertility households. Second, we focus on measuring fertility outcomes over a very short period of time. This approach allows us to assess the immediate responsiveness of fertility decisions and potential general equilibrium effects more accurately. However, it may weaken the connection between our empirical findings and the theoretical framework, which considers fertility decisions across two generations. Third, we examine the relaxation of birth quotas in the context of China's transition from the OCP, denoted as $\tilde{n} = 1$, to the universal two-child policy, where $n^* = 2$. According to Proposition 1, our model suggests that this relaxation alleviates restrictions on high-preference households, leading to an increase in second-child births. Nonetheless, the dynamic impact of this policy change extends beyond a mere rise in second-child births. This surge in the number of children could trigger equilibrium effects in the economy, which may subsequently influence the fertility desires of low-preference households. Consequently, we may observe either an increase or a decrease in first-child births. If the rise in second-child births does not sufficiently offset any decline in first-child births, the overall birth might actually decrease following the relaxation of birth quotas.

# 3 Data, Variables, and Summary Statistics

## 3.1 Childbearing outcomes

To assess childbearing outcomes, we utilized a 10% random sample from the 2015 population census conducted by China's National Bureau of Statistics (NBS). This dataset encompasses approximately 1.37 million individuals, representing a significant cross-section of the entire Chinese population. The census data offer a distinct advantage by providing comprehensive details on childbirth, particularly for women aged between 15 and 50, including precise information on the monthly timing of childbirth.[21]

Building on the timing of childbirth, we define a birth in month $t$ as a woman giving birth to a child during that specific month. We aggregate all births at the province level, applying the sampling weights provided by the NBS to ensure representativeness, enabling us to calculate total births in province $s$ in month $t$. The census data also encompass individual characteristics such

[21] While retrospective reporting of births can introduce measurement errors-especially due to issues like infanticide-the inclusion of children who were not alive during the census mitigates this concern. Census data are generally considered relatively reliable in capturing childbearing outcomes.

as age, gender, education, hukou status, total number of births, and social security participation, which allows us to compute monthly births by cohort and distinguish between first and second children. Consequently, we construct panel data on monthly second-child (first-child and total) births for all 31 provinces spanning from November 2014 to October 2015.

While the ideal outcome variable could be a monthly fertility rate that adjusts for changes in the age structure of mothers, such data are not available to us. In our empirical analysis, we primarily use the number of births as the outcome variable based on monthly birth outcomes within a calendar year from the 2015 population census. Additionally, we utilize yearly total population data to calculate birth rate variables for robustness checks. Although accounting for the age structure of mothers is crucial, we think believe that it may not pose a serious concern within a short period (e.g., one year).

## 3.2 Relaxation of birth quotas

To quantify the relaxation of birth quotas, we utilized data related to the two-child policy, which the Chinese government introduced in November 2013 and was gradually adopted by provinces throughout 2014. In **Section** 5.2, we explain why the anticipation effect is unlikely to be a significant concern in our context.

While the policy was rolled out relatively quickly across provinces, there were noticeable variations in timing. For instance, fewer than two-thirds of provinces implemented the policy within six months of the central government's announcement, while the remaining provinces delayed the adoption by more than half a year. We leverage these differences in timing to identify the causal impact of relaxing birth quotas. In **Section** 5.1, we explain why the pre-treatment trends differences across provinces are unlikely to be a significant concern in our context. Furthermore, we assess the strength of the policy's impact by considering variations in exposure to the policy based on the numbers of eligible couples (for details, see **Section 5.6**). **Figure 3a** illustrates the timing distribution of the relaxation of birth quotas across provinces.

Since provinces enacted the same policy on different dates, sometimes within the same month (e.g. February $5^{th}$ versus February $25^{th}$), we assume that if a province adopted the policy before the middle of the month,it took effect within that month. If adoption occurred after mid-month, we assume the policy took effect in the following month. We also account for 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 conceive (Olsen, Juul and Basso (1998); Gnoth, Godehardt, Godehardt, Frank-Herrmann and Freundl (2003); Eisenberg, Thoma, Li and McLain (2021)) and 10 months to complete the pregnancy. See **Appendices C.1 to C.4** for more details on the recoding of the timing of the relaxation of birth quotas.

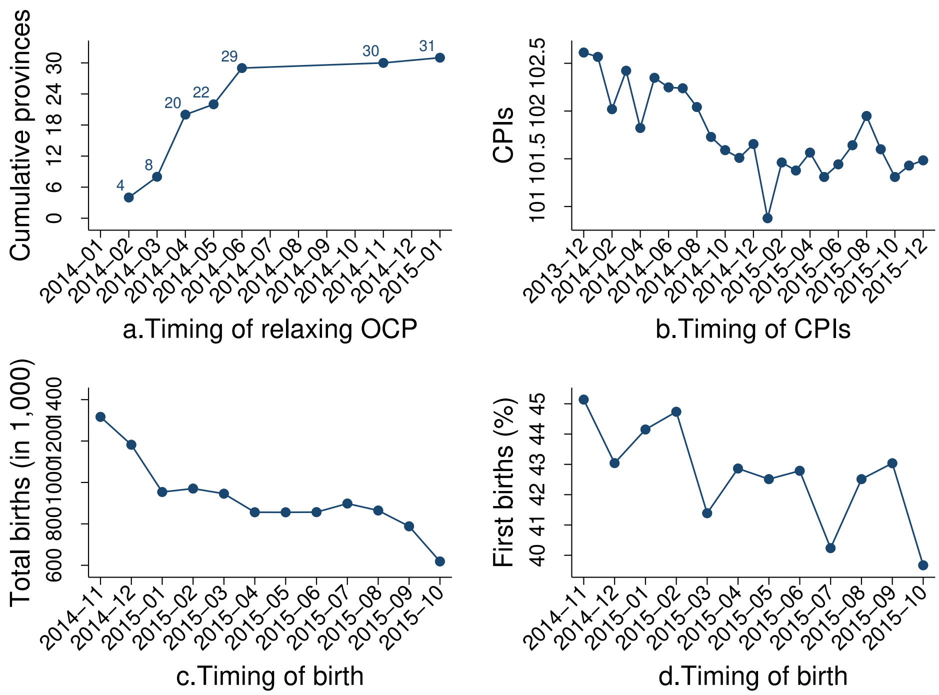


Figure 3: Descriptive summary for the relaxation of birth quotas, CPIs and births

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 the expenses associated with child-rearing costs in China poses challenges due to limited available data. Since directly calculating the overall expenditure for child-rearing in each household over time is not feasible, we employ an alternative method. Our approach involves measuring changes in the prices of final goods and services essential for raising a child. The underlying assumption is that if raising a child requires a minimum amount of final goods and services (e.g., hiring a nanny), then variations in the prices of these goods and services can effectively reflect changes in the overall cost of child rearing. This methodology, to some extent, aligns with the conceptualization of child-rearing costs as defined in our theoretical model.

Given data constraints, we have developed three distinct measures to assess changes in the prices of final goods and services relevant to child-rearing. The first measure utilizes Consumer Price Indices (CPIs), which track the average price fluctuations over time for a basket of goods and services purchased by consumers. Data on CPIs in each province from August 2014 to July 2015 are obtained from China's National Bureau of Statistics. Acknowledging that overall CPIs may not accurately capture changes in the true costs of raising a child, we examine CPIs by expenditure category to identify those that are more relevant to child rearing. **Figure 3b** illustrates the average CPIs over time, showing a gradual decrease from December 2013 to the lowest point in January 2015.

Our second measure involves Residential Real Estate Sale Price Indices. Previous studies have underscored the importance of home prices in influencing fertility decisions (Yi and Zhang (2010); Lovenheim and Mumford (2013); Dettling and Kearney (2014); Liu, Xing and Zhang (2020)). Moreover, housing costs, which constitute a substantial portion of child-rearing expenses in China, are not fully captured by CPIs. To address this, we use this measure to assess changes in the cost of child rearing associated with residential real estate sale prices. These indices are obtained from China's National Bureau of Statistics for 70 large and medium-sized cities (excluding Tibet)from August 2014 to July 2015. We further categorize residential real estate sale price indices by the type of building and floor area, enabling us to examine the heterogeneous impacts on housing costs across households with different sensitivities to such expenses.

Our third measure focuses on the cost of hiring a nanny. Recognizing that demand for certain goods and services (e.g., housing, education) steadily increases as a child grows, while the demand for others (e.g., nanny services) likely peaks during the first year of a child's life and subsequently decline. To capture changes in the cost of child rearing related to hiring a nanny, we construct a third measure——nanny wages at the individual level. Leveraging the China Migrants Dynamic Survey ($2012-2015$), a nationally representative dataset conducted by the National Health Commission of the People's Republic of China, we obtain wage information for migrant workers, including nannies. This dataset, with a total sample size of $764,288$, enables us to estimate the causal impact of relaxing birth quotas on nanny wages within a DID framework.

**Tables 1 and 2** display the summary statistics of the main variables. **Figures 3c and 3d** illustrates 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** presents the monthly second-child births between treatment and control groups over time in an event-study setup.[22] We observe a significant rebound in monthly second-child births in the treatment group after the policy relaxation, whereas no similar patterns are found for the control group.

[22]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 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.

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.

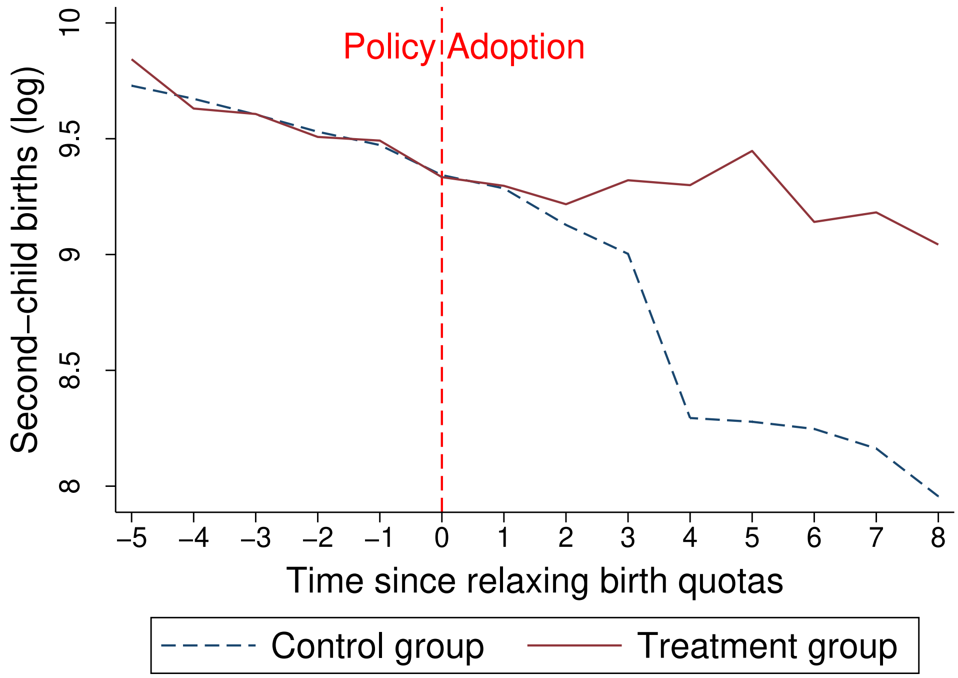


Figure 4: Number of second-child births (in log) over time

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.

