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Effects of Vietnam's two-child policy on fertility, son preference, and female labor supply

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Abstract
In 1988, facing a total fertility rate of over four births per woman, the Vietnamese government introduced a new policy that required parents to have no more than two children. Using data from the Vietnam Population and Housing Censuses from 1989, 1999, and 2009, I apply a differences-in-differences framework to assess the effects of this policy on family size, son preference, and maternal employment. I find that the policy decreased the probability that a woman has more than two children by 15 percentage points for younger women and by 7 percentage points for middle-aged women. The policy reduced the average number of living children by 0.2 births per woman. Low-education women and women in rural areas were more affected by the policy. The policy had no effects on mothers' age at first birth and gender of mothers' last birth. The reduction in fertility caused by the policy was associated with a 1.2 percentage point decrease in the proportion of sons in each family. The policy increased maternal employment by 1.3 percentage points. Instrumental variables estimates of the effects of fertility on maternal employment and child education suggest a negative relationship between the number of children and female labor supply and a trade-off between child quantity and child quality in Vietnam.

Keywords Two-child policy · Fertility · Son preference · Female labor supply · Child quality · Child quantity

JEL classification J13 · J18 · J21

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1 Introduction

High fertility rates and low economic growth are prevalent problems in developing countries. Family planning policies are considered a solution to these problems. In both China and India, the two most populous countries in the world, the governments have relied on family planning policies to limit family size.[1] In the 1970s, China implemented its "later, longer, fewer" policies, which encouraged women to marry at a later age, wait more than 3 years between births, and have at most two children (Chen and Huang 2018). The policies were technically voluntary and were in effect until 1979, when they were replaced by the one-child policy, which was not voluntary (Chen and Huang 2018). Research shows that the "later, longer, fewer" policies played a more important role in China's demographic transition than China's one-child policy (Miller and Babiarz 2016; Whyte et al. 2015; Wang 2016; Chen and Huang 2018). Chen and Huang (2018) examine the effects of the "later, longer, fewer" policies on fertility and document that the policies accounted for half of the decline in China's total fertility rate during 1969 and 1978. Chen and Fang (2018) also examine the long-term consequences of the "later, longer, fewer" policies on fertility and physical and mental well-being of affected parents. The authors document that the policies significantly decreased the number of children born by 0.85 births per couple. In addition, parents who were more exposed to the policies experienced a higher level of consumption, better physical health, and lower expenses on medical services but were more likely to be depressed. Together, these studies suggest the effectiveness of the "later, longer, fewer" policies in reducing fertility and their different effects on physical and mental well-being of affected parents.

In the 1980s, China relaxed its one-child policy and allowed couples in rural areas to have a second child if the first birth was a girl (Greenlaugh 1986). During that period, four regions in China also implemented comprehensive two-child policies in their rural areas (Wang et al. 2017). Gu and Wang (2009), Wei and Zhang (2014), and Qin and Wang (2017) examine the effects of the two-child policies in these regions and mostly find limited effects of the policies on fertility. At the start of 2016, China relaxed its one-child policy and expanded it to the two-child policy with the hopes of bringing its fertility rates back to replacement level. More time is needed to fully examine the effects of the expansion of China's current two-child policy (Wang et al. 2017).

Although a number of papers have investigated the effects of the two-child policies in China on fertility, evidence on the effects of the policies on fertility is mixed. While Whyte et al. (2015), Wang (2016), Chen and Huang (2018), and Chen and Fang (2018) document significant effects of the "later, longer, fewer" policies on fertility, Gu and Wang (2009), Wei and Zhang (2014), and Qin and Wang et al. (2017) find limited long-term effects of the two-child policy in the mid-1980s on fertility. Thus, more evidence is needed on the policy's effects. Understanding the effect of the two-child policy on family size would facilitate evaluating the effects of the policy on other substantive issues, such as child quality, parental labor supply, and ultimately economic prosperity. Studying the consequences of the two-child policy also has relevance to understanding central issues in economics such as the trade-off between child quality and quantity and

[1] Family planning in India is usually criticized for being unsystematic and ineffective (Visaria et al. 1999; Muttreja and Singh 2018).

the causal relationship between fertility and labor supply. In addition, in developing countries where there is a strong preference for sons and where sons act as social security for parents in their old age, the two-child policy may affect the gender balance as well.

In 1988, facing a total fertility rate of over four births per woman, the Vietnamese government introduced a new policy that required parents to have no more than two children. The motivation for the policy was that with fewer children per woman, population growth would be reduced and women might spend less time on home production and more time in the labor force, which might promote economic growth (Council of Ministers 1989; Brander and Dowrick 1994). With fewer children, women also have less ability to achieve their desired number of sons and thus may have a lower proportion of sons in the family. The evidence on the effect of Vietnam's two-child policy is limited to a few small studies. For example, Goodkind (1995) undertakes a field survey in two provinces in Vietnam and find that the policy was enforced more in the North than in the South. Hoa et al. (1996) conduct a cross-sectional survey in Tien Hai, a district in one of the most densely populated provinces in Vietnam, and report that most families did not seem to follow the policy. Scornet (2000) points out that a forceful nationally defined demographic policy like Vietnam's two-child policy should take into account the local social and economic context, suggesting that effects may differ across areas. The author finds that while family planning policies in the region of the Red River Delta in Vietnam have greater effects on fertility decline than social and economic modernization, the faster economic development in the South seems to play a more important role in fertility decline than family planning programs. The limited geographic coverage of these studies limits the external validity of their findings. In addition, these studies do not account for other factors, such as social and economic changes that may affect parental preferences and the cost of raising children. Thus, the findings from these studies are at best descriptive and may be misleading in terms of identifying the causal effect of the policy on fertility.

In this article, I examine the effects of Vietnam's two-child policy on family size, parents' ability to achieve son preference, and maternal labor supply. I also use the two-child policy as an instrumental variable to investigate the causal effect of the number of children on maternal employment and the trade-off between child quantity and child quality in Vietnam. To measure the effects of the policy, I use data from the Vietnam Population and Housing Censuses from 1989, 1999, and 2009 with a differences-in-differences framework. The exposure of a woman to the policy is determined by both her ethnicity and her age in 1989. Because the policy does not apply to ethnic minorities, I treat ethnic minorities as the control group and ethnic majorities as the treatment group. Women's age in 1989 identifies the length of exposure to the policy. A woman who was 25 in 1989 would be affected more by the policy than a woman who was 40 or older at that time since a 40-year-old woman should have finished her fertility by the time that the policy was in place. The fundamental assumption underlying the differences-in-differences research design is that, in the absence of the policy, changes over time (birth year) in fertility and employment are the same for both ethnic groups. Under the assumption that fertility is the only channel through which the policy could affect maternal employment and child education, exogenous variations in family size caused by the two-child policy can be used to construct instrumental variables (IV) estimates of the effect of fertility on maternal employment and child education. Below, I provide evidence that these assumptions are likely to hold in this context.

The results of my analysis suggest that the policy decreased the probability that a woman has more than two children by 15 percentage points (50%) for women under age 30 in 1989 and by 7 percentage points (11.5%) for women ages 30–39 in 1989. The policy reduced the average number of living children by 0.2 births per woman (10%). Low-education women and women in rural areas were more affected by the policy. The policy had no significant effects on mothers' age at first birth and the gender of mothers' last birth. The policy increased the probability that mothers have longer birth spacing (at least 3 years) by 10.5 percentage points (16.67%) for younger women in 1989. On average, the policy increased birth spacing between the first two births of younger mothers by 0.81 years (20.56%). The reduction in family size caused by the policy was associated with a 1.2 percentage point (2.4%) decrease in the proportion of sons in each family. The policy increased maternal employment by 1.3 percentage points (1.5%). The IV estimates of the causal effect of family size on maternal employment indicate that having an additional child decreased maternal employment by 15 percentage points (17.4%). The IV estimates of the effects of child quantity on child quality indicate that having more than two children significantly decreased the average years of education of all children in the household as well as school enrollment and education of the first two children.

In addition to providing estimates of the causal effects of Vietnam's family planning policy, my paper also contributes to the literature in a few other ways. First, my paper presents new evidence on the effectiveness of family planning policies at reducing fertility. My paper is the first to examine the effects of Vietnam's two-child policy on other outcomes, such as parents' ability to achieve son preference, sex ratio at last birth, and maternal employment. My study is also the first to use Vietnam's two-child policy as an instrument for fertility to examine the relationship between fertility and maternal employment and the quantity-quality trade-off in Vietnam. In addition, I show that reductions in fertility caused by the two-child policy are associated with a decline in parents' ability in achieving son preference, which provides evidence that the implementation of Vietnam's two-child policy did not lead to a sex ratio imbalance as the one-child policy in China did (Li et al. 2011). Moreover, the recent adoption of a two-child policy in China strengthens the contribution of my study because of the need for additional evidence on the consequences of a two-child policy in East Asia. My study provides new evidence on the trade-off between child quantity and child quality. My results suggest that having more than two children significantly decreased child education. My findings have important policy implications for developing countries where the governments rely on family planning policies to curb population growth and promote economic development. My results suggest that while the policy was effective at reducing family size, it had smaller effects on female labor supply. Thus, family planning policies may not promote economic growth through the labor supply channel as governments expect.

2 The two-child policy in Vietnam

The two-child policy was recommended by the Vietnamese government in 1981 and made law in 1988. The goal of the policy is to maintain national population growth at 2 percent (Council of Ministers 1989). The policy is applied to every family except for families of ethnic minorities. Families of ethnic minorities could have a third child if they desire (Council of Ministers 1989). Couples who already have one child but have twins or triplets in the second birth are not considered to violate the policy. The specific guidelines of the policy also included requirements on the minimum childbearing age and the birth spacing (Council of Ministers 1989). For state employees and government officials, childbearing age should be 22 or older for women and 24 or older for men. For others, the childbearing age should be 19 or older for women and 21 or older for men. The second child, if desired, should be spaced 3 to 5 years apart from the first one, except for women aged 30 and older (Council of Ministers 1989).[2]

To promote the two-child policy, the Vietnamese government imposed fines and punishments on families that violated the policy. The government denied a third child a birth certificate (Population Research Institute 1995). If families violated the policy, they would be fined about $80, which was equivalent to 10 months of income in 1995 (Population Research Institute 1995). State employees and government officials would not get promoted or would be relegated to lower-status jobs in smaller cities or, in some cases, would lose their jobs if they violated the policy (Nikkei Asian Review 2017). The government subsidized the fees of housing, healthcare, and education for the first two children, but not for the third child. Families with more than two children had to pay extra fees for housing, education, and healthcare of the third child (Council of Ministers 1989). With these fines and punishments, the policy imposed real costs on families that have a third child, and there is a plausible expectation that the policy reduced fertility.

Besides imposing fines, the government also engaged in public health strategies, such as posters and billboards that depicted happier families with fewer children. Television programs that provided information on family planning were shown several times per week and at prime time (Goodkind 1995). The government offered a reward of $20 to women who had a hysterectomy, "a procedure that approximately half of all village women were subjected" (Population Research Institute 1995). The government also supplied birth control devices and birth control pills at free of charge to eligible people and poor families that registered to practice family planning (Council of Ministers 1989).

3 Conceptual framework

In this section, I discuss economic theories on demand for children, son preference, and the effect of children on female labor supply to highlight mechanisms through which the two-child policy could affect family size, son preference, and maternal employment.

3.1 Demand for children

In developing countries where agriculture plays an important role in the economy and people cannot rely on social security or retirement plans in their old age, parents want to

[2] These features of the policy were not equally enforced as the restriction to have no more than two children since the government did not impose punishments or fines on violating these requirements.

have children as social security for their old age (Priebe 2010). In addition, having more children increases the probability that one of them may be successful (high quality) and can take care of parents in their old age. Especially in rural areas where most economic activities are farm work, having more children implies that families have more labor to work in the farm in the future. However, with harsh punishments that the government imposes on families that violate the two-child policy, families face a higher cost of having a third child. As a result, families would not want to have more than two children as they did before. In other words, the policy is likely to cause a reduction in family size.

Since the policy also requires that the second child should be spaced 3 to 5 years apart from the first child, it is reasonable to expect that the policy may increase birth spacing between the first two births. The policy may affect mothers' age at first birth as well. Women growing up under the two-child policy know that they will have only two children. As a result, they may delay having their first birth at a later age. In addition, the policy requires the minimum childbearing age of 22 for women who are state employees and government officials and of 19 for others. Thus, it is plausible to expect that the policy increase mothers' age at first birth.

In the standard quantity and quality model of fertility, parents derive utility from both the quantity and quality of children (Becker and Lewis 1973; Becker and Tomes 1976; Rosenzweig and Wolpin 1980). This theory predicts that when child quantity decreases exogenously, child quality increases. Thus, if the two-child policy decreases family size, then it is expected to increase child quality as well. In sum, the basic theory of the demand for children predicts that the two-child policy of Vietnam has the potential to not only decrease family size but also raise the human capital of children.

