Daughters, Dowries, Deliveries: The Effect of Marital Payments on Fertility Choices in India

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Abstract

This study estimates the effect of dowries on fertility in India. The future dowry associated with the birth of each child introduces a gender-specific cost to its parents. This leads families with more daughters to have higher fertility. For identification, the paper exploits a revision in anti-dowry law in combination with pre-treatment heterogeneity across the gender of the first child, maternal ethnicity and birth cohort. The resulting decrease in expected dowries attenuates the correlation between daughters and their parents' birth rates. The effect is strongest for lower birth orders and for more educated and autonomous women.

JEL Classifications: O15, J12, J13

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1 Introduction

It is widely accepted that Indian fertility is key to world demography. Over the past 5 years, the number of children born in India has accounted for about one fifth of global births (UN, 2013).[1] Moreover, the country's fertility rates are likely to translate into future population growth; India is estimated to overtake China as the world's most populous country within the next 20 years.[2] Concerns that these phenomena will lead to a scarcity of resources and societal problems (UNDP, 2006) have sparked an increased interest in the determinants of fertility choices. For each individual household, parents' desire to achieve a determined gender composition of their children has been identified as an important factor influencing reproductive decisions (Angrist and Evans, 1998; for instance).

A growing body of evidence points to parents favouring male offspring (Dahl and Moretti, 2008; among others). South East Asia in general and India in particular, have been argued to exhibit especially strong preferences for sons, based on the belief that sons are more valuable than daughters (Pande and Astone, 2007; Das Gupta et al., 2003; Clark, 2000). The presence of significant gender gaps in a number of human development indicators such as mortality (Bhargava, 2003; Arnold et al., 2002), nutrition (Jayachandran and Kuziemko, 2011; Oster, 2009; Pande, 2003), abortions (Bhalotra and Cochrane, 2010) and more recently child care (Barcellos et al., 2012) further corroborates these findings. Indeed, various studies have argued that parents adjust their reproductive behaviour in an effort to achieve their ideal mix of sons and daughters (Arokiasamy, 2002; Yount et al., 2000; Srinivasan, 1992). Empirically, this behaviour results in a correlation between the gender composition of a couple's offspring and its fertility rates; ceteris paribus, the presence of daughters in the family tends to increase fertility, whereas sons decrease birth rates. This correlation is often been interpreted as ”son preference” in fertility behaviour. Yamaguchi (1989) and Dreze and Murthi (1999) argue that this behaviour substantially increases fertility rates. Jensen (2003) meanwhile points out that these patterns ultimately decrease young girls' welfare by concentrating daughters into larger families. Whilst it is well documented that parents condition reproductive choices on the gender composition of their offspring, our understanding of the underlying mechanisms behind this behaviour

[1] 129 million births for India and 675 million births for the world

[2] Figures come from the Population Reference Bureau: http://www.prb.org/. The 2025 and 2050 estimates for India are 1.444 billion and 1.747 billion individuals. For China 1.476 billion and 1.437 billion individuals.

remains rudimentary. In particular, we are yet to identify the role pecuniary factors play in determining these patterns.

The present study addresses this gap by investigating the extent to which the correlation between the gender composition of a couple's offspring and its fertility choices is the result of gender specific economic costs of children. The focus is on one custom that is particularly widespread in India: dowries, defined as marital transfers of resources from the family of the bride to the groom or his family (see Anderson, 2007a; for a review). The prospect of these marital payments introduces an expected future cost that is conditional on the gender of each child: the birth of a girl will be associated with a negative, and the birth of a boy with a positive, income shock at the time of his or her marriage. Forward-looking parents are likely to take this into account when making reproductive decisions. Indeed, qualitative evidence has suggested a strong correlation between dowries and reproductive decisions (Diamond-Smith et al., 2008). In contrast to other expenses related to children, where the gender specific component can be hard to determine - take educational expenses for example - the focus on dowries will allow us to better approximate the effect of boy's or a girl's birth on the finances of a family.

The theoretical framework views the total number of a couple's offspring as the result of a series of sequential yes/no decisions. After every birth, parents decide whether or not to opt for a further child. Individuals are assumed to have children for two reasons. First, to increase their utility net of costs; second, to influence the flow of costs and returns associated with their children. The latter mechanism stipulates that parents aim to offset the negative income shock associated with the birth of a daughter with the revenue generated from dowries received at the time of a son's marriage. As a consequence, couples', whose offspring are mainly female, are likely to continue childbearing beyond their ideal family size. In contrast to fertility models rooted in the standard constrained optimisation setting, where parents decide on the optimal number of children at the outset of their reproductive years (see Becker and Lewis, 1973; for instance), this way of viewing fertility allows parents to revise their reproductive choices as the gender of each child is revealed birth by birth. The model has the following implications: (i) there exists a negative correlation between the expected value of the dowry and the probability of the woman experiencing a further birth; (ii) conditional on the total number of children, there exists a positive correlation between the number of daughters and the probability of the couple opting for a further birth; and (iii) the correlation mentioned in point (ii) depends positively on the expected value of the dowry.

To test the three aforementioned implications empirically, this study estimates a reduced form sequential fertility model using information on completed birth histories from three rounds of the National Family Health Survey (NFHS, 1994; 1999; 2007b). At every birth order, the parameter estimates show a strong positive correlation between the number of daughters in the family and the probability of the woman experiencing a further birth. This lends empirical support to theoretical implication (ii). On average the birth of a girl increases the conditional probability of a woman experiencing a further birth by 7 percentage points. The correlation, further, appears to be the strongest between the second and fourth birth.

To isolate the causal effect of dowries on fertility choices, the present analysis exploits a substantial revision in India's anti-dowry laws. The ineffective and widely criticised Dowry Prohibition Act of 1961 was tightened in 1985 under the Dowry Prohibition Rules. The changes encompassed more stringent monitoring of transfers associated with marriage as well as substantial increases in penalties for offenders. Evidence from the Survey of Status of Women and Fertility (SWAF, 2000) shows a marked decrease in the conditional probability of dowry payments in the years immediately after the introduction of the policy, which, in turn, is likely to decrease parents' expectation that a dowry will be transferred upon their children's marriage. This is borne out by descriptive evidence from the SWAF, which suggests that mothers became less willing to pay a dowry for children born after the implementation of the policy. With regard to fertility decisions, the theoretical model predicts that this shift in expectations will cause an increase in the probability of the couple opting for a further birth - see implication (i) - as well as an attenuation of the positive correlation between the number of daughters and birth rates - see implication (iii).

The policy change is evaluated in a difference in differences framework, which allows for a change in the intercept as well as in the slope parameter. Because the Dowry Prohibition Rules were introduced simultaneously in the whole of India in 1985, this paper exploits the heterogeneity in dowry payments before that year to identify the effect of the policy on reproductive decisions. It argues that the impact of the policy on an individual family was proportional to the average dowry paid and received by that household in the pre-treatment period. In other words, the change in the law had an especially strong effect on couples paying particularly high dowries before 1985. The specification defines treatment status along the lines of the gender of the firstborn child and employs mothers, whose firstborn is female, as treated individuals. The woman's ethnicity is employed as a further way to distinguish the treatment and control group. Evidence on dowry transfers between 1975 and 1984 from the International Crops Research Institute for the Semi-Arid Tropics (ICRISAT, 1984) suggests that these two groups paid significantly higher dowries than average. To strengthen the specification further, the empirical model takes advantage of the fact that some women had come to the end of their reproductive years by the time the changes in the policy had been implemented.

The estimates suggest that the amendment of the law impacted upon fertility behaviour significantly. The policy change is estimated to have lead to an one-off increase the probability of a woman experiencing a further birth of 4 percentage points; this is in line with prediction (i). The specification further points to an attenuation in the previously observed correlation between the number of girls in the household and fertility rate of 5 percentage points, this lends support to prediction (iii). For the treated, the policy decreased the influence of the gender composition on fertility by 20 percent. The effect appears particularly strong for children of lower birth orders and for more educated and autonomous women.

By analysing the link between dowries and fertility decisions, this study attempts to shed light on as yet unanswered questions such as: why is son preference still so widespread? What are its determinants? And how does it influence fertility decisions? Insights into these issues can uncover some of the mechanisms underlying fertility choices and are thus likely to be of interest to policy makers concerned with, for instance, fertility rates or son preferring behaviour more generally. Moreover, the present analysis aims to add to the growing knowledge base on dowries. Whilst the determinants of marital payments (Botticini and Siow, 2003; for instance) as well as their effects on brides (Bloch and Rao, 2002) have received growing attention, little is known on the ramifications of this practice on other household members.

The remainder of the paper is structured as follows: Section 2 introduces the data and gives motivating descriptive evidence. Section 3 explains the practice of dowries and the legal framework. Section 4 lays out a theoretical framework to help conceptualise the effect changes in dowry policies can have on fertility choices. Section 5 analyses the influence of a couple's gender composition on fertility; the effect of the policy change on this behaviour is explored in section 6. Section 7 shows the results of the empirical analysis, the robustness of which is shown in section 8. Section 9 concludes.

2 Fertility and Gender Preferences in India

This study starts from the qualitative observation that, when asked about their ideal number of children, women in India tend to report a preference for sons over daughters.[3]

2.1 The Data

The present analysis employs data drawn from three rounds of the National Family Health Survey (NFHS) for India (NFHS-1, NFHS-2 and NFHS-3), a nationally representative survey of Indian households. The NFHS is part of the Demographic and Health Surveys series, which is conducted in about 70 low and middle income countries around the world.[4] The questionnaires collect extensive information on health, nutrition, population and focus particular on women and children. The NFHS-1 (International Institute for Population Sciences and Macro International, 1994) was carried out in 1992 and 1993 and interviewed 89,777 ever-married women aged 13 to 49; the NFHS-2 (International Institute for Population Sciences and Macro International, 1999) was conducted in 1998 and 1999 and interviewed 89,199 ever married women aged 15 to 49; finally the NFHS-3 (International Institute for Population Sciences and Macro International, 2007b) was implemented in 2005 and 2006 and interviewed 124,385 women aged 15 to 49. Each round collected detailed information on women's complete birth histories including the number, gender and morality status of all births.

Individuals selected for the present purpose are women, who have experienced at least one birth and who have come to the end of their reproductive years, i.e. aged 36 or above. Although biologically women can still conceive in their late 30s and early 40s, the percentage of women doing so in India is very low. The NFHS-3 final report (NFHS, 2007a) indicates that fertility at ages 35 and above accounts for only 4 percent of total fertility in urban and 7 percent in rural areas.[5] Similarly, the omission of childless women is unlikely to bias the results significantly, only 3.6 percent of women

[3]The estimates reported are for women aged 35 or above at the time of interview. The same patterns, however, can be found in younger women.

[4]The data are publicly available at measuredhs.com

[5]Restricting the sample to women aged 40 and above does not alter the results significantly.

aged 35 to 40 in India have never experienced a birth (NFHS-3, 2007a). The final sample consists of 412,378 children of 99,533 mothers born between the years 1942 and 1970.

