Does Birth Control Improve Economic Growth in Developing Countries?[1]

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National family planning programs have swept much of the developing world since the 1960s. Although the main motivation of these programs has been to enhance economic growth by reducing population growth, empirical evidence on their effectiveness is extremely scarce. Depending on the unique data from China's 30-year-old one-child policy, the current article examines the effect of birth control on economic growth. After addressing endogeneity and accounting for dynamic effects, we find evidence in contrast to the motivation of family planning programs: Birth control has hindered China's economic growth, and the total loss of this amounts to 5.6 trillion U.S. dollars. (JEL J13 O47)

Key words: Economic growth, birth control, One-child policy

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1. Introduction

National family planning programs have swept much of the developing world since the 1960s (Mundial 1984). The number of countries adopting national family planning programs reached more than 115 by the 1990s (Cleland et al. 2006), and global efforts on birth control are still in progress in developing countries (Kuang and Brodsky 2016, Stover and Sonneveldt 2017). While the vast majority of countries adopted voluntary family planning programs, the programs in highly populated countries, such as China, India, Bangladesh, and Indonesia, had significant elements of coercion (Harkavy and Roy 2007, Feng et al. 2013). Ironically, although the main motivation of these programs was to enhance income growth by reducing population growth (Coale and Hoover 1958, Ehrlich 1968),[2] empirical evidence on this is extremely scarce.

We believe the lack of empirical evidence is partly because of data limitation. The economic effect of a birth rate change may last for decades and differ over time (Bloom et al. 1998). Data from developing countries in which the birth control policies changed frequently is not suitable for evaluating the long-run and dynamic effects. An even more important reason for the minimal evidence is the difficulty caused by endogenous fertility. Fertility and income are simultaneously determined in an economy (McNicoll 1984, Becker and Barro 1988), and the effect of income on fertility likely changes with the level of economic development (Galor and Weil 2000). Without obtaining exogenous fertility variation caused by a birth control policy, it is impossible to rigorously evaluate the causal effect of birth control on economic growth.

Depending on the unique data from China's one-child policy (OCP), the current

[2] The rationale for family planning programs has *extended* to women's empowerment and reproductive health and rights since the fifth international population conference held in Cairo in 1994 (Cleland et al. 2006).

article provides an empirical examination of the effect of birth control on economic growth. China implemented the OCP from 1980 to 2011, and the policy generally allowed couples with a non-agricultural *hukou* (China's household registration system) to have only one child and allowed couples with an agricultural *hukou* to have a second child if the first was a daughter. The three decades of data from the OCP enables us to evaluate the long-run and dynamic effects. In addition, the highly coercive OCP substantially reduces the concern of endogenous fertility: The extremely low birth quota left almost no room for parents to endogenously choose their number of children according to their income and, therefore, significantly blocked the reverse causality from income to fertility.[3] Furthermore, the different birth limitations among residents with agricultural and non-agricultural *hukou* provides a potential instrument variable (IV) for identifying the causal effect.

Although the strict OCP substantially reduced the concern of endogeneity, fertility was still endogenous if the OCP was violated and the violation was correlated with income. For example, the incoming of new local political leaders might affect both the strictness of the OCP and local income growth. For this reason, we identify the effect through the system generalized method of moments (SYS-GMM) estimation that regress per capita GDP growth rate on the birth rate. In the estimation, the time-invariant determinants of the violation are eliminated by taking first difference, and the potentially endogenous birth rate (caused by the time-varying determinants of the violation) is instrumented by the lagged birth rate as well as by the percentage of residents with an agricultural *hukou*. After controlling for the various potential determinants of the violation, our IVs are likely uncorrelated with the error term and,

[3] The birth limitations of 1.0 and 1.5 per couple were far below the average birth rate in other middle- and low-income countries during 1980-2010, which was 3.3 per woman (World Bank 2018).

therefore, the effect identified is unbiased.

We obtain the *net* effect of a change in fertility on economic growth by estimating and adding the current and delayed effects. This is necessary because the different interactions of an individual with the economy over his or her lifetime.[4] For example, a birth is associated with a current burden on the economy but also contributes to the future labor force. To obtain the dynamic effects, we estimate a series of the above-mentioned SYS-GMM models where the birth rate is lagged by different years. The model in which the birth rate is lagged by, for example, 10 years captures the effect of a higher birth rate 10 years ago on the current period's income growth rate. The dynamic effects were then added to obtain the net effect.

We find that a unit decline in the birth rate caused by the OCP *increases* the current period's annual growth rate of per capita GDP by 0.29% but *reduces* the annual growth rate 5, 10, 15, and 20 years later by 0.30%, 0.46%, 0.30%, and 0.1%. Therefore, in contrast to the motivation of birth control policies, our finding suggests that the net effect of birth control on economic growth is significantly negative. This finding is robust to alternative IVs, control variables, and estimation methods. Combining these marginal effects with per capita GDP, population size, and the OCP-induced declines in the birth rate, we approximate that the total loss in GDP caused by the 30-year OCP was 38.2 trillion yuan (or 5.6 trillion U.S. dollars) with a 95% confidence interval of

[4] The *net* effect can also be obtained by estimating a long-run effect model if the following two assumptions are valid: first, the time series is long enough to cover a life cycle, and second, the age structure is stable (i.e., time-invariant birth rate and death rate). While the first assumption may be valid, the second assumption is obviously invalid. Although the dynamic effects of a change in fertility on income growth have been stressed in studies interested in the demographic dividend (e.g., Bloom and Williamson 1998, Bloom et al. 2009), most fertility-growth studies simply ignore the dynamic effects by implicitly assuming a stable age structure. However, nearly all developing economics are in demographic transition and contain an unstable age structure.

36.8 to 39.6 trillion yuan. This loss is nearly as large as the total real GDP in China in 2010, which was 41.0 trillion yuan.

To the best of our knowledge, the current article is the first to evaluate the net effect of a birth control policy on economic growth while considering the long-run dynamic effects. Most of the existing studies on birth control policies focused on the effects on other family outcomes, such as birth behavior (Ahn 1994, Huang et al. 2016), human capital investment in children (Rosenzweig and Zhang 2009, Li and Zhang 2017), and sex ratio imbalance (Ebenstein 2010, Goodkind 2011). A most relevant study was conducted by Li and Zhang (2007), who examined only the *current* effect of a decline in birth rate caused by China's OCP and concluded that birth control had improved economic growth in China. By considering both the current and delayed effects, we arrive at the opposite conclusion.

This article also contributes to the broader literature on population growth and economic growth. The theoretical literature has debated the causal effects of population growth on economic growth at least since Thomas Malthus (1798), though the empirical evidence on this is mixed.[5] The surveys of postwar literature generally found no robust

[5] The classical economist Malthus (1798) argued that a higher population depressed incomes per capita by diluting the fixed factors of production. The following neoclassical models (Solow 1956, Nelson 1956, Cass 1965) inherit this legacy by assuming diminishing returns to labour but also admit the possibility that population growth improves economic growth by inducing capital accumulation. However, the endogenous growth models (Romer 1990, Grossman and Helpman 1991, Aghion and Howitt 1992) generally predict a positive effect of population growth on income growth through improving technological progress. Empirically, Kremer (1993) found that income growth has been in proportion to population size over most of human history; however, a typical post-war observation is that the highest levels of fertility can be found in the poorest countries (Galor and Weil 2000). Many scholars (e.g., Coale and Hoover 1958, Mill 1965) understood this cross-country negative correlation as evidence of the negative effect of fertility on growth. On the other hand, economists like Kuznets (1960) and Boserup (1981) believed that population size actually enhances economic growth.

effect of changes in fertility rates on income growth (e.g., Simon 1989, Kelley and Schmidt 1994, Dasgupta 1995, Headey and Hodge 2009). It has long been argued that endogenous fertility is a major cause of the inconsistency in the population-growth literature. By addressing endogenous fertility, this article find results echoing the believe of "population optimists" (Boserup 1981, Romer 1986, Krugman 1991, Simon 1992) that population growth promotes economic growth.

