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Democracy and Demography: Societal Effects of Fertility Limits on Local Leaders*

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Abstract

We investigate whether legally restricting elected leadership positions to candidates with low fertility leads to a society-wide fertility decline in a low-income democracy with high political participation. Exploiting quasi-experimental variation in fertility limits imposed on village council members to curb population growth in India, we find that rural households significantly reduce their fertility in response. Socially disadvantaged families exhibit the greatest reduction in marginal fertility immediately following passage of the limits, but the decline in the long-run “stock” of children is more evenly spread across socioeconomic groups. The limits however also increased the already male-biased sex ratio at birth in castes with strong son preference. Therefore, while restricting access to elected office may be effective in achieving certain social objectives, it can also have adverse consequences for inequality in other dimensions.

JEL Codes: J13, J16, H75, O11

Keywords: India, Panchayat Elections, Fertility Limits, Sex Ratio

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1 Introduction

Representative democracy promotes stable economic growth (Rodrik (2000), Mobarak (2005), Papaioannou and Siourounis (2008)), narrows large income inequalities (Acemoglu and Robinson (2000), Gradstein (2007)), and prevents elite capture of property rights and productive resources (Brown and Mobarak (2009), Foster and Rosenzweig (2004), Bardhan and Mookherjee (2000)), all of which are critical for welfare improvement in developing countries. The literature on political selection suggests that these positive effects of democracy may be due to the higher quality of leaders in democratic relative to autocratic systems (Besley and Coate (1997), Osborne and Slivinski (1996), Besley and Reynal-Querol (2011)) as voters choose from a wider pool of candidates.¹ Good leaders also serve as role-models and effect social change by raising aspirations.² Consequently perhaps, legal restrictions on who can become a candidate or a leader in a democratic system are usually minimal.³

Despite this theoretically open access to the candidate pool in a democracy, in practice, political networks, campaign costs, and other socioeconomic inequities create substantial entry barriers for a significant share of citizens.⁴ Moreover, democratic accountability may not work if voters have imperfect information on candidates' characteristics⁵ or if they prefer to elect leaders who can provide patronage at the expense of other constituents. Therefore, several countries have sought to improve candidate quality, and thus policy outcomes, by imposing minimum requirements for candidate characteristics, such as educational qualification and criminal records.⁶ However, little is known about the effects of these laws on leaders' performance, political representation, and constituent outcomes.

¹Democratic leaders may also perform better due to greater accountability, conditional on quality, than autocrats, especially if voter participation is sufficiently high and punitive towards poorly performing incumbents (Acemoglu and Robinson (2006), Bidner and Francois (2013), Mueller and Stratmann (2003)).

²In general, individuals in positions of authority do exert considerable influence on their followers' behaviors and outcomes (Beaman et al. (2012), Bettinger and Long (2005), Jensen and Oster (2009), Chong et al. (2012), Olivetti et al. (2013), and Bassi and Rasul (2014)).

³Most countries do have natural citizenship, minimum age, linguistic proficiency, and/ or no conflict of interest requirements.

⁴In order to improve the political representation of disadvantaged groups several countries have legislated quotas, for instance, based on gender and caste in India (Chattopadhyay and Duflo (2004), Bhalotra et al. (2013), and Kapoor and Ravi (2014)).

⁵Ferraz and Finan (2008) and Banerjee et al. (2011) show that increasing transparency can improve political accountability.

⁶For instance, Angola, Azerbaijan, the Philippines, and Turkey have minimum education requirements for Presidential candidates.

In this paper we examine the impact of a unique policy experiment in India, whereby several states legally restrict individuals with more than two children from contesting local government (*Panchayat*) elections, on fertility in the general population. Although this manipulation of the candidate pool has a policy objective, namely to lower population growth, it is not intended to directly improve leaders' performance as education requirements might. Instead, these legislations seek to improve economic outcomes by precipitating fertility decline. There are at least three different channels that may underlie any causal effect of the two-child limits on statewide fertility. If elected representatives serve as role models, their constituents may be indirectly affected by these limits as they emulate their leaders' fertility choices ("role-model channel"); in fact, this appears to be the primary mechanism the policymakers had in mind when these laws were enacted.⁷ Societal fertility decline may also be driven by citizens' desire to maintain eligibility for future Panchayat candidacy ("aspirations/incentives channel").⁸ Additionally, even if citizens do not have political ambitions, they may adjust fertility if the limits signal that similar policies may be enacted for government jobs, for instance ("anticipation channel").

The Panchayat system in India comprises village-, block-, and district-level councils that exercise considerable power in their constituencies. Starting in 1992, eleven states have enacted the fertility limits for at least some years and they remain in effect in seven major states. These laws provided a one-year grace-period from the time of announcement, during which an individual could have additional children and still remain eligible for election. However, for people with two or more children by the end of the grace-period, a subsequent birth leads to disqualification. Individuals with fewer than two children by the end of the grace-period are limited to at most two children afterwards to maintain eligibility.⁹ We exploit the quasi-experimental geographical and temporal variation in announcement of these fertility limits across states to estimate their causal impacts on citizens' demographic outcomes in a difference-in-differences framework.

We find that among couples with two children when the law is announced, the hazard of third birth declines by 2.06 to 2.65 percentage points (p.p.), which is about 10 to 11 percent

⁷Source: <http://www.nytimes.com/2003/11/07/world/states-in-india-take-new-steps-to-limit-births.html>.

⁸For more on aspirations in developing countries, see Appadurai (2004), Mookherjee et al. (2010), Ray (2006), Genicot and Ray (2014), Dalton et al. (2014), and Bernard et al. (2012).

⁹The same rules apply for dismissal of an elected member who exceeds the fertility limit while in office.

of the baseline hazard of third birth. These magnitudes are comparable to the effects of well-known fertility control programs such as the Matlab intervention in Bangladesh, which reduced child-to-mother ratios by 16 percent, and the combined Navrongo interventions in Ghana, which reduced children ever born to treated women by 9 percent ([Canning and Schultz \(2012\)](#)). There are no significant effects on marginal fertility of couples who already had three or more children at the time the laws were announced, as they are more likely to have already achieved their desired family size.¹⁰ We also examine the impact of the limits on the total “stock” of fertility and find that the probability of having three or more children in any given year declines by 1.33 p.p. or 4.42 percent and the probability of having only two children increases by 0.75 p.p. or 3.41 percent in enacting states. However, the fertility limits also adversely affect the already male-biased sex ratio among social groups with high son preference, thus increasing the number of missing girls. The human capital and income gains from fertility decline are therefore undermined by worsening gender inequality ([Doepke et al. \(2012\)](#)).

The fertility limits potentially have crucial implications for the efficacy of democratic institutions in protecting the welfare of the socially disadvantaged, who may have more children than elites due to lower access to contraception and higher risk of child mortality, and depend more on political representation to obtain resources prone to elite capture. On the other hand, low socioeconomic status families are more “treatable” by the limits as they have a higher baseline hazard of third birth relative to couples of higher socioeconomic status. We find that the declines in the hazard of third birth are concentrated in families with low socioeconomic status (i.e., couples with low wealth or no schooling), and from lower castes that rely on mandated reservations of Panchayat seats for elected political representation. Moreover, the substitution away from more than two children and towards exactly two children, while not as precisely estimated, is similar between poorer, less educated households and wealthier, literate households. This reduces concerns that public resources may become more prone to capture by local elites ([Bardhan and Mookherjee \(2000\)](#)) once the limits are in effect. Our results suggest that disadvantaged households are not only aware of the two-child laws but also respond in a manner that allows them to continue benefitting

¹⁰Average completed fertility in treated states before the law is 2.8 children.

from mandated gender- and caste-based quotas for Panchayat membership.¹¹

Back-of-the-envelope calculations using our estimates suggest that 615,390 couples or 1,230,780 individuals responded to the fertility limits in treatment states by decreasing the likelihood of having more than two children. These include couples who had two children at announcement and stopped childbearing thereafter; those who had one child at announcement and stopped after having the second child; and couples who began childbearing post-announcement and had no more than two children. While we cannot credibly say how many of these couples were driven by political aspirations or were merely emulating their role-models or were fearful of stricter fertility limits in the future, the magnitudes are consistent with the high participation of voters and candidates in Panchayat politics. At the minimum, our paper shows that policy interventions in the political sphere can have substantial impacts in a highly participatory democratic society.

Our paper make novel contributions to the literature on family planning interventions, which can promote economic growth, human capital accumulation, and women’s empowerment when effective (e.g., [Miller \(2010\)](#), [Ashraf et al. \(2013\)](#), [Joshi and Schultz \(2007\)](#), and [Rosenzweig and Zhang \(2009\)](#)). Additionally, we contribute to two other literatures: (i) on the relationship between fertility and career decisions and (ii) on the determinants of sex ratios. Improvements in labor market opportunities, especially for women, increase the opportunity cost of having children and thereby lower fertility ([Chiappori et al. \(2002\)](#), [Rosenzweig and Wolpin \(1980\)](#)). We examine a similar relationship between fertility and political careers where the change in the opportunity cost of children is caused by the two-child limits. Recent papers have also highlighted the effect of fertility decline on rising sex ratios in societies like India where sons are preferred ([Ebenstein \(2010\)](#), [Anukriti \(2014\)](#), [Jayachandran \(2014\)](#)). We augment this second literature by analyzing a new source of fertility decline and show that it too has an unintended effect on sex ratios.

The rest of the paper is organized as follows. Section 2 discusses the legislations in detail. Sections 3 and 4 describe our data and empirical strategy. Section 5 presents the estimation results, Section 6 discusses the results, and Section 7 concludes the paper.

¹¹[Beaman et al. \(2012\)](#) and [Chattopadhyay and Duflo \(2004\)](#) show that quotas-based improvements in the representation of women and lower castes in Panchayats enhances targeting of public expenditure towards these groups and reduces discrimination against them.

2 Background

India is the world’s second most populous country and houses a third of its poorest citizens (Olinto et al. (2013)). Consequently, population control remains a policy priority. Based on the recommendations of the 1992 Committee on Population, several states enacted the two-child laws for Panchayat candidates.¹² These laws aim to lower fertility through the role-model channel. However, they also incentivize individuals who intend to contest elections to plan smaller families.

India has a three-tiered system of local governance in rural areas, known as the Panchayati Raj. It comprises village-level councils (*Gram Panchayat*), block-level councils (*Panchayat Samiti*), and district-level councils (*Zila Parishad*). Regular Panchayat elections have taken place every five years in most states. The village councils are the building blocks of the Indian democratic system and exercise considerable power in their constituencies. They receive substantial funds from national and state governments,¹³ and are authorized to implement development schemes.¹⁴ Panchayats are also responsible for providing public goods such as village roads, wells, and water-works. They can collect taxes and license fees, and receive seignorage from the auction of local mineral and forestry resources.

The typical monthly salary of a Panchayat chief is about USD 50 - USD 60 and other council members are paid less. While these official wages are not substantial, the potential private returns from political rents and corrupt practices may provide a strong incentive for becoming an officeholder. According to the Association for Democratic Reforms, an average candidate spends USD 400 - USD 800 during a Panchayat election.¹⁵ However, the benefits from even one term as a Panchayat member are likely to be much higher. The average declared wealth of re-contesting candidates (for elections to the Parliament and state legislative assemblies) in 2004 was 134% higher than their wealth during the first election (Sastry (2014)). Fisman et al. (2014) show that the annual asset growth of winners in state elections

¹²In fact, the Committee recommended these restrictions for all elected positions—from Panchayats to the national Parliament.

¹³For example, in Tamil Nadu, all Panchayats received at least USD 4,900 in annual state grants in 2009-10, and 35% of them received funds in the range of USD 16,330-40,800. These are significant budgets considering that India’s annual per capita income was USD 1,570 in 2013 (Source: The World Bank).

¹⁴Panchayats are often authorized to identify local beneficiaries of major central and state development schemes, such as the National Rural Employment Guarantee Scheme.

¹⁵Source: <http://www.ndtv.com/india-news/the-rs-81-500-crore-lie-565175>

is 3-5 percentage points higher than that of runners-up. Similar statistics are not available for Panchayat candidates, but the returns are likely to also be large.