2.2 Descriptive Statistics

Women in the estimation sample show relatively low levels of education, around half of the individuals have completed primary school. The majority are Hindu (85 percent) with a minority of Muslim women (11 percent). Around 15 percent belong to a scheduled caste. These mothers give birth to, on average, 4.1 children. The mean age at first birth is around 20 years. The histogram in figure 1 shows the fraction of completed births for women born between 1942 and 1970. The mode of the distribution lies at 4 children. Most women have between 2 and 5 children. The gender composition of offspring appears relatively close to the natural rate; 47.8 percent of children born to women are female.[6] However, an inspection of the gender composition at every birth order reveals that the total number of children within a household and their gender composition are closely linked. Figure 1 also reports the average number of sons and daughters by the total number of births experienced in a woman's lifetime. Whilst, on average, families with five or less children are at home to more boys than girls, the opposite is true for larger families. This pattern is in congruence with Jensen (2003), who shows that in India girls tend to be brought up in larger families compared to boys.

The most common explanation for the aforementioned pattern is that parents adjust their reproductive behaviour to achieve a gender composition skewed towards boys. The retrospective questions elicited by the NFHS regarding the respondent's ideal number of children, sons and daughters confirm this.[7] Individuals in the sample report an ideal number of 2.8 children, 1.4 sons and 1.1 daughters.[8] Figure 2 plots the mean of the ideal number of children, boys and girls as well as the difference between the latter two by the birth year of the mother. For all years of birth, the ideal number of sons exceeds the ones of daughters. The difference between the ideal number of sons and

[6]The natural rate of girls born at birth is 48.8 percent.

[7]The relevant questions are *If you could choose exactly the number of children to have in your whole life, how many would that be?* and *How many of these children would you like to be boys and how many girls?*

[8]The number of ideal sons and daughters do not add up because in the second and third round of the NFHS women were also asked about their ideal number of children regardless of the sex.

daughters is around 0.4. Only 14 percent of respondents report not to desire any sons. However, women report to also desire to give birth to daughters; the average number of ideal daughters is constantly around 1.1. Indeed, only 19 percent of respondents report not to desire any girls. Whilst the ideal number of children decreases over time, from 3.3 to 2.5, the difference between the ideal number of sons and daughters remains remarkably stable, from 0.5 for the oldest women to 0.3 for the youngest. This finding is in line with previous research pointing to the fact that attitudes towards girls only change very slowly over time.

For the whole sample, only 3 percent of mothers report to desire more girls than boys. The majority of respondents, 64 percent, report an ideal number of sons equal to one of daughters. However, a third of all mothers reports a higher number of ideal sons compared to daughters. This descriptive and retrospective evidence suggests that women desire sons as well as daughters. However, on average, women desire at least as many sons as daughters.

3 Dowries in India

The custom of marital payments is widespread in India. Between 60 and 90 percent of women interviewed in 1993 (SWAF, 1994) reported to have paid a dowry at their own marriage.

3.1 Evidence on Dowries

Much of the research on dowries has focused on the prevalence and value of marital payments. From a theoretical point of view, Anderson (2003) maintains that the prevalence of dowries in India is a result of fast economic development combined with the rigid social system provided by the country's caste system; Tertilt (2005) considers the importance of monogamy. The value of these transfers, however, has been debated. Whilst some studies argue for substantial dowry inflation (Rao, 1993, 2000) others do not find evidence for this phenomenon (Arunachalam and Logan, 2008; Anderson, 2007b; Edlund, 2000). Other determinants of dowries identified are idiosyncrasies of the bridal couple such as physical appearance of the bride (Caldwell et al., 1983) and the groom's labour market prospects and characteristics (Deolalikar and Rao, 1998; Dalmia, 2004).

A commonly used source of information on dowry transfers is the survey carried out by the International Crops Research Institute for the Semi-Arid Tropics (ICRISAT) in 6 different villages in the Indian state of Andhra Pradesh between the years 1975 and 1984. This survey contains information on, inter alia, age, marital status, education level, primary and secondary occupations, of all household members and inventory files for animals, farm implements, farm buildings, and current physical stocks as well as on financial assets and liabilities such as bank accounts, life insurance, loans and dowries. Evidence from this source suggests that dowry payments are considerable. Summary statistics of average dowries paid and received by households are reported in table 1. The figures for the whole sample are reported in column (1). The majority of households report to have either paid or received a dowry, 81 percent. Rows (b) and (c) report the dowries paid, (d) and (e) dowries received and (f) and (g) the net out payments per household. For the whole sample, households pay larger amount of dowries than they receive, 1,584 Rupees versus 600 Rupees per annum. The 684 Rupees per year net payments translate into around 70 Rupees per person residing in the household per year.[9] Rows (h) to (i) report net dowry payments as the percentage of other household transfers and indicate that marital payments constitute a sizeable part of the family's budget making up more than half of the household's non-durable consumption expenditures and 8 percent of the income of all household members combined.

3.2 Laws Regarding Dowries in India

In an attempt to curb the prevalence of dowries, the government of India passed the Dowry Prohibition Act in 1961 prohibiting the giving and taking of dowry, defined as "any property or valuable security given or agreed to be given either directly or indirectly a) by one party to a marriage to the other party to the marriage or b) by the parents of either party to a marriage or by any other person to either part to the marriage or to any other person at or before or any time after the marriage in connection with the marriage of said parties" (Dowry Prohibition Act, 1961). According to the Act the giving or taking of dowries is punishable with an imprisonment of no less than 6 months and with a fine of 5,000 Rupees or the amount of the dowry, whichever is larger. Similarly, demanding a dowry is punishable with imprisonment of no less than 6 months and with a fine of up to 10,000 Rupees. Furthermore, the act stipulates that if a dowry is received by a person other than the woman, whose marriage the dowry

[9] All numbers are given in 1984 Rupees. Official GDP per capita in 1980 was Rupees 1,630.

is associated with, it shall be transferred in full to the bride. The law also bans any advertisement of dowries.

Despite this policy, the practice of dowries continued to flourish - see evidence in section 3.1. One reason put forward for this is the Act's lack of provisions for monitoring marital transfers. In response to this, the government of India introduced the Dowry Prohibition Rules (1985),[10] which are the focus of this empirical analysis. The purpose of this amendment was to make the Dowry Prohibition Act of 1961 more stringent and effective in a number of ways. First, the legislation establishes a set of rules in accordance with which a list of presents has to be maintained. The list of presents given to be bride is mainlined by the bride whereas the list containing presents to the groom is kept with the groom. These lists must to be in writing and contain a brief description of each present, the approximate value of the present, the name of the person who has given the present and where the person giving the present is related to the bride or bridegroom, a description of such relationship. These lists are required to be signed by both parties. Second, the Dowry Prohibition Rules raise the minimum punishment for taking or abetting the taking of dowry to 5 years of imprisonment and to a fine of 15,000 Rupees. Third, the burden of proving that no funds were exchanged now lies with the person who takes or abets the taking of dowry. Fourth, offences to the act are made non-bailable.

4 Fertility as a Sequential Stopping Decision

Much of the previous work on fertility choices is based to some extent on the Becker-Lewis framework (Becker and Lewis, 1973), where parents choose the number as well as the quality of children subject to the household's budget in a constrained optimisation setting. Nevertheless, there are two reasons why this approach does not allow parents to condition their reproductive behaviour on the gender composition of their offspring. First, parents decide on the total number of children irrespective of gender. Second, parents are assumed to decide on their optimal number of children at the outset of their reproductive years. However, the gender of the child may not be known or influenced before each birth.[11] Thus, parents aiming for a particular gender composition are likely to revise their fertility choices as the gender of each child is revealed birth by birth.

[10]Amendment Act 63 of 1984 came into force on the 2.10.1985.

[11]Sex selective abortions are not considered here. This is discussed in section 8.

One possibility of linking the inherent uncertainty regarding the gender of the child to fertility decisions is to view the total number of births a woman experiences as the result of a number of yes/no decisions. Instead of choosing the optimal number of children before commencing their reproductive years, the couple treats the birth of each child separately. In practice, after every birth, parents decide whether or not to continue childbearing. The advantage of this framework is that it allows parents to incorporate newly available information - such as, for instance, the gender of the last child - at every birth order.

It is widely accepted that the gender composition of a couple's offspring affects its fertility decisions for a variety of reasons. Schultz (1993), for instance, points out that parents may prefer sons for both economic and non-economic reasons - because of differences in farm productivity, for example, or out of a desire to raise children with culturally accepted characteristics. Although this study acknowledges the importance of the latter category of determinants, it abstracts from the possibility that the gender of each child affects its parents' utility for a number of reasons. In the first instance, the descriptive evidence in section 2 and the parameter estimates based on equation 5 suggest that parents might be aiming to match the number of daughters with an equal number of sons. This desire might stem from either a very specific set of preferences regarding the gender of children or, more likely, from an effort to balance the raising costs resulting from the gender composition of children in the household. Furthermore, the assumption of parental indifference with regard to the gender of each child can be viewed as an abstraction to achieve this study's original purpose , i.e. to investigate the extent to which the economic costs of children influence reproductive choices. The framework presented below can, in principle, also accommodate the notion that parents prefer sons to daughters. Doing so, however, would imply assumptions about the functional form of the parents' utility function, which this study intends to avoid.

4.1 Utility and Costs of Children

In the theoretical framework proposed here, parents weigh up the utility gains from a further birth against the expected costs, which depend on the gender of the child. The couple draws utility from the total number, n, as well as from the quality, q, of children, irrespective of gender. The utility function is denoted as $U(n, q)$ and the marginal utility of children is assumed to be positive and decreasing.[12] The economic

[12]The quality of children is taken as exogenous here.

costs of children consist of two components. The first is the net economic cost of every child, which is independent of its gender. It is denoted by π_n and consists of all child raising costs, such as, for instance, food or educational expenses, net of any returns the parents receive from children - for example remittances or household production. This cost component is linear in the number of children, n.

The second cost component of every child is the dowry paid or received at the time of his or her marriage, which is contingent on the child's gender. For every daughter born, g, the family may have to pay a dowry at the time of her marriage. The expected total of all dowry payments to be made by the parents is denoted as $D_g(g)$. From the parents' perspective, the birth of a girl will thus lead to an expected negative income shock with a value of

$$\pi_n + D_g'(g)$$

where $D_g'(g)$ is the partial derivative of $D_g(g)$ with respect to the number of girls, g. For every son born, b, by contrast, the parents may receive a dowry - paid by the parents of the future bride. The expected total of all dowries paid to the sons in the family is denoted by $D_b(b)$. Of this, the parents capture only a percentage, ϕ. This study assumes that the income generated by the incoming dowry associated with the marriage of a son exceeds the net child raising costs. For the parents, the birth of a boy is thus associated with an expected positive income shock with a value of

$$\phi D_b'(b) - \pi_n$$

where $D_b'(b)$ is the partial derivative of $D_b(b)$ with respect to the number of boys, b. This theoretical framework is a partial equilibrium model, which considers fertility choices in isolation. It assumes that parents make reproductive decisions independently of consumption choices. Parents do not, for instance, substitute consumption of other goods in order to afford the birth of an additional child and vice versa. Instead, the couple is assumed to set aside an exogenously fixed amount from their income to pay for fertility related costs, E_n. Total expenditures on children cannot exceed this amount.