The rest of this paper is organized as follows: Section 2 contains data and summary statistics, Section 3 details the econometric model, Section 4 presents estimation results, and the last section concludes the article.

2. Data and summary statistics

China formally implemented the OCP in 1980 to curb its surging population (Coale 1981, Zhu 2003). In general, the OCP allowed couples with a non-agricultural *hukou* to have only one child but allowed couples with an agricultural *hukou* to have a second child if the first was a girl (Feeney et al. 1989, Baochang et al. 2007).[6,7] Residents who violated the OCP not only faced a stiff fine but also faced the risk of employment loss and being unable to register their children for health care and education (Feng et al. 2013). There were three important exceptions within the OCP (Feeney et al. 1989, Baochang et al. 2007). First, residents in Xinjiang and Tibet were not subject to the

[6] *Hukou* was a system of household registration in mainland China during 1958-2014 under which each resident was classified in an agricultural or non-agricultural *hukou* (Zhao 2014).

[7] In fact, a universal one-child limitation was imposed at the beginning, but this draconian policy was met with strong resistance in China's vast rural areas and was modified in 1984 (some regions in 1982) to allow residents with an agricultural *hukou* to have a second child if the first was a girl (Hardee-Cleaveland and Banister 1988, Yi 1989).

OCP until the early 1990s. Second, couples with an agricultural *hukou* in three provincial cities (Beijing, Tianjin, and Shanghai) were only allowed to have one child. Third, residents who belonged to an ethnic minority population group (except, since the 1990s, minorities of *Zhuang* and *Manchu*) were allowed to have more than one child.[8] The OCP lasted for three decades, was significantly modified in 2011, and abolished in 2015.[9]

In order to employ the different birth statutes between residents with agricultural and non-agricultural *hukou* to identify the effect of fertility on income growth, our main analyses depended on provincial data over the three decades of the OCP. This study primarily used data from 24 of the 31 provincial districts in the mainland of China. Xinjiang, Tibet, Beijing, Tianjin, and Shanghai were excluded because they had no differences in their fertility statutes between *hukou* statuses (see the first and second exceptions of the OCP), and Hainan and Chongqing were excluded because they were separated from Guangdong and Sichuan in 1988 and 1997. However, during robustness checks, we included various alternative sample years and sample provinces. In order to reduce the confounding effect of unobservable inter-annual shocks, the total sample

[8] In addition, several other groups of population are allowed to have two children: 1) Rural couples in four provinces with a significant share of minority in population (Hainan, Ningxia, Qinghai, and Yunnan) 2) Miners who work underground, fishermen, farmers in mountainous or poor areas, and those deemed to have economic difficulties; 3) Those in a uxorilocal marriage; 4) Returning overseas Chinese; 5) Persons with the status of being the single child of a revolutionary martyr; 6) The first child has died or is physically handicapped.

[9] The OCP was significantly modified in 2011 to allow all parents who were only child themselves to have two children. This modification was important because in 2011, after three decades of the OCP, a large share of parents of prime reproductive ages (aged 25–44) were only children themselves. In 2015, the OCP was abolished and was replaced by a universal two-child policy.

period was divided into six five-year intervals and variables were calculated as five-year averages. Therefore, the data used in this analysis is primarily the averages of 1980–1984, 1985–1989, 1990–1994, 1995–1999, 2000–2004, and 2005–2009 within 24 provinces.

Table 1 presents this study's data sources. Most of the data used in this paper (rows 1–9) was derived from *China Compendium of Statistics 1949–2008 (2010)*, *National Bureau of Statistics of China*, and *China Population (and Employment) Statistics Yearbook (2009–2010)*. The data for the remaining three variables (rows 10–12), which will be used only in sensitivity checks, was derived from the *National Population Census of the PRC*, *China 1% National Population Sample Survey*, and *China Population Statistics Compendium 1949–1985 (1988)*. Although data is available for each of the sample years for most variables (so their five-year averages can be calculated), the data for the last three variables (rows 10–12) was only available for 1982, 1987, 1990, 1995, 2000, and 2005 (during this study's sample period). This data has been matched to each of the five-year intervals within the main dataset.[10]

[10] For example, the data from 1982 has been matched with the average from 1980–1984, and the data from 1990 has been matched with the average from 1990–1994.

Table 1. Data sources

(1) Real per capita GDP (yuan) (2) Investment share in GDP (%) (3) Growth of labor force share (%) (4) Trade share in GDP (%) (5) Government spending share (%)	**1980–2007**: *China Compendium of Statistics 1949–2008 (2010)* **2008–2009**: *National Bureau of Statistics of China* (http://www.stats.gov.cn/english/)
(6) Birth rate (1/1000) (7) Natural growth rate of population (fertility minus mortality, 1/1000)	**1980–2007**: *China Compendium of Statistics 1949–2008 (2010)* **2008–2009**: *China Population and Employment Statistics Yearbook (2009 –2010)*
(8) PAH: Percentage of residents with agricultural *hukou* (%) (9) Percentage of labor with primary, middle, or high school education (%)	**1980–2009**: *China Population (and Employment) Statistics Yearbook (1981 –2010)*
(10) Percentage of minority population (%)[#] (11) Youth dependency ratio (%)	**1982, 1990, 2000**: *National Population Census of the PRC (1982, 1990, and 2000)* [†] **1987, 1995, 2005**: *China 1% National Population Sample Survey (1987, 1995, and 2005)*
(12) Net out-migration rate (%)	**1981, 1985**: *China Population Statistics Compendium 1949–1985 (1988)* **1990, 2000**: *National Population Census of the PRC (1990 and 2000)* **1995, 2005**: *China 1% National Population Sample Survey (1995, 2005)*

Note: [†] The National Population Census of the PRC was conducted by the National Bureau of Statistics of the PRC in 1953, 1964, 1982, 1990, 2000, and 2010. The National Bureau of Statistics also conducts the 1% Population Sample Survey during inter-censual years ending in 5. These surveys were generally conduced in December of the survey years and therefore represents the data of the survey year. [#]These are the best minority data sources this study found. A related study (Li and Zhang 2007) did indicate that they used minority data in five year intervals from 1978–1998 from other sources; however, we failed to retrieve the data from the sources provided in their paper.

Table 2 contains the summary statistics of the variables. The key variables in our analyses are *annual growth rate of real per capita GDP*, *the birth rate*, *natural growth rate of population*, *percentage of residents with agricultural hukou*, and *percentage of minorities in the population*. The *annual growth rate of real per capita GDP* will be used as the dependent variable in the econometric estimation. China observed remarkable growth during the sample period, and the average annual growth rate was 8.5%. The *birth rate*, defined as the total number of live births per 1,000 women in a year, is the key independent variable in our main analysis. The average births per 1,000 women was 17.3 with a standard deviation of 4.7. See Figure 1 for the provincial birth rate declines during the OCP. The *natural growth rate of population* will be used as an alternative key independent variable (in place of the birth rate) to test the possibility of inferring the effect of population growth from the effect of changes in the fertility rate.

The *percentage of residents with agricultural hukou* (PAH) will be used as an IV for the birth rate. The average PAH was 78.7%. See Figure 1 for provincial changes in the PAH during the OCP. The *percentage of minorities in the population* (PMP) will be used as an alternative IV when making a comparison to the work of Li and Zhang (2007), who also used it as an IV. Our main analyses do not depend on the PMP because, as shown in the appendix, it is likely a weak IV.

Table 2 also summarizes 10 time-varying control variables that are believed to have effects on income growth (Levine and Renelt 1992, Kelley and Schmidt 2005): five-year lagged real per capita GDP (in constant 2010 price); percentage of labor with primary, middle, and high school education; net out-migration rate, which is out-migration minus in-migration; youth dependency ratio, which is the population ages 0–15 divided by the population ages 16–64; investment share in GDP; growth of labor force share; trade share in GDP; and government spending share in GDP.