The effect of the two-child policy on family size may vary across mothers' schooling and urban/rural areas. In theory, it is ambiguous whether low- or high-educated women will be affected more by the policy. Well-educated women tend to have a higher wage and thus face a higher cost of having a third child (Ebenstein 2009). Low-educated women, in contrast, may live with elderly in laws and thus face a lower marginal cost of having an additional child (Ebenstein 2009; Caceres-Delpiano 2012). With a relatively higher marginal cost of having an additional child, well-educated women tend to prefer a smaller family size. Thus, it is likely to observe a bigger effect of the policy on them. On the other hand, high-educated women are likely to have higher income. The fines and punishments imposed for having a third child may account for a relatively smaller portion of their income than for less educated, poorer women. Thus, it is also plausible to expect that high-educated women are less responsive to the policy. Due to these confounding factors, it is theoretically ambiguous whether the effect of the policy will be larger for more or less educated women.

Similarly, it is difficult to determine whether the policy will be more effective in urban or rural areas. In urban areas, women tend to work on paid jobs rather than on farm work or housework. Thus, they face a higher marginal cost of having an addition child (i.e., higher foregone earning of the childbearing time). In addition, it is more expensive to raise children in urban areas (i.e., higher childcare costs, higher education fees). On the other hand, women in rural areas are in need of having more children to work in the farm and at least one son to support them in their old age (Priebe 2010). Thus, they may keep having births despite of the fines and punishments imposed by the policy. As a result, it is likely to observe a bigger effect of the policy on women in urban areas. It is also plausible to expect that women in urban areas are less responsive to the policy. Women in urban areas tend to have higher income, and the fines and punishments imposed by the policy may be a small proportion of their income. In contrast, women in rural areas are poorer, and the fines and punishments may account for a big share of their income. Thus, these women may stop at two children. Because of the abovementioned factors, it is also theoretically ambiguous whether the policy will have a larger effect on women in urban or rural areas.

3.2 Son preference and son targeting fertility behavior

In many developing countries in East, South, and Southeast Asia, parents have a strong preference for sons (Haughton and Haughton 1998; Arnold et al. 1998; Clark 2000; Jensen 2003; Belanger 2002; Gupta et al. 2003; Jayachandran and Kuziemko 2011). One of the reasons that parents prefer boys to girls in these countries is because sons act as social security for parents in their old age (Larsen et al. 1998; Lundberg 2005; Filmer et al. 2009). Due to social and culture norms, sons will live and support their parents in the future (Larsen et al. 1998; Lundberg 2005; Filmer et al. 2009). This strong son preference has been documented as differential stopping behavior (DSB) or male-preferring stopping rules in the literature (Clark 2000). Under these rules, women will continue having births until they achieve their desired number of sons or hit their maximum number of children given their budget constraint (Clark 2000). The gender composition of current living children also identifies their mothers' decisions to have more children (Andersson et al. 2006; Jayaraman et al. 2009; Basu and Jong 2010). If a woman desires for two sons and already has both of them at the first two attempts, she will stop at two children. However, if she has achieved only one son, she may continue to three children to obtain the second son.

Vietnam's two-child policy may decrease the ability of parents to achieve son preference. The policy would decrease the proportion of sons in each family because it lowers the number of children that a woman may have and places an additional constraint on her ability to achieve the desired number of sons. Specifically, families that want at least one son may end up with no sons if they do not have one at the first two births. Similarly, families that desire two sons may have no sons or only one son if their first- and second-born children are two girls or one boy and one girl. In these cases, the proportions of sons in these families will decrease.

In addition, due to the former low quality of ultrasounds in the country, sex-selective abortions did not start after 2003 (Bélanger et al. 2003; Becquet and Guilmoto 2013; Guilmoto et al. 2018). Thus, in the 1990s, parents could not turn to sex-selective abortions as an alternative to have at least one son. As a result, it is reasonable to expect that the policy have no effects on the probability that mothers have a boy at last birth.

3.3 The effect of children on female labor supply

In the standard labor leisure model augmented to include a desire for children, women decide their fertility and their labor supply at the same time. Children require mothers' time and increase the value of household work. Therefore, the more children a woman has, the less likely she is going to work. Vietnam's two-child policy is an exogenous change in fertility. With the presence of the policy, women are less likely to have more than two children. For those affected, the desired number of children is above the realized number. With fewer children, the demands of childbearing and childrearing on women's time may fall, which potentially gives them more time for increased market work. In addition, with fewer children, mothers' productivity at home may decline, thus lowering the value of nonmarket time. Therefore, the two-child policy is expected to increase maternal employment.

The effect of fertility on labor supply may vary with mothers' schooling (Angrist and Evans 1998) and across urban and rural areas. A number of empirical studies found that more educated women were more responsive to a fertility shock than less educated women (Gronau 1986; Caceres-Delpiano 2012). Gronau (1973) also documents that as mothers obtain more education, children have a stronger effect on their mothers' value of time. Therefore, I expect that changes in family size due to the two-child policy will have a larger effect on more educated women. Women in urban areas may be more responsive to a shock in family size than women in rural areas. Women in urban areas tend to have more job opportunities and thus face a higher marginal cost of having a third child (Ebenstein 2009). On the other hand, women in rural areas work at home and have a relatively lower marginal cost of having an additional child (Ebenstein 2009). These considerations suggest that the effect of Vietnam's two-child policy on maternal labor supply may be heterogeneous and differ by rural/urban residence (Priebe 2010[3]; Caceres-Delpiano 2012[4]). In this circumstance, it is plausible to expect that the policy have a larger effect on maternal employment of high educated women and women in urban areas. As discussed above, these women face higher opportunity costs of childbearing and thus they would be more responsive to the fertility shock caused by the two-child policy. Specifically, in this context, they would be more likely to increase their labor supply in response to a reduction in family size caused by the two-child policy.

4 Data

4.1 The Vietnam population and housing censuses

The data used in this study are from the Vietnam Population and Housing Censuses from 1989, 1999, and 2009.[5] The data are from the 5%, 3%, and 15% nationally representative samples of the population (Minnesota Population Center 2017). The surveys include information on the number of children ever born and the number of surviving children of women ages 15–49 at the time of survey. The surveys also contain other relevant information on individuals' characteristics such as age, ethnicity, marital status, educational attainment, work, and current place of residency (at provincial levels).

[3] Priebe (2010) examines the effect of fertility on maternal employment in Indonesia and documents that less educated women and women in rural areas were more responsive to the presence of children.

[4] Caceres-Delpiano (2012) investigates the impact of children on maternal employment in 40 developing countries and finds that the impact of children is stronger among high-educated mothers and mothers in urban areas.

[5] Access to the Vietnam Population and Housing Census from 1979 is not publicly available. Thus, I do not use the 1979 census in this study.

The analyses are conducted using two samples drawn from these three surveys. To examine the effect of the policy on women's fertility and maternal labor supply, I use what I refer to as full sample, which includes women ages 10–49 in 1989. Since the surveys ask fertility questions for women at the childbearing age and the key measure of treatment (exposure) is age in 1989, I do not observe every birth cohort (age in 1989) in all of the three censuses. Specifically, I only observe fertility of women ages 40–49 in 1989 in the survey year 1989. Similarly, I only observe fertility of women ages 30–39 in 1989 in the survey years 1989 and 1999 (ages 40–49) and fertility of women under age 15 in 1989 in the survey years 1999 and 2009. Finally, for women ages 15 to 29 in 1989, I observe their fertility in 1989, 1999 (ages 25 to 39), and 2009 (ages 35 to 49).

To investigate the impact of the policy on the proportion of sons in each family, birth spacing, gender at mothers' last birth, and the effect of having more than two children on child education, I use a different sample – what I refer to as subsample – because this analysis requires that I know the gender of the children. Although the data asks every woman at the childbearing age about the number of children ever born and surviving, they do not ask any further information about these children for these women. The subsample includes women ages 10–49 in 1989 who are household heads or wives of household heads and have all of their children living home.[6] For women who are household heads or wives of the household head, the surveys ask questions about the gender of all children who are currently living with them. Since the surveys have no information on children living outside home, the gender of children who already left the household is unknown. By imposing these two restrictions, there is a chance that I have selected the sample nonrandomly with respect to the number and gender of the child. However, this is unlikely to bias my estimates. I show that when I use this subsample to estimate the effect of the policy on family size, I obtain very similar estimates to those that I get from full sample.[7] This indicates that the potential sample selection problem is not likely to be problematic. In addition, the subsample represents 80% of full sample.

4.2 Measures of family size, parents' ability to achieve son preference, and maternal labor supply

Family size is measured as the probability that a woman has more than two children and the number of living children. More than two children equals 1 if a woman has more than two children at the time of survey and 0 otherwise. Parents' ability to achieve son preference is measured as the proportion of sons in each family, which is the ratio of the number of sons to the number of living children. Birth spacing between the first two births is derived as the difference between age of the first child and age of the second child. Birth spacing at least 3 years equals 1 if mothers' birth spacing between the first two births is equal to or greater than 3 years and 0 otherwise. Boy at last birth equals 1 if the mother reports to have a son at last birth and 0 otherwise.

Maternal employment is measured based on the activities in the last 12 months. The dummy variable of maternal employment is coded as 1 if the mother reports that she was employed in the last 12 months and 0 otherwise. The average years of education of

[6] Appendix Table 10 shows the fraction of mothers that still have all of their children living at home across three survey years.

[7] The estimates of the effect of the policy on family size for subsample are shown in Appendix Figure 11.

children in the household are the ratio of total years of education of all children to the number of living children. The average years of education of the first two children are the ratio of years of education of the first two children to two. The dummy variables of being enrolled in school and having at least primary education equal 1 if the first/second born child is enrolled in school or has at least primary education, and 0 otherwise.

4.3 Summary statistics

Table 1 presents summary statistics for women in full sample and subsample.[8] As Table 1 indicates, women in full sample are slightly older than are women in subsample. The average age of women is 35.95 for full sample and 33.62 for subsample. Women in full sample also have a higher number of living children and a higher probability of having more than two children than women in subsample. The average number of living children is 2.17 for women in full sample and 2.08 for women in subsample. While 34% of women in full sample have more than two children, 28% of women in subsample have three children or more. About 85%–87% of women in two samples are employed. The average age at first birth of mothers in subsample is 23.07. The proportion of sons in each family is 0.51.

The average birth spacing between the first two births is 3.78 years. Approximately 61% of mothers in the subsample have birth spacing between the first two births equal or greater than 3 years. The average years of education of all children in the household are 4.06 years. On average, the first two children have 4.26 years of education. While 95% of mothers in the subsample have their first- and second-born children enrolled in school, approximately 86% have their first two children with at least primary education. Approximately 65% of women in full sample and 69% of women in subsample live in rural areas. In terms of educational levels, 33%–34% of women in both samples have less than primary education.

5 Research design

5.1 The first stage: the effect of the two-child policy on family size and parents' ability to achieve son preference

To examine the effect of the two-child policy on family size, I use a differences-in-differences framework. The analysis is performed using full sample. Since the policy does not apply to ethnic minorities, I treat ethnic minorities as the control group and ethnic majorities as the treatment group. The length of the exposure to the policy is determined by the age of women in 1989. Younger (e.g., < 30) women in 1989 should be affected more by the policy than older (e.g., > 39) women who are closer to have completed fertility by that age. In practice, I use women's age in 1989 as a continuous variable to determine the length of the exposure to the policy. As an alternative specification, I also group women in three different age groups (under age 30, 30–39, and 40 and older in 1989).

[8] I also show summary statistics for women of both ethnic groups and by education and urban/rural status separately in Appendix Tables 11, 12, 13, 14, 15, 16.

The average probabilities of having more than two children and the average number of living children for four groups (Y_i) are derived as below:

Samples	Ethnic majority women	Ethnic minority women
Younger women in 1989	E (Y_i | Majority = 1, Younger = 1)	E (Y_i | Majority = 0, Younger =1)
Older women in 1989	E (Y_i | Majority = 1, Younger = 0)	E (Y_i | Majority = 0, Younger = 0)

The change in family size for ethnic majorities is

$$[E(Y_i|\text{Majority} = 1|\text{Younger} = 1)–E(Y_i|\text{Majority} = 1|\text{Younger} = 0)]$$
$$= \text{cohort effects}_{\text{majority}} + \text{policy effects}$$

Similarly, the change in family size for ethnic minorities is

$$[E(Y_i|\text{Majority} = 0|\text{Younger} = 1)-E(Y_i|\text{Majority} = 0|\text{Younger} = 0)]$$
$$= \text{cohort effects}_{\text{minority}}$$

My differences-in-differences (DID) estimates are as follows:

$$\begin{aligned}\text{DID} = &[E(Y_i|\text{Majority} = 1|\text{Younger} = 1)–E(Y_i|\text{Majority} = 1|\text{Younger} = 0)] \\ &–[E(Y_i|\text{Majority} = 0|\text{Younger} = 1)–E(Y_i|\text{Majority} = 0|\text{Younger} = 0)] \\ &= \left(\text{cohort effects}_{\text{majority}} + \text{policy effects}\right)–\left(\text{cohort effects}_{\text{minority}}\right) \\ &= \left(\text{cohort effects}_{\text{majority}}-\text{cohort effects}_{\text{minority}}\right) + \text{policy effects}\end{aligned} \tag{1}$$

Under the assumption that cohort effects $_{\text{majority}}$ = cohort effects $_{\text{minority}}$, my DID estimates will capture the causal effect of the two-child policy on family size.