4.2 The Fertility Decision

At every birth order, the couple takes the binary decision whether or not to have an additional child. Two separate mechanisms can lead to the parents deciding in favour of another birth. First, the couple will opt for a further birth in order to increase its utility net of costs. In other words, the parents will have another child if the marginal utility of another child exceeds its expected marginal costs. If pecuniary factors enter the utility function linearly, the decision can be written as

$$MU_n - \pi_n - \frac{1}{2}D'_g(g) + \frac{1}{2}\phi D'_b(b) \geq 0 \tag{1}$$

where MU_n is the partial derivative of the utility function $U(n, q)$ with respect to n. The expected costs of the next birth are defined in section 4.1 and incorporate the uncertainty surrounding next child's gender.[13] This mechanism is in the spirit of the Becker-Lewis framework (1973) and can be thought of as parents trying to achieve their "ideal" family size.

Second, the couple will opt for a further birth in order to change the gender composition of its offspring in an effort to keep the overall economic costs of children below the threshold for child expenditures, E_n. Because the economic cost of each child crucially depends on its gender, the total economic costs of all the couple's children depend on their number as well as on their gender composition. The costs associated with the number of daughters in the household $(D_g(g) + \pi_n g)$ decrease the parents' budget, whereas the income streams associated with the number of sons $(\phi D_b(b) - \pi_n b)$ are assumed to have the opposite effect. As a consequence, only certain gender compositions will result in costs inferior to E_n and the couple will have to achieve a specific gender composition along with a determined total size of offspring to keep total child raising costs below that threshold.

Many aspects of fertility-related costs are incurred at different stages of a child's lifetime. Educational costs, for instance, commence with school enrolment. Similarly, dowries only have to be transferred at the time of a child's marriage. To incorporate the time lag between reproductive decisions and their costs, the present analysis assumes that the family will only stop childbearing once the total cost of all children falls below the budget allocated for expenditure on children. Thus, parents will opt for a further birth if the total costs of children exceed E_n

[13] The probability of the birth of a daughter is set equal to one half. Strictly speaking it is slightly lower than that, 48.8 percent.

$$D_g(g) + \pi_n g + \pi_n b - \phi D_b(b) - E_n \geq 0 \quad (2)$$

where the cost components are as defined in section 4.1. Because the rationale for the mechanisms behind fertility choices outlined in equations 1 and 2 are distinct, this study assumes that the satisfaction of one inequality is a sufficient condition for the couple to have another child.

The parents' expectations regarding future dowry transfers, $D_g(g)$ and $D_b(b)$, are defined as a function of the family's characteristics, f, some of which have been highlighted in section 3. Also, both terms are increasing in the perceived probability that a dowry will be transferred at the time of their child's marriage, p. The inclusion of this latter factor reflects the empirical observation that not all households in India pay dowries (see section 3). In their simplest form, $D_g(g)$ and $D_b(b)$ are the products of the average dowry paid or received by a household with characteristics f, D(f), and the perceived probability of a dowry being transferred:

$$D'_g(g) = D'_b(b) = pD(f)$$

This linear specification, however, can lead families with particular gender compositions to never stop childbearing. One way to exclude this eventuality is to assume that the expected dowry to be paid for every daughter decreases with the number of daughters already born. The dowry function hence becomes

$$D_g(f, g, p)$$

which is increasing at a decreasing rate with g and increasing in p. By contrast, the expected dowry income from sons, $\phi D_b(b)$, remains independent of the gender composition of the household.

4.3 Testeable Implications

Using the framework laid out in sections 4.1 and 4.2, the possibility of future dowry payments affects fertility choices twofold. First, at every birth order, future dowry payments affect the expected cost of the next child - see inequality 1. The birth of a girl implies a negative and the birth of a boy a positive income shock. Under the assumption that $D_g(g)$ and $D_b(b)$ are of similar magnitude, the net impact of dowries on the expected costs of the next birth is negative. The first empirical implication of the model thus stipulates that (i) there exists a negative correlation between the anticipated dowry payments and the probability of the couple opting for a further birth.

Second, at every birth order the number of daughters alive determines the overall cost of children; the larger the number of girls, the higher the likelihood that the expenditure on children exceeds the threshold E_n, which increases the probability of the woman experiencing a further birth - see inequality 2. Intuitively, the parents attempt to offset the future dowry payments resulting from the birth of a girl with the revenue generated by the marital transfers they receive at the time of their sons' marriage. In this model, the couple will thus respond to the birth of a daughter by increasing its fertility rates in an effort to give birth to a son. The second implication of the model thus states that (ii) there exists a positive correlation between the number of daughters at every birth order and the probability of a subsequent birth. The strength of this correlation crucially depends on the parents' perceptions regarding future dowry transfers. In terms of equation 2, high expected marital transfers translate into a large difference between dowry in- and out-flows for the household, $D_g(g) - \phi D_b(b)$. Ceteris paribus, this increases the probability of inequality 2 holding, which, in turn, increases fertility rates. The intuition behind the correlation between the number of daughters and fertility can be thought of as follows. Larger expected dowries lead to a starker difference between dowry in- and out-flows of the household. For parents with predominantly female offspring, this increase in expected transfers will decrease the set of gender compositions that result in dowry expenses below E_n. This, in turn, diminishes the probability of inequality 2 holding thus increasing birth rates. Hence, the third empirical implication of the model is that (iii) the correlation between the number of daughters in the household and the likelihood of another birth - mentioned in point (ii) - depends positively on the anticipated value of dowries.

5 Empirical Framework

To empirically test implications (i), (ii) and (iii), this study estimates a reduced form fertility equation, which may be considered as a sequential probit regression.[14]

[14]The functional form chosen here is the linear probability model.

5.1 Reduced Form Fertility Equation

The empirical specification constructs an unbalanced panel using women's complete birth histories. Each mother contributes J+1 observations where J is the total number of births experienced in her lifetime; one for every birth she experiences with the addition of one observation for her entering motherhood. The outcome of interest is the unobserved propensity of woman i to give birth at birth order j, y^*_{ij}. Empirically, this is captured by an indicator function taking the value one if the couple opts for a further birth, $\mathbb{1}(y^*_{ij} > 0)$. Thus, for every woman the dependent variable will be vector of ones followed by a zero once she stops childbearing at birth order J. A similar specification has previously been employed in the context of infant mortality (Arulampalam and Bhalotra, 2006).

Many earlier studies have modelled the determinants of fertility by either employing parity progression (Borooah and Iyer, 2004; Moursund and Kravdal, 2003; Arokiasamy, 2002) or count models (Wang and Famoye, 1997; Caudill and Mixon, 1995). The drawback of the former approach is that it only allows for analysing part of a woman's reproductive history. The latter models, by contrast, make the implicit assumption that all counts are derived from the same data generating mechanism. The empirical framework proposed here addresses the first concern by modelling a woman's complete birth history ex post and the second by allowing the gender composition to change after every birth.

Due to collinearity, the effect of the number or boys, girls and total children on fertility cannot be investigated jointly. This study employs the latter two for a number of reasons. In a first instance, the qualitative evidence in section 2.2 suggests that parents aim for a number of sons greater or equal to their daughters. As a consequence the effect the birth of a girl is likely to be unequivocally positive, whereas the opposite is not necessarily true for boys. The evidence in section 5.2 further shows that the relation between daughters and fertility rates is linear and robust to the number of sons in the family. Finally, in the theoretical framework the effect of sons on fertility is clouded by the factor ϕ. If the parents are only able to capture a small amount of the dowries paid to their son, the effect of male births on fertility will be small. By contrast, the birth of a girl will be likely to imply a large negative income shock.

Woman i's probability of experiencing a further birth at birth order j is defined as

$$y^*_{ij} = \gamma_n n_{ij} + \gamma_g g_{ij} + x'_{ij}\beta + \alpha_i + u_{ij} \tag{3}$$

where n_{ij} and g_{ij} are the number of children and girls alive for woman i at birth order j, x_{ij} a vector of exogenous covariates, α_i the woman's unobserved heterogeneity and u_{ij} a woman and birth order specific error term. Recall that prediction (ii) stipulates a positive correlation between the number of girls alive at every birth order and the probability of the woman experiencing another birth. In terms of equation 3, this implies a positive coefficient on the number of girls at every birth order, g_{ij}. A test of $\gamma_g = 0$, therefore, will test implication (i) of the theoretical model.

A possible concern is the presence of child mortality. The theoretical literature stresses that parents decide on the total number of surviving children (Becker and Lewis, 1973; Bhaskar and Gupta, 2007). The death of one child is, therefore, likely to prompt the birth of another. Similarly "excess children" will receive disproportionately little resources, which will decrease their survival chances. Mortality rates in India are not negligible; for the whole of India, the under-5 mortality rates[15] for the years 1980, 1990 and 2000 were 152, 117 and 91 deaths per 1,000 births respectively.[16] To address this issue, the present analysis considers the sex mix of surviving children rather than the gender composition of all previous births irrespective of that particular child's survival status.[17]

The coefficient γ estimated in equation 3 denotes the correlation between g_{ij} and y^*_{ij} averaged across different birth orders of children. To investigate whether this association changes with the birth order of every child, a more flexible specification is estimated as

$$y^*_{ij} = \sum_k s_{i,k-1}\gamma_k + x_{ij}\beta + \alpha_i + u_{ij} \tag{4}$$

where k is the birth order of the child. In practice, the coefficient on the number of girls, γ_k, is allowed to vary by the birth order of every child.

5.2 Functional Form

A number of concerns are connected with the modelling of the interaction between the number of daughters and the probability of the woman experiencing a further birth. First, the specification in equation 3 assumes a linear relationship between g_{ij} and y^*_{ij}.