# 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 birth quotas.[23] 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:

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

[23]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.

where $Y_{s,t_{birth}}$ is the logarithm of monthly second (first, total) births in province $s$ in month $t_{birth}$,[24] $I_s$ 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 at least 12 months, as previously 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 $\gamma$ is the coefficient of interest. We cluster standard errors at the provincial level.[25]

The identification of parameter $\gamma$ relies on assumptions akin to those in typical DID analyses. It is crucial that the decision to relax birth quotas, including the timing thereof, is not correlated with any prior trends in birth outcomes. Furthermore, the timing of the policy change should not coincide with any province-specific shocks or policies that might influence birth outcomes.

However, several challenges exist in establishing identification. Recent studies suggest that, under the assumption of a constant treatment effect across groups and over time, the two-way fixed effects regression employed above identifies the effect under the standard "common-trends" assumption. Nevertheless, it is often implausible for the treatment effect to remain constant. When the constant effect assumption is not tenable, the two-way fixed effects regression identifies weighted sums of the average treatment effects (ATEs) in each group and period. Importantly, these weights may take negative values, as demonstrated by de Chaisemartin and D'Haultfoeuille (2020), Sun and Abraham (2020), Athey and Imbens (2022), Callaway and Sant'Anna (2021), Imai and Kim (2018), and Goodman-Bacon (2018).

Addressing the second potential identification problem, which concerns the correlation between unobserved, province-specific shocks and the relaxation of birth quotas, presents a more intricate challenge. To investigate this issue, we conduct a falsification test, focusing on birth outcomes for women who migrate from other provinces, namely inter-provincial migrants. The rationale behind this test is grounded in China's hukou system, where the relaxation of birth

[24] 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.

[25] 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.

quotas primarily influences fertility decisions for women registered within the same province (i.e., local natives). According to our hypothesis, if the relaxation of birth quotas indeed operates as posited, this policy should not yield a similar and significant impact on birth outcomes for inter-provincial migrants who originate from provinces that may have relaxed birth quotas at different points in time.

To enhance the robustness of our analysis, we extend our investigation by incorporating alternative population censuses from the years 2005 and 2010. This expansion enables the construction of birth outcomes for all 31 provinces during two distinct periods: November 2004 to October 2005 and November 2009 to October 2010. These alternative datasets, representing periods before the relaxation of birth quotas, serve as a basis for conducting placebo analyses. This strategic approach is particularly relevant for addressing concerns related to seasonality effects, such as the timing of fertility decisions within a calendar year, which could potentially confound the observed policy effect. If the primary estimates are not solely driven by unobserved province-specific shocks, including seasonal effects, the placebo analyses using these counterfactual birth outcomes should not reveal similar and significant impacts of the relaxation of birth quotas.

# 5 Empirical Results

In this section, we examine the effects of relaxing birth quotas. Firstly, we assess how this relaxation influences second-child births and whether its impact varies across different demographic groups. Secondly, we investigate the effects on first-child births. Thirdly, we explore the implications for child-rearing costs. Lastly, we analyze the overall impact on total births.

## 5.1 Pretreatment trends

A critical assumption underlying the DID approach is the existence of common parallel trends in the pretreatment periods. As illustrated in **Figure 4**, we observe similar pretreatment trends. To formally test this assumption, we employ the estimator proposed by de Chaisemartin and D'Haultfoeuille (2020).[26] If the common parallel assumption holds, we would expect the relaxation of birth quotas to have no significant impacts in the pretreatment periods. **Figure 5a** displays the placebo effects in the pretreatment periods for the outcome variable.[27] The horizon-

[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.

tal axis represents time relative to the de facto relaxation of birth quotas, with negative values indicating the pretreatment periods. The vertical axis denotes changes in the outcome of interest (i.e., the logarithm of second-child births). The points represent the estimated treatment effects in the treatment periods and placebo effects in the pretreatment periods, while the error bars indicate the 95% confidence intervals of these estimates and placebos. Our estimation results reveal that the placebo effects are close to zero and not statistically significant at the conventional level in the pretreatment periods.

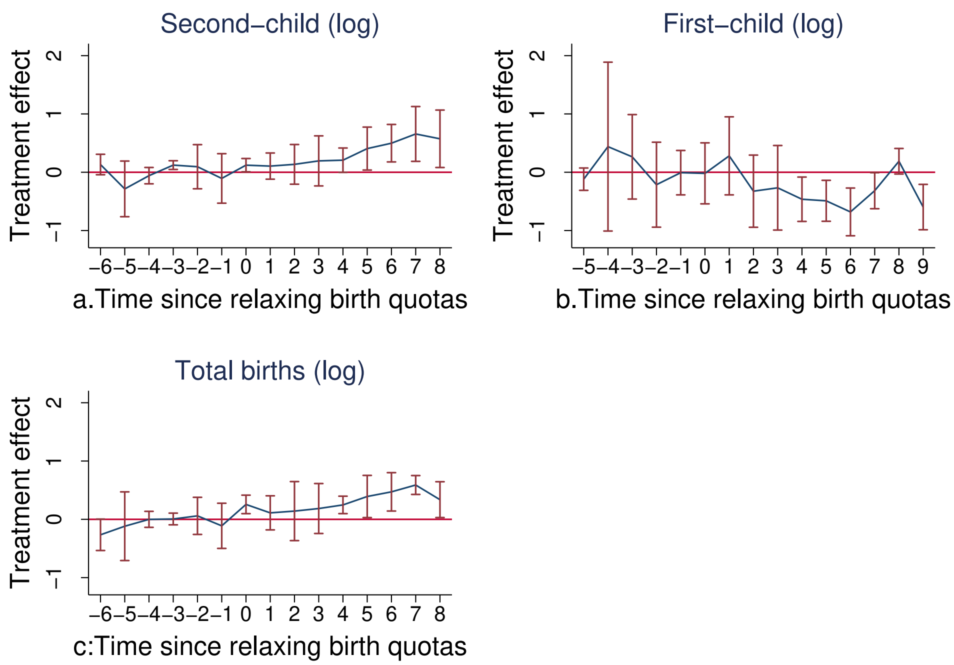


Figure 5: Effects of relaxing birth quotas on second-child, first-child and total births.

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).

## 5.2 Anticipation effect

A potential concern is the anticipation effect, where couples in regions expecting the reform could have advanced their fertility decisions before the official policy change. However, this is unlikely to be a significant issue in our case. First, at that time, it remained illegal for ineligible couples to have a second child. Those who violated the OCP faced fines and a range of penalties, including property seizures and dismissal from government employment. For couples willing to accept these penalties, there would be little incentive to alter their fertility behavior in response to the mere announcement of the national policy.

Second, there was no clear guideline on how quickly the reform would be implemented nationwide. On November 17, 2013, Peian Wang, the vice Minister of the National Health and Family Planning Commission, stated in an official interview that there would be no unified timetable for policy implementation. Provinces, autonomous regions, and municipalities were to determine their own timelines based on local conditions (http://www.nhc.gov.cn/xcs/s3574/201311/24fb42100d9945a99de21712f3f38f55.shtml). Later, on December 7, 2013, Wenzhuang Yang, director of the family planning instruction department, confirmed in an interview with China Central Television (CCTV) that the policy would likely be implemented in early 2014, once local administrations had completed preparations and amended their regulations. Yang emphasized that while provinces could set their own timelines, significant delays in implementing the policy across regions were not expected. He also advised couples to plan the timing of their second baby carefully, as the policy would be long-standing. Notably, there was considerable variation in the timing of reform adoption across provinces, with a gap of nearly a year between the earliest adopter (February 2014) and the latest (January 2015).

Third, we provide empirical evidence suggesting that the anticipation effect is not a major concern. For instance, we do not observe significant pre-treatment effects. To further account for the common trend assumption, we conducted a counterfactual analysis based on the birth outcomes of inter-provincial migrants. Fortunately, the results do not show patterns reject the common trend assumption, mitigating the concern that the anticipation effect might bias our estimates significantly. Even if some couples did adjust their fertility timing-either by advancing a second child or delaying a first-prior to the official relaxation of birth quotas, this would like lead our estimates to provide a lower bound of the policy's impact on second-child (or first-child) births (Malani and Reif (2015)).

## 5.3 Impact of relaxing birth quotas on second-child births

We utilize equation (14) to estimate the effect of relaxing birth quotas on second-child births, with results displayed in Column 1 of **Table B.1**. The coefficient of interest, noted as $-0.007$, does not achieve statistical significance at the conventional level. Addressing the issue of negative weights, as previously discussed, we examine the weights assigned to each average treatment effect in group $s$ and period $t$. Our analysis reveals a negative weight problem in the data, indicating that under the common-trends assumption, the estimator $\hat{\gamma}$ estimates a weighted sum of 199 Average Treatment Effects of the Treated (ATTs), with 159 ATTs assigned positive weights and 40 receiving negative weights. To mitigate concerns about the potential limitations of the two-way fixed effects model in capturing average treatment effects in the presence of heterogeneous treatment effects, we employ the estimator proposed by de Chaisemartin and D'Haultfoeuille (2020) to estimate the impact of relaxing birth quotas on second-child births. The main results are illustrated in **Figure 5a**.

The relaxation of birth quotas has significantly impacted second-child births, with an average increase of approximately 37.9% in monthly figures immediately following the policy change. This surge can be attributed to the pent-up demand among couples who were previously restricted from having more than one child. Due to these prior limitations, a significant number of couples had a strong desire to expand their families with a second child. This observation is supported by data from the National Health Commission of China, which reveals that the total number of second-child births rose from 5.11 million in 2013 to 6.52 million in 2015, representing a notable national increase of about 27.6%.

