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The Unintended Consequences of Relaxing Birth Quotas: Theory and Evidence

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The Unintended Consequences of Relaxing Birth Quotas: Theory and Evidence

Zhangfeng Jin; Shiyuan Pan; Zhijie Zheng*

Abstract: This study examines the consequences of relaxing birth quotas by exploiting an exogenous two-child policy adopted by local Chinese governments on different dates. Using China's 2015 population census combined with a difference-in-differences framework, we find that the adoption of a two-child policy substantially increases the number of second-child births. The impact of the policy is more pronounced among couples who have higher fertility preferences and who are less sensitive to child-rearing costs. At the same time, this policy substantially decreases the number of first-child births. Child-rearing costs are a likely underlying mechanism for this decrease. All of these findings are in line with an extended Barro-Becker model.

Keywords: two-child policy; fertility preferences; child-rearing cost; childbearing decision; China

JEL Codes: J13; J18; H23;

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1 Introduction

The one-child policy that China imposed in 1979 was one of its most influential and contentious population control policies. In the four decades that followed it, China witnessed a significant decline in its fertility rate, with total births per woman in China falling from 2.75 in 1979 to 1.69 in 2018 (World Bank 2019). The apparent success of the policy induced Chinese authorities to relax the one-child policy and to permit eligible couples to give birth to a second child (Li et al. 2019). In 2016, the country adopted a universal two-child policy, permitting all couples to have a second child. Unexpectedly, three years after this shift, China's fertility rate hit a record low. According to China's National Bureau of Statistics, the annual births decreased from 16.87 million in 2014 to 14.65 million in 2019, in which the percentage of first-child births among total births decreased from 60 to 40.5 (National Bureau of Statistics of China 2020). The causes of this decline in births, particularly in births of first children, remain unclear. Other developing countries, even without China's stringent child-limitation policies, have also experienced declines.[1] A better understanding of whether, to what extent, and how relaxing birth quotas affects childbearing decisions would help China, as well as other countries, to adopt policies to boost birth rates.

In the first part of the paper, we extend the canonical Barro-Becker model (Becker and Barro 1988; Barro and Becker 1989) by introducing mandatory family planning policies and heterogeneous households in terms of fertility preferences. In this dynamic framework, we find that relaxing birth quotas raises birth rates unambiguously in the long term but may increase or decrease them in the short term. On the one hand, intuitively, relaxing birth quotas encourages households with higher fertility preferences to have more children. On the other hand, this policy raises child-rearing costs in the short term, which in turn discourages households' fertility

[1] According to World Bank (2019), from 1979 to 2018, total births per woman decreased from 4.12 to 1.73 in Brazil and from 5.0 to 2.4 and South Africa. Total births per woman in developed countries also declined, but only slightly, falling, for example, from 1.81 to 1.73 in the United States and from 1.86 to 1.68 in the United Kingdom.

willingness universally. The model therefore predicts increased second-child births and fewer first-child births in the short term.

In the second part of the paper, we confront our theoretical predictions with data. Estimating the causal impacts of relaxing birth quotas raises a number of challenges. The first challenge is to find defensible ways to identify the causal impacts of family planning policies (Zhang 2017).[2] In this article, we take advantage of a two-child policy adopted by local Chinese governments on different dates in 2014, which permitted eligible couples to give birth to a second child.[3] Using data about the implementation of this policy, we estimate its impact using a difference-in-differences approach. The validity of this approach is based on a common-trend assumption that childbearing outcomes in provinces affected by the two-child policy would have been similar to trends in unaffected provinces if the two-child policy had not been implemented. This assumption implies, more specifically, that any unobserved province-specific shocks to childbearing outcomes over time are uncorrelated with the timing of the adoption of the two-child policy—which is likely to be true because the two-child policy was adopted by local governments on different dates within a relatively short period (i.e., the same calendar year).

Measuring childbearing outcomes presents another empirical challenge. Most existing studies make use of low-frequency annual birth data to investigate the impact of family planning policies. An obvious disadvantage is that the annual birth data are more likely to be contaminated by simultaneous social and economic changes that are not fully observable, making it difficult to identify the causal impact of family

[2] According to Zhang (2017), existing studies adopt four different empirical approaches whose validity is based on different assumptions.

[3] China's constitution provides for three de jure levels of government. Currently, however, there are five practical (de facto) levels, consisting of (1) provincial government (including province, autonomous region, municipality, and special administrative region), (2) prefecture, (3) county, (4) township, and (5) village. In this article, for the sake of simplicity, we treat "local governments" and "provincial-level governments" as such as interchangeable. Later, we introduce more details on the adoption of the two-child policy.

planning policies. Moreover, the two-child policy studied in this article was adopted by local governments on different dates within the same calendar year, implying that we cannot identify the impact of the adoption of the two-child policy through annual birth data. In this study, we take advantage of China's 2015 population census to construct the number of monthly births from November 2014 to October 2015 in each province. Importantly, the population census data also allow us to distinguish between first- and second-child births at the monthly level.

To estimate the causal impact of the adoption of the two-child policy, we make assumptions about the minimum time it takes for local couples to give birth to a second child after the policy is adopted. This is intuitive, to some extent, because giving birth to a child involves a lot of uncertainties and is also time consuming. Moreover, under the two-child policy, prospective parents are required to apply for a birth permit in advance to give birth to a second child, and the procedure for obtaining birth permits is complex and can last more than one month. In this study, we assume that giving birth to a second child generally takes three steps. The first is the application process. Based on the official procedure of granting birth permits, we assume that it takes about 1 month on average to obtain a birth permit from the local authorities. The second step is conception. Childbearing is a complicated task involving repetitive household bargaining and uncertainty (Doepke and Tertilt 2018; Doepke and Kindermann 2019). In the benchmark model, we assume that it takes about 6 months on average to prepare for pregnancy and to get pregnant with a second child, in addition to the time required to obtain the birth permit.[4] The final step is the childbearing itself. We assume that the pregnancy lasts about 10 months on average. All together, we assume that it takes a total of about 17 months, on average, for the adoption of the two-child policy to affect local second-child births. For example, for a

[4] Another important factor for this assumption is the current capability of giving birth to a child in light of the rising infertility rate in China (Yiwen 2018). A recent study shows that 25% of couples actively attempting to become pregnant suffer infertility in China (Zhou et al. 2018). Nevertheless, as a robustness check, we will use an alternative assumption for the time it takes to prepare and get pregnant with a second child (e.g., 1 month).

two-child policy that was launched in February 2014, we assume that it will begin to affect local second-child births, to a large extent, starting in July 2015.

To cope with the concern that heterogeneous treatment effects (e.g., by groups or over time) may make estimates from a two-way fixed effects model misleading, we exploit an estimator proposed by de Chaisemartin and D'Haultfoeuille (2020) to estimate the impact of the adoption of the two-child policy, which is robust when there are heterogeneous treatment effects. We also conduct several robustness checks to guarantee that our estimated impact is driven by the two-child policy but not by other factors, such as pretreatment trends differences, the timing pattern of the adoption of the two-child policy, and the assumption about the time it takes for the adoption of the two-child policy to affect local second-child births. Finally, we conduct two additional placebo analyses by focusing on individuals less affected by the adoption of the two-child policy, and by analyzing alternative census data reflecting a period prior to the adoption of the two-child policy.

Our main findings are as follows. First, we find that second-child births increase substantially after the two-child policy is adopted, suggesting that it does encourage some couples to give birth to a second child. In particular, we find that the impact is heterogeneous across groups, and is more pronounced for couples who have higher fertility preferences and who are less sensitive to child-rearing costs. Second, the same policy significantly reduces first-child births, a decrease for which child-rearing costs are a likely underlying mechanism. We provide additional evidence suggesting that the number of marriages increases significantly after the adoption of the two-child policy, mitigating against the possibility that the decrease in first-child births after the adoption of the two-child policy is driven by changes in the marriage market. All of these findings are in line with the theoretical implications of an extended Barro-Becker model.

This study makes several contributions to the literature. First, it sheds light on the effectiveness of family planning policies. Numerous studies have examined the effectiveness of family planning policies in different countries but have not reached a consensus (Olivetti and Petrongolo 2017; Farré and González 2019; de Silva and Tenreyro, forthcoming). While the policy impact varies across countries and individuals, most existing studies show that the effect of family planning policies on birth rates is limited (Freedman 1997; Gauthier 2007; Miller 2009; Bailey 2012; Miller and Babiarz 2016; De Silva and Tenreyro 2017). In contrast to family planning policies in other countries, which are usually voluntary, China provides a prime case for studying the impact of mandatory family planning policies. A precise evaluation of mandatory family planning policies could provide an upper bound for the impact of family planning policies in general (Miller and Babiarz 2016).

The second contribution is to shed light on the child-rearing cost effect (or price effect) of family planning policies in China. Existing studies do not reach a consensus on the relationship between family planning policies and fertility changes (Merli and Smith 2002; Peng 2010; Babiarz et al. 2018; Chen and Huang 2020). In particular, most existing studies focus on the direct impact of family planning policies (Chen and Huang 2020). A few also find that family planning policies are responsible for the rising sex ratio in China (Ebenstein 2010; Babiarz et al. 2018). Nevertheless, much less is known about the indirect effects of family planning policies (Medoff 2008; Thévenon and Gauthier 2011; Huang, Lei, and Zhao 2016; Gathmann and Sass 2018). Our findings suggest that failing to consider the child-rearing cost effect (or price effect) of family planning policies may lead to misleading policy implications (e.g., continue relaxing birth quotas universally rather than giving priority to reducing child-rearing costs for the young couples).

Finally, this article contributes to the theoretical analysis of relaxing birth quotas. Beginning with the seminal works of Becker and Barro (1988) and Barro and Becker

(1989), many studies have discussed endogenous childbearing decisions using macroeconomic models (Becker, Murphy, and Tamura 1990; Boldrin and Jones 2002; R. Barro and Sala-i-Martin 2004; Doepke 2004; Manuelli and Seshadri 2009; Bar and Leukhina 2010).[5] More recently, using the Barro-Becker model as a benchmark framework, a number of studies have further investigated the relationship between demographic transition and economic development (Bar and Leukhina 2010; Jones and Schoonbroodt 2010; Ashraf, Weil, and Wilde 2013; Lanz, Dietz, and Swanson 2017). That said, the way childbearing decisions respond to policies of relaxing birth quotas remains theoretically unclear. In this article, we incorporate mandatory family planning policies, which are known as birth quotas and regulate the maximum number of children each couple is permitted to give birth to, into a Barro-Becker model with heterogeneous fertility preferences, and use the extended Barro-Becker model to analyze the short- and long-term impacts of relaxing birth quotas on childbearing decisions. This line of inquiry is complementary to existing theoretical analyses.[6]

The remainder of this article is structured as follows. The next section introduces some background on mandatory family planning policies in China. Section 3 proposes an extended Barro-Becker model with mandatory family planning policies. Section 4 introduces key variables and a description of the data. Section 5 presents a model specification and identification strategy. Section 6 reports empirical findings and the final section concludes.

2 Background

The mandatory one-child policy in China, formally introduced in the late 1970s, strictly controlled, through birth permits, the number of children each couple was permitted to give birth to. As a result, most couples were permitted to give birth to one

[5] An excellent literature review can be found in Jones, Schoonbroodt, and Tertilt (2008).

[6] A thorough discussion of the relationship between family planning policies and fertility rates can be found in De Silva and Tenreyro (2017).

child. Enforcement of the one-child policy was the responsibility of the then-powerful National Family Planning Commission. Couples violating the one-child policy were subject to high penalties. Nevertheless, the system of penalties was inconsistently applied, often at the discretion of local officials, with wide variations across the country (Zeng and Hesketh 2016). Before implementing the one-child policy, the Chinese authorities adopted a mostly voluntary "later-longer-fewer" policy in the 1970s (Chen and Huang 2020).

The one-child policy was implemented more strictly for couples from urban areas than for their counterparts from rural areas. In fact, beginning in 1984, couples from rural areas were permitted to give birth to a second child in most provinces if their first child was female; this was the so-called 1.5-child policy. Two or more children were allowed for ethnic minorities. Since the early 2000s, the one-child policy has been relaxed in a piecemeal fashion to address the falling fertility rate. Since 2011, all provinces have permitted eligible couples, both of whose partners were only children, to give birth to a second child. In November 2013, a new policy began allowing couples in which at least one of the partners was an only child to give birth to a second child (Ouyang 2013). However, eligible couples were still required to apply for birth permits in advance in order to do this. Then, in December 2015, the universal two-child policy was announced, which allows all couples to give birth to a second child (Li et al. 2019). Importantly, the universal two-child policy no longer requires couples to apply for a birth permit in order to legally give birth to a second child. **Figure A1** shows the number of first-child and second-child births from 2000 to 2015. We find that the number of first-child births decreased gradually from 2000 to 2010, while the number of second-child births was relatively stable during the same period. In 2015, however, the number of first-child births decreased substantially along with a significant increase in the number of second-child births. The rest of this article explores, both theoretically and empirically, whether and how the adoption of a two-child policy contributes to these changes in second-child and first-child births.

3 Theoretical Framework

In this section, we extend the Barro-Becker framework (Becker and Barro 1988; Barro and Becker 1989) by incorporating mandatory family planning policies, which regulate birth quotas, and heterogeneous households into a dynamic general equilibrium model, and use the extended model to analyze the theoretical implications of relaxing birth quotas.

3.1 Households

We assume that the economy consists of $N(0)$ heterogeneous households and that each household begins with one person at time $t = 0$. Households are heterogeneous in fertility preference, as is captured by the parameter $\alpha > 0$.[7] The probability density and the cumulative distribution functions of α are $f(\alpha)$ and $F(\alpha)$, respectively, where $\underline{\alpha} < \alpha < \overline{\alpha}$, $F(\underline{\alpha}) = 0$ and $F(\overline{\alpha}) = 1$. Each person lives for two periods, childhood and adulthood. In every period, each adult chooses the desired number of children and raises them in that period.[8] The endogenous choice of fertility determines family size as well as the total population in the economy.

Suppose that the number of adults in household α is 1 at time $t = 0$. In the beginning of $t = 0$, an adult chooses the desired number of children $n(\alpha, 0)$ and raises them in that period. At time $t = 1$, the $n(\alpha, 0)$ children in household α grow up to be adults, and decide to give birth to and raise $n(\alpha, 1)$ children in that period. Note that the 1 adult in household α has passed away at time $t = 0$. Therefore, the number of adults in household α is $n(\alpha, 0)$ at time $t = 1$, and the number of children in household α is $n(\alpha, 0) \cdot n(\alpha, 1)$. The fertility rate of household α is

[7] In our model, the initial wealth also differs among households. Differences in initial wealth, however, only affect the initial consumption and family size but do not affect fertility decisions over time. More discussions of the impact of heterogeneous household wealth on fertility can be found in Bosi, Boucekkine, and Seegmuller (2011).

[8] For simplicity, our theoretical model considers neither gender differences nor the fertility process.

denoted as $n(\alpha, 1)$. We assume that fertility preferences are the same across members within each household. By the same token, at time t, the total number of adults in household α is $n(\alpha, 0) \cdot n(\alpha, 1) \cdots n(\alpha, t-1) = \prod_{t=0}^{t-1} n(\alpha, t) = N(\alpha, t)$, the number of newborn children is $n(\alpha, 0) \cdot n(\alpha, 1) \cdots n(\alpha, t-1) \cdot n(\alpha, t) = N(\alpha, t+1)$, and the corresponding fertility rate of household α is $n(\alpha, t)$.

The utility function of each adult at time t in household α is given by

$$U(\alpha, t) = v[c(\alpha, t)] + a[n(\alpha, t)] \cdot n(\alpha, t) \cdot U(\alpha, t+1), \quad (1)$$

where $c(\alpha, t)$ is the consumption for the adult and $v(\cdot)$ is the utility derived from the consumption. Let $v[c(\alpha, t)] = [c(\alpha, t)]^{\sigma}/\sigma$, where $\sigma < 1$, $1/(1-\sigma)$ is the elasticity of intertemporal substitution in consumption. $U(\alpha, t+1)$ is the utility attained by each child. $a[n(\alpha, t)]$ measures the degree of altruism toward each child,[9] and is given by

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

We assume that $0 < \alpha < 1$ so that the utility for each household is bounded.[10] Moreover, as in Becker and Barro (1988) and Barro and Becker (1989), $\varepsilon + \sigma < 1$ must hold as well to ensure that the utility is bounded. The higher the degree of altruism, the more likely that an adult will give birth to more children. Equation (2) suggests that, given the utility function of each child $U(\alpha, t+1)$, the degree of altruism decreases with the number of children. By substituting out for $U(\alpha, t+1)$ (and $U(\alpha, t+2)$, etc.) in equation (1), the utility function for household α at time $t = 0$ can be expressed as

$$U(\alpha, 0) = \sum_{t=0}^{\infty} \alpha^t \cdot [N(\alpha, t)]^{1-\varepsilon} \cdot [c(\alpha, t)]^{\sigma}/\sigma \quad (3)$$

where $N(\alpha, t) = \prod_{t=0}^{t-1} n(\alpha, t')$ is the size of household α at time t, and $N(\alpha, 0) = 1$.

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[9] Parents with different degrees of altruism will have different fertility preferences. In other words, the degree of altruism reflects the fertility preference. Fertility preferences are affected by such factors as individual characteristics and culture.

[10] More discussions can be found in the proof for Proposition 1 in the appendix.