In practice, I use the following regression model to apply my differences-in-differences framework:

$$\begin{aligned}Y_{ijt} = &b_0 + b_1\text{age}_{it} + b_2\text{age}^2{}_{it} + b_3\ \text{age}_{it} \times \text{majority}_i + b_4\ \text{age}^2{}_{it} \times \text{majority}_i \\ &+ \textstyle\sum_{j=10}^{48} b_{5j}\text{age in } 1989_{ij} + b_6\ \text{majority}_i + \sum_{j=10}^{48} b_{7j}\text{age in } 1989_{ij} \\ &\times \text{majority}_i + \text{province}_{ijt} + v_{ijt}\end{aligned} \tag{2}$$

in which i = 1,..., N (index of person)

j = 10,…, 49 (index of age in 1989)

t = survey years 1989, 1999, 2009

Y_{ijt} is having more than two children, the number of living children, mothers' age at first birth, birth spacing, the proportion of sons in each family, and gender at last birth. The omitted group is women ages 49 in 1989. To account for differences in parental preferences and the costs of raising children across provinces, I also include provincial dummies in Eq. (2).

Table 1 Summary statistics of women's birth cohorts 1940–1979

Variables	Full sample		Subsample	
	Mean or %	SD	Mean or %	SD
Number of living children	2.173	1.570	2.079	1.248
% who have				
No children	15.70		7.23	
One child	15.04		23	
2 children	34.91		42.31	
3 or more children	34.35		27.46	
Proportion of sons in the family			0.506	0.394
Employed	0.856	0.351	0.872	0.334
Age	35.948	8.193	33.615	7.458
Maternal age at first birth			23.067	3.914
Birth spacing between the first two births			3.82	2.62
Birth spacing between the first two births at least 3 years			0.61	0.49
Last birth is a boy			0.55	0.50
Average years of education of all children in the household			4.06	3.45
Average years of education of the first two children			4.26	3.54
1st child enrolled in school			0.95	0.22
2nd child enrolled in school			0.94	0.23
1st child has at least primary education			0.87	0.34
2nd child has at least primary education			0.86	0.35
Rural	0.654	0.476	0.690	0.462
Less than primary education	0.340	0.474	0.327	0.469
Number of obs.	3,197,622		1,852,725	

Note: Full sample includes all women ages 10–49 in 1989 in 3 survey years. Subsample includes women ages 10–49 in 1989 who satisfy the following conditions: (1) they are household heads or wives of household heads, and (2) they have all children living with them

The dummy variables indicating women's age in 1989 measure the length of exposure to the policy and an individual's birth cohort. The indicator of ethnic majority defines the treatment and control groups. The coefficients of interest are thus on the interaction terms between dummy variables of women's age in 1989 and ethnic majority (b_{7j}). Here, b_{7j} captures the relative effects of the two-child policy on family size and other related outcomes of each cohort.

To interpret b_{7j} as the causal effect of the policy, I need to assume that in the absence of the policy, cohort effects would have been the same for both ethnic groups. While this assumption is untestable, I indirectly test it by examining cohort effects of women ages 40 and older in 1989 and cohort effects of women under age 20 in 1989. Since women ages 40 and older in 1989 are too old to be affected by the policy, the estimates of these women will capture only the differences in cohort effects of both ethnic groups. Thus, if these estimates are close to zero, they imply that in the pre-policy period, changes in family size by birth year cohort are the same for both ethnic groups. On the other hand, women under age 20 in 1989 are fully affected by the policy. Thus, if I observe no differences in cohort effects of these women, it implies that cohort effects are likely to be the same for both ethnic groups in the post-policy period.

Figure 1 illustrates the coefficients of interactions in Eq. (2) and provides the first piece of evidence that cohort effects are the same for both ethnic groups. As seen from Fig. 1, the estimates of women ages 40 and older in 1989 (relative to the omitted group – women ages 49 in 1989) are close to zero and not statistically significant. Since these estimates capture differences in cohort effects of both ethnic groups, this implies that changes in family size are the same for both groups in the pre-policy period. The estimates of women under age 20 in 1989 are constant across ages, suggesting that cohort effects are likely to be the same for both ethnic groups in the post-policy period.

Although the evidence above suggests that the required common trend (birth year cohort) assumption holds, my estimates may be biased if the policy has spillover effects on ethnic minorities. If ethnic minorities follow the fertility behavior of ethnic majorities and stop at two children as well, then my estimates would be biased downward. Since the policy does not apply to ethnic minorities, ethnic majorities may have an incentive to marry ethnic minorities, which would also lead to spillover effects on ethnic minorities. Even though there are inter-ethnic marriages, they may not be a big concern here. The number of these marriages is very small. In 2009, the inter-ethnic marriages accounted for just 1.9% of all marriages (General Statistics Office of Vietnam 2010). In addition, because of the war period (1955–1975), the analysis of older women presented above may not be an adequate assessment of the parallel trend assumption. However, as the analysis of younger women suggests, my research design is valid in this context. If there were a violation of the parallel trend assumption, I would observe a change in the DID estimates of these cohorts. However, as I mention above, I do not observe this for younger women.

Given the patterns shown in Fig. 1, in the next step, I estimate the following regression equation to examine the effect of the policy on women's fertility at different ages in 1989 and to capture the main trends in the effects of the policy.

$$\begin{aligned} \mathrm{Y}_{ijt} = {} & \mathrm{c}_0 + \mathrm{c}_1\,\mathrm{age}_{\mathrm{it}} + \mathrm{c}_2\,\mathrm{age}^2_{\mathrm{it}} + \mathrm{c}_3\,\mathrm{age}_{\mathrm{it}} \times \mathrm{majority}_{\mathrm{i}} + \mathrm{c}_4\,\mathrm{age}^2_{\mathrm{it}} \times \mathrm{majority}_{\mathrm{i}} \\ & + \mathrm{c}_5\,\mathrm{d}(\mathrm{age\ in\ 1989} < 30)_{\mathrm{i}} + \mathrm{c}_6\,\mathrm{d}(\mathrm{age\ in\ 1989} < 30)_{\mathrm{i}} \times \mathrm{majority}_{\mathrm{i}} \\ & + \mathrm{c}_7\,\mathrm{d}(30 \leq \mathrm{age\ in\ 1989} \leq 39)_{\mathrm{i}} + \mathrm{c}_8\,\mathrm{d}(30 \leq \mathrm{age\ in\ 1989} \leq 39)_{\mathrm{i}} \times \mathrm{majority}_{\mathrm{i}} \\ & + \mathrm{c}_9\,\mathrm{year}_{\mathrm{it}} + \mathrm{c}_{10}\,\mathrm{province}_{\mathrm{it}} + \eta_{ijt} \end{aligned} \tag{3}$$

Instead of including dummy variables of each age in 1989 as in Eq. (2), I include indicators for each age group in 1989 in Eq. (3). The coefficients of interest are on the interactions between these dummy variables and ethnic majority (c_6 and c_7). Here, c_6 and c_7 capture the average effects of the policy on fertility and other outcomes of women under age 30 and ages 30–39 in 1989 (relative to the omitted group – women who were 40 and older in 1989). With this specification, I can include year effects in the regression. The estimates of coefficients of the interaction terms in Eq. (3) are similar with and without the inclusion of year effects.

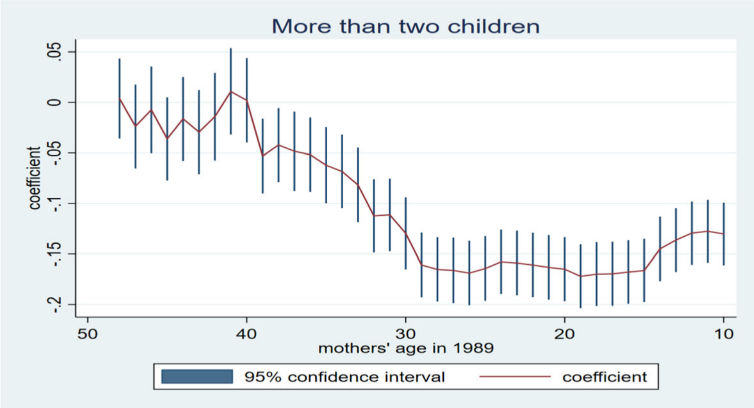

Fig. 1 Estimated coefficients of interactions between dummy variables for women's age in 1989 and ethnic majority in model of having more than two children, full sample

Ethnic minorities tend to be less educated and more likely to live in rural areas than are ethnic majorities. Thus, to account for these differences, I also include dummy variables of educational attainment and rural areas in Eq. (3). Appendix Table 23 shows robustness of the results with and without the inclusion of these controls. These results indicate that these differences do not drive my estimates. Moreover, I also estimate the event study (exposure age) analysis (Eq. 2) and the analysis with three different age groups (Eq. 3) with no control variables and obtain very similar results as those in the equations with all control variables.[9]

5.2 The reduced form: the effect of the two-child policy on maternal labor supply

To examine the effect of the two-child policy on maternal labor supply,[10] I use the same econometric framework as I do for family size. The linear probability model of my reduced form is thus as follows:

$$\begin{aligned}\mathrm{Mother's\ employment}_{ijt} = {} & \alpha_0 + \alpha_1\,\mathrm{age}_{it} + \alpha_2\,\mathrm{age}_{it}^2 + \alpha_3\,\mathrm{age}_{it} \times \mathrm{majority}_i \\ & + \alpha_4\,\mathrm{age}_{it}^2 \times \mathrm{majority}_i + \textstyle\sum_{j=10}^{48} \alpha_{5j}\mathrm{age\ in\ 1989}_{ij} \\ & + \alpha_6 \mathrm{majority}_i + \textstyle\sum_{j=10}^{48} \alpha_{7j}\mathrm{age\ in\ 1989}_{ij} \times \mathrm{majority}_i \\ & + \mathrm{province}_{it} + \eta_{ijt} \qquad (4)\end{aligned}$$

The coefficients of interest are on the interaction terms between dummy variables for women's age in 1989 and ethnic majority (α_{7j}). To interpret α_{7j} as the causal

[9] I try different age bins (age 10–19, 20–29, 30–39, 40–49) rather than the quadratic in the probability of having more than two children equation and obtain very similar results.

[10] The Vietnam Population and Housing Censuses have no information on working hours. The Censuses have information on whether an individual works for salary or not. However, more than 50% of this information are missing in the data. Thus, I could not examine the effects of the policy on these outcomes.

effect of the two-child policy on maternal labor supply, I assume that changes in the maternal employment would be the same for both ethnic groups in the absence of the policy. Figure 6 presents the coefficients of the interaction terms in Eq. (4) and provides the evidence that cohort effects are the same for both ethnic groups in the pre-policy period. The estimates of women ages 40 and older in 1989 capture the differences in cohort effects of both ethnic groups, and they are indistinguishable from zero. On the other hand, the estimates of women under age 30 in 1989 are positive and constant across women's ages, suggesting a constant effect of the policy on maternal employment of these women. Together, these results indicate that changes in maternal employment are the same for both ethnic groups in the pre- and post-policy period.

To control for differences in education and rural residency between both ethnic groups, I further control for rural indicators and educational dummies in Eq. (4). Appendix Table 24 shows the estimates of interest with and without the inclusion of these controls and suggests the robustness of my results.

5.3 Instrumental variables estimates of the effect of family size on maternal employment and child education in Vietnam

To obtain the instrumental variables estimates of the effect of family size on maternal labor supply and child education, I use estimates from the first-stage regression of the two-child policy on the number of children and probability of having more than two children to construct the predicted number of children and the predicted probability of having more than two children, which are the instruments. The regression model used for the instrumental variables procedure is as follows:

$$\begin{aligned} Z_{ijt} = {} & e_0 + e_1 \text{age}_{it} + e_2\, \text{age}^2{}_{it} + e_3 \text{age}_{it} \times \text{majority}_i + e_4 \text{age}^2{}_{it} \times \text{majority}_i \\ & + \textstyle\sum_{j=10}^{48} e_{5j} \text{age in } 1989_{ij} + e_6\, \text{majority}_i + \mathbf{e_7}\, \widehat{\text{N of children}}_{ijt} \\ & + e_8\, \text{province}_{it} + \varepsilon_{ijt} \end{aligned} \tag{5}$$

Z_{ijt} is maternal employment, average years of education of children in the household, school enrollment, and education of the first two children. Equation (5) uses the predicted number of children and the predicted probability of having more than two children from Eq. (2) instead of the actual number of children and the actual probability of having more than two children. As I show later, the two-child policy is a significant predictor of the number of children and the probability that mothers have more than two children. Thus, the instruments have good explanatory power in the first stage.

The exclusion restriction of the instrumental variables approach is that fertility should be the only channel through which the policy could affect maternal employment or child education. Although the policy can affect maternal employment or child education through other channels (e.g., delayed marriage, increase in education), these effects operate through the fertility channel. Thus, the exclusion restriction is likely to hold in this context.

6 Results

6.1 The effect of Vietnam's two-child policy on family size

Figure 1 shows coefficients of interaction terms between dummy variables for women's age in 1989 and ethnic majority in the having more than 2 children equation for women's birth cohorts 1941–1979.[11] As seen from Fig. 1, the estimates of women ages 40 and older in 1989 are close to zero and not statistically significant, suggesting no effects of the policy on fertility of these women. The estimates for women ages 30 to 39 in 1989, on the other hand, indicate a monotonic increasing effect of the policy on fertility of these mothers. The estimates range from −0.05 to −0.13, suggesting an average effect of a 9 percentage point decrease in the probability that these mothers have more than two children. As Fig. 1 further indicates, women under age 30 in 1989 are the most affected group. The estimates are significantly different from zero and constant across ages for mothers ages 15–30 in 1989. The estimates slightly decreased for mothers under age 15 in 1989, but they are still significantly different from zero. Overall, the estimates indicate a 15 percentage point decrease in the probability that mothers under age 30 in 1989 have more than two children, which is equivalent to a 50% reduction at means.

