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Classification: D1, J11, J13

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1 Introduction

Fertility rates have decreased across countries and regions at all income levels worldwide over the past half-century, as shown in Fig. 1a. By 2021, the total fertility rate (TFR) for 110 out of 217 countries fell to or below the replacement level of 2.1.[1] This trend of low fertility presents a significant challenge for policymakers globally and has sparked extensive academic interest in exploring the causes of this long-term decline (Doepke *et al.*, 2023). Beyond shifts in social, economic, and cultural conditions, birth-control policies have also played a role in this global fertility decline. Between 1976 and 2013, 160 out of 217 countries implemented such policies in various forms (De Silva and Tenreyro, 2017). Fig. 1b shows the share of countries that implemented such policies by income level over the past half-century. This share has decreased for upper-middle-income countries and increased for low-income countries.[2]

We ask: Have these policies continued to reduce fertility even after their abolition? If so, what are the underlying reasons for this persistence? Understanding the answers to these questions is essential for identifying the causes and consequences of low fertility rates globally.

To address these questions, we examine the endogenous evolution of fertility norms by exploiting the unique empirical setting provided by the rollout and abolition of China's One-Child Policy (OCP). Theoretically, we demonstrate that a one-shot birth-control policy imposed on first-generation women can establish a lasting low-fertility norm that constrains the fertility of their daughters, even after the policy is lifted. Empirically, leveraging the

[1] Data source: World Development Indicator Database. The TFR is defined as the average number of children born to a woman over her lifetime.

[2] We categorize countries into income levels based on the earliest standard released by the World Bank.

2020 population census, we find that the OCP significantly reduces the fertility of subsequent generations in the post-policy era. Extensive mechanism analyses underscore the critical role of norms in driving the intergenerational persistence of low fertility.

Our analysis begins with a model that integrates the endogenous evolution of fertility norms. It examines how birth-control policies for first-generation women impact the fertility of their daughters (second generation) even after the policies are no longer in effect for daughters. In our model, a woman's fertility is determined not just by the economic costs and benefits of childbearing, but also by social norms for fertility. Following the "conformity" framework (Akerlof, 1997), we assume that first-generation fertility establishes a norm for the second generation. Second-generation fertility deviates from the "intrinsic" optimal level, which is determined solely by economic costs and benefits, and moves closer to the fertility level of her mother (first generation). In an economy with initially high fertility rates, second-generation fertility tends to surpass the intrinsic optimal level. However, birth-control policies that reduce first-generation fertility levels can establish new norms of lower fertility, and thereby lead to lower fertility rates among the second generation.

To empirically test our theoretical predictions, we require a one-shot birth control policy that impacts a single generation and is subsequently lifted for following generations. China's OCP, and its unexpected abolition, offers an ideal context for such analysis. The OCP's rollout across provinces in the late 1970s and early 1980s provides a quasi-experimental setting by directly imposing a fertility restriction on mothers of those born in the 1970s, and thereby reducing sibling sizes for the 1970s birth cohort. Easing of the OCP began in 2014 and culminated on December 27, 2015, with a proposed amendment to the Law of Population and Family Planning. This led to the unexpected introduction of a universal two-child policy on January 1, 2016, and thereby ended the OCP era. Importantly, those born in the 1970s cohort remained within their reproductive years post-abolition.

Our empirical analyses are based on multiple datasets. The recently released 2020 population census provides valuable data on the fertility behavior of the 1970s cohort following termination of the OCP. We also use the 2010 population census, 2005 and 2015 mini-censuses, China Family Panel Studies (CFPS) from 2012 to 2018, and the 1987 Fertility In-Depth Survey (FIDS).

We employ a cohort event-study design to identify the causal effect of first-generation exposure to the OCP on the fertility of second-generation women born in the 1970s, leveraging the rollout of the OCP across provinces in the late 1970s and early 1980s. We carefully measure and validate the timing of the OCP's rollout across provinces using a variety of sources, including government documents, *Population Chronicles*, and the *Encyclopedia of Chinese Family Planning*. In this setup, a first-generation woman's exposure to the OCP within a province is determined by her daughter's age when the policy was implemented: The younger a second-generation woman was when the OCP began, the less pre-policy time her mother had to bear additional children, and thus the stronger the mother's expected policy exposure.

We find that first-generation exposure to the OCP lowers second-generation fertility. Conditional on birth-province and birth-year fixed effects, fertility is up to 0.22 lower for second-generation born 1 year after OCP implementation, compared with those born 5 years before the OCP. This effect is substantial, especially given the notably low TFR of 1.06 in

China in 2022.[3] Because the OCP lowers first-generation fertility by 0.33, a Wald estimator suggests that a one-unit decrease in first-generation women's fertility leads to a 0.67 (0.22/0.33) reduction in second-generation women.

We conduct a battery of robustness checks. We check the non-anticipation and parallel trends assumptions, examine partial OCP exposure for second-generation women, and conduct heterogeneity-robust estimators in the recent event-study literature. Our empirical results remain robust to all of these identification concerns.

In our theoretical model, intergenerational fertility persistence comes from the evolution of fertility norms. To test this mechanism, we use stated fertility desires in a hypothetical scenario without any fertility restrictions as proxies for prevailing norms and track these norms across three generations. We find that the OCP leaves a lasting imprint: low fertility norms prevail among both the second and third generations. To explore what sets this evolution in motion, we examine the first generation's fertility intentions. Although the OCP lowered first-generation women's realized fertility, it did not affect their desired fertility under the hypothetical scenario without birth-control restrictions. Therefore, it is the forced reduction in actual births that generated the low-fertility norm rather than shifts in intentions.

We examine three alternative explanations that attribute intergenerational fertility persistence to changes in second-generation childbearing costs. Specifically, first-generation OCP exposure may result in (i) an increase in second-generation human capital via the child quantity-quality trade-off (Becker and Lewis, 1973); (ii) an increase in maternal labor

[3]Data source: Qiao (2023).

supply, which potentially influences second-generation labor market engagement via the intergenerational transmission of cultural traits (Fernández and Fogli, 2006, 2009); and (iii) an increase in coresidence between the second generation and their parents (Rosenzweig and Zhang, 2014). We do not detect significant effects of first-generation OCP exposure on second-generation educational attainment, labor-market outcomes, or coresidence behavior. Overall, the analyses suggest the evolution of fertility norms as the main mechanism behind the intergenerational effect of the OCP. We also investigate supplementary mechanisms, including intrahousehold bargaining, assortative mating, and mother-in-law effects.

Our study provides the first causal evidence on the intergenerational effects of birth-control policies, and complements the macro-level analysis by De Silva and Tenreyro (2020), who evaluate the role of these policies in the global fertility decline. Using microdata, we leverage the phased implementation and subsequent repeal of China's OCP as a quasi-experimental setting. Both our findings and those of De Silva and Tenreyro (2020) show a lasting reduction in fertility due to birth-control policies.

Our research contributes to the burgeoning literature on the diffusion of cultural norms that influence female behaviors (Fernández, 2011; Bisin and Verdier, 2023). A vibrant strand of research examines the transformation of female labor during the twentieth century, highlighting the intergenerational transmission of work behaviors from mothers to daughters (Fernández and Fogli, 2009); from mothers-in-law to daughters-in-law (Fernández *et al.*, 2004); and through social interactions that extend beyond family ties (Fogli and Veldkamp, 2011; Olivetti *et al.*, 2020). Recent work by Gay (2023), which exploits detailed data on internal migrants in France, demonstrates that these three mechanisms operate simultaneously in the diffusion of cultural norms on female labor force participation.

Research on the diffusion of fertility norms has explored such diffusion through religious groups (Munshi and Myaux, 2006); migrant networks (Daudin *et al.*, 2019); and cultural lineages (Beach and Hanlon, 2023; Spolaore and Wacziarg, 2022). Notably, Rossi and Xiao (2024) observe a spillover effect of China's "Later, Longer, Fewer" (LLF) Campaign on fertility among non-targeted minorities, while Sun *et al.* (2024) identify a similar spillover effect of China's "Selective Two-Child Policy," whereby fertility behaviors in targeted families influence those in non-targeted families. These studies explore variations in fertility norms that result from shifts in culture or policy and analyze how these changes are disseminated within peer groups, with a specific emphasis on the diffusion of fertility norms among peers of the same generation. We add to the literature by studying the intergenerational spillovers in fertility behavior, using the OCP as a source of exogenous variation in the fertility of the previous generation. We highlight the role of maternal fertility in shaping fertility norms.

Inspired by this literature, we further incorporate the peer effect on fertility in our model. In this expanded framework, the fertility norm for a daughter is a weighted average of her mother's fertility and the (expected) average fertility of her peer group. This extension yields two novel insights. First, although the peer effect facilitates further diffusion of the low fertility norm among the daughter's generation, it alone cannot generate the intergenerational effects of birth-control policies. Second, the peer effect on its own leads to multiple equilibria, as shown by Akerlof (1997); however, maternal fertility acts as a focal point, and thereby pins down a unique equilibrium for the daughter's generation. To disentangle these forces, we calibrate the model using peer parameters estimated via the strategy of Li and Zhang (2009) and Rossi and Xiao (2024). The results reveal that at least 20% of our estimated effect of maternal OCP exposure on daughters' fertility is due to the maternal effect alone.

Our study serves as a bridge between two strands of the literature. The first examines the impact of public policies on cultural practices, as exemplified by Bau (2021), who demonstrates that pension policies influence post-marriage residences in societies with matrilocal or patrilocal customs. The second strand explores intergenerational cultural persistence, with studies such as those by Fernández *et al.* (2004); Fernández and Fogli (2009); and Giuliano (2007), who demonstrate how source-country traits determine fertility rates, labor force participation, and the living arrangements of second-generation immigrants in the U.S. These two strands of the literature jointly imply that a one-shot policy can have enduring effects through its influence on culture and the persistence of cultural values across generations. This implication is theoretically and empirically explored in our study.

Our analysis also relates to the extensive literature on the incidence and impact of the OCP. The OCP literature begins with assessments of its immediate impact on fertility reduction (Ahn, 1994; McElroy and Yang, 2000). While initial studies debate the extent of this impact, more recent work confirms a substantial effect of the OCP in reducing fertility among targeted populations (García, 2022). Economists have since explored the broader implications of the OCP, extending beyond its influence on fertility alone. The policy has been shown to drive sex ratio imbalance (Ebenstein, 2010; Li *et al.*, 2011); alter urban-rural population structure (Wang and Zhang, 2018); spur internal migration (Wang *et al.*, 2017; Zhang, 2017; Huang *et al.*, 2024; Guo *et al.*, 2024b); and distort the marriage market (Bhaskar *et al.*, 2023; Huang *et al.*, 2023). A strand of literature has examined the intergenerational impact of the OCP on children's education, health, and occupational choices (Liu, 2014; Zhao and Zhou, 2018; Huang *et al.*, 2021; Huang, 2022). More recently, Yin (2024) documents the intergenerational persistence of policy compliance, finding that second-generation individuals from smaller families were less likely to have unauthorized births prior to the policy's abolition.

Our study adds three specific contributions to the OCP literature. First, we compile a novel dataset that documents the phased rollout of the OCP, with the date each provincial government formally imposed a one-child mandate and applied significant penalties for unauthorized second births. Second, we leverage fertility data from the 2020 census, which offers a unique 5-year window to observe unrestricted fertility behaviour following the policy's abolition——a critical period for studying the intergenerational fertility effects of the OCP. Third, we identify the intergenerational diffusion of fertility norms as a central mechanism that determines fertility in subsequent generations.

Our study highlights two major policy implications. First, Fig. 1b shows that many less developed countries have recently implemented birth-control policies to reduce fertility (Rossi and Godard, 2022). While initially effective, our research points to their lasting impact on fertility norms by leading to sustained low fertility rates even after policy removal. This underscores the need for ongoing policy adjustments to prevent irreparable changes in fertility norms.

Second, the prevalence of low fertility norms may undermine the effectiveness of traditional pro-natal policies, such as childcare subsidies and extended maternal leave, which target individual households rather than groups of peers who share a common norm. For example, even large sums of childcare subsidies may not boost a recipient's fertility if she does not expect her peers to have more children. New public campaigns can reshape these expectations, providing a focal point to sustain higher fertility rates. In line with the classic growth theory, which advocates for the simultaneous industrialization of multiple sectors (Murphy *et al.*, 1989), our study suggests that pro-natal policies also require a "big push" to effectively reshape fertility norms.

2 Theory

This section develops the theoretical framework for our analysis. We begin by building a model in which maternal fertility acts as a social norm shaping a woman's fertility choice. We then analyze the model's dynamics to demonstrate how one-shot birth-control policies can generate lasting intergenerational effects. Finally, we extend the framework to include peer effects, examining how social interactions among peers amplify the intergenerational policy impact.

2.1 A Model of the Evolution of Fertility Norms

Consider a woman in generation t who chooses her consumption (c_t) and fertility (n_t) to maximize her utility:

$$\begin{aligned} \max_{c_t, n_t} U_t = \; & \ln(c_t) + \gamma \ln(n_t) - \sigma |\ln(n_t) - \ln(n_{t-1})|, \\ \text{s.t.} \quad & c_t + p_t n_t = y_t, \end{aligned} \tag{P1}$$

where γ represents the degree of maternal altruism toward children, p_t is childbearing cost, y_t is her income, and n_{t-1} denotes the fertility of her mother or the number of her siblings. The price of consumption is normalized to 1. In this model (P1), the mother's fertility serves as a norm that influences the daughter's fertility choices. The penalty term ($-\sigma|\ln(n_t) -$

$\ln(n_{t-1})|)$ captures the utility loss at a rate of σ for every log point deviation of the woman's fertility from her mother's fertility (n_{t-1}).

Our model's formulation is analogous to those of Akerlof (1997); De Silva and Tenreyro (2020); and Spolaore and Wacziarg (2022). In Akerlof (1997) and Spolaore and Wacziarg (2022), the norm is defined as the expected mean fertility for her generation ($\bar{n}_t$), whereas De Silva and Tenreyro (2020) define it as a weighted average of the replacement level of fertility (2.1) and the woman's mother's fertility. By contrast, we define the norm as n_{t-1}, which emphasizes the evolution of fertility norms over generations. We also consider the norm to be the weighted average of n_{t-1} and $\bar{n}_t$ when allowing for the peer effect on fertility in later sections.[4]

In the special case in which $\sigma = 0$, the fertility norm n_{t-1} does not influence the woman's fertility n_t. The optimal fertility n_t^* is only determined by the intrinsic economic costs and benefits, given by $n_t^* = \frac{\gamma}{1+\gamma}\frac{y_t}{p_t}$. This intrinsic optimal fertility declines when (i) the childbearing cost p_t rises relative to income y_t (i.e., higher p_t/y_t) or (ii) the preference for children diminishes (i.e., lower γ).

Generally, we derive the optimal fertility (n_t^*) through the following proposition, under the assumption that n_{t-1} is exogenous to the current generation:[5]

[4]De Silva and Tenreyro (2020) include the squared difference between n_t and the norm in their penalty term, while Akerlof (1997) and Spolaore and Wacziarg (2022) use the absolute value of the difference. We follow the latter, as discussed by Akerlof (1997).

[5]We exclude any dynastic interactions except for the evolution of fertility norms over generations. This assumption is justified in our empirical context, given that the birth-control policy exogenously changes the fertility of the woman's mother.

Proposition 1. *When* $\sigma \geq 0$,

$$n_t^* = \begin{cases} n_t^u, & \text{if} \quad n_{t-1} > n_t^u, \\ n_{t-1}, & \text{if} \quad n_t^l \leq n_{t-1} \leq n_t^u, \\ n_t^l, & \text{if} \quad n_{t-1} < n_t^l, \end{cases}$$

where $[n_t^l, n_t^u] = [\frac{\gamma-\sigma}{1+\gamma-\sigma}\frac{y_t}{p_t}, \frac{\gamma+\sigma}{1+\gamma+\sigma}\frac{y_t}{p_t}]$.

Proof. See Appendix Section A.1.

The proposition demonstrates how maternal fertility serves as a norm that influences the daughter's fertility. Specifically, the daughter fully mimics her mother's fertility n_{t-1} when $n_{t-1} \in [n_t^l, n_t^u]$, such that $n_t^* = n_{t-1}$. When $n_{t-1} \notin [n_t^l, n_t^u]$, she partially mimics the norm. If the mother's fertility is sufficiently high (low), such that $n_{t-1} > n_t^u$ ($n_{t-1} < n_t^l$), the daughter's optimal fertility is n_t^u (n_t^l), which is larger (smaller) than the intrinsic optimal fertility.

2.2 Model Implications for the Lasting Effect of Birth-control Policies

Fig. 2 illustrates the lasting effect of birth-control policies. Our model explains this effect. For illustration, consider a sequence of four generations, indexed by 0, 1, 2, and 3, in a society. Based on Proposition 1, we define the fertility interval for each generation as $[n_t^l, n_t^u] = [\frac{\gamma-\sigma}{1+\gamma-\sigma}\frac{y_t}{p_t}, \frac{\gamma+\sigma}{1+\gamma+\sigma}\frac{y_t}{p_t}]$, $\forall t = 0, 1, 2, 3$.

For a compact presentation of our main theoretical results, define the time path of childbearing cost-income ratio as

$$\frac{p_t}{y_t} = (1 + \Psi_t) \cdot \frac{p_{t-1}}{y_{t-1}},$$

where Ψ_t is the growth rate of the cost-income ratio $\frac{p_t}{y_t}$ at generation t for $t = 1, 2, 3$. Recursive iteration yields $\frac{p_t}{y_t} = \frac{p_0}{y_0} \cdot \Pi_{\tau=1}^{t}(1 + \Psi_\tau)$. Based on this formula, we express the fertility interval at generation t as functions of the fertility interval at the initial generation 0:

$$n_t^u = \frac{n_0^u}{\Pi_{\tau=1}^{t}(1 + \Psi_\tau)},$$
$$n_t^l = \frac{n_0^l}{\Pi_{\tau=1}^{t}(1 + \Psi_\tau)}.$$

It is evident that Ψ_t determines the rate at which the fertility interval declines.

We make assumptions to align our theoretical framework with the empirical context of China. The OCP was introduced during a prolonged period of fertility decline (Section 3.1). To reflect this, we assume that Ψ_t is positive for generations $t = 1, 2, 3$, which results in a downward shift in fertility intervals across these three generations. As an initial condition, the fertility level for generation 0 starts at $\tilde{n}_0$, which should fit the total fertility rate of approximately 6 per woman around the founding years of the People's Republic of China (Yin, 2023). If the fertility behavior of generation 0 is in equilibrium, then $\tilde{n}_0$ should lie within the range $[n_0^l, n_0^u]$. For the sake of simplicity, we also assume that Ψ_1 is large enough to reduce fertility for generation 1. A summary of these assumptions is presented in Assumption 1.

Assumption 1.

(a) The growth rate of the cost-income ratio is positive for each generation,

$$\Psi_t > 0 \ , \forall i = 1, 2, 3;$$

(b) the fertility of generation 0 starts at $\tilde{n}_0$, such that $\tilde{n}_0 \in [n_0^l, n_0^u]$; and

(c) the price change from generation 0 to 1 is large enough, such that $p_1 > \frac{\gamma+\sigma}{1+\gamma+\sigma}\frac{y_1}{\tilde{n}_0}$.

The solid line, OABC in Fig. 2, represents the "natural" equilibrium fertility rates without birth-control policies. Point O denotes the initial fertility level $\tilde{n}_0$ for generation 0. By Assumption 1a, the fertility interval $[n_t^l, n_t^u]$, where $t = 1, 2, 3$, shifts downward. By Assumption 1c, Proposition 1 states that the equilibrium fertility rates for generations 1, 2, and 3 land on the upper bounds of the fertility intervals, represented by points A, B, and C.

The dashed line, OA'B'C', represents the equilibrium fertility rates with a one-shot birth-control policy on generation 1. Comparing the dashed line with the solid line, we observe a lasting effect of the policy. Because of the policy, we assume that the fertility of generation 1 is constrained to be n^p such that $n^p < n_1^l$ (point A'). When the policy is removed after this generation, the fertility of generation 2 does not bounce back to point B, the natural equilibrium fertility rate. This is because the constrained low fertility of generation 1 establishes a new norm for generation 2. The equilibrium fertility is n_l^2 for both generations 2 and 3. Fig. 2 shows that n_2^l is smaller than both n_2^u and n_3^u.

We summarize the long-lasting impact of the birth-control policies in the following proposition:

Proposition 2. *Under Assumption 1,*

(a) Without a birth-control policy, the equilibrium fertility rates are

$$n_t^* = n_t^u = \frac{n_0^u}{\Pi_{\tau=1}^t (1+\Psi_\tau)}, \quad \forall t = 1, 2, 3.$$

(b) When a birth-control policy restricts the fertility of generation 1 at $n^p < n_2^l$, but is removed afterward, the equilibrium fertility rates at generations 2 and 3 are

$$n_2^* = n_3^* = n_2^l = \frac{n_0^l}{(1+\Psi_1)(1+\Psi_2)}.$$

Proof. See Appendix Section A.3.

Because $\sigma > 0$, $n_2^l < n_2^u$: The one-shot birth-control policy at generation 1 lowers fertility for generation 2. If we assume $n_2^l < n_3^u$, the policy further lowers fertility for generation 3.[6] The theoretical prediction of Proposition 2 holds robustly across broader scenarios. Appendix A Fig. A2 illustrates the lasting effect of the one-shot birth-control policy in "stable cost" scenario in which p_t remains constant across all generations ($\Psi_t = 0$). In this stable setting, as long as the initial fertility level exceeds the lower-bound fertility of generation 0 ($\tilde{n}_0 > n_0^l$), a one-time birth-control policy implemented in generation 1 results in reduced fertility for generations 2 and 3.[7]

Our model is adaptable to include parental investment in child human capital, and thereby incorporates a mechanism of the child quantity-quality (QQ) trade-off (Becker and

[6] The condition $n_2^l < n_3^u$ imposes a restriction on the growth rate of the cost-income ratio for generation 3, requiring $\Psi_3 < \frac{2\sigma}{(\gamma-\sigma)(1+\gamma+\sigma)}$. Also note that the theoretical prediction of Proposition 2 does not hinge on the assumption $n^p < n_2^l$. As long as $n^p < n_2^u$, the policy for generation 1 reduces equilibrium fertility for subsequent generations.

[7] For more details, please refer to Appendix Section A.4.

Lewis, 1973).[8] If this mechanism is at work, a birth-control policy could lead to reduced fertility in the first generation, increased investment per child, enhanced human capital in the second generation, and subsequently lower fertility in the second generation. In Section 6.2, we analyze this mechanism within our empirical setting but find no significant effect of the birth-control policy on the human capital of the second generation. This aligns with other studies that investigate the QQ trade-off in the Chinese context (Rosenzweig and Zhang, 2009), which indicates that this mechanism does not drive our empirical findings. For the sake of simplicity, our theoretical model abstracts from child human capital investment.

2.3 Model Extension: Incorporating the Peer Effect

We now extend our model to incorporate the peer effect in shaping fertility norms, which has been extensively investigated in the literature (Richerson and Boyd, 2008; Li and Zhang, 2009; Spolaore and Wacziarg, 2022; Beach and Hanlon, 2023; Rossi and Xiao, 2024). Specifically, the utility maximization problem for woman i in generation t becomes

$$\begin{aligned} \max_{c_{i,t}, n_{i,t}} \quad & U_{i,t} = \ln(c_{i,t}) + \gamma \ln(n_{i,t}) - \sigma |\ln(n_{i,t}) - \ln(N_{i,t})|, \\ \text{s.t.} \quad & c_{i,t} + p_t n_{i,t} = y_t, \end{aligned} \tag{P2}$$

where the fertility norm ($N_{i,t}$) is a weighted average of the mother's fertility ($n_{i,t-1}$) and the (expected) fertility of peers ($\bar{n}_t$), such that

$$N_{i,t} = \alpha n_{i,t-1} + (1-\alpha)\bar{n}_t, \tag{1}$$

[8] For a comprehensive review of the child QQ trade-off, see Guo *et al.* (2022) and Doepke (2015).

where $\alpha \in [0,1]$ represents the strength of maternal influence relative to the peer effect on the fertility norm. The solution to P2 is identical to that to P1, stated in Proposition 1, when a society starts at the steady state for generation $t-1$ — i.e., $n_{i,t-1} = n_{j,t-1}, \forall i \neq j$. Because of the peer effect, generation t is also in a steady state, $n_{i,t} = \bar{n}_t$.[9]

We examine a scenario in which the society deviates from the steady state for generation 1, due to the disruption caused by the birth-control policy. When the policy is not uniformly applied to all women in generation 1, differences in fertility emerge, which enables peer effects to play an active role in reducing these differences. For simplicity, we assume that a birth-control policy restricts the fertility rate of a proportion $\lambda \in (0,1)$ of women at generation 1 to $n^p < n_2^l$.[10] The policy does not restrict the remaining $1-\lambda$ proportion of women.

In this case, two fertility norms emerge for women in generation 2. For those who are daughters of mothers treated by the policy, the norm is $N_{r,2} = \alpha n^p + (1-\alpha)\bar{n}_2$. For the remaining women, it is $N_{ur,2} = \alpha n_1^u + (1-\alpha)\bar{n}_2$. As $n^p < n_1^u$, $N_{r,2} < N_{ur,2}$. Because they have different norms, the optimal fertility differs between these two types of women. The solution to P2 now becomes a Bayesian Nash equilibrium (Blume *et al.*, 2011), in which women in generation 2 have rational expectations that the exante expected average fertility of their peers ($\bar{n}_2$) coincides with the expost realized average fertility, such that

$$\bar{n}_2 = \lambda n_{r,2}^* + (1-\lambda) n_{ur,2}^*,$$

and $n_{r,2}^*$ ($n_{ur,2}^*$) is the equilibrium fertility for women whose mothers' fertility was (not) constrained by the policy.

[9]See Appendix Sections A.5 and A.6 for the formal proof.

[10]If the birth-control policy restricts every woman in generation 1 to n^p, incorporating the peer effect does not affect our analysis in the above section.

The proposition below characterizes, and Fig. 3 plots, equilibrium fertility for generation 2 in this case.

PROPOSITION 3. *(a) When $\alpha = 0$, the maternal effect does not exist (Fig. 3a). The equilibrium fertility for generation 2, $n^*_{r,2} = n^*_{ur,2} \in [n^l_t, n^u_t]$. Multiple equilibria exist.*

(b) When $\alpha = 1$, the peer effect does not exist (Fig. 3b). $n^*_{r,2} = n^l_2 < n^*_{ur,2} = n^u_2$.

(c) When $\alpha \in (0, \alpha_1)$, $n^*_{r,2}$ decreases with λ when $\lambda < \lambda_1$, then stays constant at n^l_2; $n^*_{ur,2}$ stays constant at n^u_2 when $\lambda < \lambda_2$, then decreases with λ (Fig. 3c).

(d) When $\alpha \in [\alpha_1, 1)$, $n^*_{r,2}$ decreases with λ when $\lambda < \lambda_3$, then stays constant at n^l_2; $n^*_{ur,2}$ stays constant at n^u_2 when $\lambda < \lambda_4$, then decreases with λ (Fig. 3d).

We have $\alpha_1 = \frac{n^u_2 - n^l_2}{n^u_1 - n^p}$, $\lambda_1 = 1 - \frac{n^l_2 - n^p}{(1-\alpha)(n^u_1 - n^p)}$, $\lambda_2 = \frac{n^u_1 - n^u_2}{(1-\alpha)(n^u_1 - n^p)}$, $\lambda_3 = 1 - \frac{\alpha(n^l_2 - n^p)}{(1-\alpha)(n^u_2 - n^l_2)}$, $\lambda_4 = \frac{\alpha(n^u_1 - n^u_2)}{(1-\alpha)(n^u_2 - n^l_2)}$, *where* $\lambda_1 > \lambda_3$, *and* $\lambda_2 < \lambda_4$.

Proof. See Appendix Sections A.5–A.8.

Based on the proposition, we have five observations. (1) The average fertility rate of generation 2 decreases with a one-shot birth-control policy that constrains fertility for λ proportion of women at generation 1, as long as the maternal effect exists ($\alpha > 0$). This holds true irrespective of the existence of the peer effect. (2) The policy does not affect the average fertility rate of generation 2 when only the peer effect exists ($\alpha = 0$). (3) The optimal fertility for those whose mothers are untreated ($n^*_{ur,2}$) decreases with the policy as long as the share of women whose mothers are treated (λ) is sufficiently large. (4) The effect of the one-shot birth-control policy extends naturally beyond generation 2, since the decreased fertility for generation 2 reduces the norms for generation 3 through the maternal effect. (5) Our results for multiple equilibria when the peer effect exists only echo those of

Akerlof (1997). However, when $\alpha > 0$, maternal fertility serves as a focal point that pins down a unique equilibrium in our extended model.

In exploring the unique institutional context characterized by the staggered rollout of the OCP among Han Chinese in the late 1970s and early 1980s, followed by its termination in 2016, our main empirical analysis below tests the prediction based on Proposition 2. We also use the differential implementation between Han and ethnic minorities to test some predictions based on Proposition 3. Finally, we use Proposition 3 to quantitatively separate maternal and peer effects in accounting for the long-lasting effect of the OCP on women whose mothers' fertility was directly constrained by the policy, but they themselves are no longer constrained by the policy after 2016.

3 Institutional Background

To test our theoretical prediction, a one-shot birth-control policy is required—one that affects a single generation and is removed for subsequent generations. China's OCP, followed by its sudden abolition in 2016, provides a unique context for such an analysis. This section first describes the rollout of the OCP in the late 1970s and early 1980s, measures the implementation timing of the policy across provinces, and finally describes its nationwide termination in 2016.

3.1 History of the Birth-control Policy in China

China launched its family planning campaign, the "Later, Longer, Fewer" Campaign, in 1971 (Zhang, 2017).[11] In response to fears of an overpopulation crisis, Xiaoping Deng, then the new leader of China, initiated a more stringent family planning program in 1979. This radical initiative, aimed at restricting most families to a single child, was one of the most extreme birth-control measures in recorded history. It encompassed a comprehensive approach that involved widespread propaganda, regulatory measures, incentives for compliance, and punitive actions, including financial and employment-related penalties. Despite its intent, the program did not fully achieve a universal one-child norm across the country. However, the collective set of these radical population control policies since 1979 has been commonly referred to as the OCP.

The OCP, since its inception, faced significant public resistance. To mitigate this, the government, starting in the 1980s, relaxed the policy for rural residents and ethnic minorities. The central committee's 1984 Document No. 7 allowed rural families, especially those with a first-born daughter, to have a second child, and ethnic minorities were generally permitted more than one child (Scharping, 2013). Despite these relaxations, the policy remained strict for urban populations. Prime Minister Li Peng in 1989 reaffirmed a population target of 1.2 billion by the end of the century, aiming to limit the natural growth rate to below 1.25%. The 2002 Law of Population and Family Planning further legalized the OCP and maintained the stringent one-child rule in urban areas through the 1990s and 2000s.[12]

[11]This slogan advocated for late marriage and childbearing, extended birth spacing (at least 3 years between two births), and limiting the number of children to no more than two. The program began with the State Council's Document 51, "Report on Better Implementing Family Planning Policy," released in 1971.
[12]See Appendix B Table B1 for more details on the OCP rollout across China.

3.2 Measuring the Rollout Timing of the OCP across Provinces

The phased implementation of the OCP across provinces offers a quasi-experimental setting to examine our theoretical predictions on the evolution of fertility norms. Although the initiative was formally announced in 1979, the central government only enacted general guidance for implementing the policy. The timing of OCP implementation varied significantly across provinces, since it depended on the Communist Party's political decision process and enforcement at the local level.

To measure the timing of OCP rollout across provinces, we first collected numerous government documents, the *Population Chronicles*, and the *Encyclopedia of Chinese Family Planning*. From these, we compiled a timeline for issuing key legal documents—approvals, notices, regulations, and provisions—pertaining to the OCP for each province. The final step is identifying the date each provincial government officially mandated one child per family and imposed heavy penalties on families with excess second-order births. These official issuance dates, along with comprehensive timelines, are presented in Appendix B Table B1, which reports the start of the OCP in each province. This table shows significant variation in OCP timing across provinces.[13]

We then validate our measure. Appendix B Fig. B1 shows that TFR declined in a province right after our identified OCP implementation year in the event study framework.[14]

We investigate the determinants of implementation timing at provincial level. We regress the timing against a set of socioeconomic and demographic variables measured in the year prior to the implementation, along with the geographic distance between the provincial capital and Beijing. This distance variable serves as a proxy for political proximity to the

[13]See Appendix Section B.2 for details on the comparisons with alternative measures of the OCP.
[14]See Appendix Section B.3 for details on the event-study estimation.

central government. Appendix B Table **??** reveals that estimates of socioeconomic and demographic variables are minimal and statistically insignificant. However, as expected, the distance from Beijing shows a positive and statistically significant correlation, which suggests that provinces farther from Beijing tend to implement the OCP later. These findings affirm that the timing of the rollout was determined by political decisions of the Communist Party at provincial level.[15] We carefully control for distance from Beijing interacted with birth cohort dummies in our empirical analysis, as discussed in Section 6.2.

3.3 Abolition of the OCP and the Short-lived Fertility Rebound

The central government began to ease the OCP in 2014 with the "Selective Two-child Policy," which allowed couples to have a second child if either spouse was an only child. This policy had limited impact, however, because the targeted population did not seem to respond to the policy. On December 27, 2015, a proposal to amend the Law of Population and Family Planning was submitted to the National People's Congress, and by January 1, 2016, the universal two-child policy was unexpectedly enacted, which effectively ended the OCP (He *et al.*, 2023).

This abrupt policy shift in 2016 led to a temporary increase in fertility rates, predominantly due to a rise in second-child births, which particularly benefited the 1970s cohort. Birth rates per 1,000 women aged 15-49 climbed from 25.9 in 2015 to 44.2 in 2017, then gradually fell back to 28.9 by 2021 (Fig. 4a). This fertility spike was driven by second-child births, which surged from 8.9 per 1,000 women in 2015 to 22.1 in 2017, before decreasing to

[15] See Appendix Section B.4 for details on the regression analysis.

12.1 in 2021. Meanwhile, the rate of first-child births remained stable, hovering around 16 during 2015-2021.[16]

In 2021, China lifted all remaining fertility restrictions and began encouraging families to have up to three children. Despite these measures, the TFR continued to decline, falling from 1.16 in 2021 to 1.06 in 2022.[17] This indicates that the post-2016 fertility rebound was transient. In response to ongoing low fertility rates, the Chinese government is now formulating and enacting a range of pro-natalist policies, though the effectiveness of these policies remains to be seen.