The labor market is competitive. Each household owns 1 unit of time in each period and earns the same wage rate $\mathrm{w}(t)$. Thus, $\mathrm{w}(t)$ is also the total labor income at time t. Parents leave a bequest of capital $k(\alpha, t+1)$ at time $t+1$ to each child born at time t. Capital $k(\alpha, t)$ earns rent at the rate $r(t)$ at time t. An adult of household α at time t spends his earnings and inheritance $w(t) + [1 + r(t)] \cdot [k(\alpha, t)]$, on own consumption $P(t) \cdot c(\alpha, t)$, on bequests to children $n(\alpha, t) \cdot k(\alpha, t+1)$, and on costs of raising children. $P(t)$ is the price of final goods (or services). We assume that raising a child costs $\beta(\alpha, t)$ (in units of real income), so that $n(\alpha, t) \cdot \beta(\alpha, t)$ is the total costs of raising children to adulthood. The budget constraint for an adult in household α at time t is

$$w(t) + [1 + r(t)] \cdot k(\alpha, t) = P(t) \cdot c(\alpha, t) + n(\alpha, t) \cdot [\beta(\alpha, t) + k(\alpha, t+1)] \quad (4)$$

And the budget constraint for an adult in household α at time 0 is

$$\frac{w(0)}{1+r(0)} + k(\alpha, 0) = \frac{P(0) \cdot c(\alpha,0)}{1+r(0)} + \frac{n(\alpha,0) \cdot [\beta(\alpha,0)+k(\alpha,1)]}{1+r(0)} \quad (5)$$

The dynastic budget constraint, which equates the present value of all resources to the present value of all expenditures, is

$$k(\alpha, 0) + \sum_{t=0}^{\infty} d\,(t) N(\alpha, t) w(t) = \sum_{t=0}^{\infty} d\,(t) \cdot [N(\alpha, t) \cdot P(t) \cdot c(\alpha, t) + N(\alpha, t+1) \cdot \beta(\alpha, t)] \quad (6)$$

where $d(t) = \prod_{t'=0}^{t} [1 + r(t')]^{-1}$.

The optimization problem is choosing consumption $c(\alpha, t)$ and the number of children $n(\alpha, t)$ to maximize utility $U(\alpha, 0)$ in equation (3), subject to the budget constraints in equation (6). The Lagrange function is :

$$L = \sum_{t=0}^{\infty} \alpha^t \cdot [N(\alpha, t)]^{1-\varepsilon} \cdot [c(\alpha, t)]^{\sigma} / \sigma + \lambda(\alpha, 0) \cdot [k(\alpha, 0) + \sum_{t=0}^{\infty} d(t) \cdot N(\alpha, t) \cdot w(t) - \sum_{t=0}^{\infty} d(t) \cdot N(\alpha, t) \cdot P(t) \cdot c(\alpha, t) - \sum_{t=0}^{\infty} d(t) \cdot N(\alpha, t+1) \cdot \beta(\alpha, t)]$$

The first-order conditions are

$$\frac{\partial L}{\partial c(\alpha,t)} = 0 \Leftrightarrow \alpha^t \cdot [N(\alpha,t)]^{1-\varepsilon} \cdot [c(\alpha,t)]^{\sigma-1} = \lambda(\alpha,0) \cdot d(t) \cdot N(\alpha,t) \cdot P(t) \quad (7)$$

$$\frac{\partial L}{\partial N(\alpha,t)} = 0 \Leftrightarrow \frac{(1-\varepsilon)\cdot\alpha^t\cdot[c(\alpha,t)]^{\sigma}}{\sigma[N(\alpha,t)]^{\varepsilon}} = \lambda(\alpha,0) \cdot [d(t)P(t)c(\alpha,t) + d(t-1)\beta(\alpha,t-1) - d(t)w(t)] \quad (8)$$

We can rewrite equation (7) as

$$\alpha^{t+1} \cdot [N(\alpha,t+1)]^{1-\varepsilon} \cdot [c(\alpha,t+1)]^{\sigma-1} = \lambda(\alpha,0) \cdot d(t+1) \cdot N(\alpha,t+1) \cdot P(t+1) \quad (9)$$

Combining equations (7) and (9) yields

$$\left[\frac{c(\alpha,t+1)}{c(\alpha,t)}\right]^{1-\sigma} = \frac{\alpha\cdot[1+r(t+1)]}{[n(\alpha,t)]^{\varepsilon}} \cdot \frac{P(t)}{P(t+1)} \quad (10)$$

Using the form of $\lambda(\alpha,0)$ from equation (7), we can rewrite equation (8) as

$$\frac{(1-\varepsilon)\cdot\alpha^t\cdot c(\alpha,t)]^{\sigma}}{\sigma[N(\alpha,t)]^{\varepsilon}} = \frac{\alpha^t\cdot[N(\alpha,t)]^{1-\varepsilon}[c(\alpha,t)]^{\sigma-1}}{P(t)N(\alpha,t)} \cdot \left[\frac{d(t-1)\beta(\alpha,t-1)}{d(t)} + P(t)c(\alpha,t) - w(t)\right]$$

$$\Leftrightarrow c(\alpha,t) = \frac{\sigma}{1-\varepsilon-\sigma} \cdot \frac{1}{P(t)}\{\beta(\alpha,t-1)\cdot[1+r(t)] - w(t)\} \quad (11)$$

Equation (11) shows that at time $t \geq 1$, the consumption per capita of household α is independent of the initial capital stock. Moreover, as in Becker and Barro (1988) and Barro and Becker (1989), we suppose that children are a net financial burden to altruistic parents: the costs of raising an additional child would exceed that child's lifetime earnings. This implication also implies $\sigma/(1-\varepsilon-\sigma) > 0$.[11] Equations (10) and (11) solve for the optimal number of children borne, such that:

$$[n(\alpha,t)]^{\varepsilon} = \{\alpha\cdot[1+r(t+1)]\}\cdot[\frac{P(t)}{P(t+1)}]^{\sigma}\cdot\left\{\frac{\beta(\alpha,t-1)\cdot[1+r(t)]-w(t)}{\beta(\alpha,t)\cdot[1+r(t+1)]-w(t+1)}\right\}^{1-\sigma} \quad (12)$$

Equation (12) shows that the optimal number of children borne for each adult of household α depends on the degree of altruism, the price of final goods, the interest rate, the wage rate and the costs of raising children in different periods $\beta(\alpha,t-1)$ and $\beta(\alpha,t)$. Moreover, the fertility rate of household α is independent of the initial capital stock of the household.

For simplicity, we assume that the costs of raising children are the same across households in the same period, namely $\beta(\alpha,t) = \beta(t)$. According to equation (12),

[11] Since $0 < \varepsilon < 1$, to ensure that $\sigma/(1-\varepsilon-\sigma) > 0$ holds, $\sigma > 0$ must also hold.

differences in the desired number of children borne are largely driven by fertility preference α. We further assume that raising a child requires $\mu > 0$ unit of final goods.[12] Therefore, the costs of raising a child are

$$\beta(t) = \mu P(t) \tag{13}$$

3.2 Firms

The production sector is competitive. A representative firm produces final goods using labor and capital. The production function is

$$Y(t) = [L(t)]^{\eta}[K(t)]^{1-\eta} \tag{15}$$

where $Y(t)$ is output of final goods, $K(t)$ is capital, and $L(t)$ is labor. Solving the profit-maximizing problems gives rise to the demand functions of labor and capital:

$$L(t) = \eta P(t)\, Y(t)/w(t) \tag{16}$$

$$K(t) = (1-\eta) P(t)\, Y(t)/r(t) \tag{17}$$

We assume that the wage rate is the numeraire in the economy, such that $w(t) = 1$.[13] The price and interest rate are expressed as

$$P(t) = \frac{[k(t)]^{\eta-1}}{\eta} \tag{18}$$

$$r(t) = \frac{1-\eta}{[\eta k(t)]} \tag{19}$$

where $k(t) \equiv \frac{K(t)}{L(t)}$ is the capital stock per capita.[14]

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[12] First, the assumption that raising a child does not require an adult to spend time helps ensure the uniqueness of the steady state equilibrium. More importantly, this assumption is enough to capture the impact of a change in mandatory family planning policies on the costs of raising children.

[13] Selecting the price of wage rate as the numeraire brings some convenience in this model. In the current setting, changes in family planning policies can affect the price of final goods and thus the costs of raising a child, which captures the child-rearing costs channel described in the empirical analysis. If, instead, choosing the price of final goods as the numeraire, we need to assume that raising a child involves time devoted by adults, in order to capture the child-rearing costs channel. The modification complicates our analysis.

[14] Given the fact that only adults participate into the labor market, it is more precise to consider the variable as capital stock per adult.

3.3 The government

We assume that the government can adopt different mandatory family planning policies, or birth quotas, to affect population growth in the economy. Suppose that the number of children that each household gives birth to at time t satisfies the following condition $n(\alpha,t) \in [\underline{n},\bar{n}]$, where $\underline{n}$ and $\bar{n}$ are the minimum and maximum of the number of children. We assume that the government strictly controls the maximum number of children to whom each household is permitted to give birth, namely $\underline{n} \le n^* \le \bar{n}$. Therefore, the number of children $n(\alpha,t)$ at time t in the economy satisfies the following condition:

$$n(\alpha,t) = \begin{cases} n^*, & n(\alpha,t) \ge n^* \\ n(\alpha,t), & n(\alpha,t) < n^* \end{cases} \tag{20}$$

3.4 Steady state equilibrium

We first define the equilibrium and then investigate the impact that relaxing mandatory family planning policies has on the long-term fertility rate.

In equilibrium, households maximize utility, firms maximize profits, and all markets clear. Define the total number of adults at time t under the government's mandatory family planning policies (n^*) as $\hat{N}(t) = \int_{\underline{\alpha}}^{\dot{\alpha}} N(\alpha,t)dF(\alpha)$. $\hat{N}(t+1) = \int_{\underline{\alpha}}^{\dot{\alpha}} N(\alpha,t+1)dF(\alpha)$ is then the total number of children. The final goods market-clearing condition is

$$c(t)\cdot\widehat{N}(t) + \mu\cdot\widehat{N}(t+1) = Y(t) \tag{21}$$

The left-hand side (LHS) of equation (21) is the aggregate demand of final goods and the right-hand side (RHS) is the aggregate supply of final goods. The labor market-clearing condition is $L(t) = \widehat{N}(t)$, where the LHS denotes the total labor supply, namely the total number of adults, and the RHS is the total labor demand. Define the total capital stocks owned by all households as $K(t) = \int_{\underline{\alpha}}^{\dot{\alpha}} N(\alpha,t)\cdot k(\alpha,t)dF(\alpha)$. The total capital market clearing condition then holds.

In a steady state, the economic growth rate is zero, and neither the price of final goods

nor the interest rate change over time. We rewrite equation (12) as

$$[n(\alpha,t)]^{\varepsilon} = \alpha \cdot (1+r) \tag{22}$$

Equation (22) shows that, in a steady state, for a given interest rate, parents with higher fertility preferences choose to give birth to more children. For simplicity, and without a loss of generality, we assume that α follows a uniform distribution, such that the probability density function and the cumulative distribution function are $f(\alpha) = \frac{1}{\bar{\alpha}-\underline{\alpha}}$ and $F(\alpha) = \frac{\alpha-\underline{\alpha}}{\bar{\alpha}-\underline{\alpha}}$, respectively.[15] The number of children in each household then follows a uniform distribution within the interval $[\underline{n}, n^*]$, where $\underline{n} = [\underline{\alpha} \cdot (1+r)]^{1/\varepsilon}$ and $n^* = [\alpha^* \cdot (1+r)]^{1/\varepsilon}$.[16] Therefore, for a given n^*, the steady state fertility rate can be expressed as

$$B = \int_{\underline{\alpha}}^{\alpha^*} n(\alpha)dF(\alpha) = \int_{\underline{\alpha}}^{\alpha^*} [\alpha \cdot (1+r)]^{1/\varepsilon} dF(\alpha) = \frac{\varepsilon \cdot (1+r)^{1/\varepsilon} \cdot [(\alpha^*)^{(1/\varepsilon+1)} - \underline{\alpha}^{(1/\varepsilon+1)}]}{(\alpha^* - \underline{\alpha})(1+\varepsilon)} \tag{23}$$

This equation shows that mandatory family planning policies affect the total fertility rate through two channels, namely α^* and r.[17] Suppose now that the government relaxes the mandatory family planning policies, that is, it increases the maximum number of children n^*. It will then induce parents with greater fertility preferences ($\alpha > \alpha^*$) to give birth to more children and in turn increase the fertility rate. This policy effect can be confirmed in equation (23) with $\partial B/\partial \alpha^* > 0$. In addition, relaxing the mandatory family planning policies also affects capital per adult by increasing total population, and in turn affects the interest rate r. This channel is considered an income effect.

To further demonstrate the impact of relaxing mandatory family planning policies on the interest rate, we now derive the expression of the steady-state interest rate. The budget constraint on the steady state can be rewritten as

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[15] In this article, we assume that the distribution does not change over time. In other words, the heterogeneity in the fertility preference only exists among households. The fertility preference of members within the same household does not change over time.

[16] It is noteworthy that the interest rate can be affected by family planning policies.

[17] In our model, the policy variable is n^*. According to equation (22), when the interest rate is fixed, the policy variable is equivalent to the fertility preferences chosen by the government.

$$\hat{N}(t) + K(t)[1 + r(t)] = P(t)c(t)\hat{N}(t) + \beta(t)\hat{N}(t+1) + K(t+1),$$

By applying the labor market clearing condition, we can show the budget constraint per adult, which is given by

$$1 + k(t)[1 + r(t)] = P(t)c(t) + \mu P(t)B(t) + B(t)k(t+1) \quad (24)$$

where $\frac{\hat{N}(t+1)}{\hat{N}(t)} = \int_{\underline{\alpha}}^{\alpha^*} n(\alpha, t)dF(\alpha) = B(t)$ is the ratio of children to adults at time t, or the total fertility rate.[18] Using equations (11) (18) (19) (22) and (23), we show that variables $c(t)$, $P(t), r(t), n(\alpha, t)$, and $B(t)$ are functions of capital per adult $k(t)$, which is determined by equation (24). Therefore, in a steady state equilibrium, we have

$$P = 1/(\eta \cdot k^{1-\eta}) \quad (25)$$

$$r = (1 - \eta)/(\eta k) \quad (26)$$

$$c = \frac{\sigma}{1-\varepsilon-\sigma}[\mu(1 + r) - 1/P] \quad (27)$$

$$1 + k(1 + r) = Pc + \mu PB + Bk \quad (28)$$

$$B = \frac{\varepsilon \cdot (1+r)^{1/\varepsilon} \cdot [(\alpha^*)^{(1/\varepsilon+1)} - \underline{\alpha}^{(1/\varepsilon+1)}]}{(\alpha^* - \underline{\alpha})(1+\varepsilon)} \quad (29)$$

Proposition 1 (Existence and uniqueness): This model admits a unique steady-state equilibrium in which the interest rate, total fertility rate, and capital stock per adult are solely determined. (See the proof in the Appendix.)

We now examine the impact of relaxing mandatory family planning policies (an increase in α^*) on the steady-state total fertility rate. First, combining all of the five equations (25)–(29) above, we get the following equation determining the capital stock per adult in a steady state:

$$\left\{k^{1-\eta} + \eta k^{2-\eta} - \frac{\sigma}{1-\varepsilon-\sigma}\left[\mu + \frac{\mu(1-\eta)}{\eta k} - \eta k^{1-\eta}\right]\right\}/\Phi = (\mu + \eta k^{2-\eta})\left(1 + \frac{1-\eta}{\eta k}\right)^{1/\varepsilon} \quad (30)$$

where $\Phi = \frac{\varepsilon[(\alpha^*)^{(1/\varepsilon+1)} - \underline{\alpha}^{(1/\varepsilon+1)}]}{(\alpha^* - \underline{\alpha})(1+\varepsilon)}$. Denote the LHS of equation (30) as $g_1(k)$, and $g_2(k)$ for the RHS. According to Proposition 1, there is a unique intersection point

[18] Given the fact that each person lives for two periods, we also consider the ratio of children to adults as the fertility rate.

between functions $g_1(k)$ and $g_2(k)$, as shown in **Figure 1** (E_1 or E_2). The impact of relaxing mandatory family planning policies is shown in Propostion 2.

Proposition 2: Relaxing mandatory family planning policies increases the total fertility rate in the long term. (See the proof in the Appendix.)

According to **Figure 1**, $g_2(k)$ is a U-shaped function of k. If E_1 is an initial steady state ($k_0 < \bar{k}$), an increase in α^* will shift $g_1(k)$ right, and reach a new steady state at E'_1, which corresponds to lower capital stock per adult. Combining equations (25)–(27), we can show that the price of final goods, consumption per capita, and interest rate will all increase accordingly. According to equation (22), the increased interest rate encourages households to give birth to more children in order to raise lifetime utility.[19] According to equation (29), the total fertility rate increases in response. In other words, relaxing mandatory family planning policies also increases the fertility rate through the income effect. Combining both the policy effect and the income effect, we find that relaxing mandatory family planning policies will raise the total fertility rate B. If E_2 is the initial steady state ($k'_0 > \bar{k}$), an increase in α^* will shift $g_1(k)$ left. As a result, E'_2 is the new steady state. This change will also lead to lower capital stock per adult in a steady state and increase the total fertility rate in the long term.

[Insert Figure 1 about here.]

3.6 Transitional dynamics

In transitional dynamics, relaxing mandatory family planning policies also affects short-term child-rearing costs and in turn affects the fertility rate. Specifically, when the government relaxes mandatory family planning policies at time t and increases

[19] Equation (22) shows that the long-term fertility rate only depends on the degree of altruism and the interest rate. The price of final goods itself does not affect childbearing decisions in the long term.

the maximum number of children n^*, adults with $\alpha > \alpha^*$ are encouraged to give birth to more children. As a result, the capital per adult at time $t+1$ decreases, while the capital per adult at time t does not. Equations (18) and (19) show that the decreased capital per adult will raise the price of final goods and the interest rate at time $t+1$.