Table 2 contains estimates of the effect of Vietnam's two-child policy on the probability of having more than two children and the number of living children for women of different age groups in 1989. Instead of including dummy variables for each age in 1989, I include indicators of three age groups as mentioned above. As Table 2 indicates, the policy was more effective at reducing the probability of having more than two children of younger women in 1989. The policy decreased the probability that a woman has more than two children by 15 percentage points (50%) for women under age 30 in 1989 and by 7 percentage points (11.5%) for women ages 30–39 in 1989. This result is in line with my expectation. Since most women in Vietnam have children in their 20s, the policy should have a bigger effect on fertility of younger women in 1989. On average, the policy decreased the number of living children of younger women by 0.18 births per woman (9%) and of middle-aged women by 0.2 births per woman (6.4%).

As discussed in Sect. 3.1, there might be different effects of the two-child policy on fertility by education and urban/rural areas. Table 3 presents estimates of the effect of the two-child policy on women's fertility by mothers' schooling.[12, 13] As Table 3 indicates, young and low-educated women were more affected by the policy than more educated women were at the same age. The policy decreased the probability of having more than two children of young, low-educated women by 16 percentage points (37%) and of more educated women by 5.1 percentage points (21.25%). The policy, on the other hand, reduced the probability of having more than two children of middle-aged, low-educated women by 4 percentage points (5.55%) and of more educated women at the same age by 5.3 percentage points (9.81%). The policy reduced the average number of living children of low-educated women under age 30 by 0.58 births per woman

[11] Coefficients of interaction terms in Fig. 1 are also shown in Appendix Table 17.

[12] I show coefficients of interactions between dummy variables for women's age in 1989 and ethnic majority in the fertility equation for women with less than primary education in Appendix Figure 7 and for women with at least primary education in Appendix Figure 8.

[13] If the policy has an impact on educational attainment, then women with less than primary education may obtain more education and thus have at least primary education.

(24%) and of those ages 30–39 by 0.3 births (8.02%) per woman. However, the policy had no significant effects on the average number of living children of more educated women at all ages.

Table 4 shows heterogeneity in the effect of the policy on family size by urban and rural areas.[14] As Table 4 indicates, the policy was more effective at reducing the probability of having more than 2 children of younger women in rural areas. Since women in rural areas tend to be less educated than women in urban areas, this finding is in line with the results shown in Table 3, which find that young, less educated women are affected more by the policy than more educated women at the same age. Specifically, the policy decreased the probability of having more than two children of younger women in rural areas by 12.8 percentage points (35%) and by 4.6 percentage points (6.5%) for women ages 30–39. On the other hand, the policy reduced the probability of having more than two children by 10.5 percentage points for younger women in urban areas and by 6 percentage points for women ages 30–39. On average, the policy reduced the number of living children of women in rural areas by 0.19–0.2 births per woman (an approximately 9-percentage decrease for younger women and a 5.6-percentage decrease for middle-aged women). However, the policy had no significant effects on the number of living children of women in urban areas.

6.2 The effect of Vietnam's two-child policy on mothers' age at first birth and birth spacing

Figure 2 shows coefficients of interaction terms between dummy variables for women's age in 1989 and ethnic majority in the mothers' age at first birth equation.[15] As Fig. 2 indicates, the estimates of women ages 40 and older in 1989 (relative to the omitted group – women ages 49 in 1989) are indistinguishable from zero in most cases. Overall, the estimates suggest a potential zero effect of the policy on age at first birth of these mothers, which is consistent with my expectation. The estimates of women under age 40 in 1989 are around −0.05, constant across ages, but are not significantly different from zero. These estimates suggest that the policy has no significant effects on mothers' age at first birth in the post-policy period.

Column (1) of Table 5 shows estimates of the effect of the policy on mothers' age at first birth for women of different ages in 1989 and suggests similar results. The estimates for younger women in 1989 are 0.13 and not significantly different from zero. Likewise, the estimates for women ages 30–39 are close to zero (approximately 0.03) and not statistically significant. Together, these results imply that women did not alter the timing of their first birth to achieve the reduction in fertility as observed above. One explanation is that they must have stopped fertility earlier or spaced children out more as directed by policy.

Figure 3 shows coefficients of interactions between dummy variables for mothers' age in 1989 and ethnic majority dummy in the birth spacing at least 3 years equation.[16] The estimates of older women fluctuate around zero and not statistically significant,

[14] I show coefficients of interactions in having more than 2 children equation for women in rural areas in Appendix Figure 9 and for women in urban areas in Appendix Figure 10.

[15] I show coefficients of the interactions in the mothers' age at first birth equation in Appendix Table 18.

[16] Coefficients of interactions in the birth spacing equation are shown in Appendix Table 19.

Table 2 The effect of the policy on the probability of having more than two children and the number of living children

Dependent variables	More than 2 children	Number of living children
Mean for women under age 30 in 1989	0.30 (0.36)	2.01 (1.41)
Mean for women ages 30–39 in 1989	0.61 (0.49)	3.14 (1.90)
Column	(1)	(2)
Under age 30 in 1989 × majority	−0.152***	−0.192***
	(0.005)	(0.031)
Age 30–39 in 1989 × majority	−0.067***	−0.196***
	(0.006)	(0.035)
Number of obs.	3,197,622	3,197,622

Note: Numbers in parentheses are standard errors (standard deviations for means)

The omitted group is those ages 40 and older in 1989

Other covariates included in the regressions are dummy variables for women under age 30 in 1989, women ages 30–39 in 1989, ethnic majority, age at the time of survey, age squared, age × majority, age squared × majority, province fixed effects, and year dummies for survey years 1999 and 2009

*$p < 0.05$. **$p < 0.01$, ***$p < 0.001$

Table 3 Heterogeneity in the effect of the policy on family size across mothers' schooling

Dependent Variables	More than 2 children		Number of living children	
Education	Less than primary education	At least primary education	Less than primary education	At least primary education
Mean for women < 30 in 1989	0.43 (0.50)	0.24 (0.43)	2.39 (1.68)	1.84 (1.22)
Mean for women 30–39 in 1989	0.72 (0.45)	0.54 (0.50)	3.74 (2.16)	2.71 (1.57)
Column	(1)	(2)	(3)	(4)
Under age 30 in 1989 × majority	−0.158***	−0.051***	−0.578***	0.071
	(0.006)	(0.014)	(0.035)	(0.064)
Age 30–39 in 1989 × majority	−0.040***	−0.053***	−0.282***	−0.083
	(0.001)	(0.015)	(0.041)	(0.069)
Number of obs.	1,083,490	2,105,403	1,083,490	2,105,403
% of the sample	35.27%	64.73%	35.27%	64.73%
% of minority	32.72%	9.93%	32.72%	9.93%

Note: Numbers in parentheses are standard errors (standard deviations for means)

The omitted group is those ages 40 and older in 1989

Other covariates included in the regressions are dummy variables for women under age 30 in 1989, women ages 30–39 in 1989, ethnic majority, age at the time of survey, age squared, age × majority, age squared × majority, province fixed effects, and year dummies for survey years 1999 and 2009

*$p < 0.05$. **$p < 0.01$, ***$p < 0.001$

Table 4 Heterogeneity in the effect of the policy on family size across urban and rural areas

Dependent variables	More than 2 children		Number of living children	
Areas	Rural	Urban	Rural	Urban
Mean for women < 30 in 1989	0.37 (0.48)	0.17 (0.38)	2.23 (1.45)	1.57 (1.21)
Mean for women 30–39 in 1989	0.71 (0.45)	0.51 (0.50)	3.56 (1.98)	2.69 (1.72)
Column	(1)	(2)	(3)	(4)
Under age 30 in 1989 × majority	−0.128***	−0.105***	−0.185***	−0.077
	(0.005)	(0.020)	(0.031)	(0.10)
Age 30–39 in 1989 × majority	−0.046***	−0.06**	−0.214***	0.027
	(0.006)	(0.020)	(0.033)	(0.108)
Number of obs.	2,090,432	1,107,190	2,090,432	1,107,190
% of the sample	66.62%	33.38%	66.62%	33.38%
% of minority	22.76%	8.02%	22.76%	8.02%

Note: Numbers in parentheses are standard errors (standard deviations for means)

The omitted group is those ages 40 and older in 1989

Other covariates included in the regressions are dummy variables for women under age 30 in 1989, women ages 30–39 in 1989, ethnic majority, age at the time of survey, age squared, age × majority, age squared × majority, province fixed effects, and year dummies for survey years 1999 and 2009

except for mothers ages 40 and 43 in 1989. The estimates suggest that potentially, there are no significant differences in the birth spacing between both ethnic groups. The estimates of women ages 21–39 indicate a monotonic increase in birth spacing of these mothers. However, none of these estimates is significantly different from zero. The estimates of mothers under age 20 in 1989 are positive and significantly different from zero, indicating an increase of more than 10 percentage points in the probability that mothers have longer birth spacing (at least 3 years) of younger women in 1989.

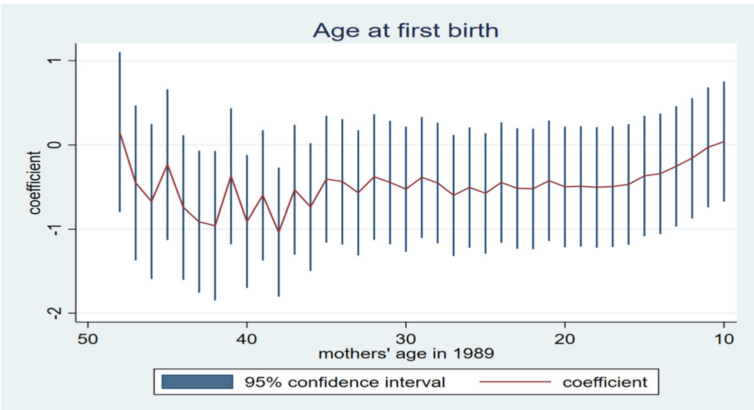

Fig. 2 Estimated coefficients of interactions between dummy variables for women's age in 1989 and ethnic majority in model of mothers' age at first birth, subsample

Table 5 The effect of Vietnam's two-child policy on mothers' age at first birth and mothers' birth spacing between first two births

Dependent variables	Mothers' age at first birth	Birth spacing at least 3 years	Birth spacing
Mean for women < 30 in 1989	23.28 (3.94)	0.63 (0.48)	3.94 (2.67)
Mean for women 30–39 in 1989	24.22 (4.09)	0.50 (0.50)	3.05 (2.01)
Column	(1)	(2)	(3)
Under age 30 in 1989 × majority	0.130	0.105***	0.806***
	(0.086)	(0.010)	(0.050)
Age 30–39 in 1989 × majority	0.031	−0.006	0.079
	(0.094)	(0.011)	(0.055)
Number of obs.	1,768,536	1,454,159	1,454,159

Note: Numbers in parentheses are standard errors (standard deviations)

The omitted group is those ages 40 and older in 1989

Other covariates included in the regressions are dummy variables for women under age 30 in 1989, women ages 30–39 in 1989, ethnic majority, age at the time of survey, age squared, age × majority, age squared × majority, province fixed effects, and year dummies for survey years 1999 and 2009

*$p < 0.05$, **$p < 0.01$, ***$p < 0.001$

Columns (2) and (3) of Table 5 present estimates of the effects of the policy on mothers' birth spacing between the first two births across mothers' ages. The estimates suggest that the policy increased the probability that younger mothers have longer birth spacing (at least 3 years) by 10.5 percentage points (16.67%). On average, the policy increased birth spacing of these mothers by 0.81 years (20.56%). However, the estimates suggest no significant effects of the policy on birth spacing of mothers ages 30–39 in 1989. The results are in line with the findings suggested by Fig. 3.

6.3 The effect of Vietnam's two-child policy on parents' ability to achieve son preference

Figure 4 presents coefficients of interaction terms between dummy variables for women's age in 1989 and ethnic majority in the proportion of sons in each family equation.[17] For this analysis, I use subsample. The estimates of women ages 40 and older in 1989 were negative and significantly different from zero in some cases. The estimates indicate that in the pre-policy period, ethnic majorities have a lower proportion of sons than ethnic minorities, which is consistent with ethnic majorities having fewer children than ethnic minorities in this period. The estimates of women ages 30–39 fluctuate around −0.05 and significantly different from zero, indicating that ethnic majority women experienced a reduction in the proportion of sons in each family in the post-policy period. However, since there are significant differences in the proportion of sons between both ethnic groups in the pre-policy period, it is hard to claim that these estimates capture the causal effect of the policy on the proportion of sons in each family. At best, these estimates indicate that there is a significant association between

[17] Coefficients of interactions in the proportion of sons in each family equation are shown in Appendix Table 20.