The impact varies over time. Relaxing birth quotas results in a 12.7% increase in second-child births in the month when the policy is de facto relaxed, with this impact rising to 92.7% by the seventh month following the policy adoption.

As previously mentioned, the relaxation of birth quotas primarily affects women registered in the same province. Conducting a falsification test using monthly births of inter-provincial migrants, we find no clear evidence that relaxing birth quotas increases second-child births for this group. Conversely, there is a significant increase in second-child births for local natives. These results are depicted in **Figures B.1a and B.1b**.

Finally, we employ counterfactual monthly second-child births from the 2005 and 2010 population censuses to conduct placebo analyses. As demonstrated in **Figures B.2a and B.2c**, we do not find obvious evidence that relaxing birth quotas affects these counterfactual second-child births. These placebo analyses thus help alleviate concerns that our estimation results are driven by unobserved province-specific shocks, such as seasonal effects. Overall, these findings support the conclusion that relaxing birth quotas substantially increases second-child births, consistent with our theoretical model.

### 5.4 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 depicted in **Figures B.3a and B.3b**, regarding childbearing age, the impact is more pronounced for women of relatively early childbearing age compared to those with late childbearing age (below the median value of the childbearing age distribution). Examining women's education levels, as illustrated in **Figures B.4a and B.4b**, the impact is more noticeable among women without a college degree than among women with a college degree or higher. In terms of homeownership, **Figures B.5a and B.5b** indicate that the impact of relaxing birth quotas on second-child births is more significant for women who do not own a home. Regarding the gender of children, it's observed that couples in China are more likely to have a second child if their first child is a girl. **Figures B.6a to B.6d** support this observation, showing that the impact is more pronounced for couples who either have no male child or have a female child at the time of the survey.

Finally, we investigate 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 the support of their children (i.e., the intergenerational altruism motive). On the other hand, having access to pensions may deter childbearing decisions if government-provided social security serves as a substitute for old-age support from children (i.e., the old-age security motive for fertility). **Figures B.7a and B.7b** present the main results. We discover 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. These finding supports the intergenerational altruism motive over the the old-age security motive for fertility.

Overall, these findings indicate that fertility preferences, rather than sensitivity to child-rearing costs, play a more significant role in second-child births. Couples with stronger fertility preferences are more likely to have a second child following the relaxation of birth quotas, irrespective of their economic conditions. Once again, these results align with our theoretical model.

### 5.5 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.[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 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 unobservable factors such as seasonal effects. These findings suggest that relaxing birth quotas substantially decreases or delays first-child births, which aligns with predictions from our theoretical model.

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,

[28] 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.

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.

## 5.6 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 fertility transitions along both the extensive and intensive margins. We now go a step further and examine whether and how child-rearing costs respond following the relaxation of birth quotas using three distinct measures to gauge changes in the price of final goods and services relevant to child-rearing as described in **Section 3.3**.

### 5.6.1 Impact of relaxing birth quotas on consumer price indices

We start by using local CPIs to quantify the average price fluctuations over time for a basket of goods and services paid by consumers. 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 $\gamma_2$ is the coefficient of interest.

Using the same approach, we estimate the dynamic impacts of relaxing birth quotas on local CPIs.[29] As shown in **Figure 6a**, a simple average of the instantaneous and dynamic effects shows

[29]Using a two-way fixed effects model, we find that the coefficient of interest is 0.27. Further investigation 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.

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

Housing costs, which constitute a substantial portion of child-rearing expense in China, are not entirely captured by CPIs. To address this, our second measure focuses on changes in the cost of child rearing associated with residential real estate sale prices. To examine the impact of relaxing birth quotas on residential real estate sale price indices, 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 $\gamma_3$ 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

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.

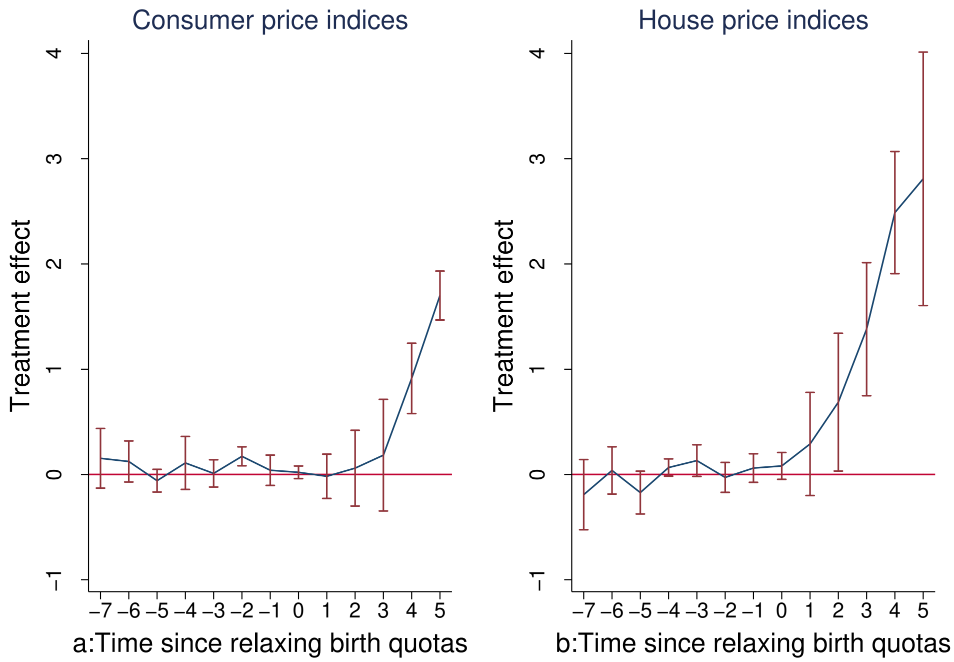


Figure 6: Effects of relaxing birth quotas on consumer price indices and residential real estate sale price indices.

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 households with different sensitivities to such expenses, we investigate the effects of relaxing birth quotas on residential real estate sale prices across three categories of floor area: below 90 $m^2$, between 90 and 144 $m^2$, and above 144 $m^2$. Our analysis reveals that the impact decreases as floor area increases. This trend holds true for both newly constructed (shown in **Figure B.10**) and second-hand (shown in **Figure B.11**) residential buildings. Essentially, disadvantaged couples who are more sensitive to housing costs, tend to face higher housing costs after the relaxation of birth quotas compared to advantaged couples, who are less sensitive to housing costs.

One concern is that higher residential real estate sale prices may lead to increased assets for homeowners, affecting both their child-rearing costs and overall wealth. If this is true, it's expected that first-child births would decline more among non-homeowners compared to homeowners. To test this hypothesis, we investigate the effect of relaxing birth quotas on first-child births based on homeownership status. As anticipated,our analysis reveals a negative impact on non-homeowners and a positive impact on homeowners, consistent with prior research on how home prices influence fertility decisions (Yi and Zhang (2010); Lovenheim and Mumford (2013); Dettling and Kearney (2014); Liu, Xing and Zhang (2020)). However, similar patterns are not observed for the impact of relaxing birth quotas on second-child births, suggesting that other factors, such as fertility preferences, may play a more significant role in determining the decision to have a second child. Further details are provided in Figures B.5a to B.5d.

### 5.6.3 Impact of relaxing birth quotas on nanny wages

As a child grows, the demand for certain goods and services like housing and education steadily increases, while the demand for others such as nanny services likely peaks during the first year of the child's life and then decline. To explore this aspect, our third measure concentrates on the cost of child rearing associated with hiring a nanny.To assess the impact of relaxing birth quotas on nanny wages, 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 $y_{i,s,t}$ represents the logarithm of the monthly wage in April for migrant $i$ in province $s$ in year $t$, $Caregiver_{i,s,t}$ 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 for the year 2012 and 2014, respectively. $K_{i,s,t}$ includes individual characteristics such as gender, age, age squared, education, marital status, hukou type, duration of migration, and duration of migration squared for migrant $i$ in province $s$ in year $t$. $\varepsilon_{i,s,t}$ denotes the error term. The parameter of interest $\gamma_{2014}$ captures the impact of relaxing birth quotas on nanny wages. The parameter $\gamma_{2012}$ represents any differences in pre-treatment trends.

Given that the survey data only provide wage information for April of each survey year, our main analysis focuses on samples from provinces that relaxed birth quotas between January and April 2014, spanning survey years from 2012 to 2014. Additionally, we conduct two placebo analyses: one by limiting samples to provinces that relaxed birth quotas after April 2014 (i.e., during the survey period without the policy in effect), and another using survey years 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. For the second analysis, the reasoning is that higher nanny wages may attract more nannies, potentially reducing wage discrepancies over time.

Our main results indicate that relaxing birth quotas leads to an average increase of approximately 9.3% in monthly nanny wages in the short term (Columns 1 to 4 of Table 3). We observe no significant differences in pre-treatment periods, which supports the common parallel assumption. Additionally, we find that the impact is more pronounced for migrant workers registered in the same city (referred to as local migrants) compared to those registered in other provinces (referred to as non-local migrants). Specifically, the increase in nanny wages for local migrants and non-local migrants is 22.4% and 7.6%, respectively (Columns 5 and 6 of Table 3). The discrepancy may be attributed to the fact that the labor supply of local nannies is less responsive to higher wages compared to non-local nannies, likely due to differences in population size. Our main findings are not supported when using two different placebo analyses, which further strengthens the argument that the higher nanny wages are driven by the 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 (0/1) | −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. $*p < 0.1$, $**p < 0.05$, $***p < 0.01$.

In summary, our analysis employs three distinct measures to assess variations in the prices of final goods and services related to child-rearing. We consistently find evidence that the relaxation of birth quotas significantly increases the expenses associated with raising a child in the local context. Importantly, our results indicate that the parallel trends assumption holds for all measures, alleviating concerns that the substantial changes in child-rearing prices are merely a continuation of pre-existing trend, rather than a consequence of demand induced by the relaxation of OCP. This noticeable short-term price effect supports that forward-looking couples are facing heightened child-rearing costs in the current period, driven by anticipated increases in second-child births. Consequently, some couples may adjust their current childbearing decisions in response to the relaxation of birth quotas.[30] These findings reinforce the validity of the mechanisms outlined in our theoretical model, contributing to a better understanding of the asymmetric effects of relaxing birth quotas on fertility transitions along both the extensive and intensive margins.