The average population per village Panchayat is about 3,100, although the size varies widely. The minimum age to contest elections is 21 years. There are no term limits on Panchayat members. In Rajasthan and Uttar Pradesh, respectively, 19% and 33% of council chiefs were under 36 years old and 56% and 51% were in the 36-50 year age-group.¹⁶ The council members are typically younger: 47% of Panchayat members in 2012 in Rajasthan were under 36 years of age and 41% were in the 36-50 year age-group. The PR Act requires that at least one-third of all member and chief positions are reserved for women.¹⁷ Similarly, positions are reserved for Scheduled Castes (SC) and Scheduled Tribes (ST) in proportion to their population share. Quotas are implemented in a stratified manner—among positions reserved for SC, ST, and “general” castes, one-third are randomly chosen for women. We also present results separately for Other Backward Castes (OBC), which are non-SC/ST castes identified by individual states or the central government as economically and socially disadvantaged, and therefore qualifying for affirmative action.

Rajasthan was the first state to introduce the two-child limit for its village councils in 1992;¹⁸ this requirement was later included in the state’s 1994 PR Act.¹⁹ Andhra Pradesh and Haryana announced their legislations in 1994,²⁰ although the latter revoked its law in 2006. Orissa announced the limit for its district councils in 1993 and for the village and block councils in 1994. Himachal Pradesh (HP), Madhya Pradesh (MP), and Chhattisgarh²¹ introduced their laws in 2000 and repealed them in 2005. In Maharashtra, the law has been in retrospective effect since 2002. Lastly, Bihar and Uttarakhand adopted the limit respectively in 2002 and 2007, but only for municipal elections. Table 1 presents a more detailed timeline for the announcement, grace-period, and implementation of these laws²² and Table A.2 shows

¹⁶In West Bengal, the average age of chiefs was 36 years in 2000 ([Chattopadhyay and Duflo \(2004\)](#)) and in Andhra Pradesh it was 43 years in 2011 ([Afridi et al. \(2014\)](#)).

¹⁷In 14 states, half of all seats are reserved for women.

¹⁸Rajasthan’s law predates the recommendations of the Committee on Population.

¹⁹The 1994 Act included a grace-period from April 23, 1994 to November 27, 1995. Effectively, this resulted in a nearly three-year grace-period since the original announcement was made in 1992.

²⁰However, since the 1994 elections in Haryana took place before the announcement and since members are elected for a period of five years, no one was disqualified during 1995-2000.

²¹Chhattisgarh inherited the law when it was carved out of MP in 2000. Since 2004, candidates below 30 years of age in Chhattisgarh are also required to be literate.

²²This information is largely based on [Buch \(2005\)](#) and [Buch \(2006\)](#).

the election years for which they were effective. The relevant clauses from each state’s PR Act are presented in Appendix B.

Candidates do not have to explicitly state their number of children when filing their nomination papers. However, they have to declare that, to the best of their knowledge, they are qualified for the Panchayat seat.²³ Table A.1 shows the number of Panchayat members that have been disqualified under these laws in Haryana, Rajasthan, MP, and AP during 2000-2004.²⁴ Newspaper reports suggest that, in some instances, the fertility limits have led to abandonment of wives, selective abortion of female fetuses, and giving up of children for adoption to avoid disqualification. Consequently, implementing states have faced criticism from women’s rights advocates and civil society organizations, as well as from the central government.²⁵ The revocation of the limits in four states may have been in response to this pressure. To summarize, eleven states have imposed fertility limits on Panchayat members for at least a few years and they remain in effect in seven states.

3 Data

We utilize three cross-sectional rounds of the National Family Health Survey (NFHS-1, 2, 3) of India that were conducted in 1992-93, 1998-99, and 2005-06. Each round is representative at the state-level and includes a complete retrospective birth history for the woman interviewed, containing information on the month and the year of birth, birth order, and mother’s age at birth. We combine these birth histories to construct an unbalanced woman-year panel; a woman enters the panel in her year of first marriage and exits in her year of survey.

For consistency across rounds, we limit the sample to women in the 15-49 age-group who were married at the time of survey.²⁶ We also drop women (i) whose marriage took place more than 20 years before the survey to avoid issues related to imperfect recall, (ii) whose husband’s age was below 15 or above 80 in the year of survey, and (iii) who had given birth to more than ten children, to prevent any composition-bias since these women are likely to

²³The Returning Officer (nominated by the Election Commission) is responsible for scrutinizing the information submitted by the nominees and any objections raised by the rival candidates, general public, or the media.

²⁴Data for the remaining states and years is not readily available and is being collected by the authors.

²⁵http://policydialogue.org/files/events/Aiyar_Key_Role_of_Panchayati_Raj_in_India.pdf

²⁶The questionnaires were administered to 13-49-year old ever-married women in NFHS-1 and 15-49-year old ever-married women in NFHS-2, 3.

be fundamentally different from rest of the sample. Lastly, we exclude mothers who have had twins since multiple births in our context are largely unplanned and do not reflect parents' fertility preferences. However, all our results are robust to the inclusion of these observations.

Our final sample comprises 156,541 women and 381,124 births from 18 major states²⁷ and covers the time period 1973-2006. We define treatment based on the year of announcement of the law, i.e., the earliest year when the law might have had an effect in a state. Since the most recent year in our sample is 2006, we cannot credibly examine the effect of revocations that took place in 2005.²⁸ However, we have a large number of post-announcement years, ranging from 4 to 13 years, to estimate the long-term effect of the fertility limits.

Table 2 presents the sample means and standard deviations for the key variables used in our analysis, separately for never-treated and treated states. We further split the treated sample into pre- and post-treatment observations. About two-thirds of women in our sample live in a rural area. A majority of them are Hindus, with a larger share of Hindus among treated relative to never-treated households. In terms of caste-composition, scheduled castes comprise about 16 percent of the sample. Educational attainment of women is low, with about half the sample being uneducated; in comparison, 26 percent of the husbands are uneducated. The pre-treatment average terminal fertility (as measured by fertility of women more than 40 years old) in treated states is 2.8.

The sample means for the three groups in Table 2 are similar along many socioeconomic dimensions. Nevertheless, to ensure that our estimates are not confounded by underlying differences between these samples, we control for religion, caste, standard of living, husband's and wife's years of schooling, and residence in an urban area in all regressions. To take into account state-specific factors, we include state fixed effects and state-specific linear time trends (or state-year fixed effects). We also show (in the next section) that the timing of announcement of the limits across states is uncorrelated with changes in these socioeconomic characteristics across states and over time.

²⁷The states of Uttarakhand, Jharkhand, and Chhattisgarh were, respectively, carved out from Uttar Pradesh (UP), Bihar, and MP in 2000. Since our data does not include districts-identifiers for all rounds, we subsume these three new states into their parent states for our analysis.

²⁸The only other source of demographic data after 2006 is the National Sample Survey (NSS) of India. However, the household roster in the NSS does not match mothers with their children.

4 Empirical Strategy

The goal of our empirical strategy is to estimate the causal effect of the two-child limits on candidates in village council elections in a state on fertility-related outcomes among residents in the same state. To do so, we utilize the quasi-experimental geographical and temporal variation in announcement of these laws across Indian states. Although eleven states have enacted such a law thus far, due to data limitations we can estimate the impact for only seven (eight) states: Rajasthan, Haryana, AP, Orissa, HP, MP (including Chhattisgarh), and Maharashtra. This is because the limits came into effect in Bihar and Gujarat after 2006, which is the most recent year in our dataset; so in our sample these states are in the control group. Gujarat announced its law in 2005, so we can potentially include it in the treatment group and use 2006 as the post-treatment year; doing so makes no difference to our results. Uttarakhand announced its law for urban municipal elections in 2002, however, we exclude it from the treatment group because Uttarakhand was a part of Uttar Pradesh until 2000 and we cannot distinguish between the two in the pre-2000 sample.²⁹ Moreover, we do not find significant effects in urban areas, and thus focus on the rural sample, and Uttarakhand has not enacted a limit for rural Panchayats. For these reasons, we keep Uttarakhand in the control group. In addition to Bihar, Gujarat, and Uttarakhand, our control group comprises nine other states. Figure 1 depicts the treatment and control states in a map.

We begin by examining how the hazard of birth evolves before and after the laws are announced, in an event-study framework. Specifically, for a woman i of age a in state s and year t we estimate the following regression specification:

$$Y_{iast} = \sum_{k=-6}^5 \alpha_k T_s * Post_{s,t+k} + \sum_{k=-6}^5 \beta_k Post_{s,t+k} + X_i' \delta + \gamma_s + \theta_t + \psi_a + \nu_s * t + \epsilon_{iast} \quad (1)$$

We include both treatment and control states in specification (1). For a treatment state, $Post_{s,t+k}$ indicates k years during which the law is in place; we assign the year before the year of announcement as the omitted year. For a control state, $Post_{s,t+k}$ indicates k years during which a fictitious law is in place; we assign the same announcement year to a control state as its neighboring treatment state. The outcome variables Y_{iast} are indicators for first,

²⁹Note that Uttar Pradesh has never enacted a two-child limit for its local politicians.

second, third, fourth, and fifth birth. In regressions where the outcome indicates a birth of order b , the sample is restricted to years after birth $(b - 1)$, until birth b , and to women whose previous $(b - 1)$ children were born before the year of announcement of the law in their respective state. For example, the hazard of third birth is estimated using years after the year of second birth and excluding the years after the year of third birth for women whose first two children were born before announcement of the fertility limits. We control for fixed effects for state, year, and woman's age (γ_s , θ_t , and ψ_a , respectively), state-specific linear trends ($\nu_s * t$), and the following covariates (X_i): five categories each for a woman's and her husband's years of schooling, indicators for religion (five categories), caste (four categories), and standard of living (three categories) of the household, and for the year of interview.

The β_k coefficients in (1) capture how the hazard of birth evolves in control states and the α_k coefficients capture the changes in the hazard of birth in treatment states after differencing out the control group. If there is no noticeable pre-trend in the differential hazard of birth across treatment and control states, we can interpret the α_k coefficients during the post-announcement years as the causal effect of the limits. In order to measure the net effect of the limits on birth hazards we pool the event study coefficients in (1) and estimate:

$$Y_{iast} = \omega + \alpha Treat_{st} + X'_i \delta + \gamma_s + \theta_t + \psi_a + \nu_s * t + \mu_{sa} + \epsilon_{iast} \quad (2a)$$

$$Y_{iast} = \omega + \alpha T_s * Post_{st} + \beta Post_{st} + X'_i \delta + \gamma_s + \theta_t + \psi_a + \nu_s * t + \mu_{sa} + \epsilon_{iast} \quad (2b)$$

Specification (2b) corresponds to (1), except that we pool the α_k and β_k coefficients into α and β , respectively. Additionally, we estimate (2a) where instead of assigning fictitious treatment years to control states, we define treatment as zero for them throughout the sample, i.e., treatment is defined in terms of the variable $Treat_{st}$ which is equal to one for women residing in the treated states if $t \geq$ the year of announcement, and zero otherwise; thus, $Treat_{st}$ is always zero in (2a) for control states. Additionally, we control for state \times mother's age fixed effects (μ_{sa}) and fixed effects for years since last birth or, in case of the hazard of first birth, years since marriage. Where the outcome indicates a birth of order b , we again restrict the sample to women whose previous $(b - 1)$ births took place before the law was announced in their respective states. These regressions, therefore, capture the effects of the fertility limits on *marginal* fertility of affected households.

While specifications (1), (2a), and (2b) tell us how marginal fertility was affected for

couples who already had a certain number of children when the limits were announced in their state, they do not measure the overall impact of the laws on the “stock” of fertility. Therefore, we re-estimate equations (2a) and (2b) using indicators for whether a woman i of age a in state s and year t reports having one, two, three, four, and five living children in year t as the outcome variables. Unlike the hazard analysis, however, in this case we do not impose any restrictions in terms of the prior number of children and use all available years for each woman. If the two-child limits are effective, we expect the likelihood of having two children to increase in the treatment states after the laws have been announced.

