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

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

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## Abstract

This paper investigates the effect of the differential pecuniary costs of sons and daughters on fertility decisions. The focus is on dowries in India, which increase the economic returns to sons and decrease the returns to daughters. The paper exploits an exogenous shift in the cost of girls relative to boys arising from a revision in anti-dowry law. The reform is found to have attenuated the widely documented positive correlation between daughters and their parents' fertility. The observed patterns can be explained by a simple model of sequential fertility decisions where the gender composition of children determines future dowry payments.

**JEL Classifications:** O15, J12, J13

**Keywords:** Dowry, Fertility, India, Son Preferences

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# 1 Introduction

India is characterised by gender gaps in many human development indicators. Young girls, in particular, remain the most disadvantaged (The World Bank, 2012) and face significant disparities in a number of welfare measures such as mortality (Bhargava, 2003; Arnold et al., 2002a), nutrition (Jayachandran and Kuziemko, 2011; Oster, 2009), abortions (Bhalotra and Cochrane, 2010) and child care (Barcellos et al., 2012). One prominent explanation amongst economists for such disparities is the fact that parents invest more heavily in boys than in girls (see Jensen, 2012; Qian, 2008; for recent examples).

Parental fertility choices are one of the first margins of investment in children and there is evidence that the gender of children affects their parents' reproductive choices (Dahl and Moretti, 2008; Angrist and Evans, 1998). Throughout Asia, families have traditionally exhibited son preferences, defined as the belief that sons are more valuable than daughters (Das Gupta et al., 2003; Clark, 2000). After the birth of a daughter, parents have an incentive to give birth to another child in the hope of giving birth to a son (Arokiasamy, 2002). Such son-preferring gender biased stopping rules (Yamaguchi, 1989) have been shown to increase overall fertility (Seidl, 1995; Dreze and Murthi, 1999) and decrease young girls' welfare (Jensen, 2003). This reproductive behaviour may, in part, be explained by the future income streams associated with sons and daughters.<sup>1</sup>

This paper explores whether changes in the economic costs of daughters vis-à-vis sons affect their parents' fertility decisions. For most costs related to children, the gender specific component can be hard to determine. The paper addresses this measurement problem by focusing on a widespread custom 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). Because of these payments, 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.<sup>2</sup> As a consequence the overall cost of children depends on their total number as well as on their gender composition. Parents are likely to internalise this association and there is, in fact, qualitative evidence that dowries affect fertility choices (Diamond-Smith et al., 2008).

The paper estimates the likelihood that a woman gives birth at a given birth order<sup>3</sup> as a function of individual controls and the gender composition of children. The model is estimated using retrospective birth histories of women drawn from three rounds of the National Family Health Survey (NFHS, 1994; 1999; 2007b). Each girl is associated with an increase in the probability of giving birth of 2 percentage points, which is found to be

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<sup>1</sup>Jensen (2010) investigates the importance of future income streams for human capital investments.

<sup>2</sup>In India, parents can capture (at least) part of their son's dowry.

<sup>3</sup>Defined as the rank of siblings by age. For example: the second child is at birth order 2.

comparable to a decrease of 1.5 years in maternal education. The effect is particularly strong for third- and fourth-born children.

To estimate the effect of dowries on fertility decisions, the paper exploits a substantial revision in anti-dowry law. The Dowry Prohibition Rules (1985) tightened the monitoring of dowries and increased the penalties for offenders. The resulting decrease in expected dowries translates into an exogenous change in the costs of girls relative to boys. The specification evaluates this policy change in a difference-in-differences framework, which compares children in the same household born before and after the reform. The identification strategy uses the fact that exposure to the Dowry Prohibition Rules varied with the relative number of boys and girls in the family at the implementation of the reform. For parents with more daughters than sons dowries to be paid exceed dowries to be received. These families are thus more likely to be affected by the policy change. The gender composition at the reform is approximated by the gender of the firstborn child, which is argued to be exogenous. Families with a firstborn daughter are the treatment and households with a firstborn son the control group. To strengthen the specification further, the paper also exploits variation in pre-1985 dowry payments across matri- and patrilineal states as well as the mother's caste.

The introduction of the Dowry Prohibition Rules is estimated to have attenuated the positive correlation between the number of girls and fertility by 4.5 percentage points. For the treated, the policy decreased the influence of the gender composition on fertility by 50 percent. The effect appears particularly strong for children of lower birth orders and for more educated and autonomous women. The paper also estimates the relative importance of various confounding factors and of sex selective abortions. Neither appears to be driving the results.