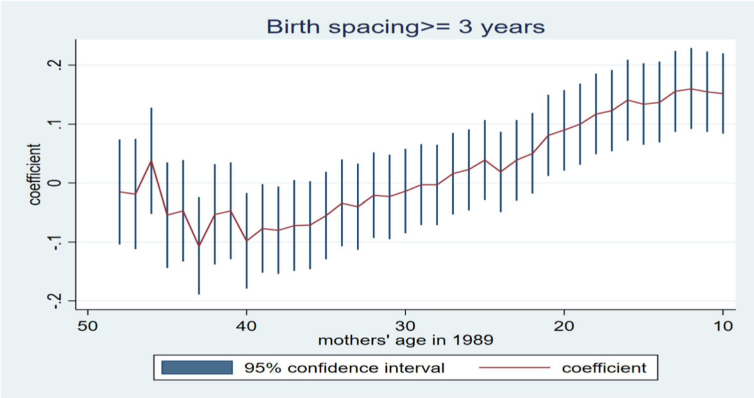

Fig. 3 Estimated coefficients of interactions between dummy variables for women's age in 1989 and ethnic majority in model of birth spacing at least 3 years equation, subsample

the reduction in family size caused by the two-child policy and the decrease in the proportion of sons in each family in the post-policy period.

Column (1) of Table 6 shows the association between the reduction in family size and the proportion of sons in each family across women's ages. The results are in line with the estimates in Fig. 3. The estimates indicate that the reduction in family size caused by the two-child policy was associated with a 1.2 percentage point (2.4%) decrease in the proportion of sons in each family of younger women in 1989. However, there is no significant association between the reduction in fertility and the proportion of sons in each family of middle-aged women. Overall, these results suggest a decline in parents' ability to achieve son preference in the post-policy period.

Figure 5 shows coefficients of interactions between dummy variables for mothers' age in 1989 and ethnic majority in the gender at last birth equation.[18] The estimates of older women are close to zero and not significantly different from zero, suggesting no significant differences in the probability of having a boy at last birth between mothers of both ethnic groups. The estimates of younger women in 1989 are also close to zero and not statistically significant, indicating no effects of the policy on the probability of having a boy at last birth of these mothers.

Column (2) of Table 6 presents estimates of the effect of the policy on gender at last birth across mothers' age and suggests similar results as Fig. 5. The estimates of both women ages 30–39 and women under age 30 in 1989 are very close to zero and not significantly different from zero. Together with the findings in Fig. 5, these estimates suggest that the policy had no significant effects on the probability that the mother has a boy at last birth.

6.4 The effect of Vietnam's two-child policy on maternal labor supply

Figure 6 shows coefficients of interaction terms between dummy variables for women's age in 1989 and ethnic majority in the maternal employment equation.[19] As seen from Fig. 6, the estimates of women ages 40 and older in 1989 are indistinguishable from

[18] Coefficients of these interactions are displayed in Appendix Table 21.
[19] Coefficients of interaction terms in Figure 6 are also presented in Appendix Table 22.

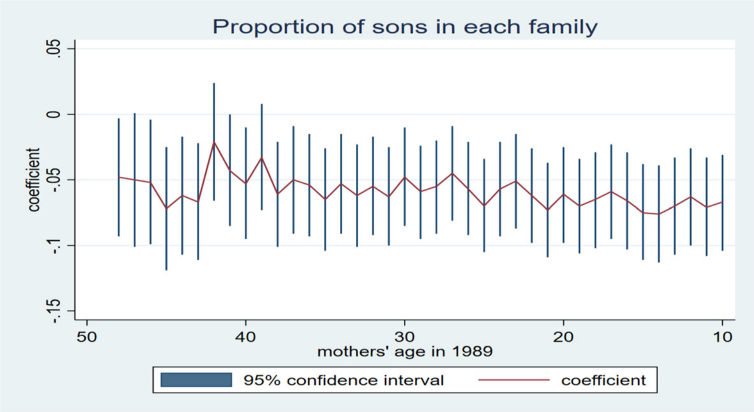

Fig. 4 Estimated coefficients of interactions between dummy variables for women's age in 1989 and ethnic majority in model of proportion of sons in each family equation, subsample

zero, suggesting that the policy had no effects on the labor supply of these women. This is also the evidence that changes in employment rates are the same for both ethnic groups in the pre-policy period, which further demonstrates the validity of my identification assumption. The estimates of women under age 40 in 1989 are positive, although not significant, suggesting a potential increase of 1.1 percentage points (1.3%) in the labor supply of these mothers in the post-policy period.

Column (1) of Table 7 presents estimates of the causal effect of the policy on maternal labor supply for women of different age groups in 1989. As the estimates indicate, the policy has a stronger effect on the labor supply of younger women in 1989,

Table 6 The effect of Vietnam's two-child policy on the proportion of sons in each family and gender of mothers' last birth

Dependent variables	Proportion of sons in each family	Last birth is a boy
Mean for women < 30 in 1989	0.499 (0.386)	0.557 (0.497)
Mean for women 30–39 in 1989	0.506 (0.322)	0.545 (0.498)
Column	(1)	(2)
Under age 30 in 1989 × majority	−0.012*	−0.009
	(0.005)	(0.009)
Age 30–39 in 1989 × majority	−0.005	−0.002
	(0.006)	(0.009)
Number of obs.	1,852,725	1,773,785

Note: Numbers in parentheses are standard errors (standard deviations for means)

The omitted group is those ages 40 and older in 1989

Other covariates included in the regressions are dummy variables for women under age 30 in 1989, women ages 30–39 in 1989, ethnic majority, age at the time of survey, age squared, age × majority, age squared × majority, province fixed effects, and year dummies for survey years 1999 and 2009

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

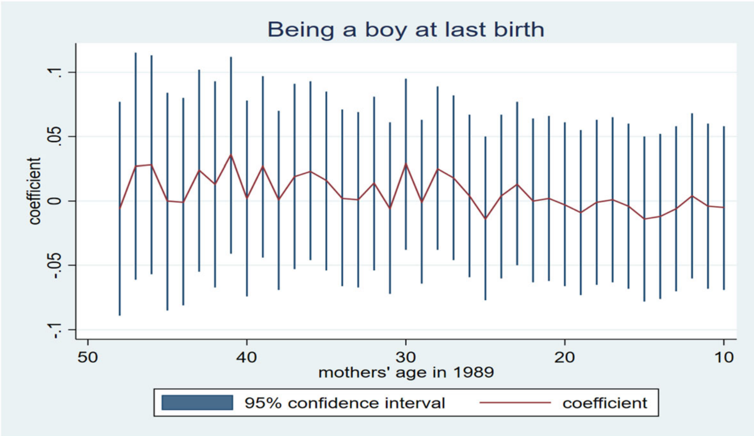

Fig. 5 Estimated coefficients of interactions between dummy variables for women's age in 1989 and ethnic majority in model of gender of last birth equation, subsample

which is consistent with my expectation. The policy increased the labor supply of women under age 30 in 1989 by 1.3 percentage points (1.5%). In contrast, the policy had no significant effects on the labor supply of women ages 30–39 in 1989.

Columns (2) and (3) of Table 7 show estimates of the effect of the policy on maternal labor supply by mothers' schooling. The estimates suggest that the policy had a stronger effect on the labor supply of young, less educated mothers, which is consistent with the results that the policy affected fertility of these mothers the most. Specifically, the policy increased the labor supply of young, low-educated women by 3 percentage points (1.7%). On the other hand, the policy had no significant effects on the labor supply of middle-aged, low-educated women. The estimates for women with at least primary education are negative and significantly different from zero. However, as Table 3 indicates, the policy had no significant effects on the number of living children of these mothers. Thus, the estimates cannot be interpreted as the effect of the policy on

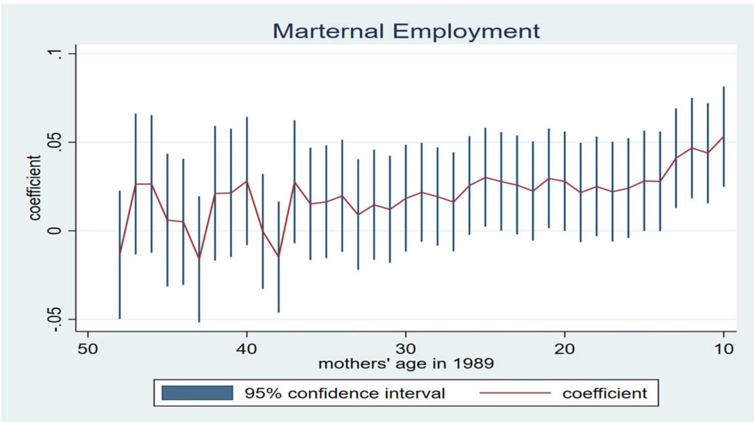

Fig. 6 Estimated coefficients of interactions between dummy variables for women's age in 1989 and ethnic majority in model of maternal employment, full sample

Table 7 The effect of Vietnam's two-child policy on maternal employment and heterogeneity in the effect of the policy across mothers' schooling and urban/rural areas

Dependent variable	Employed				
Samples	Full Sample	Less than primary education	At least primary education	Rural	Urban
Mean for women < 30 in 1989	0.86 (0.34)	0.84 (0.37)	0.87 (0.33)	0.91 (0.29)	0.77 (0.42)
Mean for women 30–39 in 1989	0.80 (0.40)	0.76 (0.43)	0.84 (0.37)	0.86 (0.35)	0.75 (0.43)
Column	(1)	(2)	(3)	(4)	(5)
Under age 30 in 1989 × majority	0.013**	0.026***	−0.039***	0.039***	−0.122***
	(0.004)	(0.005)	(0.012)	(0.004)	(0.019)
Age 30–39 in 1989 × majority	−0.001	−0.004	−0.039**	0.007	−0.072***
	(0.005)	(0.006)	(0.013)	(0.004)	(0.021)
Number of obs.	3,126,106	1,066,675	2,056,651	2,059,290	1,066,816
% of the sample	–	35.27%	64.73%	65.87%	34.13%
% minority	–	32.95%	9.93%	22.89%	7.95%

Note: Numbers in parentheses are standard errors (standard deviations for means)

The omitted group is those ages 40 and older in 1989

Other covariates included in the regressions are dummy variables for women under age 30 in 1989, women ages 30–39 in 1989, ethnic majority, age at the time of survey, age squared, age × majority, age squared × majority, province fixed effects, and year dummies for survey years 1999 and 2009

$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$

the labor supply of these mothers in this context since the channel through which the policy may affect their labor supply is through fertility.

Columns (4) and (5) of Table 7 present estimates of the effect of the policy on maternal employment by urban and rural areas. The estimates suggest that younger women in rural areas are more responsive to the reduction in family size caused by the two-child policy than middle-aged women, which is consistent with my expectation. The policy increased the labor supply of younger women in rural areas by 4 percentage points (4.4%) and had no significant effects on the labor supply of middle-aged women. The estimates for women in urban areas are negative and significantly different from zero. However, as Table 4 indicates, the policy had no significant effects on the number of living children of women in urban areas. Thus, it should have no noticeable effects on their labor supply. In other words, the estimates for women in urban areas in Table 7 cannot be interpreted as the effect of the policy on employment of these mothers.

6.5 Instrumental variables estimates of the effect of fertility on maternal employment and child education in Vietnam

Table 8 shows OLS and 2SLS estimates of the causal effect of fertility on maternal employment in Vietnam. As mentioned above, the interactions between dummy variables for women's age in 1989 and ethnic majorities serve as instruments for the

Table 8 The effect of fertility on maternal employment in Vietnam

Dependent variable	Employed
	Coefficient
OLS	
Number of living children	-0.008^{*}
	(0.0003)
R-squared	0.107
2SLS	Coefficient
Number of living children (year of birth dummies × ethnic majority)	-0.150^{*}
	(0.013)
R-squared	0.416
F-test (p value)	10,257.29 ($p < 0.001$)
Number of obs.	3,126,106

Note: Numbers in parentheses are standard errors. Year of birth dummies, province indicators, ethnic majority indicator, age, age squared, age × majority, and age squared × majority are included in the regressions. F-test and R-squared of the first stage are reported in the tables. $^{*}p < 0.05$

number of living children in the maternal employment equation. The exclusion restriction would be violated if the policy can affect maternal employment through other channels such as education, son proportion, and marriage. For example, the policy may affect maternal employment through son proportion. With the reduction in the proportion of sons in each family caused by the policy, women may have more bargaining power in the household and thus alter their labor market behaviors accordingly. Although the policy may facilitate women to obtain more education, delay marriage, have a lower son proportion, and thus participate more in the labor force, the effects of the policy on these outcomes operate through the effect of the policy on fertility. Thus, despite the potential effects of the policy on women's education, son proportion, and marriage, the exclusion restriction is still likely to hold in this context.

The upper panel presents OLS estimates of the impact of children on maternal employment. The point estimate is −0.008 and statistically significant. The estimate indicates a small negative effect of children on mothers' labor supply, which is in line with the findings of a recent study (Aaronson et al. 2017). The lower panel shows the instrumental variables estimates in Eq. (5). The point estimate is negative and statistically significant, suggesting that having an additional child decreased maternal employment by 15 percentage points (17.4%).

Compared to the estimates of other studies, my IV estimates on the effect of fertility on maternal employment in Vietnam are larger. The US estimates of fertility on maternal employment reported by Angrist and Evans (1998) are −10.4 percentage points for 1980 and − 8.4 percentage points for 1990. The estimates reported by Cruces and Galiani (2007) also range from 8.1 to 9.6 percentage points for Argentina and from 6.3 to 8.6 percentage points for Mexico. However, my estimates capture the average effect of children on the labor supply of women who comply with the two-child policy. In contrast, the estimates of other studies capture the average effect of fertility on the labor supply of women who either have multiple births or prefer a mixed-sibling gender composition. Thus, it is possible that my estimates are larger than the estimates of others.