## 5.7 Do more eligible couples magnify the policy impact?

In our 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 based on the differences in the number of couples eligible to have a second child. We estimate the treatment effect using 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 this ratio of eligible couples across reform provinces.[31] 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. $\gamma_5$ 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.

Applying the same methodology, we examine how the proportion of eligible couples influences the dynamic fertility effects of relaxing birth quotas. Our findings reveal that the positive

[30] As discussed in Section 2, the price effects emerge in the current period due to the expected rise in second-child births. However, this does not preclude the possibility that the decline in first births could outpace the increase in second births during the same period.

[31] To further address concerns about potential contamination of our estimates by cross-province differences in cohort sizes, we use the absolute number of eligible couples, rather than the ratio of eligible couples, as a proxy for exposure to changes in birth quotas. Our main results remain robust under this approach. More details are provided in Figure 9

impact of relaxing birth quotas on second-child births strengthens as the proportion of eligible couples increases, and the negative impact of relaxing birth quotas on first-child births decreases with a higher proportion of eligible couples. These findings provide further evidence that relaxing birth quotas encourages some couples to have a second child while discouraging first-child births among others. Additional evidence suggests that the positive impact of relaxing birth quotas on child-bearing costs intensified as the proportion of eligible couples rises. More details are provided in Figures 7 and 8.

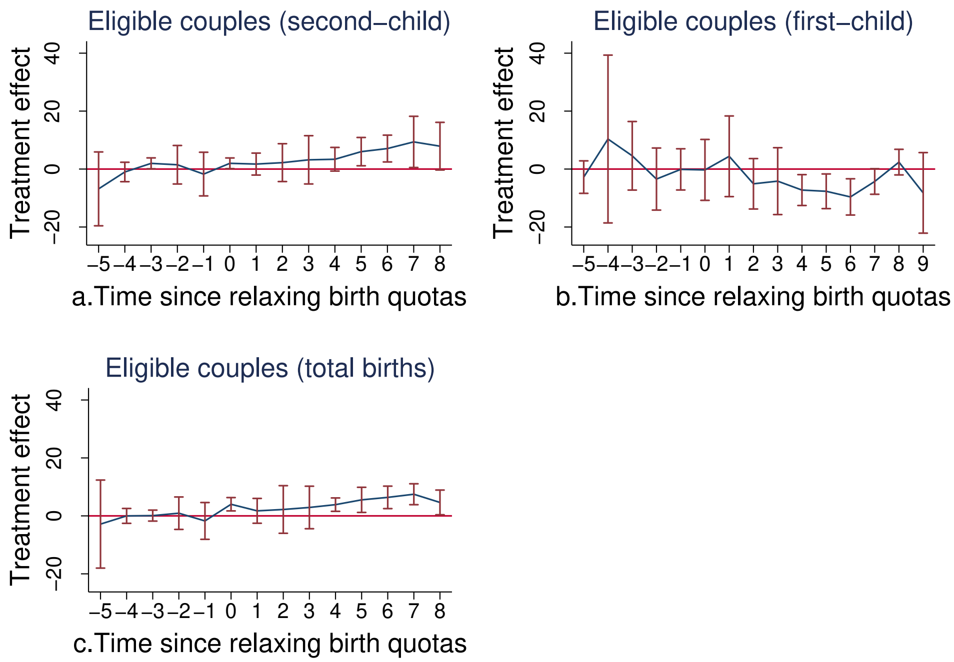


Figure 7: Effects of eligible couples on the effects of relaxing birth quotas.

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 de Chaisemartin and D'Haultfoeuille (2020).

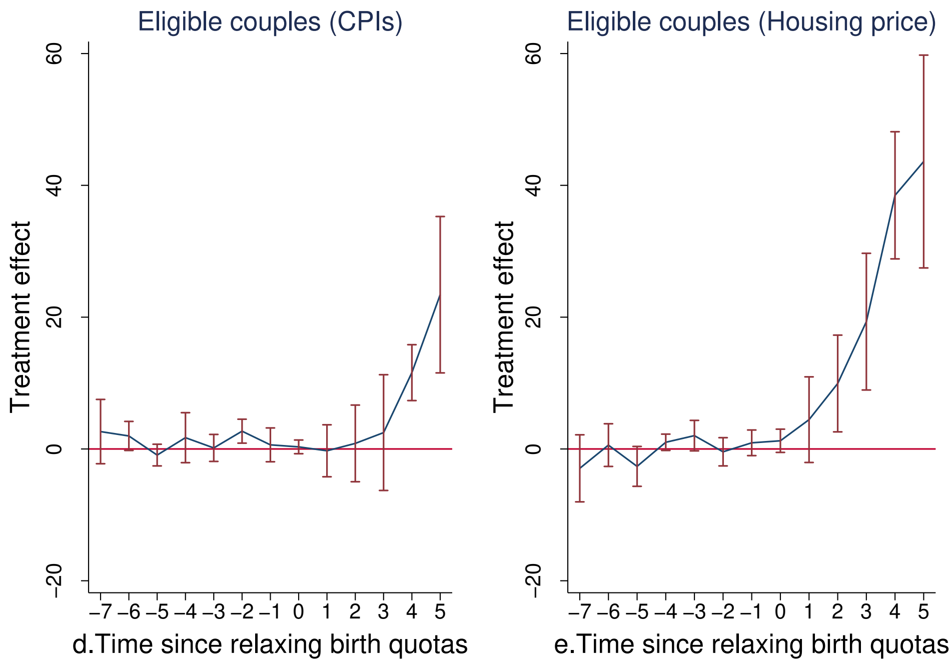


Figure 8: Effects of eligible couples on the effects of relaxing birth quotas (cont.).

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 de Chaisemartin and D'Haultfoeuille (2020).

## 5.8 Does relaxing birth quotas affect total births?

Up to this point, our analysis has demonstrated that relaxing birth quotas increases second-child births while potentially decreasing or delaying first-child births, likely due to increased child-rearing costs in the short term. To determine whether the decline in first-child births offset the rise in second-child births, we further investigate the impact of relaxing birth quotas on total births. Using the same empirical approach, we discover that relaxing birth quotas substantially boosts total births. This suggests that while there may be a decrease in first-child births, it is partially offset by the increase in second-child following the relaxation of birth quotas. This

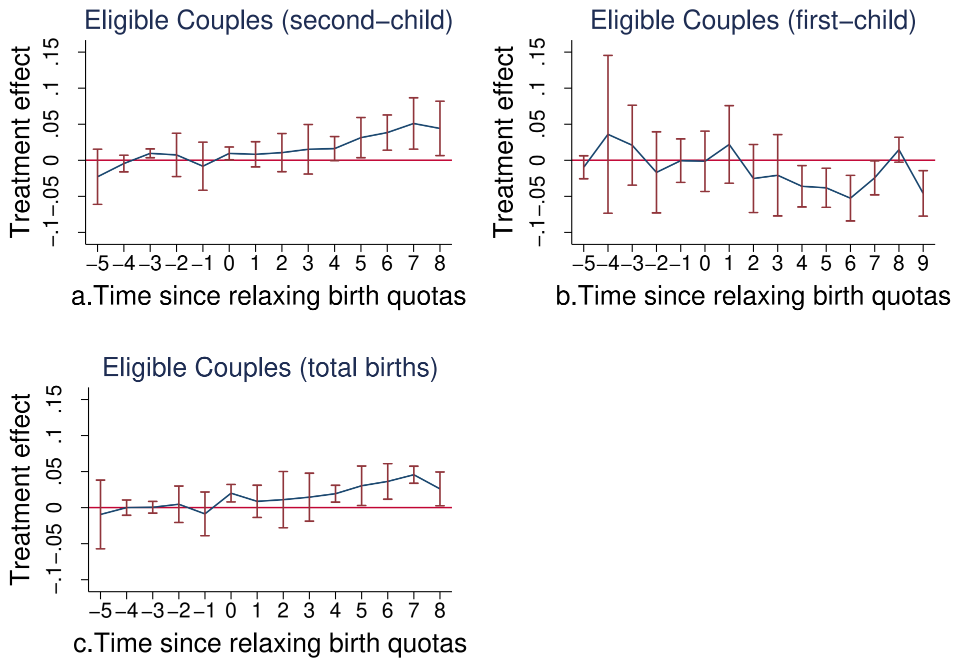


Figure 9: Effects of eligible couples (in absolute number) on the effects of relaxing birth quotas.

Notes: This graph shows how the absolute number 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 , total births and eligible couple 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).

finding further supports the existence of a short-term price effect.[32] See **Figure 5c** for details.

## 5.9 Further discussions on the significant impact of child-rearing costs

First, changes in the number of total births alone do not fully reflect the fertility transition before and after the relaxation of the OCP. According to the National Health Commission of China, total births increased from 16.4 million in 2013 to 16.55 million in 2015, representing a modest rise of 0.9% annually. In contrast, second-child births saw a notable increase of approximately

[32]Even if the number of births decreases at a certain point, a short-term price effect can still exist as long as the cumulative number of births within a given period continues to increase.

27.6%, while first-child births declined by around 16.1%. Furthermore, our estimates indicate a substantial rise in total births following the relaxation of birth quotas.

Second, we examine the immediate impact of relaxing birth quotas on child-rearing costs. The increased demand for child-rearing goods (e.g., housing) and services (e.g., nannies) is unlikely to be met with a quick adjustment in supply, resulting in higher prices. For instance, our analysis shows that the relaxation of birth quotas leads to an average short-term increase of about 9.3% in monthly nanny wages, with a pronounced effect for local nannies (22.4%) compared to non-local nannies (7.6%).

Finally, childlessness among women in China remains low, with fewer than 2% of women aged 49 being childless, according to the 2017 National Fertility Survey (Jiang *et al.* (2023)). Therefore, it is more likely that prospective parents delay having their first-child births rather than forgo it altogether in the short term. This suggests that the sharp rise in second-child births may more accurately reflect changes in local demand for child-rearing than changes in total births during this period.