[15] Under-5 mortality is defined as the deaths per 1,000 births of children before the age of 5

[16] Source: World Bank, Development Indicators, 2010. http://data.worldbank.org/

[17] The specification also includes a dummy for the death of one of the children in the family

To investigate this linearity assumption, equation 3 is re-estimated using a dummy for every girl born as

$$y_{ij}^* = \sum_k D_k \psi_k + x_{ij}\beta + \alpha_i + u_{ij} \tag{5}$$

where D_k is an indicator variable taking the value 1 if the number of surviving girls equals to k. The parameter estimates of ψ_k along with a histogram for the fraction of the number of girls alive in each family in the sample are shown in Figure 3. For values of the explanatory variable between 1 and 6, the dashed line suggests a linear relationship between the number of girls alive and the probability of the woman experiencing a further birth.[18]

Second, the specification in equation 3 assumes that parents' only criterion for reproductive decisions is the number of girls alive. Parents may, however, take other factors into account - take the ratio of sons to daughters, for example. As a consequence, controlling for these alternative factors may alter the correlation between daughters and fertility significantly. To address this concern, equation 5 is re-estimated with the inclusion of variables accounting for the gender composition of the child's siblings; a dummy for more boys than girls in the family and one for the opposite case. The parameter estimates for the number of daughters on fertility are reported in figure 3 as the dotted line. They show that the inclusion of variables approximating the gender composition of the household does not change the correlation between the number of daughters and fertility significantly. The dotted line is very similar to the dashed line for the model without these controls.

6 The Effect of Dowries on Fertility Decisions

The empirical modelling of the effect of marital payments on fertility decisions - outlined in implications (i) and (iii) - is complicated by the fact that there is little information on these transfers. In addition, decisions on marital transfers are likely to be determined jointly with reproductive and other family-related decisions. Finally, current dowry payments are only error prone measures of parents' expectations regarding future dowries. To isolate the causal effect of anticipated dowries on fertility decisions, this study exploits the major revision in anti dowry laws outlined in section 3.2.

[18]The percentage of women with more than 6 girls alive is around 1%.

6.1 How the Policy Changed Expectations

This study argues that the introduction of the Dowry Prohibition Rules in 1985 lowered parents' expectations that a dowry would be transferred at their children's marriage. In terms of the theoretical model, this decrease lowers the expected dowry transfers, $D_g(g)$ and $D_b(b)$, in equations 1 and 2, which depend positively on the expected probability of a dowry being transferred, p. The newly established monitoring devices, in particular the mandatory list of transfers, increase the likelihood of illicit marital payments being detected by the authorities. The more severe penalties associated with such a conviction, in turn, are likely to act as a further deterrent to the giving and taking of dowries.

Evidence from the Status of Women and Fertility Survey (SWAF) suggests a marked decrease in the incidence of dowry payments after the change in law. The survey is part of a series implemented in India, Malaysia, Pakistan, the Philippines and Thailand. The questionnaire collects information on health, marital status, dowry transfers and different dimensions of female autonomy.[19] Figure 4 reports the differences in the conditional probability of a dowry being paid at a woman's marriage between the years 1970 and 1994.[20] Whilst the conditional probability estimates of a dowry being paid in the years leading up to the policy are very similar, around 12% to 16% lower than the base years, there is a significant decrease in the likelihood from the year 1985 onwards. For women married in the year 1984, the conditional probability is 14% lower than the base period. By contrast, for individuals married in 1985, the likelihood fell to 30% below the reference years. In subsequent years, the probability estimates oscillate between -30% and -40% eventually dropping to -60% for the last year, 1992. This evidence lends empirical support to the claim that the tightening of the anti-dowry laws has affected the incidence of marital payments.

Further descriptive evidence from the SWAF suggests that the change in policy indeed affected parents' expectations that a dowry be demanded. One question concerns mothers and their expected future dowries. Figure 5 shows the percentage of mothers expecting to pay a dowry separated by the year of the last born child. Individuals who finished their childbearing years before the introduction of the policy are less likely to have heard about the change in the law, and are consequently expected to report a

[19]The data are publicly available at swap.pop.upenn.edu/datasets

[20]The base years are 1966 to 1972. Covariates include the years of birth of the two members of the couple, their education, their parental background and a village level fixed effect.

stronger willingness to pay dowries in the future. This is borne out by the descriptive evidence shown here, where the percentage of women intending to pay a dowry in the future increases steadily until the introduction of the policy. After the policy change, by contrast, the fraction decreases.

6.2 Policy Evaluation

The changes in expectations resulting from the introduction of the Dowry Prohibition Rules are predicted to affect fertility decisions in two distinct ways. First, implication (i) stipulates that anticipated dowry payments have a direct and negative impact on fertility decisions. Within this framework, a decrease in the anticipated sum to be transferred will lead to an increase in the probability of a further birth. In terms of equation 1, the net expected sum of dowry transfers, $\frac{1}{2}D'_g(g) + \frac{1}{2}\phi D'_b(b)$, decreases the probability of a further birth. Empirically, this translates into an upward shift in the reduced form fertility regression in equation 3. Second, implication (iii) of the theoretical model states that the positive correlation between the number of girls and the probability of a further birth - denoted by the coefficient γ_g in the reduced form regression in equation 3 - increases in line with expected dowry payments. According to the model, a decrease in the anticipated sum of these payments will attenuate this correlation. In terms of equation 2, a decrease in the total dowries paid by the household, $D_g(g)$, will weaken the effect of the number of girls, g, on fertility decisions. Empirically, this translates into a decrease in the slope parameter γ_g. The paper tests for these two effects within a difference in differences framework, which allows for a change in the slope as well as in the intercept

$$\begin{aligned} y^*_{ij} &= (D_P + D_T + D_P D_T)(1 + g_{ij}) + \\ &\quad n_{ij}\gamma_n + g_{ij}\gamma_g + x_{ij}\beta + \alpha_i + u_{ij} \end{aligned} \tag{6}$$

where D_P is an indicator variable for the child being born after the introduction of the policy in 1985, and D_T a dummy variable for the child's mother belonging to the treatment group. A test of the hypothesis that the coefficient-estimate on the post-treatment interaction $D_T D_P$ is positive investigates whether the change in law increased the intercept of the reduced form equation. Similarly, a test for the parameter on the triple interaction $D_T D_P g_{ij}$ being negative informs us whether the policy indeed changed the correlation between the gender composition and the probability of a further birth denoted by the coefficient γ_g.

Because the policy changes outlined in section 3.2 were introduced in the whole country simultaneously, geographical variation cannot be employed to identify the effect of the policy on reproductive decisions. Instead, the analysis exploits the heterogeneity of dowry payments before the implementation of the Dowry Prohibition Rules in order to argue that the effect of the policy varies according to the average amount of dowries paid by households before its introduction. In particular, it maintains that the policy change had a more pronounced impact on households paying particularly high dowries prior to 1985; these individuals are the treatment group. By contrast, households paying on average low dowries are unlikely to have been affected by the change in law; these individuals are the control group. The present analysis evaluates the Dowry Prohibition Rules by comparing mothers with a firstborn son to women whose firstborn child is female. To further strengthen the results, the specification exploits variation by the mother's caste as well as by her birth cohort.

6.2.1 Variation by Gender of Firstborn Child

Families with a firstborn daughter are more likely to have been affected by the Dowry Prohibition Rules than couples with a firstborn son for two reasons. First, the gender of the firstborn child is likely to influence the total amount of dowry transfers paid by the parents. There is a strong and positive correlation between the number of daughters in the family and the total sum of a given household's dowry expenditures. Families with a firstborn daughter are more likely to have a higher number of girls because of the gender of the child itself on the one hand, and by increasing fertility levels on the other hand. Second, the gender of the firstborn is likely to determine the timing of dowry payments. If the oldest child in the family is a girl, the negative income shock resulting from her dowry payment is likely to occur before any positive income streams deriving from, for instance, sons' dowries or remittances. The fact that age at first marriage is relatively low in India exacerbates this. The gender of the firstborn child has been increasingly employed in past studies to approximate for parental fertility as well as for the gender composition of children (Rosenblum, 2013; Jensen, 2003; for instance).

In terms of the theoretical framework, inequality 2 states that households with a large number of daughters are more likely to have expenditures on children in excess of E_n. The decrease in expected dowries is likely to bring some of these families below E_n and thus have a stronger impact on fertility than on households with a predominantly male offspring. The gender of the first child can be seen as an approximation of the number of daughters alive at the introduction of the Dowry Prohibition Rules. Column (1) in table 2 shows that in the sample at hand the gender of the firstborn child strongly influences the gender composition of children born to women in India. The parameter estimates indicate that the birth of a girl at birth order one increases the ratio of girls at the end of the woman's reproductive life by 32%.[21]

Descriptive evidence from the ICRISAT data on dowry transfers between 1975 and 1984 lends weight to the positive correlation between the women residing in the household and the total amount of dowry payments by the household. Columns (2) and (3) of Table 1 show that household in which a daughter married during the sample period show a higher proportion of dowry transfers, 87 percent, compared to households where such an event did not take occur, 74 percent. Furthermore, the marriage of a daughter increases the net amount of dowries paid out by the household; 1,115 Rupees per year compared to 721 Rupees, see row (f). When compared to the households' consumption on non-durables in row (h), the difference is stark, 72 versus 33 percent. Note that these figures consider the overall gender composition of the household and not just the gender composition of the offspring, and can thus only be seen as suggestive evidence.

A possible concern with this empirical strategy is the exogeneity of the sex of the firstborn child. The presence of sex selective abortions in India at birth order one is debated. Whilst Jha et al. (2006) argue for the presence of this practice even at the first birth, Retherford and Roy (2003) and more recently Poertner (2010) and Rosenblum (2013) have shown that sex ratios at birth lie within normal limits. Descriptive evidence from the three rounds of the NFHS appears to lend support to the exogeneity of the sex of the first born child. Table 2 shows the sex ratio at birth and differences in socio-economic characteristics for women with a firstborn son versus a firstborn daughter. In Row (a), the percentage of girls born at birth order one for the sample at hand, 47.9, is very close to the one predicted by the natural rate shown in row (b). Columns (2) and (3) report the characteristics of parents by the gender of their firstborn child. The differences appear negligible. The only statistical significantly different variables are the percentage of women with primary education and the mother's age at first birth. However, the magnitudes of these differences appear very small, 1.1% for maternal education and 0.1 years for the age at first birth. Column (5) of Table 2

[21]The magnitude of the estimate is in line with what reported by Rosenblum (2013)

shows the parameter estimates of the regression of the firstborn's gender on parental characteristics. The mother's education and age at birth are significantly correlated with the gender of the firstborn son. Akin to above, the parameter estimates do not appear very large, 0.007 for the former and 0.001 for the latter. The low R-squared further points to the exogeneity of the firstborn's gender. A possible reason for these findings is that ultrasound technology was not widely available in the mid 1980s. This notion will be explored further in section 8.

6.2.2 Variation by Ethnicity

There are two reasons why dowries weigh more heavily on the budget of low caste compared to that of high caste families. On the one hand, individuals belonging to lower castes tend to exhibit worse socio-economic outcomes. The government of India explicitly recognises this and classifies scheduled castes, tribes and other castes as historically disadvantaged groups of society, which are to be given special provisions.[22] Academic work has also pointed to similar differences (Deshpande, 2000; for instance). On the other hand, lower caste status is seen as an unattractive feature in a bride (see Anderson, 2003; for a theoretical model). As a consequence, parents of lower castes are likely to have to pay higher dowries than higher caste members to attract husbands of the same quality, ceteris paribus. In fact, the growing importance of dowries for the budget of lower caste households has been documented. Whilst the practice originated among members of the highest caste Brahmins, and Rao (1993) finds a positive correlation between a household's caste rank and dowries, more recent work points out that lower caste members started paying disproportionately high dowries in an effort to copy the higher castes (Srinivas, 1997).