The two-child laws may also affect the sex ratio of births. For instance, parents who do not have the desired number of sons when the law is announced and who can afford to have an additional birth without violating the limit may be more likely to practice sex-selection for that additional birth. Therefore, we estimate the impact of the fertility limits on the sex ratio at second and higher parity births for couples whose first child was born before the limits were announced in their state. Specifically, we estimate the following specifications where the outcome variable is an indicator for the child being male:

$$Male_{iast} = \alpha + \beta Treat_{st} + X_i' \delta + \gamma_s + \theta_t + \psi_a + \nu_s * t + \mu_{sa} + \phi Girl_i + \epsilon_{iast} \quad (3a)$$

$$Male_{iast} = \alpha + \beta_1 T_s * Post_{st} + \beta_2 Post_{st} + X_i' \delta + \gamma_s + \theta_t + \psi_a + \nu_s * t + \mu_{sa} + \phi Girl_i + \epsilon_{iast} \quad (3b)$$

These specifications are similar to (2a) and (2b) except that we also control for the sex of the first child. We focus on second and higher parity births as prior literature on India has shown that, despite the availability of prenatal sex-determination technology, sex of the first birth is plausibly random (Bhalotra and Cochrane (2010), Das Gupta and Bhat (1997), Visaria (2005)) and most instances of sex-selection occur for higher-order births. This finding is consistent with recent survey data that suggests that Indian parents do not always prefer having a son over a daughter—Jayachandran (2014) finds that although the vast majority of families want to have a son if they can only have one child, at a family size of two they prefer having one daughter and one son over having two sons. As desired and actual fertility in India are well above one, it is reasonable to assume that parents are not averse to having one daughter, despite a strong desire for at least one son. In fact, Table 2 shows that the sex ratio at first birth in our sample is “normal” (i.e., between 0.516 and 0.519) in the never-treated states and in the treatment states (both pre- and post-treatment). It is also well-established

that parents whose first child is a girl are more likely to practice sex-selection at higher-parity births (e.g., [Pörtner \(2010\)](#), [Rosenblum \(2013\)](#), [Anukriti et al. \(2016\)](#)) since they desire at least one son. Therefore, we control for the sex of the first child in (3a) and (3b). Although we explicitly show that the fertility limits did not change the sex ratio of first births, to ensure that our results are not biased by sex-selection at first parity, we restrict the sample to couples whose first child was born before the limits were announced in their state.

The inclusion of state and year fixed effects in our specifications controls for all time-invariant state-level variables and state-invariant time effects that might affect the outcomes of interest. The state-specific time trends account for differential linear trends in fertility and sex-selection patterns across states over time (e.g., due to differential growth rates of state GDP or availability of abortion and other health services). The inclusion of state \times mother’s age fixed effects controls for any confounding differences in the age composition of mothers across states. Since treatment varies at the state level, we cluster standard errors by state to allow for correlations in the error terms of women in each state. As the total number of treatment and control states in our sample is 18, we also report standard errors based on a clustered (by state) wild bootstrap- t procedure described in [Cameron et al. \(2008\)](#) to address econometric issues pertaining to a small number of clusters.³⁰

The underlying identifying assumption in our analysis is that the state-year variation in the timing of law announcement is uncorrelated with other time-varying determinants of the outcomes of interest. Although we control for state-specific linear trends in our regressions, we also explicitly test if the timing of announcement is correlated with other socioeconomic characteristics that vary by state and time. In Table 3 we present the coefficients from regressions that use various maternal, paternal, and household characteristics as dependent variables in the estimation of equation (2a) with state and year fixed effects, and state-specific time trends, but without any other controls for the rural sample. None of the 21 coefficients in Table 3 are significant, thus eliminating any concerns about endogenous timing of announcements.³¹ Moreover, during the time-period we examine, there were no other state-specific programs in the treatment states that promoted smaller families and whose timing coincided with the fertility limits.

³⁰We use the STATA code written by [Busso et al. \(2013\)](#) that computes the errors by assessing the fraction of bootstrap test statistics (in 1,000 repetitions) greater in absolute value than the sample test statistic.

³¹The same is true if we include the urban sample in this regression.

5 Results

Next we present regression estimates of the causal impact of the fertility limits on marginal fertility, on the total number of living children in a year, and on the sex ratio of second and higher parity births. We find no significant effects for urban areas and hence focus only on the rural sample for the rest of the paper. To the extent that most of the fertility limits in our sample period were enacted for rural Panchayats, it is unsurprising that we do not find substantial effects in the urban sample. The urban estimates are available upon request.

5.1 Effects on marginal fertility

We start by examining the evolution of the hazard of birth (for birth orders one to four) in treatment and control states using specification (1). The estimated α_k coefficients for the annual hazard of births over a 11-year period are presented in Figure 2. The corresponding regression coefficients with standard errors clustered by state are shown in Appendix Table A.3. The vertical line in Figure 2 indicates the year before announcement for treatment states and the year before the fictitious announcement year for control states. The difference between treatment and control states is quite flat for all birth orders in the years before announcement, indicating that pre-trends in the hazard of birth are parallel between treated and control states. In fact, none of the lagged α_k coefficients in Appendix Table A.3 are significant, supporting the parallel trends assumption. However, after the fertility limits are announced there is an immediate increase in the relative hazard of second births in treated states, which declines in the following years. Moreover, the hazard of third birth shows a visible and sustained decline during all post-announcement years. There is no visible impact on the hazards of first and fourth births.³²

This decline in the hazard of third birth suggests that the fertility limits induced couples in treatment states who had two children at the time of announcement, and were thus at the margin of electoral eligibility, to sacrifice third births. Although this does not necessarily imply a decrease in completed fertility, the fact that the hazard of third birth continues to decrease for at least five years after the announcements indicates that our results are less likely to be driven by a mere postponement of third births. On the other hand, the shape of the second birth hazard does suggest a shift in timing of second births. Although the fertility

³²We also find no effect on the hazard of fifth birth.

limits do not impose any restrictions on the timing of second birth, it appears that couples who had one child at announcement have their second child sooner than they would have in the absence of these eligibility requirements. This response could, for example, be driven by fear or anticipation of even stricter limits on fertility in the near future. The lack of a significant decline in the hazard of fourth births is intuitive as these couples already had three children before announcement and had, therefore, most likely achieved their desired fertility; so they were less likely to have a fourth birth even in the absence of the new laws. Similarly, an insignificant impact on the hazard of first birth makes sense as desired fertility is above one in India and the fertility limits do not impose any constraints on first births.

Next we present the net impacts of the limits on marginal fertility estimated using specifications (2a) and (2b). Table 4 presents results for the hazard of third birth while the estimates for first, second, and fourth births are in Appendix Table A.4. The hazard of third birth is presumably the most relevant margin at which we expect the fertility limits to operate—at births below birth order three the policy does not bind, and at birth orders beyond three households have largely satisfied their desired fertility by the time the policy is announced. Panel A of Table 4 corresponds to specification (2a) and panel B corresponds to (2b). The standard errors in brackets are clustered by state and in parentheses are wild-cluster bootstrapped by state. In column (1), where we control for fixed effects for state, year, and years since second birth, and for the vector of socioeconomic characteristics, we find that the limits decreased the hazard of third birth, but insignificantly. The effect remains negative and becomes significant in column (2) once we control for state-specific linear time trends. We find that the probability of third birth declines by 2.13 percentage points (p.p.) in column (2) of panel A and by 2.65 p.p. in column (2) of panel B. These effects are large, respectively translating into 10 percent and 11 percent declines from the baseline probability of 21.31 percent of having a third birth in a given year before announcement. The inclusion of state \times mother’s age fixed effects in column (3) changes the estimated decline in third births marginally to 2.06 p.p. in panel A and to 2.63 p.p. in panel B. Appendix Table A.4 shows no significant net effects on the hazard of first, second, and fourth births.

In Table 5, we present results from estimating (2a) and (2b) inclusive of state-specific linear time trends and state \times mother’s age fixed effects separately by socioeconomic group. Columns (1)-(4) of panel A show that among caste groups, the greatest declines in the hazard of third birth are in SC and OBC families, of 4.10 p.p. and 2.88 p.p., respectively. This rep-

resents a 16.57 percent decline from the baseline pre-announcement hazard of third birth for SC households, and a 14.47 percent corresponding decline for OBC households. As described in Section 2, both these caste groups are legally recognized as socially and economically disadvantaged by central and state governments, and Panchayat positions are reserved for candidates from these castes in every state to ensure adequate political representation. As such, it is unsurprising that these households show the greatest response to the law at the margin, as they have the most to lose from reduced access to political power if they violate the fertility limits. In comparison, there is no significant decline in third births among the “general” or Upper castes (and for ST families). In a similar vein, columns (5)-(6) reveal a statistically significant decline in third births among poorer households (low SLI) of 2.14 p.p. (8.64 percent), whereas wealthier households with a high SLI score show no visible decline in the probability of having a third birth in column (6).

In the last four columns, we split the sample by husband’s and wife’s years of schooling (no schooling versus some schooling). While the coefficients in columns (7) and (8) of panel A are negative, they are both insignificant based on bootstrapped standard errors. On the other hand, in columns (9) and (10), the hazard of third birth significantly declines both in households where the father/ husband has no schooling and some schooling. In line with the pattern of results in the rest of panel A, the decline is larger if the father has no schooling (2.94 p.p. or 11.29 percent) than if he has some schooling (1.59 p.p. or 8.29 percent). The smaller decreases in the hazard of third birth for relatively higher socioeconomic status families is also due to the fact that these families have a lower baseline probability of third birth even in the absence of the limits, and hence are “less treatable” than lower socioeconomic status households. For instance, the baseline hazard of third birth is only 11 percent for high SLI mothers as compared to a 25 percent probability for low SLI mothers.

The estimates in panel B of Table 5 are slightly larger in magnitude and more precisely estimated than in Panel A, but the overall pattern of heterogeneity is similar. The hazard of third birth declines significantly by 4.67 p.p. (16.50 percent) for SCs, by 2.52 p.p. (10.20 percent) for OBCs, and by 2.96 p.p. (11 percent) for low SLI families. The effects are also significant in the last four columns, with the coefficients being larger in magnitude in families where the father has no schooling.

5.2 Effects on the total number of children

The results thus far have focused on the marginal effect of the limits on an additional birth conditional on a woman already having a certain number of children. However, they do not tell us the extent of substitution from, say, having four children to having only two children. In order to evaluate the overall impact of the laws on total “stock” of fertility, we re-estimate specifications (2a) and (2b) using indicators for whether a woman reports having one, two, three, four, and five living children in a given year as the outcome variables. Unlike the hazard analysis, in these regressions we do not impose any restrictions in terms of the prior number of children and use all available years for each woman. If the two-child limits are effective, we expect the likelihood of having two children to increase and the likelihood of more than two children to decrease after the laws have been announced in treatment states relative to control states and relative to pre-treatment years. These regressions capture the marginal effects on couples who had begun childbearing before the laws were announced as well as the behavioral response of new parents who started having children in the post-announcement years.

Table 6 presents the estimated effects of the fertility limits on these outcomes. In panel A, we also show results from specification (2a) estimated only for the sample of treatment states. The coefficients of $Treat_{st}$ in panel A imply that after the limits were announced, the probability of having two living children in a given year increased significantly by 0.75 p.p. or 3.41 percent in treatment states. The rest of the columns in panel A indicate that this increase in the likelihood of two children is a result of substitution from higher fertility levels—the probability of having three, four, or five living children declined respectively by 0.42 p.p. (2.48 percent), 0.47 p.p. (5.62 percent), and 0.28 p.p. (8.70 percent). There is no significant impact on the likelihood of one child. Panels B and C in Table 6 show a similar pattern of results, although the coefficients are not always as significant as those in panel A.

Since our dataset is a retrospective unbalanced panel, one concern is that the women in post-announcement years may be systematically younger than those in the pre-announcement period. Although our specifications control for age fixed effects as well as for state \times age fixed effects, in order to test the robustness of our findings we re-estimate the prior set of regressions for women who were no older than age 33 in a given year. This age restriction reduces sample selection due to changes in mother cohort composition across NFHS rounds. As Figure A.1 displays, the average age of mothers evolves smoothly over time across birth

parities with the age restriction. Appendix Table A.5 shows that the estimated effects on the number of children remain similar to those in Table 6 despite the age restriction.