By applying equation (12), we can show that the total fertility rate along the transitional path is

$$B(t) = \frac{\varepsilon \cdot [1+r(t+1)]^{1/\varepsilon} \cdot [(\alpha^*)^{(1/\varepsilon+1)} - \underline{\alpha}^{(1/\varepsilon+1)}]}{(\alpha^* - \underline{\alpha})(1+\varepsilon)} \cdot [\frac{P(t)}{P(t+1)}]^{\varepsilon/\sigma}$$

$$\cdot \left\{\frac{\mu P(t-1) \cdot [1+r(t)] - 1}{\mu P(t) \cdot [1+r(t+1)] - 1}\right\}^{(1-\sigma)/\varepsilon}$$

Therefore, in the short term, in addition to inducing adults with higher fertility preferences to have more children, relaxing mandatory family planning policies also affects the fertility rate by affecting the price of final goods and the interest rate. Straightforwardly, the raised price of final goods $P(t+1)$ increases child-rearing costs and tends to reduce the birth rate. As for the fertility impact of changing the interest rate, on the one hand, a higher interest rate $r(t+1)$ encourages households to give birth to more children by creating a higher return on investing in children, which is captured as $[1+r(t+1)]^{1/\varepsilon}$. On the other hand, a higher interest rate discourages parents to give birth to more children because child-rearing costs in the current period increase, which is captured as $\mu P(t)[1+r(t+1)]$. We can therefore see child-rearing costs rise with the relaxation of mandatory family planning policies. Combining the above effects leads to an ambiguous effect on the total fertility rate, either a positive or negative effect. Proposition 3 summarizes the above results.

Proposition 3: Relaxing mandatory family planning polices increases child-rearing costs in the short term but increases or decreases the fertility rate in the short term. (See the proof in the Appendix.)

4 Data, Variables, and Summary Statistics

4.1 Birth outcomes

To construct the variable for birth outcomes, we use a 10% random sample of the 2015 population census conducted by China's National Bureau of Statistics. The census data include about 1.37 million individuals, which is representative of the total population in China. An obvious advantage is that the census data contain rich details on childbirth for women who are between 15 and 50 years of age. Specifically, these data provide information on the timing of childbirth at the monthly level. Based on the timing of childbirth, we define a birth at month t as a woman giving birth to a child at month t. We then aggregate all births at the province level, combined with sampling weights, to calculate the total births that occurred in province s at month t. The census data also include individual characteristics such as age, gender, education, *hukou* status (rural/urban), and other information, such as the number of total births and social security participation, which allows us to calculate monthly births by cohorts. We make use, in particular, of information on the number of total births to distinguish between first-child births and second-child births.[20] As a result, we have panel data on the monthly births of a second (first) child for all 31 provinces from November 2014 to October 2015.

4.2 Relaxation of birth quotas

To quantify the relaxation of birth quotas, we use data about a two-child policy adopted by local Chinese governments on different dates in 2014. The two-child policy permitted eligible couples, in which at least one of the marital partners is an only child, to give birth to a second child after obtaining a birth permit from local authorities. We collect official documents on the adoption of the two-child policy.

[20] We define first birth, or births of a first child, as being experienced by women who do not have any children prior to this birth. We define second birth, or the birth of a second child, as being experienced by women who have had at least one child prior to this birth.

This policy was announced at the national level in November 2013 and was gradually adopted by province-level governments on different dates in 2014. For example, provinces like Zhejiang, Jiangxi, and Anhui adopted the two-child policy in late January 2014. Meanwhile, provinces like Hebei and Shandong adopted the two-child policy in late May 2014. By contrast, provinces like Tibet and Xinjiang adopted the two-child policy in November and December 2014. More details on the timing of the adoption of the two-child policy can be found in **Figure A2**.

Given the fact that parents-to-be are required to apply in advance for a birth permit in order to legally give birth to a second child, and that childbearing is a complicated and time-consuming task involving conception and pregnancy, we make the following assumptions on the minimum time it takes to give birth to a second child after the adoption of the two-child policy. First, based on the official procedure of granting birth permits, we assume that it takes about 1 month for prospective parents to obtain a birth permit from local authorities in order to be legally allowed to give birth to a second child. Second, we assume that it takes about 6 months on average to prepare for pregnancy and to get pregnant with a second child, after obtaining the birth permit. Third, we assume that the pregnancy lasts about 10 months on average. As a result, we assume that the time lag between the date when a province launched the two-child policy and the date when the two-child policy de facto affected the number of local births is 17 months. In other words, we assume that it takes at least 17 months on average for the two-child policy to affect local second-child births. For example, for a two-child policy that was launched in March 2014, we assume that it will begin to affect local second-child births, to a large extent, starting in August 2015. Given the fact that our assumption about the length of time taken to prepare for and get pregnant with a second child is to some extent arbitrary, we use an alternative assumption for a robustness check. **Figure A3** shows the de facto distribution of the adoption of the two-child policy based on the assumptions above.

4.3 Child-rearing costs

To capture changes in the average local costs of raising a child, we collect data on consumer price indices (CPIs), measuring the average change in prices over time that consumers pay for a basket of goods and services in each province from August 2014 to July 2015, as provided by China's National Bureau of Statistics.[21] In addition to the overall CPI, we also collect CPI by expenditure categories such as food, healthcare, education, transportation, entertainment, and housing. **Figure A4** shows the overall CPI on average over time. We find that the overall CPI decreased gradually beginning in December 2013 and reached its lowest point in January 2015.

4.4 Additional data sets

To further explore the mechanism behind the effects of relaxing mandatory family planning policies, we complement our main analysis with two additional data sets. The first is the per capita expenditure of households in each province over time. Since China's National Bureau of Statistics only provides per capita expenditure of households accumulated at the quarterly level in each province, we create the variable of per capita expenditure of households at the monthly level from October 2013 to September 2014 using the cubic spline interpolation method. We use this data set to further explore the extent to which the adoption of the two-child policy affects the average costs of living, complementing the analysis using CPIs. The second data set we use, for a further exploration, is the number of marriages in each province over time. Since China's Ministry of Civil Affairs only provides the number of marriages accumulated at the quarterly level in each province, we again resort to the cubic spline interpolation method to create the variable of the number of marriages at the monthly level from October 2013 to September 2014. We use this data set to explore the extent to which the adoption of the two-child policy affects the marriage market.

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[21] That said, CPIs may overstate or understate the true rise in local living costs. The reliability of our estimates is largely dependent on the association between CPIs and the average local cost of living.

Table A1 shows the summary statistics of the main variables. We find that there were a monthly average total of 29,861 births at the provincial level from November 2014 to October 2015, and a monthly average total of 17,062 second-child births in the same period. In other words, second-child births account for more than half of total births during this period. **Figures A5** and **A6** show the number of total births at the monthly level and the percentage of first-child births among total births at the monthly level, respectively. **Figure A7** uses a descriptive approach to compare the monthly second-child births of the treatment and control groups over time. We find that monthly second-child births rebounded significantly in the treatment group 17 months after the policy was announced in a given province. By contrast, we do not find similar patterns for the control group.

5 Model Specification and Identification Strategy

For each birth outcome, the following regression specification evaluates the impact of the adoption of the two-child policy as a quasi-experiment. The treatment is the degree of exposure to the two-child policy. We use a difference-in-differences (DID) setup:

$$Y_{s,t_{birth}} = \alpha' \mathrm{I}_s + \beta' \mathrm{I}_{t_{birth}} + \gamma \left(I_{s,t_{birth} \geq t_{policy+r}} \right) + \varepsilon_{s,t_{birth}} \quad (31)$$

where $Y_{st_{birth}}$ is the logarithm of monthly births 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, $I_{s,t_{birth} \geq t_{policy+r}}$ is a dummy for an observation r months after the adoption of the two-child policy in province s (specifically, an interaction of an indicator variable of being in a province s where the two-child policy is adopted and an indicator for being in the r^{th} month after the reform adoption), or an observation after the de facto policy adoption. r is the time interval between the date when a province launched the two-child policy and the date when the two-child policy de facto affected

the number of local births. In the benchmark model, we assume that $r = 17$, meaning that it takes about 17 months (including time for the birth application, conception, and pregnancy) on average for the two-child policy to affect local second-child births. The error term is $\varepsilon_{s,t_{birth}}$, α and β are vectors of coefficients to be estimated, and γ is the coefficient of interest.

The assumptions underlying the identification of parameter γ are similar to all DID analyses, in that the decision about whether and when to adopt the two-child policy must be uncorrelated with any prior trends in birth outcomes, and that the timing of when the two-child policy is adopted cannot coincide with any province-specific shocks or policies that might influence the birth outcomes.

Recent studies show that the two-way fixed effects regression above identifies the effect under the standard "common-trends" assumption when the treatment effect is constant across groups and over time. However, it is often implausible for the treatment effect to be constant. If the constant effect assumption is violated, the two-way fixed effects regression identifies weighted sums of the average treatment effects (ATE) in each group and period, with weights that may be negative (de Chaisemartin and D'Haultfoeuille 2020; Abraham and Sun 2018; Athey and Imbens 2018; Callaway and Sant'Anna 2019; Imai and Kim 2018; Goodman-Bacon 2018).

Specifically, according to de Chaisemartin and D'Haultfoeuille (2020), parameter γ in equation (1) is equal to a weighted sum of the treatment effect in each treated (s,t) cell: $\gamma = E\left(\sum_{(s,t):\, D_{s,t=1}} W_{s,t}\Delta_{s,t}\right)$. $\Delta_{s,t}$ is the ATE in group s and period t, and the weights $W_{s,t}s$ sum to one but may be negative. Negative weights arise because $\hat{\gamma}$ is a weighted sum of several DID estimates, which compare the evolution of the outcome between consecutive time periods across pairs of groups. However, the "control group" in some of those comparisons may be treated at both periods, which

may lead to the negative weights.[22] Due to the negative weights, the linear regression estimand may, for instance, be negative while all the average treatment effects are positive. To cope with the potential negative weights problem, we resort to an estimator proposed by de Chaisemartin and D'Haultfoeuille (2020). The estimator also allows us to check the common-trends assumption of the difference-in-differences setup when there are heterogeneous treatment effects (e.g., by groups or over time). If the common trends assumption holds, we should not find any significant treatment effects in the pretreatment periods.

The second potential identification problem of unobserved provincial specific shocks correlated with the adoption of the two-child policy is more difficult to investigate. However, there is much variation in the timing of the adoption of the two-child policy even within the same calendar year, as shown in **Figures A2** and **A3**, making it very unlikely that there are unobserved provincial specific shocks that are systematically correlated with the timing of the adoption of the policy and only influence the affected cohorts. We use permutation tests to formally support the contention that unobserved shocks correlated with the timing of the policy adoption are not biasing our estimates. We perform permutation tests for the outcomes of interest that randomly reassign the timing across provinces. We do this in two ways: first, we randomly assign the policy adoption timing among months in 2014 across provinces, and second, we randomly

[22] To illustrate this, we consider a simple example of a staggered adoption design with two groups and three periods: group 1 is untreated at periods 1 and 2 and treated at period 3, while group 2 is untreated at period 1 and treated both at periods 2 and 3. In this example, $\gamma = (DID_1 + DID_2)/2$, with $DID_1 = E(Y_{2,2}) - E(Y_{2,1}) - \left(E(Y_{1,2}) - E(Y_{1,1})\right)$, $DID_2 = E(Y_{1,3}) - E(Y_{1,2}) - \left(E(Y_{2,3}) - E(Y_{2,2})\right)$. The first DID compares the evolution of the mean outcome from period 1 to 2 in group 2 and in group 1. The second one compares the evolution of the mean outcome from period 2 to 3 in group 1 and in group 2. The control group in DID_2 (group 2) is treated both in pre and in the post period. Therefore, under the common trends assumption, it follows from de Chaisemartin and D'Haultfoeuille (2020) that DID_1 is equal to the ATE in group 2 in period 2 $\left(i.e., DID_1 = E[\Delta_{2,2}]\right)$, and DID_2 is equal to the ATE in group 1 in period 3, minus the change in the group2's ATE between periods 2 and 3 $\left(i.e., DID_2 = E[\Delta_{1,3}] - \left(E[\Delta_{2,3}] - E[\Delta_{2,2}]\right)\right)$. Intuitively, the mean outcome of groups 1 and 2 may follow different trends from period 2 to 3 either because group 1 becomes treated, or because group2's ATE changes. Negative weights arise because $\hat{\gamma}$ uses treated observations as controls.

assign the timing so that the same percentage of provinces adopted the two-child policy in each month as shown in **Figure A3**.

We also go further and use an alternative 2005 population census to construct monthly second-child births for all 31 provinces from November 2004 to October 2005. Given the fact that these second-child births occurred prior to the adoption of the two-child policy, we use them to conduct a placebo analysis. If our main estimates are not driven by other unobserved provincial specific shocks, then using these counterfactual monthly second-child births we should not find that the adoption of the two-child policy has similar and significant impacts.

Finally, we estimate the impact of the adoption of the two-child policy on birth outcomes for minority women, respectively. As we have shown before, the one-child policy mainly applies to women with Han ethnicity. If the two-child policy really works according to our hypothesis, we should therefore not find that adopting it has a similar and significant impact on monthly second-child births for women of minority ethnicities.

6 Empirical results

This section reports the effects of the adoption of the two-child policy. First, we show its impact on second-child births. Second, we show whether the impact is heterogeneous across groups. Third, we explore whether, to what extent, and how it affects first-child births. Finally, we conduct additional robustness checks.

6.1 Pretreatment trends

A crucial assumption that justifies the difference-in-differences approach is the presence of common parallel trends in the pretreatment periods. **Figure A7** shows that the pretreatment trends are similar. To test this assumption more formally, we use the estimator proposed by de Chaisemartin and D'Haultfoeuille (2020). If the common

parallel assumption is valid, we should find that the adoption of the two-child policy has not had any significant impacts in the pretreatment periods. **Figure 2** shows the treatment effect in the pretreatment periods for the outcome variable. We find consistent evidence that the treatment effect is close to zero and not statistically significant at the conventional level in the pre-treatment periods.

[Insert Figure 2 about here.]

6.2 Impact of the adoption of the two-child policy on second-child births

Using equation (31), we estimate the impact of the adoption of the two-child policy on monthly second-child births. Column 1 of **Table 1** reports the main results. We find that the coefficient of interest is 0.264, suggesting that the adoption of the two-child policy leads to an increase of about 30.2% ($[e^{0.264} - 1] \cdot 100\%$) in monthly second-child births on average.

[Insert Table 1 about here.]

In addition to presenting the pretreatment trends, **Figure 2** shows dynamic impacts on childbearing, as well. We find a significant increase in second-child births in the month of the de facto adoption of the two-child policy, or 17 months after the official launch of the two-child policy. We also find heterogeneous impacts over time. On average, we find that the adoption of the two-child policy increases monthly second-child births by 25.4% during the month of the de facto policy adoption, and that its impact further increases to 70.2% three months after that. A simple average of the instantaneous and dynamic effects shows that the adoption of the two-child policy leads to an increase in second-child births of about 32.1% on average, which is very close to the benchmark results (30.2%). We also check the negative weight problem, using the test proposed by de Chaisemartin and D'Haultfoeuille (2020), but find that it is not obvious in the data.[23]

[23] Specifically, under the common-trends assumption, gamma estimates a weighted sum of 54 ATTs. 54 ATTs receive a positive weight, and 0 receive a negative weight.

[Insert Figure 2 about here.]

As discussed above, of primary concern in our identification strategy is the existence of unobserved province-specific shocks that differ systematically with treatment exposure. To test whether the timing pattern of the adoption of the two-child policy is driving our results, we perform permutation tests for all of our outcomes that randomly reassign the timing of the adoption of the two-child policy across provinces (Lovenheim and Willén 2019). We do this in two ways: first, we randomly assign the two-child policy dates between January 2014 and December 2014 across provinces, and second, we randomly assign the two-child policy dates to match the timing distribution shown in **Figure A3**. We perform these permutations 300 times for each outcome and calculate the percentage of times the simulated estimate is larger than the actual estimate. These results therefore represent p-values of the null hypothesis that any combination of the two-child policy dates across provinces would generate the same pattern of treatment effects. We reject the null at the 1-percent level for every outcome in both panels. These results suggest that our baseline estimates are not driven by the timing pattern of the adoption of the two-child policy.

In addition, we use the counterfactual monthly second-child births from the 2005 population census to conduct a placebo analysis. If the main estimates are not driven by other unobservable province-specific shocks, we should not find any significant impact of the adoption of the two-child policy on these counterfactual monthly second-child births. The results are reported in column 2 of **Table 1**. We find that the coefficient of interest is 0.07. Again using the estimator proposed by de Chaisemartin and D'Haultfoeuille (2020), we find that the adoption of the two-child policy increases the counterfactual monthly second-child births by 4% upon de facto policy adoption, and that the impact further increases to 6% three months after that. More details are reported in **Figure A8**. All of these estimates are much smaller than our benchmark results and are not statistically significant at the conventional level. This placebo analysis therefore further alleviates the concern that our estimation results are

driven by unobserved province-specific shocks.

Finally, to check whether the increase in monthly second-child births is driven by the adoption of the two-child policy rather than other unobservable provincial specific shocks, we estimate its impact for minority women. Given the fact that minority women were not restricted by the one-child policy, we should not find any significant impact of the two-child policy on monthly second-child births for women of minority ethnicities if the two-child policy works. The results are shown in Column 4 of **Table 1**. We find that the adoption of the two-child policy increases monthly second-child births by 11.1% on average for minority women, which is not statistically significant at the conventional level. By contrast, we find that the adoption of the two-child policy significantly increases monthly second-child births by 32.4% on average for the Han. Overall, we find consistent evidence that the adoption of the two-child policy substantially increases monthly second-child births, which is consistent with our theoretical model.