Table 2. Summary statistics of variables

Variables	*N*	Mean	S.D.
Key variables			
Annual growth rate of real per capita GDP (%, constant 2010 price)	144	8.5	3.2
The birth rate (1/1000)	141	17.3	4.7
Natural growth rate of population (birth rate minus death rate, 1/1000)	141	11.0	4.6
PAH: Percentage of residents with agricultural *hukou* (%)	144	78.7	10.1
PMP: Percentage of minorities in the population (%)	144	11.3	14.3
Control variables			
Five-year lagged real per capita GDP (*yuan*, constant 2010 price)	144	8589	7482
Percentage of labor with primary education (grades 1–6, %)	144	43.2	10.3
Percentage of labor with middle education (grades 7–9, %)	144	30.3	8.2
Percentage of labor with high education (grades 10–12, %)	144	9.2	2.9
Net out-migration rate (out-migration minus in-migration, %)	144	0.6	6.6
Youth dependency ratio (%)	144	41.8	11.5
Investment share (% of GDP)	144	39.5	11.5
Growth of labor force share (%)	143	1.8	3.5
Trade share (% of GDP)	143	11.6	14.3
Government spending share (% of GDP)	144	13.3	5.4

Note: The number of observations for most variables are 144 (i.e., 24*6), but we have three missing values for the birth rate and the natural growth rate of population (Guangdong for 1980 and 2000 and Shaanxi for 2000), one missing value for the growth of labor force share (Inner Mongolia for 1980), and one missing value for trade share (Shaanxi for 1980).

Figure 1 presents the correlation between the provincial PAH and the birth rate in 1979 (Panel A) and 1981–2010 (Panel B). We find no obvious correlation between these factors just before the OCP in 1979. However, the birth rate varies from 14.6 to 27.6 across provinces, and the PAH was around 85% in most provinces (except the four northeast ones, which can be seen in the geographic distribution in Figure A2).[11] However, there is an obvious positive correlation between these factors for the 1981–2010 average. Although both the birth rate and the PAH dropped significantly compared to 1979, provinces with a higher PAH generally had higher birth rates. This reflects the fact that the OCP dramatically reduced fertility rates and that fertility statutes were less severe for individuals with an agricultural *hukou*.

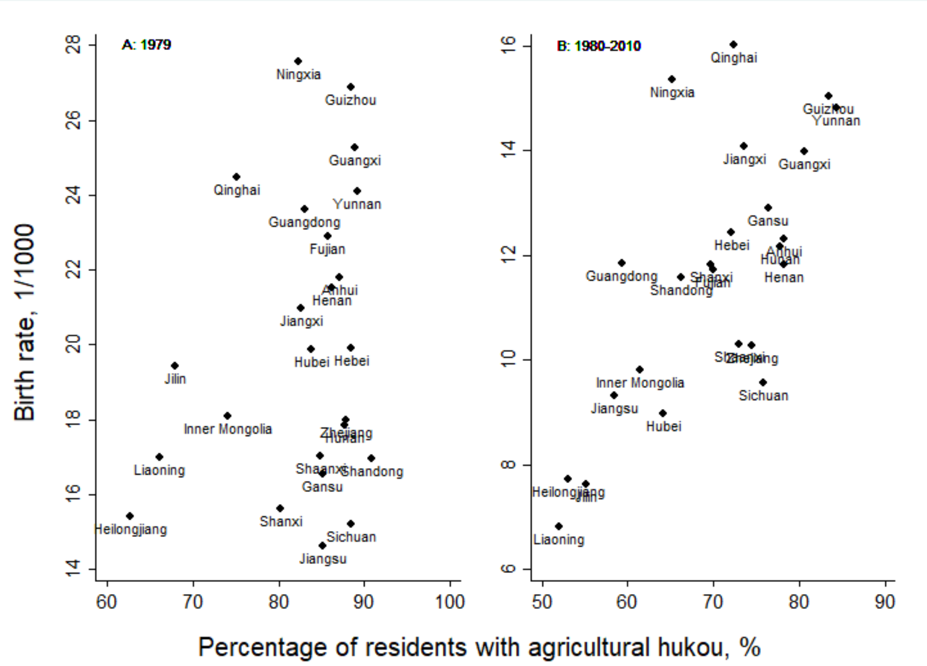

Figure 1. Correlation between the percentage of residents with agricultural *hukou* (PAH), the birth rate in 1979 (Section A), and the 1980–2010 average (Section B)

Figure 2 shows a positive correlation between the PAH and the growth rate of per capita GDP for most provinces during the OCP. Specifically, Panel A shows that the

[11] This lack of correlation was possible because a significant share of the labour with agricultural *hukou* were actually working in non-agricultural sectors, which implies that many determinants of the birth rate that were correlated with working sectors, such as income and education, were not necessarily correlated with *hukou* status. Although the *hukou* was initially set in 1958 based on a family's production sector (agricultural or non-agricultural) (Zhao 2014), massive rural-urban migration had been observed since then, and many migrants were not able to change their *hukou* status (Yang 1993, Solinger 1999). For example, the average share of labour in agriculture was 50.7% during 1980–2010 (National Bureau of Statistics of China 2011), but the average PAH during the same period was as high as 78.8%.

entire sample indicates a weak positive correlation between the PAH and the growth rate. Panels B and C find obvious positive correlations in the eight eastern coastal provinces and the 10 central provinces, respectively (see Figure A2 for the geographic distribution). Panel D finds no obvious correlation for the remaining six western provinces where minorities constitute a large share of the population.[12] This is potentially because there were no differences in fertility statutes among minorities with different *hukou* statuses. Figures 1 and 2 together provide suggestive evidence that provinces with a higher birth rate (caused by higher PAHs) experienced faster economic growth during the OCP.

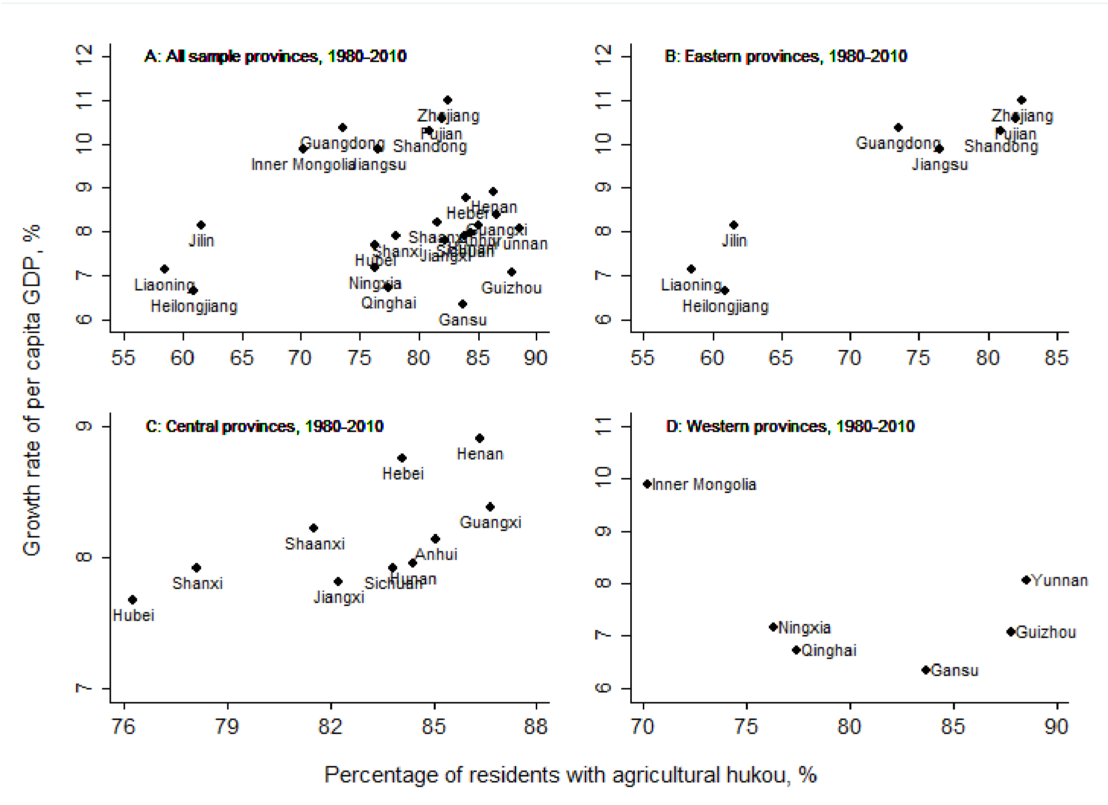

Figure 2. Correlation between the percentage of residents with agricultural *hukou* (PAH) and the growth rate of per capita GDP (1980–2010)

[12] The average share of minorities in the six provinces was 27.4% but was 4.7% in other sample provinces. See Figure A2 for details.