Next we examine heterogeneity in the effects on the total number of children by socioeconomic characteristics of the parents using specification (2a). To avoid issues related to a large number of subgroup estimates for five different outcomes, for our heterogeneity results we use an indicator for more than two living children as the outcome variable. Panel A in Table 7 focuses only on treatment states whereas panel B shows the estimates for all states. In panel A, the probability of having more than 2 children declines significantly by 1.33 p.p. or 4.42 percent in treatment states, although the coefficient is insignificant in panel B. Although both high and low SLI families exhibit a significant decline, the magnitude of the effect is larger for low SLI households (1.9 p.p. versus 1 p.p. for high SLI). Similarly, couples where the wife has no schooling display a slightly larger decline (1.6 p.p.) than couples where the wife has some schooling (1.1 p.p.), although both are significant. In terms of husband's schooling, the coefficients are quite similar for both groups, but only significant if the husband has some schooling. In terms of caste affiliation, the only significant coefficient is for STs. On the whole, Table 7 is roughly consistent with the differential effects on hazard rates across socioeconomic groups that we described in the previous sub-section.

5.3 Effects on the sex ratios at birth

Next we examine the heterogeneous effects on the sex ratio of second and higher parity births by household caste using specifications (3a) and (3b). We restrict the sample to women whose first child was born before announcement of the limits and control for the gender of this child, as the firstborn child's gender is known to be exogenous and is a strong predictor of future fertility and sex-selective behavior.³³ These results are displayed in Table 8. Columns (1)-(5) correspond respectively to the entire sample, and sub-samples of SCs, STs, OBCs, and upper castes. Panel A focuses on the sample of treatment states and shows that the limits have no significant impact on the sex ratio of second and higher parity birth in the total sample or in the SC, ST, and upper caste samples. However, in column (4) there is a large and highly significant 5.28 p.p. or a 10.32 percent increase in the sex ratio for the OBC sample over a baseline probability of 51.17 percent. We find the same pattern of results in

³³In Appendix Table A.6, we verify that the fertility limits did not affect the sex ratio of first births.

panels B and C with OBC households exhibiting a significant 5.57 p.p. (10.74 percent) to a 5.70 percentage point (11.00 percent) increase in the probability that a birth is male.

The castes that are included in the OBC category vary across states; for instance, the *Jat* caste group is included in the OBC category in Rajasthan, but not in Haryana. Although OBC status is meant to improve the socioeconomic situation of historically backward classes, quite often politically dominant castes are able to lobby for OBC status in order to benefit from caste-based quotas in public employment and education. This often leads to socioeconomically privileged castes having OBC status in particular states, such as the influential landowning *Reddy* and *Kamma* castes in Andhra Pradesh (Deshpande and Ramachandran (2013)).³⁴ Hence it is not too surprising that OBC households respond to the law with greater sex-selection, as sex-selective behavior has been shown to be concentrated in non-SC and non-ST households with higher socioeconomic status (e.g., see Bhalotra and Cochrane (2010)). Further, OBCs constitute significant fractions of the populations in our treatment states, such as Haryana (28.1 percent), Rajasthan (47.5 percent), Madhya Pradesh (41.2 percent), and Maharashtra (27.1 percent)³⁵, that have highly adverse sex ratios of respectively 861, 922, 920, and 922 females per 1000 males in the 2001 Census of India.

6 Discussion

Given that average baseline terminal fertility in the treatment states is 2.8, the two-child limits impose a binding constraint on the fertility of a large fraction of individuals who wish to maintain electoral eligibility in these states. According to our sample, nearly 30 percent of couples where the wife is 15-49 years old had more than two children at baseline in treatment states. We find that the probability of having more than 2 children declines significantly by 1.33 p.p. in treatment states, i.e., 1.33 percent of couples where the wife is 15-49 years old decreased their fertility due to the limits. According to the 2001 Census of India, the number of married rural women in the 15-49 age group in treatment states is 46,269,920.³⁶ Thus, 615,390 (which is 1.33 percent of 46,269,920) couples responded to the limits. If we treat

³⁴In recent years, the *Jat* community in Haryana and the *Patels* in Gujarat have been proactively seeking OBC status.

³⁵This information is from the 64th round of the National Sample Survey of India (2007-08).

³⁶We use the number of married women in the 15-49 age group in treatment states (= 71,184,492) and multiply it with 0.65 (the share of rural women in our sample) to get the corresponding number for the rural sector, i.e., 46,269,920.

husbands and wives as individuals with separate political ambitions, the number doubles, i.e., 1,230,780 individuals responded to the limits.

We can do the same calculation using our estimated effects on the hazard of third birth. We find that the probability of having a third child declines by 2.06 to 2.65 p.p. in treatment states, i.e., 2.06 to 2.65 percent of couples where the wife is 15-49 years old and who had two children in treatment states decreased their marginal fertility due to the limits. According to the 2001 Census of India, the number of married rural women in the 15-49 age group in treatment states who had two children is 10,562,944.³⁷ Thus, 217,597 to 279,918 (which is 2.06 to 2.65 percent of 10,562,944) couples responded to the limits. If we treat husbands and wives as individuals with separate political ambitions, the numbers double, i.e., 435,194 to 559,836 individuals gave up a third child due to the limits. The total number of couples who adjusted fertility (in the previous paragraph) is higher since it also includes couples who started childbearing after the limits were announced and those who had one child at announcement and stopped after having the second child.

We can also translate these numbers into response per Panchayat seat. Suppose that the fertility response is entirely driven by political aspirations (and not by the role-model or anticipation channels). Our treatment states had 912,597 seats across all three tiers of the Panchayat system in 2004. The 1,230,780 figure thus translates into a response rate of 1.35 individuals per seat. The Association for Democratic Reforms reports an average of 2.43 candidates per Panchayat seat. This implies that slightly more than half of all potential contestants per seat in each election cycle in treatment states altered their fertility. This is a lower bound. In reality, not all incumbents would get disqualified by these limits. According to survey data from South India (Besley et al. (2003)), about 35 percent of village council heads or members had more than two children in 2002. Assuming the same fertility distribution for enacting states and assuming that all incumbents who have two or less children would be able to get re-elected, 1,230,780 individuals adjusted fertility for 593,188 seats that became “vacant” after the fertility limits, which is 2.07 people per seat or 85 percent of contestants per seat. Even if we assume that only one of the spouses runs for office, 28 to 43 percent of contestants per seat adjusted fertility due to these laws.

³⁷We use the number of married women in the 15-49 age group with two children in treatment states (= 16,250,683) and multiply it with 0.65 (the share of rural women in our sample) to get the corresponding number for the rural sector, i.e., 10,562,944.

Although we cannot credibly distinguish between the relative importance of various channels, these impacts are large, and consistent with the high participation of voters and candidates in Panchayat politics, making both the aspirations and role-model channels for our estimated effects plausible. Voter turnout in Panchayat elections routinely exceeds 70 percent. In the 2014 World Values Survey, 53 percent of the respondents (69 percent among the “lower class”) say that politics is “very important” or “rather important” in their life and about 48 percent of the respondents are members of a political party.³⁸ Thus our results have significant implications for the understanding of the relationship between political participation and social change in low-income democracies. Our paper is extremely policy-relevant as several Indian states have also recently instituted minimum educational qualifications for Panchayat candidates in a bid to improve educational attainment of the constituents and the quality of elected officials.

7 Conclusion

We find that the two-child limits on candidates in Panchayat elections decrease fertility among constituents, but also lead to an unintended increase in the already male-biased sex ratio in certain socioeconomic groups. These effects are caused by constituents’ political ambitions or anticipation of stricter restrictions in other non-political arenas rather than the role-model influence of their leaders. Political aspirations may not only reflect the desire to effect positive social change, but could also be driven by rent-seeking behavior. The potential income from political rents and corrupt practices may be a strong incentive for becoming an officeholder in low-income countries. While we cannot separately identify these “altruistic” and “selfish” components of political aspirations, we show that these ambitions may be substantial and represent a previously ignored channel of demographic change.

Our findings reiterate that population control measures that ignore son preference can worsen the sex ratio at birth. Similar limits have been proposed for members of state legislative assemblies and the national parliament in India. If incentives for local leadership are stronger than state or national leadership ambitions, the proposed limits may be less

³⁸About 72 percent say that a democratic political system is a “very good” or “fairly good” way of governing the country. According to the 2005 India Human Development Survey, in 28 percent of households a member attended a public meeting called by the local council in the last year and in 10 percent of households someone from or close to the household is a member of the local council.

effective than the laws we examine.³⁹

Fertility restrictions on elected members also have implications for political representation of various socioeconomic groups. The two-child limits impose a more severe constraint on couples with weaker access to contraception or higher demand for children, increasing their risk of disqualification and reducing their political representation. Given that we do not find a significant long-run fertility response among families in the bottom third of the wealth distribution, the limits may be more likely to diminish their political representation. Since a large proportion of the poor belong to lower castes, the limits could also impede the progress made by caste-based affirmative action if only the “creamy-layer” of the lower-castes are able to meet the eligibility criteria. The limits could also undermine gender-based quotas as aspiring female leaders may not have autonomy over their fertility due to intra-household gender disparities. Indeed, women comprise the overwhelming majority of individuals in Table A.1 that were disqualified for violating the limits.

Recently, some Indian states have enacted similar restrictions to meet policy goals in the areas of education and sanitation. As of 2014, individuals are barred from Panchayat membership in Rajasthan if they have less than primary schooling or do not have a functional toilet in their home.⁴⁰ Although 50% of the Panchayat seats in Rajasthan are reserved for women, the female literacy rate is only 45.8% (2011 Census of India).⁴¹ Moreover, lower castes face considerable discrimination in access to sanitation and education. Our results show that these new restrictions are likely to stifle political representation of disadvantaged groups which may cause social conflict. The effects of such barriers to local leadership on political representation, discrimination, and aspirations are key to poverty reduction, and merit further investigation.

³⁹Genicot and Ray (2014) formalize a related idea as follows: “...the “best” aspirations are those that lie at a moderate distance from the individual’s current economic situation standards, large enough to incentivize but not so large as to induce frustration.”

⁴⁰The minimum schooling requirements for block and district councils are eight and ten years, respectively.

⁴¹For tribal women, the literacy rate is even lower (25.22%).

References

- ACEMOGLU, D. AND J. A. ROBINSON (2000): “Why Did the West Extend the Franchise? Democracy, Inequality, and Growth in Historical Perspective,” *Quarterly Journal of Economics*, 1167–1199.
- (2006): *Economic Origins of Dictatorship and Democracy*, Cambridge Univ Press.
- AFRIDI, F., V. IVERSEN, AND M. SHARAN (2014): “Women Political Leaders, Corruption and Learning: Evidence from a large Public Program in India,” *IGC Working Paper*.
- ANUKRITI, S. (2014): “The Fertility-Sex Ratio Trade-off: Unintended Consequences of Financial Incentives,” *IZA Discussion Paper No. 8044*.
- ANUKRITI, S., S. BHALOTRA, AND H. TAM (2016): “Missing Girls: Ultrasound Access and Excess Female Mortality,” *Working Paper*.
- APPADURAI, A. (2004): “The Capacity to Aspire: Culture and the Terms of Recognition,” in *Culture and Public Action*, ed. by V. Rao and M. Walton, Stanford University Press.
- ASHRAF, Q. H., D. N. WEIL, AND J. WILDE (2013): “The Effect of Fertility Reduction on Economic Growth,” *Population and development review*, 39, 97–130.
- BANERJEE, A., S. KUMAR, R. PANDE, AND F. SU (2011): “Do Informed Voters make Better Choices? Experimental Evidence from Urban India,” *Working Paper*.
- BARDHAN, P. AND D. MOOKHERJEE (2000): “Capture and Governance at Local and National Levels,” *American Economic Review: Papers and Proceedings*, 90, 135–139.
- BASSI, V. AND I. RASUL (2014): “Persuasion: A Case Study of Papal Influence on Fertility Preferences and Behavior,” *Working Paper*.
- BEAMAN, L., E. DUFLO, R. PANDE, AND P. TOPALOVA (2012): “Female Leadership Raises Aspirations and Educational Attainment for Girls: A Policy Experiment in India,” *Science*, 335, 581–586.
- BERNARD, T., S. DERCON, AND A. T. TAFFESSE (2012): “Beyond Fatalism: An Empirical Exploration of Self-Efficacy and Aspirations Failure in Ethiopia,” *IFPRI Discussion Paper 01101*.
- BESLEY, T. AND S. COATE (1997): “An Economic Model of Representative Democracy,” *Quarterly Journal of Economics*, 112, 85–114.
- BESLEY, T., R. PANDE, L. RAHMAN, AND V. RAO (2003): “The Politics of Public Good Provision: Evidence from Indian Local Governments,” *Journal of the European Economic Association*, 2, 416–426.
- BESLEY, T. AND M. REYNAL-QUEROL (2011): “Do Democracies Select More Educated Leaders,” *American Political Science Review*, 105, 552–566.