## 5.10 Competing hypotheses

### 5.10.1 Marriage market

In this subsection, we start by exploring a competing mechanism-the marriage market-as a potential driver of the asymmetric effects of relaxing birth quotas on fertility transitions along both the extensive and intensive margins. To estimate the impact on marriage outcomes, we construct monthly data on the number of marriages across provinces. The model specification is as follows:[33]

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

where $Marriage_{s,t_{birth}}$ represents 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. We assume $r = 0$ to refelct the immediate impact of the relaxation of birth quotas on local marriage markets. This assumption is based on the notion that marrying someone is less time-consuming than giving birth to a child, and forward-looking individuals would anticipate immediate changes in the returns of getting married upon the relaxation of birth quotas. The error term is $\varepsilon_{s,t_{birth}}$, $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are vectors of coefficients to be estimated and $\gamma_6$ is

[33] Given that China's Ministry of Civil Affairs offers data on the number of marriages at the quarterly level for each province, we use the cubic spline interpolation method to calculate the number of marriages at the monthly level from October 2013 to September 2014.

the coefficient of interest, capturing the effect of the relaxation of birth quotas on the number of marriages.

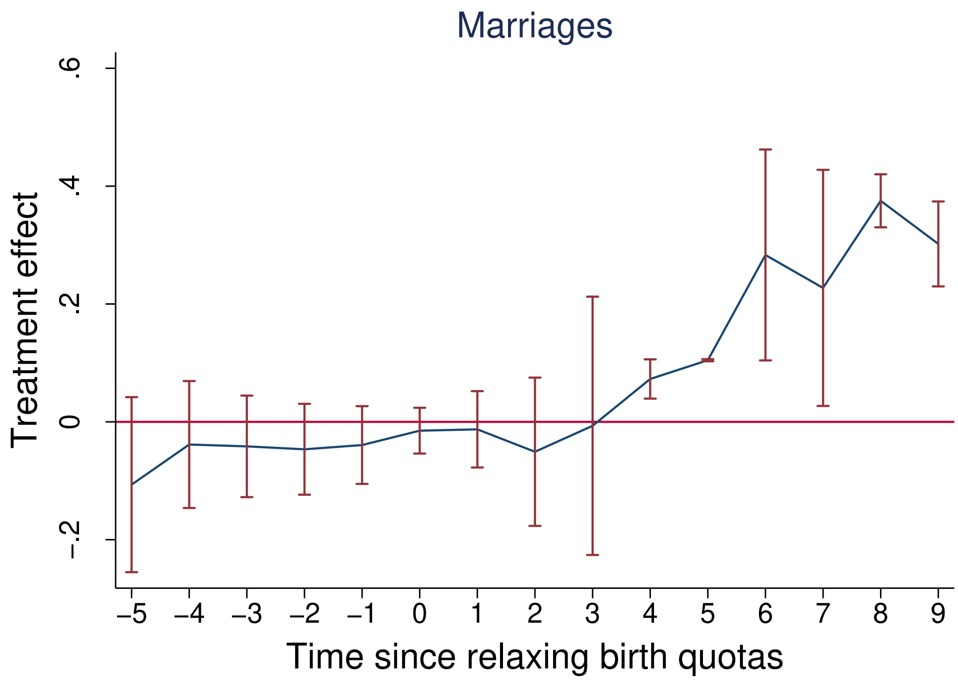


Figure 10: Impact of relaxing birth quotas on local marriages

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).

Using a similar methodology, we estimate the dynamic impacts of relaxing birth quotas on local marriages.[34] Our main results are reported in **Figure 10**. Our analysis reveals a significant increase in the number of marriages, which aligns with findings from a previous study by Huang, Pan and Zhou (2023), who reported a notably lower marriage rate induced by the OCP. Specifically, we observe that the number of marriages increases by 7.4% in the fourth month fol-

[34] Using a two-way fixed effects model, we find that the coefficient of interest is $-0.08$. Further 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.

lowing the relaxing birth quotas, escalating 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. Therefore, the positive impact of relaxing birth quotas on the marriage market strengthens our earlier argument that higher child-rearing costs likely contribute to the decline in first-child births following the relaxation of birth quotas.

### 5.10.2 Fertility decisions within the same household

Another potential factor contributing to the asymmetric effects observed in fertility transitions after the relaxation of birth quotas is the dynamics of internal family support systems. In environments where market-provided childcare is limited, reliance on grandparents or grandparents-in-law for childcare becomes common. This reliance can significantly influence fertility decisions within a household. For instance, if elder siblings decide to have a second child, the limited availability of grandparental childcare might compel younger siblings to delay their plans for a first child. This scenario underscores the critical role of intra-household resource allocation in shaping individual fertility choices under constraints. However, if intra-household resource allocation were the primary influence on fertility decisions, the significant price effects previously discussed would likely not be as pronounced. Additionally, the partial two-child policy analyzed in this paper, which applies specifically to couples where only one partner is an only child, further diminishes the likelihood that the second-child births of elder siblings would impede the first-child births of younger siblings.

## 6 Conclusion

This study investigates the economic consequences of relaxing birth quotas. An extended Barro-Becker model predicts asymmetric effects on fertility transitions along the extensive and intensive margins. By employing a difference-in-differences framework to examine the transition from a one-child to a two-child policy, our analysis reveals a substantial increase in second-child births following the relaxation of birth quotas. Notably, this effect is more pronounced among couples covered by pensions, suggesting intergenerational altruism as a driving force for fertility decisions, rather than solely the motive of old-age security.

Furthermore, our findings present new evidence of a decrease or delay in first-child births following the relaxation of birth quotas. Subsequent analyses highlight a significant increase in the cost of child rearing, aligning with predictions from the extended Barro-Becker model, which suggests that increased child-rearing expenses contribute to the asymmetric effects on fertility transitions. This underscores the importance of implementing policies addressing the financial burden on prospective parents to stimulate fertility rates.

It is crucial to acknowledge the observed spill-over effects, which underscore the inadequacy of a one-size-fits-all approach in shaping fertility through birth quotas. Policymakers should recognize the nuanced and differential impacts on fertility decisions along extensive and intensive margins, and tailor interventions accordingly to address specific needs and challenges associated with each. Our research advocates for targeted measures that alleviate economic constraints on parenthood to effectively tackle demographic challenges and foster sustainable population growth.

## References


Aaronson, D., Lange, F. and Mazumder, B. (2014). Fertility transitions along the extensive and intensive margins. *American Economic Review*, **104** (11), 3701–3724.

Agersnap, O., Jensen, A. S. and Kleven, H. (2020). *The Welfare Magnet Hypothesis: Evidence From an Immigrant Welfare Scheme in Denmark*. Tech. Rep. 4.

Athey, S. and Imbens, G. W. (2022). Design-based analysis in difference-in-differences settings with staggered adoption. *Journal of Econometrics*, **226** (1), 62–79.

Barro, R. J. and Becker, G. S. (1989). Fertility choice in a model of economic growth. *Econometrica*, pp. 481–501.

Baudin, T., De La Croix, D. and Gobbi, P. E. (2015). Fertility and childlessness in the United States. *American Economic Review*, **105** (6), 1852–1882.

Becker, G. S. and Barro, R. J. (1988). A reformulation of the economic theory of fertility. *The quarterly journal of economics*, **103** (1), 1–25.

Callaway, B. and Sant'Anna, P. H. C. (2021). Difference-in-differences with multiple time periods. *Journal of Econometrics*, **225** (2), 200–230.

Cameron, A. C., Gelbach, J. B. and Miller, D. L. (2008). Bootstrap-based improvements for inference with clustered errors. *The Review of Economics and Statistics*, **90** (3), 414–427.

Chandra, A., Gruber, J. and McKnight, R. (2010). Patient cost-sharing and hospitalization offsets in the elderly. *American Economic Review*, **100** (1), 193–213.

Chen, Y. and Fang, H. (2021). The long-term consequences of China's "Later, Longer, Fewer" campaign in old age. *Journal of Development Economics*, **151**, 102664.

Choukhmane, T., Coeurdacier, N. and Jin, K. (2023). The One-Child Policy and Household Saving. *Journal of the European Economic Association*, **21** (3), 987–1032.

Cigno, A. (1992). Children and pensions. *Journal of Population Economics*, **5** (3), 175–183.

de Chaisemartin, C. and D'Haultfoeuille, X. (2020). Two-way fixed effects estimators with heterogeneous treatment effects. *American Economic Review*, **110** (9), 2964–96.

Dettling, L. J. and Kearney, M. S. (2014). House prices and birth rates: The impact of the real estate market on the decision to have a baby. *Journal of Public Economics*, **110**, 82–100.

Doepke, M., Hannusch, A., Kindermann, F. and Tertilt, M. (2023). The economics of fertility: A new era. In *Handbook of the Economics of the Family*, vol. 1, Elsevier, pp. 151–254.

Ebenstein, A. (2010). The "missing girls" of China and the unintended consequences of the one child policy. *Journal of Human Resources*, **45** (1), 87–115.

Eisenberg, M. L., Thoma, M. E., Li, S. and McLain, A. C. (2021). Trends in time-to-pregnancy in the USA: 2002 to 2017. *Human Reproduction*.

Gnoth, C., Godehardt, D., Godehardt, E., Frank-Herrmann, P. and Freundl, G. (2003). Time to pregnancy: results of the German prospective study and impact on the management of infertility. *Human reproduction*, **18** (9), 1959–1966.

Gobbi, P. E. (2013). A model of voluntary childlessness. *Journal of Population Economics*, **26**, 963–982.

Goodman-Bacon, A. (2018). *Difference-in-differences with variation in treatment timing*. Tech. rep.

Huang, W., Lei, X. and Sun, A. (2021). Fertility Restrictions and Life Cycle Outcomes: Evidence from the One-Child Policy in China. *Review of Economics and Statistics*, **103** (4), 694–710.

—, — and Zhao, Y. (2016). One-child policy and the rise of man-made twins. *Review of Economics and Statistics*, **98** (3), 467–476.

—, Pan, Y. and Zhou, Y. (2023). One-child policy, marriage distortion, and welfare loss. *Review of Economics and Statistics*, pp. 1–47.

Hwang, J. (2023). Later, fewer, none? Recent trends in cohort fertility in South Korea. *Demography*, **60** (2), 563–582.

Imai, K. and Kim, I. S. (2018). On the use of two-way fixed effects regression models for causal inference with panel data.

Jiang, Q., Zhang, C., Zhuang, Y., Jiang, Y. and Zhang, X. (2023). Rising trend of childlessness in china: analysis of social and regional disparities with 2010 and 2020 census data. *BMJ open*, **13** (5), e070553.

Li, H., Yi, J. and Zhang, J. (2011). Estimating the effect of the one-child policy on the sex ratio imbalance in China: Identification based on the difference-in-differences. *Demography*, **48** (4), 1535–1557.