3. Methods

This study estimates the effect of fertility change on economic growth using a standard growth model:[13]

(1)
$$g_{it} = \alpha \log y_{i(t-\tau)} + \beta BR_{it} + X_{it}\gamma + u_i + \varepsilon_{it} \ , \quad i = 1, 2, ..., N; \quad t = 1, 2, ..., T$$

in which g_{it} is the annual growth rate of real per capita GDP in year t and province i; $\log y_{i(t-\tau)}$ is the log of real per capita GDP lagged by τ years (τ is the interval of the data), which is usually included in a growth regression to account for the effect of initial income on the growth rate; BR_{it} is the birth rate; X_{it} are other usually included time-varying determinants of economic growth, such as education, the share of investment in GDP, the share of trade in GDP, youth dependency ratio, net migration rate, and the growth rate of labor force (Levine and Renelt 1992, Kelley and Schmidt 2005); u_i denotes location fixed effects used to account for all observable and unobservable time-unvarying determinants of growth; α, β, and γ are coefficients; and ε_{it} is the error term.

To account for endogeneity, we estimate model (1) by the system generalized method of moments (SYS-GMM) (Arellano and Bover 1995, Blundell and Bond 1998). We take the first difference to remove u_i and its associated omitted-variable bias. The differencing results in

(2)
$$\Delta g_{it} = \alpha \Delta \log y_{i(t-\tau)} + \beta \Delta BR_{it} + \Delta X_{it}\gamma + \Delta \varepsilon_{it} \quad i = 1, 2, ..., N; \ t = 1, 2, ..., T \ .$$

[13] This kind of model has been widely used in empirical growth literature (e.g., Mankiw et al. 1992, Levine and Renelt 1992, Islam 1995, Li and Zhang 2007)

Endogenous variables (i.e., $\Delta \log y_{i(t-\tau)}$ and ΔBR_{it})[14] in the differenced equation are then instrumented by their lagged levels. Since the lagged values of $\log y_{i(t-\tau)}$ (i.e., $\log y_{i(t-2\tau)}, \log y_{i(t-3\tau)}, \ldots, \log y_{i0}$) are orthogonal to the error term $\Delta\varepsilon_{it}$ and correlated with $\Delta \log y_{i(t-\tau)}$, we can use them as IVs to produce a GMM estimator of α. Similarly, we use lags of the birth rate (i.e., $BR_{i(t-2\tau)}, BR_{i(t-3\tau)}, \ldots, BR_{i0}$)[15] as IVs for ΔBR_{it} based on the assumption that these lags are correlated to ΔBR_{it} but uncorrelated to $\Delta\varepsilon_{it}$. The latter assumption is equal to that omitted determinants of the current economic growth are uncorrelated with past birth rates. This assumption is likely true because *predictable* determinants of the birth rate have been controlled by the first difference, which removes all fixed effects, and the control variables included; *unpredictable* shocks to the current economic growth are unlikely to be correlated with past birth rates.

The above estimation method, in which GMM is applied to the first-differenced equation, is usually called the first-differenced GMM (DIF-GMM). The DIF-GMM estimates are biased upwards and inefficient when N is finite (Arellano and Bover 1995). As proposed by Blundell and Bond (1998), this study addresses the small sample bias by the SYS-GMM estimator that combines equations of the first differences (2) instrumented by lagged levels, with an additional set of equations in levels (1) instrumented by lagged first differences. We will use the SYS-GMM for our main estimation but will also report the DIF-GMM estimates for comparison.

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[14] First difference makes the predetermined variable $\log y_{i(t-\tau)}$ endogenous; the error term $\Delta\varepsilon_{it} \equiv \varepsilon_{it} - \varepsilon_{i(t-\tau)}$ is correlated with the regressor $\Delta \log y_{i(t-\tau)} \equiv \log y_{i(t-\tau)} - \log y_{i(t-2\tau)}$.

[15] Here we use only lags 2 and higher (i.e., lag 1 is excluded) as the instruments because the lag 1 of the birth rate is likely correlated to the error term in the differenced equation.

The SYS-GMM estimator of the equation (1) captures only the *current* effect of a change in the birth rate on the growth rate in the same period. However, a change in current birth rate may also have effects on future growth rates. We now extend model (1) to capture both the current and the delayed effects:

$$g_{it} = \alpha \log y_{i(t-\tau)} + \beta^q BR_{i(t-q\tau)} + X_{it}\gamma + u_i + \varepsilon_{it} \quad (3)$$
$$i = 1, 2, ..., N; \; t = 1, 2, ..., T.$$

The only difference from (1) is that the birth rate is lagged by $q\tau$ years in (3), where $q = 0, 1, 2, ..., Q$. Model (3) is the same as model (1) if $q = 0$. When $q \geq 1$, the coefficient β^q captures the effect of a higher birth rate $q\tau$ years ago on the growth rate in the current year. Therefore, by estimating separate models where the birth rate is lagged by different years (including zero), we can obtain both the current and delayed effects of a higher birth rate. We estimate model (3) using the SYS-GMM procedure previously introduced, and the instruments for $\Delta BR_{i(t-q\tau)}$ are the lagged values of the lagged birth rate (i.e., $BR_{i(t-q\tau-2\tau)}, BR_{i(t-q\tau-3\tau)}, \ldots, BR_{i0}$). The sum of the current and the delayed effects reflects the net effect of a higher birth rate on economic growth.

The above SYS-GMM estimation only includes internal IVs for the birth rate—the lagged values of the instrumented variables. To raise the explanatory power, we also include an external IV: the PAH. As previously detailed, a higher PAH corresponds to a higher statutory birth rate. In addition, in a SYS-GMM estimation that removes fixed effects and controls for various time-varying factors, we assume that the PAH does not have a partial effect on economic growth except that through the birth rate. A similar IV was used in previous studies on the effect of the OCP on child quality (e.g., Qian 2009, Liu 2014). Tests presented in Table A1 of the appendix show that both the internal and external IVs have significant explanatory power regarding the birth rate. Our main analyses will include both the internal and the external IVs. But as robustness checks, we will also estimate models that include only one of them.[16]

4. Results

In this section, we estimate the baseline model (3), provide various robustness tests, and calculate the total effect of the OCP on economic growth in China.

4.1 Baseline Results

The baseline SYS-GMM estimates are reported in Table 3. From columns (1) to (5), we report the regressions where the birth rate is lagged by 0, 5, 10, 15, and 20 years, respectively.[17] The model is estimated by the Blundell and Bond (1998) two-step procedure where the covariance matrix is corrected for finite-sample bias using the method of Windmeijer (2005).[18] In each regression, the (current or lagged) birth rate is instrumented by its lags as well as the PAH (in the same period as the birth rate).

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[16] The statutory birth rate difference between *Han* Chinese and minorities, as previously mentioned, suggests that the PMP is also a potential external IV. Our main analyses do not include the PMP as an IV because, as detailed in Appendix A1, minorities comprised only a small share of the population and most minorities were living in several non-presentative western provinces where the population and economic densities were much lower than the remaining provinces. A test presented in Table A1 finds that the PMP has no significant explanatory power on the birth rate during 1980-2009, suggesting that the PMP is a weak IV. However, in a robustness check that replaces the PAH with the PMP as an IV we still find a negative current effect and a positive delayed effect of a higher birth rate.