4 Data Sources

This section outlines our data sources. We draw on the most extensive datasets that contain information on fertility in China: the 1982, 2010 and 2020 population censuses, 2005 and 2015 mini-censuses, China Family Panel Studies (CFPS) from 2012 to 2018, and 1987 Fertility In-Depth Survey (FIDS). The 2020 census has particular significance because it marks the first census conducted after the abolition of the OCP, and thus allows for the analysis of fertility behavior in the post-policy era. Together, these datasets provide a unique opportunity to examine both desired and actual fertility across multiple generations, before and after the OCP's termination.

[16]Data source: *China Labor and Population Statistical Yearbook*, published by the National Bureau of Statistics: 2007-2021.
[17]Data source: Qiao (2023).

4.1 Population Censuses

The National Bureau of Statistics (NBS) of China has conducted a population census every 10 years and a 1% population survey (mini-census) midway between each census since 1990. Both the census and mini-census provide nationally representative data that contain rich demographic and socioeconomic information.[18] The census includes age, gender, birthplace, ethnicity, fertility (for women aged 15 to 64), education, and labor market outcomes. The mini-census contains more detailed information; for example, the 2005 mini-census includes data on the number of siblings for each respondent.[19]

Our main dataset draws on the 2020 census. It offers up-to-date, nationally representative data on the actual fertility of women aged 15 to 64 following the OCP's termination. We analyze a 1‰ random sample from this census, accessed through the Peking University Development Research Center portal under NBS authorization.[20] We also use a 1% random sample from the 1982 census, a 20% random sample from the 2005 mini-census, a 0.35% random sample from the 2010 census, and a 20% random sample from the 2015 mini-census for robustness and mechanism analyses. Table 1 reports summary statistics for samples constructed using the census data.

4.2 The Chinese Family Panel Studies

The CFPS is a biennial survey that captures China's societal, economic, educational, and health conditions (Xie, 2012). The baseline was launched in 2010 and includes 33,600 adults

[18] See Appendix Section B.5 for details on the sampling design of the censuses.

[19] In the Chinese census, questions related to the number of children are asked only to women, as is the case in many countries.

[20] Data analysis was conducted in a secure lab at the NBS-Peking University Development Research Center.

and 8,990 children from 14,960 households across 25 provinces. Its probability-proportional-to-size sampling ensures the survey's national representativeness.[21] The CFPS conducted follow-ups in 2012, 2014, 2016, 2018, and 2020. A distinctive feature of the CFPS is its collection of data on desired fertility or fertility intention. In the 2012 and 2018 waves, children aged 10-15 were asked how many children they would like to have in a hypothetical scenario without any birth-control restrictions.[22] This question was also posed to adults aged 16 and above in the 2014 and 2018 waves.[23] We use data from these three waves to examine the impact of the OCP on desired fertility for the second and third generations.[24]

4.3 China Fertility In-Depth Survey

The 1987 FIDS complements the CFPS in our analysis by offering insights into desired fertility for the first generation whose fertility was directly constrained by the OCP, some of whom had died prior to the CFPS baseline survey in 2010. The NBS conducted the FIDS in April 1987 across several provinces. The 1987 FIDS gathered detailed data on female respondents' characteristics, marriage, fertility, breastfeeding history, contraception knowledge and usage, and basic information on their husbands. Notably, it included a question on desired fertility in a hypothetical scenario without any birth-control restrictions—similar to the CFPS—that targeted all adults.[25]

[21]See Appendix Section B.5 for details on the sampling design of the CFPS.
[22]Question KA202 in the individual (youth) module.
[23]Question M501 in the individual (adult) module. This question was also designated the inquiry regarding the ideal number of children in the questionnaire. In 2018, adults were directly asked how many children they would like to have.
[24]The sample, variables, and summary statistics for analysis based on the CFPS are detailed in Appendix Section B.6.
[25]The sample, variables, and summary statistics for analysis based on the 1987 FIDS are detailed in Appendix Section B.7.

5 Empirical Design

This section presents the empirical design for identifying the intergenerational impact of birth-control policies. We begin by constructing the analytic sample from the 2020 census, focusing on the second-generation women most affected by the policy's rollout. Next, we introduce our primary econometric specification: a cohort event-study design that exploits the staggered rollout of the OCP. Finally, we discuss the validity of our identification strategy, including tests for anticipatory effects, parallel trends, and potential confounding factors.

5.1 Defining Generations and Constructing the Analytical Sample

Our empirical analysis tracks three generations of women. Fig. 5a situates the three generations along a timeline that marks the implementation and abolition of the OCP. We define women born between 1969 and 1980 as the second generation ($t = 2$), their mothers as the first generation ($t = 1$), and their children as the third ($t = 3$), to match the structure of the theoretical model in Fig. 2. We chose the 1969-1980 range for two reasons. First, based on the fertility history between 2015 and 2020 by province and birth cohort, our statistical analysis shows that there is a structural break in cohort 1969: Women born before 1969 seldom experienced childbirth after the 2016 abolition.[26] Second, women born after 1980 saw minimal variation in their mothers' OCP exposure across provinces. Restricting the sample to 1969–1980 cohorts maximizes quasi-experimental variation.[27] We conduct sensitivity analysis of alternative cohort ranges in Appendix Section C.1.

[26]Appendix C Fig. C1 shows that births in 2016-2020 are nearly 0 for cohorts born before 1969, but they surge for later cohorts born after 1969. We further conduct a formal test on structural breaks following Bailey *et al.* (2021), and details are presented in Appendix Section C.1.

[27]After the OCP took effect, China's birth sex ratio became significantly male-biased, revealing widespread prenatal sex selection (Ebenstein, 2010; Li *et al.*, 2011; Chen *et al.*, 2013). Our sample therefore consists of second-generation women who survived that selection process; consequently, the estimated intergenerational effects pertain to families with comparatively weaker son preference.

The estimation sample includes urban Han Chinese women, the group for whom the one-child rule was strictly enforced; as discussed in Section 3, ethnic minorities and rural residents were largely exempt from the one-child rule. The resulting 2020 census sample of the second generation comprises 75,438 urban Han women born in 1969–1980. As reported in Table 1 Panel A, their average completed fertility by 2020 is 1.46 children.[28]

5.2 The Cohort Event-study Specification

We aim to estimate the effect of a first-generation woman's exposure to the OCP on a second-generation woman's fertility, which we observe in the 2020 census. We create a set of ordinal indicators that proxy for the expected policy exposure for first-generation women. Let T_p denote the year the OCP was introduced in province p. For a second-generation woman born in province p and year (cohort) c, her age was l $(= T_p - c)$ when the policy began. For each integer l, we define

$$D^l_{cp} = \mathbb{I}[l = T_p - c],$$

where D^l_{cp} is equal to one if the second-generation woman was exactly l years old when the policy was introduced in her province of birth. The older a second-generation woman was at OCP implementation, the longer the window her mother had to bear additional children before the one-child limit became binding. Thus, larger values of l correspond to lower expected policy exposure for first-generation women. We further explain this exposure measure in Section 5.3 below.

[28]See Appendix Section C.2 for further details on the parity distribution across generations. Appendix C Fig. C3 plots the shares of first-, second-, and third-or-higher-born women. Appendix C Fig. C4 depicts the parity distribution among second-generation women in the 2020 census.

We exploit the staggered rollout of the OCP in an event-study design. Fig. 5b depicts provincial rollout years against second-generation birth cohorts. Heilongjiang and Jilin adopted the policy first, in 1979; most other provinces followed by 1982. Only four provinces—Yunnan, Guangdong, Hainan, and Xinjiang—waited until after 1984, with Xinjiang not adopting until 1992.[29]

Our baseline event-study specification is

$$Y_{icp} = \alpha_p + \theta_c + \delta_p \cdot c + \sum_{l=8, l \neq 5}^{-1} \beta_l \cdot D_{cp}^{l} \cdot NotLastOCP_p + \epsilon_{icp}, \tag{2}$$

where Y_{icp} measures fertility in the 2020 census for a woman i in the second generation; α_p, θ_c, and $\delta_p \cdot c$ capture, respectively, province fixed effects, cohort fixed effects, and province-specific cohort linear trends; and ϵ_{icp} is the error term. Standard errors are clustered at province level in the estimation of Eq. 2.

We exclude the indicator for $l = 5$ because the 1971 "Later, Longer, Fewer" campaign encouraged a 4-year interval between births, and thus women who were 5 years old when the OCP started was enacted serve as a natural benchmark.[30] Accordingly, all coefficients β_l are interpreted relative to this $l = 5$ reference group.[31] The term $NotLastOCP_p$ is equal to one for provinces that adopted the OCP on or before 1982 and zero for the four "last-treated" provinces that adopted the policy after 1984. Following Bailey *et al.* (2021) and Goodman-Bacon (2021), we interact $NotLastOCP_p$ with D_{cp}^{l}. This interaction includes

[29] Xinjiang's adoption year 1992 lies beyond the axis range of Fig. 5b and is therefore omitted.

[30] The "Later, Longer, Fewer" (Wan Xi Shao) campaign, launched by China's National Family Planning Committee in 1973, promoted delayed marriage, increased spacing between pregnancies—specifically encouraging intervals of 4 years—and having fewer children.

[31] For a detailed justification for omitting age 5, please refer to Appendix Section C.3. We also bin $l \geq 8$ as a single indicator D_{cp}^{8} to prevent multicollinearity concerns.

the four last-treated provinces in the control group, and thus improves the efficiency of our event-study estimators.[32]

To address cross-provincial heterogeneities in life-cycle fecundity patterns, we include province-specific cohort trends. This is crucial, because although the OCP was uniformly terminated in 2016 across all provinces, women of varying ages, whose mothers experienced different degrees of OCP exposure, may have exhibited distinct fertility behaviors between 2016 and 2020. Whereas cohort fixed effects capture common life-cycle fecundity patterns, our addition of province-specific linear trends accounts for any heterogeneities across provinces.[33]

5.3 Measuring Policy Exposure of the First Generation

We use a set of indicators for second-generation ages at OCP onset, D^l_{cp}, to proxy for their mothers' expected exposure to the OCP.[34] The younger a second-generation woman was when the OCP began, the less pre-policy time her mother had to bear additional children, and thus the stronger the mother's expected policy exposure. Recent literature often constructs a continuous exposure index directly from the mother's age at OCP launch (e.g., Chen and Huang, 2020; Chen and Fang, 2021; Rossi and Xiao, 2024). However, implementing that approach for intergenerational analysis requires matching mothers to adult daughters in the 2020 census, which is infeasible because grown daughters rarely coreside with their mothers. Our measure based on second-generation age therefore offers the only practicable proxy.

[32] Appendix Section C.4 shows that our estimation is similar but less precise without interacting D^l_{cp} with $NotLastOCP_p$.

[33] Appendix Section C.5 presents more discussion and robustness analyses of the specification of the province-specific cohort linear trend.

[34] See Appendix Section C.6 for further details on an alternative continuous approach to proxy for expected exposure to the OCP of the second-generation women's mothers.

Conceptually, the two measures are very similar. Consider second-generation women born in 1976 and 1977 in provinces that adopted the OCP in 1981 (the 9 provinces ranging from Anhui to Tianjin in Fig. 5b). At the policy start, the 1977 cohort was 4 years old ($l = 4$, "treated"), whereas the 1976 cohort was 5 ($l = 5$, reference). Linking mothers and daughters in the 1982 census—when the girls still lived with their mothers—shows virtually identical maternal ages at birth: an average of 26.380 for the 1976 cohort versus 26.378 for the 1977 cohort.[35] A Kolmogorov–Smirnov test cannot reject the equality of the two distributions ($p = 0.362$; Appendix C Fig. C2a).

A first-generation mother's age in 1981 equals her daughter's age plus her age at birth. Because first-generation age at birth is identically distributed across the two second-generation birth cohorts (Appendix C Fig. C2a), the distribution of first-generation age in 1981 for 1977-born daughters is merely a 1-year leftward shift relative to that for the 1976-born daughters (Appendix C Fig. C2b). Hence D^4_{cp} corresponds to first-generation women who were, on average, 30.378 in 1981, while the reference group ($l = 5$) corresponds to first-generation women who were 31.380. This is precisely the comparison that a measure based on first-generation age would make.

Neither approach compares fully exposed mothers directly with those completely unexposed, and thus prevents a direct estimation of the overall policy effect on second-generation fertility. Nevertheless, by first estimating Eq. 2 with second-generation fertility as the dependent variable (reduced-form estimation), and subsequently re-estimating Eq. 2 with second-generation sibling size as the dependent variable (first-stage estimation), we can

[35]The census lists every household member. We match each mother to her coresident children and keep only families in which all children remain at home, which enables us to assign exact birth order (Li *et al.*, 2011; Guo *et al.*, 2025). Nearly all second-generation women—born in the 1970s—satisfy this requirement in the 1982 census.

construct a Wald estimator using the corresponding estimated β_l's. This Wald estimator identifies the impact of the first-generation fertility reduction induced by the OCP on second-generation fertility.

5.4 Identification

This subsection uses an illustrative example to clarify the sources of identification and discuss the identifying assumptions behind the two steps in our empirical analysis. In the first step, we identify the effect of first-generation exposure to the OCP on the fertility behavior of the second generation, and thereby document the policy's intergenerational impact. In the second step, we explore whether this intergenerational impact operates through the evolution of fertility norms, the theoretical mechanism we proposed in Section 2.

Sources of identification. Return to the 9 provinces that adopted the OCP in 1981. The raw fertility difference between 1977- and 1976-born women for this group contains two components: (a) a treatment effect at $l = 4$ and (b) any secular change in fertility between the two cohorts. Even with identical maternal ages at birth, the two cohorts can differ in other maternal characteristics—education, for example—that affect fertility and contribute to component (b).

Our event-study specification removes component (b) by subtracting the same cohort difference in a suitable control group, e.g., the fertility gap between 1977- and 1976-born women in provinces that adopted the OCP in 1982 (10 provinces ranging from Fujian to Zhejiang in Fig. 5b). This 2×2 difference-in-differences is one of many such contrasts that Eq. 2 aggregates to estimate β_4 (Sun and Abraham, 2021). We now discuss the identifying assumptions underlying in two steps.

Assumptions for identifying the intergenerational impact of the OCP. In the first step, we are interested in the estimates of $\beta_l' s$ ($l < 5$) in Eq. 2, which capture the causal effects of first-generation exposure to the OCP on the fertility of second-generation women. We leverage variations in both the birth years of second-generation women and the timing of OCP implementation across provinces. These two factors jointly influence the likelihood of these women having younger siblings, and thereby capture the fertility constraint imposed by the OCP on their mothers.

Consistent estimation of β_l in our event-study framework requires two key assumptions, as outlined by Sun and Abraham (2021) and Miller (2023). The first is the absence of anticipation effects, which means that mothers should not have altered their fertility prior to the OCP's implementation in anticipation of the policy. The second assumption is parallel trends in potential fertility outcomes. This assumes that in the absence of the OCP, fertility changes across birth cohorts should be the same between provinces that implemented the policy early and those that implemented it later. The second assumption is subtle in our study context, because we estimate the effect of OCP exposure of the mothers' generations on the fertility of the daughters' generation. This requires that the timing of OCP implementation does not correlate with factors that vary by province and birth cohort and that could influence fertility rates across both generations, conditional on covariates. These two assumptions will be thoroughly examined in Section 6.2.

A particularly important concern in our context is the intergenerational correlation in OCP exposure. In particular, provinces that implemented the OCP earlier may tend to enforce the policy more strictly in subsequent years. This could trigger two challenges. First, if the timing of the OCP rollout correlates with subsequent changes over time in OCP enforcement within the same province, the common trend hypothesis is violated. Second, if the timing of the OCP rollout correlates with cross-province variations in the degree of OCP enforcement, the treatment effect of OCP rollout on second-generation fertility could be heterogeneous across provinces, and thus introduce a misspecification in Eq. 2. In Section 6.2, we directly examine intergenerational correlation in OCP exposure, conduct a placebo test using second-generation fertility before OCP relaxation, and implement heterogeneity-robust estimators in the recent event-study literature (Sun and Abraham, 2021; De Chaisemartin and D'Haultfœuille, 2023; De Chaisemartin and D'Haultfoeuille, 2024).

To further investigate the parallel trends assumption, we conduct balance tests using first-generation women's predetermined characteristics from the 1982 census, when nearly all second-generation women still lived with their mothers, which enables us to match mothers to daughters. Our identification strategy requires that these predetermined characteristics remain balanced across the OCP rollout schedule. For instance, if provinces adopting the OCP earlier also experienced faster increases in first-generation women's education—which could subsequently affect both their own and their daughters' fertility—our estimates would be biased. To address this concern, we re-estimate Eq. 2, using first-generation characteristics (education, occupational status, and age at birth) as dependent variables, and find coefficients statistically indistinguishable from zero (Appendix C Fig. C6).[36] This result supports the validity of our identifying assumption.

Exploring the role of norms. In the second step, we examine the mechanism that underlies our identified intergenerational impact of the policy. Our theoretical framework

[36]See Appendix Section C.7 for further details on the empirical design.

proposes the evolution of fertility norms as the primary driving force, although conceptually, several alternative mechanisms could also be at play. The OCP directly influenced fertility among the first generation, and subsequently affected both first-generation post-policy characteristics and second-generation characteristics. Consequently, some characteristics might be inherently unbalanced between treatment and control groups. For example, second-generation women in the treatment group have fewer siblings and lower average birth orders, which implies that their mothers—first-generation women—have lower parity on average.[37] Also, through the child quantity-quality (QQ) trade-off (Becker and Lewis, 1973), second-generation women in the treatment group might have better education due to fewer siblings. These differential characteristics could lead to distinct fertility patterns between second-generation women in the treatment and control groups. However, such differences do not undermine our identification of the policy's intergenerational impact; rather, they influence our interpretation of *how* first-generation exposure to the OCP affects the fertility of second-generation women.

We employ two strategies to investigate the mechanism that underlies the intergenerational effect in Section 7 below. First, we directly assess the policy's intergenerational impact on fertility norms—proxied by desired fertility—among the second and third generations. Second, we comprehensively evaluate alternative mechanisms by examining second-generation women's education, labor supply, and living arrangements. We also explore the role of household bargaining in shaping norm evolution.

[37] In other words, within our sample of second-generation women, mothers (first-generation women) with higher parity are disproportionately represented in the control group relative to the treatment group.

6 Empirical Results

This section reports our main empirical findings. We first present the baseline event-study estimates, demonstrating that maternal exposure to the OCP significantly reduces fertility across generations. We then assess the validity of our research design through a battery of robustness checks, including tests for anticipatory effects, parallel trends, and confounding policy shocks. Finally, we examine the sensitivity of our results to early-life policy exposure, the limited observation window following the policy's abolition, and heterogeneous treatment effects.

6.1 Main Results

Before presenting our main results for the fertility of women in the second generation, we first estimate the OCP's impact on their mothers' fertility. This impact corresponds to the effect of birth-control policies on the first generation in our model. Using urban Han Chinese women born in 1975–1980 from the 2005 mini-census, we estimate Eq. 2 with the number of siblings as the dependent variable to proxy for first-generation women's fertility.[38] Fig. 6a reports that estimates of β_l are near zero for $l > 5$, which indicates no significant pre-trend. The β_l estimate decreases from -0.07 for $l = 4$ to -0.33 for $l = -1$, consistent with the

[38] Table 1 Panel B provides summary statistics for this sample. We include women born after 1974, since the 2005 mini-census requires only those aged 30 or below to report their sibling size. The 2005 mini-census is the only census wave that requires individuals to report their sibling size.

literature.[39] This suggests that the OCP significantly reduced fertility in the first generation. Table 2 Column (1) reports the estimates and corresponding standard errors.

We now present our main results. Based on our main analytic sample of urban Han Chinese women born during 1969-1980 in the 2020 census, Fig. 6b plots our estimates of β_l. The figure shows no pre-trend, since the estimates of β_l are nearly zero when $l > 5$. The β_l estimate decreases from -0.05 for $l = 4$ to -0.22 for $l = -1$, which are both significant at the 10% level. Women whose mothers' generation were exposed to the OCP had between 0.05 and 0.22 fewer children compared with those whose mothers were not exposed. Table 2 Column (2) reports the result, which is consistent with our prediction by Proposition 2. The estimated effect is substantial, especially considering a TFR of 1.06 in 2022.

By combining the findings from Figs. 6a and 6b, we calculate the intergenerational persistence in fertility, as implied by Proposition 1. Within a local average treatment effect framework (Imbens and Angrist, 1994), the estimates from Fig. 6a serve as first-stage results that show the OCP's impact on first-generation women's fertility. Fig. 6b provides reduced-form estimates of the OCP's influence on second-generation women's fertility. The ratio of these estimates offers a causal interpretation of how changes in first-generation women's fertility affect the fertility behaviors of second-generation women (Angrist and Krueger, 1992). Consequently, we infer that for a long-term horizon ($l = -1$), a one-unit decrease in first-generation women's fertility reduces second-generation women's fertility by 0.67 (0.22/0.33).

[39]The magnitude of β_l are consistent with the literature. Ahn (1994), the first study examining the immediate fertility impact of the OCP, finds that the likelihood of having a second child declines by 19 to 44 percentage points in urban areas of Hebei, Shaanxi, and Shanghai. McElroy and Yang (2000) employ survey data with broader geographic coverage and suggest that the OCP reduces rural fertility by 0.33 children. Liu (2014) employs alternative survey data and finds that exemptions to the OCP led mothers to have 0.404 more children. Using data from the representative 2000 China population census, Huang (2022) assesses the OCP's fertility effects and finds that it reduces maternal fertility by 0.29 children. A recent study by García (2022) similarly suggests that the OCP reduces the total fertility rate by 0.3 children.

6.2 Robustness

Anticipatory effects. Figs. 6a and 6b show that the estimates of β_l for $l > 5$ are small and statistically insignificant, which suggests no anticipatory effects on fertility in both generations. This result is confirmed by Appendix B Fig. B1, which demonstrates no pre-trend in the TFR preceding the year of OCP implementation. We further assess anticipatory effects by analyzing birth spacing among first-generation women in provinces that implemented the OCP after 1981. These first-generation women are grouped based on their daughters' birth dates: Group 1 (pre-1975 births) and Group 2 (1975-1978 births). Mothers in Group 1 were not influenced by the policy. We aimed to discern whether first-generation women in Group 2 adjusted their birth intervals in anticipation of the policy change. Upon comparing birth-spacing distributions in Appendix C Fig. C7, we observe similar patterns for first-generation women in both groups, which indicates no anticipatory behavior.

Common trends for potential outcomes. This is the most important identifying assumption for our event-study design. We comprehensively investigate the robustness of our results to this assumption by conducting two sets of analyses.

First, we control for four socioeconomic and policy shocks in our regression: the Great Famine (1959-1961), the Cultural Revolution (1966-1976), the "Later, Longer, Fewer" Campaign during the 1970s, and implementation of the Compulsory Education Law (around 1986).[40] These shocks varied in intensity across provinces, and exposure to these events also differed among the cohorts we are studying. Our estimate might capture the effects of these shocks if they correlate with both fertility and our measured OCP exposure simultaneously.[41]

[40]We do not consider the rural household responsibility system reform in the late 1970s, because we focus on urban women. Our construction of measurements for different policy shocks is detailed in Appendix Section C.8.

[41]Appendix C Fig. C10 shows the estimated results.

Second, we control for interactions between the cohort dummies and a series of variables that measure pre-OCP socioeconomic and political conditions in the regression. Our event-study framework has well accounted for any cohort-invariant and linearly evolved cross-province heterogeneity. Nonetheless, the heterogeneity in these conditions could impact not only the level and linear changes, but also nonlinear changes in fertility across cohorts of different provinces. As shown in Appendix B Table **??**, whereas OCP timing does not significantly correlate with pre-OCP socioeconomic conditions, it indeed correlates with the geographic distance between the provincial capital and Beijing. Therefore, it is necessary to control for interactions between the cohort dummies and distance from Beijing.[42]

Early-life OCP exposure of the second generation. One particular concern with our estimation is that all women in our estimation sample were subject to the OCP before 2014. If OCP exposure for women in the first generation in the 1970s and 1980s, net of province and cohort fixed effects and province-specific linear trends, correlates with the intensity of the OCP for women in the second generation in the 1990s and 2000s, our negative estimates of β_l ($l < 5$) would mechanically reflect the OCP effect in early life for women in the second generation. However, this is unlikely for two reasons. First, policy enforcement has remained stable since the mid-1990s (Section 3.1), leaving little room for cross-cohort variations in OCP exposure for second-generation women. Second, we find no reason to expect a correlation between the timing of OCP implementation for women in the first generation and changes in OCP intensity for second-generation women across birth cohorts.

To further alleviate this concern, we conduct two analyses. First, we examine the correlation between the OCP exposure of first-generation women and intensity of the OCP

[42]Appendix C Fig. C11 shows the estimated results.

experienced by second-generation women. Specifically, we re-estimate Eq. 2, using second-generation women's OCP intensity as the dependent variable. Intensity is measured by the average fine rate the second-generation woman experienced during her early life (6–20) and reproductive ages (20–45) based on her birth province and cohort.[43] Appendix C Fig. C12 shows that the estimates of β_l ($l < 5$) are statistically insignificant.[44]

Second, we examine the correlation between first-generation women's OCP exposure and second-generation women's fertility before OCP termination. We re-estimate Eq. 2 using the same birth cohorts as in our main sample, but measure their fertility prior to 2013 based on the 2015 mini-census.[45] If there is a significant correlation between first-generation women's OCP exposure and the OCP intensity experienced by second-generation women, estimates of β_l ($l < 5$) should be significantly negative. However, the results, reported in Column 3 of Table 2, do not support this prediction (Appendix C Fig. C13). This is because second-generation women's fertility was directly restricted by the policy before OCP termination, which likely masked their intended fertility levels.

A short window of childbearing after OCP abolition. A key challenge arises from the fact that the OCP was only fully repealed in 2016, which limits our observation of fertility to a 5-year window until the 2020 census. Cohorts born between 1969 and 1980 were already in their mid-30s to late 40s by 2016, and thus entering later reproductive stages. This short window raises concerns about the ability of these cohorts to reflect changes in fertility in the post-OCP era. Also, post-abolition fecundity likely varied across these cohorts. To

[43] Under the OCP, parents in China who exceeded their fertility limit were forced to pay a fine, which varied across provinces. See Appendix Section C.9 for the details of the policy.

[44] Appendix Section C.9 reports the details.

[45] Table 1 Panel C reports the average fertility of these women before policy relaxation in the 2015 mini-census, which shows an average of 1.37 children by 2013. We measure the fertility of these women in 2013, which corresponds to the period before the central government initiated relaxation of the OCP in 2014 and introduced the "Selective Two-child Policy," as outlined in Section 3.3.

mitigate these concerns, our main analysis includes province-specific trends for each cohort to address heterogeneous life-cycle fertility patterns. Despite these efforts, uncertainty persists regarding the adequacy of the 5-year observation period to capture genuine fertility behavior in the absence of birth-control policies.

We conduct two analyses to address this concern. First, analyzing national aggregate fertility data reveals a notable trend in second-child birth rates per 1,000 women in our study period, as depicted in Fig. 4b. There was a significant rise from 5.7 in 2015 to 10.2 in 2017, with a subsequent decline to almost zero by 2021. The peak in 2017, when rates hit 10.2—approximately half the average for fertile women aged 15-49, as detailed in Fig. 4a—suggests that women in our sample were actively seizing the last chance to fulfill their fertility desires.

Second, we re-estimate Eq. 2 using the number of children born after OCP termination as the dependent variable.[46] Appendix C Fig. C14 shows the estimated coefficients, which are comparable to our baseline results in Fig. 6b. The intergenerational effects we document are driven almost entirely by births that occurred after OCP relaxation.

Heterogeneous treatment effects. Even though the timing of the OCP rollout doesn't predict cross-cohort variations in OCP exposure among second-generation women in the same province, these women on average experienced greater exposure to the OCP during their early lives. If early-life OCP experience alters the responsiveness of second-generation fertility to fertility norms, Eq. 2 would be misspecified by not allowing for heterogeneous treatment effects across provinces. To address this issue, we employ the estimators proposed by Sun and Abraham (2021) and De Chaisemartin and D'Haultfoeuille (2024). If provinces

[46]See Appendix Section C.10 for further detail on the empirical analyses.

that implemented the OCP at different times exhibit distinct treatment effects given the same relative year, the coefficients obtained using these two estimators would no longer exhibit the same pattern as those estimated using Eq. 2. Appendix C Fig. C15 shows that our results remain robust.[47]

When treatment effects are heterogeneous, standard two-way fixed-effect regressions may also suffer from confounding variations arising from alternative treatments (De Chaisemartin and D'Haultfœuille, 2023). We follow De Chaisemartin and D'Haultfœuille (2023) and re-estimate our baseline results with their multi-treatment estimator. The procedure aggregates valid 2×2 difference-in-differences in which each alternative treatment is held fixed at its initial value. Appendix C Fig. C16 reports the results after separately controlling for four alternative treatments: the LLF campaign, compulsory education law, Great Famine, and Cultural Revolution. Our results remain robust, although standard errors widen because the estimator relies on smaller effective samples.

7 Mechanisms

This section explores the mechanisms through which first-generation exposure to the OCP influences second-generation women's fertility. First, we trace fertility norms across three generations, and detect a lasting impact of first-generation women's fertility behavior on fertility norms among the second and third generations. Second, we examine changes in childbearing costs as an alternative explanation. Third, we quantify the relative contributions of maternal and peer effects to the policy's enduring impact. Fourth, we investigate

[47] The observed uniformity in behavioral responses across various provinces may stem from widespread reluctance among women to limit their family to a single child during that period. This interpretation is supported by the empirical analysis presented in Section 6.2 using the 1987 FIDS.

intrahousehold bargaining, assortative mating, and mother-in-law effects. Finally, we assess whether the OCP also has lasting effects on childlessness and marital behavior.

7.1 Evolution of Norms

Our model suggests that the OCP restricts first-generation women's fertility and thus establishes a new fertility norm for future generations. To examine this norm evolution mechanism, we use women's desired fertility as a proxy for perceived fertility norm.

Norm evolution in the second and third generations. We use desired fertility in the CFPS to trace norm evolution across the second and third generations. The CFPS solicits desired fertility based on the following survey question: "What's your ideal number of children under the hypothetical scenario without any birth-control restrictions?"[48] We re-estimate Eq. 2 using second-generation desired fertility as the dependent variable. The results, depicted in Fig. 7a, show that the estimates of β_l are significantly negative, which suggests that first-generation exposure to the OCP establishes a low fertility norm for the second generation.

We then extend our analysis of desired fertility to the third generation, who are children of second-generation women. The third generation are teenagers during the survey years and will not be affected by the OCP during their entire reproductive lifespan, which offers an alternative context to test the norm evolution mechanism. We re-estimate Eq. 2 using desired fertility of the third generation as the dependent variable. Results presented in Fig. 7b show that grandmothers' OCP exposure decreases the desired fertility for their

[48]We provide detailed information on sample construction and summary statistics for the CFPS desired fertility data in Appendix Section B.6. The literature debates how well stated fertility desires predict actual fertility (Rosenzweig and Wolpin, 1993; Schoen *et al.*, 1999; Joyce *et al.*, 2002; Morgan and Rackin, 2010; Doepke and Kindermann, 2019). We use fertility desires as a proxy for perceived fertility norms, and not for realized fertility behavior itself.

grandchildren. The OCP appears to create low fertility norms that last beyond the second generation.

Behavior versus intention as the trigger of norm evolution. To explore what sets the norm evolution in motion, we examine the first generation's fertility intentions in the 1987 FIDS, which includes a question on desired fertility similar to the CFPS. The FIDS also contains the birth year of the respondents' first-born children, which enables us to calculate child age at OCP implementation.

Based on the 1987 FIDS, we re-estimate Eq. 2 using the desired fertility of first-generation women as the dependent variable. The estimates of β_l ($l > 0$) are presented in Appendix D Fig. D3, and are found to be near zero and statistically insignificant. By contrast, the policy has reduced first-generation women's actual fertility, as shown in Section 6.1 Fig. 6a. This result aligns with our theoretical prediction, which suggests that it is the actual fertility of mothers—rather than their fertility intentions—that triggers the evolution of fertility norms for succeeding generations.

7.2 Childbearing Costs

In the theoretical analysis, we derive the evolution of fertility norms as the driving force behind the lasting impact of the OCP, assuming similar changes in childbearing costs for women across cohorts and provinces. This norm influences the fertility of these subsequent generations by influencing their preferences, rather than through changes in the costs in the budget constraint. However, the OCP itself could have altered these costs, depending on the extent of first-generation exposure to the policy. If this was true, our main findings could be attributed to these OCP-induced shifts in childbearing costs, rather than to the evolution of fertility norms.

The literature suggests three mechanisms through which the OCP could change childbearing costs. The first relates to the theory of the child QQ trade-off (Becker and Lewis, 1973), which we discussed in Section 2.2. The QQ theory predicts that the OCP's role in reducing the first generation's fertility might lead to more investment per child, and consequently enhance the human capital of women in the second generation. This increase in human capital could, in turn, affect the fertility of these women. The second relates to women's labor supply: The OCP-induced decrease in fertility may increase the labor supply of women in the first generation. This may alter the labor supply of women in the second generation through the intergenerational transmission of cultural traits that affects female labor supply (Fernández and Fogli, 2006, 2009). In addition, if the OCP increases the human capital of women in the second generation through the child QQ trade-off, the OCP will also change their labor supply. The potential change in the women's labor supply may affect their fertility. Third, the literature on economics, sociology, and demography suggests that the OCP would change the living arrangements of the first and second generations (Rosenzweig and Zhang, 2014). Because fertility decreases, mothers in the first generation are more likely to coreside with their daughters (second-generation women). Since first-generation women can assist in caring for children, the cost of childbearing might decrease for women in the second generation. All three factors may influence women's fertility in the second generation through the change in childbearing costs in our theoretical model.

We empirically examine the three potential mechanisms. We re-estimate Eq. 2 using women's education, labor market outcomes, and coresidence status with parents or parents-in-law as dependent variables. To avoid potential behavioral changes stemming from relaxation of the OCP, we measure the three dependent variables using 2010 census data. Summary statistics are presented in Table 1, Panel D. The results remain robust when dependent variables are replaced by those measured in the 2015 mini-census and 2020 census. Appendix D Figs. D4-D6 show that the estimates of β_l ($l < 5$) are all small and statistically insignificant.[49] Our results regarding women's education align with prior research. Comprehensive studies examining the OCP in China have concluded that its effect on children's human capital is minimal (Rosenzweig and Zhang, 2009; Liu, 2014; Li and Zhang, 2017; Guo *et al.*, 2025). These results suggest that our main finding is less likely to be driven by the three alternative mechanisms.