The empirical results can be explained by a simple model of fertility and marriage choice. Parents have children sequentially. After every birth, they decide whether or not to have another child. The cost of each child depends in part on its gender; at the marriage of a son, parents receive a dowry and at the marriage of a daughter they pay a dowry. In the face of these income streams, parents have an incentive to marry sons before daughters. One possible explanation for postponing the marriages of girls is that parents employ the resources received from their daughter-in-law to fund their own daughter's dowry. Consequently, the expected benefits of a daughter are independent of the gender composition of her older siblings. The expected value of a male birth, by contrast, increases with the number of daughters in the family. The paper derives an optimal stopping rule with the following testable implications: (i) conditional on birth order, there exists a positive correlation between the number of daughters and the probability of parents having another child; and (ii) the correlation mentioned in point (i) depends positively on the expected value of the dowry.

This study aims to add to the growing knowledge base on dowries and marriage institutions. Both the economic rationale behind dowries (Ambrus et al., 2010; Bloch et al., 2004; Botticini and Siow, 2003) and the effect of marital institutions on women (Brown, 2009; Bloch and Rao, 2002) have received growing attention. Little is known, however, on how these practices influence reproductive choices. By relating dowry payments to fertility decisions, this paper puts forward an explanation for son preferring stopping behaviour, which does not rely on parental preferences regarding a child's gender.

The remainder of the paper is structured as follows: Section 2 lays out a theoretical framework of the effect of dowries on fertility choices and derives its predictions. Section 3 introduces the data and gives motivating descriptive evidence. Section 4 explains the practice of dowries and the legal framework. Section 5 lays out the empirical strategy the results of which are discussed in section 6. Section 7 addresses empirical concerns and section 8 concludes.

## 2 Fertility as a Sequential Stopping Decision

To help conceptualise why and how dowries may impact fertility decisions, this paper sets up a simple model of fertility choices. Parents maximise returns to their children by deciding how many children to have and by determining the marriage order of their children. This model assumes that all children marry and that parents are in a position to determine the order in which their children marry. The empirical validity of both assumptions is discussed in section 3.3.

### 2.1 Probabilistic Mechanism and Payoff Structure

Parents have children repeatedly. After every child, parents have the option to stop or to continue to have another child. Parents must stop after a finite number of children and if they stop after child  $n$  they receive reward  $R_n$ . This reward consists of one random and two deterministic variables. The set up is similar to the repeated coin-tossing game analysed by Chow and Robbins (1965) among many others.

Parents draw utility from the total number of children,  $n$ , independent of their gender. The utility function  $U(n)$  is concave. Also, for each child, parents incur a cost,  $p$ . This consists of all child raising costs, such as, for instance, food or educational expenses, minus of any returns the parents receive from their children - from for example household production. This component is constant, therefore, the total cost of  $n$  children is  $np$ . The utility and cost of children are deterministic.

The random reward sequence consists of the sum of dowry payments,  $\sum_{j=1}^n s_j \delta^j d_e$  where  $j$  is the ranking of children's marriages and  $s_j$  is independently and identically distributed with common probability distribution

$$P(s_j = 1) = P(s_j = -1) = \frac{1}{2}$$

So,  $s_j = 1$  if the  $j^{\text{th}}$  child getting married is a boy and  $s_j = -1$  if it is a girl. Further,  $d_e$  is the dowry parents expect to be transferred at their child's marriage. Dowry payments are discounted by  $\delta^j$ .

After the parents stop childbearing, they decide on the order in which their children marry. This ranking determines the discount factor associated with each dowry. Because  $0 < \delta < 1$  it follows that  $\delta^1 < \delta^2 < \dots < \delta^n$ . Parents choose the optimal ranking strategy,  $\sigma$ , to maximise the net value of all dowry transfers.

The term  $\delta$  does not necessarily denote a discount factor in the strict sense. Rather, this parameter reflects the more general fact that the timing of children's marriages matters to their parents. One possible reason for this is the presence of credit constraints. If parents only have limited access to credit, they can use the resources transferred by their sons's bride to fund their own daughter's dowry. Thus, parents influence the order of marriage not because they value earlier transfers more. Instead, it is the most cost effective strategy to finance their own daughter's marriage. To keep the model simple and tractable, these considerations are only implicitly incorporated in the discount factor  $\delta$ .