6.3 Which groups are more responsive to the two-child policy?

We now further explore which groups are more responsive to the adoption of the two-child policy. Our main results are reported in **Table 1**. First, we examine its impact in relation to land-contracting rights.[24] On the one hand, the one-child policy was implemented less strictly for couples who were granted land-contracting rights (those who were from rural areas) compared with their counterparts who were not granted land-contracting rights (those who were from urban areas). For example, prior to the adoption of the two-child policy, couples from rural areas were permitted to have a second child if their first child was female. On the other hand, couples from rural areas are traditionally more willing to give birth to more children, or exposed to higher fertility preferences, than their counterparts from urban areas. As a result, it is

[24] The right to contract land is closely related to one's place of birth (rural or urban areas).

empirically uncertain which of the two groups are more responsive to the adoption of the two-child policy. **Table 1** reports the main results. We find that the adoption of the two-child policy increases monthly second-child births by 36.4% on average among couples from rural areas. By contrast, the adoption of the two-child policy increases monthly second-child births by only 16.4% on average for couples from urban areas, which is not statistically significant at the conventional level.

Second, we examine the impact of the adoption of the two-child policy in relation to childbearing age. We find that the adoption of the two-child policy leads to an increase of 30.7% on average in monthly second-child births for couples of relatively late childbearing age (above the median value of the childbearing age distribution), whereas the same policy leads to an increase of only 12.7% on average in monthly second-child births for couples of relatively early childbearing age (below the median value of the childbearing age distribution), the latter of which is not statistically significant at the conventional level.

Third, we examine the impact of the adoption of the two-child policy in relation to women's education. We find that the adoption of the two-child policy leads to a 35.1% increase on average in monthly second-child births for women with a college degree or above, whereas the same policy leads to a 20.4% increase on average in monthly second-child births for women without a college degree.

Finally, we examine the impact of the adoption of the two-child policy in relation to pension accessibility. On the one hand, couples with access to pensions are expected to have a relatively higher likelihood of being able to bear the costs of child-rearing than couples without access to pensions. On the other hand, having access to pensions may crowd out childbearing if government-provided social security is substitutable for old age support from children. We find that the adoption of the two-child policy leads to an increase of 26.4% on average in monthly second-child births for couples

covered by pensions, whereas the same policy leads to an increase of only 9.7% on average in monthly second-child births for couples without access to pensions. The latter increase is not statistically significant at the conventional level. When we further examine the impact of the adoption of the two-child policy in relation to medical insurance accessibility, we likewise find that the impact is more pronounced for women who are covered by medical insurance.

In summary, we find evidence that the impact of the adoption of the two-child policy varies significantly across groups. The impact is more pronounced for women who are relatively older, more educated, covered by social security, and from rural areas. All of these findings suggest that fertility preferences and sensitivity to child-rearing costs are likely important factors determining second-child births, and that couples who have relatively higher fertility preferences and who are less sensitive to child-rearing costs are more likely to give birth to a second child after the adoption of the two-child policy. These findings are again consistent with our theoretical model.

6.4 Does the two-child policy affect first-child births?

Our previous findings are in line with the theoretical implication that relaxing birth quotas induces couples with relatively higher fertility preferences, particularly those less sensitive to child-rearing costs, to give birth to more children. In this sub-section, we further investigate the impact of the adoption of the two-child policy on first-child births. Although the two-child policy itself does not affect first-child births *directly*, it may affect first-child births *indirectly*, through two opposing channels. On the one hand, the higher price of final goods, caused by the adoption of the two-child policy, increases the average local child-rearing cost, which in turn affects first-child births (this is the child-rearing cost effect or price effect). On the other hand, the higher interest rate, caused by the adoption of the two-child policy, encourages households to give birth to more children through a higher return on investing in children (the

income effect). As a result, it is theoretically ambiguous regarding the impact of the adoption of the two-child policy on first-child births.

To empirically investigate the impact of the adoption of the two-child policy on monthly first-child births, we make an alternative assumption that it takes about 3 months to get pregnant. The rationale for the relatively shorter time needed to prepare for pregnancy and to get pregnant for the first birth is that women who are giving birth to a first child are relatively younger and in turn have higher chance of getting pregnant on average than their counterparts who are giving birth to a second child. Also, unlike second-child births, prospective parents do not need to apply for a birth permit to give birth to a first child. As a result, we assume that it takes only about 13 months in total for the adoption of the two-child policy to affect local first-child births ($r = 13$).

Using the same empirical approach, we find that the coefficient of interest is –0.03, which is not statistically significant at the conventional level. We also find that the negative weight problem does exist in the data.[25] To reduce the concern that the two-way fixed effects model outlined above may not capture the average treatment effects when there are heterogeneous treatment effects, we once again use the estimator proposed by de Chaisemartin and D'Haultfoeuille (2020) to estimate the impact of the adoption of the two-child policy on monthly first-child births. **Figure 3** shows the main results. We find that the adoption of the two-child policy reduces monthly first-child births by 31.6% on the month of de facto policy adoption, an impact that further increases to 80.6% in the seventh month of de facto policy adoption. A simple average of the instantaneous and dynamic effects shows that the adoption of the two-child policy leads to a decrease of about 40.7% on average in monthly first-child births. The impact is also heterogeneous over time. As a result,

[25] Specifically, under the common-trends assumption, gamma estimates a weighted sum of 169 ATTs. 137 ATTs receive a positive weight, and 32 receive a negative weight.

these findings suggest that the adoption of the two-child policy substantially decreases the number of first-child births. In other words, the child-rearing cost effect (or price effect) dominates the impact of the adoption of the two-child policy on first-child births.

[Insert Figure 3 about here.]

6.5 Does the two-child policy affect child-rearing costs?

We now take a step further and examine the changes in child-rearing costs after the adoption of the two-child policy. It is challenging to try to measure average local child-rearing cost directly. As we explain in section 4, we use consumer price indices (CPIs) to capture average changes in the local costs of living. We assume the costs of raising a child are closely associated with average local living costs. If the child-rearing cost effect (or price effect) does exist, we should find evidence that the adoption of the two-child policy leads to significantly higher local CPIs.

To capture the impact of the adoption of the two-child policy on local CPIs, we collect monthly data on local CPIs from August 2014 to July 2015, and assume that it takes about 10 months on average for the adoption of the two-child policy to affect local CPIs ($r = 10$). The model specification is as follows:

$$CPI_{st_{birth}} = \alpha' \mathrm{I}_s + \beta' \mathrm{I}_{t_{birth}} + \gamma_2 \left(I_{s,t_{birth} \geq t_{policy+r}} \right) + \varepsilon_{st_{birth}} \quad (32)$$

where $CPI_{st_{birth}}$ 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, $I_{s,t_{birth} \geq t_{policy+r}}$ is an indicator for an observation r months after the adoption of the two-child policy, or an observation after the de facto policy adoption. In this model, we assume that $r = 10$, meaning that it takes about 10 months on average for the two-child policy to affect

local CPIs.[26] The error term is $\varepsilon_{s,t_{birth}}$, α and β are vectors of coefficients to be estimated, and γ_2 is the coefficient of interest.

To reduce the concern that the two-way fixed effects model described above may not capture the average treatment effects when there are heterogeneous treatment effects, we make further use of the estimator proposed by de Chaisemartin and D'Haultfoeuille (2020) to estimate the impact of the adoption of the two-child policy on local CPIs.[27] Our main results are shown in **Figure 4**. A simple average of the instantaneous and dynamic effects shows that the adoption of the two-child policy leads to an increase of about 1.1 unit in local CPIs. The impact is also heterogeneous over time. In particular, the impact increases slightly in the beginning and then speeds up over time, suggesting that it takes a certain number of months to prepare for pregnancy and to get pregnant after the adoption of the two-child policy. Specifically, we find that local CPIs increase by 0.5 units 13 months after the adoption of the two-child policy, and that that impact increases to 1.3 units 17 months after the policy is adopted.

[Insert Figure 4 about here.]

When further analyzing the impact of the adoption of the two-child policy on CPIs by industrial category, we find that the impact is more pronounced for industries such as food, healthcare, education, and housing. For example, local CPIs for food increase by 4 units 17 months after the adoption of the policy. We also find evidence that the CPIs for commodities such as tobacco, alcohol, and clothes witnessed a decline after the adoption of the two-child policy. All of the evidence suggests that the adoption of

[26] As we have discussed before, it takes a certain number of months to prepare for pregnancy and to get pregnant after the adoption of the two-child policy. If this is true, it is likely that the impact of the adoption of the two-child policy on local CPIs will be relatively small in the short term and then gradually increase over time.

[27] Using a two-way fixed effects model, we find that the coefficient of interest (γ_2) is 0.16. We also find that the negative weight problem is obvious in the data. Specifically, under the common-trends assumption, gamma estimates a weighted sum of 170 ATTs. 138 ATTs receive a positive weight, and 32 receive a negative weight.

the two-child policy significantly affects the average price of final goods and services, particularly those relevant to childbearing and child-rearing, and in turn contributes to the increase in child-rearing costs after the adoption of the two-child policy. More details can be found in **Figures A9** to **A14**.

6.6 Does the two-child policy affect the number of marriages?

We have shown that the child-rearing cost effect (or price effect) is a likely underlying mechanism through which the adoption of the two-child policy adversely affects first-child births. That said, it is possible that the two-child policy may crowd out first-child births through other alternative channels. An important one is the marriage market. Since children can be sources of joy or future supporters, the adoption of the two-child policy might increase expected marriage gains, and therefore marriage outcomes, which are closely related to first-child births, should be altered accordingly. In this sub-section, we further explore the causal impact of the adoption of the two-child policy on the number of marriages. We construct monthly data on the number of marriages from October 2013 to September 2014 using data provided by China's Ministry of Civil Affairs.

To capture the impact of the adoption of the two-child policy on the number of marriages, we assume that the adoption of the two-child policy affects local marriage markets immediately ($r = 0$). The model specification is as follows:

$$Marriage_{st_{birth}} = \alpha' \mathrm{I}_s + \beta' \mathrm{I}_{t_{birth}} + \gamma_3 \left(I_{s,t_{birth} \geq t_{policy+r}} \right) + \varepsilon_{st_{birth}} \quad (33)$$

where $Marriage_{st_{birth}}$ is 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, $I_{s,t_{birth} \geq t_{policy+r}}$ is an indicator for an observation r months after the adoption of the two-child policy, or an observation after the de facto policy adoption. In this model, we assume that $r = 0$, meaning that the adoption of the two-child policy affects local

marriage markets immediately.[28] The error term is $\varepsilon_{s,t_{birth}}$, α and β are vectors of coefficients to be estimated, and γ_3 is the coefficient of interest.

To reduce the concern that the two-way fixed effects model described above may not capture the average treatment effects when there are heterogeneous treatment effects, we make use of the estimator proposed by de Chaisemartin and D'Haultfoeuille (2020) to estimate the impact of the adoption of the two-child policy on local marriages.[29] Our main results are reported in **Figure 5**. We find that the adoption of the two-child policy significantly increases the number of marriages, echoing the previous study by Huang and Zhou (2015), who find that the one-child policy induced a significantly lower marriage rate. Specifically, we find that the number of marriages increases by 7.4% in the fourth month of the adoption of the two-child policy, and that that impact increases to 35.1% in the ninth month of the policy adoption. A simple average of the instantaneous and dynamic effects shows that the adoption of the two-child policy leads to an increase of about 13.4% in the number of local marriages. As a result, the positive impact of the adoption of the two-child policy on the marriage market does not weaken our previous argument that the child-rearing cost effect (or price effect) is responsible for the decrease in first-child births after the adoption of the two-child policy.

[Insert Figure 5 about here.]

6.7 Robustness checks

In this section, we conduct several robustness checks. First, we relax the assumption of the time taken to prepare for pregnancy and to get pregnant with a second child. We

[28] Unlike our previous exercises that set r at 10 to 17 months, we argue that it takes shorter time for the adoption of the two-child policy to affect local marriage market. This is intuitive, to some extent, because, conceptually, marrying someone is much less time-consuming than giving birth to a child.

[29] Using a two-way fixed effects model, we find that the coefficient of interest (γ_3) is -0.08. We also find that the negative weight problem is obvious in the data. Specifically, under the common-trends assumption, gamma estimates a weighted sum of 229 ATTs. 169 ATTs receive a positive weight, and 60 receive a negative weight.

assume that it takes about 1 month to do this for a second child after the adoption of the two-child policy. Based on this alternative assumption ($r = 12$), we estimate the impact of the adoption of the two-child policy on second-child births. To reduce the concern that the two-way fixed effects model may not capture the average treatment effects when there are heterogeneous treatment effects, we make further use of the estimator proposed by de Chaisemartin and D'Haultfoeuille (2020) to estimate the impact of the adoption of the policy.[30] Our main results are reported in **Figure 6**. A simple average of the instantaneous and dynamic effects shows that the adoption of the two-child policy leads to an increase of about 37.9% in second-child births. We also find that the impact is heterogeneous over time. In particular, the impact is relatively small and is not statistically significant at the conventional level until 17 months after the adoption of the two-child policy, which further supports our assumption in the benchmark model that it takes about 6 months to prepare for pregnancy and to get pregnant with a second child.

[Insert Figure 6 about here.]

Second, we consider whether our main results are contaminated by migrants from other provinces. We re-estimate our main results for local natives and inter-provincial migrants, respectively. Given the fact that the two-child policy mainly applies to the local registered population, we should not find a significant impact of the adoption of the two-child policy on monthly births of inter-provincial migrants if the two-child policy works. The results are shown in **Table 1**. We find that the adoption of the two-child policy reduces monthly second-child births by 8.5% for inter-provincial migrants, which is not statistically significant at the conventional level. By contrast, the adoption of the two-child policy significantly increases monthly second-child births by 31.5% for local natives on average. This evidence thus confirms that our main results are not driven by migration.

[30] Using a two-way fixed effects model, we find that the coefficient of interest is -0.007, which is not statistically significant at the conventional level. We also find that the negative weight problem is obvious in the data. Specifically, under the common-trends assumption, gamma estimates a weighted sum of 199 ATTs. 159 ATTs receive a positive weight, and 40 receive a negative weight.

Third, we use the number of births in levels, rather than the logarithm of births, as outcome variables and check whether our main results are affected by different outcome variables. Our main results are reported in **Table 2**. Our findings show that most of our main results are still robust when we use alternative outcome variables.[31]

Finally, we use per capita expenditure of households as an alternative proxy for child-rearing costs, which further considers the potential changes in the composition of household expenditures, and check whether per capita expenditure of households responds to the adoption of the two-child policy. We make a final use of the estimator proposed by de Chaisemartin and D'Haultfoeuille (2020) to estimate the impact of the adoption of the policy. Our main results are reported in **Figure 7**. Our findings show that households' per capita expenditure increases by 20% in the ninth month of the adoption of the two-child policy, further strengthening our argument that the adoption of the two-child policy leads to higher child-rearing costs by raising the average costs of living.

[Insert Tables 1 and 2 about here]

7 Conclusion

To the best of our knowledge, this is the first study to investigate the childbearing consequences of relaxing birth quotas, theoretically and empirically. We exploit a plausibly exogenous two-child policy adopted by local Chinese governments on different dates and estimate the effects of the adoption of the two-child policy in a different-in-differences setup. We find that the adoption of the two-child policy significantly increases second-child births, particularly for households with higher fertility preferences and those that are less sensitive to child-rearing costs. However, the same policy substantially reduces first-child births. We provide additional

[31] An exception is that we find a larger impact for women without a college degree or above using the outcome variable in levels.

evidence suggesting that child-rearing costs are a likely underlying mechanism for the decrease. All of these empirical findings are supported by an extended Barro-Becker model.

This study still has some limitations. First, our findings only reveal short-term impacts of the two-child policy on births. The policy's long-term impacts remain empirically unclear. Our results should therefore be interpreted cautiously. Second, the main assumption in our article, that it takes about 6 months to prepare for pregnancy and to get pregnant with a second child, is to some extent arbitrary. That said, using an alternative assumption that it takes about one month to prepare for pregnancy and to get pregnant with a second child in the robust analysis, we do find some evidence suggesting that it takes some time for the adoption of the two-child policy to significantly affect local second-child births. To some extent, our article provides a lower bound of the impact of the adoption of the two-child policy. Third, due to data limitations, we do not consider other alternative channels (e.g., labor market outcomes) through which relaxing birth quotas may crowd out other couples' decisions about whether to give birth to a first child. That said, we do provide evidence that the adoption of the two-child policy increases the number of marriages, which is likely to strengthen rather than weakening the child-rearing cost channel.

Overall, this article shows that the adoption of the two-child policy does facilitate second-child births, especially for those couples who have higher fertility preferences and who are less sensitive to child-rearing costs, but also crowds out first-child births. These findings suggest that policy makers should give priority to reducing the child-rearing costs borne by prospective parents rather than simply relaxing or even abolishing birth quotas universally when adopting policies to boost fertility rates.