[17] Recall that this study uses five-year interval data from 1980–2009. Since the SYS-GMM is estimated in a first-difference framework, the longest lag allowed is 20 years (four periods).

[18] The Blundell-Bond estimators have one- and two-step variants (Blundell and Bond 1998). Although two-step estimation is asymptotically more efficient, the reported two-step standard errors tend to be severely downward biased (Arellano and Bond 1991). Windmeijer (2005) developed a finite-sample correction to the two-step covariance matrix to correct this downward bias.

These baseline regressions include only the two most important control variables: lagged log GDP and education level; more controls will be used to test the robustness of the model in the following section.

Table 3. SYS-GMM estimates of the effect of birth rate on GDP growth rate

Dependent variable: Annual growth rate of real per capita GDP (%)					
	Lagged years of the birth rate				
Independent variables	(1) 0–5	(2) 5–10	(3) 10–15	(4) 15–20	(5) 20–25
Birth rate (1/1000)	**-0.29***** **(0.06)**	**0.30***** **(0.05)**	**0.46***** **(0.13)**	**0.30***** **(0.12)**	**0.10** **(0.22)**
Five-year lagged LogGDP	1.68*** (0.31)	4.44*** (0.47)	5.43*** (0.54)	4.11*** (1.00)	21.40 (13.73)
Primary education (%)	0.09 (0.07)	0.33*** (0.03)	0.52*** (0.15)	0.90*** (0.27)	-1.43 (1.88)
Middle school education (%)	-0.22*** (0.03)	-0.13*** (0.04)	-0.34*** (0.02)	-0.17** (0.08)	0.03 (0.24)
High school education (%)	0.14*** (0.02)	0.16*** (0.02)	0.09*** (0.03)	0.60*** (0.16)	0.17 (0.67)
Hansen test of over-identification restriction					
J-statistics	19.2	22.3	15.8	13.2	2.5
P-value	0.95	0.81	0.78	0.43	0.74
Arellano-Bond test for second-order serial correlation					
z-statistics	1.61	1.51	0.65	--	--
P-value	0.11	0.13	0.51	--	--
Provinces	25	25	25	25	25
Observations	144	121	98	73	48

Notes: 1) The birth rate is instrumented by all of its lags as well as the PAH, and the lagged LogGDP is instrumented by all of its lags; 2) We adopt the forward orthogonal deviations transformation (see Arellano and Bover 1995 for details) instead of the first difference transformation to avoid the magnifying of gaps in unbalanced panels (See Table 2); 3). The Arellano-Bond test is not reported in columns (4) and (5) because the number of observations is too small for the test when delaying the birth rate by 15 and 20 years. 4) The Huber-White heteroskedastic consistent standard errors are reported in parentheses. 5). *, **, and *** represent significance levels of 10%, 5%, and 1%.

As reported in Table 3, although the current effect (Column 1) of a *higher* birth rate on GDP growth rate is negative, all delayed effects are positive (Columns 2–5). Most estimated coefficients of the birth rate are statistically significant at 1% level, except where the birth rate is lagged by 20 years (Column 5). Specifically, other things being equal, a unit increase in the birth rate will reduce the *annual* growth rate of per capita GDP by 0.29% in 0–5 years, but will increase the annual growth rate by 0.30%, 0.46%, 0.30%, and 0.1% in the following 5–10, 10–15, 15–20, and 20–25 years, respectively. The negative current effect of a higher birth rate is substantially smaller than the positive delayed effects, which implies a negative net effect of *birth control* on the growth of per capita GDP.

These findings are intuitive when referencing back to the literature regarding the costs and benefits of population growth. The net effect of population growth on economic growth depends on the cost of childrearing (especially the time cost for mothers) (Galor and Weil 1996), the capital dilution effect (Malthus 1798, Solow 1956), and the induced capital accumulation and innovation (because of population pressure and a larger number of innovators) (Boserup 1981, Romer 1986, Krugman 1991, Simon 1992). In the first several years of childrearing, the time cost is very high; the induced capital accumulation and innovation may occur immediately if parents save more and work harder to feed more children; the capital dilution effect occurs only when the new cohort enters the labor force; and the larger labor force will also incentivize capital accumulation and innovation. Therefore, our findings suggest that the negative time cost dominates initially; with the decline of the time cost as children growing, the positive effect of the induced capital accumulation and innovation dominates; finally, when the new cohort enter labor force, the negative capital dilution effect obscures the positive effect of the induced capital accumulation and innovation.

For control variables, the estimated coefficients of the *five-year lagged per capita GDP* are positive for all regressions, which indicates inter-provincial divergence in per capita income in China during the sample period. This finding is consistent with the widely observed expansive trend of inter-provincial income disparities during this period (e.g., Jones et al. 2003, Zhang and Zou 2012).[19] The estimated coefficients of the *shares of labor force with different levels of education* are generally statistically significant, and the overall effect of education is positive, which reflects that human capital investment enhances economic growth.

Table 3 also reports the *J*-statistics and *P*-value of the Hansen overidentification restriction test. The *P*-values of the Hansen tests are all larger than 0.1 in columns (1)–(5), which suggests that there is no evidence to reject the validity of our IVs in each regression. We also report the Arellano-Bond tests (Arellano and Bond 1991) for the second-order serial correlation in the first-differenced residuals; no second-order serial correlation in the first-differenced residuals is a necessary condition for consistent estimates. The *P*-value of the Arellano-Bond test is larger than 0.1 for each regression in which the test was conducted,[20] which suggests that the null hypothesis, which is the absence of the second-order serial correlation, has not been rejected.

4.2 Robustness check

We check the robustness of the above findings to measures of population growth, alternative IVs, control variables, estimation methods, and subsamples. All checks considered here include the same model setting as the baseline model, except for the one specified in each check. To save space, we report only the estimates of the birth

[19] Theoretical models predict that it is possible for income growth to decrease with income level (Solow 1956), independent of income level (Lucas Jr 1988), and increase with income level (Romer 1986), depending on the relative effects of diminishing returns to capital and the externalities of technical change.

[20] The Arellano-Bond test is not conducted for regressions that contain a 15 or 20 year delayed birth rate because the number of observations is too small for the test.

rate, and full results are available in the online appendix.

The first check replaces the key explanatory variable, the birth rate, with the natural growth rate of the population, which is the birth rate minus the death rate. Our main analyses infer the effect of population growth on economic growth through the effect of the birth rate. To determine if this inference is reliable, Row (1) of Table 4 presents the estimates from the model that replace the birth rate with the natural growth rate of population. The main finding is consistent with the baseline model: a higher natural growth rate of population has a negative effect on economic growth only within a 5-year period and the effects become positive after that, and the following positive effects overcome the initial negative effect.

Rows (2) to (4) test the sensitivity to the IVs of the birth rate. The baseline GMM estimation includes both the internal and external IVs. Row (2) presents the model uses only the internal IVs (lags of the birth rate), and Row (3) uses only the external IV (the PAH). The initial negative effect of a higher birth rate is still smaller than the following positive effects, although the effect is no longer statistically significant when the birth rate is lagged by 15 years. Row (4) replaces the external IV (the PAH) by the percentage of minorities in the population (PMP) to make a comparison with a similar study (Li and Zhang 2007) that used the PMP as the IV. Consistent with Li and Zhang (2007), we also find a negative *current* effect (note that they did not report the delayed effects). The delayed effects are mostly insignificant except where the birth rate is lagged by 5–10 years. This is presumably because the PMP is a weak IV (see the appendix for evidence).