In terms of the theoretical framework, one can think of the low budget available to lower castes as translating into higher dowry payments $D_g(g)$ without necessarily generating higher dowry revenues, $\phi D_b(b)$. Evidence from the ICRISAT confirms that marital payments weigh more on the budget of lower caste than higher caste households. Columns (4) and (5) of table 1 show that, lower caste households exhibit a higher probability of paying dowries, 84 versus 80 percent. Furthermore, although the caste of the household does not appear to influence the inflow of dowry payments, lower caste households appear to show considerably higher out payments of dowries. This

[22]The affirmative action of the "Reservation Policy" in India is an example of policies aimed explicitly at the lower castes.

translates into considerably larger net dowry payments for these households, 2,539 Rupees per year compared to 216 Rupees per year for higher caste households.

6.2.3 Variation by Birth Cohort of Mother

To further strengthen the two sets of results laid out in sections 6.2.1 and 6.2.2, this study exploits the fact that some women had come to the end of their reproductive years by the time the Dowry Prohibition Rules were introduced in 1985. Women born between the years 1954 and 1970 were aged 15 to 31 at the implementation of the policy change. These individuals are assumed to be strongly affected by the policy. By contrast, mothers born between 1942 and 1948 were aged 37 to 43 at the time of the policy change and thus less likely to have been affected by the change in the law. Referring only to information from the latter group, the empirical analysis evaluates a fictional policy in the year 1971, using the two methodologies outlined above to test the hypothesis that the policy had a negligible effect on fertility decisions.

7 Results

The empirical models laid out in sections 5 and 6.2 are estimated using completed birth histories drawn from three rounds of the National Family Health Survey for India. Section 2.1 describes the data employed in detail. The dependent variable is the unobserved probability of the woman experiencing a further birth at every birth order. The specifications control for the number of girls and children alive and sequentially add a time trend, maternal and paternal characteristics as well as a mother-level fixed effect.

7.1 Gender Composition of Children and Fertility Choices

The estimates of the empirical framework outlined in equation 3 show a strong positive correlation between the number of daughters alive at every birth order and the probability of a subsequent birth. This lends support to prediction (ii), i.e. the positive correlation between the number of daughters alive at every birth order and the probability of a subsequent birth. A birth (and survival) of a girl is associated with a 5 to 7 percentage point increase in the probability of the woman experiencing a further birth. This finding is in congruence with the large number of studies reported in the introduction pointing out the strong influence of gender preferences on fertility decisions in India and Southeast Asia.

The relative magnitudes of γ_n (the coefficient on the number of children) and γ_g (the coefficient on the number of girls) shed further light on the way in which the gender composition of a couple's offspring influences its fertility decisions. For the preferred specification in column (5), the magnitudes of the two parameter estimates are of similar size, -0.101 for γ_n and 0.068 for γ_g. This suggests that the negative influence of a further birth on subsequent fertility is almost completely outweighed by the positive effect of the delivery of a girl. A possible explanation for this fact is that parents aim for a particular number of boys as pointed out by Yamaguchi (1989) and Jensen (2003).

The very small parameter estimates on the square of the number of children alive at every birth order (between 0.000 and 0.005) suggest that the correlation between this covariate and y^*_{ij} is close to linear. Finally and as expected, the mother's age at the birth of every child, shows a consistently negative correlation with the likelihood of that woman experiencing a further birth.

Figure 6 shows the parameter estimates based on equation 4 in which the parameter on the number of girls, γ_g, is allowed to vary with the birth of every child. The plot suggests that the correlation between the number of girls alive and fertility decisions varies considerably throughout a woman's reproductive life. The correlation shows an inversely U shaped pattern where the association between g_{ij} and y^*_{ij} increases for birth orders one to three from close to 0 to 0.09. After this birth order, the coefficient estimates begin to decrease reaching a very small estimate again for birth order eight. A possible explanation for these results is that the gender composition of the offspring matters most to parents as they approach their ideal family size, which lies between 2 and 4.

7.2 Dowries and Fertility Decisions

In line with predictions (i) and (iii), the decrease in expected dowry payments resulting from the introduction of the Dowry Prohibition Rules is found to lead to a one-off increase in fertility as well as to an attenuation in the previously observed correlation between the gender composition of a couple's offspring and its fertility choices.

7.2.1 Results based on the Gender of the Firstborn Child

Rows (a), (b) and (c) of table 4 report the coefficients on the difference in differences dummy variables. The estimates reported in row (a) suggest that fertility decreased for the control group over time; the figures in row (b) that before the introduction of the policy, families with a firstborn girl are more likely to give birth. The interaction between the post and treatment dummies is positive and significant. The decrease in expected dowries resulting from the policy change is thus estimated to have increased the chances of a further birth by between 2 and 4 percentage points, which supports prediction (i), i.e. the positive relation between the expected level of dowries to be paid and the probability of the couple opting for a further birth.

The figures reported in row (d) of table 4 suggest that before the policy change the control group exhibited a positive correlation between g_{ij} and y^*_{ij}. The parameter estimates for the number of girls alive range from 0.02 to 0.03. Rows (e) to (g) report the coefficients on the number of daughters alive (variable g_{ij} in equation 6) interacted with the difference in differences indicator variables. The estimates in row (e) indicate that, for the control group, the influence of the number of girls increased over time. The positive coefficients in row (f) point to this correlation in the pretreatment period to have been stronger for couples with a firstborn daughter. The magnitude of the difference is estimated between 0.04 and 0.06. This finding lends support to the claim that the gender of the first child adequately distinguishes treated from untreated individuals. Recall that treatment status depends on the amount of dowries paid before the introduction of the policy. If a sub-sample pays particularly high dowries before the change in the law, one would expect these individuals to exhibit a stronger correlation between g_{ij} and y^*_{ij} (see equation 3). The positive and significant coefficient estimates of row (f) show this to be the case for couples with a firstborn daugher.

The coefficient estimates on the triple interaction between the number of girls alive, g_{ij}, the post and the treatment dummies shown in row (g) report how the coefficient γ_g changed as a result of the introduction of the Dowry Prohibition Rules; this coefficient can be thought of as the difference in the differences in the slope parameter γ_g, i.e. the correlation between the number of girls alive and fertility rates. The estimates show that the correlation decreased by between 4 and 5 percentage points, which lends support to prediction (iii), i.e. that the positive correlation between the gender composition and fertility choices is increasing in expected future dowries. The coefficient estimates in rows (d) to (g) for the preferred specification in column (5) can be used to calculate that, for the control group, the coefficient γ_g increased from 0.027 to 0.055. For the treatment group, by contrast, it decreased from 0.085 to 0.068.

7.2.2 Results based on Ethnicity

Table 5 reports the difference in differences estimates employing the woman's ethnicity to distinguish treatment and control group. Rows (a) and (b) suggest that fertility increased for the control group over time but do not show a substantial difference between treatment and control before the policy. The parameter estimates in row (c) indicate that the policy increased the probability of a further birth by between 4 and 6 percentage points.

Row (d) confirms the positive correlation between g_{ij} and y^*_{ij} for the control group prior to the introduction of the Dowry Prohibition Rules. Row (e) suggests that this correlation decreased over time. The coefficient estimates in row (f) show that before the policy was introduced, the correlation between the number of girls alive at every birth order and the probability of a further birth was significantly larger for lower caste members. The estimates for this difference range from 0.01 to 0.27. As stated in the previous section, these estimates support the claim that treated and untreated individuals are adequately distinguished. The estimates in row (g) suggest that the policy decrease the coefficient γ_{ij} by between 1 and 2 percentage points, which - akin to above - supports prediction (iii). Post estimation calculations further show that, for the control group the coefficient γ_{ij} decreased from 0.103 to 0.052. For the treatment group the decrease was more pronounced, from 0.130 to 0.058.

7.2.3 Results based on older cohorts

Table 6 reports the estimates of equation 6 carried out for women born between 1942 and 1948 with a fictional policy in the year 1971. Row (a) shows that, for women with a firstborn daughter, the fictional policy did not change the intercept. The estimates of the five different specifications are small in size and not significantly different from zero. Row (b) confirms the positive correlation between the number of daughters alive and the probability of a further birth for the control group in the pre-treatment period. As row (c) shows, however, the fictional policy did not affect the slope parameter associated with this variable. The triple interaction terms are small and not significantly different from zero. The estimates using the woman's ethnicity shown in rows (d) to (f) paint a very similar picture. For this specification, however, the fictional policy of 1971 affected the intercept of the estimating equation by between 0.05 and 0.07. The slope of the coefficient γ_g, by contrast, remains unchanged by the placebo treatment.

7.3 Treatment Heterogeneity

Figure 7 reports the parameter estimates of an extended version of equation 6, which allows for the impact of the policy to vary with the birth order of the child. The schedule shows the changes in the slope parameter only and indicates that the strongest impact of the policy occurs at birth orders 2 and 3. For the control group, the coefficient γ_g hardly changed at birth orders 2 and 3. For the treatment group, by contrast, the parameter estimates dropped from 0.09 to 0.03 at birth order 2 and from 0.14 to 0.1 at birth order 3. Since the frequency of these birth orders is very high (see figure 6), this behaviour has a strong impact on the overall sample.

Table 7 reports how the impact of the policy change varies by the characteristics of the mother. Columns (1) to (4) show the impact of the Dowry Prohibition Rules distinguishing by the mother's region of residence. The parameter estimates suggest that the largest change in behaviour occurred in the West and North of India. In these two locations, the policy lead to an increase in the intercept of 0.057 and 0.042 respectively. The parameter γ_g further decreased by -0.053 in the West and -0.036 in the North. By contrast, in the South of India women hardly reacted to the policy. This last finding is in congruence with previous research highlighting the less widespread custom of dowries in the South. Columns (5) and (6), moreover, suggest a clear positive correlation between the impact of the Dowry Prohibition Rules and the mother's education. Although the change in the intercept is similar for women with and without completed primary education (0.042 and 0.04, respectively), the slope parameter for the former group decreased by -0.06 whereas for the latter it was only attenuated by -0.036. One possible explanation for this finding is that women with higher levels of education are likely to have higher levels of autonomy. The resulting improved agency is likely to enable these individuals to respond more effectively to the new circumstances by affecting decisions taken by the household as a whole.