It's essential to recognize that, theoretically, the fertility norm mechanism is largely unaffected by OCP-induced changes in childbearing costs, as long as these alterations are not substantial. While we have examined three alternative mechanisms influenced by changes in childbearing costs due to the OCP, potential variations in these costs are vast and often challenging to measure. Nevertheless, the assumption regarding common price changes (Assumption 1a) can be adjusted to accommodate variations in childbearing costs based on OCP treatment status.[50] The strength of the fertility norm mechanism roots in Akerlof (1997), which suggests that within a context of "conformity," a "marginal change" in the cost of childbearing would have "literally no effect" on fertility behavior, and thus highlights the norm mechanism's robustness.

[49] Appendix D Table D1 reports estimates and corresponding standard errors.

[50] In our model, the upper fertility bound for second-generation women is given by $u_2^u = \frac{\gamma+\sigma}{1+\gamma+\sigma}\frac{y_2}{p_2}$. The OCP may lead to an increase in p_2 and a corresponding decrease in n_2^u. However, if the OCP-induced rise in p_2 is modest, such that $n_2^u \geq n_1$, the fertility of second-generation women will mirror that of the first generation ($n_2^* = n_1$), indicating that the fertility norm mechanism continues to operate dominantly.

7.3 Maternal vs. Peer Effects

So far, we have assumed that the OCP had a uniform effect on the fertility of all mothers for daughters born in the same province and year. Yet this assumption might not always hold. For example, imagine two women born in 1978 in a province where the OCP was implemented in 1979. If one woman's mother was past childbearing age by 1978, while the other's was not, the OCP's fertility restrictions would only apply to the latter. This variation implies that the OCP's impact on mothers' fertility—and consequently on the daughters' perceived fertility norms—can differ even within the same province and cohort. Under such circumstances, peer influences might also shape fertility norms, which means that our estimates of β_l ($l > 0$) capture both maternal and peer effects, as Proposition 3 suggests.[51] In this part, we aim to distinguish quantitatively between maternal and peer influences on the lasting effect of the OCP.

A method to separate maternal and peer effects. For simplicity, we assume that the proportion of mothers affected by the OCP is λ for a given relative year l across all provinces, which is not observed by researchers. Our estimate of β_l is thus represented as

$$\beta_l = \lambda(n^*_{r,2} - n^u_2) + (1 - \lambda)(n^*_{ur,2} - n^u_2), \tag{3}$$

where n^u_2 denotes the hypothetical fertility of urban Han Chinese women in the second generation in the absence of the OCP, and $n^*_{r,2}$ and $n^*_{ur,2}$ represent the realized fertility of women whose mothers were and were not, respectively, affected by the OCP.

[51]It is important to note that the maternal effect is a prerequisite. By Proposition 3, mothers' OCP exposure would not impact daughters' fertility in the absence of the maternal effect, regardless of the peer effect's presence.

Our objective is to establish a lower bound (in absolute terms) for the impact of a mother's OCP exposure on her daughter's fertility through the maternal effect *alone*, denoted by $n_2^l - n_2^u$, where n_2^l represents the hypothetical fertility of Han Chinese women whose mothers were restricted by the policy but in the absence of the peer effect, as suggested by Proposition 2:

$$\begin{aligned} n_2^l - n_2^u &\leq n_{r,2}^* - n_2^u \\ &\leq \lambda(n_{r,2}^* - n_2^u) \\ &= \beta_l - (1-\lambda)(n_{ur,2}^* - n_2^u) \\ &\leq \beta_l - (n_{ur,2}^* - n_2^u), \end{aligned} \tag{4}$$

where the first inequality is due to $n_2^l \leq n_{r,2}^*$, as implied by Proposition 3; the second and third inequalities follow from $0 < \lambda < 1$; and the equality is based on Eq. 3. Therefore, $\beta_l - (n_{ur,2}^* - n_2^u)$ serves as a lower bound for $n_2^l - n_2^h$ in absolute value, since both terms are negative.

We have estimated β_l but not $n_{ur,2}^* - n_2^u$, which represents the effect that arises from the diffusion of low fertility norms within the second generation *alone*. Whereas extensive literature focuses on the diffusion of low fertility norms in peer groups (Spolaore and Wacziarg, 2022; Rossi and Xiao, 2024), estimating its impact on urban Han Chinese women is infeasible. This is due to the limitations of census data, which, for a given birth cohort within the same province for urban Han Chinese women, do not differentiate between women whose mothers were and were not treated by the OCP.

Analysis based on minorities. In the following analysis, we estimate the peer effect, $n^*_{ur,2} - n^h_2$, using minority women. This analysis presupposes that, were it not for the OCP, the fertility preferences of Han and ethnic minorities would be similar. While the credibility of this assumption remains an open question, it is a premise frequently adopted in prior research (Li and Zhang, 2009; Rossi and Xiao, 2024; Li and Zhang, 2007; Wang and Zhang, 2018; Zhang, 2017).[52] Accepting this assumption enables us to estimate a lower bound for the maternal effect's contribution to the lasting effect of the OCP. Specifically, we re-estimate Eq. 2, this time focusing on urban minority women born from 1969 to 1980 in the 2020 census. Their fertility is used as the dependent variable. For these women, we assign the same policy exposure values to their mothers as we did for the mothers of urban Han women born in corresponding provinces and years.[53]

The results in Appendix E Table E2 inform the peer effect. Minority women, whose Han counterparts' mothers were impacted by the OCP, have lower fertility than those whose peers' mothers were not under the OCP. The β_1 estimate reveals that in 2020, fertility among minority women, who were aged 1 at OCP implementation and born in provinces that implemented the policy earlier was 0.066 lower than that of women born in the same year in provinces that adopted the policy later.[54] Our results align with those of Rossi and Xiao (2024).

We apply Eq. 4 to compute the lower bound of the maternal effect using estimates from the ethnic minority sample. For example, considering the age at OCP 1 ($l = 1$), the absolute

[52]We also conduct a balance test between urban Han women and urban minority women in our sample and find no evidence of violations. For more details, see Appendix Section E.1.1.
[53]For more details on the empirical analysis using the ethnic minority sample, who are culturally similar to the Han Chinese, see Appendix Section E.1.2.
[54]Due to the small sample size of urban minority women in the 2020 census, the coefficients of β's are not estimated with precision; this limitation also hinders our investigation of heterogeneous effects based on the share of Han Chinese across regions, as implied by Proposition 3.

lower bound for $n_2^l - n_2^h$ is 0.02 ($= 0.08 - 0.06$), which constitutes 20% of the total effect (β_1). It is important to note that the estimated coefficient of 0.06 may overestimate the peer effect within the second generation.[55] This coefficient encompasses two elements: the direct peer effect from Han Chinese to minorities within the second generation, and an indirect effect that combines the peer effect from Han Chinese to ethnic minorities in the first generation and the maternal effect from minority mothers to their daughters. The latter component includes a maternal effect. Therefore, the estimate of 0.06 based on minority women may overestimate the peer effect.[56]

Analysis based on rural-to-urban migrants. We next estimate the peer effect, $n^*_{ur,2} - n_2^h$, using rural-to-urban migrants.[57] We re-estimate Eq. 2, this time focusing on rural-to-urban migrant women born from 1969 to 1980 in the 2020 census. Their fertility is used as the dependent variable. For these women, we assign the same policy exposure values to their mothers as we did for the mothers of urban Han women born in corresponding provinces and years. Column (2) of Appendix E Table E2 presents the estimated results. We find that rural-to-urban migrant women whose urban peers' mothers experienced greater exposure to the OCP exhibit significantly lower fertility than those whose peers' mothers were subject to less intense exposure. The size of the peer effect estimate using rural-to-urban migrants (column 2) closely mirrors that using urban minorities (column 1).

[55]This overestimation may not be of great concern, since it could result in an underestimation of the component of maternal influence in the overall effect.

[56]The above analysis quantifies a lower bound on the contribution of the maternal effect to the OCP effect on daughters' fertility. Interestingly, we are also able to establish a lower bound for a structural parameter: the strength of the maternal effect relative to the peer effect, α, in the woman's utility function (Eq. 1). See Appendix Section E.1.3 for more details.

[57]For more details on the empirical analysis using the rural-to-urban migrants, see Appendix Section E.1.4. We also conduct a balance test between urban non-migrant and rural-to-urban migrant women in our sample and find no evidence of violations. For more details, see Appendix Section E.1.1.

7.4 Intrahousehold Bargaining

Up to this point, we have modeled the household as a unitary decision maker. We now treat fertility as a collective outcome of intrahousehold bargaining between spouses (Chiappori, 1992; Lundberg and Pollak, 1993; Doepke and Kindermann, 2019). As Fong (2002) notes, in urban China, daughters born under the OCP seldom have brothers—a shift that reduces son preference and strengthens daughters' status. In general, the OCP should have increased the bargaining power of second-generation women relative to their husbands. Consequently, the small-family ideal embraced by these women would carry greater weight in household negotiations, which reinforces the persistency of this norm across successive generations of women.

In particular, the OCP reduced family size and enabled first-generation households to accumulate additional wealth that could be transferred to their daughters (second-generation women). In Chinese society, this transfer typically occurs at marriage in the form of a dowry, and larger dowries strengthen a bride's bargaining position within the household (Zhang and Chan, 1999). The CFPS reports the value of dowry payment, which allows us to empirically examine this bargaining-power channel.[58]

We re-estimate Eq. 2 using three dependent variables: (a) an indicator for whether the woman received a dowry; (b) the value of the dowry payment; and (c) an indicator for whether she was listed as the household head, a proxy for her bargaining power within the household. Appendix E Fig. E1 shows that first-generation exposure to the OCP increases both the likelihood of receiving a dowry and its value, and that larger dowries are in turn associated with a higher probability of the woman being recorded as household head. The

[58]Appendix Section E.2 details the derivation of dowry data in the CFPS. Appendix E Table E5 shows the summary statistics for these variables.

results underscore a channel of female empowerment that bolsters our proposed mechanism of norm evolution across generations of females.

7.5 Mother-in-law Effects

Because fertility is typically a joint decision, the intergenerational effect could begin even earlier, at the stage of spouse selection. Women who favor small families may seek to marry men with similarly modest fertility goals. A bride's childbearing behavior could also be shaped by her mother-in-law, and at times more strongly than by her own mother (Fernández *et al.*, 2004; Fernández and Fogli, 2009; Gay, 2023). We therefore examine assortative mating and the influence from the mother-in-law.

First, we examine assortative mating. Using matched couples in the 2010 census, we construct seven husband-side variables: four education dummies (completed primary school, junior high school, senior high school, or college) and three labor-market measures (labor-force participation, employment last week, and hours worked last week). We also construct the variable of desired fertility for husbands from the CFPS.[59] Re-estimating Eq. 2 with each of these variables reveals no statistically significantly correlation between a woman's maternal OCP exposure and her husband's education, labor market behaviors, or fertility desires (Appendix E Fig. E2). Observable assortative mating therefore seems limited.

Second, we explore the influence of the mother-in-law. If a mother-in-law experienced the OCP, her reduced fertility could reinforce the daughter-in-law's perceived low fertility norm. We follow Gay (2023) and use husbands' birthplace and birth year to proxy for the mother-in-law's OCP exposure. Specifically, we construct the same set of ordinal exposure

[59] Appendix E Table E3 shows the summary statistics of the second-generation husbands.

indicators for the mother-in-law that we use for the wife's own mother. The analysis relies on matched couples in the 2020 census, when second-generation fertility is observed after the policy's abolition. Appendix E Table E4 shows a minimal influence of the mother-in-law. A likely reason is the customary spousal age gap: In our sample, second-generation women marry men who are on average 2.5 years older. Consequently, 86% of women in our sample wed men born at least 5 years before the OCP rollout, leaving most mothers-in-law essentially unexposed to the policy—and therefore with little scope to transmit an OCP-driven low-fertility norm.

These findings should be interpreted cautiously. Our measures capture only .observable attributes of the husband and the mother-in-law, and leave many unobserved characteristics—and their potential influence—outside the analysis.

7.6 Childlessness and Marriage

After the OCP entrenched a low fertility norm, childlessness may also have become more socially acceptable.[60] Consistent with this notion, Guo *et al.* (2024a) report that lifetime childlessness in China rose from 1.1% for women born in the 1940s to more than 4% for those born in the 1970s—a nearly fourfold increase to which the new norm might have contributed, although it is unlikely to be the sole explanation.

We re-estimate Eq. 2 using an indicator for childlessness as the dependent variable. Appendix E Table E6 presents the summary statistics for these women. Appendix E Fig. E3a indicates that first-generation exposure to the OCP does not have a statistically significant

[60] We are grateful to an anonymous referee for suggesting this point.

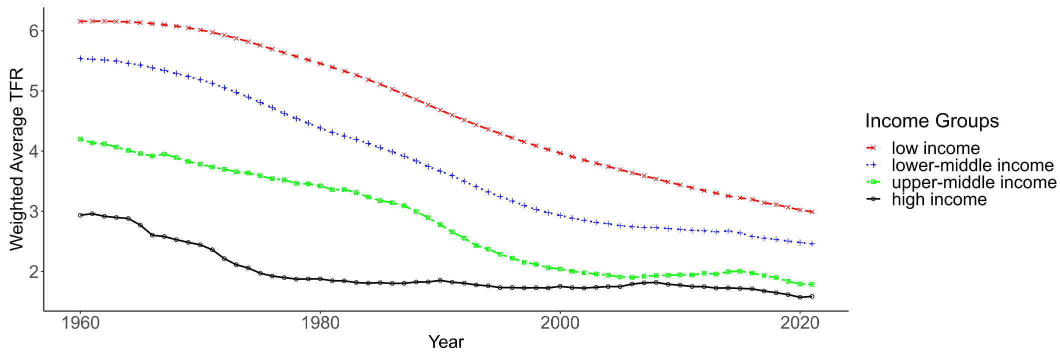
influence on the second-generation childlessness rate. Prior work stresses that the determinants of the extensive and intensive fertility margins can differ markedly (Aaronson *et al.*, 2014; Baudin *et al.*, 2015; Myong *et al.*, 2021). Our evidence is consistent with this view: A norm that lowers the number of children does not necessarily influence the decision regarding whether to have children or not.

We next investigate marital behavior by re-estimating Eq. 2 using two alternative dependent variables: an indicator for ever having married and age at first marriage. Appendix E Fig. E3b shows that first-generation exposure to the OCP has no statistically significant effects on the probability of ever marrying for second-generation women. By contrast, Appendix E Fig. E3c indicates that the same exposure delays marriage, which raises second-generation women's age at first marriage. We do not analyze age at first birth, because the 2020 census does not record child birth order for our analytical birth cohorts. Overall, the results suggest that the low-fertility norm created by the OCP relaxes the pressure to marry early. This allows women to postpone marriage, even though the overall propensity to marry remains unchanged.

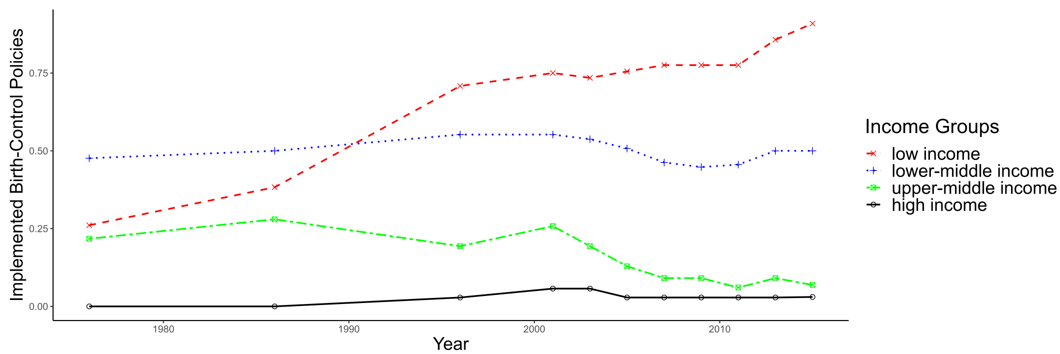
8 Conclusion

In this paper, we develop a model that incorporates the endogenous evolution of fertility norms to examine how birth-control policies can have lasting effects on fertility even after they are lifted. Using the staggered rollout and unexpected end of China's OCP as a natural experiment, we show that lower fertility among mothers leads to lower fertility among daughters—mainly through newly formed low-fertility norms—rather than through changes in childbearing costs or other economic factors. These results suggest that once low-fertility norms take hold, simply removing restrictions may not be enough to raise birth rates. This underscores the need for more coordinated pro-natal policies that can reshape people's fertility intentions.

We suggest two possible directions for future research. First, our model assumes that a mother's fertility affects her daughter's fertility mainly through its influence on fertility norms. Future work could extend this framework by allowing for additional pathways, such as changes in daughters' educational attainment and labor force participation. Although our robustness checks do not support these alternative mechanisms in this context, they may play a role in other settings. Second, our theoretical framework could be applied to investigate the intergenerational effects of other socioeconomic policies that may shape social norms. Such extensions could offer valuable insights into the broader societal consequences of policy interventions.



(a) Total Fertility Rate of Countries, by Income Group



(b) Share of Countries that Implemented Birth-control Policies, by Income Group

Figure 1: Global Patterns of Total Fertility Rates and Birth-control Policies. Figure (a) illustrates weighted average total fertility rates within each income group across time. Figure (b) shows the share of countries implementing birth-control policies within each income group across time. Fertility data are from the World Bank's World Development Indicators, and countries are classified into four income groups based on their earliest available income classification data from the World Bank. Population policy types are extracted from the United Nations World Population Policies Database. In Figure (b), we define countries implementing policies to lower fertility as "implementing birth-control policies."

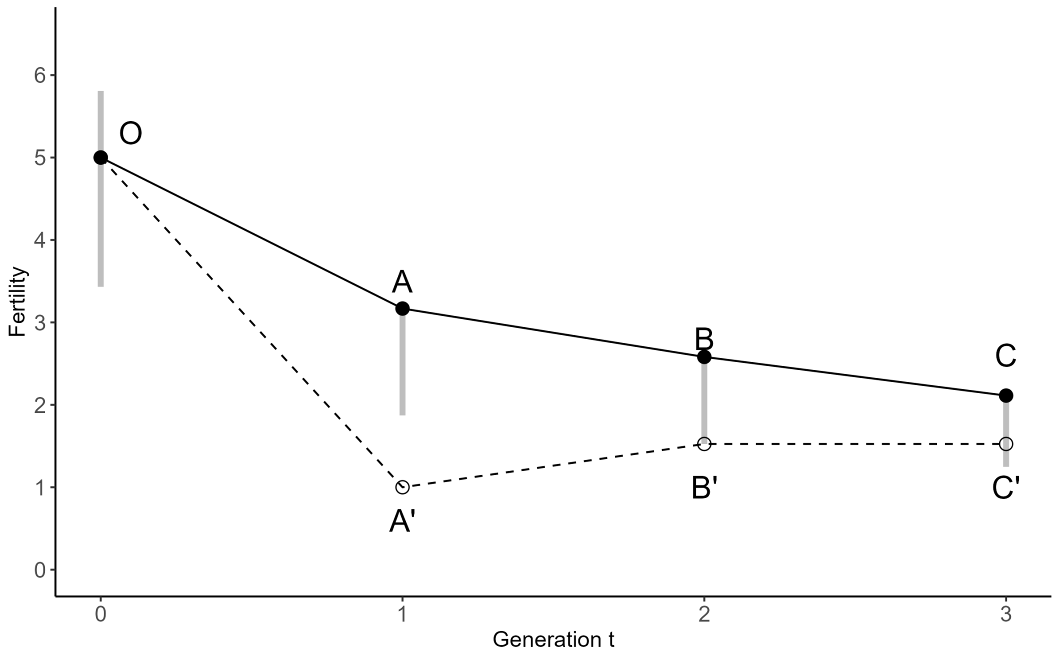
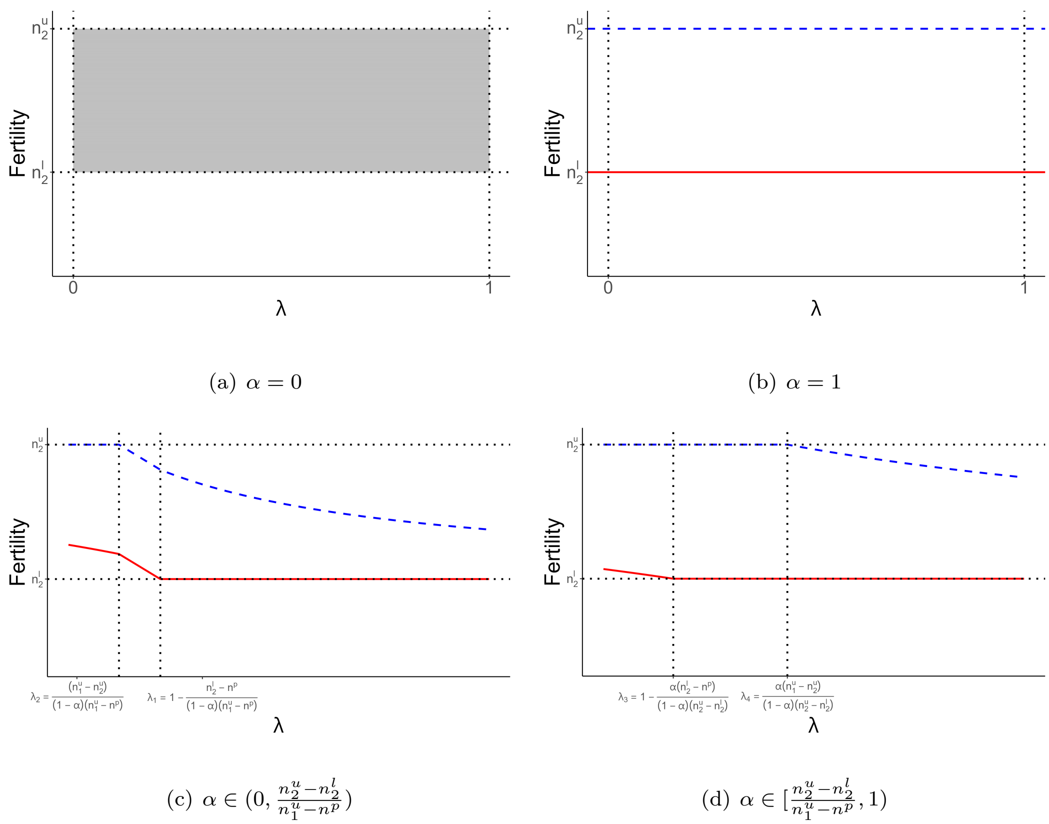


Figure 2: Theoretical Predictions of the Intergenerational Effects of One-shot Birth-control Policies for Generation 1, without Peer Effect. The x-axis denotes generations, the y-axis shows fertility, and vertical gray lines represent the fertility interval $[n_t^l, n_t^u]$ for each generation t. The solid line OABC denotes the fertility trend when the birth-control policy never comes into existence. The dashed line OA'B'C' depicts the fertility trend when a mandatory birth-control policy restricts fertility for generation 1, but not for other generations.



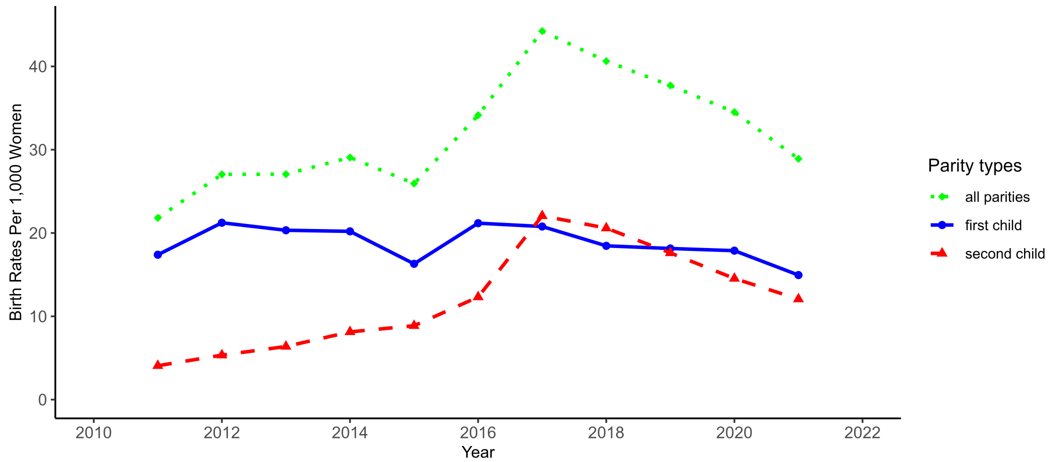
(a) $\alpha = 0$

(b) $\alpha = 1$

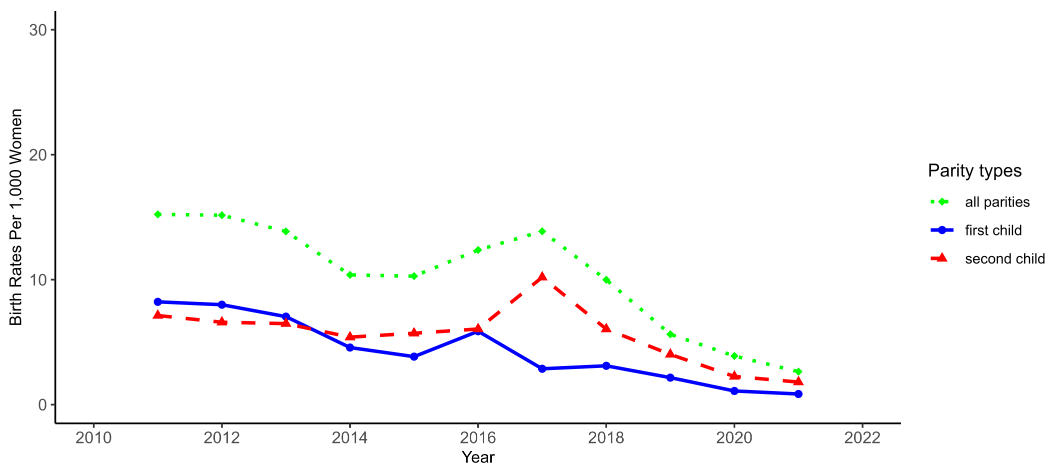
(c) $\alpha \in (0, \frac{n_2^u - n_2^l}{n_1^u - n^p})$

(d) $\alpha \in [\frac{n_2^u - n_2^l}{n_1^u - n^p}, 1)$

Figure 3: Equilibrium Fertility of Generation 2, with both Maternal and Peer Effects. These figures show how the equilibrium fertility of different types of women in generation 2 changes with varying shares of constrained women in generation 1 (λ). Figures (a)–(d) show various circumstances corresponding to different ranges of α, where α denotes the weight of mother's fertility ($n_{i,t-1}$) in determining the fertility norm. The x-axis represents the value of λ ranging from 0 to 1, and the y-axis represents fertility levels ranging from n_2^l to n_2^u. In Figure (a), as $\alpha = 0$, the peer effect results in the same fertility level for all individuals in generation 2. Multiple equilibria exist, and the equilibrium fertility level can take any value between n_2^l and n_2^u. The shaded area represents all possible equilibrium fertility for λ ranging from 0 to 1. In Figures (b), (c), and (d), the equilibrium is unique for given values of α and λ; solid red lines represent equilibrium fertility for individuals with restricted mothers, while dashed blue lines represent equilibrium fertility for those with unrestricted mothers.

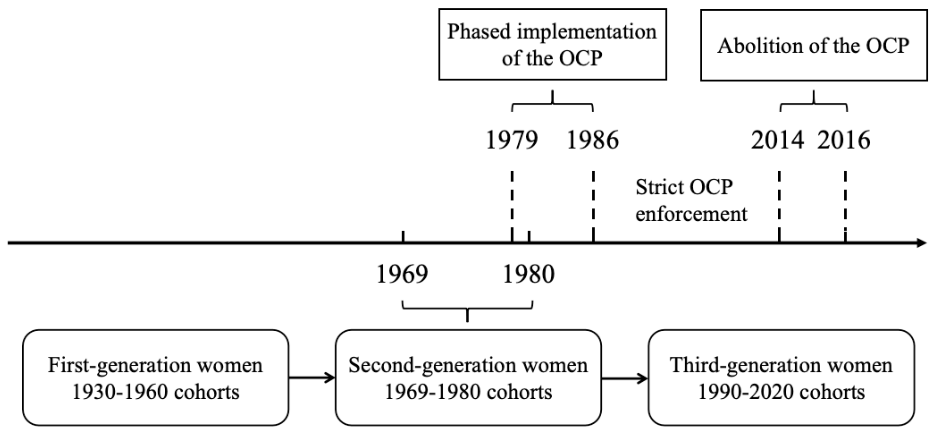


(a) Overall Trends

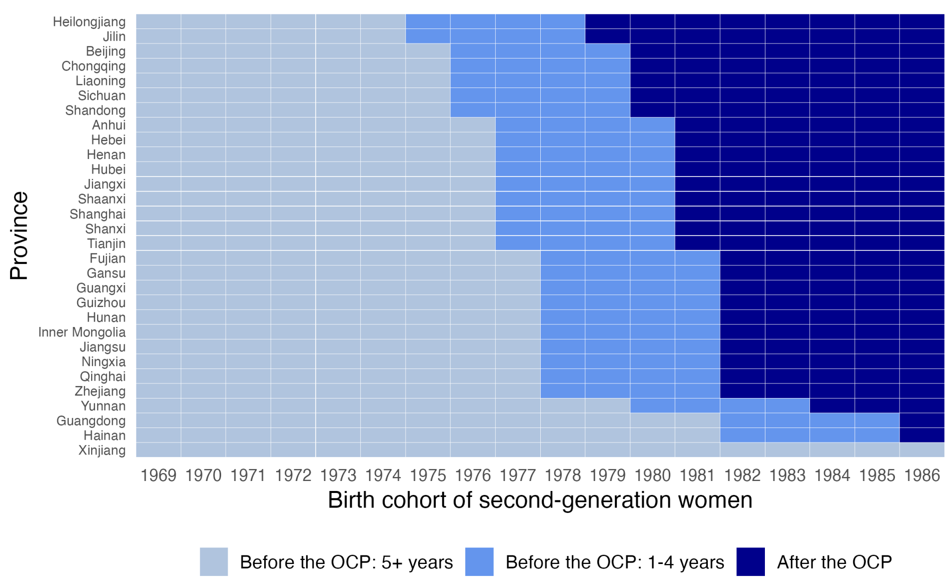


(b) Birth Cohorts Born during 1969-1980

Figure 4: Birth Rates among Urban Chinese Women, by Birth Parity. These two figures depict the birth rates among urban Chinese women for first-born and second-born children from 2011 to 2021. Figures (a) and (b) illustrate the trends in birth rates per 1,000 urban women aged 15-49 during survey years and per 1,000 urban women born during 1969-1980, respectively. Data are from *China Population and Employment Statistical Yearbook 2008-2022*.

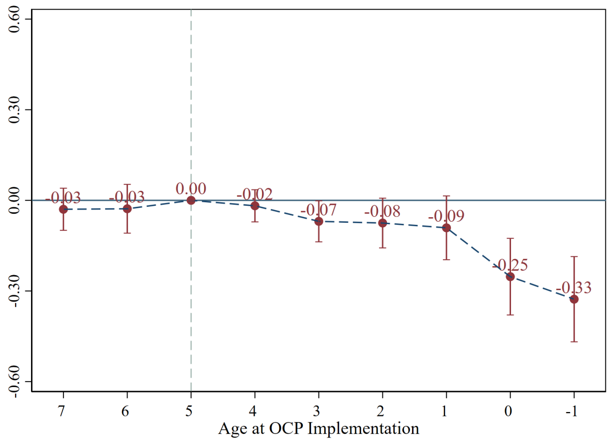


(a) Timeline: OCP and Three Generations

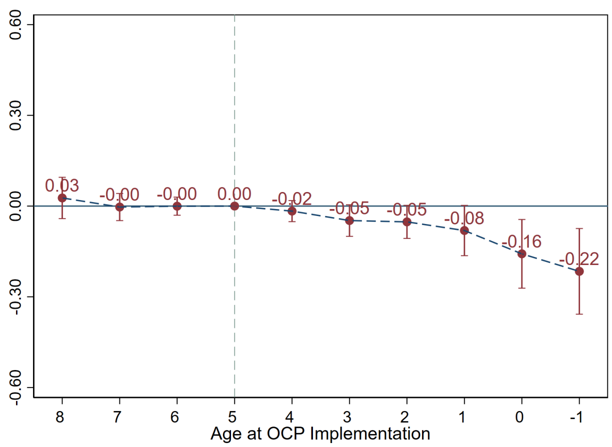


(b) Birth Cohort of Second-generation Women Relative to OCP Rollout

Figure 5: Timeline and OCP Rollout. Figure (a) positions the three generations of women along a timeline marking the initial implementation and eventual abolition of the OCP. In Figure (b), the x-axis represents the birth years of second-generation women, while the y-axis lists province names. Gray cells indicate second-generation cohorts born more than 5 years before implementation of the OCP, light blue cells represent those born within 1–4 years before OCP implementation, and dark blue cells denote those born after OCP implementation.

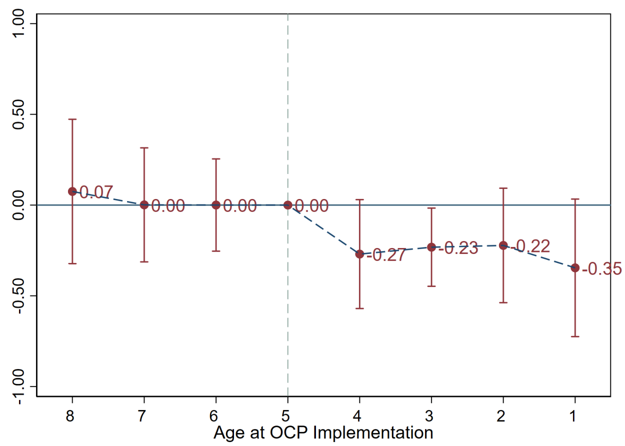


(a) Second-generation Women's Sibling Size

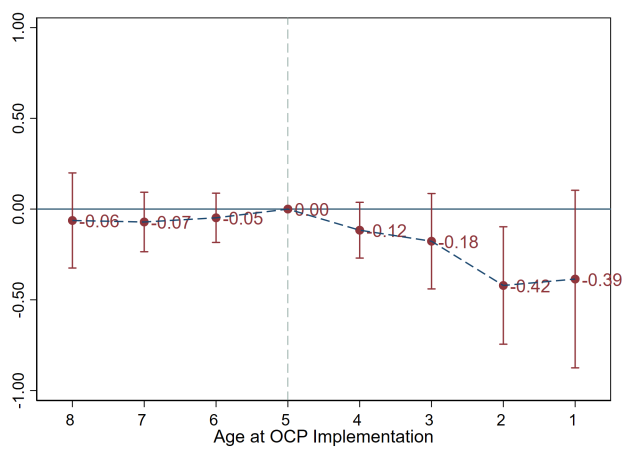


(b) Second-generation Women's Fertility

Figure 6: Impacts of the OCP Rollout on Second-generation Women's Sibling Size and Fertility. These figures plot coefficients estimated using Eq. 2. The dependent variables are the sibling size of second-generation women and the number of children they had ever born by 2020. Figure (a) shows the impact of the OCP rollout on second-generation women's sibling size. Data are from the 2005 mini-census and the sample includes urban Han women born during 1975-1980. Figure (b) shows the impact of the OCP rollout on second-generation women's fertility. Data are from the 2020 census and the sample includes urban Han women born during 1969 -1980. All specifications control for province fixed effects, cohort fixed effects, and province-specific cohort linear trends. Robust standard errors are clustered at province level. Vertical bars represent confidence intervals at the 95% level.