Table 4. Robustness checks

Dependent variable: Annual growth rate of real per capita GDP (%)					
	Lagged years of the birth rate				
	(1) 0–5	(2) 5–10	(3) 10–15	(4) 15–20	(5) 20–25
Sensitive to the measure of population growth					
(1) Replace the birth rate by the natural growth rate of population	-0.25***	0.25***	0.36***	0.44***	0.13
	(0.03)	(0.07)	(0.11)	(0.14)	(0.23)
Sensitive to IVs					
(2) No external IV	-0.24***	0.24***	0.38***	0.17	-0.19
	(0.04)	(0.05)	(0.11)	(0.14)	(0.21)
(3) No internal IV	-0.67***	0.43***	1.13***	0.13	0.29
	(0.11)	(0.13)	(0.36)	(0.22)	(0.26)
(4) Use the percentages of minority population as the external IV	-0.28***	0.16**	0.09	-0.04	-0.25
	(0.04)	(0.06)	(0.06)	(0.10)	(0.39)
Sensitive to control variables					
(5) Include six more controls	-0.42***	0.69***	0.31*	0.30	1.45
	(0.12)	(0.15)	(0.18)	(0.22)	(11.24)
(6) Include year dummies	0.15	0.27*	0.59***	0.62***	0.02
	(0.11)	(0.14)	(0.18)	(0.21)	(0.24)
Sensitive to estimation methods					
(7) DIF-GMM	-0.40***	0.51***	0.13	0.78*	-0.13
	(0.04)	(0.07)	(0.18)	(0.42)	(0.13)
(8) Panel Regressions with fixed effects	-0.28**	0.41***	0.18	0.32	0.32**
	(0.12)	(0.12)	(0.19)	(0.23)	(0.14)
Sensitive to sub-samples					
(9) Drop provinces with a high share of minorities	-0.31***	0.23*	0.19*	0.36**	0.06
	(0.05)	(0.09)	(0.12)	(0.12)	(0.24)
(10) Drop sample before 1985	-0.50***	0.66***	0.43*	0.73**	2.98
	(0.04)	(0.08)	(0.13)	(0.07)	(0.37)

Notes: Heteroskedasticity robust *t*-statistics are reported in parentheses. *, **, and *** represent significance levels of 10%, 5%, and 1%.

Rows (5) and (6) check the sensitivity to omitted variables. Because we have a limited number of observations and adopt an IV approach, the baseline model includes only the most important covariables. As a robustness check, Row (5) includes six additional controls: net out-migration rate, youth dependency ratio, investment share, the growth of labor force share, trade share in GDP, and government spending share in GDP. We still find that the initial negative effect is smaller than the following positive effects, but the effect delayed by 15 years is no longer statistically significant. Row (6) includes a full set of year dummies, although we have accounted for inter-annual fluctuations by using five-year average data. We find that the initial effect becomes positive and statistically insignificant but the following effects are still positive and mostly statistically significant.

Rows (7) and (8) check the robustness against two alternative estimation methods, the first-differenced GMM (DIF-GMM) and the panel model with fixed effects. Compared with the SYS-GMM estimates in our baseline analyses, the DIF-GMM estimates are more vulnerable to the problems of weak instrument and finite sample biases (Bond 2002), while the panel fixed effect estimates cannot address the endogenous problem as that in the GMM frameworks. Using these two alternative estimation methods we still find that the initial negative effect of a higher birth rate is smaller than the following positive effect, although more lagged effects are statistically insignificant.

Rows (9)–(10) test the robustness against subsamples. Row (9) drops five provinces (Yunnan, Ningxia, Guizhou, Guangxi, and Qinghai) where more than one-third of the population were minorities considering that most minorities were not subjected to the OCP. Row (10) drops the sample before 1985 in consideration of the fact that the OCP during imposed a universal one-child limitation in some regions from

1980–1984 (see footnote 10). The subsample checks support our main finding that higher a birth rate reduces growth rates only during the first five years and the following positive effects overcome the initial negative effect.

Finally, Figure 3 shows the sensitivity of the results to the data from each individual province. We drop one province from the regression each time and plot the subsample estimates in the figure. Although dropping single provinces does have effects on the size of the estimates, the pattern of the effect is consistent: the initial effects (within five years) are negative and significant, the following effects (from 5–20 years) are positive and significant, and the effects from 20–25 years are generally positive and statistically insignificant.

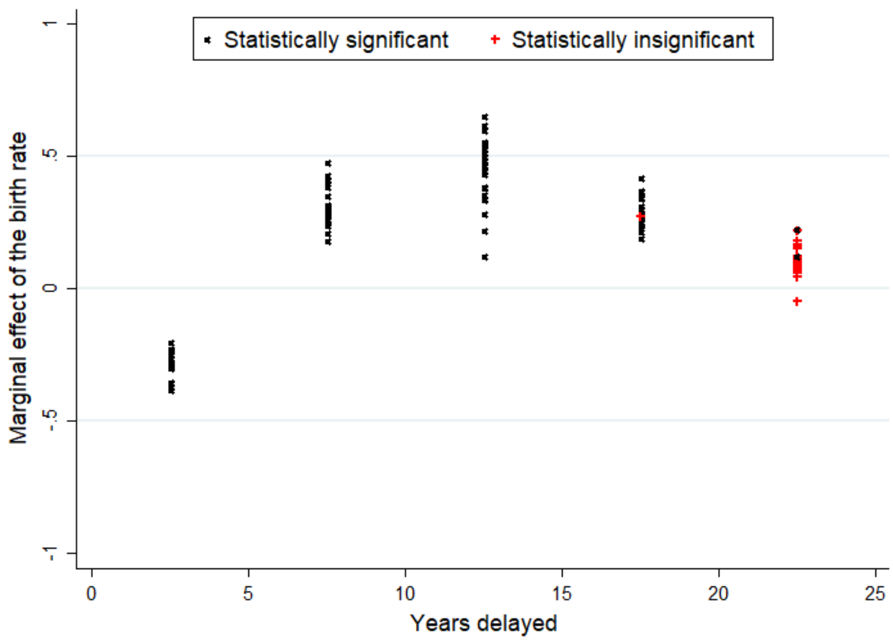

Figure 3. Sensitive to Sample Provinces

Note: We estimate the baseline model that excludes one province each time. Statistically significant (at 10% level) estimates are marked with a black x, and statistically insignificant estimates are marked with a red +.

4.3 The total effect of the OCP

The total economic effect of the OCP can be obtained by combining the marginal effect of fertility changes with OCP-induced fertility changes. However, the OCP-induced fertility changes are unknown, because fertility declines anyway as per capita income grows. For this reason, we propose approximating the OCP-induced fertility rate changes using the difference between Chinese fertility rates and those of other *developing* East Asian and Pacific economies. Without the OCP, China may have had the same fertility rate as these geographically adjacent economies due to similarly economic development levels and cultures. We find that the birth rate in China was consistently lower than that of the average of all *developing* East Asian and Pacific economies during the OCP, with a mean difference of 2.33 and a standard deviation of 0.37.[21]

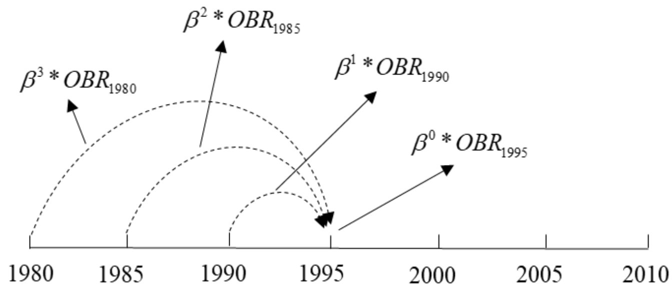

Figure 4. The effect of the OCP on the growth rate in 1995

Note: OBR_t denotes the changes in the birth rate caused by the OCP in year t. β^p is the estimated coefficient of the birth rate lagged by p periods. The data are in five-year intervals.