Table 9 The effects of more than 2 children on child education

Outcomes	Average years of education		First child enrolled in school	Second child enrolled in school	First child with at least primary education	Second child with at least primary education
Samples	All children	First two children	Children ages 6 and older		Children ages 14 and older	
Mean (more than 2 children)	0.35 (0.48)	0.35 (0.48)	0.40 (0.50)	0.53 (0.50)	0.48 (0.50)	0.63 (0.48)
Mean (outcome)	4.06 (3.45)	4.26 (3.54)	0.95 (0.21)	0.94 (0.23)	0.87 (0.34)	0.855 (0.35)
Panel A: OLS estimates						
More than 2 children	0.045^{***}	1.604^{***}	0.028^{***}	0.040^{***}	-0.080^{***}	-0.073^{***}
	(0.006)	(0.007)	(0.003)	(0.001)	(0.002)	(0.002)
R-squared	0.578	0.561	0.202	0.201	0.202	0.196
N of obs.	1,773,928	1,773,928	1,538,264	1,082,795	849,113	431,294
Panel B: 2SLS estimates						
More than 2 children (year of birth × ethnic majority)	-3.90^{***}	-2.058^{***} (0.019)	-0.168^{***}	-0.227^{***}	-0.473^{***}	-0.595^{***}
	(0.018)		(0.002)	(0.002)	(0.004)	(0.006)
R-squared	0.226	0.227	0.202	0.191	0.211	0.179
F-test	2514.47	2531.55	2106.01	1710.55	1295.61	628.36
(p value)	$p < 0.001$	$p < 0.001$	$p < 0.001$	$p < 0.001$	$p < 0.001$	$p < 0.001$
N of obs.	1,773,928	1,773,298	1,538,264	1,082,795	849,113	431,294

Note: All regressions controlled for age, age squared, age × majority, age squared × majority, majority dummy, birth year dummies, and provincial dummies. Numbers in parentheses are standard errors (standard deviations for means). F-test and R-squared of the first stage are reported in the table. $^{***}p < 0.001$

Table 9 shows OLS and 2SLS estimates of the impact of having more than two children on child education. Panel A presents OLS estimates, and Panel B presents 2SLS estimates. The OLS estimates suggest that having more than two children is associated with an increase of 0.05 years of education of all children, an increase of 1.6 years of education of the first two children, and a 2.8–4 percentage point increase in the probability that the first-/second-born child is enrolled in school. However, the estimates in the at least primary education indicate the opposite findings. These estimates are negative and significantly different from zero, suggesting that having more than two children decreases the probability that the first-/second-born child has at least primary education by 7.3–8 percentage points, respectively. Overall, the OLS estimates suggest mixed evidence on the existence of the trade-off between child quantity and child quality in Vietnam.

In contrast, the IV estimates indicate significant effects of having more than two children on decreasing child education in the household. The estimates are negative and significantly different from zero. The estimates suggest that having more than two children decreases average years of education of all children in the household by 3.90 years, average years of education of the first two children by 2.06 years, and the probability that the first-/second-born child is enrolled in school by 16.8 and 22.7 percentage points, respectively. Moreover, having more than two children reduces the probability that the first-/second-born child has at least primary education by 47.3 and 59.5 percentage points.

These results are in line with the results of other studies. Using regional differences in the enforcement intensity of China’s one-child policy as an instrumental variable for family size, Li and Zhang (2017) find that having an additional child reduces educational level of firstborn children by 0.07 to 0.12, and the probability that firstborn children attend junior secondary school by 11.1–13.3 percentage points. Exploring the tightening of China’s one-child policy in late 1979, Qin et al. (2017) apply a research discontinuity design to examine the trade-off between child quantity and child quality and document that being a single child in the family increases the child’s educational attainment in adulthood by 5.6 years of schooling. Compared to the estimates of those studies, my estimates are larger. However, my estimates capture the effect of family size on child education of families that comply with the two-child policy. Thus, it is reasonable that I may obtain larger estimates.

7 Falsification tests

In this section, I present the results of falsification tests in which I use the probability of having at least one child and the probability of getting married as alternative outcomes. The two-child policy should have no effects on the probability that women have at least one child since it only imposes fines and punishments on parents who have more than two children. Similarly, the policy should not affect the probability of getting married of women ages 30 and over in 1989. Most of these women should have gotten married by the time that the policy was in place. The policy may have a small effect on the probability of getting married of women under age 30 in 1989. Since these women know that they will have only two children, they may adjust their marriage accordingly. For example, they may delay marriage at a later age.

Appendix Table 25 shows the coefficients of interest of alternative outcomes. As the table indicates, the policy had no effect on the probability that a woman has at least one child. The policy also did not affect the probability of getting married of women ages 30 and over in 1989. These results are consistent with my expectations, and together they suggest the validity of my research design. The policy had a small feedback effect on the probability of getting married of women under age 30 in 1989. The point estimate is 0.007 and statistically significant, suggesting that under the two-child policy, younger women tend to get married at an earlier age. This is contradictory with my expectation that they would delay their marriage. However, the magnitude of this estimate is very close to zero. Thus, it does not impose a threat to the validity of my identification assumption.

8 Discussion and conclusion

Vietnam's two-child policy remains one of the most controversial policies that the Vietnamese government has implemented. The government is still debating whether to abandon the policy and to allow its people to have as many children as they desire. This paper explores differences in how the policy affected women of different ages and ethnicities to provide the estimates of the causal effects of the policy on family size, son proportion, and maternal employment. Using the policy as an instrument for family size, I further investigate the effects of fertility on maternal employment and the trade-off between child quantity and child quality in Vietnam.

Using data from the Vietnam Population and Housing Censuses from 1989, 1999, and 2009, I use a differences-in-differences framework and reach the following findings. First, the policy decreased the probability that a woman has more than two children by 15 percentage points (50%) for women under age 30 in 1989 and by 7 percentage points (11.5%) for women ages 30–39 in 1989. This result is similar to the findings of the study on China's one-child policy, which finds that China's one-child policy reduced the likelihood that mothers have a second child by 11 percentage points (50%) (Li et al. 2005). Low-education women and women in rural areas are affected more by the policy. On average, the policy reduced the number of living children by 0.2 births per woman (10%). The result is consistent with the findings of other studies on the effects of "later, longer, and fewer" policies in China. Chen and Fang (2018) examine the effects of these policies and find that a one-unit exposure to the policies led to a reduction of 0.2–0.3 births in the number of living children.

Second, the policy had no significant effects on mothers' age at first birth and the gender at mothers' last birth, which contradicts the findings of Li et al. (2011) that China's one-child policy causes sex ratio imbalance. The estimates on birth spacing further indicate that the policy increased the probability that younger mothers have longer birth spacing (at least 3 years) by 10.5 percentage points (16.67%) and raised their birth spacing by 0.81 years (20.56%). Chen and Huang (2018) investigate the effects of "later, longer, and fewer" policies in China and document that mothers who were exposed to the policies experienced a 1 year increase in birth spacing. The reduction in family size caused by the policy was associated with a 1.2 percentage point (2.4%) decrease in the proportion of sons in each family, indicating a decrease in parents' ability to achieve son preference in the post-policy period.

Third, the policy increased the labor supply of younger women in 1989 by 1.3 percentage points (1.5%). The instrumental variables estimates of the causal effect of fertility on maternal labor supply indicate that having an additional child decreased maternal employment by 15 percentage points (17.4%). This result is in line with the results of previous studies that examine the causal relationship of fertility on maternal employment in other developing countries. Caceres-Delpiano (2012) use the event of multiple births as an instrumental variable for family size and find that children negatively affect maternal employment in over 40 developing countries. Similarly, Cruces and Galiani (2007) use parental preferences for a mixed-sibling gender composition as instrumental variables and document that having an additional child led to a 6–10 percentage point decrease in maternal employment in Argentina and Mexico.

Fourth, my findings suggest that having more than two children decreases schooling of the first- and second-born children by 2.06 years, the probability that the first-/second-born child is enrolled in school by 16.8 and 22.7 percentage points, respectively. In addition, having an additional child reduces the probability that the first-/second-born child has at least primary education by 47.3 and 59.5 percentage points.

These results are the evidence that there exists a trade-off between child quantity and child quality in Vietnam. These results are consistent with the findings of previous studies on the quantity-quality trade-off in other developing countries. Using exogenous variation in fertility caused by China's one-child policy, Liu (2014) documents negative effects of the number of children on child height. Park and Chung (2012) exploit the exposure to a randomized family planning program in Matlab, Bangladesh, and find that having an addition child reduces children's schooling by 1.2 years and the probability that the first- and second-born children are enrolled in school by 11–13 percentage points. Using Chinese twins as an instrument for family size, Li et al. (2008) find that having an additional child decreases children's educational level by 0.04 and the probability that children are enrolled in school by 3 percentage points. Using variations in fertility under son preference as an instrument, Lee (2008) documents that having an additional child significantly decreased per-child investment in education in South Korea and that these negative effects are stronger when fertility is high.

Overall, my paper contributes to the literature by providing new evidence on the effectiveness of family planning policies at reducing fertility. My results suggest that while Vietnam's two-child policy was effective at decreasing family size, it had smaller effects on female labor supply. My findings also provide new evidence on the existence of the trade-off between child quantity and child quality in developing countries. These findings have important policy implications for developing countries where the governments rely on family planning policies to curb population growth, foster human capital, and promote economic development.

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Compliance with ethical standards

Conflict of interest The authors declare that they have no conflict(s) of interest.

Appendix

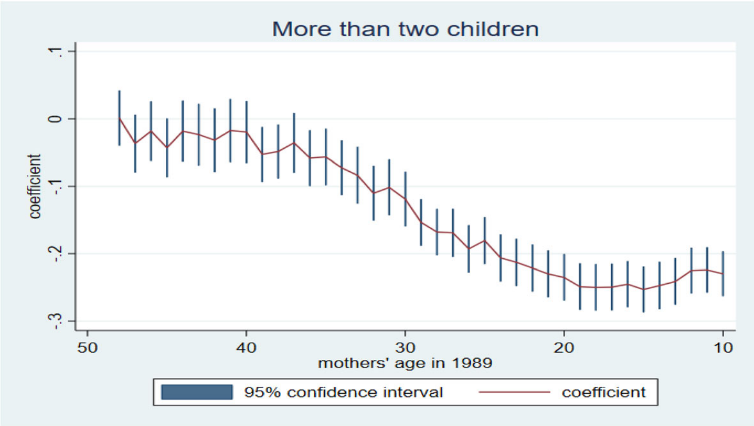

Fig. 7 Estimated coefficients of interactions between dummy variables for women's age in 1989 and ethnic majority in model of having more than two children, subsample of women with less than primary education

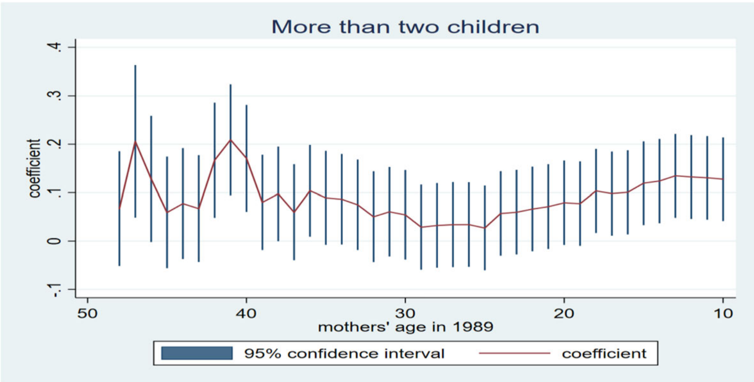

Fig. 8 Estimated coefficients of interactions between dummy variables for women's age in 1989 and ethnic majority in model of having more than two children, subsample of women with at least primary education

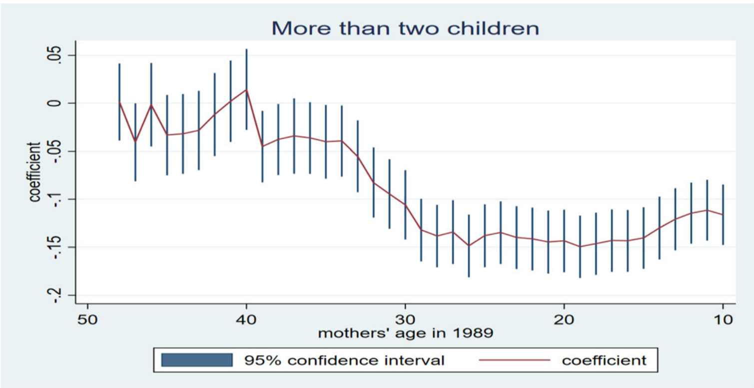

Fig. 9 Estimated coefficients of interactions between dummy variables for women's age in 1989 and ethnic majority in model of having more than two children, subsample of women in rural areas