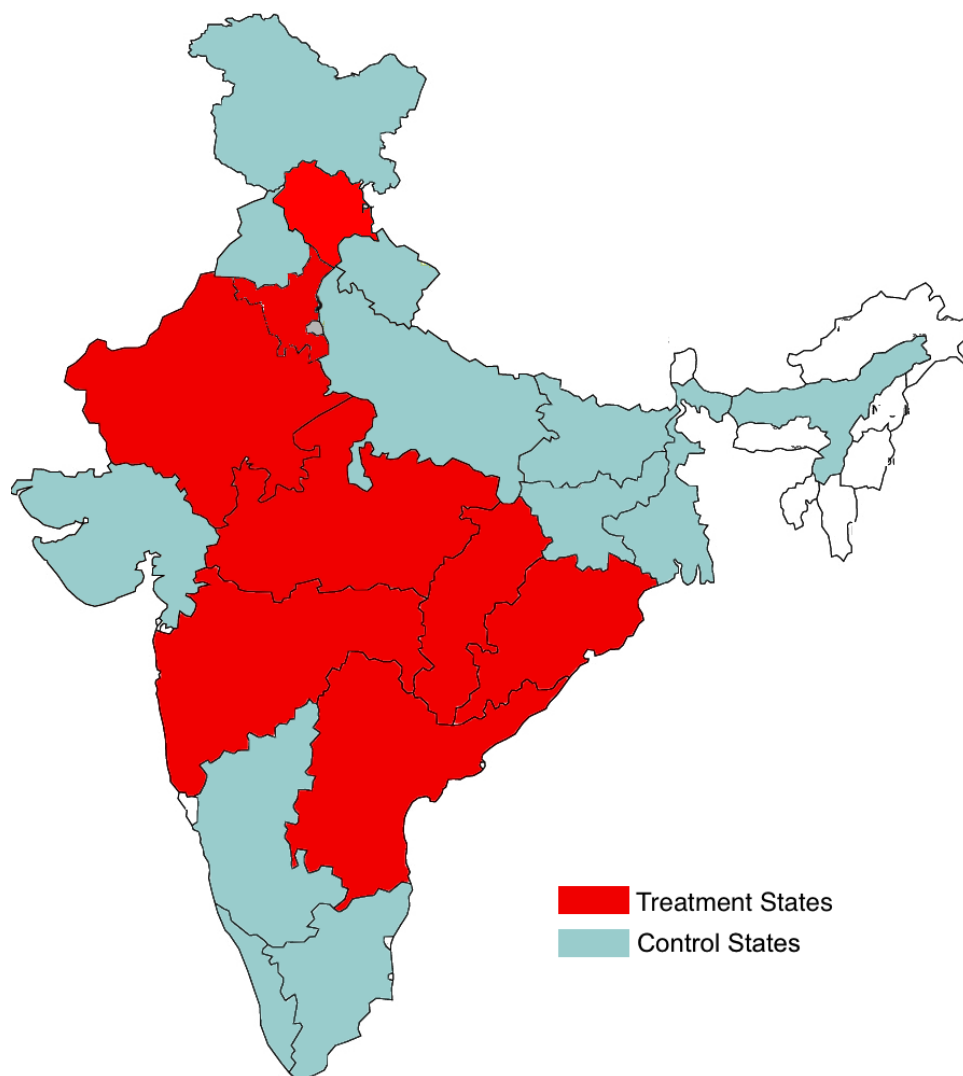
- BETTINGER, E. P. AND B. T. LONG (2005): “Do Faculty Serve as Role Models? The Impact of Instructor Gender on Female Students,” *American Economic Review*, 95, 152–7.
- BHALOTRA, S., I. CLOTS-FIGUERAS, AND L. IYER (2013): “Path-Breakers: How Does Women’s Political Participation Respond to Electoral Success?” *HBS Working Paper 14-035*.
- BHALOTRA, S. AND T. COCHRANE (2010): “Where Have All the Young Girls Gone? Identification of Sex Selection in India,” *IZA Discussion Paper No. 5381*.
- BIDNER, C. AND P. FRANCOIS (2013): “The Emergence of Political Accountability,” *The Quarterly Journal of Economics*, 128, 1397–1448.
- BROWN, D. S. AND A. M. MOBARAK (2009): “The Transforming Power of Democracy: Regime Type and the Distribution of Electricity,” *American Political Science Review*, 103, 193–213.
- BUCH, N. (2005): “Law of Two-Child Norm in Panchayats: Implications, Consequences and Experiences,” *Economic and Political Weekly*, XL.
- (2006): *The Law of Two Child Norm in Panchayats*, Concept Publishing Company.
- BUSO, M., J. GREGORY, AND P. KLINE (2013): “Assessing the Incidence and Efficiency of a Prominent Place Based Policy,” *American Economic Review*, 103, 897–947.
- CAMERON, A. C., J. B. GELBACH, AND D. L. MILLER (2008): “Bootstrap-Based Improvements for Inference with Clustered Errors,” *Review of Economics and Statistics*, 90, 414–27.
- CANNING, D. AND T. P. SCHULTZ (2012): “The Economic Consequences of Reproductive Health and Family Planning,” *The Lancet*, 380, 165–171.
- CHATTOPADHYAY, R. AND E. DUFLO (2004): “The Impact of Reservation in the Panchayati Raj: Evidence from a Nationwide Randomized Experiment,” *Economic and Political Weekly*, 39, 979–986.
- CHIAPPORI, P.-A., B. FORTIN, AND G. LACROIX (2002): “Marriage Market, Divorce Legislation, and Household Labor Supply,” *Journal of Political Economy*, 110, 37–72.
- CHONG, A., S. DURYEA, AND E. LA FERRARA (2012): “Soap Operas and Fertility: Evidence from Brazil,” *American Economic Journal: Applied Economics*, 4, 1–31.
- DALTON, P. S., S. GHOSAL, AND A. MANI (2014): “Poverty and Aspirations Failure,” *The Economic Journal*, Forthcoming.
- DAS GUPTA, M. AND P. BHAT (1997): “Fertility Decline and Increased Manifestation of Sex Bias in India,” *Population Studies*, 51, 307–315.

- DESHPANDE, A. AND R. RAMACHANDRAN (2013): “How Backward are the Other Backward Classes? Changing Contours of Caste Disadvantage in India,” *Center for Development Economics Working Papers*.
- DOEPKE, M., M. TERTILT, AND A. VOENA (2012): “The Economics and Politics of Women’s Rights,” *Annual Review of Economics*, 4.
- EBENSTEIN, A. (2010): “The “Missing” Girls of China and the Unintended Consequences of the One Child Policy,” *Journal of Human Resources*, 45, 87–115.
- FERRAZ, C. AND F. FINAN (2008): “Exposing Corrupt Politicians: The Effects of Brazil’s Publicly Released Audits on Electoral Outcomes,” *Quarterly Journal of Economics*, 123, 703–745.
- FISMAN, R., F. SCHULZ, AND V. VIG (2014): “The Private Returns to Public Office,” *Journal of Political Economy*, 122.
- FOSTER, A. D. AND M. R. ROSENZWEIG (2004): “Democratization and the Distribution of Local Public Goods in a Poor Rural Economy,” *Brown University Working Paper*.
- GENICOT, G. AND D. RAY (2014): “Aspirations and Inequality,” *NBER Working Paper 19976*.
- GRADSTEIN, M. (2007): “Inequality, Democracy and the Protection of Property Rights,” *The Economic Journal*, 117, 252–269.
- JAYACHANDRAN, S. (2014): “Fertility Decline and Missing Women,” *NBER Working Paper 20272*.
- JENSEN, R. AND E. OSTER (2009): “The Power of TV: Cable Television and Women’s Status in India,” *Quarterly Journal of Economics*, 124, 1057–94.
- JOSHI, S. AND T. P. SCHULTZ (2007): “Family Planning as an Investment in Development: Evaluation of a Program’s Consequences in Matlab, Bangladesh,” *Yale University Economic Growth Center Discussion Paper*.
- KAPOOR, M. AND S. RAVI (2014): “Why So Few Women in Politics? Evidence from India,” *Brookings Working Paper*.
- MILLER, G. (2010): “Contraception as Development? New Evidence from Family Planning in Colombia,” *The Economic Journal*, 120, 709–736.
- MOBARAK, A. M. (2005): “Democracy, Volatility, and Economic Development,” *Review of Economics and Statistics*, 87, 348–361.
- MOOKHERJEE, D., S. NAPEL, AND D. RAY (2010): “Aspirations, Segregation, and Occupational Choice,” *Journal of the European Economic Association*, 8, 139–168.
- MUELLER, D. C. AND T. STRATMANN (2003): “The Economic Effects of Democratic Participation,” *Journal of Public Economics*, 87, 2129–2155.

- OLINTO, P., K. BEEGLE, C. SOBRADO, AND H. UEMATSU (2013): “The State of the Poor: Where are the Poor, Where is Extreme Poverty Harder to End, and What is the Current Profile of the World’s Poor?” *Economic Premise*.
- OLIVETTI, C., E. PATACCINI, AND Y. ZENOU (2013): “Mothers, Friends and Gender Identity,” *NBER Working Paper 19610*.
- OSBORNE, M. J. AND A. SLIVINSKI (1996): “A Model of Political Competition with Citizen Candidates,” *Quarterly Journal of Economics*, 111, 65–96.
- PAPAIIOANNOU, E. AND G. SIOUROUTIS (2008): “Democratization and Growth,” *Economic Journal*, 118, 1520–1551.
- PÖRTNER, C. C. (2010): “Sex Selective Abortions, Fertility and Birth Spacing,” *University of Washington, Department of Economics, Working Paper UWEC-2010-4-R*.
- RAY, D. (2006): “Aspirations, Poverty, and Economic Change,” in *What Have We Learnt About Poverty*, ed. by A. Banerjee, R. Bénabou, and D. Mookherjee, Oxford University Press.
- RODRIK, D. (2000): “Institutions for High-Quality Growth: What They Are and How to Acquire Them,” *Studies in Comparative International Development*, 35, 3–31.
- ROSENBLUM, D. (2013): “The effect of fertility decisions on excess female mortality in India,” *Journal of Population Economics*, 26, 147–180.
- ROSENZWEIG, M. AND K. WOLPIN (1980): “Life-Cycle Labor Supply and Fertility: Causal Inferences from Household Models,” *Journal of Political Economy*, 88.
- ROSENZWEIG, M. R. AND J. ZHANG (2009): “Do Population Control Policies Induce More Human Capital Investment? Twins, Birth Weight and China’s “One-Child” policy,” *The Review of Economic Studies*, 76, 1149–1174.
- SASTRY, T. (2014): “Towards Decriminalisation of Election and Politics,” *Economic and Political Weekly*, XLIX, 34–41.
- VISARIA, L. (2005): “Female Deficit in India: Role of Prevention of Sex Selective Abortion Act,” *mimeo*.
- VISARIA, L., A. ACHARYA, AND F. RAJ (2006): “Two-Child Norm: Victimising the Vulnerable?” *Economic and Political Weekly*, XLI.