We combine the OCP-induced changes in fertility rates with the current and delayed marginal effects of the birth rate (the baseline estimates from Table 3) to calculate the

[21] To make it comparable, we exclude high-income East Asian and Pacific economies. The economies included for comparison are Cambodia, Indonesia, North Korea, Lao PDR, Malaysia, Mongolia, Myanmar, Philippines, Thailand, and Vietnam. The birth rate is calculated as the five-year average to match the Chinese data. The data was derived from the World Bank (2018).

yearly effect of the OCP on the growth rate of per capita GDP. Figure 4 uses 1995 as an example to illustrate this calculation. The growth rate in 1995 was affected by the OCP-induced birth rate declines in 1980, 1985, 1990, and 1995 (note that the data is in five-year intervals). Therefore, the 1995 effect is

(4) $$E_{1995} = 5\beta^3 * OBR_{1980} + 5\beta^2 * OBR_{1985} + 5\beta^1 * OBR_{1990} + 5\beta^0 * OBR_{1995},$$

where OBR_t is the OCP-induced changes in the birth rate in year t; β^0, β^1, β^2, and β^3 are the estimated coefficients of the birth rate that is lagged by 0, 5, 10, and 15 years; and we multiply each coefficient by 5 to reflect that the data is in five-year intervals.

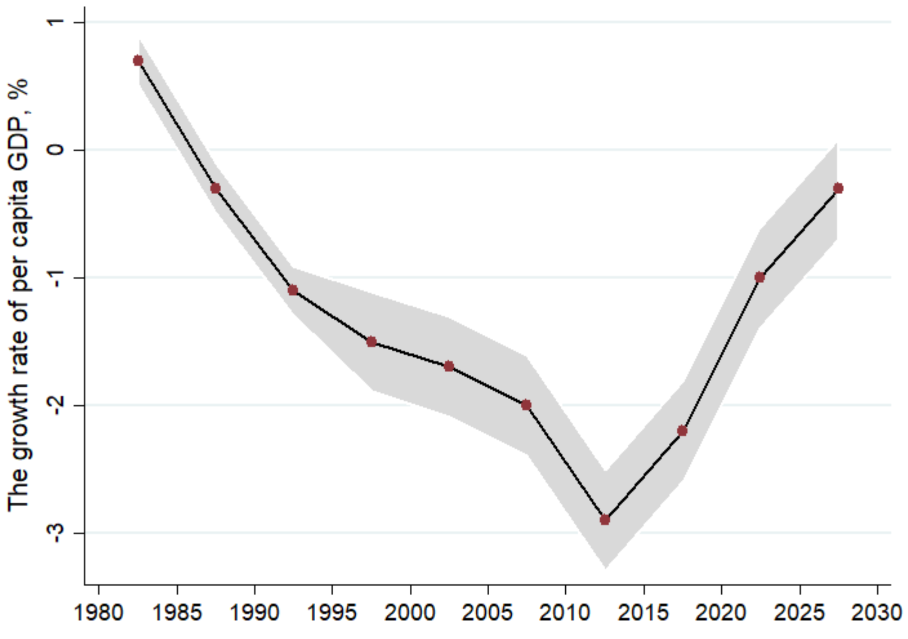

Figure 5. Yearly effect of the OCP on the growth rate of per capita GDP

Note: The 95% confidence interval (shadows in the figure) is calculated from the delta method. All values are calculated as five-year averages.

Figure 5 presents the yearly effect of the OCP on the *growth rate* of per capita GDP. The yearly effect is calculated up until the years 2025–2029 due to the delayed effect of the OCP. A positive yearly effect was only found during 1980–1984, which was 0.7%. The effect became negative during 1985–1989 (-0.3%), and this negative effect grew until it reached -2.9% in 2010–2014 and shrank until it reached -0.3% in 2025–2029. Based on these estimates, we find an average yearly effect of -1.2%. This is about 14.5% of the average per capita GDP growth rate during the OCP, which was 8.5%.

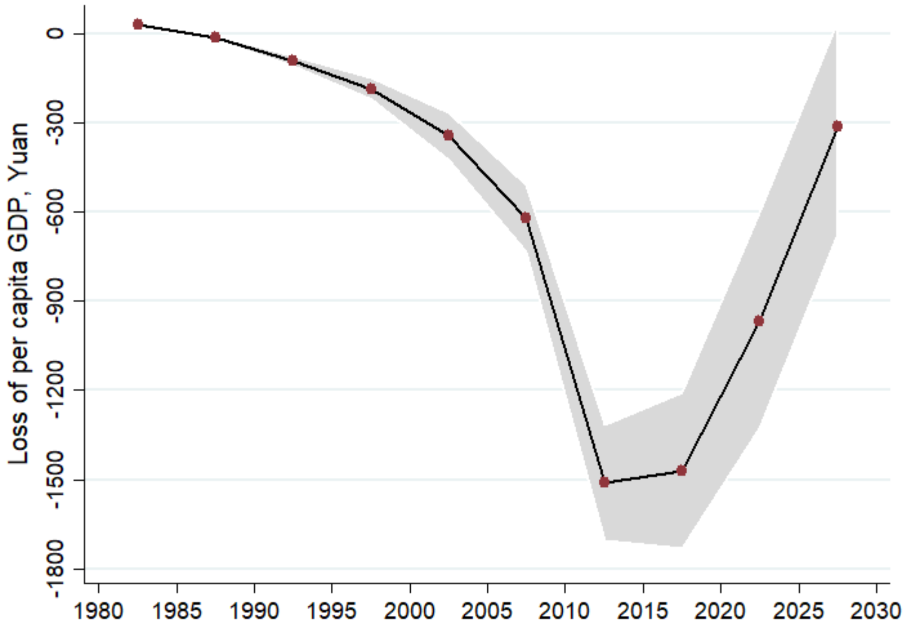

Figure 6. The effect of the OCP on per capita GDP (yuan, constant 2010 price)

Note: To calculate the yearly effect on the *level* of per capita GDP, we assume the per capita GDP without the OCP as y_0 and that with the OCP as y. We denote the effect of the OCP on yearly growth rate (presented in Figure 5) as r, and then we have $y = y_0(1+r)$. The yearly effect on per capita GDP is then calculated as $y - y_0$. The value of y after 2018 is predicted by linear extrapolation based on the data from 2011–2018 derived from the National Bureau of Statistics of China (2019). All values are calculated as five-year averages.

Figure 6 presents the yearly effect on the *level* of per capita GDP, which is calculated by combining the yearly effect on the growth rate with yearly per capita GDP (see the note on Figure 6 for further details). There was an extremely small initial positive yearly effect of only 27 yuan. The following negative effect increased from -14 to -1,512 yuan and then declined to -315 yuan. The yearly losses account for only a small share of the per capita GDP in China. For example, the largest yearly loss observed during 2010–2014 accounted for only 3.0% of the per capital GDP during the same period, which was 50,028 yuan.

However, the small yearly per capita effect adds up to an enormous total effect. When the yearly per capita effect is multiplied by China's population size, and then summed up over 1980–2029, a total loss of 38.2 trillion yuan (or 5.6 trillion U.S. dollars)[22] can be found with a 95% confidence interval of 36.8 to 39.6 trillion yuan. This loss is only slightly smaller than the real GDP in China in 2010, which was 41.0 trillion yuan.

5. Concluding remarks

This article estimates the effect of birth control on economic growth in China. After addressing endogeneity and accounting for long-run dynamic effects, We find that a unit decline in the birth rate caused by the OCP *increases* the current period's annual growth rate of per capita GDP by 0.29% but *reduces* the annual growth rate 5, 10, 15, and 20 years later by 0.30%, 0.46%, 0.30%, and 0.1%. The total loss in GDP caused by the 30-year OCP in China was as large as 38.2 trillion yuan with a 95% confidence interval of (-39.6, -36.8).

These findings provide important empirical evidence for evaluating national birth control policies that have swept much of the developing world since the 1960s.

[22] The population size is 1.06, 1.14, 1.21, 1.27, 1.31, 1.34, 1.37, 1.41, 1.45, and 1.49 billion in each of the 10 five-year intervals from 1980–2029. The population sizes after 2018 are predicted by linear extrapolation based on yearly data from 2011–2018 (National Bureau of Statistics of China 2019).