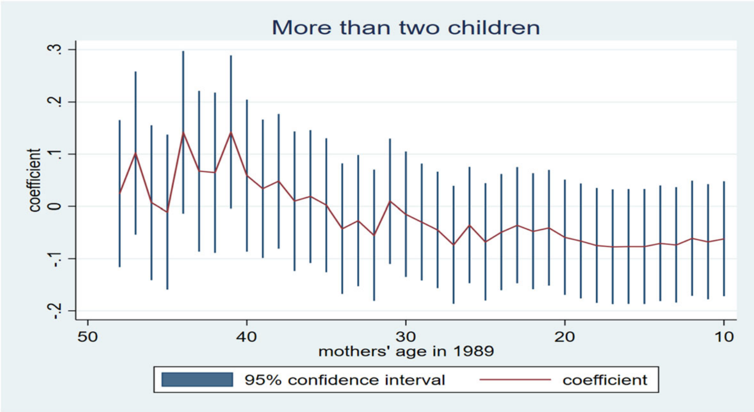

Fig. 10 Estimated coefficients of interactions between dummy variables for women's age in 1989 and ethnic majority in model of having more than two children, subsample of women in urban areas

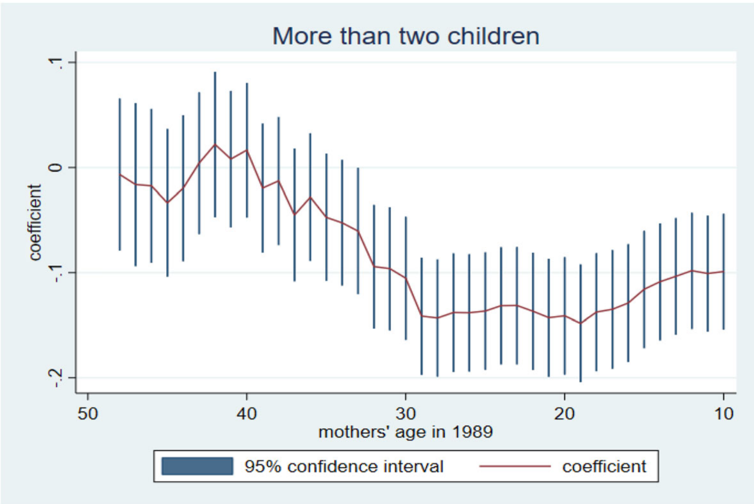

Fig. 11 Estimated coefficients of interactions between dummy variables for women's age in 1989 and ethnic majority in model of having more than two children, subsample

Table 10 Fractions of mothers that still have all of their children living at home

Censuses	Census 1989	Census 1999	Census 2009
Ethnic minority	80.73%	84.98%	79.29%
Ethnic majority	84.09%	84.18%	78.01%
Mothers' age at the time of survey			
Under age 30	96.58%	98.32%	97.15%
Age 30–39	83.63%	84.29%	78.31%
Age 40 and older	54.02%	60%	52.34%
Under age 30			
Ethnic minority	94.68%	98.03%	97.53%
Ethnic majority	96.91%	98.38%	96.99%
Age 30–39			
Ethnic minority	80.73%	84.98%	79.29%
Ethnic majority	84.09%	84.18%	78.01%
Age 40 and older			
Ethnic minority	52.30%	58.03%	50.50%
Ethnic majority	58.03%	60.27%	52.79%

Table 11 Summary statistics for full sample of women's birth cohorts 1940–1979: ethnic majority versus ethnic minority

Censuses	Census 1989		Census 1999		Census 2009	
	Ethnic majority	Ethnic minority	Ethnic majority	Ethnic minority	Ethnic majority	Ethnic minority
Number of children	1.83 (2.05)	2.23 (2.31)	1.93 (1.63)	2.40 (1.93)	2.19 (1.19)	2.8 (1.60)
Percent having						
No child	38.62	33.69	22.2	18.31	7.79	5.41
One child	14.65	13.2	20.45	16.36	15.01	9.74
2 children	15.87	14.37	27.25	24.31	45.04	34.81
At least 3 children	30.86	38.74	30.1	41.02	32.16	50.04
Percent employed	0.76 (0.42)	0.91 (0.29)	0.79 (0.41)	0.91 (0.29)	0.867 (0.33)	0.96 (0.20)
Mothers' age	25.35	24.69	35.36	34.86	45.12	44.41
Rural	(10.42)	(10.58)	(10.40)	(10.55)	(10.25)	(10.41)
	0.55	0.8	0.45	0.69	0.68	0.88
	(0.5)	(0.4)	(0.5)	(0.46)	(0.47)	(0.33)
Less than primary education	0.36 (0.48)	0.66 (0.47)	0.26 (0.44)	0.54 (0.50)	0.28 (0.45)	0.62 (0.48)
N of Obs.	706,134	119,591	517,049	82,362	2,436,388	575,442

Note: Numbers in parentheses are standard deviations

Table 12 Summary statistics for subsample of women's birth cohorts 1940–1979: ethnic majority versus ethnic minority

Censuses	Census 1989		Census 1999		Census 2009	
	Ethnic majority	Ethnic minority	Ethnic majority	Ethnic minority	Ethnic majority	Ethnic minority
Number of children	2.66	3.15	2.24	2.66	1.85	2.25
	(1.69)	(1.89)	(1.33)	(1.59)	(1.00)	(1.30)
Percent having						
No child	5.93	4.60	5.79	5.07	8.54	4.98
One child	19.92	14.88	22.56	17.40	24.53	21.32
2 children	26.98	22.07	37.75	31.40	47.21	40.81
At least 3 children	47.17	58.45	33.90	46.13	19.72	32.89
Proportion of sons	0.487	0.489	0.498	0.497	0.493	0.509
	(0.357)	(0.333)	(0.373)	(0.351)	(0.46)	(0.333)
Percent employed	0.82	0.93	0.79	0.92	0.87	0.96
	(0.39)	(0.25)	(0.41)	(0.26)	(0.34)	(0.19)
Mothers' age	32.21	31.90	33.94	32.60	34.24	32.44
	(6.93)	(7.39)	(7.13)	(7.23)	(7.47)	(7.65)
Mothers' age of first birth	23.288	22.659	23.315	22.528	23.748	22.868
	(3.744)	(4.076)	(3.806)	(3.779)	(4.102)	(3.977)
Rural	0.58	0.82	0.47	0.72	0.68	0.90
	(0.49)	(0.38)	(0.50)	(0.45)	(0.47)	(0.29)
Less than primary education	0.36	0.67	0.25	0.54	0.23	0.63
	(0.48)	(0.47)	(0.43)	(0.50)	(0.42)	(0.48)
N of obs.	248,858	38,044	262,071	45,423	1,411,727	424,808

Note: Numbers in parentheses are standard deviations

Table 13 Summary statistics for subsample of women with less than primary education of birth cohorts 1940–1979: ethnic majority versus ethnic minority

Censuses	Census 1989		Census 1999		Census 2009	
	Ethnic majority	Ethnic minority	Ethnic majority	Ethnic minority	Ethnic majority	Ethnic minority
Number of children	2.63 (2.47)	2.66 (2.50)	2.51 (2.04)	2.82 (2.14)	2.39 (1.44)	3.13 (1.72)
Percent having						
No child	29.07	28.35	19.43	16.22	9.57	5.21
One child	11.35	12.01	16.15	13.54	13.91	6.99
2 children	12.8	12.84	19.57	19.22	34.49	27.45
At least 3 children	46.78	46.8	44.85	51.02	42.03	60.35
Percent employed	0.69 (0.46)	0.91 (0.28)	0.68 (0.47)	0.88 (0.31)	0.74 (0.44)	0.90 (0.30)
Age	27.78	24.99	37.52	35.58	47.72	44.56
	(11.47)	(10.39)	(11.42)	(11.37)	(11.21)	(10.86)
Rural	0.67	0.88	0.59	0.81	0.78	0.92
	(0.47)	(0.33)	(0.49)	(0.39)	(0.41)	(0.27)
% of the sample	76.1	23.9	74.48	25.52	65.53	34.47
N of obs.	484,806	152,288	251,931	86,343	1,328,975	699,169

Note: Numbers in parentheses are standard deviations

Table 14 Summary statistics for subsample of women with at least primary education of birth cohorts 1940–1979: ethnic majority versus ethnic minority

Censuses	Census 1989		Census 1999		Census 2009	
	Ethnic majority	Ethnic minority	Ethnic majority	Ethnic minority	Ethnic majority	Ethnic minority
Number of children	1.38 (1.61)	1.43 (1.68)	1.76 (1.41)	1.91 (1.48)	2.11 (1.09)	2.22 (1.14)
Percent having						
No child	43.36	43.76	22.01	19.79	7.17	5.77
One child	16.47	15.39	22.26	20.01	15.38	14.57
2 children	17.57	17.2	30.42	30.83	48.64	47.79
At least 3 children	22.6	23.65	25.31	29.37	28.81	31.87
Percent employed	0.79 (0.41)	0.88 (0.33)	0.80 (0.40)	0.88 (0.32)	0.84 (0.37)	0.92 (0.27)
Age	25.09	24.53	34.04	33.73	43.17	42.64
	(8.93)	(8.29)	(9.48)	(9.13)	(9.14)	(8.71)
Rural	0.48	0.62	0.39	0.53	0.63	0.78
	(0.50)	(0.49)	(0.49)	(0.50)	(0.48)	(0.41)
% of the sample	91.54	8.46	90.78	9.22	65.53	34.47
N of obs.	426,080	37,178	357,225	34,366	1,623,833	185,979

Note: Numbers in parentheses are standard deviations

Table 15 Summary statistics for subsample of women in urban areas of birth cohorts 1940–1979: ethnic majority versus ethnic minority

Censuses	Census 1989		Census 1999		Census 2009	
	Ethnic majority	Ethnic minority	Ethnic majority	Ethnic minority	Ethnic majority	Ethnic minority
Number of children	1.63	1.51	1.72	1.83	1.88	2.03
	(1.91)	(1.99)	(1.48)	(1.66)	(1.08)	(1.38)
Percent having						
No child	40.2	47.15	24.18	25.13	10.6	12.65
One child	16.41	13.22	22.58	20.81	20.3	18.54
2 children	16.92	14.64	29.33	27.38	48.97	42.5
At least 3 children	26.47	24.99	23.91	26.68	20.13	26.31
Percent employed	0.68 (0.47)	0.70 (0.46)	0.71 (0.45)	0.77 (0.42)	0.70 (0.46)	0.76 (0.43)
Age	25.57	25.30	35.57	35.50	44.99	45.09
	(10.43)	(10.26)	(10.34)	(10.50)	(10.17)	(10.30)
Less than primary	0.27	0.41	0.22	0.35	0.23	0.42
education	(0.44)	(0.49)	(0.42)	(0.48)	(0.42)	(0.49)
% of the sample	92.65	7.35	91.6	8.4	91.63	8.37
N of obs.	313,224	24,256	285,705	25,627	774,025	70,569

Note: Numbers in parentheses are standard deviations

Table 16 Summary statistics for subsample of women in rural areas of birth cohorts 1940–1979: ethnic majority versus ethnic minority

Censuses	Census 1989		Census 1999		Census 2009	
	Ethnic majority	Ethnic minority	Ethnic majority	Ethnic minority	Ethnic majority	Ethnic minority
Number of children	2.0 (2.15)	2.43 (2.37)	2.20 (1.78)	2.66 (2.0)	2.33 (1.22)	2.90 (1.60)
Percent having						
No child	36.66	30	19.77	15.24	6.47	4.45
One child	13.2	13.19	17.83	14.35	12.53	8.57
2 children	15	14.29	24.71	22.92	43.20	33.79
At least 3 children	35.14	42.52	37.69	47.49	37.80	53.19
Percent employed	0.82 (0.38)	0.96 (0.20)	0.82 (0.38)	0.94 (0.24)	0.86 (0.35)	0.93 (0.26)
Age	24.98 (10.46)	24.31 (10.63)	35.10 (10.45)	34.57 (10.58)	45.17 (10.29)	44.31 (10.42)
Less than primary education	0.48 (0.50)	0.76 (0.43)	0.39 (0.49)	0.68 (0.47)	0.38 (0.49)	0.71 (0.45)
% of the sample	80.08	19.9	80.03	19.97	91.63	8.37
N of obs.	391,849	95,335	231,344	56,735	1,662,459	504,801

Note: Numbers in parentheses are standard deviations

Table 17 The estimates of the effect of the two-child policy on the probability of having more than 2 children – results from full sample

Dependent variable		More than 2 children		
Birth year × majority	Age in 1989	Coefficient	95% CI	
1941	48	0	−0.04	0.04
1942	47	−0.02	−0.07	0.02
1943	46	−0.01	−0.05	0.04
1944	45	−0.04	−0.08	0.01
1945	44	−0.02	−0.06	0.03
1946	43	−0.03	−0.07	0.01
1947	42	−0.01	−0.06	0.03
1948	41	0.01	−0.03	0.05
1949	40	0	−0.04	0.04
1950	39	−0.05	−0.09	−0.02
1951	38	−0.04	−0.08	−0.01
1952	37	−0.05	−0.09	−0.01
1953	36	−0.05	−0.09	−0.01
1954	35	−0.06	−0.1	−0.02
1955	34	−0.07	−0.1	−0.03
1956	33	−0.08	−0.12	−0.04
1957	32	−0.11	−0.15	−0.08
1958	31	−0.11	−0.15	−0.08
1959	30	−0.13	−0.17	−0.09
1960	29	−0.16	−0.19	−0.13
1961	28	−0.17	−0.2	−0.13
1962	27	−0.17	−0.2	−0.13
1963	26	−0.17	−0.2	−0.14
1964	25	−0.16	−0.2	−0.13
1965	24	−0.16	−0.19	−0.13
1966	23	−0.16	−0.19	−0.13
1967	22	−0.16	−0.19	−0.13
1968	21	−0.16	−0.2	−0.13
1969	20	−0.16	−0.2	−0.13
1970	19	−0.17	−0.2	−0.14
1971	18	−0.17	−0.2	−0.14
1972	17	−0.17	−0.2	−0.14
1973	16	−0.17	−0.2	−0.14
1974	15	−0.17	−0.2	−0.13
1975	14	−0.14	−0.18	−0.11
1976	13	−0.14	−0.17	−0.1
1977	12	−0.13	−0.16	−0.1
1978	11	−0.13	−0.16	−0.1
1979	10	−0.13	−0.16	−0.1