The total number of children is thus determined by the stopping rule that maximises the expected revenue of births conditional on the couple's optimal ordering of marriages ( $\sigma$ )

$$\rho = \sup_{n \in \mathbb{N}} E[R_n | \sigma]$$

## 2.2 Optimal Stopping Rule

The optimal stopping rule is determined by backward induction. Consider the marriage order first. The condition  $0 < \delta < 1$  implies that parents have an incentive to marry sons before daughters. More formally, if  $b$  is the number of boys and  $g$  the number of girls in family  $i$  then  $\sigma$  is a permutation on 1 to  $n$  such that  $\sigma_j = k$  if the  $k^{\text{th}}$  child is the  $j^{\text{th}}$  boy

$$\sigma_j = \begin{cases} k & \text{if } 1 \leq j \leq b \text{ and } k^{\text{th}} \text{ child is } j^{\text{th}} \text{ boy} \\ k & \text{if } b + 1 \leq j \leq b + g \text{ and } k^{\text{th}} \text{ is the } (j - b)^{\text{th}} \text{ girl} \end{cases}$$

The optimal stopping rule for reproductive choices is that parents have another child as long as the expected benefits of doing so exceed the expected costs. For parents with  $n$  children  $\rho$  is

$$U'(n) - p + E_D \geq 0 \quad (1)$$

where the utility from another child is  $U'(n) = U(n+1) - U(n)$ ,  $p$  the marginal cost and  $E_D$  the expected change in dowry transfers resulting from another child. The latter term is the sum of expected dowry payments after  $n+1$  births minus dowry payments after  $n$  births. Hence,  $E_D = E(\sum_{j=1}^{n+1} s_j \delta^j d_e) - \sum_{j=1}^n s_j \delta^j d_e$ .

To formalise  $E_D$ , consider family  $i$  with  $b$  boys and  $g$  girls. For this family net dowry payments are

$$\sum_{j=1}^b \delta^j d_e - \sum_{j=b+1}^{b+g} \delta^j d_e \quad (2)$$

The expected dowry of another birth is

$$\begin{aligned} & \frac{1}{2} \left( \sum_{j=1}^{b+1} \delta^j d_e - \sum_{j=b+2}^{b+g+1} \delta^j d_e \right) + \frac{1}{2} \left( \sum_{j=1}^b \delta^j d_e - \sum_{j=b+1}^{b+g+1} \delta^j d_e \right) \\ & = \sum_{j=1}^b \delta^j d_e - \sum_{j=b+2}^{b+g+1} \delta^j d_e \end{aligned} \quad (3)$$

The expected change in dowry,  $E_D$ , is the difference between equations 3 and 2, which simplifies to

$$E_D = d_e \delta^{b+1} (1 - \delta^g) \quad (4)$$

Thus, reformulating equation 1, parents will have another birth if the following inequality holds

$$U'(n) - p + d_e \delta^{b+1} (1 - \delta^g) \geq 0 \quad (5)$$

which clearly shows that the expected benefits from another birth depend on the gender composition of children alive. Note that with zero dowry payments ( $d_e = 0$ ), equation 5 simplifies to the standard marginal equality condition postulated by the Becker-Lewis (1973) model.

Intuitively, the model presented here can be seen as follows. The expected revenue of a daughter does not depend on the gender composition of children in the family. Optimising



parents marry the youngest daughter after all other children. The expected revenue of a son, by contrast, increases with the number of girls already in the family. Families with many daughters face higher dowry payments (net of dowry receipts) in the future. These families would benefit disproportionately from the resources received as the result of marrying a son.

### 2.3 Testable Implications

The framework laid out in sections 2.1 and 2.2 gives rise to two implications regarding the relation between dowries and fertility decisions: (i) at every birth order there exists a positive correlation between the number of daughters in the family and the propensity to have another child. Consider equation 5, since  $0 < \delta < 1$ , as  $g$  increases  $(1 - \delta^g)$  increases and so does  $E_D$ . Thus, the higher the number of girls in the family (conditional on the total number of children), the higher the expected benefits from another birth. This leads to higher birth rates. The positive correlation between female offspring and fertility rates has been widely documented (see Das Gupta et al., 2003; for instance). The second implication is that: (ii) the correlation mentioned in point (i) is increasing in the expected dowries,  $d_e$ . This can be seen from the functional form of  $E_D$  where  $d_e$  enters multiplicatively.