Li, H.-t., Xue, M., Hellerstein, S., Cai, Y., Gao, Y., Zhang, Y., Qiao, J., Blustein, J. and Liu, J.-m. (2019). Association of China's universal two child policy with changes in births and birth related health factors: national, descriptive comparative study. *Bmj*, **366**.

Liao, P.-J. (2013). The one-child policy: A macroeconomic analysis. *Journal of Development Economics*, **101**, 49–62.

Liu, H. (2014). The quality–quantity trade-off: evidence from the relaxation of China's one-child policy. *Journal of Population Economics*, **27** (2), 565–602.

Liu, J., Xing, C. and Zhang, Q. (2020). House price, fertility rates and reproductive intentions. *China Economic Review*, **62**, 101496.

Lovenheim, M. F. and Mumford, K. J. (2013). Do family wealth shocks affect fertility choices? Evidence from the housing market. *Review of Economics and Statistics*, **95** (2), 464–475.

Malani, A. and Reif, J. (2015). Interpreting pre-trends as anticipation: Impact on estimated treatment effects from tort reform. *Journal of Public Economics*, **124**, 1–17.

National Bureau of Statistics of China (2023). National data.

Nugent, J. B. (1985). The old-age security motive for fertility. *Population and development review*, pp. 75–97.

Olsen, J., Juul, S. and Basso, O. (1998). Measuring time to pregnancy. Methodological issues to consider. *Human reproduction (Oxford, England)*, **13** (7), 1751–1753.

Ouyang, Y. (2013). China relaxes its one-child policy.

Qian, N. (2009). *Quantity-quality and the one child policy: The only-child disadvantage in school enrollment in rural China*. Tech. rep., National Bureau of Economic Research.

Rendall, M. S. and Bahchieva, R. A. (1998). An old-age security motive for fertility in the United States? *Population and development review*, pp. 293–307.

Roodman, D., Nielsen, M. Ø., MacKinnon, J. G. and Webb, M. D. (2019). Fast and wild: Bootstrap inference in Stata using boottest. *The Stata Journal*, **19** (1), 4–60.

Rosenzweig, M. R. and Zhang, J. (2009). Do population control policies induce more human capital investment? Twins, birth weight and China's "one-child" policy. *The Review of Economic Studies*, **76** (3), 1149–1174.

Rossi, P. and Godard, M. (2022). The old-age security motive for fertility: evidence from the extension of social pensions in Namibia. *American Economic Journal: Economic Policy*, **14** (4), 488–518.

Scharping, T. (2003). Birth Control in China, 1949-2000: Population Policy and Demographic Development.

Sun, L. and Abraham, S. (2020). Estimating dynamic treatment effects in event studies with heterogeneous treatment effects. *Journal of Econometrics*.

Wang, F., Zhao, L. and Zhao, Z. (2017). China's family planning policies and their labor market consequences. *Journal of Population Economics*, **30** (1), 31–68.

Xinhua (2020). Across China: China moves to ease parents' concerns by bettering infant care services.

Yi, J. and Zhang, J. (2010). The effect of house price on fertility: evidence from Hong Kong. *Economic Inquiry*, **48** (3), 635–650.

YuWa Population Research Institute (2022). *Report on the cost of raising a child in China (in Chinese)*. Tech. rep.

Zhang, J. (2017). The evolution of China's one-child policy and its effects on family outcomes. *Journal of Economic Perspectives*, **31** (1), 141–160.

# 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 $\alpha$ 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 $\alpha$ is independent of initial capital stock. Moreover, as in Becker and Barro (1988) and Barro and Becker (1989), 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 $\alpha^h$-type household is gradually rising and reaches one for $j \to +\infty$, while the fraction of the $\alpha^l$-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 $\alpha^h$-type household is increasing while the size of the $\alpha^l$-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 (0.102) | 0.016 (0.110) | 0.209 (0.225) | −0.113 (0.095) | 0.220 (0.176) | 0.030 (0.129) | 0.021 (0.103) | −0.001 (0.112) | 0.030 (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.

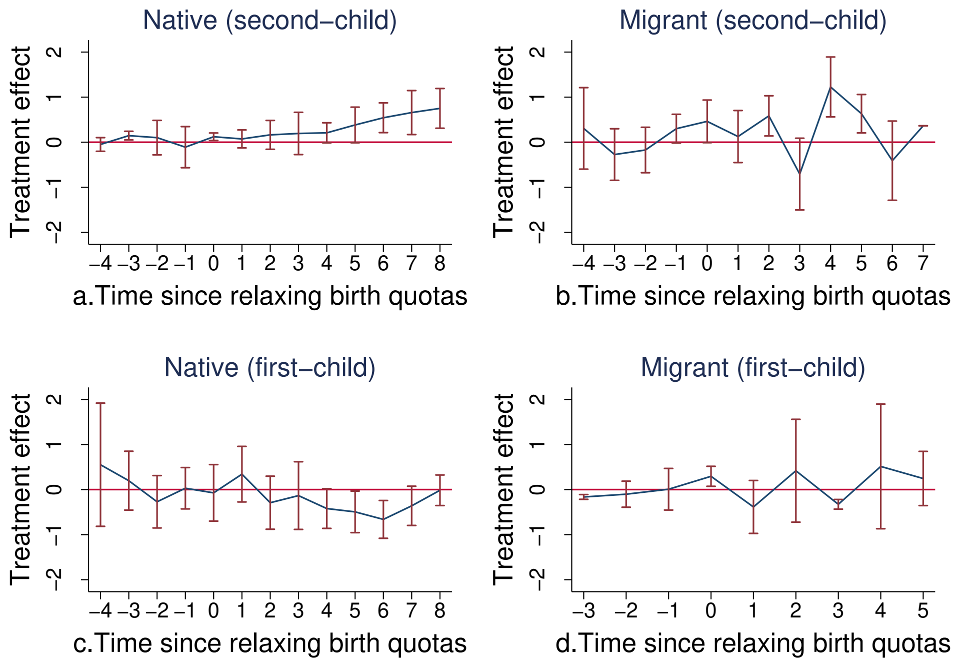


Figure B.1: Migration status and the effects of relaxing birth quotas on second-child and first-child births

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).

# Appendix C

## C.1 Recoding the timing of relaxing birth quotas

As already mentioned, using monthly birth data for estimation implies that assumptions are required for the minimum time it takes for the relaxation of birth quotas to noticeably affect local second-child (first-child) births. There are two main reasons. First, without recoding the timing of relaxing birth quotas, we cannot estimate the effect of relaxing birth quotas on childbearing

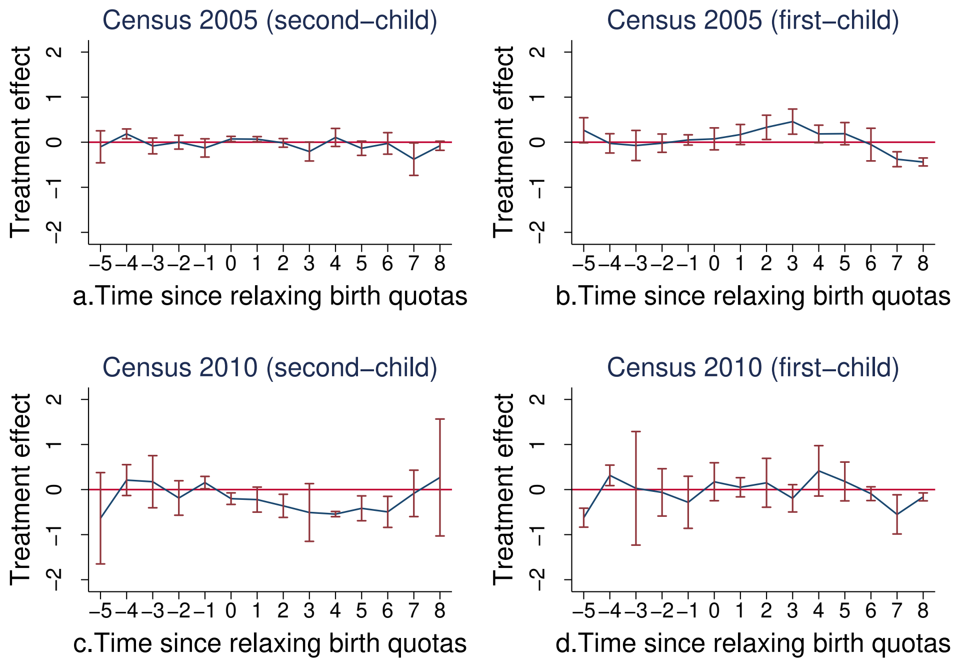


Figure B.2: Placebo analysis on the impact of relaxing birth quotas using censuses 2005 and 2010

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).

decisions in a DID setup. This is because the census data provide monthly births from November 2014 to October 2015, while most provinces relaxed birth quotas before November 2014. In other words, the timing of monthly births mostly does not overlap with the timing of relaxing birth quotas. As a result, the untreated provinces will be too few to estimate the effect of relaxing birth quotas on childbearing decisions in a DID setup. Second, even if we have early data on monthly births that overlap the timing of relaxing birth quotas, estimating the effect of relaxing birth quotas on childbearing decisions without recoding the timing of relaxing birth quotas remains problematic because its effect on childbearing decisions would be largely contaminated

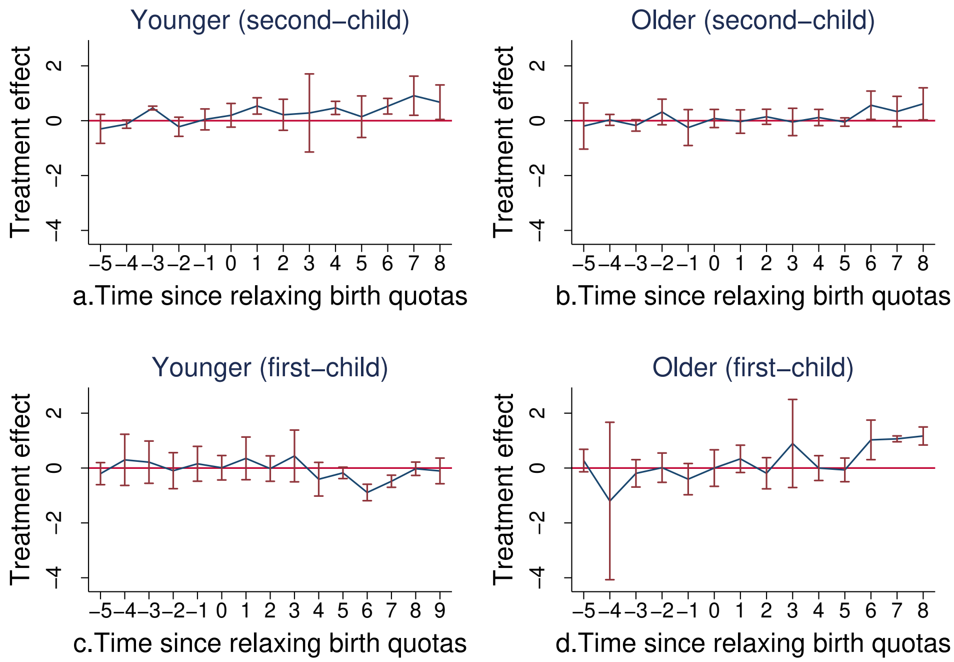


Figure B.3: Childbearing age and effects of relaxing birth quotas on second-child and first-child births.