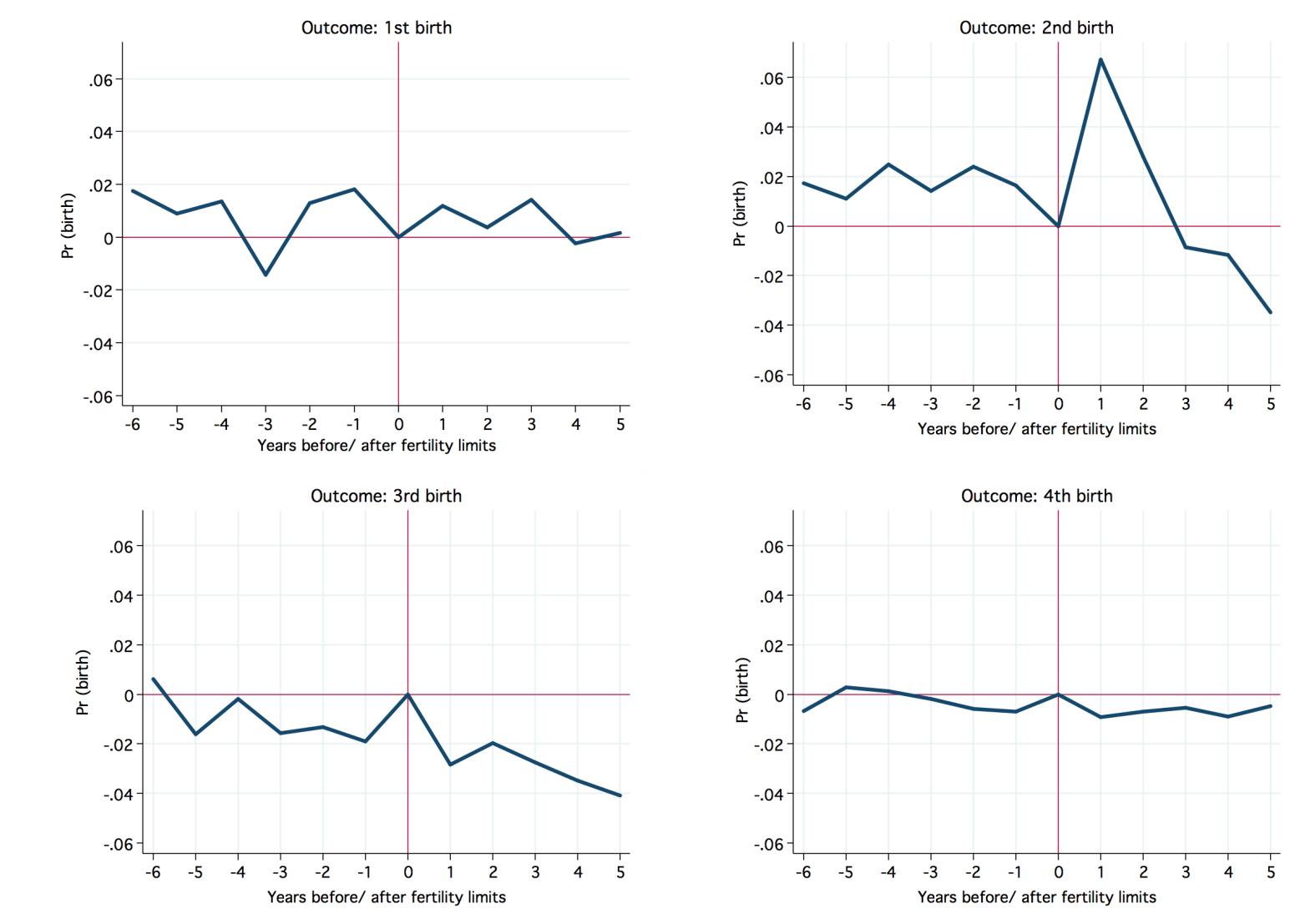
8 Figures and Tables

Figure 1: Treatment and control states



NOTES: Treatment states are those that have enacted a fertility limit at some point during our sample period. Control or never-treated states are those that did not.

Figure 2: Differential hazards of birth in treatment states relative to control states



NOTES: This figure plots the α_k coefficients from specification (1). The outcome variables are indicators for births of various orders. Each graph represents a different birth order. For the regression where birth of order b is the dependent variable, the sample is restricted to years after birth $(b - 1)$ and until birth b for mothers whose $(b - 1)^{th}$ child was born before the year of announcement in their respective states. The vertical line (at $k = 0$) indicates the year before announcement. The corresponding coefficient estimates are in Appendix Table A.3.

Table 1: Timeline for fertility limits across states

State	Announced	Grace Period	In effect	End
Rajasthan	Oct 1992	Apr 23, 1994 - Nov 27, 1995	Nov 27, 1995 -	
Orissa	Sep 1993/1994*	Apr 1994 - Apr 21, 1995	Apr 22, 1995 -	
Andhra Pradesh	Mar 1994	May 30, 1994 - May 30, 1995	Jun 1995 -	
Haryana	Apr 1994	Apr 21, 1994 - Apr 24, 1995	Apr 25, 1995 - Dec 31, 2004	Jul 21, 2006 (retro. impl. Jan 1, 2005)
Himachal Pradesh	Jan - Apr 2000	Apr 18, 2000 - Apr 18, 2001	Apr 2001 - Apr 2005	May 30, 2005
Madhya Pradesh	Jan - Mar 2000**	Mar 29, 2000 - Jan 26, 2001	Jan 2001 - Nov 2005	Nov 20, 2005
Chhattisgarh	2000	2000 - Jan 2001	Jan 2001- 2005	2005 (earliest mention) ⁴²
Maharashtra	2003***	Sep 21, 2002 - Sep 20, 2003	Sep 2003 -	
Uttarakhand (municipal only)	2002			
Gujarat	2005	Aug 2005 - Aug 11, 2006	Aug 11, 2006 -	
Bihar (municipal only)	Jan 2007	Feb 1, 2007 - Feb 1, 2008	Feb 1, 2008 -	

NOTES:

*For district councils in 1993 and for village and block councils in 1994.

**Notified on May 31, 2000. This created problems since people whose third child was born in Jan 2001 contested their disqualification for birth within 8 months of the new law.

***In retrospective effect from Sep 21, 2002.

Table 2: Summary statistics

Variable	Never treated		Treated			
	Mean	Std. Dev.	$Post_{st} = 0$		$Post_{st} = 1$	
			Mean	Std. Dev.	Mean	Std. Dev.
	(1)	(2)	(3)	(4)	(5)	(6)
Urban	0.368	0.482	0.350	0.477	0.387	0.487
Hindu	0.810	0.392	0.910	0.286	0.879	0.326
Muslim	0.166	0.372	0.072	0.258	0.080	0.272
Sikh	0.047	0.211	0.016	0.124	0.009	0.095
Christian	0.041	0.199	0.015	0.120	0.017	0.131
SC	0.165	0.371	0.153	0.360	0.177	0.382
ST	0.049	0.215	0.126	0.332	0.111	0.314
OBC	0.233	0.422	0.180	0.384	0.335	0.472
<i>Wife's years of schooling:</i>						
Zero	0.491	0.500	0.559	0.496	0.480	0.500
5-10 years	0.242	0.428	0.218	0.413	0.251	0.434
10-12 years	0.091	0.287	0.076	0.265	0.091	0.287
12-15 years	0.049	0.216	0.031	0.173	0.045	0.208
≥ 15 years	0.050	0.219	0.036	0.189	0.060	0.238
<i>Husband's years of schooling:</i>						
Zero	0.261	0.439	0.285	0.451	0.262	0.439
5-10 years	0.303	0.460	0.302	0.459	0.308	0.462
10-12 years	0.156	0.362	0.162	0.369	0.150	0.357
12-15 years	0.091	0.288	0.071	0.257	0.081	0.273
≥ 15 years	0.096	0.295	0.083	0.275	0.115	0.319
Low SLI	0.446	0.497	0.476	0.499	0.380	0.485
High SLI	0.227	0.419	0.209	0.407	0.244	0.430
Mother's age at birth	24.885	6.085	23.090	5.624	26.355	6.308
Birth = 1	0.214	0.410	0.234	0.423	0.152	0.359
1st birth is male	0.516	0.500	0.516	0.500	0.519	0.500
N	1,082,017		496,957		222,567	

NOTES: $Post_{st} = 1$ for years \geq announcement year of the law. SC, ST, and OBC indicate Scheduled Caste, Scheduled Tribe, and Other Backward Class women, respectively. Low and High SLI (standard of living index) are equal to one if the household belongs to the bottom-third or the top-third of household wealth distribution in India. This table includes both rural and urban sub-samples, however, the rest of the paper focuses on the rural sample. The summary statistics for the rural sub-sample are available upon request.

Table 3: Correlations between law announcements and socioeconomic variables

Dependent Variable ↓	Coefficient of $Treat_{st}$ (1)	Std. Error (2)
SC	-0.004	[0.008]
ST	0.009	[0.008]
OBC	-0.008	[0.010]
Upper caste	0.003	[0.011]
Hindu	0.012	[0.009]
Muslim	0.003	[0.006]
Sikh	0.001	[0.002]
Christian	0.001	[0.007]
Low SLI	0.009	[0.008]
Med SLI	-0.001	[0.006]
High SLI	-0.007	[0.005]
<i>Wife's years of schooling:</i>		
Zero	-0.005	[0.007]
5-10 years	0.009	[0.010]
10-12 years	0.002	[0.002]
12-15 years	0.001	[0.004]
≥ 15 years	-0.002	[0.002]
<i>Husband's years of schooling:</i>		
Zero	0.003	[0.008]
5-10 years	-0.002	[0.008]
10-12 years	-0.001	[0.003]
12-15 years	0.002	[0.005]
≥ 15 years	-0.000	[0.003]
N	1,143,057	

NOTES: Each coefficient is from a separate regression that includes state, year, and state x mother's age fixed effects, and state-specific linear time trends. Standard errors are in brackets and are clustered by state. SC, ST, and OBC indicate Scheduled Caste, Scheduled Tribe, and Other Backward Class households, respectively. Low, Med, and High SLI (standard of living index) are equal to one if the household belongs to the bottom-third, middle-third, or the top-third of household wealth distribution in India. *** 1%, ** 5%, * 10%.

Table 4: Net effect on the hazard of third birth

3rd birth = 1	(1)	(2)	(3)
Panel A:			
$Treat_{st}$	-0.0143 [0.0096] (0.0095)	-0.0213 [0.0088]** (0.0098)**	-0.0206 [0.0078]** (0.0093)**
Baseline mean		0.2131	
Panel B:			
$T_s * Post_{st}$	-0.0196 [0.0114] (0.0117)	-0.0265 [0.0117]** (0.0131)*	-0.0263 [0.0103]** (0.0120)**
$Post_{st}$	0.0099 [0.0101]	0.0077 [0.0090]	0.0083 [0.0082]
Baseline mean		0.2431	
N		182,082	
State FE	x	x	x
Year FE	x	x	x
Years since 2nd birth FE	x	x	x
X_{it}	x	x	x
Linear state trends		x	x
State x Age FE			x

NOTES: This table reports the coefficients of $Treat_{st}$ from specifications (2a) and (2b). The outcome variable is an indicator for third birth. The sample is restricted to years after second birth and until third birth for mothers whose second child was born before the year of announcement. Standard errors in brackets are clustered by state and in parentheses are wild-cluster bootstrapped by state. The baseline mean is calculated for observations where $Treat_{st} = 0$ in panel A and for observations where $Post_{st} = 0$ in panel B. *** 1%, ** 5%, * 10%.

Table 5: Heterogeneity in the net effect on the hazard of third birth

3rd birth = 1	SC	ST	OBC	Upper	Low SLI	High SLI	Wife has schooling	Wife has no schooling	Husband has schooling	Husband has no schooling
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Panel A:										
$Treat_{st}$	-0.0410 [0.0126] *** (0.0168) ***	0.0029 [0.0215] (0.0207)	-0.0288 [0.0104] ** (0.0142) *	-0.0098 [0.0124] (0.0119)	-0.0214 [0.0100] ** (0.0098) *	-0.0024 [0.0123] (0.0119)	-0.0160 [0.0103] (0.0095)	-0.0182 [0.0097] * (0.0103)	-0.0159 [0.0077] * (0.0083) *	-0.0294 [0.0119] ** (0.0146) **
Baseline mean	0.2475	0.2521	0.1990	0.2015	0.2478	0.1143	0.1468	0.2629	0.1917	0.2604
Panel B:										
$T_s * Post_{st}$	-0.0467 [0.0157] *** (0.0180) ***	-0.0094 [0.0314] (0.0300)	-0.0252 [0.0117] ** (0.0132) *	-0.0182 [0.0162] (0.0162)	-0.0296 [0.0123] ** (0.0133) **	-0.0062 [0.0131] (0.0127)	-0.0227 [0.0114] * (0.0119) *	-0.0246 [0.0119] * (0.0126) *	-0.0221 [0.0099] ** (0.0107) **	-0.0336 [0.0171] * (0.0178) *
$Post_{st}$	0.0083 [0.0110]	0.0171 [0.0286]	-0.0046 [0.0114]	0.0125 [0.0122]	0.0119 [0.0113]	0.0054 [0.0100]	0.0097 [0.0073]	0.0090 [0.0114]	0.0090 [0.0076]	0.0058 [0.0147]
Baseline mean	0.2831	0.2690	0.2470	0.2257	0.2669	0.1543	0.1830	0.2777	0.2257	0.2769
N	28,074	17,868	38,288	97,852	110,159	17,200	72,165	109,917	122,979	59,103

NOTES: This table reports the coefficients from specifications (2a) and (2b). The outcome variable is an indicator for third birth. The sample is restricted to years after second birth and until third birth for mothers whose second child was born before the year of announcement. Standard errors in brackets are clustered by state and in parentheses are wild-cluster bootstrapped by state. SC, ST, OBC, and Upper indicate Scheduled Caste, Scheduled Tribe, Other Backward Class, and upper caste households, respectively. Low and High SLI (standard of living index) are equal to one if the household belongs to the bottom-third or the top-third of household wealth distribution in India. The last four columns split the sample into wives and husbands who have zero and non-zero years of schooling. The baseline mean is calculated for observations where $Treat_{st} = 0$ in panel A and for observations where $Post_{st} = 0$ in panel B. *** 1%, ** 5%, * 10%.