Table 18 The estimates of the effect of the two-child policy on mothers' age at first birth – results from subsample

Dependent variable	Mothers' age at first birth			
Birth year × majority	Age in 1989	Coefficient	95% CI	
1941	48	0.152	−0.795	1.099
1942	47	−0.451	−1.371	0.468
1943	46	−0.673	−1.593	0.247
1944	45	−0.233	−1.128	0.662
1945	44	−0.743	−1.602	0.116
1946	43	−0.912	−1.755	−0.069
1947	42	−0.959	−1.846	−0.072
1948	41	−0.371	−1.178	0.436
1949	40	−0.910	−1.699	−0.120
1950	39	−0.600	−1.373	0.174
1951	38	−1.037	−1.804	−0.270
1952	37	−0.533	−1.303	0.238
1953	36	−0.738	−1.496	0.020
1954	35	−0.407	−1.160	0.346
1955	34	−0.436	−1.181	0.309
1956	33	−0.569	−1.312	0.174
1957	32	−0.380	−1.125	0.364
1958	31	−0.446	−1.179	0.287
1959	30	−0.526	−1.269	0.217
1960	29	−0.386	−1.103	0.331
1961	28	−0.453	−1.168	0.263
1962	27	−0.600	−1.319	0.119
1963	26	−0.506	−1.220	0.208
1964	25	−0.576	−1.291	0.139
1965	24	−0.447	−1.162	0.268
1966	23	−0.517	−1.233	0.198
1967	22	−0.522	−1.238	0.193
1968	21	−0.426	−1.141	0.290
1969	20	−0.498	−1.214	0.217
1970	19	−0.493	−1.207	0.222
1971	18	−0.504	−1.219	0.212
1972	17	−0.496	−1.211	0.220
1973	16	−0.470	−1.185	0.246
1974	15	−0.368	−1.083	0.346
1975	14	−0.342	−1.057	0.372
1976	13	−0.254	−0.968	0.460
1977	12	−0.157	−0.871	0.556
1978	11	−0.028	−0.742	0.685
1979	10	0.040	−0.674	0.754

Table 19 The estimates of the effect of the policy on mothers' birth spacing – results from subsample

Dependent variable	Birth spacing at least 3 years			
Birth year × majority	Age in 1989	Coefficient	95% CI	
1941	48	−0.015	−0.104	0.074
1942	47	−0.019	−0.112	0.075
1943	46	0.038	−0.052	0.128
1944	45	−0.054	−0.144	0.035
1945	44	−0.047	−0.133	0.039
1946	43	−0.107	−0.189	−0.024
1947	42	−0.053	−0.138	0.032
1948	41	−0.047	−0.129	0.035
1949	40	−0.098	−0.179	−0.017
1950	39	−0.077	−0.152	−0.002
1951	38	−0.080	−0.154	−0.006
1952	37	−0.072	−0.149	0.005
1953	36	−0.071	−0.146	0.003
1954	35	−0.055	−0.129	0.019
1955	34	−0.034	−0.107	0.040
1956	33	−0.040	−0.113	0.033
1957	32	−0.021	−0.093	0.052
1958	31	−0.023	−0.095	0.048
1959	30	−0.014	−0.085	0.058
1960	29	−0.003	−0.071	0.066
1961	28	−0.003	−0.071	0.065
1962	27	0.016	−0.053	0.085
1963	26	0.023	−0.046	0.091
1964	25	0.039	−0.029	0.107
1965	24	0.019	−0.049	0.087
1966	23	0.039	−0.030	0.107
1967	22	0.050	−0.018	0.119
1968	21	0.081	0.012	0.150
1969	20	0.090	0.021	0.158
1970	19	0.100	0.031	0.169
1971	18	0.117	0.049	0.186
1972	17	0.123	0.054	0.192
1973	16	0.141	0.072	0.209
1974	15	0.134	0.065	0.203
1975	14	0.137	0.069	0.206
1976	13	0.156	0.087	0.224
1977	12	0.160	0.092	0.229
1978	11	0.155	0.087	0.223
1979	10	0.152	0.084	0.220

Table 20 The estimates of the effect of the two-child policy on proportion of sons in each family – results from subsample

Dependent variable	Proportion of sons in each family			
Birth year × majority	Age in 1989	Coefficient	95% CI	
1941	48	−0.048	−0.093	−0.003
1942	47	−0.050	−0.101	0.001
1943	46	−0.052	−0.099	−0.004
1944	45	−0.072	−0.119	−0.025
1945	44	−0.062	−0.107	−0.017
1946	43	−0.067	−0.111	−0.022
1947	42	−0.021	−0.066	0.024
1948	41	−0.043	−0.085	0.000
1949	40	−0.053	−0.095	−0.010
1950	39	−0.033	−0.073	0.008
1951	38	−0.061	−0.101	−0.021
1952	37	−0.050	−0.091	−0.009
1953	36	−0.054	−0.093	−0.015
1954	35	−0.065	−0.104	−0.026
1955	34	−0.053	−0.091	−0.015
1956	33	−0.062	−0.101	−0.023
1957	32	−0.055	−0.092	−0.017
1958	31	−0.063	−0.100	−0.025
1959	30	−0.048	−0.085	−0.010
1960	29	−0.059	−0.095	−0.024
1961	28	−0.055	−0.091	−0.020
1962	27	−0.045	−0.081	−0.009
1963	26	−0.057	−0.092	−0.021
1964	25	−0.070	−0.105	−0.034
1965	24	−0.057	−0.093	−0.021
1966	23	−0.051	−0.087	−0.015
1967	22	−0.062	−0.098	−0.026
1968	21	−0.073	−0.109	−0.037
1969	20	−0.061	−0.098	−0.025
1970	19	−0.070	−0.106	−0.034
1971	18	−0.065	−0.102	−0.029
1972	17	−0.059	−0.095	−0.023
1973	16	−0.066	−0.103	−0.029
1974	15	−0.075	−0.111	−0.038
1975	14	−0.076	−0.113	−0.039
1976	13	−0.070	−0.107	−0.033
1977	12	−0.063	−0.100	−0.026
1978	11	−0.071	−0.108	−0.033
1979	10	−0.067	−0.104	−0.031

Table 21 The estimates of the effect of the policy on gender of last birth – results from subsample

Dependent variable	Last birth is a boy			
Birth year × majority	Age in 1989	Coefficient	95% CI	
1941	48	−0.006	−0.089	0.077
1942	47	0.027	−0.061	0.115
1943	46	0.028	−0.057	0.113
1944	45	0.000	−0.085	0.084
1945	44	−0.001	−0.081	0.080
1946	43	0.024	−0.055	0.102
1947	42	0.013	−0.067	0.093
1948	41	0.036	−0.041	0.112
1949	40	0.002	−0.074	0.078
1950	39	0.027	−0.044	0.097
1951	38	0.001	−0.069	0.070
1952	37	0.019	−0.053	0.091
1953	36	0.023	−0.046	0.093
1954	35	0.016	−0.054	0.085
1955	34	0.002	−0.066	0.071
1956	33	0.001	−0.067	0.069
1957	32	0.014	−0.054	0.081
1958	31	−0.006	−0.072	0.061
1959	30	0.029	−0.038	0.095
1960	29	−0.001	−0.064	0.063
1961	28	0.025	−0.038	0.089
1962	27	0.018	−0.046	0.082
1963	26	0.004	−0.059	0.067
1964	25	−0.014	−0.077	0.050
1965	24	0.004	−0.060	0.067
1966	23	0.013	−0.050	0.077
1967	22	0.000	−0.063	0.064
1968	21	0.002	−0.062	0.066
1969	20	−0.003	−0.066	0.061
1970	19	−0.009	−0.073	0.055
1971	18	−0.001	−0.065	0.063
1972	17	0.001	−0.063	0.065
1973	16	−0.004	−0.068	0.060
1974	15	−0.014	−0.078	0.050
1975	14	−0.012	−0.076	0.052
1976	13	−0.006	−0.070	0.058
1977	12	0.004	−0.060	0.068
1978	11	−0.004	−0.068	0.060
1979	10	−0.005	−0.069	0.058

Table 22 The estimates of the effect of the policy on maternal employment – results from full sample

Dependent variable	Employed			
Birth year × majority	Age in 1989	Coefficient	95% CI	
1941	48	−0.014	−0.050	0.023
1942	47	0.026	−0.013	0.066
1943	46	0.027	−0.012	0.065
1944	45	0.006	−0.031	0.043
1945	44	0.005	−0.030	0.041
1946	43	−0.016	−0.052	0.020
1947	42	0.021	−0.017	0.059
1948	41	0.021	−0.015	0.058
1949	40	0.028	−0.008	0.064
1950	39	0.000	−0.033	0.032
1951	38	−0.015	−0.046	0.017
1952	37	0.028	−0.007	0.062
1953	36	0.015	−0.016	0.047
1954	35	0.016	−0.015	0.048
1955	34	0.020	−0.012	0.051
1956	33	0.009	−0.022	0.040
1957	32	0.015	−0.016	0.046
1958	31	0.012	−0.018	0.042
1959	30	0.018	−0.012	0.049
1960	29	0.022	−0.006	0.050
1961	28	0.019	−0.008	0.047
1962	27	0.016	−0.012	0.044
1963	26	0.026	−0.002	0.053
1964	25	0.030	0.002	0.058
1965	24	0.028	0.000	0.056
1966	23	0.026	−0.002	0.054
1967	22	0.023	−0.005	0.050
1968	21	0.030	0.002	0.058
1969	20	0.028	0.000	0.056
1970	19	0.022	−0.006	0.050
1971	18	0.025	−0.003	0.053
1972	17	0.022	−0.006	0.050
1973	16	0.024	−0.004	0.052
1974	15	0.028	0.000	0.057
1975	14	0.028	0.000	0.056
1976	13	0.041	0.013	0.069
1977	12	0.047	0.018	0.075
1978	11	0.044	0.016	0.072
1979	10	0.053	0.025	0.082

Table 23 The effect of the two-child policy on family size with further controls

Outcomes	More than 2 children			Number of living children		
Under age 30 in 1989 × majority	−0.152***	−0.144***	−0.133***	−0.192***	−0.166***	−0.123***
	(0.005)	(0.005)	(0.005)	(0.031)	(0.031)	(0.03)
Age 30–39 in 1989 × majority	−0.067***	−0.067***	−0.063***	−0.196***	−0.196***	−0.180***
	(0.006)	(0.006)	(0.006)	(0.035)	(0.035)	(0.035)
Number of obs.	3,197,622	3,197,622	3,188,893	3,197,622	3,197,622	3,188,893
Controls						
Year indicators	Y	Y	Y	Y	Y	Y
Rural indicators	N	Y	Y	N	Y	Y
Education indicators	N	N	Y	N	N	Y

Note: Numbers in parentheses are standard errors

Other covariates included in the models are indicators of women under age 30 in 1989, women ages 30–39 in 1989, ethnic majority, age, age squared, age × majority, age squared × majority, and province fixed effects

$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$

Table 24 The effect of the two-child policy on proportion of sons in each family and maternal employment with further controls

Outcomes	Proportion of sons in each family			Employed		
Under age 30 in 1989 × majority	−0.013*	−0.012*	−0.011*	0.013**	0.019***	0.015***
Age 30–39 in 1989 × majority	(0.005)	(0.006)	(0.006)	(0.004)	(0.004)	(0.004)
	−0.0051	−0.005	−0.0048	−0.0008	−0.0008	−0.001
	(0.006)	(0.007)	(0.007)	(0.005)	(0.005)	(0.005)
Number of obs.	1,852,725	1,852,725	1,850,465	3,126,106	3,126,106	3,123,326
Controls						
Year indicators	Y	Y	Y	Y	Y	Y
Rural indicators	N	Y	Y	N	Y	Y
Education indicators	N	N	Y	N	N	Y

Note: Numbers in parentheses are standard errors

Other covariates included in the models are indicators of women under age 30 in 1989, women ages 30–39 in 1989, ethnic majority, age, age squared, age × majority, age squared × majority, and province fixed effects

$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$

Table 25 The estimates of the effect of the two-child policy on other household outcomes

Outcomes	Having at least one child	Getting married
Under age 30 in 1989 × majority	0.0001	0.007**
	(0.003)	(0.003)
Age 30–39 in 1989 × majority	0.003	0.004
	(0.004)	(0.004)
Number of Obs.	3,197,622	3,196,708

Note: Numbers in parentheses are standard errors

Other covariates included in the models are indicators of women under age 30 in 1989, women ages 30–39 in 1989, ethnic majority, age, age squared, age × majority, age squared × majority, year indicators, and province fixed effects

$^{*}p < 0.05$, $^{**}p < 0.01$

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