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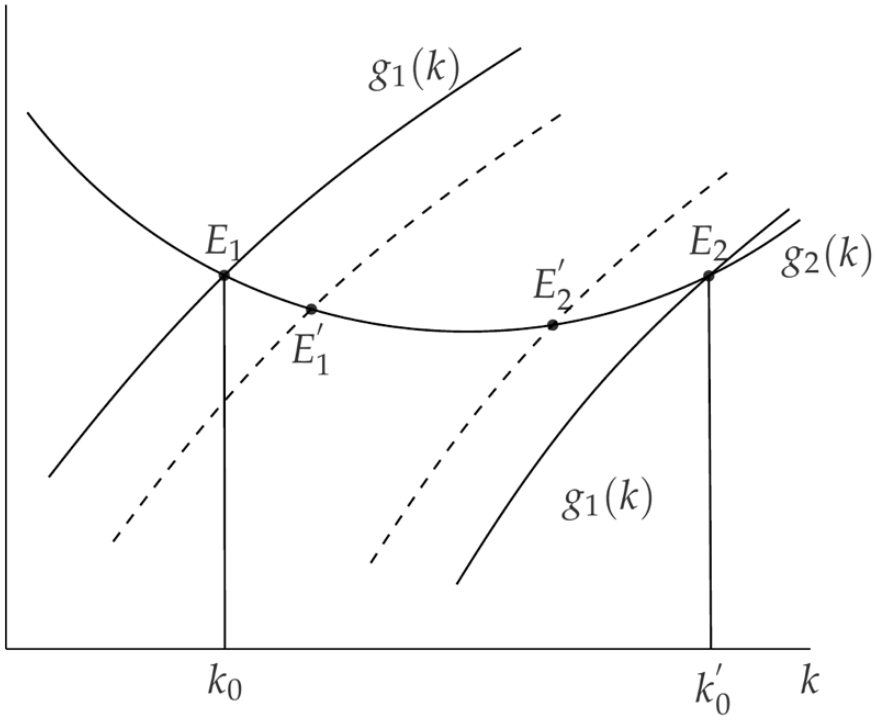

Figure 1 Impact of relaxing birth quotas on fertility rate

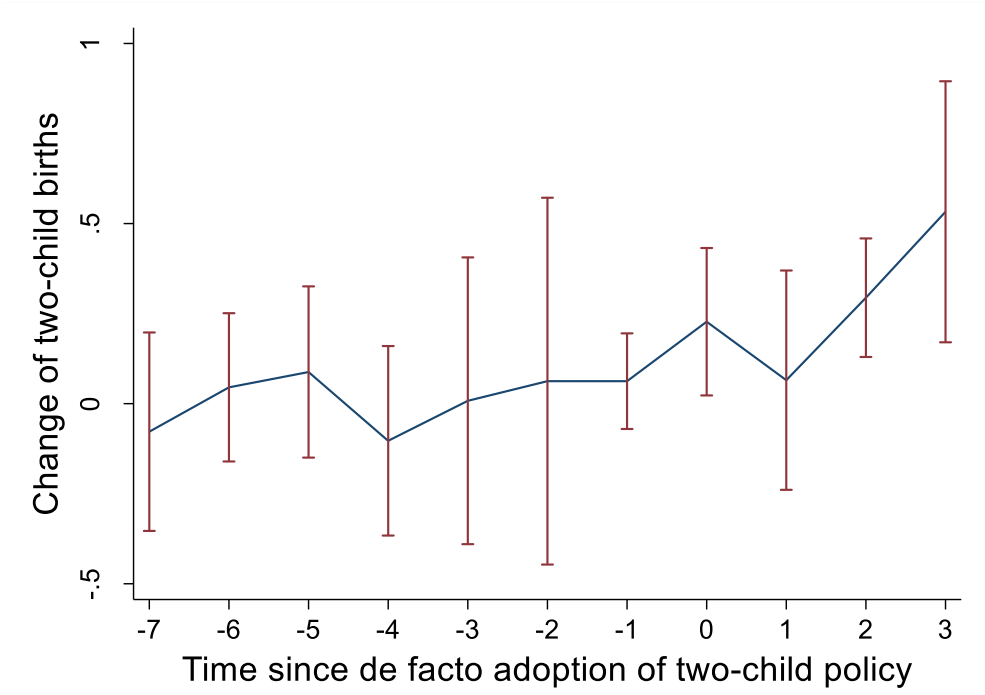

Figure 2 Impact of the adoption of the two-child policy on second-child births

Notes: We follow de Chaisemartin and D'Haultfoeuille (2020) to estimate the treatment effect using DID designs with multiple groups and periods. We use Stata code *did_multiplegt* to draw the figure.

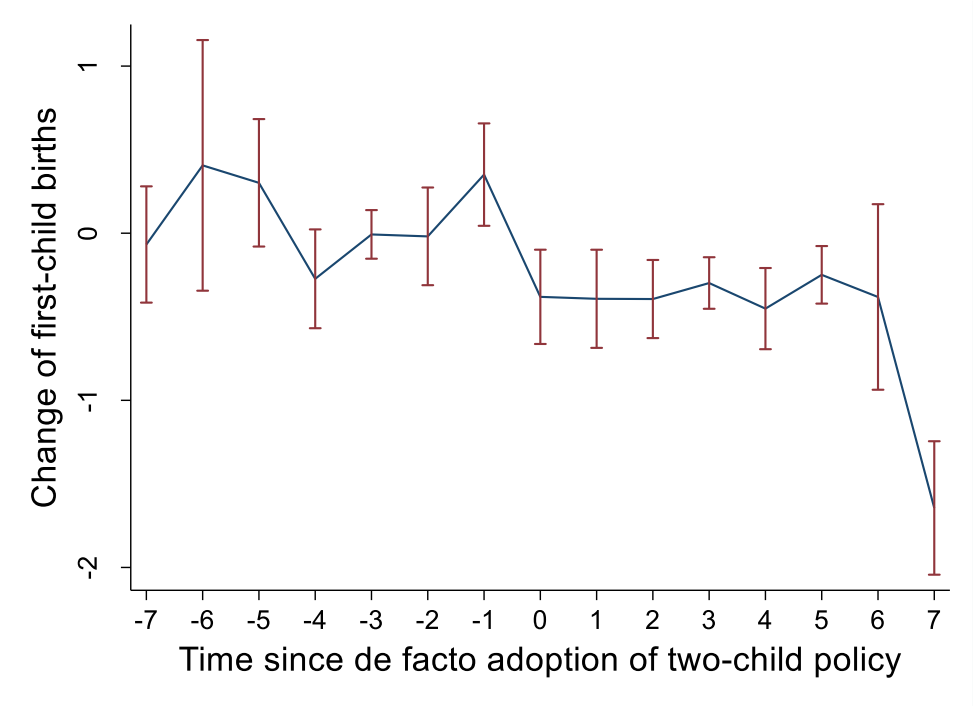

Figure 3 Impact of the adoption of the two-child policy on first-child births

Notes: We follow de Chaisemartin and D'Haultfoeuille (2020) to estimate the treatment effect using DID designs with multiple groups and periods. We use Stata code *did_multiplegt* to draw the figure.

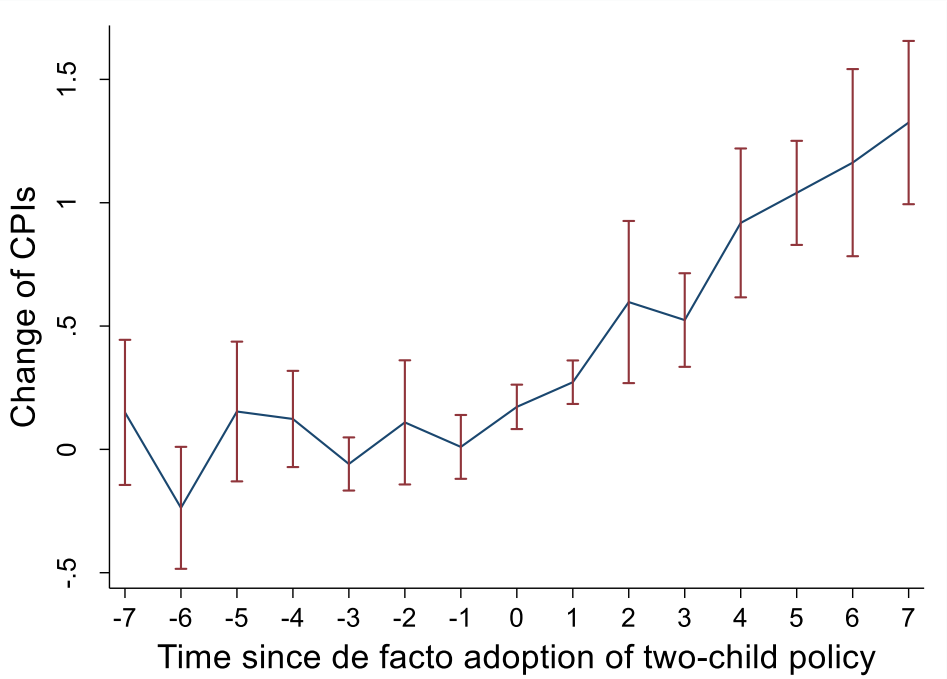

Figure 4 Impact of the adoption of the two-child policy on local CPIs

Notes: We follow de Chaisemartin and D'Haultfoeuille (2020) to estimate the treatment effect using DID designs with multiple groups and periods. We use Stata code *did_multiplegt* to draw the figure.

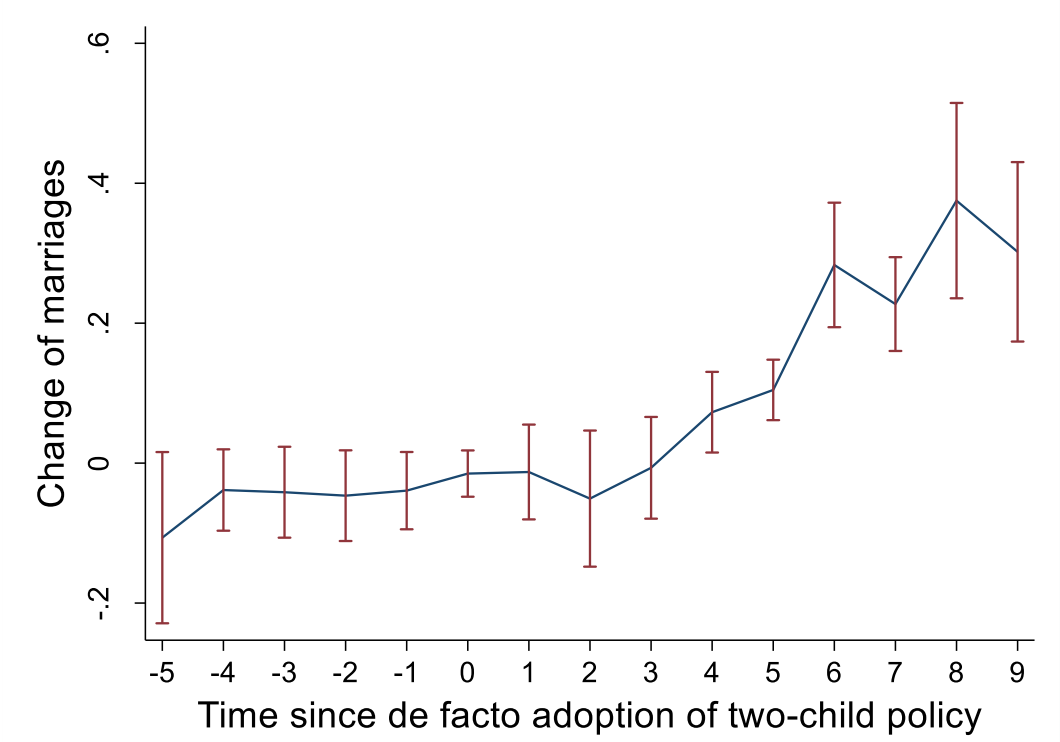

Figure 5 Impact of the adoption of the two-child policy on local marriages

Notes: We follow de Chaisemartin and D'Haultfoeuille (2020) to estimate the treatment effect using DID designs with multiple groups and periods. We use Stata code *did_multiplegt* to draw the figure.

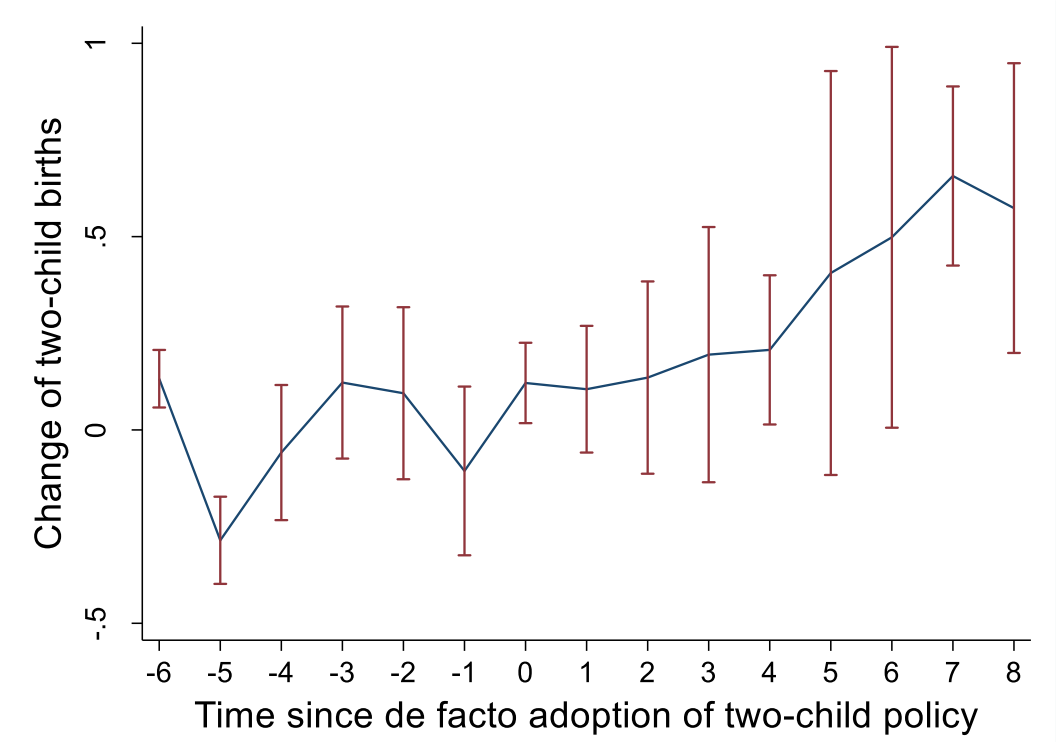

Figure 6 Impact of the adoption of the two-child policy on second-child births using alternative time (i.e. one month) it takes to prepare for pregnancy and to get pregnant with a second child

Notes: We follow de Chaisemartin and D'Haultfoeuille (2020) to estimate the treatment effect using DID designs with multiple groups and periods. We use Stata code *did_multiplegt* to draw the figure.

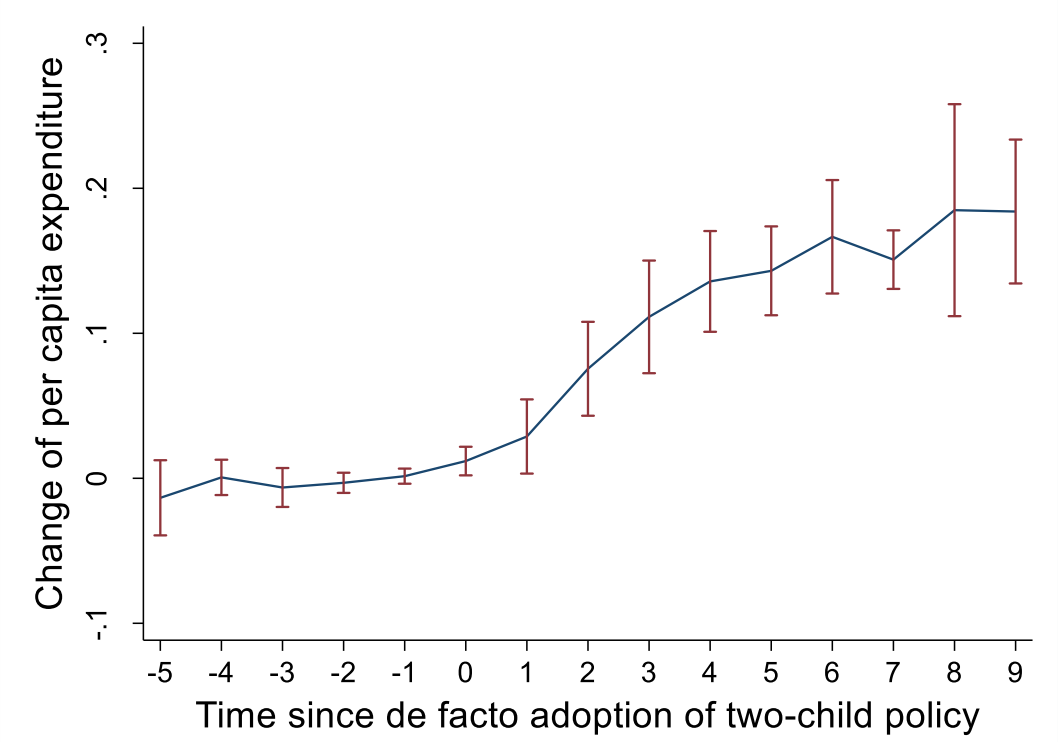

Figure 7 Impact of the adoption of the two-child policy on local per capita expenditure of households

Notes: We follow de Chaisemartin and D'Haultfoeuille (2020) to estimate the treatment effect using DID designs with multiple groups and periods. We use Stata code *did_multiplegt* to draw the figure.