(a) Second Generation



(b) Third Generation

Figure 7: Impact of the OCP Rollout on Desired Fertility for the Second Generation and Third Generation. These figures plot coefficients estimated using Eq. 2. The dependent variable is desired fertility. Figure (a) shows the impact of the OCP rollout on desired fertility for the second generation. Figure (b) shows the impact of the OCP rollout on desired fertility for the third generation. Data are from CFPS 2012, 2014, and 2018, and samples are restricted to urban Han mothers born during 1969 - 1980. Control variables include province fixed effects, cohort fixed effects, and province-specific cohort linear trends. Robust standard errors are clustered at province level. Vertical bars represent confidence intervals at the 95% level.

Table 1: Summary Statistics of Censuses

Variables	(1) *Mean*	(2) *Std.Dev.*	(3) *N*
Panel A: Second-generation Fertility in Census 2020			
Number of children ever born by 2020	1.456	0.220	75,438
Panel B: First-generation Fertility in Mini-census 2005			
Sibling size of second-generation women	1.357	0.440	32,636
Panel C: Second-generation Fertility before OCP Relaxation in Mini-census 2015			
Number of children ever born by 2013	1.374	0.136	59,561
Panel D: Second-generation Variables before OCP Relaxation in Census 2010			
Human capital accumulation			
Primary school completion (Yes = 1)	0.994	0.009	198,203
Junior high school completion (Yes = 1)	0.876	0.083	198,203
Senior high school completion (Yes = 1)	0.389	0.123	198,203
College completion (Yes = 1)	0.077	0.051	198,203
Labor market performance			
Working at least one hour in the past week (Yes = 1)	0.758	0.055	198,203
Participating in the labor market (Yes = 1)	0.810	0.053	198,203
Total working hours in the past week	46.447	11.795	148,251
Living arrangement			
Coresiding with parents or parents-in-law (Yes = 1)	0.204	0.403	195,433

Notes: This table presents summary statistics for fertility outcomes and related variables in China population censuses. Panel A shows the number of children ever born to second-generation women in the 2020 census. Panel B shows second-generation women's sibling size, which represents first-generation women's fertility. Sibling size data are only available in the 2005 mini-census. Panel C shows the number of children born to second-generation women before the OCP relaxation in 2013, as observed in the 2015 mini-census. Panel D shows before-relaxation outcomes of second-generation women, including human capital accumulation, labor market performance, and living arrangements in the 2010 census. Panels A, C, and D include urban Han women born during 1969–1980. Panel B include urban Han women born during 1975–1980, because individuals born before 1975 do not report sibling size in the 2005 mini-census.

Table 2: Impact of the OCP Rollout on Fertility Outcomes

	(1)	(2)	(3)
Sample	*Mini-census 2005*	*Census 2020*	*Mini-census 2015*
Dependent Variable	*First-generation Fertility (Second-generation Sibling Size)*	*Second-generation Fertility by 2020*	*Second-generation Fertility by 2013 before OCP Relaxation*
Age at OCP = 8	-	0.027	-0.028
	-	(0.035)	(0.043)
Age at OCP = 7	-0.030	-0.003	0.006
	(0.034)	(0.023)	(0.033)
Age at OCP = 6	-0.028	-0.000	-0.041
	(0.040)	(0.015)	(0.031)
Age at OCP = 5	-	-	-
	-	-	-
Age at OCP = 4	-0.018	-0.017	-0.026
	(0.026)	(0.018)	(0.041)
Age at OCP = 3	-0.070**	-0.048*	-0.018
	(0.033)	(0.027)	(0.045)
Age at OCP = 2	-0.075*	-0.052*	-0.013
	(0.040)	(0.028)	(0.062)
Age at OCP = 1	-0.091*	-0.081*	-0.022
	(0.052)	(0.042)	(0.073)
Age at OCP = 0	-0.253***	-0.158**	-0.049
	(0.062)	(0.058)	(0.098)
Age at OCP = -1	-0.327***	-0.216***	-0.043
	(0.069)	(0.072)	(0.134)
Observations	32,636	75,438	59,561
Mean of dep. var.	1.357	1.456	1.374

Notes: This table shows the coefficients estimated using Eq. 2. Column (1) shows the impact of the OCP rollout on the sibling size of second-generation women in the 2005 mini-census. Samples are restricted to second-generation Han women with urban hukou born during 1975–1980, because individuals born before 1975 do not report the number of siblings in the 2005 mini-census. Control variables include province and cohort fixed effects. Column (2) shows the impact of the OCP rollout on the number of children ever born to second-generation women in the 2020 census. Column (3) shows the impact of the OCP rollout on the number of children born to second-generation women before the OCP relaxation in 2013, as observed in the 2015 mini-census. Samples in Columns (2) and (3) are restricted to urban Han women born between 1969 and 1980. Control variables include province fixed effects, cohort fixed effects, and province-specific linear cohort trends. Robust standard errors clustered at province level are in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

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The Intergenerational Impact of Birth-Control Policies on Fertility: The Role of Norms

Online Appendix: The Intergenerational Impact of Birth-Control Policies

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A Full Derivation of the Theory

A.1 Proof of Proposition 1

The utility maximization problem of generation t is

$$U_t = \ln(c_t) + \gamma \ln(n_t) - \sigma |\ln(n_t) - \ln(n_{t-1})|,$$

$$s.t. \quad c_t + p_t n_t = y_t,$$

where γ represents the degree of maternal altruism toward children, p_t is the cost of child-bearing, y_t is income of generation t in adulthood, and $\sigma \geq 0$ measures the utility loss for every log point deviation of the woman' s fertility from her mother's fertility. Actually, the utility function is a piecewise function:

$$U_t = \begin{cases} \ln(c_t) + \gamma \ln(n_t) - \sigma[\ln(n_{t-1}) - \ln(n_t)], & \text{if} \quad n_t < n_{t-1} \\ \ln(c_t) + \gamma \ln(n_t) - \sigma[\ln(n_t) - \ln(n_{t-1})], & \text{if} \quad n_t \geq n_{t-1}. \end{cases}$$

We will arrive at the optimal solution via a case-by-case discussion.

Case 1. Suppose that mother's fertility is higher than daughter's optimal fertility ($n_{t-1} > n_t^*$); then daughter's utility function takes the form $\ln(c_t)+(\gamma+\sigma)\ln(n_t)-\sigma\ln(n_{t-1})$. When $n_{t-1} > n_t^*$, the utility function is continuously differentiable at n_t^*. The first-order conditions (FOCs) determine optimal fertility:

$$n_t^* = \frac{\gamma + \sigma}{1 + \gamma + \sigma} \frac{y_t}{p_t} \equiv n_t^u.$$

This case requires mother's fertility to be sufficiently high ($n_{t-1} > n_t^u$), such that mother's high fertility norm induces daughter's fertility n_t^* to hit an upper bound n_t^u.

Case 2. Symmetrically, when mother's fertility is sufficiently low ($n_{t-1} < n_t^*$), daughter's utility function becomes $\ln(c_t) + (\gamma - \sigma)\ln(n_t) + \sigma \ln(n_{t-1})$. The FOCs give optimal fertility:

$$n_t^* = \frac{\gamma - \sigma}{1 + \gamma - \sigma} \frac{y_t}{p_t} \equiv n_t^l.$$

This case requires mother's fertility to be sufficiently low ($n_{t-1} < n_t^l$).

Case 3. When mother's fertility lies between the lower and upper bounds ($n_t^l \leq n_{t-1} \leq n_t^u$), we will prove $n_t^* = n_{t-1}$ using contradiction. Suppose $n_t^* < n_{t-1}$. By the reasoning in case 1, we have $n_t^* = n_t^u$. So $n_t^u < n_{t-1}$, contradiction. Suppose $n_t^* > n_{t-1}$. By the reasoning in case 2, we have $n_t^* = n_t^l$. So $n_t^l > n_{t-1}$, contradiction. Thus we must have $n_t^* = n_{t-1}$. In this case, mother's fertility dictates daughter's fertility. When optimal fertility lies strictly within the range ($n_t^l < n_{t-1} < n_t^u$), small changes in intrinsic costs and benefits (p_t/y_t, γ) do not affect daughter's fertility. The "conformity" to social norms arises, because when n_t increases and passes over n_{t-1}, the marginal benefit of n_t drops discontinuously (from $\frac{\gamma+\sigma}{n_t}$ to $\frac{\gamma-\sigma}{n_t}$ in our case), but the marginal cost of n_t rises continuously (Akerlof, 1997).

To sum up, the case-by-case discussion gives

$$n_t^* = \begin{cases} n_t^u & \text{if} \quad n_{t-1} > n_t^u, \\ n_{t-1} & \text{if} \quad n_t^l \leq n_{t-1} \leq n_t^u, \\ n_t^l & \text{if} \quad n_{t-1} < n_t^l. \end{cases}$$

n_{t-1} establishes the norm of daughter's fertility, but the impact is limited to a range set by the bounds, which encloses the intrinsic optimum if $\sigma > 0$. This completes the proof of Proposition 1.

A.2 Desired Fertility under Zero Childbearing Cost

Under the hypothetical scenario of zero childbearing cost, the utility maximization problem for generation t is

$$\max_{n_t} U_t = \ln(y_t) + \gamma \ln(n_t) - \sigma |\ln(n_t) - \ln(n_{t-1})|,$$

where parents consume all income and freely choose the number of children. When $\sigma > \gamma$, we can show $n_t^* = n_{t-1}$. That is to say, under zero childbearing cost, desired fertility would fully capture the fertility norm.

We prove $n_t^* = n_{t-1}$ by contradiction. (i) Suppose $n_t^* < n_{t-1}$, the utility function for woman i, becomes $U_t = \ln(y_t) + (\gamma + \sigma) \ln(n_t) - \sigma \ln(n_{t-1})$. Because $\gamma + \sigma > 0$, the utility function increases with n_t. Consider $n_t^* < n_t^* + \epsilon < n_{t-1}$, then $U_t(n_t^*) < U_t(n_t^* + \epsilon)$, which contradicts the assumption that n_t^* is the optimal choice. (ii) Suppose $n_t^* > n_{t-1}$. The utility function becomes $U_t = \ln(y_t) + (\gamma - \sigma) \ln(n_t) - \sigma \ln(n_{t-1})$. Because $\gamma - \sigma < 0$, the utility function decreases with n_t. Consider $n_t^* > n_t^* - \epsilon > n_{t-1}$, then $U_t(n_t^*) < U_t(n_t^* - \epsilon)$, which contradicts the assumption that n_t^* is the optimal choice.

A.3 Proof of Proposition 2

Suppose that the cost of having a child consistently and progressively increases from generations 1 to 3:

$$\Psi_t > 0 \ , \forall i = 1, 2, 3.$$

Similar to the simple scenario above, the fertility of generation 0 begins at $\tilde{n}_0 \in [n_0^l, n_0^u] = [\frac{\gamma-\sigma}{1+\gamma-\sigma}\frac{y_0}{p_0}, \frac{\gamma+\sigma}{1+\gamma+\sigma}\frac{y_0}{p_0}]$: $n_0^* = \tilde{n}_{-1}$ (points O in Fig. 2). We also assume $p_1 > \frac{\gamma+\sigma}{1+\gamma+\sigma}\frac{y_1}{\tilde{n}_0}$ such that $n_1^u < \tilde{n}_0$.

As denoted by the dashed path OABC in Fig. 2, in the absence of birth-control policies, the fertility transition would have followed the upper bounds for generations 1, 2, and 3. A one-shot birth-control policy in generation 1 triggers the fertility transition (the solid path OA'B'C'). When $n^p < n_2^l$, the fertility of generation 2 slightly rebounds but only reaches the lower bound—that is, $n_2^* = n_2^l$.

In reality, a fertility transition results from a combination of rising child prices and declining maternal altruism toward children (γ). Proposition 2 is robust to variations in γ as long as changes in γ do not overturn the conditions on the fertility intervals ($n_1^u < \tilde{n}_0, n_2^l < n_3^u, n^p < n_2^l$). The lasting fertility effect of a one-shot birth-control policy does not hinge on the source of fertility decline.

A.4 The "Stable Cost" Scenario

Appendix Fig. A2 illustrates the lasting effect of the one-shot birth-control policy in a "stable cost" scenario. Suppose the cost-income ratio $\frac{p_t}{y_t}$ remains the same across generations 0 to 3, such that $\Psi_t = 0 \ , \forall i = 1, 2, 3$. The solid line, OABC in Fig. A2, represents the "natural" equilibrium fertility rates without birth-control policies. Point O denotes the initial fertility level $\tilde{n}_0$, such that $\tilde{n}_0 > n_0^l$. Under this scenario, the fertility interval $[n_t^l, n_t^u]$ remains the same across all four generations. By Proposition 1, the equilibrium fertility rate remains the same across generations 1, 2, and 3, $n_t^* = \tilde{n}_0, \ \forall i = 1, 2, 3$, represented by points A, B, and C.

The dashed line, OA'B'C', represents the equilibrium fertility rates with a one-shot birth-control policy on generation 1. Comparing the dashed line with the solid line, we observe a lasting effect of the policy. Because of the policy, the fertility of generation 1 is constrained to be n^p such that $n^p < n_1^l$ (point A'). When the policy is removed after this generation, the fertility of generation 2 does not bounce back to point B, the natural equilibrium fertility rate. This is because the constrained low fertility of generation 1 establishes a new norm for generation 2. The equilibrium fertility is n_2^l for both generations 2 and 3, which demonstrates a lasting fertility-reduction effect of a one-shot birth-control policy.

A.5 Model Extension: Incorporating the Peer Effect

Our theory highlights the evolution of fertility norms behind the persistent fertility-reduction effect of past birth-control policies. The literature has extensively investigated the peer effect in shaping fertility norms, i.e., the spread of new fertility norms within the same generation (Richerson and Boyd, 2008; Spolaore and Wacziarg, 2022; Beach and Hanlon, 2023). There is little doubt that the peer effect serves as an important channel through which new fertility norms spread across time and space. We incorporate the peer effect in our theory, and find a robust prediction of the persistent fertility-reduction effect of past birth-control policies for subsequent generations. Furthermore, the peer effect helps spread the new fertility norm to descendants of initially unrestricted lineages. Importantly, when the maternal effect is absent and the peer effect functions alone, a one-shot birth-control policy that restricts just one generation would have no effect on the fertility of subsequent generations. Thus, the maternal effect is pivotal for the persistent fertility-reduction effect of one-shot birth-control policies.

To incorporate the peer effect, suppose the fertility norm now consists of the mother's fertility and the expected fertility of peers. The utility maximization problem of lineage i at generation t is

$$\begin{aligned} \max_{c_{i,t}, n_{i,t}} \; U_{i,t} = \;& \ln(c_{i,t}) + \gamma \ln(n_{i,t}) - \sigma |\ln(n_{i,t}) - \ln(N_{i,t})|, \\ \text{s.t.} \quad & \alpha n_{i,t-1} + (1-\alpha)\bar{n}_t = N_{i,t}, \\ & c_{i,t} + p_t n_{i,t} = y_t, \end{aligned} \tag{P2}$$

where $N_{i,t}$ denotes the fertility norm, which is a weighted average of mother's fertility ($n_{i,t-1}$) and the expected fertility of peers ($\bar{n}_t$). The weight $\alpha \in [0, 1]$ represents the relative strength of maternal versus peer effects. When $\alpha = 1$, the peer effect is absent: P2 collapses to P1. When $\alpha = 0$, the maternal effect is nonexistent, and the peer effect operates in isolation; this scenario will be discussed further below. In a typical case of $\alpha \in (0, 1)$, both maternal and peer effects are in effect.

Equilibrium Fertility in the Absence of a Birth-control Policy

When there is no birth-control policy and $n_{i,t-1}$ is the same for each i, the solution of P2 is almost identical to that of P1.

Lemma 1. *In P2, when* $\sigma \geq 0$ *and conditional on the same* $n_{i,t-1} = n_{t-1}$ *for each* i*, the optimal fertility of woman* i *at generation* t *is*

$$n^*_{i,t} = \begin{cases} n^u_t & \text{if} \quad n_{t-1} > n^u_t, \\ n_{t-1} & \text{if} \quad n^l_t \leq n_{t-1} \leq n^u_t, \\ n^l_t & \text{if} \quad n_{t-1} < n^l_t. \end{cases}$$

where $[n^l_t, n^u_t] = [\frac{\gamma-\sigma}{1+\gamma-\sigma}\frac{y_t}{p_t}, \frac{\gamma+\sigma}{1+\gamma+\sigma}\frac{y_t}{p_t}]$.

Proof. See Appendix Section A.6.

Equilibrium Fertility under a Birth-control Policy

Suppose that a birth-control policy restricts the fertility level of a proportion $\lambda \in (0,1)$ of mothers in generation 1 to $n^p < n^l_2$.[1] For the remaining $1-\lambda$ proportion of mothers, the birth-control policy does not bind. We denote the equilibrium fertility of generation-t individuals in restricted lineages as $n^*_{r,t}$, and those in the unrestricted lineages as $n^*_{ur,t}$.

In generation 2, two sets of fertility norms emerge: Daughters with restricted mothers adhere to a fertility norm $N_{r,2}$:

$$N_{r,2} = \alpha n^p + (1-\alpha)\bar{n}_2,$$

and daughters with unrestricted mothers adopt a fertility norm $N_{ur,2}$:

$$N_{ur,2} = \alpha n_{ur,1} + (1-\alpha)\bar{n}_2.$$

[1]If the birth-control policy uniformly restricts every mother in generation 1, introduction of the peer effect would not yield fundamentally different theoretical results.

Because $n_{ur,1} \in [n_1^l, n_1^u]$, we must have $n^p < n_{ur,1}$; thus $N_{r,2} < N_{ur,2}$. For simplicity of notation, suppose that the fertility of unrestricted generation-1 mothers stays at the upper bound $n_{ur,1} = n_1^u$. In the face of different fertility norms (n^p vs. n_1^u), daughters in generation 2 may opt for different levels of optimal fertility.

Individuals optimize by solving for P2 in the expectation of the optimal fertility choices of others. A Nash equilibrium is sustained by the condition of rational expectation:

$$\bar{n}_t = \lambda n^*_{r,t} + (1-\lambda) n^*_{ur,t}.$$

In the Nash equilibrium, no individual has an incentive to deviate given the optimal fertility of others. Lemmas 2 and 3 characterize the equilibrium fertility of generation 2.

LEMMA 2. *Suppose the maternal effect is moderate relative to the peer effect,* $\alpha \in (0, \frac{n_2^u - n_2^l}{n_1^u - n^p})$. *A birth-control policy restricts the fertility level of a proportion* $\lambda \in (0,1)$ *of generation-1 mothers to* $n_{r,1} = n^p$*, which is lower than the lower bound of generation 2,* $n^p < n_2^l$*. The remaining* $1-\lambda$ *proportion of unrestricted generation-1 mothers attain* $n_{ur,1} = n_1^u$*. The strength of* λ *pins down equilibrium fertility for generation 2.*

If $\lambda \in (0, \frac{n_1^u - n_2^u}{(1-\alpha)(n_1^u - n^p)}]$*, then*

$$\begin{cases} n^*_{r,2} &= N_{r,2} = \alpha n^p + (1-\alpha)\frac{\alpha\lambda n^p + (1-\lambda) n_2^u}{1-\lambda+\alpha\lambda}, \\ n^*_{ur,2} &= n_2^u, \\ \bar{n}_2 &= \frac{\alpha\lambda n^p + (1-\lambda) n_2^u}{1-\lambda+\alpha\lambda}. \end{cases}$$

If $\lambda \in (\frac{n_1^u - n_2^u}{(1-\alpha)(n_1^u - n^p)}, 1 - \frac{n_2^l - n^p}{(1-\alpha)(n_1^u - n^p)})$, *then*

$$\begin{cases} n^*_{r,2} & = N_{r,2} = \alpha n^p + (1-\alpha)(\lambda n^p + (1-\lambda) n_1^u), \\ n^*_{ur,2} & = N_{ur,2} = \alpha n_1^u + (1-\alpha)(\lambda n^p + (1-\lambda) n_1^u), \\ \bar{n}_2 & = \lambda n^p + (1-\lambda) n_1^u. \end{cases}$$

If $\lambda \in [1 - \frac{n_2^l - n^p}{(1-\alpha)(n_1^u - n^p)}, 1)$, *then*

$$\begin{cases} n^*_{r,2} & = n_2^l, \\ n^*_{ur,2} & = N_{ur,2} = \alpha n_1^u + (1-\alpha)\frac{\lambda n_2^l + (1-\lambda)\alpha n_1^u}{\lambda + \alpha - \lambda\alpha}, \\ \bar{n}_2 & = \frac{\lambda n_2^l + (1-\lambda)\alpha n_1^u}{\lambda + \alpha - \lambda\alpha}. \end{cases}$$

LEMMA 3. *Suppose the maternal effect is strong relative to the peer effect,* $\alpha \in [\frac{n_2^u - n_2^l}{n_1^u - n^p}, 1)$. *A birth-control policy restricts the fertility level of a proportion* $\lambda \in (0, 1)$ *of generation-1 mothers to* $n_{r,1} = n^p$, *which is lower than the lower bound of generation 2,* $n^p < n_2^l$. *The remaining* $1 - \lambda$ *proportion of unrestricted generation-1 mothers attain* $n_{ur,1} = n_1^u$. *The strength of* λ *pins down equilibrium fertility for generation 2.*

If $\lambda \in (0, 1 - \frac{\alpha(n_2^l - n^p)}{(1-\alpha)(n_2^u - n_2^l)})$, *then*

$$\begin{cases} n^*_{r,2} & = N_{r,2} = \alpha n^p + (1-\alpha)\frac{\alpha\lambda n^p + (1-\lambda) n_2^u}{1 - \lambda + \alpha\lambda}, \\ n^*_{ur,2} & = n_2^u, \\ \bar{n}_2 & = \frac{\alpha\lambda n^p + (1-\lambda) n_2^u}{1 - \lambda + \alpha\lambda}. \end{cases}$$

If $\lambda \in [1 - \frac{\alpha(n_2^l - n^p)}{(1-\alpha)(n_2^u - n_2^l)}, \frac{\alpha(n_1^u - n_2^u)}{(1-\alpha)(n_2^u - n_2^l)}]$, *then*

$$\begin{cases} n_{r,2}^* &= n_2^l, \\ n_{ur,2}^* &= n_2^u, \\ \bar{n}_2 &= \lambda n_2^l + (1-\lambda) n_2^u. \end{cases}$$

If $\lambda \in (\frac{\alpha(n_1^u - n_2^u)}{(1-\alpha)(n_2^u - n_2^l)}, 1)$, *then*

$$\begin{cases} n_{r,2}^* &= n_2^l, \\ n_{ur,2}^* &= N_{ur,2} = \alpha n_1^u + (1-\alpha)\frac{\lambda n_2^l + (1-\lambda)\alpha n_1^u}{\lambda + \alpha - \lambda\alpha}, \\ \bar{n}_2 &= \frac{\lambda n_2^l + (1-\lambda)\alpha n_1^u}{\lambda + \alpha - \lambda\alpha}. \end{cases}$$

Proofs of Lemmas 2 and 3 are relegated to Appendix Section A.7.

Lemmas 1–3 immediately predict that when $\alpha \in (0,1]$ and $\lambda \in (0,1]$, the one-shot birth-control policy in generation 1 reduces the fertility of generation 2. When there is no birth-control policy, by Lemma 1, the fertility of generation 2 stays at the upper bound. The one-shot birth-control policy in generation 1 reduces the average fertility of generation 2 to $\bar{n}_2$, which is, by Lemmas 2 and 3, strictly lower than the upper bound n_2^u.

Changing Policy Coverage (λ)

Fig. 2a and Fig. 2b illustrate how equilibrium fertility changes with λ when $\alpha \in (0, \frac{n_2^u - n_2^l}{n_1^u - n^p})$ and $\alpha \in [\frac{n_2^u - n_2^l}{n_1^u - n^p}, 1)$, respectively. Solid red lines represent equilibrium fertility for individuals with restricted mothers ($n_{r,2}^*$), and dashed blue lines represent equilibrium fertility for those with unrestricted mothers ($n_{ur,2}^*$). In both figures, when λ increases from zero, $n^*_{r,2}$ immediately declines until it reaches the lower bound n^l_2; $n^*_{ur,2}$ initially sticks to the upper bound n^u_2, but eventually starts to decline when λ passes a threshold value. Based on Lemmas 2 and 3, we derive Lemmas 4–7 to depict the evolution of $n^*_{r,2}$, $n^*_{ur,2}$, and $\bar{n}_2$ in response to changes in λ.

LEMMA 4. *When λ increases from zero to $\underline{\lambda}$, $n^*_{r,2}$ strictly decreases to the lower bound n^l_2. Further increases in λ have no effect on $n^*_{r,2}$.*

1. *If $\alpha \in (0, \frac{n^u_2 - n^l_2}{n^u_1 - n^p})$, $\underline{\lambda} = 1 - \frac{n^l_2 - n^p}{(1-\alpha)(n^u_1 - n^p)}$.*
2. *If $\alpha \in [\frac{n^u_2 - n^l_2}{n^u_1 - n^p}, 1)$, $\underline{\lambda} = 1 - \frac{\alpha(n^l_2 - n^p)}{(1-\alpha)(n^u_2 - n^l_2)}$.*

In both cases, $\underline{\lambda}$ strictly increases as α decreases or as n^p increases.

LEMMA 5. *When λ increases from 0 to $\bar{\lambda}$, $n^*_{ur,2}$ remains at the upper bound n^u_2. When λ passes the threshold $\bar{\lambda}$, $n^*_{ur,2}$ strictly decreases with λ.*

1. *If $\alpha \in (0, \frac{n^u_2 - n^l_2}{n^u_1 - n^p})$, $\bar{\lambda} = \frac{n^u_1 - n^u_2}{(1-\alpha)(n^u_1 - n^p)}$.*
2. *If $\alpha \in [\frac{n^u_2 - n^l_2}{n^u_1 - n^p}, 1)$, $\bar{\lambda} = \frac{\alpha(n^u_1 - n^u_2)}{(1-\alpha)(n^u_2 - n^l_2)}$.*

In both cases, $\bar{\lambda}$ strictly increases as α increases or as the gap between the upper bounds of period 1 and period 2 ($n^u_1 - n^u_2$) increases.

LEMMA 6. *When λ increases from zero to one, the average fertility of generation 2 decreases: $\frac{\partial \bar{n}_2}{\partial \lambda} < 0$.*

Proofs of Lemmas 4–6 are relegated to Appendix Section A.8.

Changing Strength of Maternal versus Peer Effects (α)

A comparison of Fig. 3a and Fig. 3b suggests that the relative strength of maternal versus peer effects (α) affects the convergence of fertility between generation-2 individuals with restricted and unrestricted mothers. When the peer effect becomes more prominent (smaller α), the gap in equilibrium fertility between the two groups, $n^*_{ur,2} - n^*_{r,2}$, becomes smaller. We can prove Lemma 7:

LEMMA 7. *When the peer effect is more prominent (smaller α), the gap in equilibrium fertility between generation-2 individuals with restricted mothers and unrestricted mothers becomes smaller:* $\frac{\partial(n^*_{ur,2} - n^*_{r,2})}{\partial\alpha} \geq 0$.

Proof. See Appendix Section A.8.

We next explore scenarios in which α takes on extreme values of zero or one. When $\alpha = 1$, P2 collapses to P1, in which the peer effect is absent and mother's fertility exclusively shapes the fertility norm. As shown in Fig. 3c, the equilibrium fertility of individuals with restricted parents remains at the lower bound; for those with unrestricted parents, equilibrium fertility remains at the upper bound.

When $\alpha = 0$, the fertility norm is determined solely by the average fertility of peers. Consequently, past birth-control policies find no way to affect the fertility of the current generation. Akerlof (1997) has shown that the peer effect alone would result in multiple equilibria. As shown in Fig. 3d, when $\alpha = 0$, equilibrium fertility can take any value between n^u_2 and n^l_2. Interestingly, if we introduce a minimal degree of maternal effect by assigning α a value slightly larger than zero, multiple equilibria collapse into a unique equilibrium. The maternal effect turns out to provide an anchoring point for the multiple equilibria that arise from the peer effect alone.

Lasting Effects under the Peer Effect

The lasting fertility-reduction effect of the one-shot birth-control policy extends beyond generation 2. To illustrate, we again suppose that the economy consists of four generations: 0, 1, 2, and 3. The fertility of generation 0 begins at a relatively high level $n^*_{i,0} = \tilde{n}_0 \in [n^l_0, n^u_0]$. Child price rises or child preference diminishes in generation 1 and remains stable thereafter, and these variations are strong enough to reduce fertility in generation 1: $n^*_{i,0} > n^u_1 = n^u_2 = n^u_3$. In the absence of birth-control policies, by Lemma 1 equilibrium fertility from generations 0 to 3 follows path OABC in Fig. A1.

Paths OA'B'C' and OA"B"C" in Fig. A1 trace equilibrium fertility for restricted and unrestricted lineages ($n^*_{r,t}$ and $n^*_{ur,t}$), respectively. The one-shot birth-control policy in generation 1 lowers the average fertility for generations 2 and 3. Under the peer effect, equilibrium fertility tends to converge over generations between restricted and unrestricted lineages. The speed of convergence is higher when the peer effect is more pronounced. The average fertility of generation t decreases as the policy coverage expands, i.e., $\frac{\partial \bar{n}_t}{\partial \lambda} < 0, \forall t \geq 1$.

Our model prediction for the fertility-reduction effect of birth-control policies is consistent with that of De Silva and Tenreyro (2020), who suggest that birth-control policies alter family-size norms and significantly accelerate the decline in fertility toward replacement level. Despite the acceleration effect, birth-control policies do not affect the ultimate fertility level.[2] Different from De Silva and Tenreyro (2020), we predict a lasting effect of birth-control policies on reducing the fertility level. While De Silva and Tenreyro (2020) focus on episodes of fertility transition toward replacement-level fertility, our model is better suited to the recent experience of upper-middle income countries, as shown in Fig. 1c.

[2] De Silva and Tenreyro (2020) define fertility norms as a weighted average between the fertility of the previous generation and long-term replacement-level fertility. Birth-control policies increase the weight attributed to replacement-level fertility, and thereby lower the fertility norm.

A.6 Proof of Lemma A.1

We expand our model to include the peer effect by integrating the expected fertility of peers into the fertility norm. The utility maximization problem of individual i in generation t is

$$\begin{aligned} \max_{c_{i,t},n_{i,t}} U_{i,t} = \; & \ln(c_{i,t}) + \gamma \ln(n_{i,t}) - \sigma|\ln(n_{i,t}) - \ln(N_{i,t})|, \\ \text{s.t.} \quad & \alpha n_{i,t-1} + (1-\alpha)\bar{n}_t = N_{i,t}, \\ & c_{i,t} + p_t n_{i,t} = y_t, \end{aligned} \tag{P2}$$

where $N_{i,t}$ denotes the fertility norm, which is a weighted average of mother's fertility ($n_{i,t-1}$) and the expected fertility of peers ($\bar{n}_t$). The weight $\alpha \in [0,1]$ represents the relative strength of maternal versus peer effects. The utility function is a piecewise function:

$$U_{i,t} = \begin{cases} \ln(y_t - p_t n_{i,t}) + \gamma \ln(n_{i,t}) - \sigma(\ln(N_{i,t}) - \ln(n_{i,t})), & \text{if } N_{i,t} > n_{i,t} \\ \ln(y_t - p_t n_{i,t}) + \gamma \ln(n_{i,t}) - \sigma(\ln(n_{i,t}) - \ln(N_{i,t})), & \text{if } N_{i,t} \leq n_{i,t} \end{cases}$$

We will arrive at the optimal solution via a case-by-case discussion. For simplicity, suppose $n_{i,t-1} = n_{t-1}$ is the same for each individual i.

Case 1. Suppose that fertility norm $N_{i,t}$ is higher than the individual's optimal fertility ($N_{i,t} > n^*_{i,t}$). Then the individual's utility function takes the form $\ln(c_{i,t}) + (\gamma+\sigma)\ln(n_{i,t}) - \sigma\ln(N_{i,t})$. When $N_{i,t} > n^*_{i,t}$, the utility function is continuously differentiable at $n^*_{i,t}$. The FOCs determine optimal fertility:

$$n^*_{i,t} = \frac{\gamma+\sigma}{1+\gamma+\sigma}\frac{y_t}{p_t} \equiv n^u_t.$$

The condition $N_{i,t} > n^*_{i,t}$ would be satisfied if and only if $n_{t-1} > n^u_t$.

Case 2. Symmetrically, when the fertility norm $N_{i,t}$ is sufficiently low ($N_{i,t} < n^*_{i,t}$), the individual's utility function becomes $\ln(c_{i,t}) + (\gamma - \sigma)\ln(n_{i,t}) + \sigma \ln(N_{i,t})$. The FOCs give optimal fertility

$$n^*_{i,t} = \frac{\gamma - \sigma}{1 + \gamma - \sigma}\frac{y_t}{p_t} \equiv n^l_t.$$

The condition $N_{i,t} < n^*_{i,t}$ would be satisfied if and only if $n_{t-1} < n^l_t$.

Case 3. When the fertility norm lies between the lower and upper bounds ($n^l_t \leq N_{i,t} \leq n^u_t$), we will prove $n^*_{i,t} = N_{i,t}$ using contradiction. Suppose $n^*_{i,t} < N_{i,t}$. By the reasoning in case 1, we have $n^*_{i,t} = n^u_t$. So $n^u_t < N_{i,t}$, contradiction. Suppose $n^*_{i,t} > N_{i,t}$. By the reasoning in case 2, we have $n^*_{i,t} = n^l_t$. So $n^l_t > N_{i,t}$, contradiction. Thus we must have $n^*_{i,t} = N_{i,t}$. In equilibrium, since every lineage has the same $n_{i,t-1} = n_{t-1}$, we must have $N_{i,t} = n_{t-1}$ for each i.

To sum up, the above case-by-case discussion gives

$$n^*_{i,t} = \begin{cases} n^u_t & \text{if} \quad n_{t-1} > n^u_t, \\ n_{t-1} & \text{if} \quad n^l_t \leq n_{t-1} \leq n^u_t, \\ n^l_t & \text{if} \quad n_{t-1} < n^l_t, \end{cases}$$

where $[n^l_t, n^u_t] = [\frac{\gamma-\sigma}{1+\gamma-\sigma}\frac{y_t}{p_t}, \frac{\gamma+\sigma}{1+\gamma+\sigma}\frac{y_t}{p_t}]$, and $n_{i,t-1} = n_{t-1}$ is the same for each i. We have finished the proof of Lemma 1.