Although the main motivation of national family planning programs was to enhance income growth by reducing population growth, empirical evidence on this is extremely scarce. By launching the nationwide OCP, China established an unprecedented level of governmental birth control. However, evidence from the OCP still offers valuable lessons for other countries that still with high rates of population growth and adopt some sort birth control policies.

This article also contributes to estimate the causal effect of population growth on economic growth. The theoretical literature has debated on this at least since Thomas Malthus (1798), though the empirical evidence is mixed. By addressing endogenous fertility, this article finds results echoing the believe of “population optimists” that population growth promotes economic growth.

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Appendix: Explanatory power of IVs

As previously mentioned, there were two major statutory differences under the OCP. First, couples with a non-agricultural *hukou* were generally allowed to have only one child, but couples with an agricultural *hukou* were allowed to have a second child if the first child was a girl. Second, couples who belong to an ethnic minority population group (except, since the 1990s, minorities of *Zhuang* and *Manchu*) were allowed to have more than one child. These statutory fertility differences suggest two potential IVs for the birth rate in the SYS-GMM estimation: the percentage of residents with agricultural *hukou* (PAH) and the percentage of minorities in the population (PMP). This section explains why we only use the PAH in our main analysis.

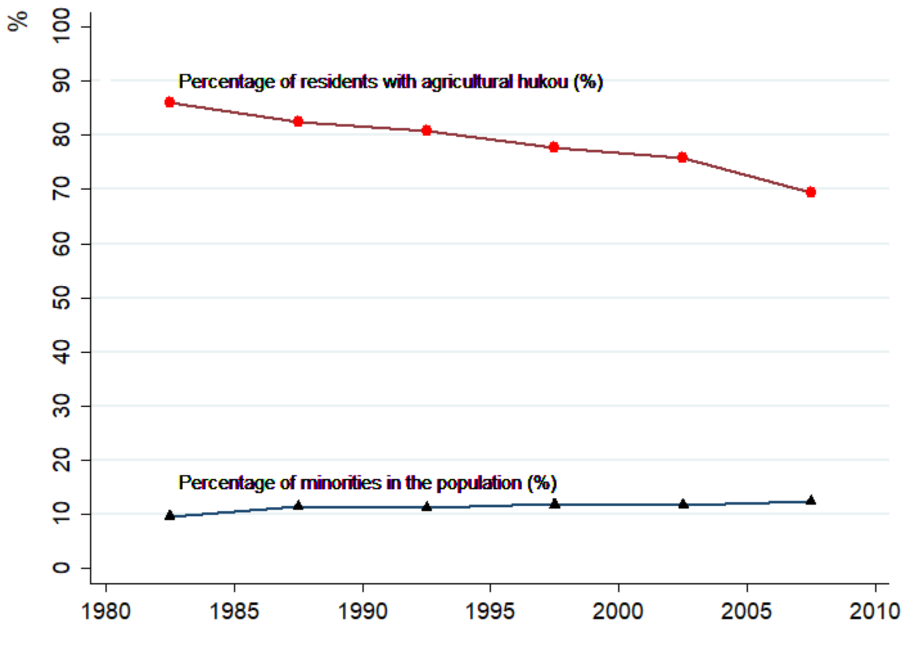

Figure A1. Country averages of the PAH (red dot) and PMP (black triangle) from 1980 to 2010

As shown in Figure A1, minorities comprised approximately 11% of the population in China, and the PMP changed only 2.7% over the 30 years of the OCP. The PMP drops to 8.4% if we exclude the two minority groups (Zhuang and Manchu) that were subject to the sample OCP as Han Chinese since the 1990s.[23] The low levels and small variations of the PMP suggest that it is likely a weak IV. Column (2) of Table A1 confirms this concern: we find no statistically significant effect of the first difference of the PMP on the first difference of the birth rate.[24] In contrast, Figure A1 also shows that the PAH was generally larger than 70% and changed 16.5% during the OCP, and Column (1) of Table A1 shows that the first difference of the PAH has a statistically significant effect on the first difference of the birth rate.

An equally important concern of using the PMP as an IV is that most of the minorities were living in several non-representative western provinces. The upper panel of Figure A2 shows that nearly all provinces with a PMP higher than 20% were western provinces. The lower panel shows that, although these western minority provinces cover about half of China's territory, they comprise only a small proportion of the total population in China (about 10% in total). Similarly, these western provinces comprised only 9.5% of the total GDP in China (not shown in the figure). It seems difficult to infer the fertility-growth causality in China from these western minority provinces that have much lower population and economic densities.

[23] In 2010, Zhuang accounted for 14.2% of the minority population of China while Manchu accounted for 9.3% (National Population Census of the PRC, 2010).

[24] Note that the SYS-GMM model is a first differenced model and that the external IV should also be in the first difference.

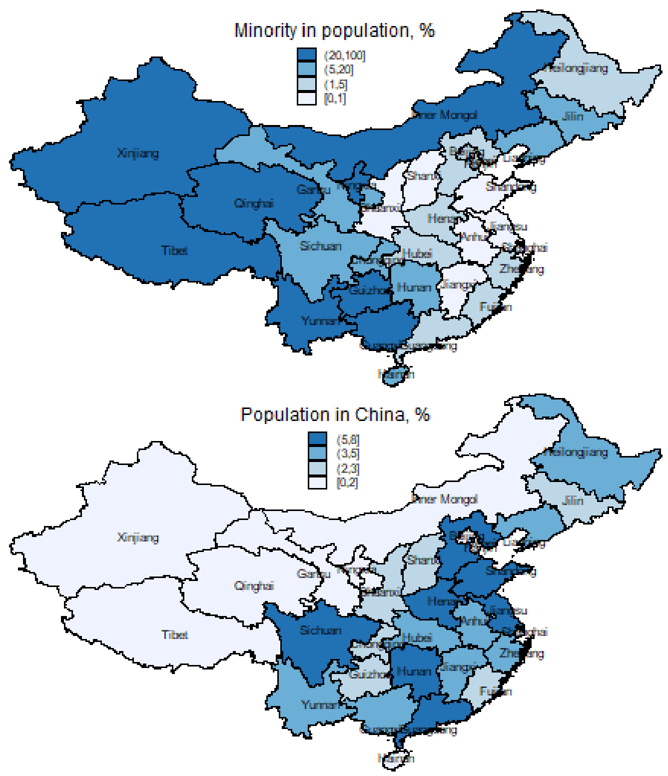

Figure A2. The percentage of minorities in each province (upper panel) and the percentage of each province's population in China (lower panel), 1981–2010

Note: The maps show the 31 mainland Chinese provincial districts.

Table A1 tests the explanatory power of the IVs. Considering that the SYS-GMM has no first-stage regression, these tests can be seen as the first stage. Because the instruments are applied to the first differenced equation in the SYS-GMM estimation, we test the explanatory power by regressing the first difference of the birth rate on the IVs. Specifically, Columns (1) and (2) regress the first difference of the birth rate on the first difference of the PAH and the first difference of the PMP, respectively; and Column (3) regresses the first difference of the birth rate on the birth rates lagged by more than one period. Although we find statistically significant (at least at the 1% level) explanatory power for the PAH and the lagged birth rates, the PMP has no significant explanatory power on the birth rate.

Table A1. The explaining power of IVs for the first difference of the birth rate

	Dependent variable: first difference of the birth rate		
Independent variables	(1)	(2)	(3)
First difference of the PAH	0.27***		
	(0.10)		
First difference of the PMP		-0.28	
		(0.26)	
10-year lagged birth rate			-0.57***
			(0.10)
15-year lagged birth rate			0.31*
			(0.18)
20-year lagged birth rate			0.15
			(0.10)
25-year lagged birth rate			-0.00
			(0.08)
Joint significance test of IVs (P-value)	0.01	0.29	<0.001
Observations	115	115	22

Notes: Standard error clustering at the province level is reported in parentheses. * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$.