Table 1 Impact of relaxing birth quotas on the number of monthly second-child births (in logarithm)

	(1) Second birth b/se	(2) Census 2005 b/se	(3) Han b/se	(4) Minority b/se	(5) Rural b/se	(6) Urban b/se	(7) Old b/se	(8) Young b/se	(9) College b/se
Treatment effect	0.264***	0.070	0.281***	0.106	0.311***	0.152	0.268**	0.120	0.301**
	(0.073)	(0.072)	(0.081)	(0.215)	(0.102)	(0.123)	(0.109)	(0.107)	(0.143)
Observations	371	371	360	199	361	368	370	353	291
Adjusted R^2	0.864	0.924	0.866	0.700	0.835	0.691	0.807	0.734	0.462
N_clust	31	31	31	26	31	31	31	31	31
	(10) No college b/se	(11) Pension b/se	(12) No pension b/se	(13) Medicare b/se	(14) No medicare b/se	(15) Native b/se	(16) Migrant b/se		
Treatment effect	0.186**	0.235**	0.093	0.285***	-0.205	0.274***	-0.089		
	(0.077)	(0.104)	(0.149)	(0.084)	(0.229)	(0.079)	(0.221)		
Observations	362	365	340	370	163	371	174		
Adjusted R^2	0.877	0.851	0.632	0.865	0.329	0.883	0.650		
N_clust	31	31	31	31	31	31	31		

Note: this table reports the impact of relaxing birth quotas using outcomes in logarithm. Column 1 shows the impact of relaxing birth quotas on monthly second-child births. Column 2 shows a placebo analysis using population census in 2005. Columns 3 and 4 show the impact of relaxing birth quotas by ethnicity. Columns 5 and 6 show the impact of relaxing birth quotas by land contracting right (rural/urban). Columns 7 and 8 show the impact of relaxing birth quotas by childbearing age. Columns 9 and 10 show the impact of relaxing birth quotas by education degree. Columns 11 and 12 show the impact of relaxing birth quotas by access to pension. Columns 13 and 14 show the impact of relaxing birth quotas by access to medical insurance. Columns 15 and 16 show the impact of relaxing birth quotas by migration status.

Table 2 Impact of relaxing birth quotas on the number of monthly second-child births (in levels)

	(1) Second birth	(2) Han	(3) Minority	(4) Rural	(5) Urban	(6) Old	(7) Young	(8) College	(9) No college
	b/se	b/se	b/se	b/se	b/se	b/se	b/se	b/se	b/se
Treatment effect	4494.732***	3988.847**	505.886	4135.303***	359.429	3448.008**	1046.725	613.883*	3880.850**
	(1616.922)	(1776.657)	(332.506)	(1457.403)	(599.927)	(1378.437)	(619.302)	(361.058)	(1620.713)
Observations	372	372	372	372	372	372	372	372	372
Adjusted R^2	0.865	0.872	0.780	0.833	0.704	0.834	0.697	0.437	0.858
N_clust	31	31	31	31	31	31	31	31	31
	(10) Pension	(11) No pension	(12) Medicare	(13) No medicare	(14) Native	(15) Migrant			
	b/se	b/se	b/se	b/se	b/se	b/se			
Treatment effect	4599.571***	-104.838	4566.676***	-71.944	4399.009***	95.723			
	(1555.536)	(467.952)	(1627.571)	(174.455)	(1576.559)	(224.827)			
Observations	372	372	372	372	372	372			
Adjusted R^2	0.856	0.684	0.862	0.259	0.864	0.646			
N_clust	31	31	31	31	31	31			

Note: this table reports the impact of relaxing birth quotas using outcomes in levels. Column 1 shows the impact of relaxing birth quotas on monthly second-child births. Columns 2 and 3 show the impact of relaxing birth quotas by ethnicity. Columns 4 and 5 show the impact of relaxing birth quotas by land contracting right (rural/urban). Columns 6 and 7 show the impact of relaxing birth quotas by childbearing age. Columns 8 and 9 show the impact of relaxing birth quotas by education degree. Columns 10 and 11 show the impact of relaxing birth quotas by access to pension. Columns 12 and 13 show the impact of relaxing birth quotas by access to medical insurance. Columns 14 and 15 show the impact of relaxing birth quotas by migration status.

Appendix

Table A1 Summary statistics

Variables	N	Mean	Median	Std. Dev.	min	max
Monthly second-child births	372.00	17062.00	13656.85	13731.51	0.00	75567.24
Monthly second-child births (for people from rural)	372.00	10317.26	7433.91	9905.70	0.00	63096.05
Monthly second-child births (for people from urban)	372.00	6744.74	5242.15	5572.66	0.00	32850.81
Monthly second-child births (for college)	372.00	2424.38	1687.54	2587.30	0.00	14227.99
Monthly second-child births (for no college)	372.00	14637.62	11503.95	12645.93	0.00	72435.29
Monthly second-child births (for old women)	372.00	10929.86	8135.93	9523.78	0.00	59647.27
Monthly second-child births (for young women)	372.00	6132.14	4490.26	5325.93	0.00	26971.66
Monthly second-child births (for women with pension)	372.00	12368.95	9794.39	11064.35	0.00	65058.43
Monthly second-child births (for women without pension)	372.00	4693.05	3287.09	4597.44	0.00	25136.28
Monthly second-child births (for han)	372.00	15021.90	9920.99	13639.28	0.00	75567.24
Monthly second-child births (for minorities)	372.00	2040.10	501.49	3748.31	0.00	23855.63
Monthly second-child births (for native)	372.00	15962.11	12735.53	13259.95	0.00	74028.93
Monthly second-child births (for migrant)	372.00	1099.89	0.00	1924.65	0.00	11396.63
Monthly first-child births	372.00	12799.36	10386.18	10001.90	0.00	48436.91
Overall CONSUMER PRICE INDEX	372.00	101.52	101.50	0.64	99.90	103.60
CONSUMER PRICE INDEX (food)	372.00	102.22	102.30	1.33	97.60	105.60
CONSUMER PRICE INDEX (cigarette)	372.00	100.38	99.90	1.98	95.70	106.00
CONSUMER PRICE INDEX (clothes)	372.00	102.93	102.65	2.07	97.20	115.20
CONSUMER PRICE INDEX (equipment)	372.00	101.07	101.00	0.86	99.20	104.00
CONSUMER PRICE INDEX (healthcare)	372.00	101.59	101.50	0.94	98.30	105.00
CONSUMER PRICE INDEX (transportation)	372.00	98.94	99.00	1.27	94.90	103.20
CONSUMER PRICE INDEX (entertainment)	372.00	101.68	101.50	1.60	96.60	108.90
CONSUMER PRICE INDEX (house)	372.00	101.12	101.10	1.24	96.90	105.10

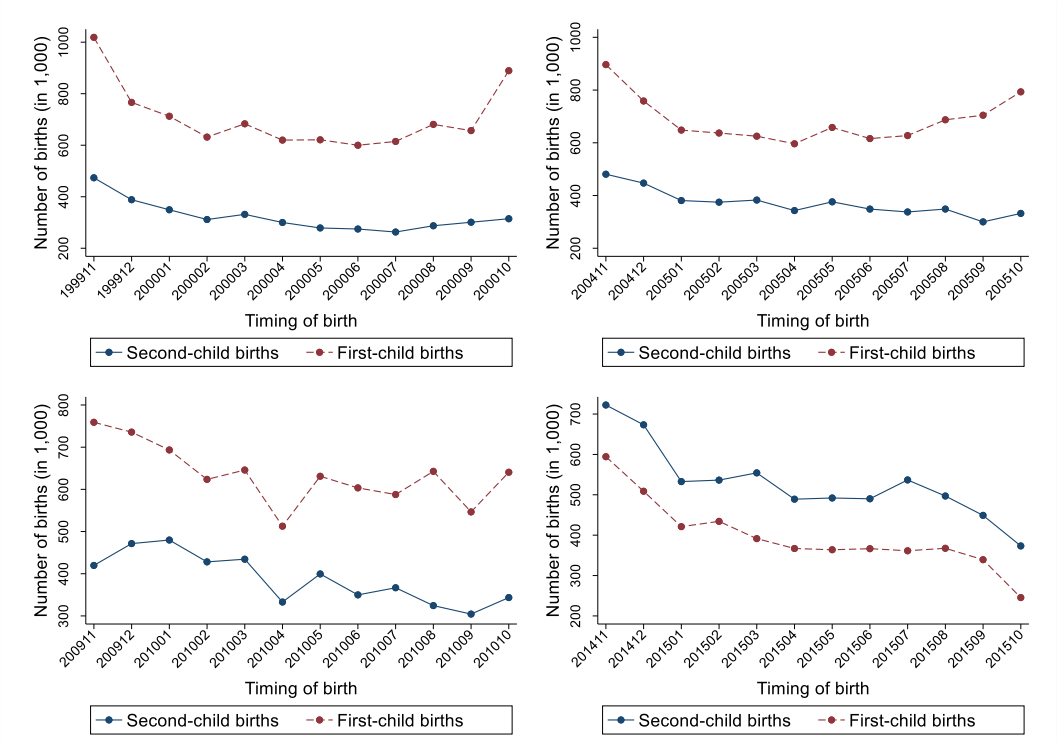

Figure A1 Relative number of first-child births and second-child births over time (2000-2015)

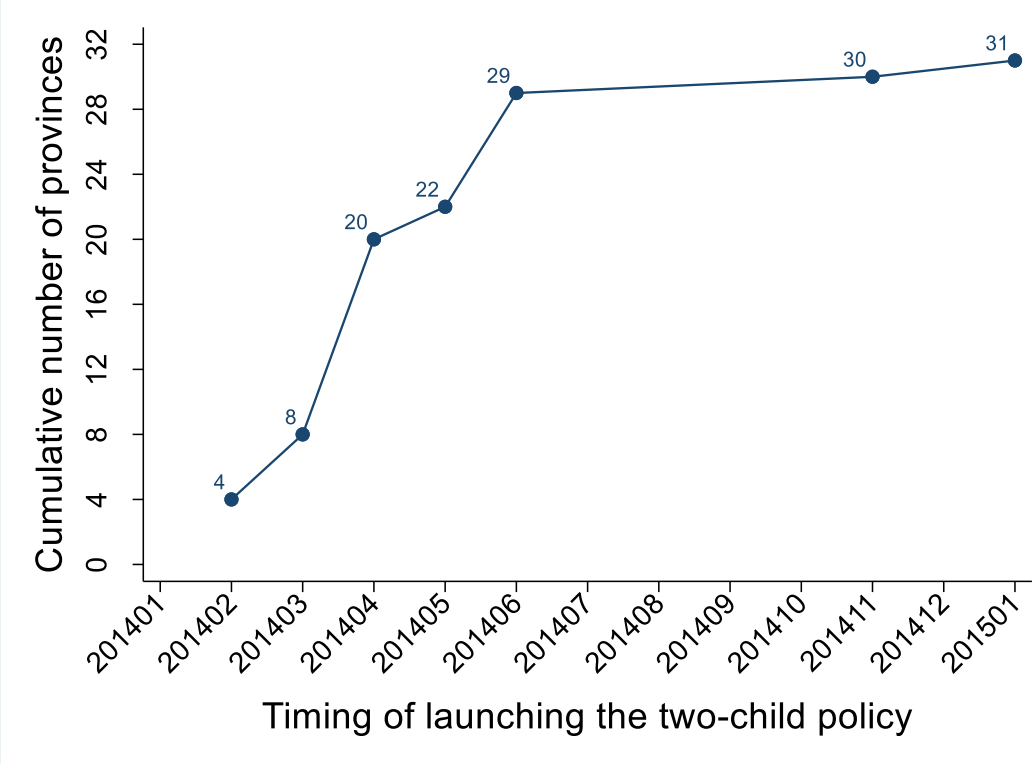

Figure A2 Timing of launching the two-child policy across provinces

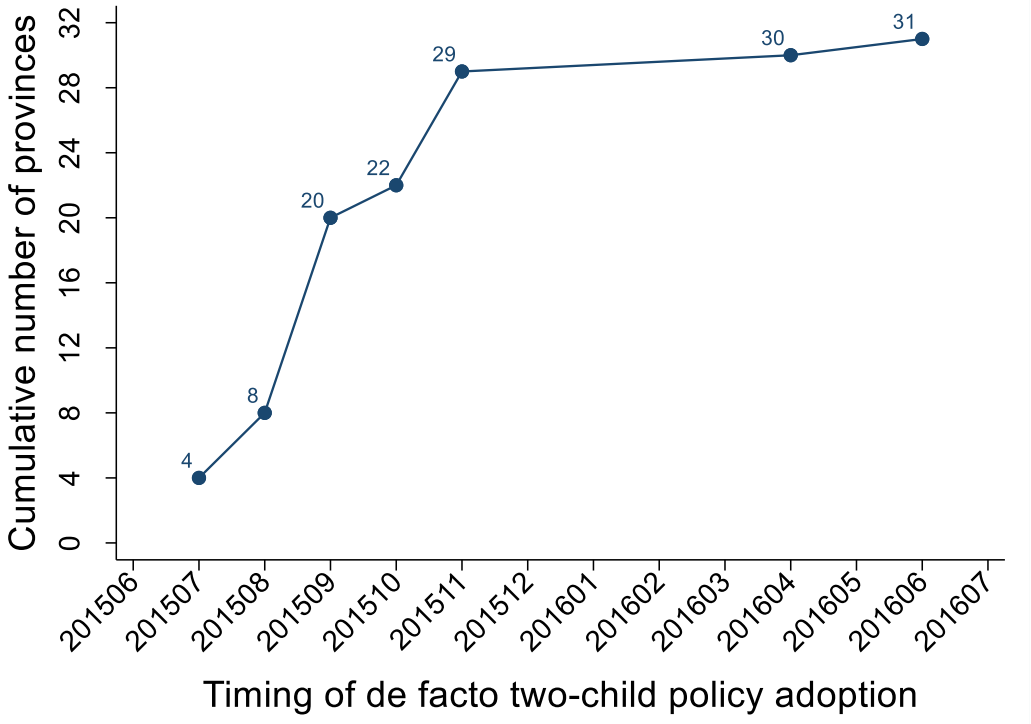

Figure A3 Timing of de facto two-child policy adoption across provinces

Notes: Timing of de facto two-child policy adoption = timing of launching the two-child policy+17 months