In order to investigate the relative importance of female autonomy on the efficiency of the policy, this study employs two measurements previously argued to approximate this concept. Abadian (1996) argues that women who marry at an early age and have a large age difference to their husband show low female autonomy. The results reported in columns (7) to (10) suggest that more autonomous women responded more strongly to the Dowry Prohibition Rules. Women whose age at marriage lies above the sample average show larger intercept and slope changes, see columns (7) and (8). Similarly, women with age differences to their husbands that lie below the sample mean show the same pattern.

8 Robustness Checks

8.1 Placebo Treatments

To address the concern that the post 1985 dummy is correlated with changes in fertility behaviour unrelated to the influence of the Dowry Prohibition Rules a number of falsification tests are carried out. Columns (1) to (4) of Table 8 specify two different placebo treatments between the years 1976 and 1984 as well as between 1979 and 1984. In addition to these variables, the specification retains the post 1985 dummies and their interactions, which will test whether the post-treatment effect is robust to different reference time periods. Rows (a) and (b) show that, although the placebo treatment changes the intercept of the reduced form fertility equation, the shift appears negligible when compared results of the post 1985 dummies reported in rows (c) and (d); 0.016 and 0.021 versus 0.052 and 0.048 for the specification employing the gender of the firstborn and 0.024 and 0.036 versus 0.070 and 0.073 for the specification employing the mother's caste. Furthermore, rows (f) and (g) show that the placebo treatment did not change the slope parameter γ_g significantly. The parameter estimates are small in size and not statistically significant. The estimates for the post 1985 dummy in rows (h) and (i), by contrast, remain large, negative and significant; column (3) being the exception.

8.2 Sex Selective Abortions

A further point of concern is the presence of sex selective abortions. This practice gives parents the possibility of responding to changes in expected dowries by influencing the gender of each child rather than by changing their fertility patterns. As mentioned before, the extent of this practice is debated.[23] Figure 8 plots the percentage of girls born per birth year in the post period.[24] The upper panel divides the sample by the

[23]See section 6.2.1.

[24]Recall that the natural percentage of girls at birth is 48.8%.

gender of the firstborn child, the bottom panel by the woman's ethnicity.[25] In both panels, it is hard to detect any systematic differences in the sex ratio at birth between treatment and control group. Furthermore, in the upper panel the only year in which there is a statistical difference between the treatment and control group is 1994; for the bottom panel, only for the years 1989 and 1998. This suggests that neither the gender of the first born or the mother's caste introduce differences in sex selective abortions.

8.3 Different Post Treatment Periods

Children in the estimation sample are born between the years 1970 and 2001. This relatively long time window raises the concern that factors independent of the policy change influence the results. Columns (5) to (10) of table 8 report the difference in differences estimates employing different time windows for the post treatment period. The three time spans are 1985 to 1995, 1985 to 1990 and 1985 to 1986. The parameter estimates show that the estimation is robust to different definitions of the post-treatment window. In both specifications, the intercept shift remains positive and significant. Similarly, the change in the slope continues to be negative, significant and of similar magnitude.

These results can also be interpreted as evidence in favour of the claim that sex selective abortions are unlikely to bias the results. As mentioned by Bhalotra and Cochrane (2010) sex detection technology was scarce and expensive in the mid-1980s in India. The fact that the results remain stable to the employment of children born only in this decade suggests that parents adopted to the change in policy by adjusting their fertility behaviour rather than by employing sex selective abortions.

8.4 Common Time Trends

The causal interpretation of the difference in differences estimates relies on the assumption of orthogonality between the Dowry Prohibition Rules and the error term. Figure 9 explores the fertility behaviour of women born between the years 1942 and 1953 to give suggestive evidence in favour of the aforementioned assumption. These individuals were aged 32 to 43 at the implementation of the policy change and are thus assumed to only have been marginally affected by it. The negative slope of both curves highlights the well-documented fertility decline in India. The upper panel of figure 9

[25] For the upper hand panel only children of birth order 2 or above are considered.

opposes women with a firstborn son to mothers, whose firstborn is female. Although individuals in the latter group show higher fertility levels, the gender of the first child does not appear to change the rate of change in fertility behaviour significantly. The lower panel contrasts women from lower and higher castes. Akin to before, whilst the former show consistently higher fertility levels, the slope of both schedules appears similar.

9 Conclusion

The main results of this paper suggest that the widely documented correlation between a couple's gender composition and its fertility choices is, in part, a reflection of gender differences in the economic costs of children. Two ramifications of these findings appear of particular interest. First, the relative importance of child raising costs for parents' reproductive behaviour raises the question whether economic factors also influence other aspects of raising children. Whilst researchers are devoting increasing interest to inheritance rights or political representation of women, human development aspects such as nutrition, weight, height and other health outcomes have remained underexplored. Second, many previous explanations of the presence of son preferring stopping rules in fertility behaviour argued for these being the result of deeply rooted attitudes that somehow boys are more valuable than girls. The results put forward here, by contrast, argue that a large part of this behaviour can explained by the relatively simple economic intuition that sons are cheaper to raise than girls. Moreover, if dowries affect reproductive behaviour it stands to reason that other factors influencing the net cost of children - may it be the cost or returns - can potentially influence the same processes. This is an encouraging finding for practitioners because it can constitute a new set of instruments to influence fertility decisions taken by households. Furthermore, from a political perspective, dowries have been widely criticised for their negative influence on brides. This analysis highlights a further negative unintended consequence of this already widely criticised custom. As a consequence, any policy counteracting dowries will have a spillover effect on couple's reproductive behaviour thus further increasing the need to counteract this practice.

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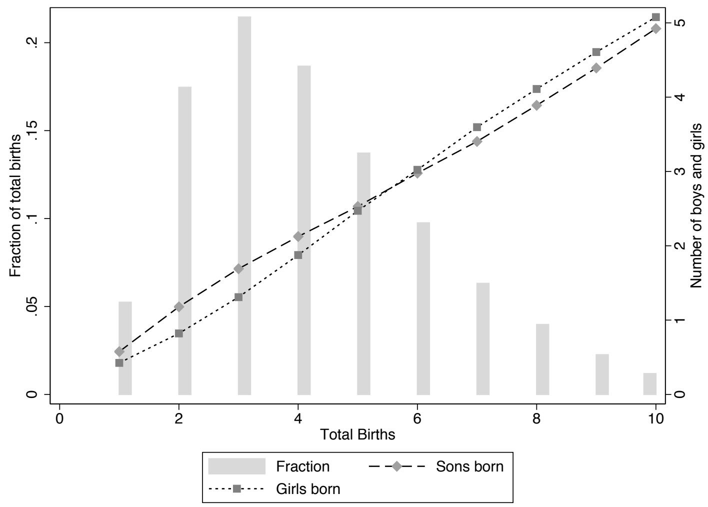
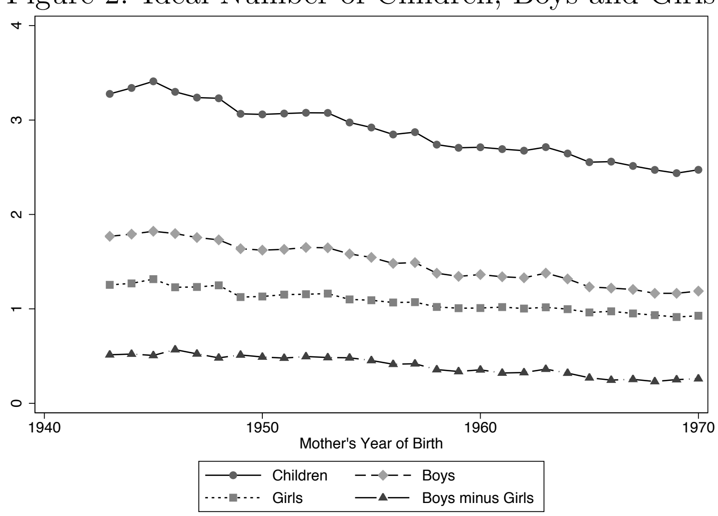
Figures

Figure 1: Mean Number of Births Per Woman



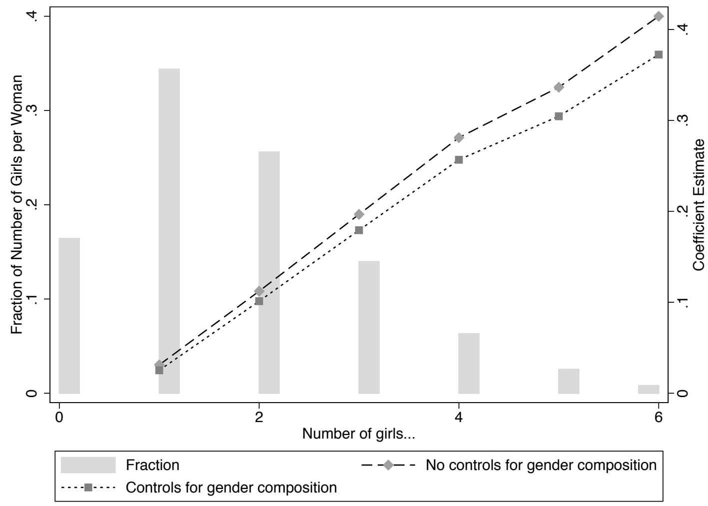
Notes: Sample consists of women born between 1942 and 1970, aged 36 to 49 years at interview and drawn from NFHS-1, NFHS-2 and NFHS-3; own calculations.

Figure 2: Ideal Number of Children, Boys and Girls



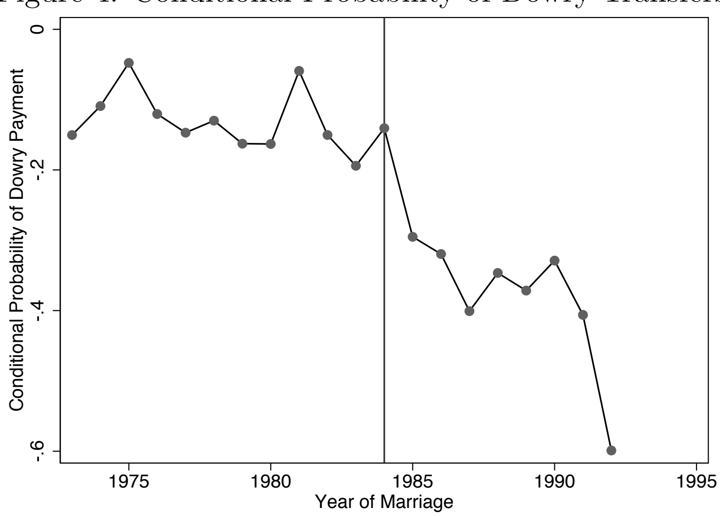
Notes: Sample consists of women born between 1942 and 1970, aged 36 to 49 years at interview and drawn from NFHS-1, NFHS-2 and NFHS-3; own calculations.

Figure 3: Coefficient Estimates for Number of Girls



Notes: Parameter estimates reported are from linear probability model; dependent variable takes value 1 if woman experiences another birth; estimations control for mother level fixed effect; sample consists of women born between 1953 and 1970, aged 36 to 49 years at interview and drawn from NFHS-1, NFHS-2 and NFHS-3; estimates are weighted using national weights.