Table 6: Effects on the number of living children

	Kids = 1	Kids = 2	Kids = 3	Kids = 4	Kids = 5
	(1)	(2)	(3)	(4)	(5)
Panel A:					
	Only treatment states				
$Treat_{st}$	0.0066	0.0075	-0.0042	-0.0047	-0.0028
	[0.0039]	[0.0035]*	[0.0021]*	[0.0025]	[0.0013]*
	(0.0046)	(0.0042)*	(0.0023)*	(0.0030)	(0.0017)*
N			459,293		
Baseline mean	0.2394	0.2199	0.1693	0.0836	0.0322
Panel B:					
$Treat_{st}$	0.0008	0.0090	-0.0018	-0.0052	-0.0024
	[0.0055]	[0.0068]	[0.0055]	[0.0026]*	[0.0020]
	(0.0054)	(0.0065)	(0.0053)	(0.0030)*	(0.0053)
N			1,143,057		
Baseline mean	0.2351	0.2351	0.1711	0.0878	0.0379
Panel C:					
$T_s * Post_{st}$	0.0001	0.0139	-0.0009	-0.0080	-0.0014
	[0.0063]	[0.0100]	[0.0073]	[0.0037]**	[0.0026]
	(0.0062)	(0.0104)	(0.0072)	(0.0042)*	(0.0026)
$Post_{st}$	0.0009	-0.0067	-0.0013	0.0040	-0.0014
	[0.0043]	[0.0062]	[0.0037]	[0.0025]	[0.0026]
N			1,143,057		
Baseline mean	0.2425	0.2220	0.1650	0.0855	0.0364

NOTES: This table presents the regression estimates corresponding to specification (2a) in panels A and B, and to specification (2b) in panel C using indicators for 1/ 2/ 3/ 4/ 5 living children as the outcome variables. No sample restrictions are imposed, except that the sample is limited to treatment states in panel A. Each column within a panel is a different regression. Standard errors in brackets are clustered by state and in parentheses are wild-cluster bootstrapped by state. The baseline mean is calculated for observations where $Treat_{st} = 0$ in panels A and B, and for observations where $Post_{st} = 0$ in panel C. *** 1%, ** 5%, * 10%.

Table 7: Heterogeneous effects on the likelihood of ≥ 3 living children

	All	SC	ST	OBC	Upper	Low SLI	High SLI	Wife has schooling	Wife has no schooling	Husband has schooling	Husband has no schooling
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
A. Treatment states only											
$Treat_{st}$	-0.0133 [0.0043]** (0.0072)**	0.0045 [0.0126] (0.0119)	-0.0402 [0.0185]* (0.0276)**	-0.0068 [0.0062] (0.0062)	-0.0168 [0.0118] (0.0111)	-0.0190 [0.0088]* (0.0135)	-0.0109 [0.0039]** (0.0062)*	-0.0110 [0.0037]** (0.0057)**	-0.0155 [0.0078]* (0.0115)	-0.0136 [0.0033]*** (0.0063)***	-0.0131 [0.0097] (0.0125)
N	459,293	78,174	77,278	105,475	198,366	291,535	33,708	149,776	309,517	292,311	166,982
Baseline mean	0.3007	0.3124	0.3013	0.2860	0.3023	0.3064	0.2561	0.2674	0.3152	0.2938	0.3124
B. All states											
$Treat_{st}$	-0.0087 [0.0065] (0.0069)	0.00002 [0.0108] (0.0104)	-0.0206 [0.0147] (0.0194)	-0.0097 [0.0068] (0.0072)	-0.0141 [0.0129] (0.0133)	-0.0100 [0.0064]	-0.0111 [0.0062]*	-0.0121 [0.0069]	-0.0081 [0.0063]	-0.0110 [0.0070]	-0.0049 [0.0067]
N	1,143,057	202,619	123,071	267,024	550,343	722,793	90,528	416,265	726,792	747,865	395,192
Baseline mean	0.3189	0.3399	0.3180	0.3129	0.3144	0.3317	0.2498	0.2592	0.3527	0.3027	0.3495

NOTES: This table presents the regression estimates corresponding to specification (2a) using an indicator for ≥ 3 living children in a given year as the outcome variable. No sample restrictions are imposed, except that the sample is limited to treatment states in panel A. Each coefficient is from a different regression. Standard errors in brackets are clustered by state and in parentheses are wild-cluster bootstrapped by state. SC, ST, OBC, and Upper indicate Scheduled Caste, Scheduled Tribe, Other Backward Class, and upper caste households, respectively. Low and High SLI (standard of living index) are equal to one if the household belongs to the bottom-third or the top-third of household wealth distribution in India. The last four columns split the sample into wives and husbands who have zero and non-zero years of schooling. The baseline mean is calculated for observations where $Treat_{st} = 0$. *** 1%, ** 5%, * 10%

Table 8: Sex ratio of second and higher parity births

Male = 1	All	SC	ST	OBC	Upper
	(1)	(2)	(3)	(4)	(5)
Panel A:					
	Only treatment states				
$Treat_{st}$	0.0086	-0.0265	0.0071	0.0528	-0.0048
	[0.0103]	[0.0419]	[0.0082]	[0.0111]***	[0.0210]
	(0.0111)	(0.0368)	(0.0066)	(0.0252)**	(0.0196)
N	61,490	11,054	11,627	12,677	26,132
Baseline mean	0.5211	0.5235	0.5142	0.5117	0.5267
Panel B:					
$Treat_{st}$	0.0109	-0.0249	0.0071	0.0557	0.0061
	[0.0060]*	[0.0232]	[0.0157]	[0.0119]***	[0.0148]
	(0.0078)	(0.0218)	(0.0154)	(0.0221)**	(0.0143)
N	165,016	31,169	18,757	35,858	79,232
Baseline mean	0.5186	0.5215	0.5177	0.5185	0.5178
Panel C:					
$T_s * Post_{st}$	0.0088	-0.0231	-0.0137	0.0570	0.0022
	[0.0078]	[0.0208]	[0.0241]	[0.0145]***	[0.0168]
	(0.0077)	(0.0201)	(0.0241)	(0.0202)***	(0.0164)
$Post_{st}$	0.0032	-0.0026	0.0288	-0.0017	0.0060
	[0.0082]	[0.0144]	[0.0272]	[0.0100]	[0.0123]
N	165,016	31,169	18,757	35,858	79,232
Baseline mean	0.5181	0.5214	0.5174	0.5184	0.5170

NOTES: This table reports the coefficients from specifications (3a) and (3b). The sample is restricted to second and higher order births to women whose first child was born before the law was announced in her state. Standard errors in brackets are clustered by state and in parentheses are wild-cluster bootstrapped by state. SC, ST, OBC, and Upper indicate Scheduled Caste, Scheduled Tribe, Other Backward Class, and upper caste households, respectively. In panels A and B, the baseline mean is calculated for observations where $Treat_{st} = 0$ and in panel C for observations where $Post_{st} = 0$. *** 1%, ** 5%, * 10%.

A Appendix Figures and Tables

Table A.1: Panchayat members disqualified during 2000-04, for selected states

State	Number of disqualifications (excluding rejected nominations)
Haryana	1,350
Rajasthan	548
Madhya Pradesh	1,140
Chhattisgarh	766
Andhra Pradesh	94*

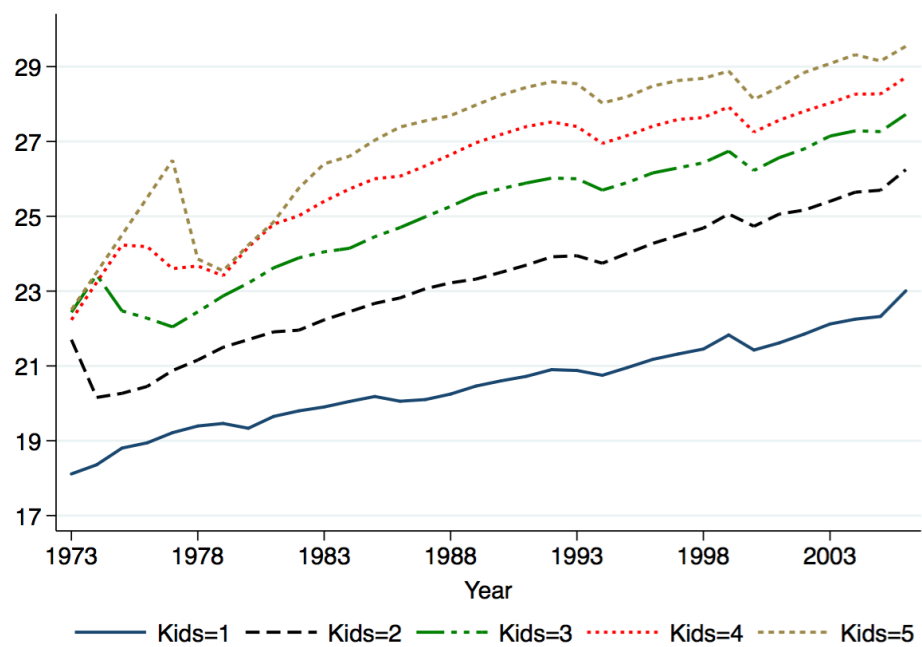
NOTES: *Data available for 15 out of 23 districts. Source: [Buch \(2005\)](#) and [Visaria et al. \(2006\)](#).

Table A.2: Timeline of Panchayat elections

State	Election Years	
	Without the limits	With the limits
Rajasthan	1995	2000, 2005, 2010, 2015
Haryana	1994, 2010	2000, 2005
Andhra Pradesh		1995, 2001, 2006, 2011
Orissa		1997, 2002, 2007, 2012
Himachal Pradesh	1995, 2005, 2010-11	2000
Madhya Pradesh	1994, 2010	2000*, 2005
Chhattisgarh	2010	2000, 2005
Maharashtra	1995, 2000	2007, 2010, 2013
Uttarakhand	2003, 2008, 2014	
Jharkhand	2010	
Gujarat	2001, 2005-06	2010-11
Bihar	2006	2011

NOTES: *Although the fertility limits were officially introduced after the Panchayat elections were over in 2000, the new government began disqualifying elected representatives earlier ([Visaria et al. \(2006\)](#)).

Figure A.1: Average mother's age at birth, by year and number of living children



NOTES: This graph plots the trends in average age of mothers who have a given number of children in a year. The sample is restricted to mothers \leq age 33 in a given year.