Notes: This graph shows the impact of relaxing birth quotas on second-child births of (a) 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 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 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).

by biological effects due to the fertility process. We further discuss how to recode the timing of relaxing birth quotas 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

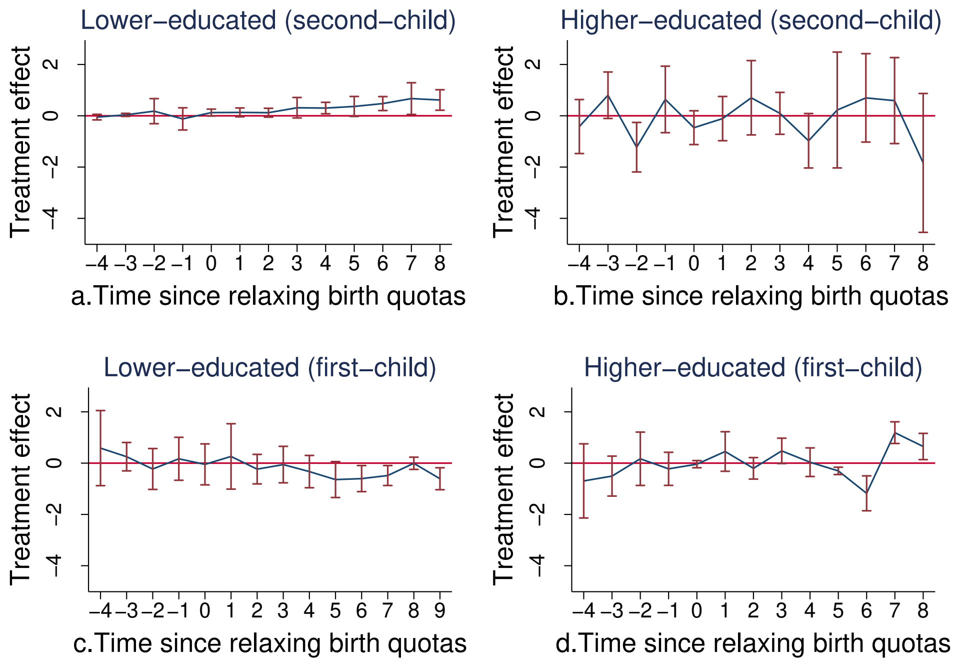


Figure B.4: Education and effects of relaxing birth quotas on second-child and first-child births.

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).

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 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.

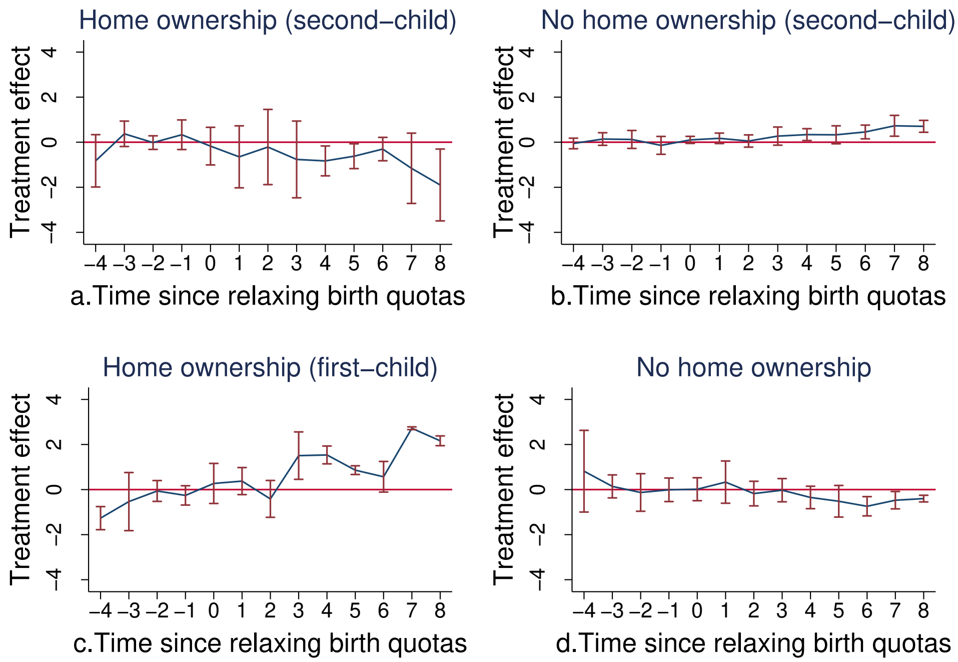


Figure B.5: Homeownership and the effects of relaxing birth quotas on second-child and first-child births.

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).

## 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 birth quotas on local births.

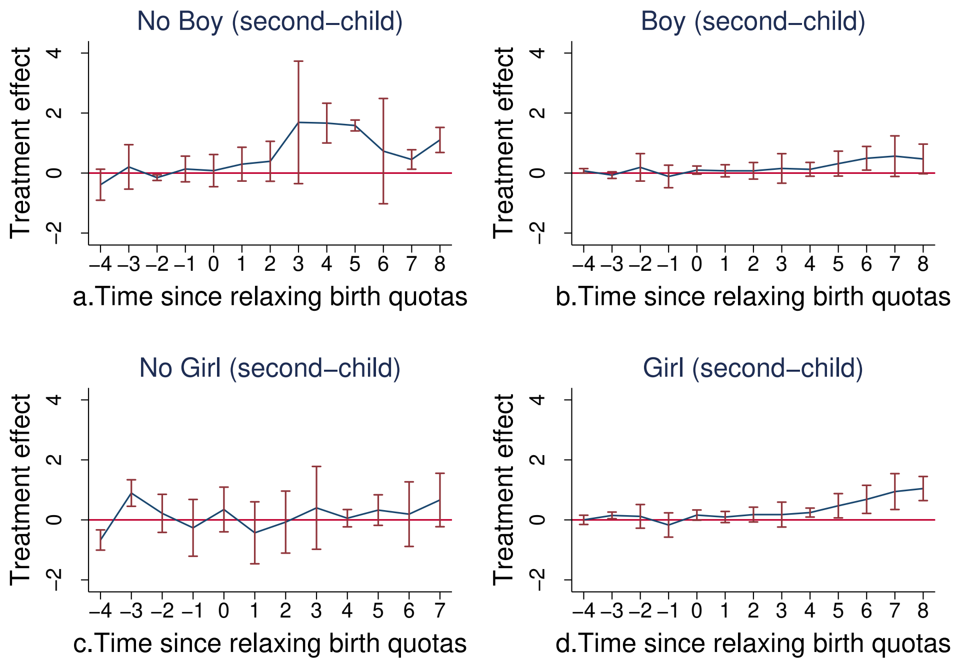


Figure B.6: Gender of children and the effects of relaxing birth quotas on second-child.

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).

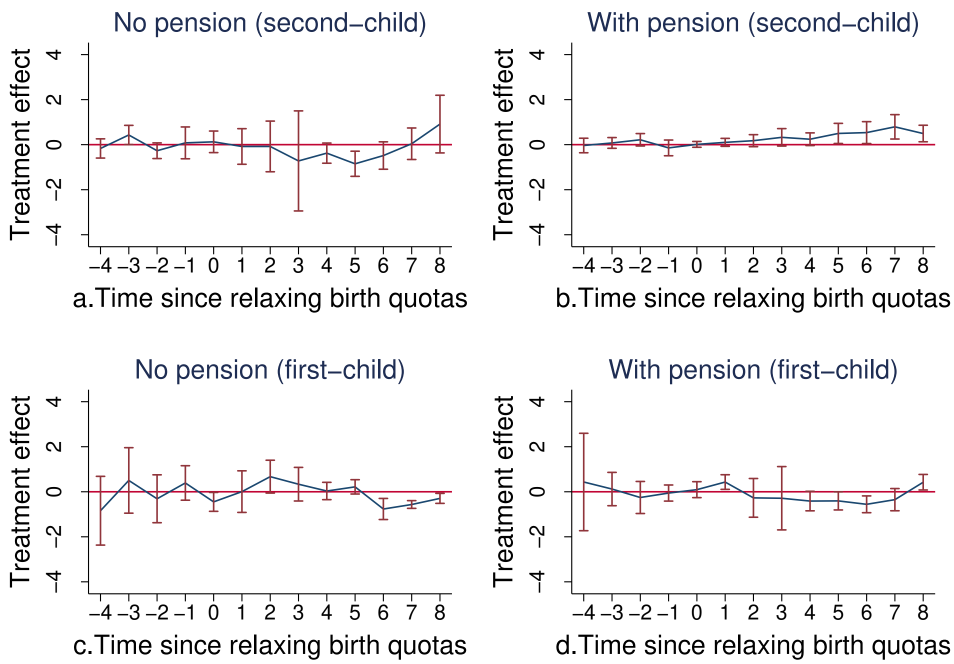


Figure B.7: Access to pension and the effects of relaxing birth quotas on second-child and first-child births.

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).