A.7 Proofs of Lemmas A.2–A.3

To uncover the implications of the peer effect, we introduce heterogeneity to $n_{i,t-1}$ by assuming that a proportion $\lambda \in (0,1)$ of generation-1 women are restricted by the one-shot birth-control policy to give birth to only n^p children, while the remaining $1-\lambda$ of generation-1 women are unrestricted and give birth to n_1^u children. The birth-control policy is binding: $n^p < n_1^u$. In generation 2, two sets of fertility norms emerge: Daughters with restricted mothers are exposed to a fertility norm

$$N_{r,2} = \alpha n^p + (1-\alpha)\bar{n}_2,$$

while daughters with unrestricted mothers are exposed to a fertility norm

$$N_{ur,2} = \alpha n_1^u + (1-\alpha)\bar{n}_2.$$

Because $n^p < n_1^u$, we must have $N_{r,2} < N_{ur,2}$. In the face of distinct fertility norms, individuals in generation 2 may opt for different levels of optimal fertility.

We denote the equilibrium fertility of generation-t individuals in the restricted lineage as $n^*_{r,t}$, and those in the unrestricted lineage as $n^*_{ur,t}$.

A Nash equilibrium is sustained by a condition of rational expectation:

$$\bar{n}_t = \lambda n^*_{r,t} + (1-\lambda)n^*_{ur,t}.$$

That is to say, individuals optimize by solving for P2 in expectation of the optimal fertility choices of others. In the Nash equilibrium, no individual has an incentive to deviate given the optimal fertility of others.

We obtain the model solution $(n^*_{r,2}, n^*_{ur,2})$ by solving for P2 of both restricted and unrestricted lineages under the condition of rational expectation. As in previous derivations, we define $n^l_t = \frac{\gamma-\sigma}{1+\gamma-\sigma}\frac{y_t}{p_t}$ and $n^u_t = \frac{\gamma+\sigma}{1+\gamma+\sigma}\frac{y_t}{p_t}$. Because $N_{r,2} < N_{ur,2}$, we discuss four cases, as follows.

Case 1. Suppose that the fertility norm of individuals with a restricted mother, $N_{r,2}$, lies between the lower and upper bounds ($n^l_2 \leq N_{r,2} \leq n^u_2$), while the fertility norm of individuals with an unrestricted mother, $N_{ur,2}$, lies above the upper bound ($N_{ur,2} > n^u_2$). By a reasoning similar to Proposition 1, individual optimization gives

$$n^*_{r,2} = N_{r,2},$$

$$n^*_{ur,2} = n^u_2.$$

The rational expectation condition further requires

$$\begin{aligned} \bar{n}_2 &= \lambda N_{r,2} + (1-\lambda)n^u_2 \\ &= \lambda(\alpha n^p + (1-\alpha)\bar{n}_2) + (1-\lambda)n^u_2 \\ \Rightarrow \bar{n}_2 &= \frac{\alpha\lambda_2 n^p + (1-\lambda)n^u_2}{1-\lambda+\alpha\lambda}. \end{aligned}$$

Substituting the above expression of $\bar{n}_2$ into $N_{r,2}$ and $N_{ur,2}$ in the inequalities $n_2^l \leq N_{r,2} < n_2^u$ and $N_{ur,2} > n_2^u$, we obtain a parameter restriction

$$\lambda < \min\{\frac{n_1^u - n_2^u}{(1-\alpha)(n_1^u - n^p)}, 1 - \frac{\alpha(n_2^l - n^p)}{(1-\alpha)(n_2^u - n_2^l)}\}.$$

Case 2. Suppose that the fertility norm of individuals with a restricted mother, $N_{r,2}$, lies between the lower and upper bounds ($n_2^l \leq N_{r,2} \leq n_2^u$), while the fertility norm of individuals with an unrestricted mother, $N_{ur,2}$, also lies between the lower and upper bounds ($n_2^l < N_{ur,2} < n_2^u$). Using a logic similar to that in Proposition 1, individual optimization gives

$$n_{r,2}^* = N_{r,2},$$
$$n_{ur,2}^* = N_{ur,2}.$$

The rational expectation condition requires

$$\begin{aligned}
\bar{n}_2 &= \lambda N_{r,2} + (1-\lambda)N_{ur,2} \\
&= \lambda(\alpha n^p + (1-\alpha)\bar{n}_2) + (1-\lambda)(\alpha n_{ur,1} + (1-\alpha)\bar{n}_2) \\
\Rightarrow \bar{n}_2 &= \lambda n^p + (1-\lambda)n_{ur,1}.
\end{aligned}$$

Substituting the above expression of $\bar{n}_2$ into $N_{r,2}$ and $N_{ur,2}$ in the inequalities $n_2^l \leq N_{r,2} < n_2^u$ and $n_2^l \leq N_{ur,2} < n_2^u$, we obtain parameter restrictions

$$\begin{aligned}
&\alpha \in (0, \frac{n_2^u - n_2^l}{n_1^u - n^p}), \\
&\lambda \in (\frac{n_1^u - n_2^u}{(1-\alpha)(n_1^u - n^p)}, 1 - \frac{n_2^l - n^p}{(1-\alpha)(n_1^u - n^p)}).
\end{aligned}$$

Case 3. Suppose that the fertility norm of individuals with a restricted mother, $N_{r,2}$, lies below the lower bound ($N_{r,2} \leq n_2^l$), while the fertility norm of individuals with an unrestricted mother, $N_{ur,2}$, lies above the upper bounds ($N_{ur,2} > n_2^u$). Individual optimization gives

$$n_{r,2}^* = n_2^l,$$
$$n_{ur,2}^* = n_2^u.$$

The rational expectation condition requires

$$\bar{n}_2 = \lambda n_2^l + (1-\lambda) n_2^u$$

Substituting the above expression of $\bar{n}_2$ into $N_{r,2}$ and $N_{ur,2}$ in the inequalities $N_{r,2} \leq n_2^l$ and $N_{ur,2} \geq n_2^u$, we obtain a parameter restriction

$$\alpha \in [\frac{n_2^u - n_2^l}{n_1^u - n^p}, 1)$$
$$\lambda \in [1 - \frac{\alpha(n_2^l - n^p)}{(1-\alpha)(n_2^u - n_2^l)}, \frac{\alpha(n_1^u - n_2^u)}{(1-\alpha)(n_2^u - n_2^l)}].$$

Case 4. Suppose that the fertility norm of individuals with a restricted mother, $N_{r,2}$, lies below the lower ($N_{r,2} \leq n_2^l$), while the fertility norm of individuals with an unrestricted mother, $N_{ur,2}$, lies between the lower and upper bounds ($n_2^l < N_{ur,2} < n_2^u$). Individual optimization gives

$$n^*_{r,2} = n^l_2,$$

$$n^*_{ur,2} = N_{ur,2}.$$

The rational expectation condition requires

$$\begin{aligned}\bar{n}_2 &= \lambda n^l_2 + (1-\lambda)N_{ur,2} \\ &= \lambda n^l_2 + (1-\lambda)(\lambda n_{ur,1} + (1-\lambda)\bar{n}_2) \\ \Rightarrow \bar{n}_2 &= \frac{\lambda n^l_2 + (1-\lambda)\alpha n^u_1}{\lambda + \alpha - \lambda\alpha}.\end{aligned}$$

Substituting the above expression of $\bar{n}_2$ into $N_{r,2}$ and $N_{ur,2}$ in the inequalities $N_{r,2} \leq n^l_2$ and $n^l_2 \leq N_{ur,2} \leq n^u_2$, we obtain parameter restrictions

$$\lambda > \max\{1 - \frac{n^l_2 - n^p}{(1-\alpha)(n^u_1 - n^p)}, \frac{\alpha(n^u_1 - n^u_2)}{(1-\alpha)(n^u_2 - n^l_2)}\}.$$

Case 5. Suppose that the fertility norm of individuals with both restricted and unrestricted mothers lies above the upper bound ($N_{ur,2} > N_{r,2} > n^u_2$). Individual optimization gives

$$n^*_{r,2} = n^u_2,$$

$$n^*_{ur,2} = n^u_2.$$

The rational expectation condition requires

$$\bar{n}_2 = \lambda n^u_2 + (1-\lambda)n^u_2 = n^u_2.$$

Substituting the above expression of $\bar{n}_2$ into $N_{r,2}$ and $N_{ur,2}$ in the inequalities $N_{ur,2} > N_{r,2} > n_2^u$, we have

$$N_{r,2} = \alpha n^p + (1-\alpha)n_2^u > n_2^u,$$

which contradicts that $n^p < n_2^u$. Hence, this case cannot establish a fertility equilibrium.

Case 6. Suppose that the fertility norm of individuals with both restricted and unrestricted mothers lies below the lower bound ($n_2^l > N_{ur,2} > N_{r,2}$). Individual optimization gives

$$n_{r,2}^* = n_2^l,$$
$$n_{ur,2}^* = n_2^l.$$

The rational expectation condition requires

$$\bar{n}_2 = \lambda n_2^l + (1-\lambda)n_2^l = n_2^l.$$

Substituting the above expression of $\bar{n}_2$ into $N_{r,2}$ and $N_{ur,2}$ in the inequalities $n_2^l > N_{ur,2} > N_{r,2}$, we have

$$N_{ur,2} = \alpha n_1^u + (1-\alpha)n_2^l < n_2^l,$$

which contradicts that $n_1^u > n_2^l$. Hence, this case cannot establish a fertility equilibrium.

Summing up the above case-by-case discussions, we reach Lemmas 2 and 3 in Appendix Section A.5.

A.8 Proofs of Lemmas A.4–A.7

We further conduct comparative analyses of the equilibrium fertility.

First, we discuss the impacts of λ on the average fertility in generation 2. The equilibrium fertility when $\alpha \in (0, \frac{n_2^u - n_2^l}{n_1^u - n^p})$ is summarized in Lemma 2. Within each of the three λ ranges, the average fertility strictly decreases as λ increases. Furthermore, when comparing the three ranges of λ, the smaller ranges consistently yield strictly higher average fertility values. Therefore, the average fertility in generation 2 strictly decreases as λ increases within $(0, 1)$. The equilibrium fertility when $\alpha \in [\frac{n_2^u - n_2^l}{n_1^u - n^p}, 1)$ is summarized in Lemma 3. Similarly, we can again obtain that the average fertility in generation 2 strictly decreases as λ increases within $(0, 1)$. In summary, we have $\frac{\partial \bar{n}_2}{\partial \lambda} < 0$. Furthermore, as the average fertility strictly decreases with increasing λ, this implies that the fertility norms of individuals both with and without restricted mothers also strictly decrease as λ rises.

We further discuss how the equilibrium fertility of individuals with and without restricted mothers changes as λ increases. First, the equilibrium fertility of individuals with restricted mothers strictly decreases with increasing λ until it reaches the cutoff value $\underline{\lambda}$, at which point $n_{r,2}^*$ hits the lower bound n_2^l. Further increases in λ have no impact on $n_{r,2}^*$. Specifically, as shown in Lemma 2, when $\alpha \in (0, \frac{n_2^u - n_2^l}{n_1^u - n^p})$, $\underline{\lambda} = 1 - \frac{n_2^l - n^p}{(1-\alpha)(n_1^u - n^p)}$. As shown in Lemma 3, when $\alpha \in [\frac{n_2^u - n_2^l}{n_1^u - n^p}, 1)$, $\underline{\lambda} = 1 - \frac{\alpha(n_2^l - n^p)}{(1-\alpha)(n_2^u - n_2^l)}$. In summary, $\underline{\lambda}$ increases as α decreases or as n^p increases. Second, the equilibrium fertility of individuals with unrestricted mothers initially remains at the upper bound n_2^u as λ increases from 0. When λ reaches the cutoff value $\bar{\lambda}$, then $n_{ur,2}^*$ begins to strictly decrease as λ further increases. When $\alpha \in (0, \frac{n_2^u - n_2^l}{n_1^u - n^p})$, $\bar{\lambda} = \frac{n_1^u - n_2^u}{(1-\alpha)(n_1^u - n^p)}$. When $\alpha \in [\frac{n_2^u - n_2^l}{n_1^u - n^p}, 1)$, $\bar{\lambda} = \frac{\alpha(n_1^u - n_2^u)}{(1-\alpha)(n_2^u - n_2^l)}$. In summary, $\bar{\lambda}$ increases as α increases or as the gap between the upper bounds in period 1 and period 2 widens.

Second, we analyze the impact of α on the gap between the equilibrium fertility. As α increases, which indicates a greater influence of parental fertility on fertility norms, $N_{r,2}$ decreases while $N_{ur,2}$ increases. This results in a larger gap in the equilibrium fertility between individuals with restricted mothers and unrestricted mothers, i.e., $\frac{\partial(n^*_{ur,2}-n^*_{r,2})}{\partial\alpha} \geq 0$.

In subsequent generations, the optimal fertility levels of all individuals, whether they have restricted ancestors or not, tend to converge due to the peer effect. Therefore, we have $\bar{n}_t < n^u_t$, which suggests that even though only a λ ($\lambda \in (0,1)$) proportion of women were subject to a one-shot birth-control policy in generation 1, the equilibrium fertility for all subsequent generations decreases. Moreover, as $\bar{n}_t = \lambda_t N_{r,t} + (1-\lambda_t)N_{ur,t}$, $\frac{\partial\lambda_t}{\partial\lambda} > 0$, $\frac{\partial N_{r,t}}{\partial\lambda} < 0$, and $\frac{\partial N_{ur,t}}{\partial\lambda} < 0$; therefore, we have $\frac{\partial\bar{n}_t}{\partial\lambda} < 0$, which verifies that the equilibrium fertility of subsequent generations decreases as λ increases.

A.9 Model Extension: Quadratic Utility Function

Our theoretical conclusions do not rely on the choice of functional form of the utility function. Our theoretical conclusions are based on the general solution form rather than the specific expressions of upper and lower bounds. Thus, changing the utility function without altering the specific expressions of the upper and lower bounds will not affect the deduction of our conclusions.

For example, our conclusion is also robust to this quadratic utility functional form following Akerlof (1997) and Spolaore and Wacziarg (2022). At time t, generation t is in adulthood, when a woman consumes c_t and rears n_t children. The utility maximization problem of generation t can be simplified to

$$U_t = -an_t^2 + bn_t + c - \sigma|n_t - n_{t-1}|, \tag{P1'}$$

where a represents the intrinsic cost of childbearing and b represents the intrinsic benefits of childbearing. Generation t would lose σ in utility if her fertility n_t deviates from her mother's fertility n_{t-1} by one point.

When $\sigma \geq 0$ and conditional on n_{t-1}, the optimal fertility of generation t is

$$n_t^* = \begin{cases} n_t^u, & \text{if} \quad n_{t-1} > n_t^u, \\ n_{t-1}, & \text{if} \quad n_t^l \leq n_{t-1} \leq n_t^u, \\ n_t^l, & \text{if} \quad n_{t-1} < n_t^l, \end{cases}$$

where $[n_t^l, n_t^u] = [\frac{b-d}{2\sigma}, \frac{b+d}{2\sigma}]$.

The only difference between this solution and Proposition 1 is the specific expression of the upper bound n_t^u and the lower bound n_t^l. The general form of the optimal fertility for generation t is consistent with Proposition 1.

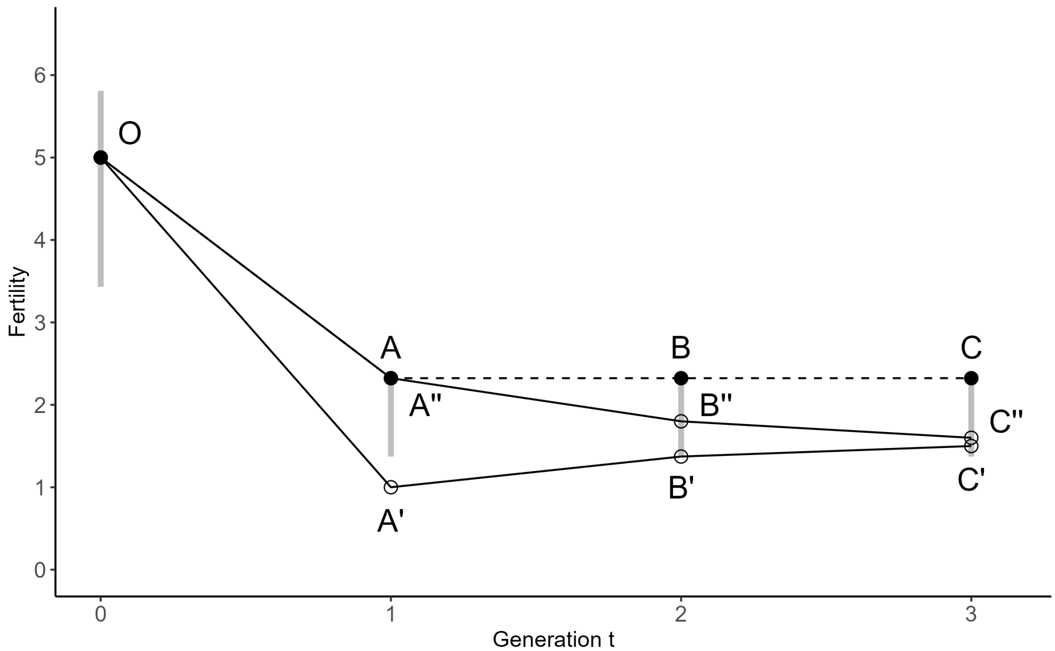


Figure A1: Theoretical Predictions for the Intergenerational Effects of One-shot Birth-control Policies for Generation 1, under the Peer Effect. The x-axis denotes generations, the y-axis fertility levels, and vertical gray lines represent the fertility interval $[n_t^l, n_t^u]$ for each generation t. The dashed line denotes a path of fertility transition when the birth-control policy never comes into existence. Solid lines depict the fertility transition paths for two lineages when a one-shot birth-control policy restricts fertility for specific lineages in generation 1. The solid line depicts the fertility transition paths when a compulsory birth-control policy restricts fertility for a segment of generation 1, but not for other generations. The solid lines trace fertility trends for women with first-generation ancestors treated (OA'B'C') and untreated (OA"B"C").

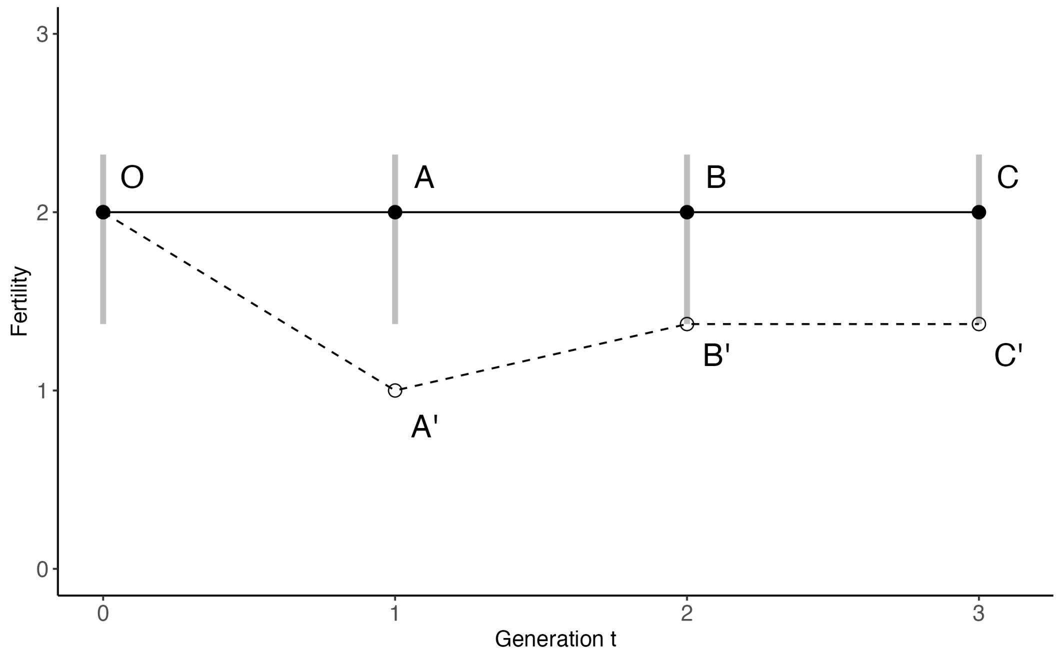


Figure A2: A "Stable Cost" Scenario. This figure demonstrates the lasting effect of a one-shot birth-control policy at generation 1, when the cost-income ratio $\frac{p_t}{y_t}$ remains the same across generations. The x-axis denotes generations, the y-axis shows fertility levels, and vertical gray lines represent the fertility interval $[n_t^l, n_t^u]$ for each generation t. The solid OABC line denotes the fertility trend when the birth-control policy never comes into existence. The dashed OA'B'C' line depicts the fertility trend when a compulsory birth-control policy restricts fertility for generation 1, but not for subsequent generations.

B Further Details on the Institutional Background and Data

B.1 Before Compulsory OCP: "Award One, Punish Three"

Before implementation of the compulsory OCP—China's national fertility system, commonly known as "Award One, Punish Three"—bore similarities to policies in Singapore and certain other countries. In October 1978, the central government articulated the principle that "one child is optimal for a family, two children are sufficient, and three children are prohibited," as documented in the "Report on the First Meeting of the State Council Leading Group on Family Planning." Essentially, couples who comitted to cease childbearing after one child received a set of rewards as their child grew up. Economic penalties were imposed only on couples with more than two children.

The Award One, Punish Three policy proved ineffective during 1979-1982. For instance, in 1981, 30.9% of mothers aged between 20 and 45 with one child in Guangdong—where only the Award One, Punish Three policy rather than the compulsory OCP was implemented—successfully had another baby during 1979-1982. In contrast, only 9.1% of mothers with one child in Liaoning, where the compulsory OCP was enforced, managed to give birth to a second child during the same period.[3] Numerous couples signed the pledge in advance to receive the reward, but then broke the contract to have a second child. Banister (1984) cites a 1981 scholarly report from Wuhan that assessed the national situation: A considerable number of those who received the certificate and the award put it aside for safekeeping. They did not dare spend it, because they planned to return it when they had a second child.[4]

There are at least two reasons for the failure of Award One, Punish Three. On the one hand, there were no severe penalties for families who broke the contract and had a second child. On the other hand, couples were eager to have more children as a means of investing

[3]While provincial governments had the authority to formulate regulations based on local social, economic, political, and cultural conditions, a small proportion of provinces, including Liaoning, opted to impose severe penalties on unauthorized second-order births instead of adhering solely to the Award One, Punish Three principle. Data are calculated using the 1982 census.

[4]Cheng Du. "China's population, some problems and their solutions." Wuhan daxuexuebao, shehui kexue ban (Wuhan University Journal, Social Sciences Edition), No. 3 (1981), p. 8.

in future support when they would be older, since the social pension system had not yet been developed.[5]

B.2 Comparisons with Alternative Measures of the OCP

Our measure of OCP timing builds on the pioneering work of Ebenstein (2010); Edlund *et al.* (2013); and Huang (2022). Ebenstein (2010) is the first to construct a province-level measure of the strength of OCP implementation. Based on Scharping's (2003) comprehensive volume on birth-control policies in China, Ebenstein (2010) compiles a province-year panel of birth-control fines levied on above-quota births in 1979–2000. The fines are measured as multiples of annual household income to facilitate cross-province and over-year comparisons. Ebenstein (2010)'s seminal work paves the way for studies that examine the socioeconomic consequences of OCP implementation (García, 2022; Huang *et al.*, 2021, 2023; Guo *et al.*, 2025). As pointed out by Huang *et al.* (2021), variations in fines primarily capture the escalating strength of the OCP around 1990. Complementary to the fines measure, our measure of OCP timing documents the initial rollout of the OCP in the late 1970s and early 1980s.

Edlund *et al.* (2013) are the first to construct measures that capture the timing of birth-control policies in Chinese provinces. Specifically, drawing on Peng's (1996) *Encyclopedia of China's Family Planning Program*, Edlund *et al.* (2013) compile the launch years of three programs: (1) family-planning science and technology research institutes; (2) family-planning education centers; and (3) family-planning associations. The three programs reflect the broader enforcement strength of China's birth-control policies, including policy phases

[5] Our model partially accounts for this pattern, since the intrinsic benefits and costs of childbearing for first-generation women remain unchanged, which renders it challenging to alter the fertility level.

that preceded the explicit mandate of one child per couple. Different from Edlund *et al.* (2013), our timing measure exclusively reflects the year in which each province implemented the one child per couple mandate and imposed strict penalties on a second birth.

Huang (2022) introduces an innovative continuous measure of OCP coverage for rural residents that combines policy timing with four types of policy exemptions, such as those for ethnic minorities. Different from Huang (2022), our measures uniquely reflect the mandatory one child per couple policy that applied to all urban Han Chinese before the 2014 policy relaxation. This focus on urban Han Chinese better serves our purpose, because a stable policy environment before 2014 offers minimal scope for cross-cohort variations in early-life OCP exposure among second-generation women—a potential confounder in our identification strategy. Upon thorough review of the OCP literature, we are surprised to find that we are the first to leverage the phased rollout of the OCP within an event-study framework.

B.3 The Immediate Effect of the OCP Rollout

We use total fertility rates in each Chinese province before 1982, as computed by Coale and Chen (1987), to investigate the immediate impact of the OCP rollout.[6] We further drop the period before 1972 to avoid the impact of the Later, Longer, Fewer Campaign. We employ the following standard specification to conduct the event study:

$$Y_{p,t} = \sum_{l \neq -1, l \geq -3} \beta_l \mathbb{I}\{t - T_p = l\} + \theta_t + \alpha_p + \epsilon_{p,t}, \tag{B1}$$

where $Y_{p,t}$ is the variable that indicates the total fertility rate in province p during year t, and T_p denotes the year province p implemented the OCP. We include year fixed effects (θ_t) and

[6]Coale and Chen (1987) only calculate the total fertility rates of each province before 1982.

province fixed effects (α_p). We also control for predetermined economic and demographic conditions, and allow for nonparametric trends for provinces with different conditions.

Fig. B1 demonstrates the impact of the OCP on period fertility rate. The insignificant coefficients for $\beta_l < 0$ demonstrate that the fertility rate is consistent across provinces before the rollout of the OCP, conditional on the covariates. The significantly negative coefficients for $\beta_l \geq 0$ show that the total fertility rate undergoes a sharp decline 1 year after implementation of the OCP. Notably, the policy effect is subject to a 1-year delay, owing to the mandatory 9-month gestation period for pregnancies. Mothers who were already pregnant in the year of policy implementation are less likely to be affected and can proceed with their planned childbirth. Moreover, the treatment effect remains consistent for both relative year 1 and relative year 2, which indicates the stability of policy intensity during that period.

B.4 Determinants of OCP Rollout

We examine the potential correlation between the year of OCP rollout and predetermined demographic, economic, and political factors. We collect data on potential correlated factors in each province and investigate their correlation with the time of OCP implementation.

First, the demographic status in one province may be correlated with OCP implementation time. For example, a rising birth rate may increase pressure on these provincial governments to implement the reform. Conversely, provinces with a larger population and shared culture may resist implementing the OCP and thus implement later. Therefore, we include total population, urban and rural population, and birth and death rates from the previous year as independent variables separately.

Second, provinces with more aggressive implementation of the OCP may have also actively promoted economic development. Therefore, we use the GDP growth rate of each province to proxy for economic development and the share of primary and secondary industry in GDP to proxy for the local economic structure. Also, we use the share of fiscal expenditure in GDP to proxy for the behavior of local provincial governments.

Third, provinces with closer political ties to the central government may implement the OCP earlier. We use the distance from Beijing to proxy for this factor.

We regress the dummy variable indicating *whether the province implemented mandatory OCP in a particular year* on lagged demographic and socioeconomic variables. We also control for year and province fixed effects, along with province-specific linear trends. We restrict the macro-level data from 1971 to 1990 to fully capture potential correlations. Panels A and B of Table **??** reveal that the effects of demographic and socioeconomic variables are small and statistically insignificant. In Panel C, we regress the OCP implementation year in each province on the geographic distance between the provincial capital and Beijing, with controls for demographic and socioeconomic factors. As expected, distance positively and significantly predicts the implementation year, indicating that provinces farther from Beijing tend to implement the OCP later. Consistent with the historiographic narrative of Peng (1996) and Scharping (2013), our findings suggest that the timing of the rollout was influenced by political decisions at the provincial level.

B.5 Sampling Design of the Data Sources

Our analysis primarily relies on data from the 2020 Population Census, the mini-census surveys, and the China Family Panel Studies (CFPS), all of which employ stratified sampling designs to ensure national representativeness.

Census and Mini-census: The National Bureau of Statistics (NBS) of China conducts a population census every 10 years and a 1% population survey (mini-census) midway between censuses; This practice was implemented in 1990. Our main dataset is drawn from a 1‰ random sample of the 2020 census. For robustness and mechanism analyses, we also use a 20% random sample from the 2005 mini-census, a 0.35% random sample from the 2010 census, and a 20% random sample from the 2015 mini-census. To ensure representativeness, these random samples incorporate urban-rural stratification, which guarantees coverage proportional to both urban and rural populations. The mini-census sampling design further employs stratified, cluster, and probability-proportional-to-size (PPS) sampling methods based on the census sampling frame, with urban-rural stratification preserved to maintain national representativeness.

CFPS Survey: The CFPS employs a multi-stage, stratified probability sampling design that covers 25 provinces, municipalities, and autonomous regions, which together represent approximately 94.5% of mainland China's population. The sampling procedure is stratified by geography and urban-rural status to capture regional economic and demographic heterogeneity. Households are then randomly selected within sampled villages or communities, which ensures adequate representation of both urban and rural populations.

Ethnicity (Han) Representation: Although ethnicity is not used as a stratification variable—given the small population sizes of many minority groups and the resulting infeasibility of such stratification—the large and randomly drawn samples provide proportional representation of the Han population. We confirmed this by comparing the Han population share in our samples with official statistics from the NBS (91.1%), and found no significant discrepancies.[7] Thus, our restriction to urban Han women yields a sample that remains broadly representative of the national Han urban population.

B.6 Details on the CFPS Sample

We first pool six waves of the CFPS and restrict the data to urban Han women born during 1969-1980 to construct the second-generation samples. We then further construct the third-generation samples using information on the parent-child relationship provided by the CFPS.[8] Table B3 Panel A reports that the desired number of children for the second generation is 1.735, while the desired fertility level for the third generation has declined to 1.532. This indicates a continuing decline in the intention to have children over time in China.

B.7 Details on the FIDS Sample

We use data from the Fertility In-Depth Survey (FIDS) conducted in 1987, which contains information on the fertility intention of first-generation women. This study is part of a program that administered in-depth surveys on the population's fertility and related factors in various provinces and municipalities of China. The 1987 survey was conducted across six

[7] Data source: $https : //www.stats.gov.cn/xxgk/sjfb/zxfb2020/202105/t20210511_1817195.html$

[8] We use the accurate household structure provided by the CFPS to identify children of the second generation (i.e, third-generation individuals). In all, mothers born between 1969 and 1980 are referred to as second-generation mothers, and children whose mothers were born between 1969 and 1980 as the third generation.

provinces: Liaoning, Beijing, Guangdong, Guizhou, Gansu, and Shandong, and represent an average across these regions.

We designate women with urban Han daughters born between 1969 and 1980 as our first-generation samples. As shown in Table B3 Panel B, the number of desired births of second-generation women is 5.108, which is higher than the number of desired births of second- and third-generation women calculated using the CFPS.

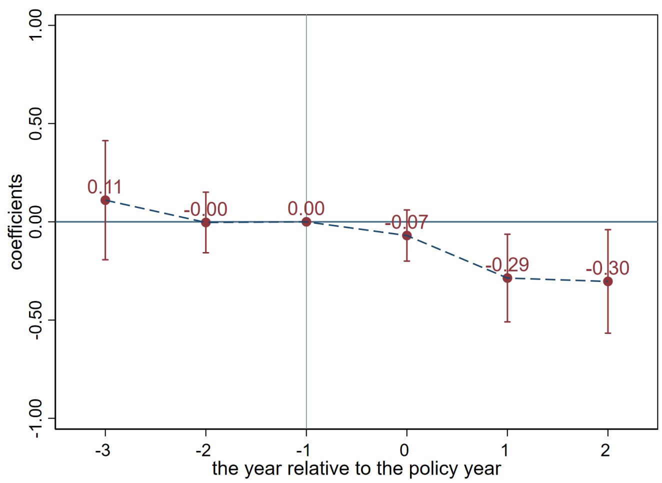


Figure B1: Immediate Impact of the OCP Rollout on TFR during 1973-1982. This figure plots the coefficients estimated using Eq. B1. The dependent variable is TFR. Data on TFR prior to 1982 are sourced from Coale and Chen (1987). The analytical period includes the years 1973 to 1982. Control variables include province fixed effects, year fixed effects, and province-specific linear trends. Robust standard errors are clustered at province level. Vertical bars represent confidence intervals at the 95% level.

Table B1: OCP Document

Provinces	Date	Document	Details
Beijing	11/01/1979	Interim Provisions of Beijing on Population and Family Planning	Encourage a couple (including one or both remarried couples) to have only one child, and give praise and reward to couples who have only one child in their lifetime. For couples who have three or more children, social support fees are levied on children who have exceeded their family planning.
	11/28/1980	**On the Revision of Interim Provisions of Beijing on Population and Family Planning on Certain Issues**	Systematically control the birth of a second child, and specify that only those who meet one of several conditions can apply for the birth of a second child: ...
Tianjin	01/27/1979	Document 6 (1979)	Encourage couples to have only one child, ...
	10/- -/1979	Document 85 (1979)	Encourage a couple (including one or both remarried couples) to have only one child, and give praise and reward to couples who have only one child in their lifetime. For couples who have three or more children, social support fees are levied on children who have exceeded their family planning.
	04/17/1981	**Decision on Encouraging Late Childbearing and Controlling the Second Birth**	Systematically control the birth of a second child, and specify that only those who meet one of several conditions can apply for the birth of a second child: ...
Hebei	07/- -/1979	Document 122 (1979)	Generally encourage a couple to have only one child during their lifetime...
	04/13/1981	**Supplementary Provisions on Certain Issues of Family Planning**	Generally mandate having fewer and better children: A couple is mandated to have only one child. Only those with several of six special circumstances can have a second child...
	05/01/1982	Interim Provisions of Hebei on Family Planning	Expand the special circumstances to have a second child from six articles to twelve articles...
Shanxi	10/01/1979	Interim Provisions of Shanxi on Family Planning	Generally encourage a couple to have only one child during their lifetime...
	04/05/1981	**Report on Family Planning in the Province in 1980 and Opinions for the Next Two Years**	All government officials and employees of state-owned enterprises are restricted to having only one child...
	12/01/1982	Provisions of Shanxi on Family Planning	Expand the control over the birth of a second child and specify that only those who meet one of several conditions can apply for the birth of a second child: ...