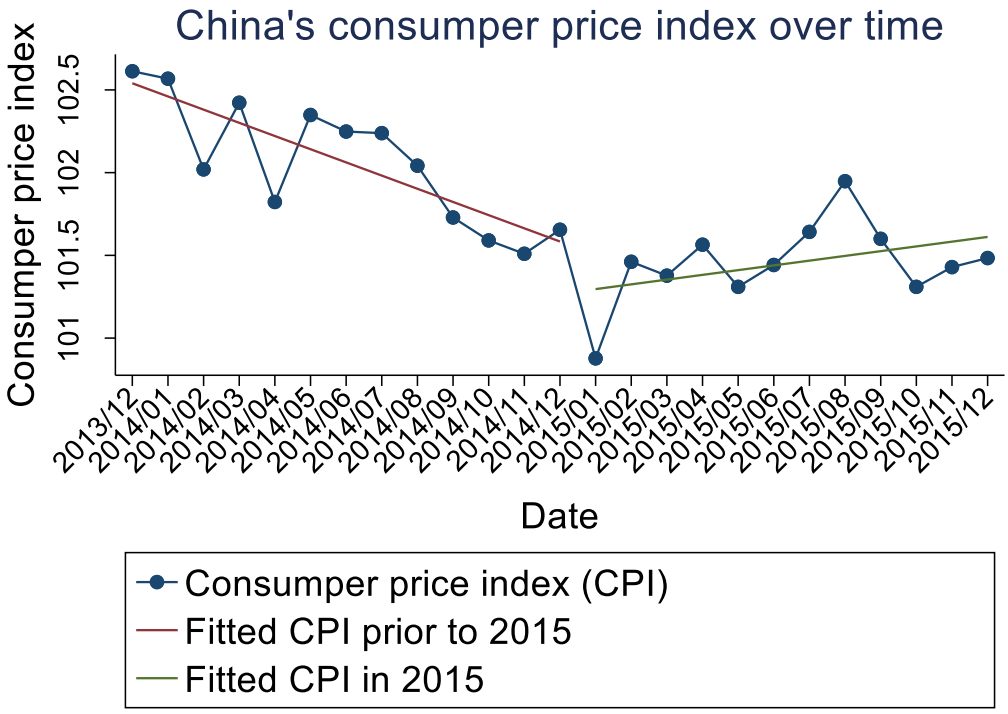

Figure A4 Local CPIs over time

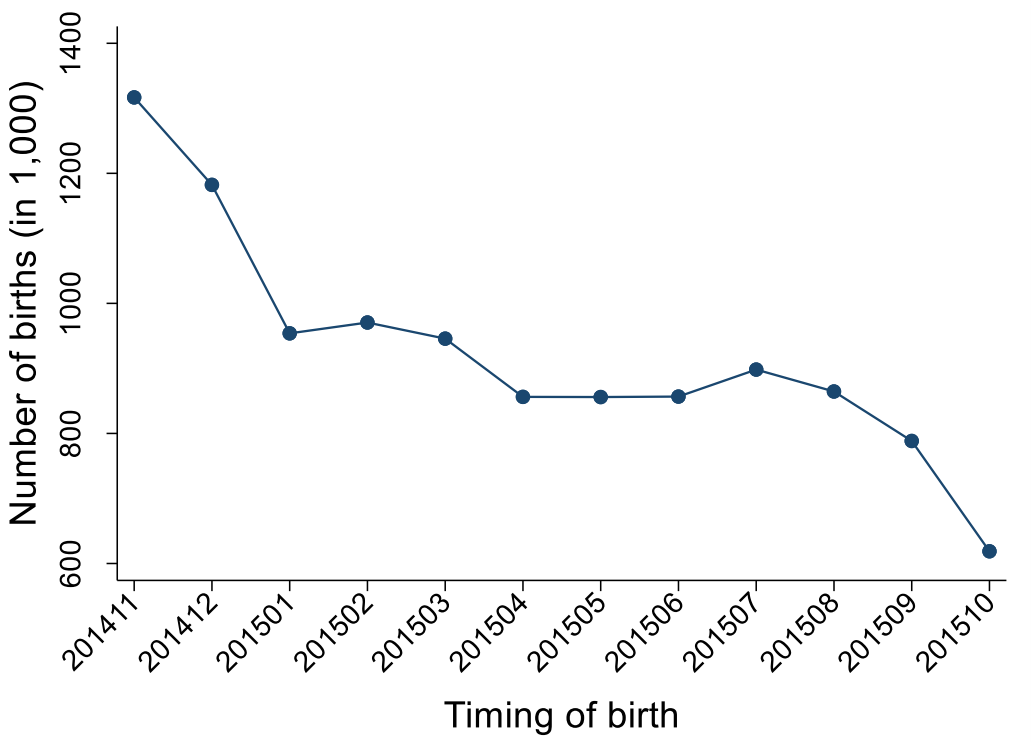

Figure A5 Number of monthly total births over time

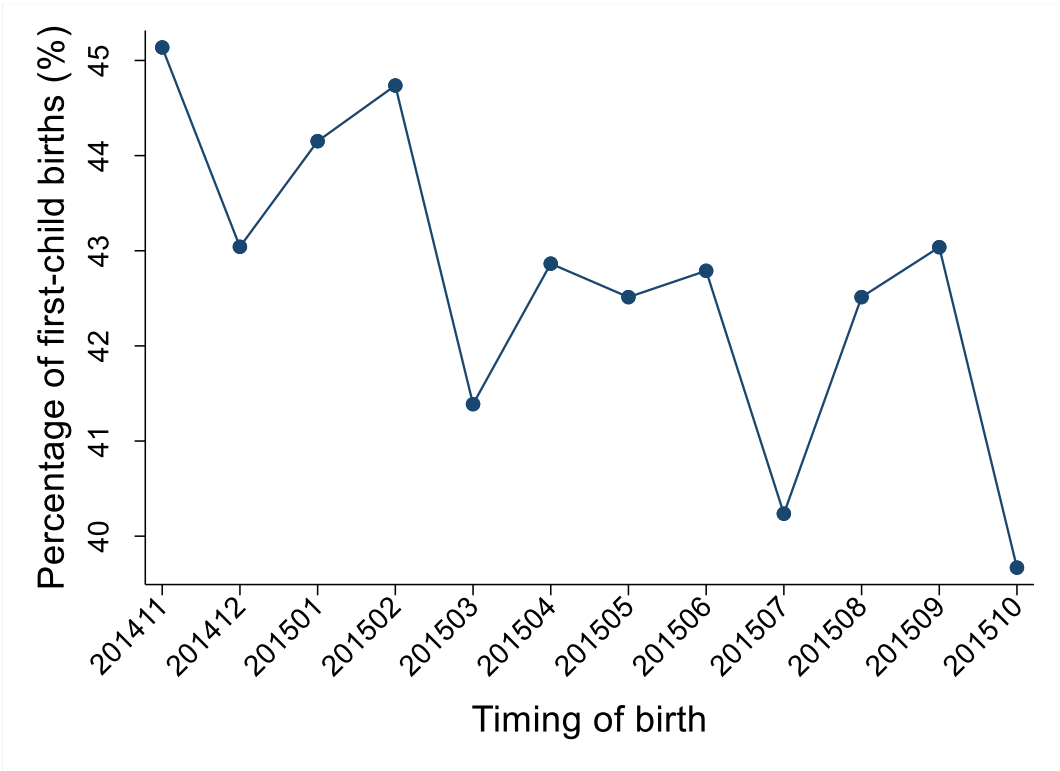

Figure A6 Percentage of first-child births in total births at the monthly level over time

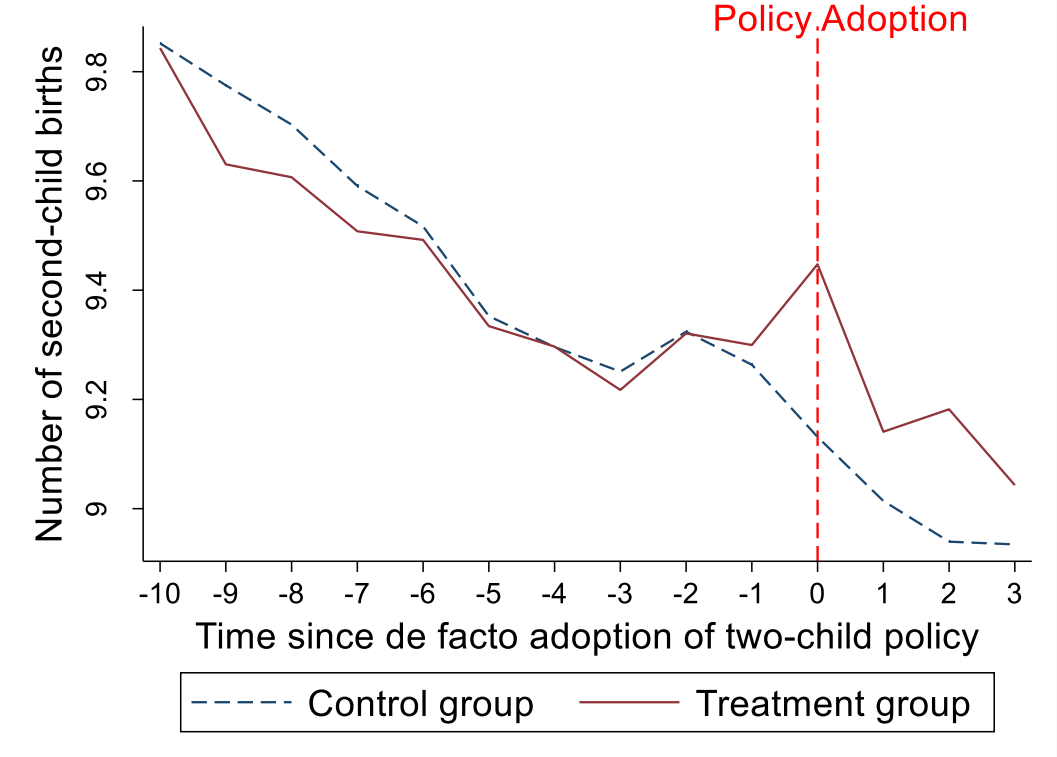

Figure A7 Number of second-child births (in log) between treatment and control groups over time

Note: The treatment group consists of provinces that launched the two-child policy between February 2014 and April 2014. The control group consists of provinces that launched the two-child policy after May 2014. We use the monthly second-child births (in log) between November 2014 and October 2015 as the outcome variable. Time since de facto adoption of two-child policy refer to 17 months after the adoption of the two-child policy.

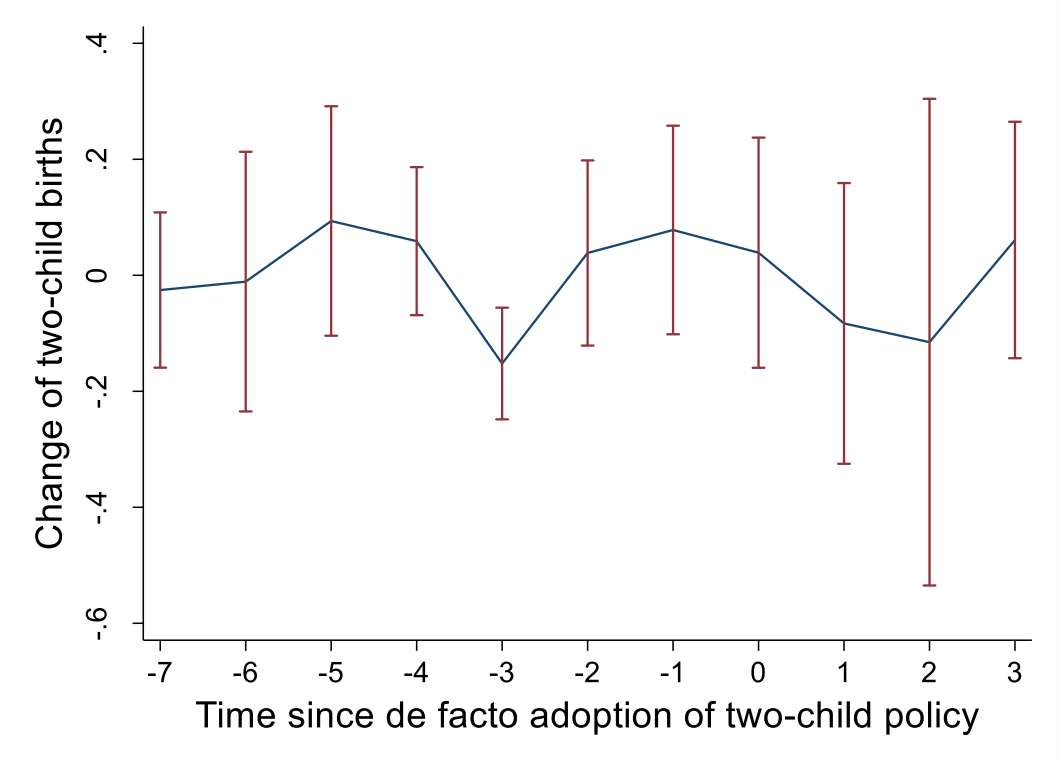

Figure A8 Placebo analysis on the impact of the adoption of the two-child policy on second-child births using census 2005

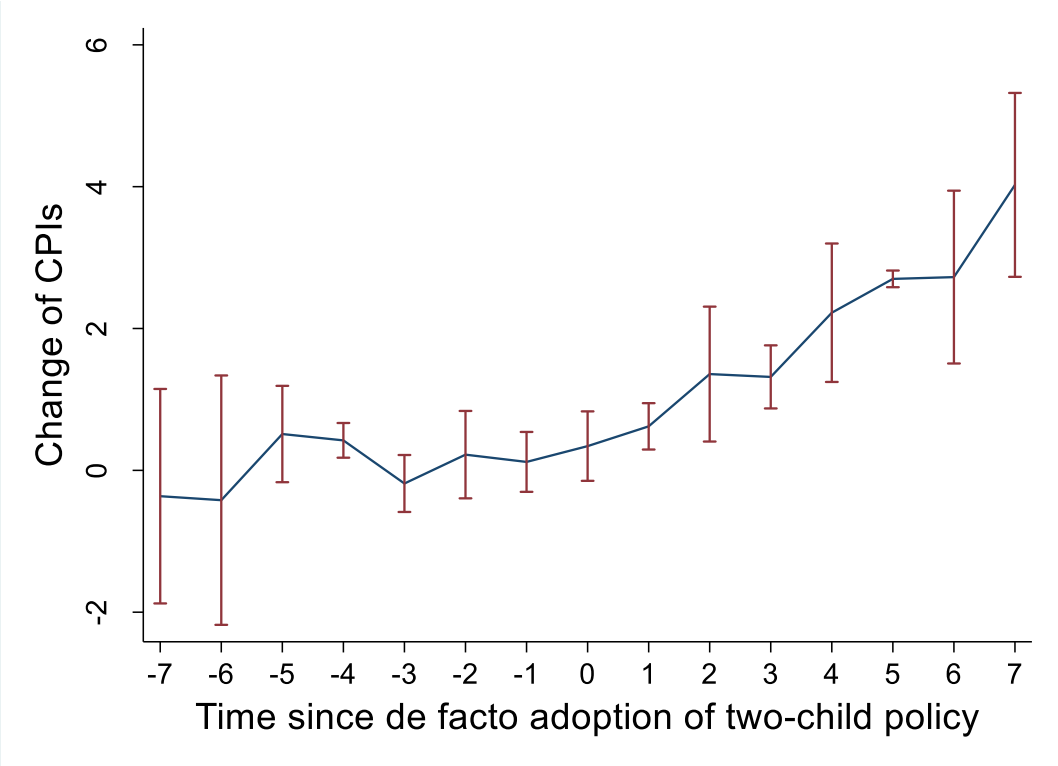

Figure A9 Impact of the adoption of the two-child policy on CPI of food etc.

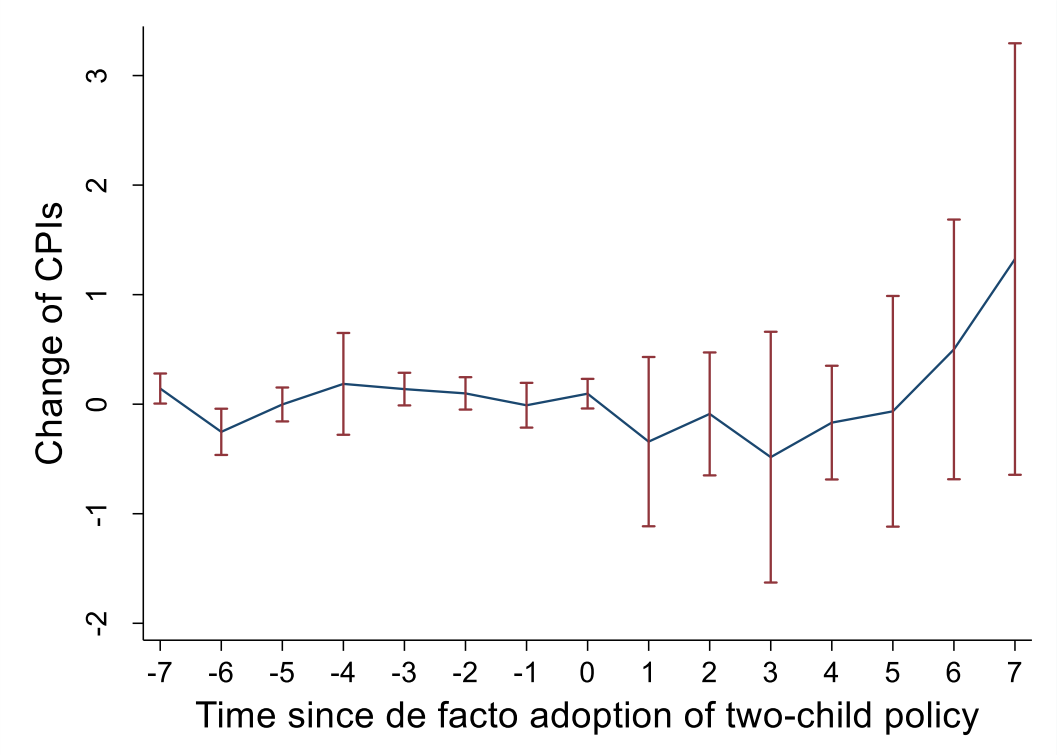

Figure A10 Impact of the adoption of the two-child policy on CPI of cigarette, alcohol etc.

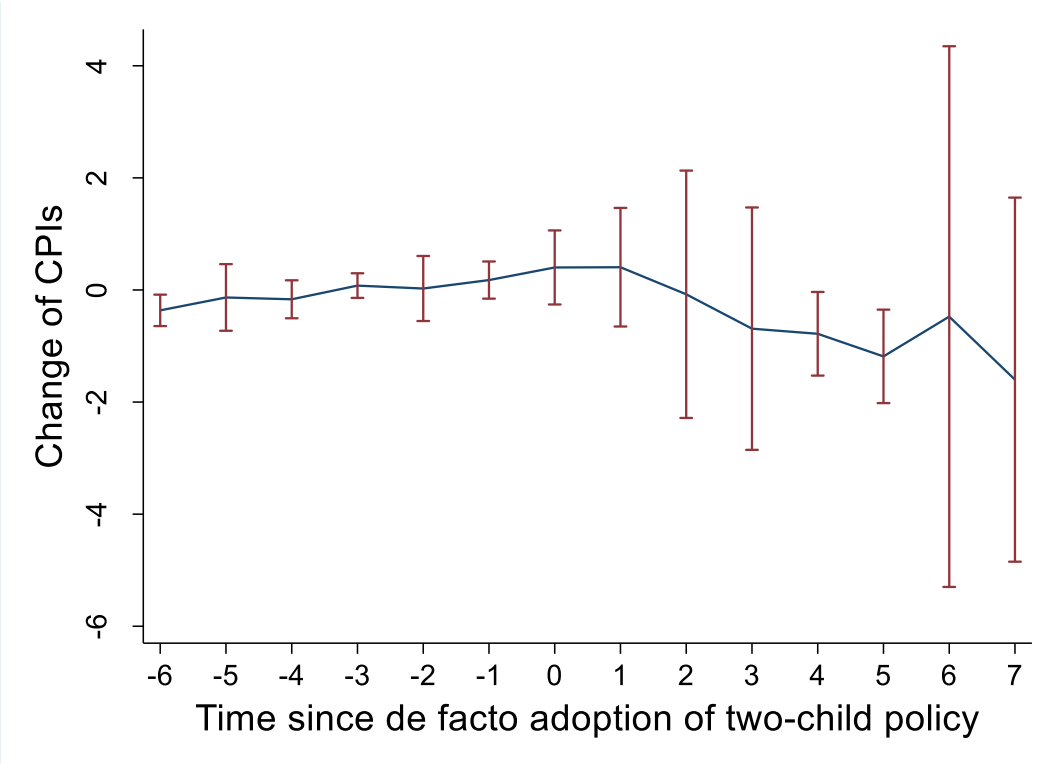

Figure A11 Impact of the adoption of the two-child policy on CPI of clothes etc.