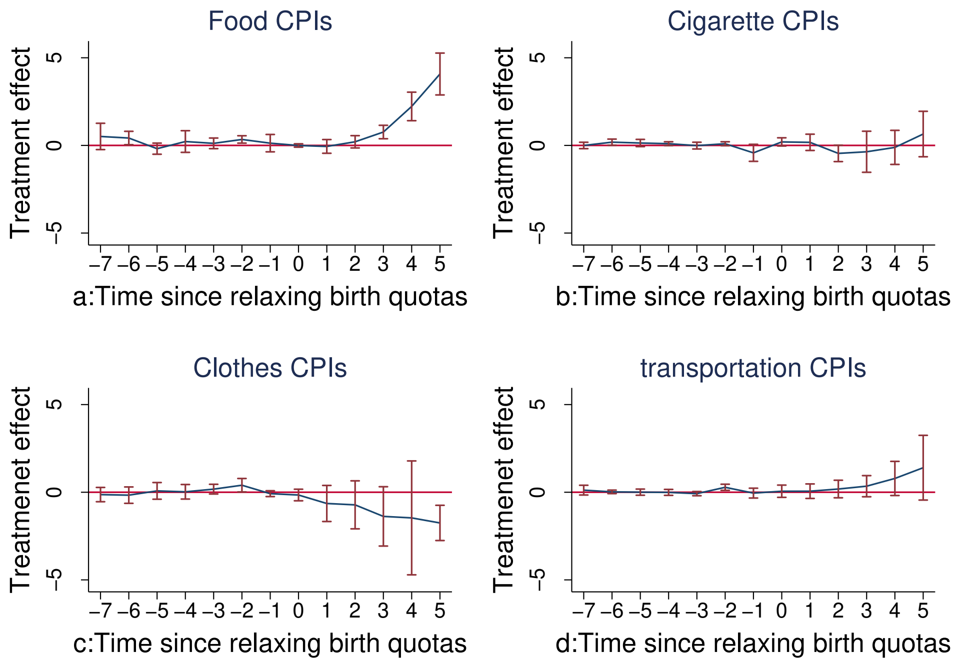


Figure B.8: Effects of relaxing birth quotas on CPIs by expenditure categories

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).

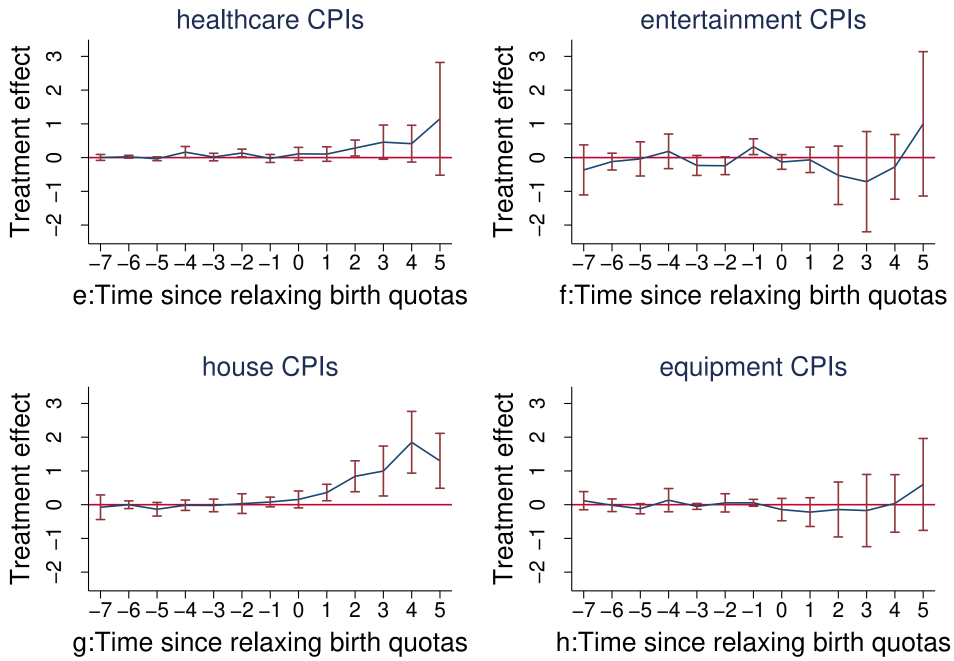


Figure B.9: Effects of relaxing birth quotas on CPIs by expenditure categories (cont.)

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).

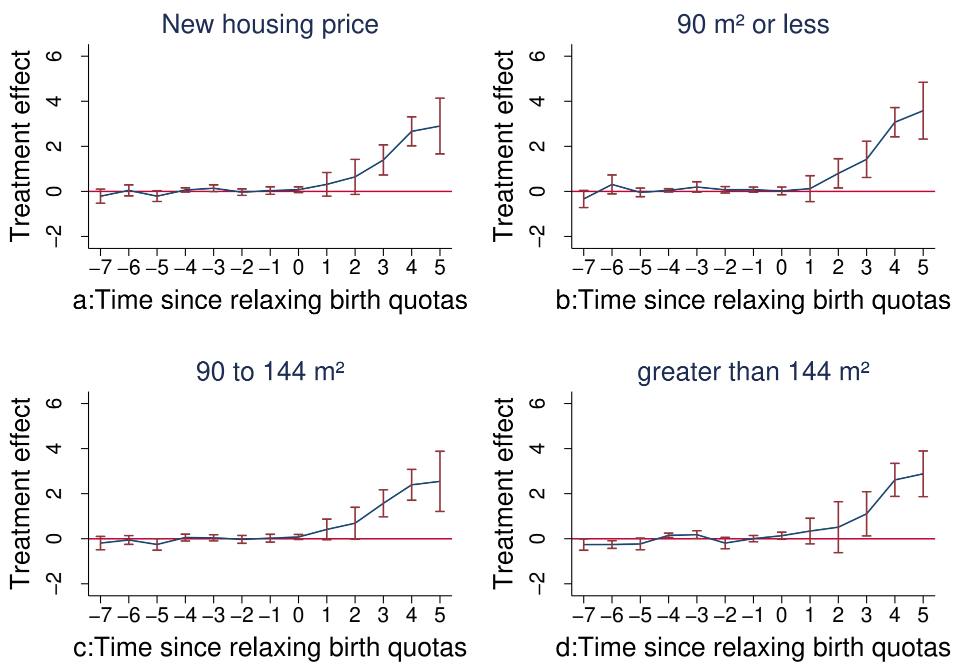


Figure B.10: Heterogeneous effects of relaxing birth quotas on sale prices of newly constructed residential buildings

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).

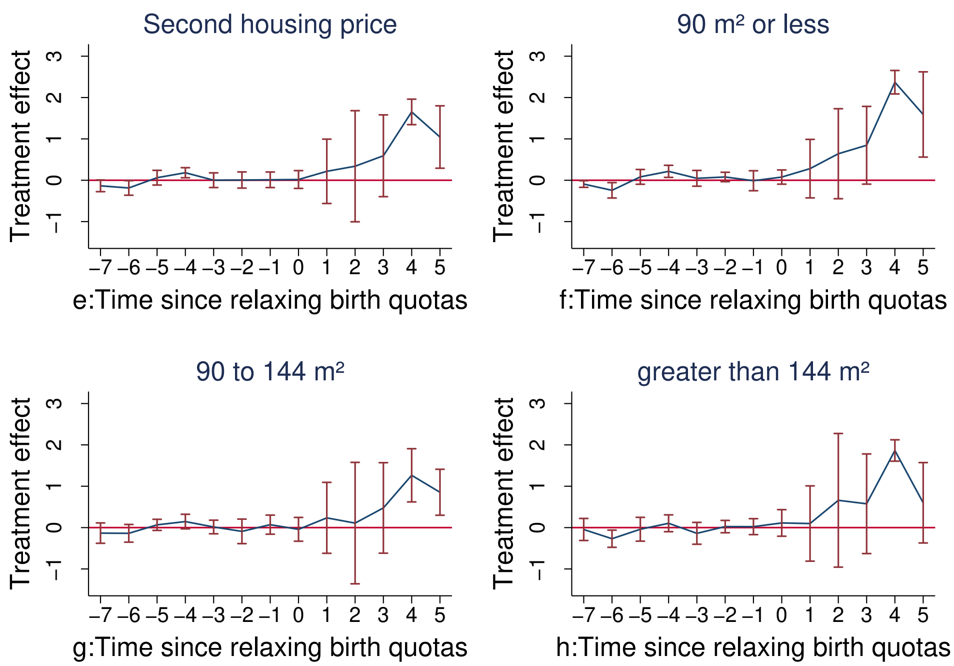


Figure B.11: Heterogeneous effects of relaxing birth quotas on sale prices of second-hand residential buildings.

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).

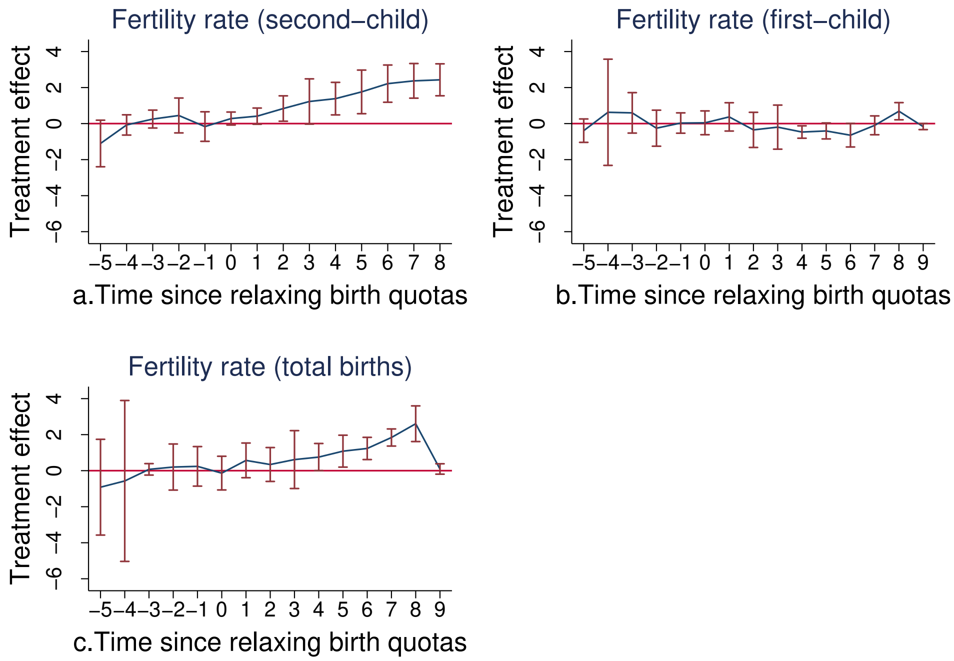


Figure B.12: Impact of relaxing birth quotas on second-child, first-child and total births fertility rates

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).