Provinces	Date	Document	Details
Inner Mongolia	**06/15/1982**	**Document 143 (1982)**	All state cadres and workers and urban residents are mandated to have only one child per couple.
Liaoning	06/16/1979	Document 131 (1979)	Generally encourage a couple to have only one child during their lifetime... Couples who have the ability to bear children and voluntarily have only one child will be praised and rewarded.
	04/03/1980	**Supplementary Provisions on Issues of Family Planning**	A couple is mandated to have only one child, and those who are allowed to have two children under special circumstances must be spaced at least several years apart.
	06/24/1982	Supplementary Provisions on Issues of Family Planning	Expand the special circumstances that allow having a second child...
Jilin	**09/26/1979**	**Document 299 (1979)**	A couple is mandated to have only one child, and those who are allowed to have two children under special circumstances must be spaced at least several years apart. The parents of all babies born outside the local government's birth plan each year will be penalized.
	03/22/1980	Approval of the Relaxation of the Korean Ethnicity Family Planning in Yanbian Korean Autonomous Prefecture	The number of children born to Korean families can be relaxed to two children per family.
	08/18/1984	Document 111 (1984)	Expand the special circumstances that allow having a second child...
Heilongjiang	**09/- -/1979**	**Document 263 (1979)**	Generally mandate a couple to have one child for a lifetime. In principle, the birth of a second child is no longer allowed except for some special cases ...
	- -/- -/1981	Notice on Accelerating Economic and Cultural Construction in Minority Areas	The number of children born to ethnic minorities can be relaxed to two children per family.
	- -/- -/1983	Document 10 (1983)	Expand the range of people who can have a second child, such as a rural family with one child ...
Shanghai	09/01/1979	Provisions on Family Planning	Provide incentives for parents who promised to bear one child for their lifetime ...
	08/–/1981	**Provisions of Shanghai on the Implementation of Family Planning**	Generally mandate a couple to have one child for their lifetime. In principle, the birth of a second child is no longer allowed except for some special cases ...

Provinces	Date	Document	Details
Jiangsu	07/31/1979	Document 106 (1979)	Reward couples with one child for their lifetime ...
	10/–/1979	Jiangsu Birth Planning Office Issued Document 36 (1979)	Further clarify the range of authorized births ...
	06/–/1982	**Document 86 (1982)**	Generally mandate a couple to have one child for their lifetime. In principle, the birth of a second child is no longer allowed except for some special cases ... Couples violating family planning policies by having a second child will result in wage deductions.
Zhejiang	- -/- -/1979	Document 108 (1979)	Encourage a couple (including one or both remarried couples) to have only one child, and give praise and reward to couples who have only one child in their lifetime. For couples who have three or more children, social support fees are levied on children who have exceeded their family planning.
	03/04/1982	**Interim Regulations of Zhejiang on Population and Family Planning**	Systematically control the birth of a second child, and specify that only those who meet one of several conditions can apply for the birth of a second child: ...
Anhui	04/09/1979	Document 38 (1979)	Encourage a couple (including one or both remarried couples) to have only one child, and give praise and reward to couples who have only one child in their lifetime. For couples who have three or more children, social support fees are levied on children who have exceeded their family planning.
	05/09/1981	**Interim Provisions of Anhui on Implementing Family Planning**	Systematically control the birth of a second child, and specify that only those who meet one of several conditions can apply for the birth of a second child: ...
Fujian	06/01/1979	Document 32 (1979)	Encourage a couple (including one or both remarried couples) to have only one child, and give praise and reward to couples who have only one child in their lifetime. For couples who have three or more children, social support fees are levied on children who have exceeded their family planning.
	–/–/1981	Opinions on A Number of Issues in Family Planning	"Rewarding one, restricting the interval, and not allowing three"
	03/08/1982	**Notice on Family Planning**	Systematically control the birth of a second child, and specify that only those who meet one of several conditions can apply for the birth of a second child: ...
	05/–/1982	**Provisions on Several Policies on Family Planning**	Further clarify the range of authorized second births for several special cases for rural and urban residents separately ...
Jiangxi	**04/24/1981**	**Several Provisions of Jiangxi on Family Planning (trial)**	Systematically control the birth of a second child, and specify that only those who meet one of several conditions can apply for the birth of a second child: ...
	01/18/1983	Interim Provisions of Jiangxi on Problems of Family Planning	Expand the range of people who can have a second child ...

Provinces	Date	Document	Details
Shandong	**03/31/1980**	**Interim Provisions of Shandong on Family Planning**	Systematically control the birth of a second child, and specify that only those who meet one of several conditions can apply for the birth of a second child: ...
	07/–/1982	Notice on the Issue of Controlling the Second Birth	Expand the range of people who can have a second child ...
Henan	**07/- -/1981**	**Interim Provisions of Henan on Family Planning**	Systematically control the birth of a second child, and specify that only those who meet one of several conditions can apply for the birth of a second child: ...
	06/–/1982	Opinions On the implementation of the Central Committee of the Communist Party of China and the State Council <Directive on Further Improvement of Family Planning>	Expand the range of people who can have a second child ...
Hubei	09/- -/1979	Document 140 (1979)	Encourage a couple (including one or both remarried couples) to have only one child, and give praise and reward to couples who have only one child in their lifetime. For couples who have three or more children, social support fees are levied on children who have exceeded their family planning.
	06/13/1981	**Document 82 (1981)**	Systematically control the birth of a second child, and specify that only those who meet one of several conditions can apply for the birth of a second child: ...
Hunan	05/26/1979	Document 58 (1979)	Encourage a couple (including one or both remarried couples) to have only one child, and give praise and reward to couples who have only one child in their lifetime. For couples who have three or more children, social support fees are levied on children who have exceeded their family planning.
	05/10/1982	**Provisions of Hunan on Family Planning**	Systematically control the birth of a second child, and specify that only those who meet one of several conditions can apply for the birth of a second child: ...
Guangdong	02/02/1980	Regulations of Guangdong on Population and Family Planning	Encourage a couple (including one or both remarried couples) to have only one child, and give praise and reward to couples who have only one child in their lifetime. For couples who have three or more children, social support fees are levied on children who have exceeded their family planning.
	06/01/1986	**Regulations of Guangdong on Population and Family Planning (revision)**	Systematically control the birth of a second child, and specify that only those who meet one of several conditions can apply for the birth of a second child: ...

Provinces	Date	Document	Details
Guangxi	11/06/1979	Provisions of Guangxi on Problems of Family Planning Work	Encourage a couple (including one or both remarried couples) to have only one child, and give praise and reward to couples who have only one child in their lifetime. For couples who have three or more children, social support fees are levied on children who have exceeded their family planning.
	11/25/1982	**Interim Provisions of Guangxi on Family Planning**	Systematically control the birth of a second child, and specify that only those who meet one of several conditions can apply for the birth of a second child: ...
Hainan		Implementing the Provisions of Guangdong	
Chongqing		Implementing the Provisions of Sichuan	
Sichuan	**08/04/1980**	**Document 177 (1980)**	Mandate couples who are non-agricultural employees, including national cadres, to have only one child ...
	07/–/1981	Document	Mandate all couples to have only one child...
Guizhou	06/16/1979	Interim Provisions of Guizhou on Family Planning	Encourage a couple (including one or both remarried couples) to have only one child, and give praise and reward to couples who have only one child in their lifetime. For couples who have three or more children, social support fees are levied on children who have exceeded their family planning.
	03/15/1982	**Interim Provisions of Guizhou on Family Planning (Revision)**	Systematically control the birth of second child, and specify that only those who meet one of several conditions can apply for the birth of a second child: ...
Yunnan	07/09/1979	Document 157 (1979)	Encourage a couple (including one or both remarried couples) to have only one child, and give praise and reward to couples who have only one child in their lifetime. For couples who have three or more children, social support fees are levied on children who have exceeded their family planning.
	04/17/1982	Resolution on Further Improving Family Planning	Impose fines on an unauthorized third birth...
	08/26/1984	**Party Committee of the Provincial Family Planning Commission's Opinions on Implementing Central Government's No.7 Document (1984)**	Systematically control the birth of a second child, and specify that only those who meet one of several conditions can apply for the birth of a second child: ...

Provinces	Date	Document	Details
Tibet	05/08/1992	Interim Provisions of Tibet on Family Planning	Systematically control the birth of a second child, and specify that only those who meet one of several conditions can apply for the birth of a second child: ...
Shaanxi	07/01/1979	Document 101 (1979)	Encourage a couple (including one or both remarried couples) to have only one child, and give praise and reward to couples who have only one child in their lifetime. For couples who have three or more children, social support fees are levied on children who have exceeded their family planning.
	05/01/1981	**Interim Provisions of Shaanxi on Family Planning**	Systematically control the birth of a second child, and specify that only those who meet one of several conditions can apply for the birth of a second child: ...
	10/31/1981	Supplementary Provisions of Shaanxi on Family Planning	Expand the special circumstances to have a second child ...
Gansu	07/14/1979	Document 152 (1979)	Encourage one child per family in urban areas ...
	03/16/1982	**Document 105 (1982)**	Systematically control the birth of a second child, and specify that only those who meet one of several conditions can apply for the birth of a second child: ...
Qinghai	09/–/1979	Notice of Certain Provisions on Population Planning and Family Planning in the Province (Trial Run)	Encourage one child per family in urban areas ...
	05/31/1982	**Interim Provisions of Qinghai on Family Planning**	Systematically control the birth of a second child, and specify that only those who meet one of several conditions can apply for the birth of a second child: ...
Ningxia	06/–/1980	Trial Provisions of Certain Issues of Family Planning	Encourage one child per family in urban areas ...
	08/18/1982	**Interim Regulations of Ningxia on Family Planning**	Systematically control the birth of a second child, and specify that only those who meet one of several conditions can apply for the birth of a second child: ...
Xinjiang	07/01/1988	Interim Provisions of Xinjiang on Minority Family Planning	Encourage one child per family in urban areas ...
	07/01/1992	**Regulations of Xinjiang on Population and Family Planning**	Systematically control the birth of a second child, and specify that only those who meet one of several conditions can apply for the birth of a second child: ...

Table B2: Correlations of OCP Implementation Year with Macro Indices

Dependent Variables	(1) *Whether the OCP was Implemented in Year t*	(2) *Implemented Year*
Panel A: Demographic factors		
Population	0.000	
	(0.010)	
Urban Population	0.044	
	(0.078)	
Rural Population	0.033	
	(0.078)	
Birth Rate	0.002	
	(0.005)	
Death Rate	-0.001	
	(0.015)	
Panel B: Socioeconomic factors		
GDP Growth Rate	-0.015	
	(0.026)	
Share of Primary Industry in GDP	-0.002	
	(0.008)	
Share of Secondary Industry in GDP	0.005	
	(0.005)	
Share of Fiscal Expenditure in GDP	0.004	
	(0.004)	
Panel C: Political ties		
Geographic Distance from Beijing		0.002**
		(0.001)

Note: Panels A and B present correlations between time-varying province-level demographic and socioeconomic factors and whether the OCP was implemented in year t (column (1)). Panel C relates province-level political ties (proxied by geographic distance from Beijing) to the province-specific OCP implementation year (column (2)). In Panels A and B, regressions include province fixed effects, year fixed effects, and province-specific linear time trends; the sample is restricted to 1971–1990 and independent variables are one-year lags. Robust standard errors clustered at the province level are in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table B3: Summary Statistics of CFPS and FIDS Samples

	(1)	(2)	(3)
Variables	*Mean*	*Std.Dev.*	*N*
Panel A: CFPS 2012, 2014, and 2018			
Desired number of children of second-generation women	1.741	0.541	1,591
Desired number of children of third generation	1.531	0.643	1,498
Panel B: FIDS 1987			
Desired number of children in 1987	5.108	1.516	5,114

Notes: Data for Panel A are from the China Family Panel Study (CFPS) 2012, 2014, and 2018. Samples in Panel A are restricted to urban Han women born during 1969-1980, along with their children. For Panel B, data are from the 1987 FIDS, and samples are restricted to the mothers of urban Han women born during 1969-1980.

C Additional Robustness Checks

C.1 Sample Construction: Pinning Down Fertile Cohorts

As illustrated in Section 5, we adopt a data-driven approach to inform the choice of birth cohorts following Bailey *et al.* (2021). The idea is to use the outcome variable to find the location of the kink that maximizes the within R^2. We use the 2020 census and restrict the samples to urban Han women born between 1965 and 1975, which is a period expected to manifest the kink in fertility patterns. Specifically, we define different birth years as the kink k and estimate the following equation:

$$Y_{pc} = \theta_p + D'_{pc}\rho_1 + aD'_{pc}\rho_2 + \epsilon_{pc}, \tag{C1}$$

where D'_{pc} is a vector of dummy variables, $I(1965 \leq c < k)$, $I(k \leq c < 1975)$, and k's are different kinks as previously defined. Y_{pc} is the variable that indicates the number of children ever born in 2020 and θ_p is province fixed effects.

In keeping with the results of this data-driven procedure, we find the kink is at the 1969 cohort and use this sample for our main analysis.

We also conduct sensitivity analyses of different birth cohorts. First, we expand our sample to include women born during 1970–1981. Also, we introduce a dummy variable that indicates whether the relative year equals 5 to enhance estimation efficiency. Second, we narrow our sample to women born during 1971–1980. Consistent estimates across samples of various birth cohorts, as illustrated in Fig. C9a and Fig. C9b, demonstrate the robustness of our specifications.

C.2 Parity Distribution Across Generations

We begin by documenting the birth-order distribution of the urban Han women born in 1969–1980—the second generation in our analysis. Leveraging the 1982 census, in which these women still lived with their mothers, we can create mother-child links and assign each daughter's birth order. Fig. C3 plots the shares of first-, second-, and third-or-higher-born women. For cohorts born in 1974 or earlier, the distribution is stable: Each category accounts for roughly one-third of births. Starting with the 1975 cohort, share of the third-or-higher-order births declines sharply, while that of first-borns rises correspondingly. Interestingly, the proportion of second-borns remains virtually unchanged for the 1975–1980 cohorts, when the OCP had yet to be implemented across most provinces.

Fig. C4 depicts the parity distribution among second-generation women in the 2020 census. A slight majority (52.8%) have exactly one child, 36.0% have two, while only 6.5% have three or more. It appears that more than half of second-generation women comply with the one-child norm. Notably, 4.7% are childless—a share more than triple the 1.5% observed for women born between 1940 and 1960 (Guo *et al.*, 2024).[9] The jump in childlessness suggest that remaining child-free became more socially acceptable in the post-OCP era—an issue we examine further in Section 7.6 of the main text.

[9]Since the 2020 census did not collect fertility information for unmarried women, and given the extremely low incidence of childbirth among unmarried women in China, we treat all unmarried women in our analysis as childless. Excluding these individuals, the childlessness rate among married women in our sample from the 2020 census is 2.9%.

C.3 Robustness: Omitting Age 5 at OCP

As shown in Section 5, the primary goal of our specification is to identify the second-generation women's birth cohort for which the policy had a discernible impact, based on the birth interval length experienced by their mothers.

First, we provide support for identification of the unaffected birth interval using descriptive evidence from the 2005 mini-census. Fig. C5 plots the average proportion of second-generation women who are only children (only-child rate hereafter) over their birth year relative to the OCP rollout. The only-child rate was relatively low and stable for women who were born 5 years before the OCP rollout. This indicates that the majority of second-generation women born during that time were able to have another sibling, as their parents originally intended. However, for women who were born within 4 years before the OCP, the only-child rate experienced a significant increase. This increase can be attributed to their inability to have a second sibling due to the OCP rollout and primarily signifies the existence of the 4-year natural birth interval.

Second, we provide empirical evidence by examining the distribution of birth spacing prior to the OCP rollout. We focus on mothers in provinces that implemented the OCP after 1981 and calculate the intervals between first and second births using data from the 1990 census. The solid blue line in Fig. C7 represents the distribution of birth spacing for mothers who gave birth before 1974 and were not affected by implementation of the OCP. The graph reveals that over 90% of these families were able to have another child within a 4-year interval following the birth of their first child. In contrast, the dashed red line represents the distribution of birth spacing for mothers who gave birth between 1975 and 1978, within 4 years prior to OCP implementation. The distribution indicates that no adjustments were made by these families in their fertility plans in anticipation of future fertility restrictions imposed by the OCP.

Our analyses suggest a natural birth spacing of up to 4 years between the first and second births. Mothers in provinces treated later may shorten their birth interval in anticipation of impending birth restrictions.

C.4 Robustness: "Last-treated Group"

In Section 5, we use the last-treated provinces as the control group following Lafortune *et al.* (2018); Bailey *et al.* (2021); and Goodman-Bacon (2021) by interacting the indicator of relative year with the dummy $Early_P$. We further provide the rationale behind this specification and robustness analysis.

Since OCP implementation time ranges from 1979 to 1992 over different provinces, a traditional event study fundamentally requires a timing-based data structure. However, the traditional event study is likely to suffer from the problem of multicollinearity by saturating the unit and time fixed effects (Miller, 2023; Borusyak *et al.*, 2021). Thus, the specification may be less efficient for our estimation. Following Lafortune *et al.* (2018); Bailey *et al.* (2021); Goodman-Bacon (2021); and Sun and Abraham (2021), we transmit the last-treated provinces to the never-treated provinces by restricting our samples to a relatively small time horizon. Our data structure combines variation in event dates and includes both treated and untreated units, which provides us with two sources of identification: the comparison of treated and control units and the timing of the policy.

We offer robustness analysis for the choice of last-treated groups. We use a traditional timing-based event study design to re-estimate Eq. 2. As Fig. C8 shows, we find coefficients similar to our baseline estimates, while the estimates are less efficient.

C.5 Robustness: Quadratic Trends

The range of birth cohorts for individuals in this study is intentionally kept relatively short, since we aim to minimize the impact of fecundity rate bias. We extend our samples to include 1965–1980 women's birth cohorts and include provincial quadratic trends in women's birth cohorts. As shown in Fig. C9c, estimates are also similar in magnitude to our baseline estimates.

C.6 Robustness: Continuous Approach

We adapt the method of Chen and Fang (2021) and Rossi and Xiao (2024) to build a continuous index that captures the share of a mother's fecund years subject to the OCP. The continuous measure is

$$Exposure_{pc} = \sum_{a=0}^{8} AFR_p(a) \cdot \mathbb{I}[c + a \geq T_p],$$

where p stands for province, c for second-generation cohort, and T_p for the year when the OCP was introduced in province p. The indicator $\mathbb{I}[c+a \geq T_p]$ equals one if, when the second-generation women was age a, the policy was already in force. Multiplying this indicator by the pre-policy age-specific fertility rate $AFR_p(a)$ yields the portion of a mother's fertility schedule constrained by the OCP. Because almost all additional births occur within 8 years of the previous child, we sum over $a = 0$ to $a = 8$.

We re-estimate Eq. 2 replacing the set of dummies $D^l_{cp}s$ with the continuous index $Exposure_{pc}$. Consistent with our baseline results, Table C1 shows that greater first-generation exposure lowers the sibship size of second-generation women (column 1), which in turn translates into lower fertility among those women (column 2).

C.7 Robustness: Parallel Trends in Family Background

Using the 1982 census, we examine the correlation between OCP implementation and the family background of second-generation women. We identify mother-daughter pairs based on self-reported relationships for those living together in the census data. Further restrictions are applied to include only mothers whose self-reported number of children matches the number of living children in the household. The sample is then restricted to mothers who gave birth to a daughter between 1972 and 1980, because the proportion of matched mother-daughter pairs is relatively low for those who gave birth before 1972. Table C2 presents the summary statistics for these mothers.

Using this sample, we re-estimate Eq. 2 with variables that indicate the distribution of family background as dependent variables. The estimated results are shown in Fig. C6.

C.8 Robustness: Additional Controls

As discussed at the end of Section 6.2, we will introduce additional controls that affect intrinsic benefits and costs in childbearing in detail in this section.

First, we control for four socioeconomic and policy shocks in our regression: the Great Famine (1959–1961), Cultural Revolution (1966–1976), Later, Longer, Fewer (LLF) Campaign during the 1970s, and Compulsory Education Law (CEL) (around 1986). These events may affect the intrinsic benefits of childbearing. The first event is the Great Famine, which significantly reduced people's mental health (St Clair *et al.*, 2005). Following the literature on the Great Famine, we use the historical provincial mortality rate in 1960 reported by the NBS to proxy for famine severity. The second event is the Cultural Revolution, which decreased people's trust. We follow Walder (2014) and construct a measure of exposure to the Cultural Revolution by province. We add interaction terms between birth cohort dummies and proxies to Eq. 2. We also apply the method of De Chaisemartin and D'Haultfœuille (2023) to ensure robustness. To implement this approach, we use residuals obtained by partialling out province-specific cohort linear trends as the dependent variable. Due to the lack of variation in the timing of the Great Famine (1959–1961) and the Cultural Revolution (1966–1976) across provinces, we retained all samples. Instead of excluding samples, we classified provinces into high- and low-policy-exposure groups based on relevant indicators. Also, we incorporated nonparametric trends across units with varying levels of policy intensity.

The third event is the LLF Campaign, which significantly restricted women's completed births during 1969 and may correlate with OCP implementation time. We generate a dummy variable $Longer_i$ on whether second-generation individuals are affected by the LLF Campaign. The fourth event is implementation of the CEL around 1986, which improved women's education levels and thus may decrease fertility (Cygan-Rehm and Maeder, 2013). We use information on CEL implementation in each province. We use the data and generate a dummy variable CEL_i for whether second-generation individuals are affected by the law. We also use the method proposed by De Chaisemartin and D'Haultfœuille (2023) for robustness. To apply De Chaisemartin and D'Haultfœuille (2023)'s methods, we use the residuals obtained by partialling out province-specific cohort linear trends as the dependent variable. Moreover, the sample is further restricted to women who experienced invariant policy intensity,[10] and nonparametric trends across different units with varying levels of policy intensity are also incorporated.

Second, we control for interactions between birth cohort dummies and predetermined variables. As for potential confounding of the socioeconomic conditions, we include interactions between birth cohort dummies and provincial economic/population growth rates prior to the policy during 1969-1979. Data for economic/population growth rates during 1969-1979 are collected from the *Compilation of statistical data of the 60 years of the People's Republic of China*. We also consider cultural confounding factors. With respect to cultural factors, Zhang (2019) argues that raising children for old-age support is a long-lasting cultural tradition in China, rooted in the values of filial piety and kinship-based clan culture. As such, clan culture may be highly correlated with completed births within a given family and can be transmitted over generations. If clan culture is correlated with OCP implementation time in each province, we may overestimate the intergenerational effects of first-generation fertility restrictions. Following Zhang (2019), we construct a proxy variable for clan culture using genealogy books compiled during the Ming and Qing dynasties. We group the original data into different provinces and then rescale the data using the population in each province in 1920. We take the logarithm of the variable and add interactions of the logarithm and

[10]Since most provinces adopted the LLF before our sampling period, we excluded those that adopted it during the sampling period. Similarly, as most provinces adopted the CEL after the sampling period, we excluded those that adopted it within the sampling period.

birth year to our analysis, as shown in Eq. 2. For the political factors, we use the distance from Beijing to proxy for political ties across different provinces.

The results presented in Figs C10-C11 show estimated results of including various control variables in our analysis. We find that our results remain robust after including these additional covariates. Data for control variables are collected from the *Compilation of statistical data of the 60 years of the People's Republic of China*.

C.9 Correlation Between OCP Exposure of Mothers and Early-life OCP Exposure of Daughters

Early-life exposure of the second-generation women is likely to start from the beginning of the 1990s, since the oldest cohort incorporated in our baseline analysis was born in 1969. The most significant policy change they experienced was legalization of the OCP in the 2000s and its gradual relaxation during the 2010s. These changes have relatively minor variations in policy changes in different provinces, since they are largely nationwide in scope. Thus, changes in the early-life exposure of second-generation women across different cohorts are similar for different provinces, conditional on linear trends.

Other than early-life exposure during ages 20-45, some scholars also argue that the exposure of second-generation women during ages 6-20 also matters in determining their fertility. We thus test whether the exposure for second-generation women during ages 6-20 is correlated with the initial OCP rollout using the same method. Fig. C12a shows estimated results. Again, OCP exposure does not appear to correlate over generations.

C.10 Robustness: Using Children ever Born after Policy Relaxation as Dependent Variable

We re-estimate Eq. 2 with two minor revisions: (i) We use the number of children ever born since 2014 rather than 2016, since the policy relaxation that allowed only-child respondents to have a second child began in 2014. This adjustment ensures that our analysis aligns more closely with the policy timeline and captures the relevant treatment effects associated with children born post-2014.

(ii) We further refine the sample selection based on the criterion whereby the number of children living in the surveyed household matches the self-reported number of children. Some women in our sample do not reside with their newborn children due to migration. Also, we restrict the sample to cities in which the concordance rate between the children count in the surveyed household and the self-reported number of children exceeds 85%.

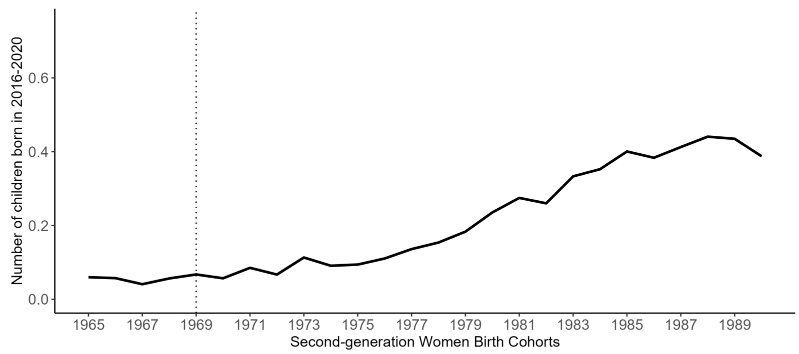
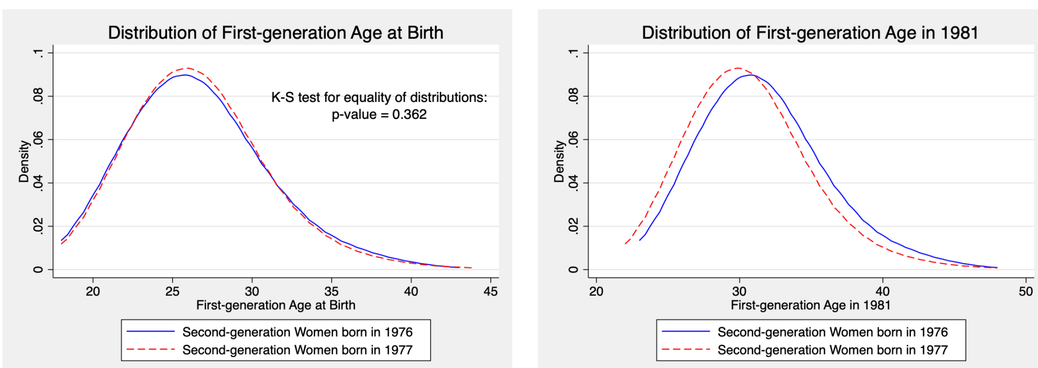


Figure C1: Change in Second-generation Fertility by Cohort between 2015 and 2020. The x-axis represents the birth cohorts of second-generation women, while the y-axis indicates the change in average fertility levels for each cohort between 2015 and 2020. The analysis is based on data from the 2020 census and 2015 mini-census. We calculated the average number of children per birth cohort for each survey year and then computed the differences in fertility across these cohorts.



(a) Distribution of First-generation Age at Birth (b) Distribution of First-generation Age in 1981

Figure C2: Distributions of First-generation Women's Age at Birth and Age at OCP implementation. The data include urban Han women born in 1976 and 1977 in the 1982 census in provinces that adopted the OCP in 1981 (the nine provinces ranging from Anhui to Tianjing in Fig. 5b). Because second-generation women were very young in 1982, they generally coreside with their mothers, allowing the creation of mother-daughter links between first and second generations. Figure (a) shows kernel density estimates of first-generation women's age at birth by second-generation birth cohort. Figure (b) shows kernel density estimates of first-generation women's age in 1981 by second-generation birth cohort.

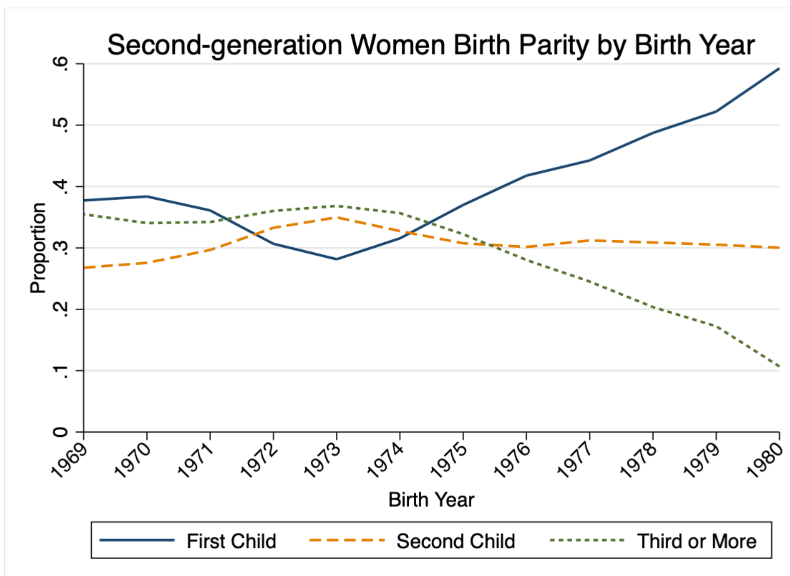


Figure C3: Trends of Second-generation Women's Birth Order. This figure shows the birth order of second-generation women across different birth cohorts in the 1982 census. The sample is restricted to urban Han women born between 1969 and 1980.

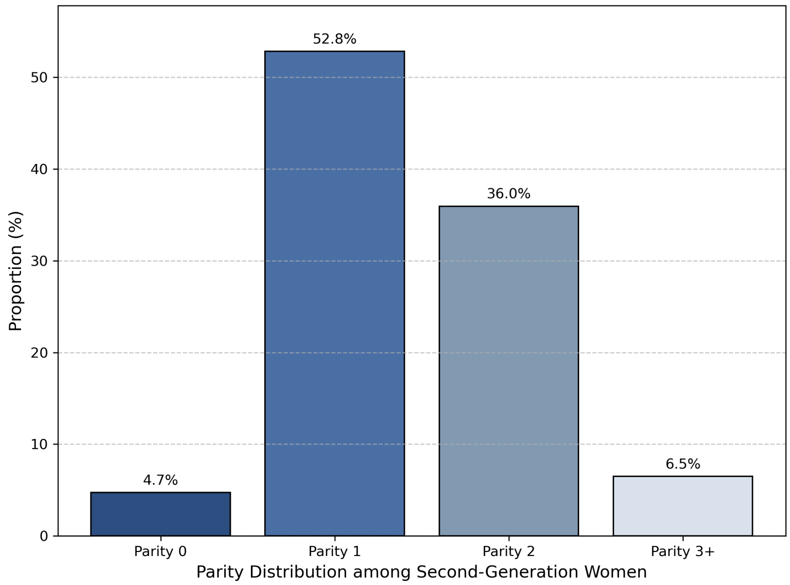


Figure C4: Parity distribution among second-generation women. This figure shows the fertility distribution of second-generation women in 2020 census. The sample is restricted to urban Han women born between 1969 and 1980.

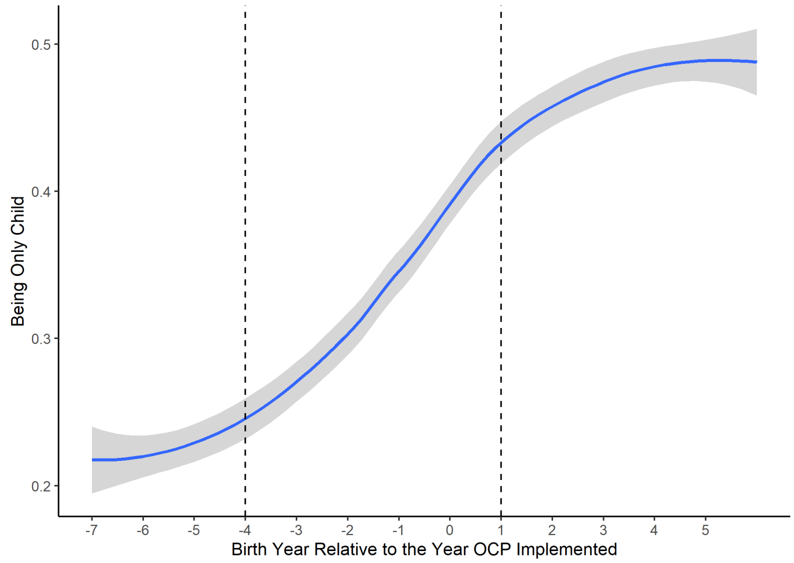
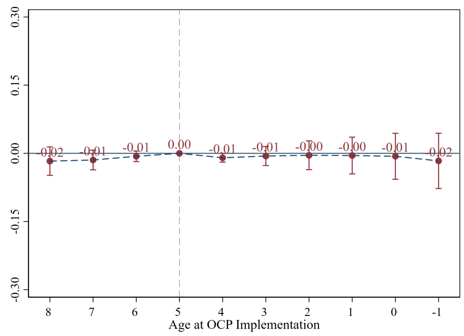
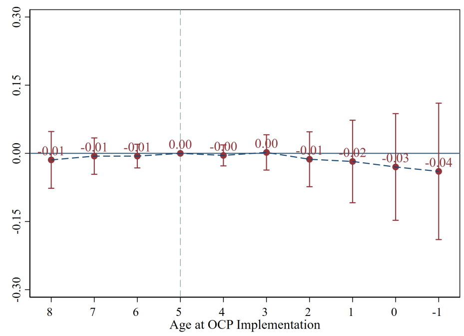


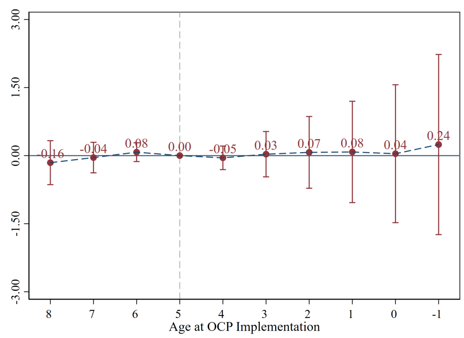
Figure C5: Only-child Rate of Second-generation Women over Years. This figure shows the trend in the probability of second-generation women being an only child over relative years, where "relative year" is defined as the birth year offset from the year of OCP implementation. Data are from the 2005 mini-census and samples are restricted to urban Han women born between 1975 and 1980. The blue line is fitted using the method of *LOESS* and the gray area represents the 95% confidence interval for the estimated smooth line.