Figure 4: Conditional Probability of Dowry Transfers

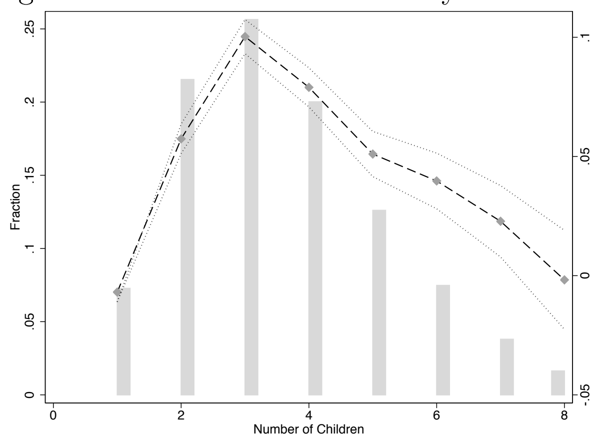


Notes: Parameter estimates reported are from linear probability model; dependent variable takes value 1 if woman paid dowry at her marriage; sample consists of women interviewed in 1993 drawn from the SWAF; bride level covariates include education, religion, caste; groom level covariates include education; parental level covariates include education of parents; base years 1969 to 1972; village level and year of birth fixed effects included.

Figure 5: Fraction of Mothers Intending to Pay Dowries in Future

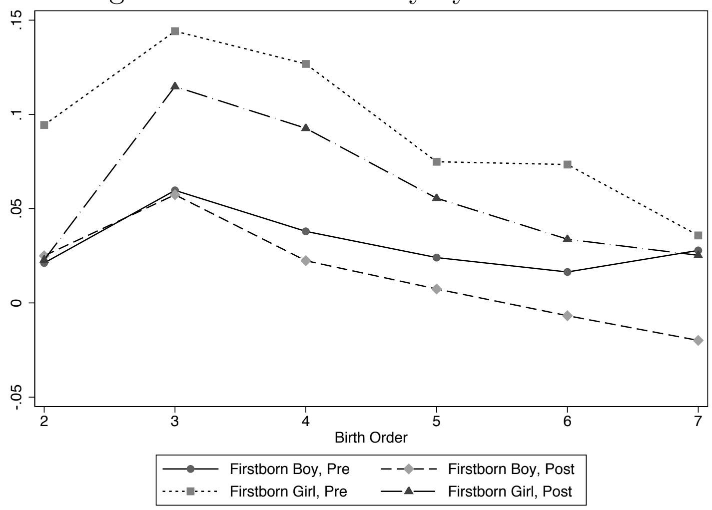
Notes: sample consists of women interviewed in 1993 drawn from the SWAF; own calculations.

Figure 6: Coefficient Estimates by Birth Order



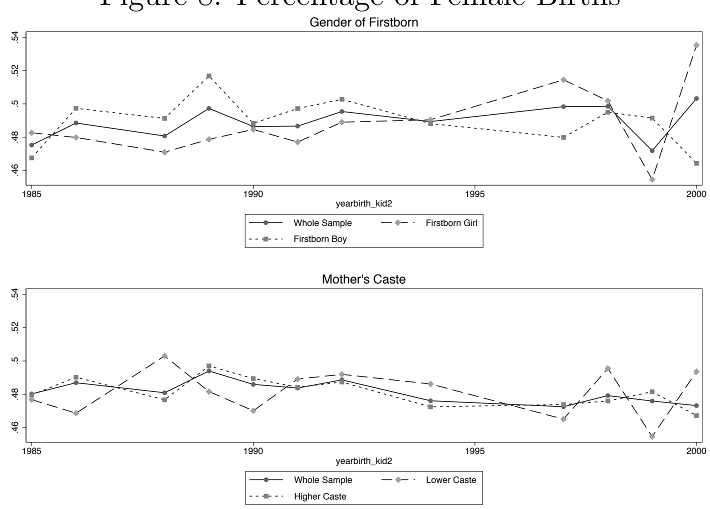
Notes: Parameter estimates reported are from linear probability model; dependent variable takes value 1 if woman experiences another birth; estimations control for mother level fixed effect; sample consists of women born between 1953 and 1970, aged 36 to 49 years at interview and drawn from NFHS-1, NFHS-2 and NFHS-3; estimates are weighted using national weights; 95% confidence interval reported as dashed line.

Figure 7: Effect of Policy by Birth Order



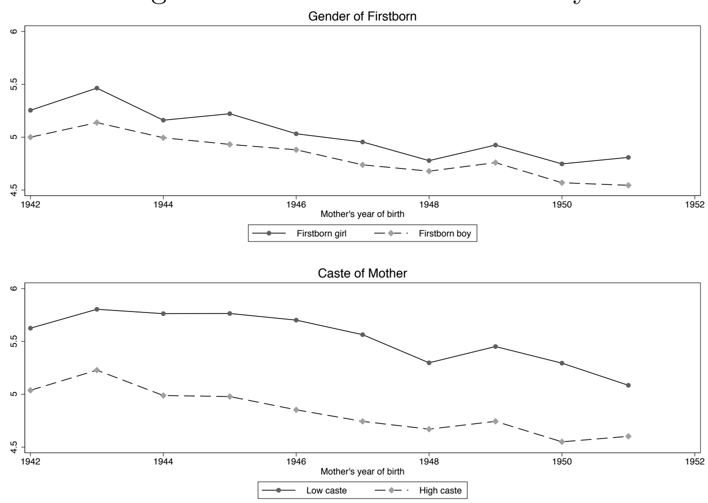
Notes: Parameter estimates reported are from linear probability model; dependent variable takes value 1 if woman experiences another birth; estimations control for mother level fixed effect; sample consists of women born between 1953 and 1970, aged 36 to 49 years at interview and drawn from NFHS-1, NFHS-2 and NFHS-3; estimates are weighted using national weights.

Figure 8: Percentage of Female Births



Notes: Variable reported is percentage of female births; sample consists of children born between 1985 and 2000 to women born between 1953 and 1970, aged 36 to 49 years at interview; drawn from NFHS-1, NFHS-2 and NFHS-3; own calculations.

Figure 9: Time Trends in Fertility



Notes: Variable reported is mean number of births experienced; sample consists of women born between 1942 and 1953, aged 36 to 49 years at interview; drawn from NFHS-1, NFHS-2 and NFHS-3; own calculations.

Tables

Table 1: Dowries paid in rural India 1975 - 1984

		(1) All	(2) Daughter Married	(3) No Daughter Married	(4) Low Caste	(5) High Caste
				Percentage		
(a)	**Household pays dowry**	80.9	86.5	73.5	84.0	80.2
				Average Rupees per year		
	Dowries paid by household					
(b)	**Per year**	1584	1736	1351	3091	840
(c)	**Per year, per person**	130	127	135	224	84
	Dowries received by household					
(d)	**Per year**	600	581	629	551	624
(e)	**Per year, per person**	63	52	81	39	76
	Net dowries of household					
(f)	**Per year**	984	1155	721	2539	216
(g)	**Per year, per person**	66	75	54	185	8
				Dowries as percentages of		
(h)	**Total household consumption**	57.8	72.4	33.3	179	4.8
(i)	**Total household income**	8.1	15.0	2.2	20.7	2.1
	Households	127	77	50	42	85

Note: Dowries reported in 1985 Rupee prices; net dowries defined as dowries paid minus dowries received by the household; sample consists of households observed between 1975 and 1984, source: ICRISAT; own calculations.

Table 2: Determinants of Gender Composition and Gender of Firstborn Child

		(1) Regression Coefficients	(2) Firstborn Boy	(3) Firstborn Girl	(4) Difference	(5) Regression Coefficients
		Percent of girls born	Percentage at birth			
(a)	**Estimation sample**	46.2	52.1	47.9		
(b)	**Natural rate**		51.2	48.8		
		Dependent variable: Ratio of girls born	Percentages and means		Ttest	Dependent variable: Firstborn is female
(c)	**Firstborn is female**	0.316***				-
		(0.014)				
(d)	**Mother has primary education**	-0.009***	50.0%	51.1%	**	0.007**
		(0.002)				(0.003)
(e)	**Mother is muslim**	0.012**	11.3%	10.9%		-0.009
		(0.005)				(0.005)
(f)	**Mother belongs to lower caste**	0.002	15.6%	15.6%		0.002
		(0.003)				(0.005)
(g)	**Age at first birth**	0.000	20.4 years	20.5 years	***	0.001**
		(0.000)				(0.001)
(h)	**Father has primary education**	0.004**	47.4%	47.5%		-0.000
		(0.002)				(0.004)
	State fixed effect	yes				yes
	Observations	66,245				66,245
	R-Squared	0.327				0.000

Notes: Parameter estimates reported are from ordinary least squares model; dependent variable in column (1) is girls as fraction of all children born; dependent variable in column (5) takes value 1 if firstborn child is female; sample consists of women born between 1953 and 1970, aged 36 to 49 years at interview and drawn from NFHS-1, NFHS-2 and NFHS-3; estimates are weighted using national weights; standard errors are reported in parentheses and are clustered at the mother level; ***, **, and * indicate significance at the 1%, 5% and 10% levels.

Table 3: Estimates of Reduced Form Fertility Decision

		(1)	(2)	(3)	(4)	(5)
		Dependent variable: Probability of further birth				
(a)	**Number of girls alive**	0.048***	0.047***	0.051***	0.051***	0.068***
		(0.001)	(0.001)	(0.001)	(0.001)	(0.002)
(b)	**Age at birth**		−0.020***	−0.019***	−0.019***	−0.035***
			(0.000)	(0.000)	(0.000)	(0.000)
(c)	**Total children alive**	−0.152***	−0.108***	−0.115***	−0.115***	−0.101***
		(0.001)	(0.001)	(0.001)	(0.001)	(0.002)
(d)	**Total children alive squrd**	0.005***	0.005***	0.004***	0.004***	−0.000
		(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
	Mother level controls	no	no	yes	yes	no
	Father level controls	no	no	no	yes	no
	Mother level fixed effect	no	no	no	no	yes
	Observations	355,203	355,203	347,830	347,573	355,203
	Mothers	72,247	72,247	70,733	70,683	72,247
	R-Squared	0.199	0.240	0.260	0.260	0.391

Notes: Parameter estimates reported are from linear probability model; dependent variable takes value 1 if woman experiences another birth; sample consists of women born between 1953 and 1970, aged 36 to 49 years at interview and drawn from NFHS-1, NFHS-2 and NFHS-3; mother level covariates include education, religion, caste and birth cohort; father level covariates include education; estimates are weighted using national weights; standard errors are reported in parentheses and are clustered at the mother level; ***, **, and * indicate significance at the 1%, 5% and 10% levels.