Table A.3: Effects on hazards of birth

Coefficients of $T_s * Post_{s,t+k}$	Outcome: Birth = 1			
	1st (1)	2nd (2)	3rd (3)	4th (4)
$t - 6$	0.0174 [0.0249]	0.0173 [0.0259]	0.0062 [0.0226]	-0.0068 [0.0274]
$t - 5$	0.0088 [0.0204]	0.0111 [0.0189]	-0.0161 [0.0143]	0.0029 [0.0189]
$t - 4$	0.0136 [0.0177]	0.0247 [0.0197]	-0.0019 [0.0140]	0.0013 [0.0141]
$t - 3$	-0.0144 [0.0225]	0.0141 [0.0184]	-0.0156 [0.0142]	-0.0018 [0.0115]
$t - 2$	0.0129 [0.0185]	0.0238 [0.0168]	-0.0132 [0.0137]	-0.0059 [0.0116]
$t - 1$	0.0182 [0.0272]	0.0163 [0.0245]	-0.019 [0.0201]	-0.0069 [0.0157]
t	0	0	0	0
$t + 1$	0.0119 [0.0220]	0.0670** [0.0268]	-0.0283 [0.0218]	-0.0093 [0.0286]
$t + 2$	0.0036 [0.0185]	0.028 [0.0270]	-0.0197 [0.0242]	-0.007 [0.0125]
$t + 3$	0.0142 [0.0209]	-0.0085 [0.0255]	-0.0275** [0.0112]	-0.0055 [0.0125]
$t + 4$	-0.0024 [0.0212]	-0.0117 [0.0270]	-0.0348*** [0.0114]	-0.0089 [0.0135]
$t + 5$	0.0016 [0.0179]	-0.0348 [0.0342]	-0.0408** [0.0159]	-0.0047 [0.0133]
N	121,093	77,765	88,476	74,473

NOTES: This table presents the regression estimates corresponding to Figure 2 or specification (1). Each column is from a different regression. The outcome variables are indicators for births of various orders. For the regression where birth of order b is the dependent variable, the sample is restricted to years after birth $(b-1)$ and until birth b for mothers whose $(b-1)^{th}$ child was born before the year of announcement. Standard errors in brackets are clustered by state. The year before announcement is the omitted year. *** 1%, ** 5%, * 10%.

Table A.4: Net effects on birth hazards

	(1)	(2)	(3)
A. 1st birth = 1			
$Treat_{st}$	0.0008 [0.0062] (0.0059)	0.0004 [0.0062] (0.0061)	-0.0006 [0.0060] (0.0058)
$T_s * Post_{st}$	0.0063 [0.0069] (0.0067)	-0.0004 [0.0101] (0.0097)	-0.0017 [0.0099] (0.0098)
$Post_{st}$	-0.0121 [0.0088]	0.0013 [0.0099]	0.0017 [0.0101]
N		323,174	
B. 2nd birth = 1			
$Treat_{st}$	0.0057 [0.0094] (0.0099)	-0.0011 [0.0120] (0.0114)	0.0015 [0.0104] (0.0108)
$T_s * Post_{st}$	0.0130 [0.0121] (0.0127)	-0.0020 [0.0164] (0.0155)	0.0023 [0.0147] (0.0143)
$Post_{st}$	-0.0113 [0.0117]	0.0014 [0.0160]	-0.0013 [0.0155]
N		186,219	
C. 4th birth = 1			
$Treat_{st}$	0.0066 [0.0093] (0.0092)	0.0003 [0.0054] (0.0054)	-0.0019 [0.0042] (0.0041)
$T_s * Post_{st}$	0.0004 [0.0116] (0.0110)	-0.0040 [0.0089] (0.0090)	-0.0077 [0.0077] (0.0078)
$Post_{st}$	0.0114 [0.0071]	0.0066 [0.0074]	0.0087 [0.0069]
N		139,408	
State FE	x	x	x
Year FE	x	x	x
Years since last birth FE	x	x	x
X_{it}	x	x	x
Linear state trends		x	x
State x Age FE			x

NOTES: This table reports the coefficients of $Treat_{st}$ from specifications (2a) and (2b). The outcome variables are indicators for births of various orders. For regressions where birth of order b is the dependent variable, the sample is restricted to years after birth $(b - 1)$ and until birth b for mothers whose $(b - 1)^{th}$ child was born before the year of announcement. In Panel A, FE for years since last birth are replaced with FE for years since marriage. Standard errors in brackets are clustered by state and in parentheses are wild-cluster bootstrapped by state. *** 1%, ** 5%, * 10%.

Table A.5: Effects on the number of living children, mother's age ≤ 33

	Kids = 1	Kids = 2	Kids = 3	Kids = 4	Kids = 5
	(1)	(2)	(3)	(4)	(5)
Panel A:					
$Treat_{st}$	0.0030 [0.0061] (0.0022)	0.0091 [0.0061] (0.0021)	-0.0021 [0.0057] (0.0058)	-0.0063** [0.0022]*** (0.0031)	-0.0024** [0.0020] (0.0030)
Baseline mean	0.2462	0.2350	0.1651	0.0812	0.0332
Panel B:					
$T_s * Post_{st}$	0.0029 [0.0069] (0.0067)	0.0125 [0.0095] (0.0097)	0.0009 [0.0077] (0.0076)	-0.0088 [0.0035]** (0.0042)**	-0.0014 [0.0028] (0.0028)
$Post_{st}$	0.0001 [0.0046]	-0.0048 [0.0064]	0.0040 [0.0025]	0.0036 [0.0030]	-0.0013 [0.0029]
Baseline mean	0.2508	0.2241	0.1609	0.0795	0.0322
N	1,060,282				

NOTES: This table presents the regression estimates corresponding to specification (2a) in panel A and to specification (2b) in panel B using indicators for 1/ 2/ 3/ 4/ 5 living children as the outcome variables. No sample restrictions are imposed, except that the sample is limited to women ≤ 33 years of age in a given year. Each column within a panel is a different regression. Standard errors in brackets are clustered by state and in parentheses are wild-cluster bootstrapped by state. The baseline mean is calculated for observations where $Treat_{st} = 0$ in panel A and for observations where $Post_{st} = 0$ in panel B. *** 1%, ** 5%, * 10%.

Table A.6: Sex ratio of first births

Male = 1	All	SC	ST	OBC	Upper
	(1)	(2)	(3)	(4)	(5)
Panel A: Only treatment states					
$Treat_{st}$	-0.0081 [0.0110]	0.0334 [0.0332]	0.0120 [0.0445]	-0.0474 [0.0401]	0.0081 [0.0217]
	34,018	5,818	5,783	7,511	14,906
Baseline mean	0.5152	0.5062	0.5103	0.5162	0.5198
Panel B:					
$Treat_{st}$	-0.0007 [0.0076]	0.0325 [0.0295]	-0.0063 [0.0339]	-0.0304 [0.0228]	0.0093 [0.0162]
	86,023	15,245	9,265	19,345	42,168
Baseline mean	0.5150	0.5128	0.5096	0.5173	0.5158
Panel C:					
$T_s * Post_{st}$	-0.0042 [0.0122]	0.0179 [0.0379]	-0.0367 [0.0486]	-0.0263 [0.0219]	0.0092 [0.0199]
$Post_{st}$	0.0056 [0.0126]	0.0226 [0.0275]	0.0435 [0.0491]	-0.0053 [0.0175]	0.0003 [0.0185]
N	86,023	15,245	9,265	19,345	42,168
Baseline mean	0.5144	0.5224	0.5056	0.5151	0.5128

NOTES: This table reports the coefficients from specifications (2a) and (2b). The sample is restricted to first births and the outcome variable is an indicator for the birth being male. Standard errors in brackets are clustered by state. SC, ST, OBC, and Upper indicate Scheduled Caste, Scheduled Tribe, Other Backward Class, and upper caste households, respectively. In panels A and B, the baseline mean is calculated for observations where $Treat_{st} = 0$ and in panel C for observations where $Post_{st} = 0$. *** 1%, ** 5%, * 10%.

B State-wise Regulations

1. Rajasthan:⁴³

According to the the Rajasthan Panchayati Raj Act, 1994, “...Every person registered as a voter in the list of voters of a Panchayati Raj Institution shall be qualified for election as a Panch or, as the case may be, a member of such Panchayati Raj Institution unless such person-...(1) has more than two children.”...“The birth during the period from the date of commencement of the Act (23rd April, 1994), hereinafter in this proviso referred to as the date of such commencement, to 27th November, 1995, of an additional child shall not be taken into consideration for the purpose of the disqualification mentioned in Clause (1) and a person having more than two children (excluding the child, if any, born during the period from the date of such commencement to 27th November, 1995) shall not be disqualified under that clause for so long as the number of children he had on the date of commencement of this Act does not increase.”

2. Haryana:

According to the 1994 Act⁴⁴, “...No person shall be a Sarpanch or a Panch or a Gram Panchayat or a member of a Panchayat Samiti or Zila Parishad or continue as such who- (q) has more than two living children: Provided that a person having more than two children on or upto the expiry or one year of the commencement of this Act, shall not be deemed to be disqualified.”

Prior to revocation.⁴⁵ “Person shall be disqualified for being elected to a Gram Panchayat, Panchayat Samiti or Zila Parishad if:

...(xvii) has more than two living children; provided that this disqualification of more than two living children shall not apply for the persons who had more than two living children before 21st April, 1995 unless he had additional child after the said date.”

The Haryana government amended Section 175(q) of the Haryana Panchayati Raj Act, 1994, retrospectively with effect from January 1, 2005 to omit the section (q).⁴⁶

⁴³Source: <http://www.rajpanchayat.gov.in/common/toplinks/act/act.pdf>

⁴⁴Source: <http://www.panchayat.gov.in/documents/10198/350801/The%20Haryana%20Panchayati%20%20Raj%20Act%201994.pdf>

⁴⁵Source: <http://secharyana.gov.in/html/faq1.htm>

⁴⁶Source: <http://hindu.com/2006/07/22/stories/2006072207150500.htm>

3. Andhra Pradesh:⁴⁷

According to Section 19 (3) of the Andhra Pradesh Panchayati Raj Act, 1994, “...A person having more than two children shall be disqualified for election or for continuing as member:

Provided that the birth within one year from the date of commencement of the Andhra Pradesh Panchayat Raj Act, 1994 hereinafter in this clause referred to as the date of such commencement, of an additional child shall not be taken into consideration for the purposes of this clause;

Provided further that a person having more than two children (excluding the child if any born within one year from the date of such commencement) shall not be disqualified under this clause for so long as the number of children he had on the date of such commencement does not increase;

Provided also that the Government may direct that the disqualification in this section shall not apply in respect of a person for reasons to be recorded in writing.”⁴⁸

4. Orissa:⁴⁹

A person shall be disqualified for being elected to a PR institution if he “...has more than one spouse living or has more than two children. The last named disqualification shall not apply if the person had had more than two children before 21.04.1995 unless he begot an additional child after the said date. Rule 25 of O.G.P. Act gives full description of the disqualifications.”

5. Madhya Pradesh:⁵⁰

“...condition to disqualify an office bearer of the Panchayat for holding the post: (1) that he must have more than two living children, and (2) out of whom one is born on or after the 26th day of January, 2001...”

The Population Policy of Madhya Pradesh states that “persons having more than two children after January 26, 2001 would not be eligible for contesting elections for *panchayats*, local bodies, *mandis* or cooperatives in the state. In case they get elected, and in the meantime they have the third child, they would be disqualified for that post.”

6. Chhattisgarh:⁵¹

⁴⁷Source: <http://www.ielrc.org/content/e9412.pdf>

⁴⁸Further explanation at: http://www.apsec.gov.in/RLBS_GPs/CLARIFICATIONS%202013/877%20-%20Qualification.pdf.

⁴⁹Source: <http://secorissa.org/download/FAQ2.pdf>

⁵⁰Source: <http://www.indiankanoon.org/doc/1285129/>

⁵¹Source: <http://www.the-laws.com/Encyclopedia/Browse/ShowCase.aspx?CaseId=023002211000>

“Section 36: Disqualification for being office bearer of Panchayat:- 36(1) No person shall be eligible to be an office bearer of Panchayat who:...(m) has more than two living children one of whom is born on or after the 26th day of January, 2001.”

7. Maharashtra:

“...(j-1) No person shall be a member of a Panchayat or continue as such, who has more than two children:

Provided that, a person having two children on the date of commencement of the Bombay Village Panchayats and the Maharashtra Zila Parishads and Panchayat Samitis (Amendment) Act 1995 (hereinafter in this clause referred to as “the date of such commencement”) shall not be disqualified under this clause so long as the number of children he had on the date of such commencement does not increase;

Provided further that, a child or more than one child born in a single delivery within the period of one year from the date of such commencement shall not be taken into consideration for the purpose of disqualification mentioned in this clause.

... For the purposes of clause (j-1):

Where the couple has only one child on or after that date of such commencement, any number of children born out of a single subsequent delivery shall be deemed to be one entity.

“Child” does not include an adopted child or children....”