(a) Ever Receiving Education



(b) White-collar Job



(c) Age at Birth

Figure C6: OCP Rollout and First-generation Women's Characteristics. These figures plot the coefficients estimated using Eq. 2. Dependent variables are first-generation characteristics including an indicator for ever receiving any education (Figure a), an indicator for white-collar occupation (Figure b), and age at birth (Figure c). The data are from the 1982 census, with the sample restricted to mothers who gave birth to a daughter between 1972 and 1980. Further restrictions are applied to mothers for whom the self-reported number of children matches the number of living children in the household. Since *Hukou* status is not available in the 1982 census, the sample is further restricted to mothers participating in non-agricultural work, using employment status as a proxy. Control variables in both figures include province fixed effects, cohort fixed effects, and province-specific cohort linear trends. Robust standard errors in both figures are clustered at province level. Vertical bars represent confidence intervals at the 95% level.

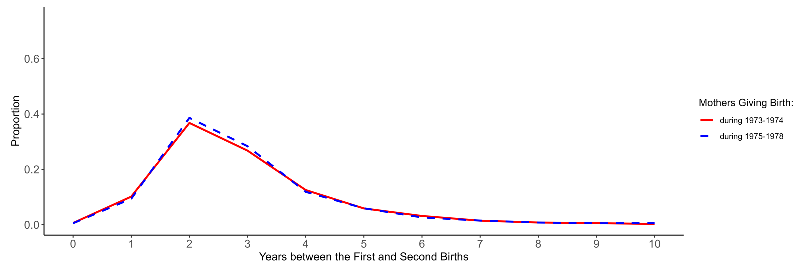


Figure C7: Distribution of Birth Spacing between First and Second Birth before the OCP Rollout. This figure shows the distribution of birth spacing between first and second birth in provinces that implemented the OCP post-1981. The solid red line represents the birth spacing distribution for urban Han women with births in 1973-1974, and the dashed blue line for those with births in 1975-1978. The analysis is based on data from the 1990 census and is restricted to provinces in which the OCP was implemented after 1981. We identified the birth sequence of each mother using information on household relations.

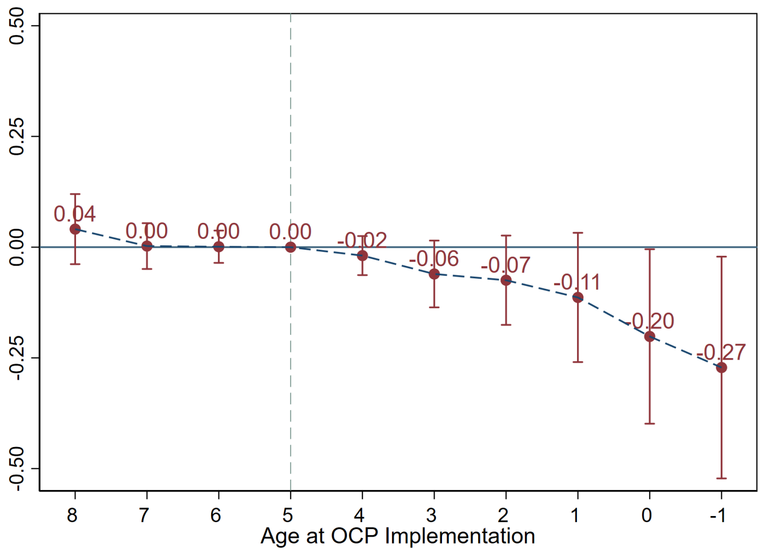
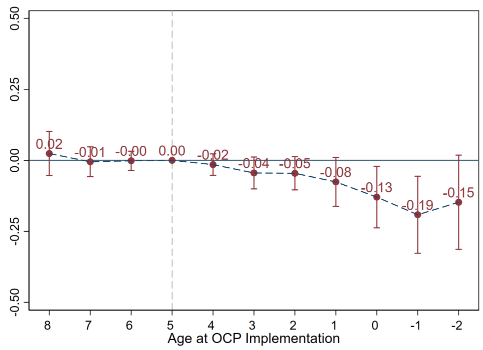
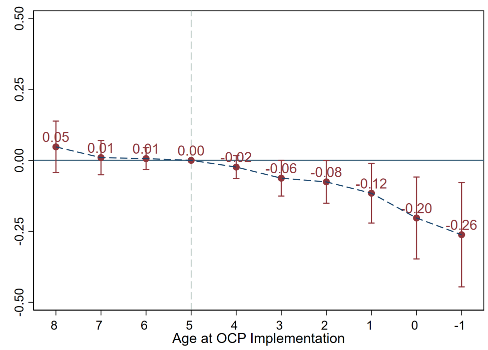


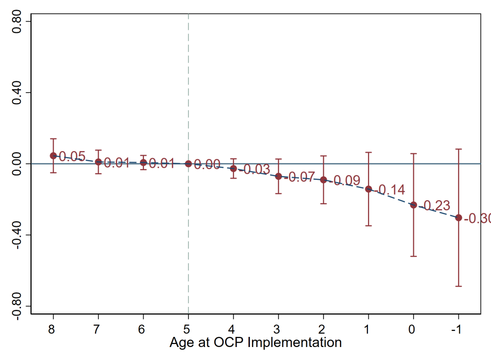
Figure C8: Impact of the OCP Rollout on Second-generation Women's Fertility in the 2020 Census: Alternative Empirical Specifications. This figure plots coefficients estimated using Eq. 2, eliminating $NotLastOCP_p$. Control variables in both figures include province fixed effects, cohort fixed effects, and province-specific cohort linear trends. Robust standard errors in both figures are clustered at province level. Vertical bars represent confidence intervals at the 95% level.



(a) Using Samples of 1969 - 1981 Birth Cohorts

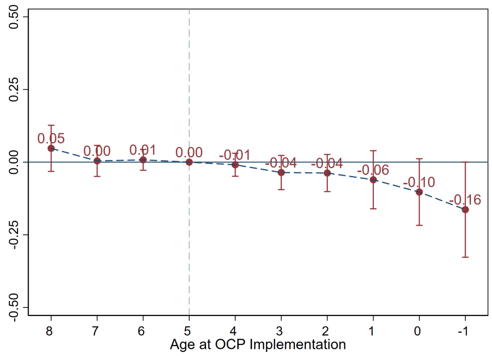


(b) Using Samples of 1971 - 1980 Birth Cohorts

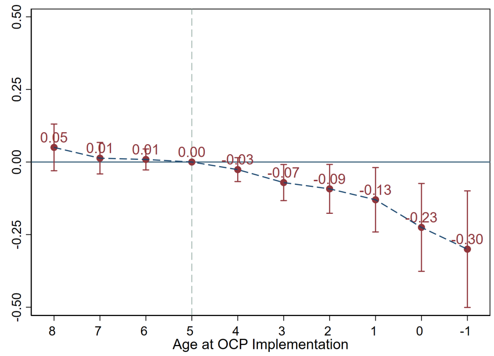


(c) Using Samples of 1965-1980 Birth Cohorts and Controlling for Quadratic Province-specific Cohort Trends

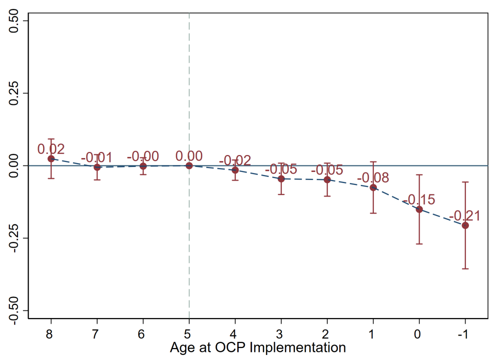
Figure C9: Impact of the OCP Rollout on Second-generation Fertility in the 2020 Census: Alternative Cohort Intervals. These figures all plot coefficients estimated using Eq. 2. The dependent variable is the number of children ever born by 2020. Figures (a)-(c) show the coefficients estimated using samples restricted to alternative cohort intervals. Data are from the 2020 census. Samples are restricted to urban Han women, with each figure representing different birth cohorts: those born between 1969 and 1981 in Figure (a), between 1971 and 1980 in Figure (b), and between 1965 and 1980 in Figure (c). Control variables in all figures include province fixed effects, cohort fixed effects, and province-specific cohort linear trends. Specifically, in Figure (c) we further control for province-specific cohort quadratic trends. Robust standard errors in Figures (a) and (b) are clustered at province level, while in Figure (c) standard errors are computed using the wild bootstrap method with 10,000 repetitions and also clustered at province level.



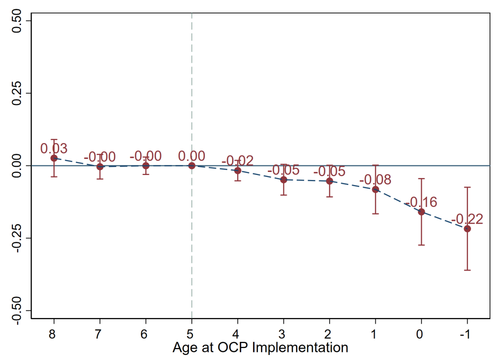
(a) Controlling for Great Famine Proxies



(b) Controlling for Cultural Revolution Proxies

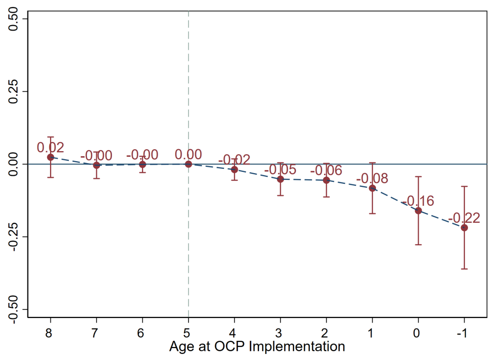


(c) Controlling for Later, Longer, Fewer Proxies

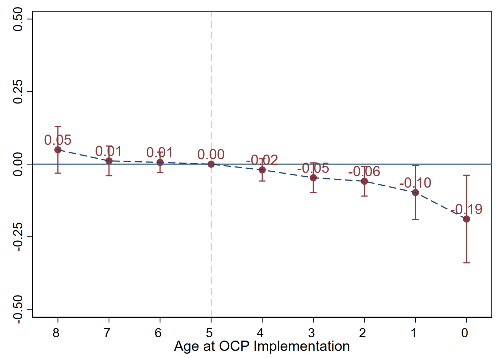


(d) Controlling for Compulsory Education Law Proxies

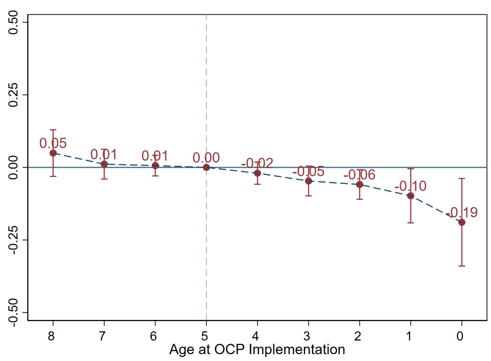
Figure C10: Impact of the OCP Rollout on Second-generation Fertility in the 2020 Census: Controlling for Policy Shocks. Figures (a)-(d) plot coefficients that are estimated using Eq. 2, controlling for different policy shocks by adding proxies that correspond to the birth cohorts from various provinces: Figure (a) controls for the Great Famine (1959-1961), Figure (b) for the Cultural Revolution (1966-1976), Figure (c) for the Later, Longer, Fewer Campaign during the 1970s, and Figure (d) for the Compulsory Education Law (around 1986). The dependent variable in all figures is the number of children ever born by 2020. Data are from the 2020 census, and samples are further restricted to urban women born during 1969-1980. Control variables in all figures include province fixed effects, cohort fixed effects, and province-specific cohort linear trends. Robust standard errors in all figures are clustered at province level. Vertical bars represent confidence intervals at the 95% level.



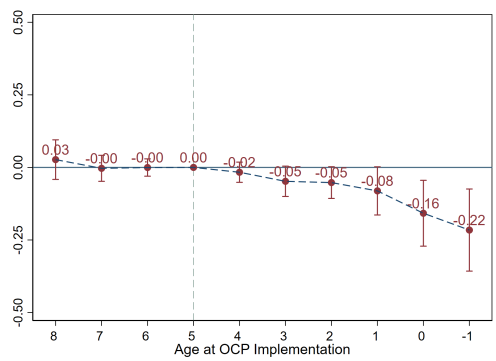
(a) Controlling for Economic Development Proxies



(b) Controlling for Clan Culture Proxies

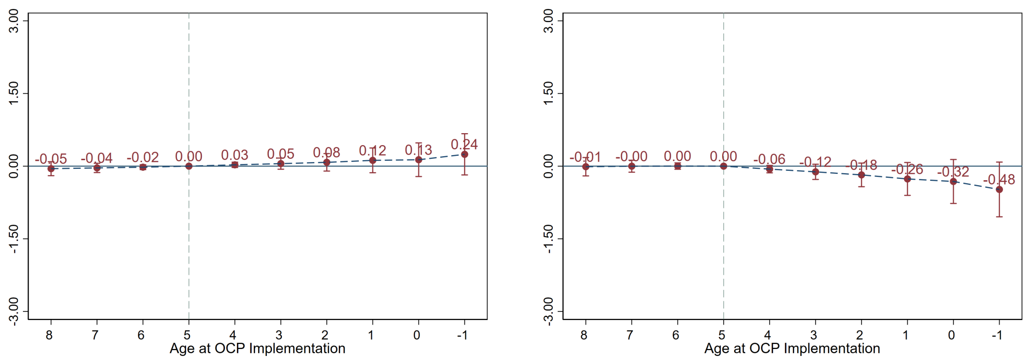


(c) Controlling for Distance from Beijing



(d) Controlling for Socioeconomic Indicators that Vary across Provinces and Cohorts

Figure C11: Impact of the OCP Rollout on Second-generation Fertility in the 2020 Census: Controlling for Pre-OCP Conditions. Figures (a)-(d) plot coefficients that are estimated using Eq. 2, controlling for interactions between the cohort dummies and a series of variables that measure different pre-OCP conditions: Figure (a) controlling for socioeconomic factors, Figure (b) cultural factors, Figure (c) political factors, and Figure (d) socioeconomic indicators that vary across provinces and cohorts. The dependent variable in all figures is the number of children ever born by 2020. Data are from the 2020 census, and samples are restricted to urban women born during 1969-1980. Control variables in all figures include province fixed effects, cohort fixed effects, and province-specific cohort linear trends. Robust standard errors in all figures are clustered at province level. Vertical bars represent confidence intervals at the 95% level.



(a) Second-generation OCP Exposure in Ages 6–20 (b) Second-generation OCP Exposure in Ages 20–45

Figure C12: Correlation Between OCP Rollout and Second-generation OCP Exposure. Figures (a) and (b) show the coefficients estimated using Eq. 2, with second-generation women's OCP exposure as the dependent variable. Figures (a) and (b) use different measures as the dependent variables: Figure (a) uses the mean fine rate for ages 6 to 20 in birth province p for cohort c, while Figure (b) uses the same measure for ages 20 to 45. Fine rate data prior to 2000 were gathered by Ebenstein (2010), while post-2000 fine rate data were compiled manually from provincial regulations. These analyses are confined to samples that represent exposure to women born between 1969 and 1980. Control variables of both figures include province fixed effects, cohort fixed effects, and province-specific cohort linear trends. Robust standard errors in both figures are clustered at province level. Vertical bars represent confidence intervals at the 95% level.

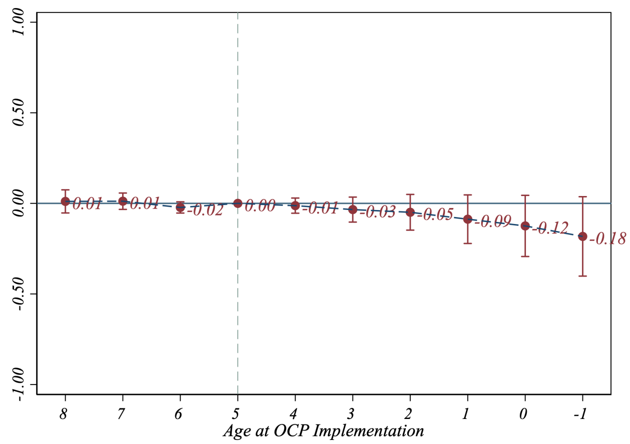


Figure C13: Impact of the OCP Rollout on Second-generation Fertility before OCP Relaxation. This figure shows the coefficients estimated using Eq. 2. The dependent variable is the number of children ever born by 2013. Data are from the 2015 mini-census and samples are restricted to urban Han women born during 1969-1980. Control variables include province fixed effects, cohort fixed effects, and province-specific cohort linear trends. Robust standard errors are clustered at province level. Vertical bars represent confidence intervals at the 95% level.

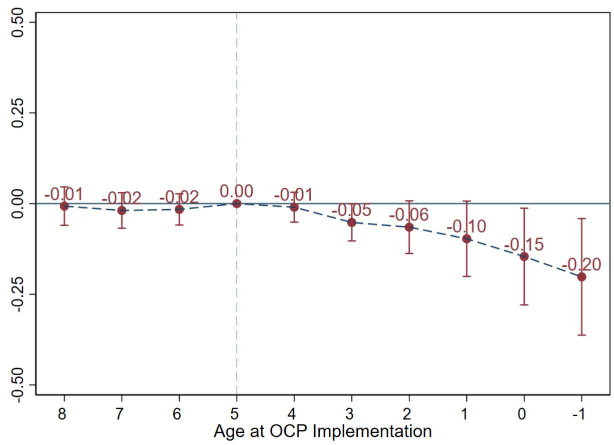
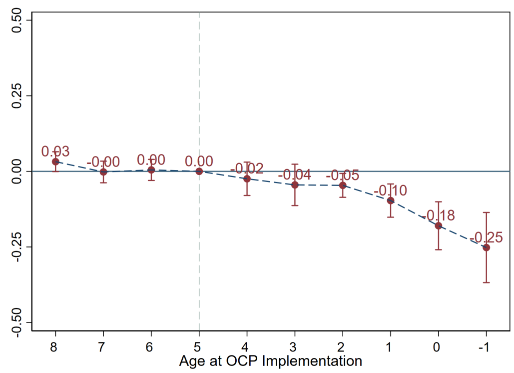
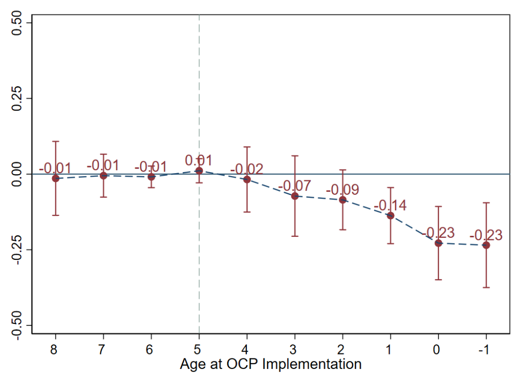


Figure C14: Impact of the OCP Rollout on Post-relaxation Fertility of the Second Generation. This figure shows estimates of Eq. 2 using the number of children born after OCP relaxation as the dependent variable. The data come from the 2020 census and the sample is restricted to urban Han women born between 1969 and 1980. Control variables include province fixed effects, cohort fixed effects, and province-specific cohort linear trends. Robust standard errors are clustered at province level. Vertical bars represent confidence intervals at the 95% level.

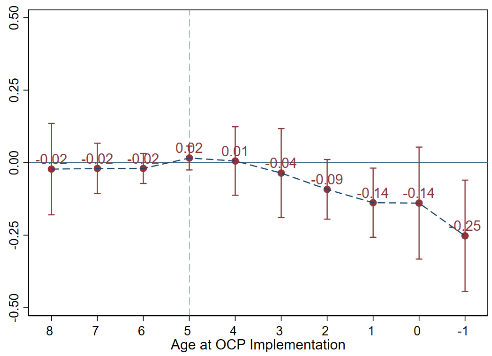


(a) Estimated Using Sun and Abraham (2021)

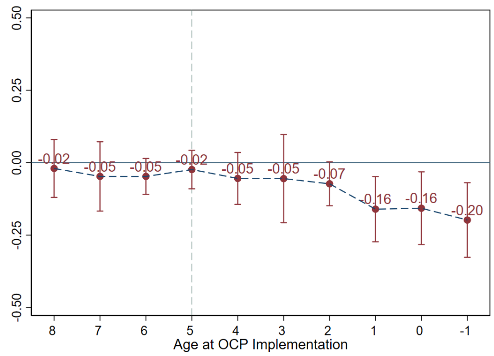


(b) Estimated Using De Chaisemartin and D'Haultfoeuille (2024)

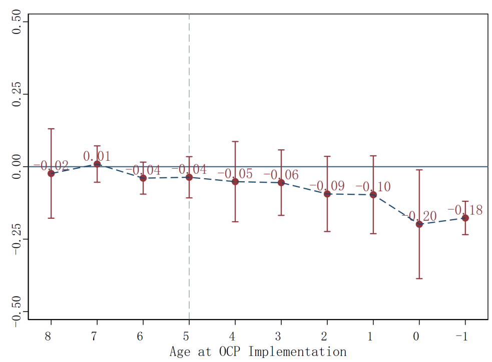
Figure C15: Impact of the OCP Rollout on Second-generation Fertility after the OCP Relaxation: Robustness to Heterogeneous Treatment Effects. Figures (a) and (b) show coefficients estimated using the estimators proposed by Sun and Abraham (2021) and De Chaisemartin and D'Haultfoeuille (2024), respectively. In Figure (a), the dependent variable is the number of children ever born by 2020. Control variables include province-specific cohort linear trends. Robust standard errors are clustered at province level. In Figure (b), the dependent variable is the residuals obtained by partialling out province-specific cohort linear trends. Standard errors are computed using bootstrap with 500 replications. Data in both figures are from the 2020 census, with the sample restricted to urban Han women born between 1969 and 1980. Vertical bars represent 95% confidence intervals.



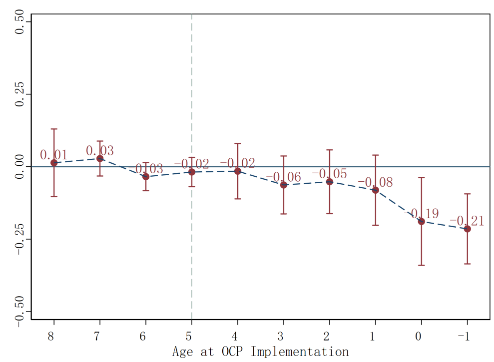
(a) Later, Longer, Fewer



(b) Compulsory Education Law



(c) Great Famine



(d) Cultural Revolution

Figure C16: Impact of the OCP Rollout on Second-generation Women's Fertility: Robustness to Multiple Treatments. Figures (a)-(d) plot the estimates based on De Chaisemartin and D'Haultfœuille (2023), holding each alternative treatment at its initial value: Figure (a) for the Later, Longer, Fewer Campaign of the 1970s, Figure (b) for the Compulsory Education Law (circa 1986), Figure (c) for the Great Famine, and Figure (d) for the Cultural Revolution. The dependent variable in all figures is the residuals of second-generation fertility, obtained by partialling out province-specific cohort linear trends. The data are from the 2020 census, and the sample is restricted to urban women born between 1969 and 1980. Following De Chaisemartin and D'Haultfœuille (2023), the sample is restricted to women who experienced the invariant intensity of each alternative treatment within the sampling period. The estimation also incorporates nonparametric trends across different provinces with varying levels of policy intensity. Standard errors are computed using bootstrap with 500 replications. Vertical bars in all figures represent 95% confidence intervals.

Table C1: Robustness Checks Using a Continuous Measure of OCP Exposure

	(1)	(2)
Sample	*Mini-census 2005*	*Census 2020*
Dependent Variable	*First-generation Fertility (Second-generation Sibling Size)*	*Second-generation Fertility by 2020*
Exposure	-0.117***	-0.044*
	(0.023)	(0.022)
Observations	32,636	75,438
Mean	1.374	1.455

Notes: This table shows the estimates of Eq. 2, replacing the set of D^l_{cp} with $Exposure_{pc}$. In column (1), the outcome variable is the sibling size of second-generation women in the 2005 mini-census. Samples are restricted to second-generation Han women with urban hukou born during 1975–1980, because individuals born before 1975 do not report the number of siblings in the 2005 mini census. Control variables include province and cohort fixed effects. In column (2), the outcome variable is the number of children ever born to second-generation women in the 2020 census. Samples are restricted to urban Han women born between 1969 and 1980. Control variables include province fixed effects, cohort fixed effects, and province-specific linear cohort trends. Robust standard errors clustered at province level are in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

Table C2: Summary Statistics of Mothers of Second-generation Women

	(1)	(2)	(3)
Variables	*Mean*	*Std.Dev.*	*N*
Have Ever Been Educated (Yes = 1)	0.910	0.287	96,849
White-collar job (Yes = 1)	0.344	0.475	96,849
Age at Birth	26.612	3.803	96,835

Notes: This table presents summary statistics for the characteristics of the mothers of second-generation women. The data are from the 1982 census, with the sample restricted to mothers who gave birth to a daughter between 1972 and 1980. Further restrictions are applied to mothers for whom the self-reported number of children matches the number of living children in the household. Since *Hukou* status is not available in the 1982 census, the sample is further restricted to mothers participating in non-agricultural work, using employment status as a proxy.

D Supplementary Figures and Tables

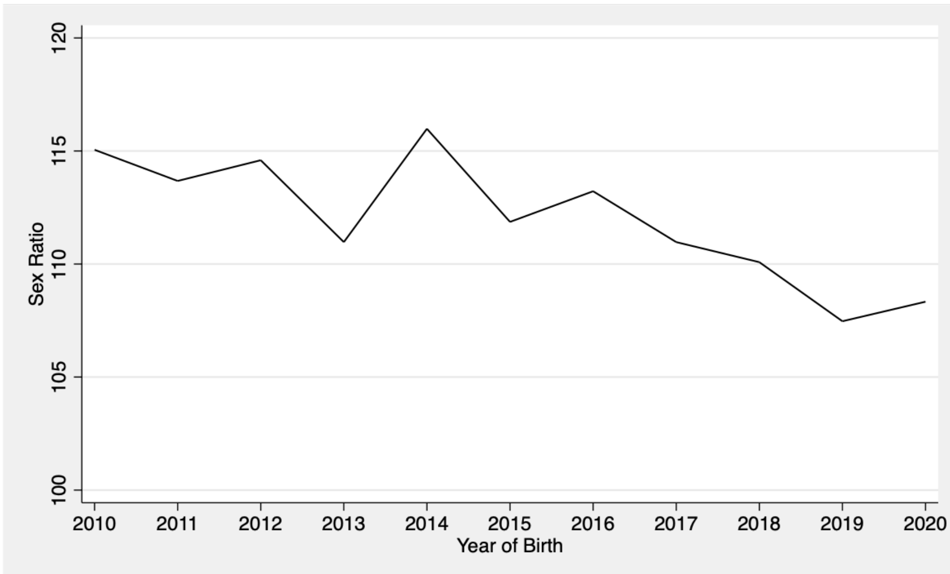


Figure D1: Sex Ratio of New Births. We calculate the sex ratios of new births from 2010 to 2020 using the 2020 census.

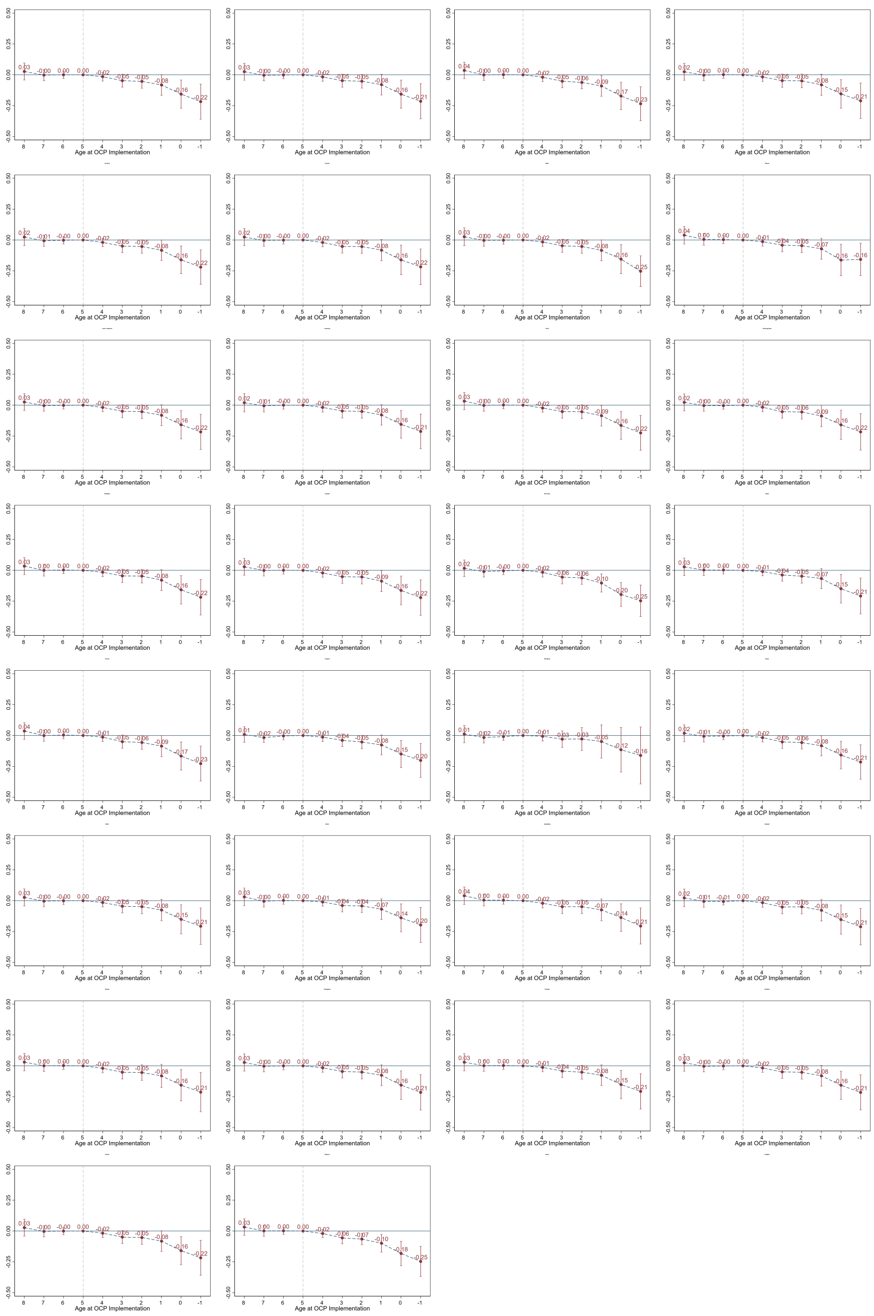


Figure D2: Impact of the OCP Rollout on Second-Generation Fertility after Policy Relaxation: Leave-One-Out Robustness. These figures show treatment effects estimated using Eq. 2, each time excluding one province. The dependent variable is the number of children ever born by 2020. Data are from the 2020 census, limited to urban Han women born between 1969 and 1980. All models control for province fixed effects, cohort fixed effects, and province-specific cohort trends. Robust standard errors are clustered at province level. Vertical bars indicate 95% confidence intervals.

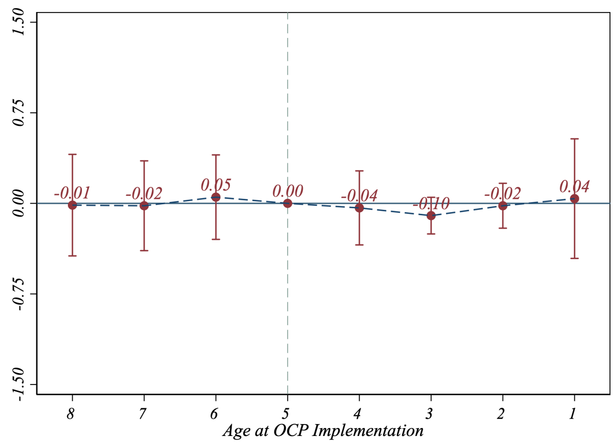
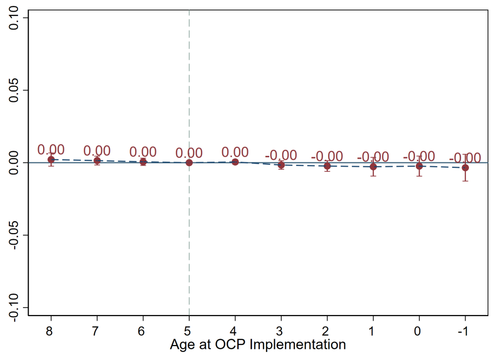
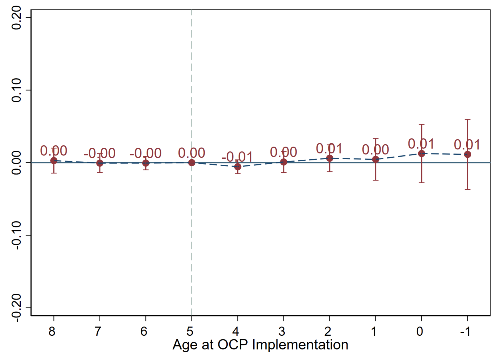


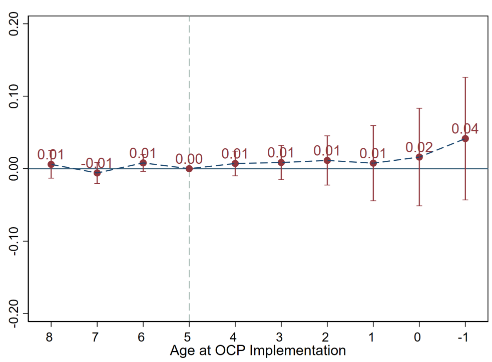
Figure D3: Impact of the OCP Rollout on First-generation Fertility Outcomes in FIDS. This figure plots the coefficients estimated using Eq. 2, with the number of desired children in 1987 as the dependent variable. Data in all figures are sourced from the 1987 FIDS, and samples are restricted to urban Han mothers born between 1969 and 1980. Control variables in all figures include province fixed effects, cohort fixed effects, province-specific cohort linear trends, and mothers' birth year fixed effects. Robust standard errors in all figures are clustered at province level. Vertical bars represent 95% confidence intervals.