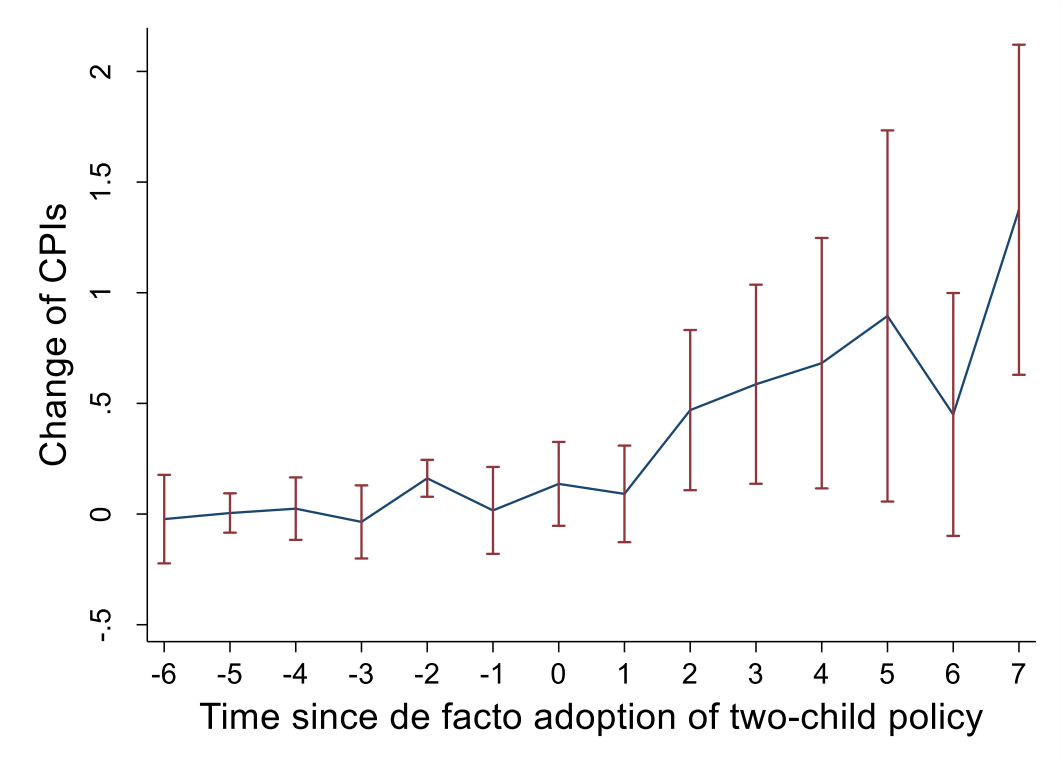

Figure A12 Impact of the adoption of the two-child policy on CPI of healthcare etc.

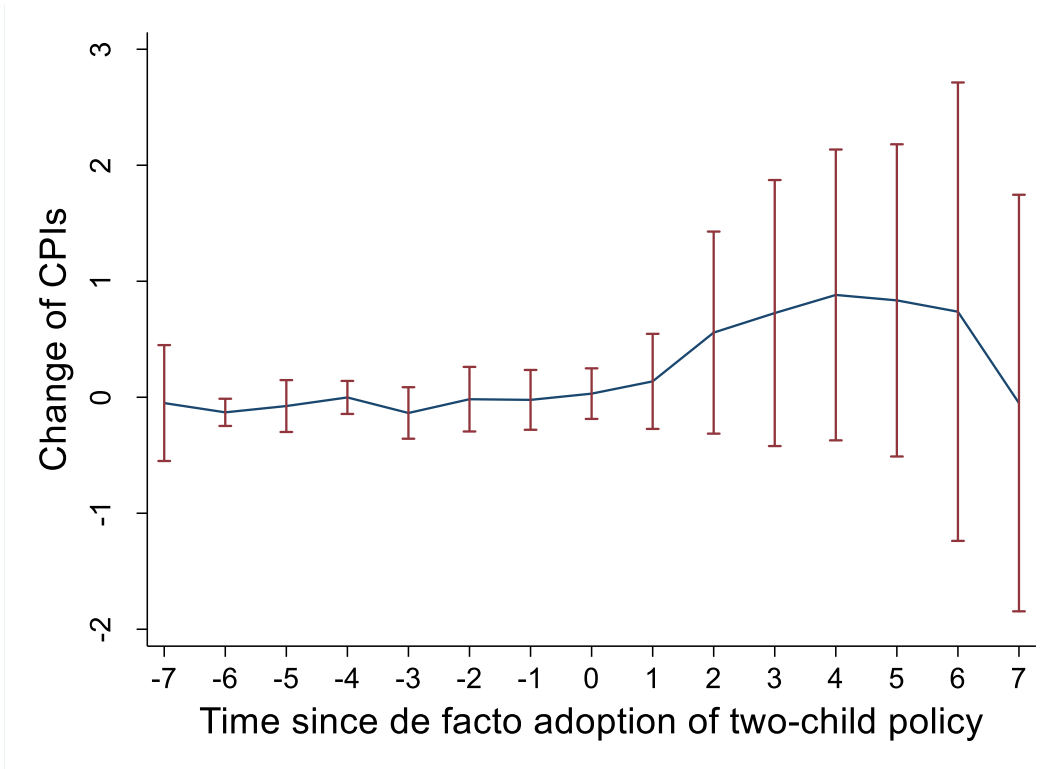

Figure A13 Impact of the adoption of the two-child policy on CPI of housing etc.

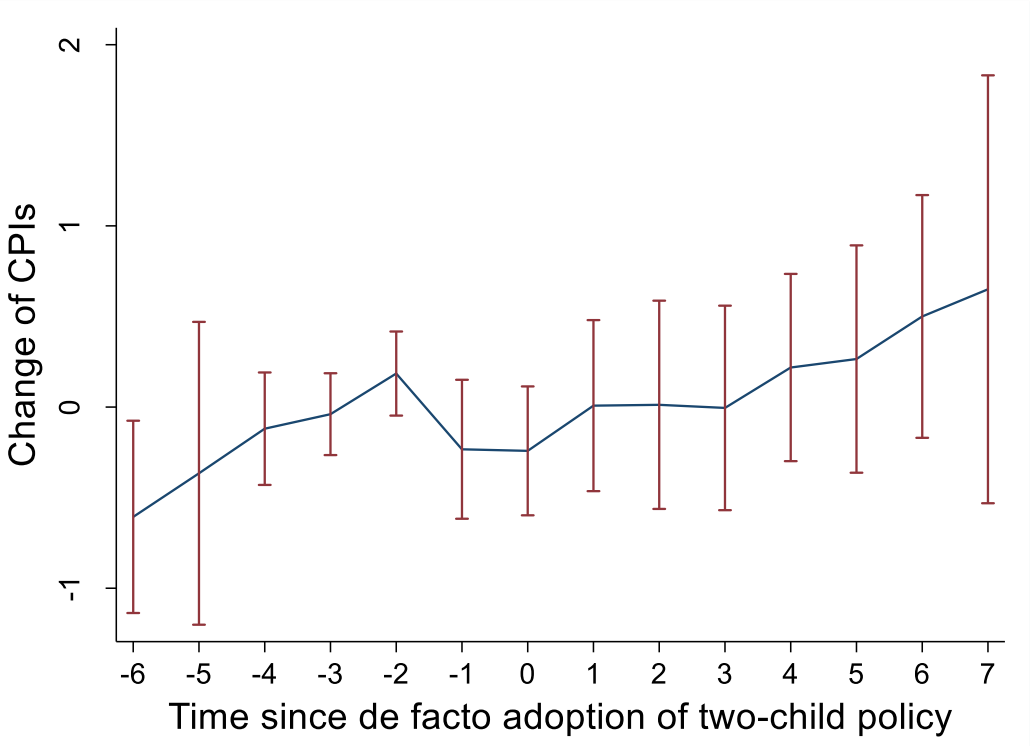

Figure A14 Impact of the adoption of the two-child policy on CPI of entertainment, education etc.

Appendix A. Proof for the Proposition 1

Denote B_1 as the solution for B given interest rate r in the equation (29), then we have $dB_1/dr > 0$, $\lim_{r\to 0}(B_1) = \Phi$, and $\lim_{r\to \infty}(B_1) = \infty$, where $\Phi = \frac{\varepsilon[(\alpha^*)^{(1/\varepsilon+1)} - \underline{\alpha}^{(1/\varepsilon+1)}]}{(\alpha^* - \underline{\alpha})(1+\varepsilon)}$. To get B and r, we further resort to the equation (28). Denote B_2 as the solution for B given interest rate r in the equation (28), then we have $dB_2/dr < 0$. When satifying $r \to 0$, we have $k \to \infty$ and $P \to 0$, then we have $\lim_{r\to 0} B_2 > 1$ from equation (28). When satifying $r \to \infty$, we have $k \to 0$ and $P \to \infty$, then we have $\lim_{r\to \infty} B_2 = -\infty$ from euqation (28). Figure A15 shows these two equations above. Therefore, if $\Phi < 1$, equations (28) and (29) do not exist a solution satifying $B = B_1 = B_2$, and the model does not exist a steady-state solution. If $\Phi \geq 1$, equations (28) and (29) show that the model exists a unique steady-state solution satifying $B = B_1 = B_2$.

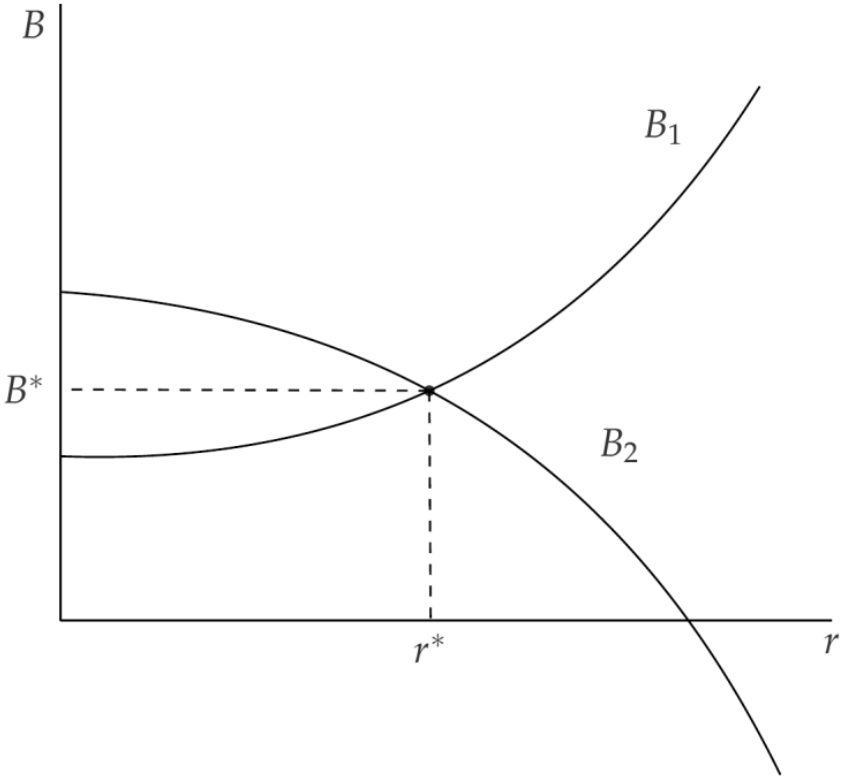

Figure A15 Existence of a steady-state solution

To further calculate the value of Φ, we have the utility fuction as $U(\alpha, 0) = c^{\sigma} \cdot$

$\left\{\frac{1}{1-\alpha\cdot[n(\alpha)]^{1-\varepsilon}}\right\}$. The bounded utility assumption suggests that $1>\alpha\cdot[n(\alpha)]^{1-\varepsilon}$. Combined with equation (22), we can find that there is a maximum interest rate r, satisfying $r<\bar{r}=min\{r(\alpha^*),r(\underline{\alpha})\}=min\{(1/\alpha^*)^{1/(1-\varepsilon)}-1,(1/\underline{\alpha})^{1/(1-\varepsilon)}-1\}=(1/\alpha^*)^{1/(1-\varepsilon)}-1$, given the condition $\alpha^*>\underline{\alpha}$. Since interest rate is postive, and $0<\varepsilon<1$, the condition $\alpha^*<1$ is satisfied. According to the defintion of Φ, we have $0=\lim_{\varepsilon\to 0}\Phi<\Phi<\lim_{\varepsilon\to\infty}\Phi=1$. Moreoever, according to equation (27), the utility is positive, indicating that consumtpion is positive. Thus, there exists a minimum interest rate r satisfying $r>\bar{r}=1/(\mu P)-1$. Therefore, given the upper and lower bound of interest rate, the model exists a unique steady-state solution.

Appendix B. Proof for the Proposition 2

First, it is obvious that $g_1(k)$ is a monotonically increasing function of k. Second, $\frac{dg_2(k)}{dk}\gtrless 0\Leftrightarrow\eta(2-\eta)k^{1-\eta}\left(1+\frac{1-\eta}{\eta k}\right)^{1/\varepsilon}+(\mu+\eta k^{2-\eta})\left(1+\frac{1-\eta}{\eta k}\right)^{1/\varepsilon-1}\left(-\frac{1-\eta}{\varepsilon\eta k^2}\right)\gtrless 0$

$$\Leftrightarrow\eta(2-\eta)k^{1-\eta}\left(1+\frac{1-\eta}{\eta k}\right)\gtrless\frac{1-\eta}{\varepsilon\eta k^2}(\mu+\eta k^{2-\eta})$$

$$\Leftrightarrow\eta(2-\eta)[k+(1-\eta)/\eta]\gtrless\frac{(1-\eta)}{\varepsilon\eta}[\mu k^{\eta-2}+\eta]$$

It is obvious that the left side of the inequation is a monotonically increasing function of k, and when $k\to 0$, the left side can be expressed as $(1-\eta)(2-\eta)$. When $k\to\infty$, the left side is equal to infinity. The right side of the inequation is a monotonically decreasing function of k. When $k\to 0$, the right side is equal to infinity; when $k\to\infty$, the right side is equal to zero. Therefore, there exists a $\bar{k}\in(0,\infty)$ so that $dg_2(k)/dk=0$. Therefore, $g_2(k)$ decreases with k firstly and then increases with k afterwards, as shown in Figure 1.

Moreover, according to L'Hopital's rule, we can have $\lim_{k\to 0}g_1(k)=-\infty$, $\lim_{k\to\infty}g_1(k)=\infty$, as well as $\lim_{k\to 0}g_2(k)=\infty>\lim_{k\to 0}g_1(k)$. Therefore, there will be a unique

intersection between $g_1(k)$ and $g_2(k)$ at $k \in (0, \bar{k})$ or $k \in (\bar{k}, \infty)$. In other words, the equation (30) exists a unique solution, which is also the unique steady-state solution of the model.

Appendix C. Proof for the Proposition 3

Denote the function $G(t) = [B(t)]^{\varepsilon}$. Given $\varepsilon > 0$, we have $sign(dG(t)/dk(t+1)) = sign(dB(t)/dk(t+1))$. Therefore, we have $\frac{dG(t)}{dk(t+1)} \gtrless 0$

$$\Leftrightarrow -\frac{1-\eta}{\eta[k(t+1)]^2}[k(t+1)]^{\sigma(1-\eta)}\left\{\mu P(t)\left[1+\frac{1-\eta}{\eta k(t+1)}\right]-1\right\}^{\sigma-1}$$

$$+\left[1+\frac{1-\eta}{\eta k(t+1)}\right]\sigma(1-\eta)[k(t+1)]^{\sigma(1-\eta)-1}\left\{\mu P(t)\left[1+\frac{1-\eta}{\eta k(t+1)}\right]-1\right\}^{\sigma-1}$$

$$+\left[1+\frac{1-\eta}{\eta k(t+1)}\right][k(t+1)]^{\sigma(1-\eta)}(1$$

$$-\sigma)\mu P(t)\frac{1-\eta}{\eta[k(t+1)]^2}\left\{\mu P(t)\left[1+\frac{1-\eta}{\eta k(t+1)}\right]-1\right\}^{\sigma-2} \gtrless 0$$

$$\Leftrightarrow -\frac{1-\eta}{\eta k(t+1)}\left\{\mu P(t)\left[1+\frac{1-\eta}{\eta k(t+1)}\right]-1\right\}+\sigma(1$$

$$-\eta)\left[1+\frac{1-\eta}{\eta k(t+1)}\right]\left\{\mu P(t)\left[1+\frac{1-\eta}{\eta k(t+1)}\right]-1\right\}$$

$$+(1-\sigma)\left[1+\frac{1-\eta}{\eta k(t+1)}\right]\left\{\mu P(t)\frac{1-\eta}{\eta k(t+1)}\right\} \gtrless 0$$

$$\Leftrightarrow \sigma[\eta k(t+1)+1-\eta]\left\{\mu P(t)\left[1+\frac{1-\eta}{\eta k(t+1)}\right]-1\right\}+(1$$

$$-\sigma)\mu P(t)\left[1+\frac{1-\eta}{\eta k(t+1)}\right] \gtrless \mu P(t)\left[1+\frac{1-\eta}{\eta k(t+1)}\right]-1$$

$$\Leftrightarrow \sigma[\eta k(t+1)-\eta]\left\{\mu P(t)\left[1+\frac{1-\eta}{\eta k(t+1)}\right]-1\right\}$$

$$+\sigma\left\{\mu P(t)\left[1+\frac{1-\eta}{\eta k(t+1)}\right]-1\right\} \gtrless \sigma\mu P(t)\left[1+\frac{1-\eta}{\eta k(t+1)}\right]-1$$

$$\Leftrightarrow \sigma[\eta k(t+1)-\eta]\left\{\mu P(t)\left[1+\frac{1-\eta}{\eta k(t+1)}\right]-1\right\}+1-\sigma \gtrless 0$$

According to equation (11), since the consumption must be positive, we then have $\mu P(t)\left[1+\frac{1-\eta}{\eta k(t+1)}\right]-1>0$ for all $k(t+1)>0$. Moreover, since $0<\sigma<1$, thus the sign of $dB(t)/dk(t+1)$ depends on $\eta[k(t+1)-1]$. In particular, there exists

a $k^* \in (0,1)$ such that for $k \in (0, k^*)$, $dG(t)/dk(t+1) < 0$; and for $k \in (k^*, \infty)$, $dG(t)/dk(t+1) > 0$. When $dG(t)/dk(t+1) < 0$ and thus $dB(t)/dk(t+1) < 0$, a lower $k(t+1)$ increases the fertility rate $B(t)$. When $dG(t)/dk(t+1) > 0$ and thus $dB(t)/dk(t+1) > 0$, a lower $k(t+1)$ decreases the fertility rate $B(t)$.