Table 4: Difference in Differences Estimates: Gender of First Born

		(1)	(2)	(3)	(4)	(5)
		Dependent variable: Probability of further birth				
(a)	**Post 1985**	−0.154***	−0.088***	−0.084***	−0.084***	−0.098***
		(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
(b)	**Firstborn is female**	0.010***	0.011***	0.013***	0.013***	−
		(0.001)	(0.002)	(0.002)	(0.002)	
(c)	**Firstborn is female * post 1985**	0.026***	0.025***	0.023***	0.023***	0.040***
		(0.007)	(0.007)	(0.007)	(0.007)	(0.008)
(d)	**Number of girls alive**	0.017***	0.021***	0.022***	0.022***	0.027***
		(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Interactions of number of girls alive with						
(e)	**Post 1985**	0.028***	0.025***	0.025***	0.025***	0.028***
		(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
(f)	**Firstborn is female**	0.044***	0.040***	0.043***	0.043***	0.056***
		(0.003)	(0.003)	(0.003)	(0.003)	(0.004)
(g)	**Firstborn is female * post 1985**	−0.039***	−0.037***	−0.036***	−0.036***	−0.045***
		(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
	Mother level controls	no	no	yes	yes	no
	Father level controls	no	no	no	yes	no
	Mother level fixed effect	no	no	no	no	yes
	Observations	250,339	250,339	246,346	246,092	250,339
	Mothers	47,518	47,518	46,810	46,761	47,518
	R-Squared	0.219	0.238	0.254	0.254	0.391

Notes: Parameter estimates reported are from linear probability model; dependent variable takes value 1 if woman experiences another birth; sample consists of women born between 1953 and 1970, aged 36 to 49 years at interview and drawn from NFHS-1, NFHS-2 and NFHS-3; mother level covariates include education, religion, caste and birth cohort; father level covariates include education; estimates are weighted using national weights; standard errors are reported in parentheses and are clustered at the mother level; ***, **, and * indicate significance at the 1%, 5% and 10% levels.

Table 5: Difference in Differences Estimates: Ethnicity

		(1)	(2)	(3)	(4)	(5)
		Dependent variable: Probability of further birth				
(a)	**Post 1985**	−0.084***	−0.011***	0.028***	0.027***	0.056***
		(0.002)	(0.002)	(0.003)	(0.003)	(0.004)
(b)	**Lower caste**	0.011***	−0.000	−0.003	−0.004 * *	
		(0.002)	(0.002)	(0.002)	(0.002)	–
(c)	**Lower caste * post 1985**	0.040***	0.038***	0.036***	0.037***	0.056***
		(0.005)	(0.005)	(0.005)	(0.005)	(0.009)
(d)	**Number of girls alive**	0.067***	0.064***	0.071***	0.071***	0.103***
		(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Interactions of number of girls alive with						
(e)	**Post 1985**	−0.025***	−0.024***	−0.030***	−0.030***	−0.051***
		(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
(f)	**Lower caste**	0.011***	0.011***	0.013***	0.013***	0.027***
		(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
(g)	**Lower caste * post 1985**	−0.020***	−0.017***	−0.014***	−0.014***	−0.021***
		(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
	Mother level controls	no	no	yes	yes	no
	Father level controls	no	no	no	yes	no
	Mother level fixed effect	no	no	no	no	yes
	Observations	348,165	348,165	347,830	347,573	348,165
	Mothers	70,803	70,803	70,733	70,683	70,803
	R-Squared	0.216	0.243	0.261	0.262	0.395

Notes: Parameter estimates reported are from linear probability model; dependent variable takes value 1 if woman experiences another birth; sample consists of women born between 1953 and 1970, aged 36 to 49 years at interview and drawn from NFHS-1, NFHS-2 and NFHS-3; mother level covariates include education, religion, caste and birth cohort; father level covariates include education; estimates are weighted using national weights; standard errors are reported in parentheses and are clustered at the mother level; ***, **, and * indicate significance at the 1%, 5% and 10% level.

Table 6: Difference in Differences Estimates: Older Cohorts

		(1)	(2)	(3)	(4)	(5)
		Dependent variable: Probability of further birth				
	Gender of firstborn					
(a)	**Firstborn is female * post85**	−0.014	−0.015	−0.014	−0.015	0.002
		(0.013)	(0.014)	(0.014)	(0.014)	(0.016)
(b)	**Number of girls alive**	0.022***	0.025***	0.026***	0.025***	0.039***
		(0.006)	(0.006)	(0.006)	(0.006)	(0.007)
	Interactions of number of girls alive with					
(c)	**Firstborn is female * post85**	−0.012	−0.007	−0.007	−0.008	−0.008
		(0.010)	(0.010)	(0.010)	(0.010)	(0.011)
	Ethnicity					
(d)	**Lower caste * post85**	0.052***	0.059***	0.059***	0.059***	0.074***
		(0.017)	(0.017)	(0.017)	(0.018)	(0.021)
(e)	**Number of girls alive**	0.041***	0.042***	0.043***	0.043***	0.056***
		(0.003)	(0.003)	(0.004)	(0.004)	(0.005)
	Interactions of number of girls alive with					
(f)	**Lower caste * post85**	−0.013	−0.015	−0.014	−0.014	−0.017
		(0.010)	(0.010)	(0.010)	(0.010)	(0.012)
	Mother level controls	no	no	yes	yes	no
	Father level controls	no	no	no	yes	no
	Mother level fixed effect	no	no	no	no	yes
	Observations	41,646	41,646	41,529	41,330	41,646
	Mothers	6,713	6,713	6,694	6,662	6,713
	R-Squared	0.212	0.234	0.240	0.240	0.372

Notes: Parameter estimates reported are from linear probability model; dependent variable takes value 1 if woman experiences another birth; sample consists of women born between 1942 and 1948, aged 36 to 49 years at interview and drawn from NFHS-1, NFHS-2 and NFHS-3; mother level covariates include education, religion, caste and birth cohort; father level covariates include education; estimates are weighted using national weights; standard errors are reported in parentheses and are clustered at the mother level; ***, **, and * indicate significance at the 1%, 5% and 10% levels.

Table 7: Treatment Heterogeneity

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	North	South	East	West	Primary education	No Primary education	Above av. age at marriage	Below av. age at marriage	Below av. age difference	Above av. age difference
	Dependent variable: Probability of further birth									
(a) **Post 1985 * firstborn is female**	0.042***	0.025	−0.007	0.057***	0.042***	0.040***	0.046***	0.037***	0.045***	0.035***
	(0.012)	(0.019)	(0.020)	(0.015)	(0.014)	(0.009)	(0.012)	(0.010)	(0.011)	(0.011)
(b) **Number of girls alive**	0.051***	−0.007	0.030***	0.042***	0.019***	0.038***	0.002	0.038***	0.030***	0.025***
	(0.005)	(0.009)	(0.009)	(0.006)	(0.007)	(0.004)	(0.006)	(0.004)	(0.005)	(0.005)
Interactions of Number girls alive with										
(c) **Post 1985 * firstborn is female**	−0.036***	−0.029 * *	−0.039***	−0.053***	−0.060***	−0.036***	−0.073***	−0.033***	−0.051***	−0.040***
	(0.008)	(0.013)	(0.013)	(0.010)	(0.010)	(0.006)	(0.009)	(0.006)	(0.007)	(0.007)
Mother level controls	no	no	no	no	no	no	no	no	no	no
Father level controls	no	no	no	no	no	no	no	no	no	no
Mother level fixed effect	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
Observations	68,814	41,880	46,561	57,929	98,295	151,940	110,224	140,115	118,824	131,515
Mothers	12,012	9,186	8,735	10,829	21,360	26,137	22,863	24,655	22,233	25,285
R-Squared	0.384	0.453	0.400	0.406	0.456	0.383	0.420	0.385	0.395	0.389

Notes: Parameter estimates reported are from linear probability model; dependent variable takes value 1 if woman experiences another birth; sample consists of women born between 1953 and 1970, aged 36 to 49 years at interview and drawn from NFHS-1, NFHS-2 and NFHS-3; mother level covariates include education, religion, caste and birth cohort; father level covariates include education; estimates are weighted using national weights; standard errors are reported in parentheses and are clustered at the mother level; ***, **, and * indicate significance at the 1%, 5% and 10% levels.

Table 8: Robustness Checks

		(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
		Placebo: 1976-84	Placebo: 1979-84	Placebo: 1976-84	Placebo: 1979-84	Treatment: 1985-2000	Treatment: 1985-95	Treatment: 1985-86	Treatment: 1985-2000	Treatment: 1985-95	Treatment: 1985-86
		Dependent variable: Probability of further birth									
(a)	**Placebo * firstborn is female**	0.016***	0.021***								
		(0.003)	(0.004)								
(b)	**Placebo * lower caste**			0.024***	0.036***						
				(0.005)	(0.006)						
(c)	**Treatment * firstborn is female**	0.052***	0.048***			0.038***	0.038***	0.050***			
		(0.008)	(0.008)			(0.009)	(0.009)	(0.011)			
(d)	**Treatment * lower caste**			0.070***	0.073***				0.049***	0.060***	0.083***
				(0.009)	(0.009)				(0.009)	(0.010)	(0.014)
(e)	**Number of girls alive**	0.079***	0.066***	0.121***	0.133***	0.024***	0.022***	0.028***	0.091***	0.072***	0.061***
		(0.008)	(0.005)	(0.004)	(0.003)	(0.003)	(0.004)	(0.004)	(0.002)	(0.002)	(0.002)
		Interactions of number girls alive with:									
(f)	**Placebo * firstborn is female**	−0.015	−0.009								
		(0.012)	(0.008)								
(g)	**Placebo * lower caste**			0.006	−0.010*						
				(0.008)	(0.006)						
(h)	**Treatment * firstborn is female**	−0.057***	−0.050***			−0.042***	−0.036***	−0.032***			
		(0.012)	(0.008)			(0.005)	(0.006)	(0.008)			
(i)	**Treatment * lower caste**			−0.006**	−0.018***				−0.017***	−0.022***	−0.035***
				(0.003)	(0.006)				(0.005)	(0.006)	(0.008)
	Observations	250,339	250,339	348,165	348,165	245,222	228,695	198,924	324,543	277,341	215,073
	Mothers	47,518	47,518	70,803	70,803	47,518	47,518	47,518	69,270	64,504	55,087
	R-Squared	0.392	0.392	0.399	0.399	0.361	0.318	0.272	0.348	0.308	0.270

Notes: Parameter estimates reported are from linear probability model; dependent variable takes value 1 if woman experiences another birth; sample consists of women born between 1953 and 1970, aged 36 to 49 years at interview and drawn from NFHS-1, NFHS-2 and NFHS-3; mother level covariates include education, religion, caste and birth cohort; father level covariates include education; estimates are weighted using national weights; standard errors are reported in parentheses and are clustered at the mother level; ***, **, and * indicate significance at the 1%, 5% and 10% levels.