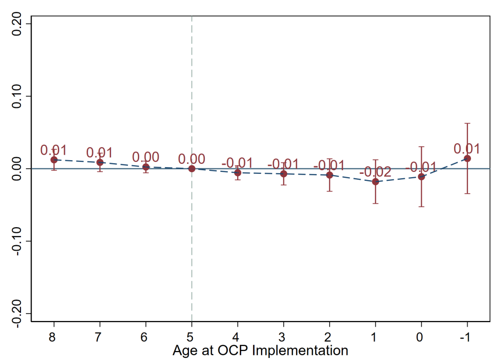
(a) Primary School Completion (Yes = 1)



(b) Junior High School Completion (Yes = 1)

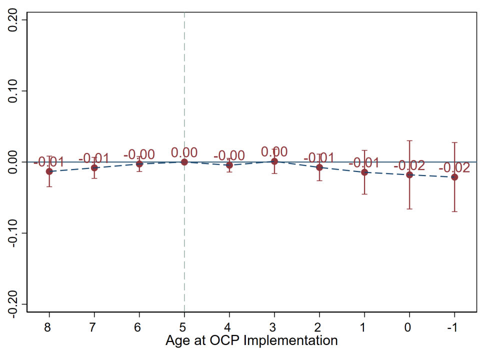


(c) Senior High School Completion (Yes = 1)

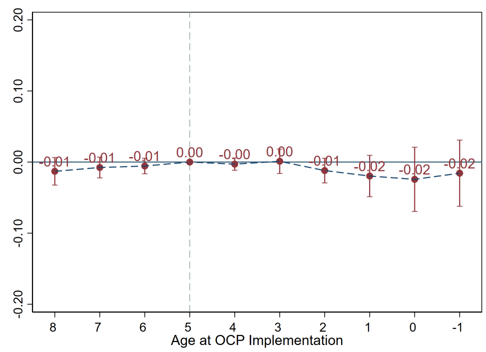


(d) College Completion (Yes = 1)

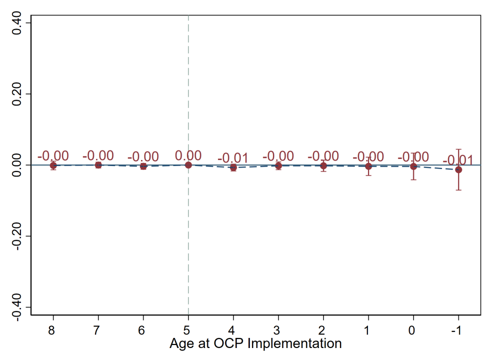
Figure D4: Impact of the OCP Rollout on Second-generation Women's Human Capital Accumulation before OCP Relaxation. Figures (a)-(d) show coefficients estimated using Eq. 2, with second-generation women's educational attainment as the dependent variable. Figures (a) to (d) use different measures of educational attainment as dependent variables: Figure (a) for primary school completion, Figure (b) for junior high school completion, Figure (c) for senior high school completion, and Figure (d) for college completion. Data in all figures are from the 2010 census and samples are restricted to urban Han women born during 1969-1980. Control variables in all figures include province fixed effects, cohort fixed effects, and province-specific cohort linear trends. Robust standard errors in all figures are clustered at province level. Vertical bars represent confidence intervals at the 95% level.



(a) Participating in the Labor Market (Yes = 1)

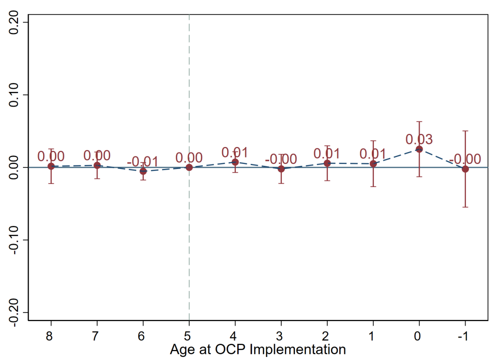


(b) Working in the Past Week (Yes = 1)

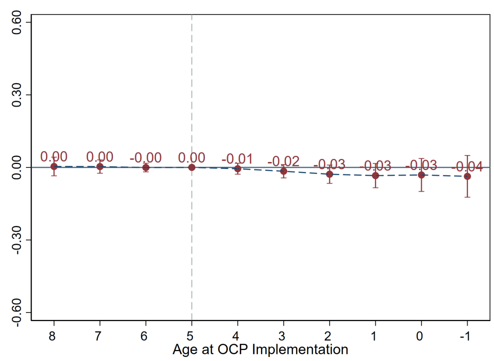


(c) Ln(Total Working Hours in the Past Week + 1)

Figure D5: Impact of the OCP Rollout on Second-generation Women's Labor Market Outcomes before OCP Relaxation. Figures (a) to (c) plot the coefficients estimated using Eq. 2, with second-generation women's labor market outcome as the dependent variable. Figures (a) to (c) use different measures of labor market outcomes: participation (Figure a), working status (Figure b), and log working hours in past week (Figure c). Data are from the 2010 census and samples are restricted to urban Han women born between 1969 and 1980. Control variables in all figures include province fixed effects, cohort fixed effects, and province-specific cohort linear trends. Robust standard errors in all figures are clustered at province level. Vertical bars represent confidence intervals at the 95% level.



(a) Coresiding with Parents or Parents-in-law (Yes = 1)



(b) Coresiding with People Older than 60

Figure D6: Impact of the OCP Rollout on Second-generation Women's Living Arrangement before OCP Relaxation. These figures plot the coefficients estimated using Eq. 2, with second-generation women's living arrangement as the dependent variable. Dependent variables include binary indicators for women's living arrangement: coresidence with parents or parents-in-law (Figure a) and coresiding with people older than 60 (Figure b). Data are from the 2010 census and samples are restricted to urban Han women born between 1969 and 1980. Control variables in both figures include province fixed effects, cohort fixed effects, and province-specific cohort linear trends. Robust standard errors in both figures are clustered at province level. Vertical bars represent confidence intervals at the 95% level.

Table D1: Impact of the OCP Rollout on Human Capital Accumulation, Labor Market Outcomes, and Living Arrangements

	Human capital accumulation				*Labor market outcomes*			*Living arrangements*
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Dependent Variables	Primary school (Yes = 1)	Junior high school (Yes = 1)	Senior high school (Yes = 1)	College or above (Yes = 1)	Working status (Yes = 1)	Labor participation (Yes = 1)	Logarithm of working hours	Coresiding with parents or parents-in-law (Yes = 1)
Age at OCP 8	0.002	0.003	0.006	0.012*	-0.013	-0.013	-0.002	0.003
	(0.002)	(0.008)	(0.009)	(0.007)	(0.009)	(0.010)	(0.006)	(0.010)
Age at OCP 7	0.001	-0.001	-0.006	0.009	-0.008	-0.008	-0.000	0.001
	(0.001)	(0.006)	(0.007)	(0.006)	(0.007)	(0.007)	(0.004)	(0.008)
Age at OCP 6	0.001	-0.001	0.008	0.002	-0.006	-0.003	-0.003	-0.007
	(0.001)	(0.004)	(0.006)	(0.004)	(0.005)	(0.005)	(0.004)	(0.006)
Age at OCP 5	-	-	-	-	-	-	-	-
	-	-	-	-	-	-	-	-
Age at OCP 4	0.000	-0.006	0.007	-0.006	-0.003	-0.004	-0.006	0.006
	(0.001)	(0.005)	(0.008)	(0.005)	(0.004)	(0.005)	(0.005)	(0.007)
Age at OCP 3	-0.002	0.001	0.009	-0.007	0.001	0.001	-0.002	-0.001
	(0.001)	(0.007)	(0.011)	(0.007)	(0.008)	(0.008)	(0.005)	(0.009)
Age at OCP 2	-0.002	0.006	0.011	-0.009	-0.012	-0.008	-0.001	0.004
	(0.002)	(0.009)	(0.016)	(0.011)	(0.008)	(0.009)	(0.008)	(0.012)
Age at OCP 1	-0.003	0.005	0.008	-0.018	-0.020	-0.014	-0.003	0.008
	(0.003)	(0.014)	(0.025)	(0.015)	(0.014)	(0.015)	(0.012)	(0.015)
Age at OCP 0	-0.002	0.013	0.016	-0.011	-0.024	-0.018	-0.003	0.027
	(0.003)	(0.019)	(0.032)	(0.020)	(0.021)	(0.023)	(0.018)	(0.019)
Age at OCP -1	-0.003	0.011	0.042	0.014	-0.016	-0.021	-0.012	0.002
	(0.004)	(0.023)	(0.040)	(0.023)	(0.022)	(0.023)	(0.027)	(0.025)
Observations	198,203	198,203	198,203	198,203	198,203	198,203	148,251	195,433
Mean of dep. var.	0.994	0.876	0.389	0.077	0.758	0.810	3.801	0.204

Note: This table shows estimated results of the impact of OCP rollout on various outcomes of second-generation women. Data are from the 2010 census, and samples are all restricted to urban Han women born during 1969-1980. The table shows results estimated using Eq. 2, and each column shows results estimated using different variables as dependent variables. Control variables in all columns include province fixed effects, cohort fixed effects, and province-specific cohort linear trends. Robust standard errors clustered at province level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1

E Further Details of the Mechanism Analyses

E.1 Social Interaction

E.1.1 Balance Tests

We compare the fertility behaviors of urban Han and urban ethnic minority women who were largely unaffected by the OCP. Specifically, we restrict the sample to women born before 1973—at least 6 years prior to the policy's implementation—to ensure that their mothers were largely unexposed to the OCP. We estimate the following equation:

$$Y_{icp} = \alpha_p + \theta_c + \phi \text{Minority}_{icp} + \lambda \cdot X_{icp} + \epsilon_{icp},$$

where Y_{icp} denotes the fertility outcome of woman i, born in cohort c and province p, measured in 2020; Minority_{icp} is an indicator that equals one if the individual belongs to an ethnic minority group and zero otherwise; and X_{icp} is a vector of individual-level control variables. The definitions of all variables follow those used in Eq. 2.

Panel A of Table E1 presents the results. We find no statistically significant differences in fertility outcomes—measured either by the number of children ever born by 2020 or by the probability of ever giving birth—between Han and minority women in urban areas. The results suggest that the two groups had similar fertility preferences in the pre-policy period.

We also apply the same approach to urban women and rural-to-urban migrant women. Panel B of Table E1 presents the estimated results. No significant differences in fertility behavior between urban Han women and rural-to-urban migrant women were detected, which supports the assumption that they share similar fertility preferences.

E.1.2 Empirical Analysis Using Ethnic Minority Sample

The sample mainly consists of urban ethnic minority women with two minor changes. First, we exclude ethnic minorities from prefectures where, for the 1969–1973 birth cohort, their average living arrangements differ substantially from those of Han women in the same region, given that we use the minority sample as a proxy for Han Chinese mothers unaffected by the OCP.

This exclusion filters out ethnic minorities less culturally aligned with the Han Chinese population. Second, we exclude women who belong to ethnic minorities whose overall population is too small to offer a precise inference.

We re-estimate Eq. 2 using the sample of ethnic minorities with five changes. First, we omit province-specific cohort trends, given that ethnic minorities are exempt from stringent one-child restrictions. This renders the use of province-specific cohort trends unnecessary for accommodating the life-cycle fecundity pattern. Second, we introduce controls for the share of Han Chinese individuals within the residential region and its quadratic form in the regression. The optimal fertility for ethnic minorities (those whose mothers are untreated) decreases with the policy only when the share of women whose mothers are treated is sufficiently large. The treatment effect in the untreated group is nonlinearly determined by the proportion of women whose mothers were treated, which is proportional to the share of ethnic minorities. Third, we incorporate city fixed effects in the analysis to account for potential heterogeneity in residential habits and the proportion of ethnic minorities at city level. Fourth, we exclude women aged 0 or -1 at OCP implementation from our analysis, since the samples from those age groups are insufficient for robust estimation. Fifth, we cluster standard errors at province and ethnicity level.

E.1.3 Estimation of the Weight α

The analysis in the main text quantifies a lower bound on the contribution of the maternal effect in shaping fertility norms to the OCP effect on daughters' fertility. Interestingly, we are also able to establish a lower bound for a structural parameter: the strength of maternal relative to peer effects, α, in the woman's utility function (Eq. 1):

$$\begin{aligned}\alpha &= \frac{N_{ur,2} - N_{r,2}}{n_1^u - n^p} \\ &\geq \frac{n^*_{ur,2} - n^*_{r,2}}{n_1^u - n^p} \\ &= \frac{\lambda(n^*_{r,2} - n_2^u) + (1-\lambda)(n^*_{ur,2} - n_2^u) - (n^*_{ur,2} - n_2^u)}{n_1^u - (\lambda n^p + (1-\lambda)n_1^u)}.\end{aligned} \tag{E1}$$

The first equality is based on the fact that $N_{r,2} = \alpha n^p + (1-\alpha)\bar{n}_2$ and $N_{ur,2} = \alpha n_1^u + (1-\alpha)\bar{n}_2$, where $N_{r,2}$ ($N_{ur,2}$) is the fertility norm for those Han women whose mothers' fertility is (not) constrained by the OCP, as defined in Section 2.3. The second inequality is due to the fact that $N_{r,2} \leq n^*_{r,2}$ and $N_{ur,2} \geq n^*_{ur,2}$, as implied by Proposition 3. The third equality is straightforward.

We note that (1) our estimate of β_l based on the main sample using women's fertility as the dependent variable (Fig. 6b) is $\lambda(n^*_{r,2} - n_2^u) + (1-\lambda)(n^*_{ur,2} - n_2^u)$, as indicated by Eq. 3; (2) the same estimate based on the minority sample (Table E2) is $n^*_{ur,2} - n_2^u$; and (3) the same estimate based on the main sample using women's sibling size (Fig. 6a) as the dependent variable is $n_1^u - (\lambda n^p + (1-\lambda)n_1^u)$. Combining these three estimates, we are able to compute the lower bound of α. Taking $l = 3$ as an example, we have $\alpha \geq 0.165 (= \frac{0.81-0.66}{0.91})$.

E.1.4 Empirical Analysis Using Rural-to-urban Migrants Sample

The sample primarily consists of rural-to-urban Han migrant women born between 1969 and 1980. We exclude observations from prefectures where, during the period 1969–1973, the average education level, number of children ever born, and household wealth of rural women differ substantially from those of urban women in the same region. This is because we use the rural-to-urban Han migrant sample as a proxy for Han Chinese mothers who were not affected by the OCP.

We re-estimate Eq. 2 using the rural-to-urban migrant sample with four key changes. First, we calculate women's age at the time of OCP implementation in our event-study design based on the timing of OCP implementation in their current province of residence, rather than in their province of birth. This is because our goal is to estimate the peer effect of urban women on rural migrants, and therefore we need to align with the policies that affected urban women's mothers. Second, we incorporate city fixed effects in the analysis to account for potential heterogeneity in residential habits and the proportion of rural-to-urban migrants at city level. Third, we control for the timing of migration to capture the duration of peer effects and include education level to account for individual fertility preferences. Fourth, we exclude women aged 0 or -1 at OCP implementation from our analysis, since the samples from those age groups are insufficient for robust estimation.

E.2 The Bargaining Power Mechanism

To examine the potential bargaining power mechanism, we draw on rich information from the CFPS regarding intra-household bargaining and dowry practices. We pool six waves of the CFPS (2010–2020) and restrict the sample to urban Han women born between 1969 and 1980 to construct our second-generation analytic sample.

For the dowry indicator, we assign a value of 1 if the respondent reported receiving a dowry in any survey wave, and 0 otherwise. For variables related to dowry amount, we compute the average across survey years to smooth out transitory fluctuations. Bargaining power is proxied by whether the woman is listed as the household head—a question included only in the CFPS 2012 wave. We therefore define a binary indicator for household headship status based on responses in that year. Summary statistics for these variables are provided in Table E5.

We re-estimate Eq.2 using three dependent variables: (i) an indicator for whether the respondent received a dowry, (ii) the logarithm of the dowry amount, and (iii) the bargaining power indicator. Estimation results, presented in Fig. E1, provide evidence consistent with the proposed mechanism. Specifically, we find that first-generation exposure to the OCP increases both the likelihood and the amount of dowry payments, and that larger dowries are in turn associated with a higher probability of women being listed as household heads. This pattern highlights a potential pathway through which intergenerational shifts in gender norms may lead to increased female empowerment via enhanced intra-household bargaining power.

E.3 Attachment to Party Doctrine

Attachment to party doctrine would lower second-generation fertility through two channels. (a) Direct exposure to propaganda: Second-generation women received the propaganda by themselves and thus have lower fertility. (b) Intergenerational transmission of doctrine attachment: First-generation women who internalized party doctrine could pass that attachment to their daughters, and thus influence the latter's fertility choices.

Our event-study design already nets out channel (a). Within a province, second-generation cohorts experienced similar fertility propaganda, so province fixed effects capture any fertility variation arising from province-level propaganda intensity. Cohort fixed effects and province-specific cohort linear trends further absorb residual cohort differences in responsiveness to such propaganda.

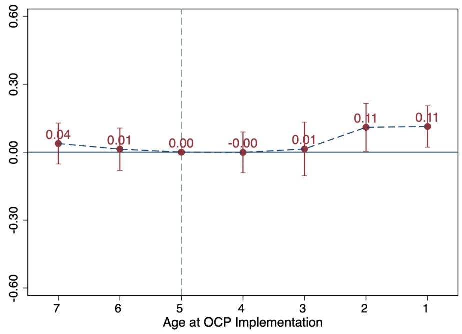
Channel (b) could potentially bias our estimates. If early OCP implementation is positively correlated with strong doctrine attachment among first-generation women, and if such attachment was transmitted to their daughters, then our estimates may be confounded by the intergenerational transmission of doctrine attachment. To address this concern, we conduct three empirical exercises.

(i) Fertility before OCP relaxation. If first-generation OCP exposure correlates with strong doctrine attachment, transmission of attachment would lower second-generation fertility when the OCP was still in effect. We re-estimate Eq. 2 using second-generation fertility before the OCP relaxation as the dependent variable. Column (3) of Table 2 shows that first-generation OCP exposure does not affect second-generation fertility before OCP relaxation. The results do not support intergenerational transmission in attachment to the low-fertility doctrine.

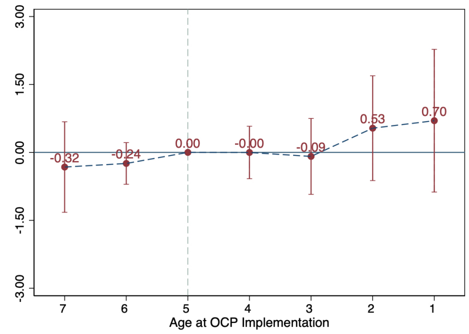
(ii) Propaganda reversal. The OCP relaxation accompanies a shift of propaganda from "one child only" to "encouraging two." If attachment to current propaganda were the key mechanism, the transmission of attachment would now raise second-generation fertility when the party recommends higher fertility. If this is true, the transmission of attachment inclination would counteract the diffusion of low-fertility norm, and render a lower bound estimate for the role of norms.

(iii) Broader attachment. We explore attachment to party doctrine beyond family-planning propaganda. As an external proxy for doctrine attachment, we examine the propensity to work for the party or in government institutions. We re-estimate Eq. 2 using an indicator for party or government job as the dependent variable. Fig. E4 shows that first-generation OCP exposure does not affect second-generation women's propensity to work for the party or in government institutions.

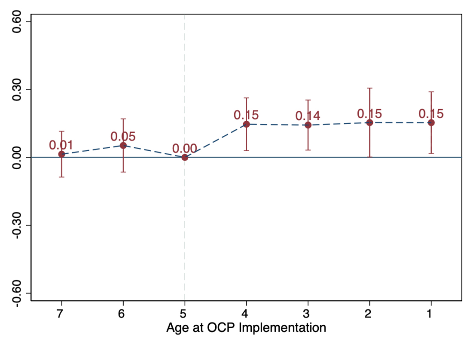
Overall, the correlation between early OCP rollout and strong attachment does not appear to confound the role of norms.



(a) Whether receiving a dowry (Yes=1)

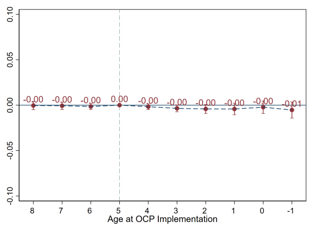


(b) Amount of the dowry (log)

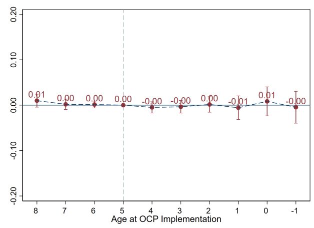


(c) Household head (Yes = 1)

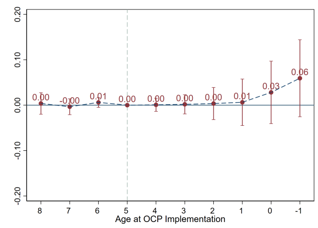
Figure E1: Impact of the OCP Rollout on Second-generation Women's Received Dowry and Household Bargaining Power. These figures plot the coefficients estimated using Eq. 2, with variables related to the dowry and bargaining power of second-generation women as the dependent variables. Dependent variables include an indicator for whether the woman received a dowry (Figure a), the amount of the dowry payment (Figure b), and an indicator for whether she was recorded as the household head (Figure c). Data are from the 2010-2020 CFPS and samples are restricted to urban Han women born between 1969 and 1980. Control variables in both figures include province fixed effects and cohort fixed effects. Robust standard errors in both figures are clustered at province level. Vertical bars represent confidence intervals at the 95% level.



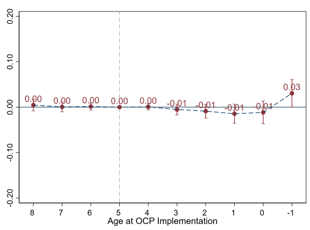
(a) Primary School Completion (Yes = 1)



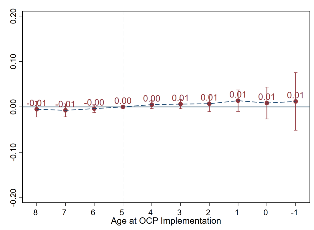
(b) Junior High School Completion (Yes = 1)



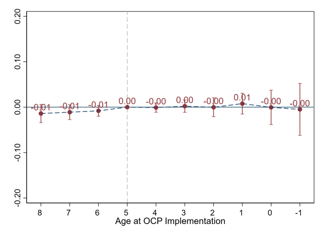
(c) Senior High School Completion (Yes = 1)



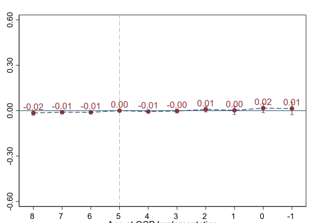
(d) College Completion (Yes = 1)



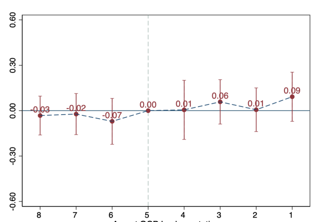
(e) Participating in the Labor Market (Yes = 1)



(f) Working in the Past Week (Yes = 1)

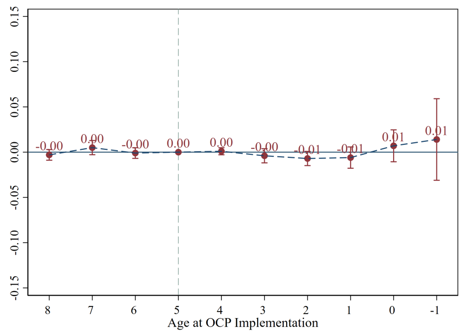


(g) Ln(Total Working Hours in the Past Week + 1)

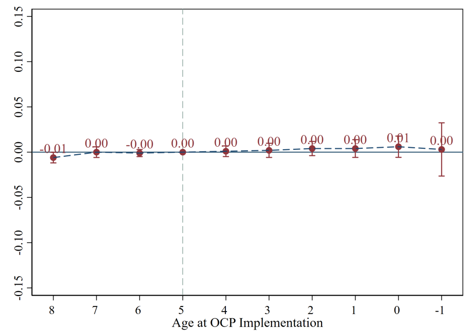


(h) Number of desired children

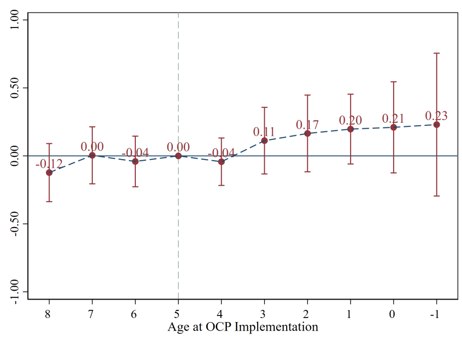
Figure E2: OCP Rollout and Second-generation Husbands' Characteristics. Figures (a)–(h) show coefficients estimated using Eq. 2, with second-generation women's husbands' characteristics as the dependent variables. Figures (a)–(d) use a series of indicators for educational attainment as dependent variables: Figure (a) for primary school completion, Figure (b) for junior high school completion, Figure (c) for senior high school completion, and Figure (d) for college completion. Figures (e)–(g) examine different measures of labor market outcomes: Figure (e) for labor-force participation, Figure (f) for working status, and Figure (g) for log working hours in past week. Figure (h) uses the number of desired children as the dependent variable. The data source is the 2010 census for Figures (a)–(g), and CFPS 2014, 2018 for Figure (h). Samples are restricted to urban Han women born during 1969-1980. Control variables in all figures include province fixed effects, cohort fixed effects, and province-specific cohort linear trends. Robust standard errors in all figures are clustered at province level. Vertical bars represent confidence intervals at the 95% level.



(a) Childlessness (Yes = 1)



(b) Ever Married (Yes = 1)



(c) Age at First Marriage

Figure E3: Impact of the OCP Rollout on Second-generation Women's Other Marriage and Fertility Decisions. These figures plot the coefficients estimated using Eq. 2, with second-generation women's marriage and fertility decisions as the dependent variables. Dependent variables include childlessness (Figure a), ever married (Figure b), and age at first marriage (Figure c). Data are from the 2020 census and samples are restricted to urban Han women born between 1969 and 1980. Control variables in both figures include province fixed effects and cohort fixed effects. Robust standard errors in both figures are clustered at province level. Vertical bars represent confidence intervals at the 95% level.

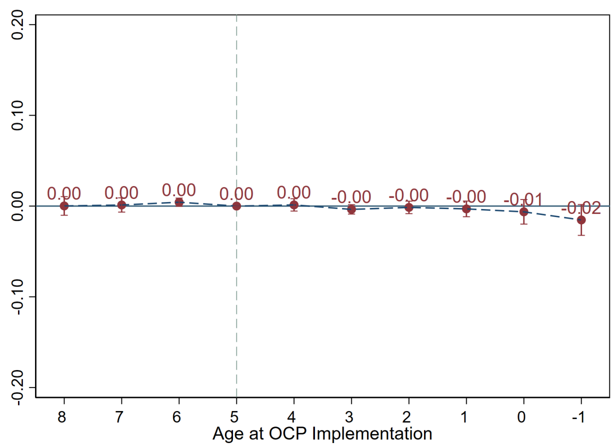


Figure E4: Impact of the OCP Rollout on Second-generation Women's Occupation. This figure plots the coefficients estimated using Eq. 2, with variables indicating whether women work in government or for the party as the dependent variables. Data are from the 2010 census and samples are restricted to urban Han women born between 1969 and 1980. Control variables in both figures include province fixed effects and cohort fixed effects. Robust standard errors in both figures are clustered at province level. Vertical bars represent confidence intervals at the 95% level.

Table E1: Balance Test across Different Samples

	(1)	(2)
Dependent Variable	*Number of children ever born by 2020*	*Have ever given birth (Yes =1)*
Panel A: Urban Han vs. Urban Ethnic Minorities		
Being an ethnic minority (Yes =1)	0.027	-0.002
	(0.041)	(0.006)
Observations	32,110	32,110
Mean of Dep. Var.	1.408	0.938
Panel B: Urban Non-migrants vs. Rural-to-urban Migrants		
Being a rural-to-urban migrant (Yes =1)	0.020	-0.002
	(0.022)	(0.004)
Observations	13,085	13,085
Mean of Dep. Var.	1.408	0.947

Notes: This table presents the coefficients from balance tests across different samples. Data are from the 2020 census. Dependent variables are the number of children ever born by 2020 and whether the individual has ever given birth to a child, respectively. Panel A presents the difference in fertility-related outcomes between Han Chinese and ethnic minorities, using a sample of urban residents, including ethnic minorities. The sample is further restricted to women born before 1973. The independent variable is whether the woman is an ethnic minority. Control variables include individual-level characteristics, as well as fixed effects for birth cohort and birth province. Panel B presents the difference in fertility-related outcomes between migrants and non-migrants, using a sample of urban Han women from areas with similar characteristics between migrants and non-migrants. The sample is further restricted to women born before 1973. The independent variable is whether the woman is a rural-to-urban migrant. Control variables include individual and residential city-level characteristics, as well as fixed effects for birth cohort and birth province.The samples in both panels are further restricted to women born before 1973, whose mothers' exposure to the OCP is minimal. Robust standard errors clustered at province level are in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

Table E2: Estimates Using Ethnic Minorities and Rural-to-urban Migrants

	(1)	(2)
Samples	*Ethnic Minorities*	*Rural-to-urban Migrants*
Dependent Variable	*Number of Children ever Born by 2020*	
Age at OCP = 6	0.010	0.012
	(0.043)	(0.038)
Age at OCP = 5	-	-
	-	-
Age at OCP = 4	0.007	-0.024
	(0.058)	(0.049)
Age at OCP = 3	-0.004	-0.017
	(0.042)	(0.052)
Age at OCP = 2	-0.042	-0.036
	(0.063)	(0.048)
Age at OCP = 1	-0.066	-0.065
	(0.044)	(0.053)
Observations	5,512	10,398
Mean of Dep. Var.	1.511	1.599

Notes: This table shows the estimates of Eq. 2 using different samples. Data are from the 2020 census. In column (1), samples are further restricted to urban women of ethnic minorities who were born between 1969 and 1980. In column (2), samples are further restricted to women who were rural-to-urban migrants. This is defined as residents who had rural *Hukou* and resided in urban areas. The dependent variables of both columns are the number of children ever born by 2020. Control variables include individual-level characteristics, province fixed effects, and cohort fixed effects. Robust standard errors clustered at province level are in parentheses. *** $p<0.01$, ** $p<0.05$, * $p<0.1$

Table E3: Summary Statistics of Second-generation Husbands

	(1)	(2)	(3)
Variables	*Mean*	*Std.Dev.*	*N*
Panel A: Census 2020			
Number of children born to the matched wives by 2020	1.615	0.711	60,714
Panel B: Census 2010			
Human capital accumulation			
Primary school completion (Yes = 1)	0.990	0.100	175,791
Junior high school completion (Yes = 1)	0.845	0.362	175,791
Senior high school completion (Yes = 1)	0.396	0.489	175,791
College completion (Yes = 1)	0.089	0.284	175,791
Labor market performance			
Working at least one hour in the past week (Yes = 1)	0.829	0.377	175,693
Participating in the labor market (Yes = 1)	0.879	0.326	175,693
Total working hours in the past week	47.810	12.000	143,631
Panel C: CFPS 2014, 2018			
Fertility desire			
Number of desired children	1.798	0.665	2,142

Notes: This table presents summary statistics for the characteristics of second-generation husbands. Panel A reports number of children born to the matched wives by 2020 from the 2020 census. Panel B reports data on human capital accumulation and labor market performance from the 2010 census. Panel C provides insights into fertility desires, based on the CFPS. All panels focus on husbands of urban Han women born between 1969 and 1980.

Table E4: Estimates Using the Matched Samples of Husbands and Wives

	(1)	(2)	(3)
Sample	*Matched samples of husbands and wives*		
Dependent Variable	*Second-generation Fertility by 2020*		
Age at OCP = 8	0.032	0.038	0.039
	(0.033)	(0.032)	(0.034)
Age at OCP = 7	-0.003	-0.001	0.001
	(0.021)	(0.021)	(0.023)
Age at OCP = 6	0.001	0.003	0.004
	(0.015)	(0.016)	(0.016)
Age at OCP = 5	-	-	-
	-	-	-
Age at OCP = 4	-0.013	-0.017	-0.018
	(0.016)	(0.015)	(0.015)
Age at OCP = 3	-0.047*	-0.047*	-0.048*
	(0.024)	(0.023)	(0.024)
Age at OCP = 2	-0.051*	-0.054*	-0.057**
	(0.027)	(0.026)	(0.027)
Age at OCP = 1	-0.070	-0.070	-0.074*
	(0.043)	(0.043)	(0.043)
Age at OCP = 0	-0.156**	-0.154**	-0.159**
	(0.058)	(0.058)	(0.060)
Age at OCP = -1	-0.192***	-0.188**	-0.195***
	(0.067)	(0.071)	(0.071)
Husbands' Birth Year FE		YES	YES
Husbands' Birth Province FE		YES	YES
Maternal OCP Exposure of Husbands			YES
Observations	52,584	51,280	51,280
Mean of Dep. Var.	1.482	1.482	1.482

Notes: This table shows the estimates of Eq. 2 using the samples of second-generation women whose husbands are also surveyed. Data are from the 2020 census. Control variables in all columns include birth province fixed effects and birth year fixed effects. In column (2), husbands' birth year and birth province fixed effects

Table E5: Summary Statistics of the Samples for Additional Mechanism Analyses

	(1)	(2)	(3)
Variables	*Mean*	*Std.Dev.*	*N*
Panel A: CFPS 2010-2020			
Whether being the household head	0.217	0.413	1,293
Whether receiving a dowry	0.420	0.494	1,914
Amount of the dowry	13977.426	23538.971	773
Panel B: Women in Government or Party			
Whether women work in government or for the party	0.024	0.154	156,335

Notes: This table presents summary statistics for additional mechanism analyses. Panel A presents summary statistics for women's bargaining power and related variables in the China Family Panel Study (CFPS) 2010-2020. Samples are restricted to urban Han women born during 1969–1980. For variables concerning parental transfer amount, we calculate the average values across different survey years. For the variable of whether receiving a dowry, a value of 1 is assigned if the respondent reported receiving a dowry in any given year and 0 if they never received a dowry. To measure bargaining power, we use whether the respondent is listed as the household head (this question was only included in the CFPS 2012 survey, so we define the bargaining power indicator based on responses to this question). Panel B reports the sample of women working in government or for the party using the 2010 census. The sample is restricted to urban Han women born between 1969 and 1980.

Table E6: Summary Statistics of Variables Related to Fertility and Marriage

	(1)	(2)	(3)
Variables	*Mean*	*Std.Dev.*	*N*
Childlessness	0.029	0.1687	75,447
Have ever married (Yes = 1)	0.981	0.1358	76,893
Age at first marriage	23.662	4.2867	75,447

Notes: This table presents summary statistics for outcomes related to fertility and marriage in the 2020 census. Samples are further restricted to urban Han women born during 1969–1980. The samples for respondents reporting childlessness and age at first marriage are further restricted to women who have ever married.

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