To illustrate implications (i) and (ii), consider two families with  $k$  children. Family A with  $b_A$  boys and  $g_A$  girls and family B with  $b_B$  boys and  $g_B$  girls.<sup>4</sup> Without loss of generality assume  $g_A > g_B$ . Because  $k$  is held constant, the expected future dowry transfers of another birth are respectively

$$E_D(A) = d_e \delta^{k+1} (\delta^{g_A} - 1) \quad (6)$$

$$E_D(B) = d_e \delta^{k+1} (\delta^{g_B} - 1) \quad (7)$$

subtracting equation 6 from 7 yields  $d_e \delta^{k+1} (\delta^{-g_A} - \delta^{-g_B})$ . Since the assumption was that  $g_A > g_B$ , then  $\delta^{-g_A} < \delta^{-g_B}$  which implies that  $E_D(A) > E_D(B)$ . This confirms (i). Furthermore the difference between 6 from 7 becomes larger for increasing values of  $d_e$ , therefore, (ii) follows.

## 3 Data and Summary Statistics

This paper is motivated by two facts: women report a higher ideal number of sons than daughters and families with more girls than boys have above average fertility rates.

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<sup>4</sup>Because the number of children is equal, the marginal utility and price of the next birth is the same for both families. These two terms are hence omitted from this illustration.

### 3.1 The Data

This study 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.<sup>5</sup> The questionnaires collect extensive information on health, nutrition and the complete birth histories of interviewed women. The NFHS-1 (IIPS 1994) was carried out in 1992 and 1993 and interviewed 89,777 ever-married women aged 13 to 49; the NFHS-2 (IIPS 1999) was conducted in 1998 and 1999 and interviewed 89,199 ever married women aged 15 to 49; finally the NFHS-3 (IIPS 2007b) was implemented in 2005 and 2006 and interviewed 124,385 women aged 15 to 49.

Individuals selected for estimation 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. The relatively low age cut off used here is chosen to keep sample sizes large enough for the sub-group analysis carried out in section 6.2.2.<sup>6</sup> Similarly, the omission of childless women is unlikely to bias the results significantly, only 4 percent of women 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, who were born between the years 1942 and 1970. Women in the 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.

### 3.2 The Sample and Summary Statistics

Individuals in the sample have given birth to, on average, 4.1 children. The mean age at first birth is around 20 years. The mode of the distribution lies at 3 children. Most women have between 2 and 5 children, the percentages of mothers with 2, 3, 4 and 5 children are 18, 22, 19 and 14 percent respectively. The gender composition of offspring is relatively close to the natural rate; 47.8 percent of children born to women are female.<sup>7</sup>

The NFHS elicits retrospective questions on respondents' ideal number of children, sons

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<sup>5</sup>The data are publicly available at [measuredhs.com](http://measuredhs.com)

<sup>6</sup>Restricting the sample to women aged 40 and above does not alter the results significantly.

<sup>7</sup>The natural rate of girls born at birth is 48.8 percent.

and daughters.<sup>8</sup> Individuals in the sample report an ideal number of 2.8 children, 1.4 sons and 1.1 daughters.<sup>9</sup> Figure 1 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 cohorts, the ideal number of sons exceeds the ones of daughters by around 0.4. This difference appears remarkably stable over time. Whilst the ideal number of children decreases from 3.3. to 2.5, the difference between ideal sons and daughters goes down relatively little, from 0.5 to 0.3.

From an intuitive point of view, the effect of preferences for sons on fertility is ambiguous. If son preferences manifest themselves as an aversion of girls (Diamond-Smith et al., 2008), families with more daughters decrease fertility to avoid another female birth. If, on the other hand, son preferences are reflected as a desire to have at least as many boys as girls in the household, parents with many daughters increase their fertility in the hope of male births. Descriptive evidence from the NFHS supports the latter. Figure 2 reports the correlations between the gender compositions of children and their parents' fertility rates. The base category are parents with an equal amount of sons and daughters.<sup>10</sup> Compared to these individuals, families with more daughters have higher fertility levels. The conditional probability of another birth increases by 10 percentage points for families with only girls and by 3 percentage points for families with more girls than boys. The correlation between the number of boys and fertility, by contrast, is ambiguous. Families with only sons have a 6 percentage point higher probability to have another child. Parents with more boys than girls, on the other hand, are 7 percentage points less likely to have another child.

### 3.3 Marriages in India

The theoretical model makes two assumptions about marital behaviour in India. First, all children marry. Marriage is central to Indian social life and holds value in both Hinduism and Islam. Evidence from the NFHS (1994) further confirms that marriage is virtually universal in India. Only 1 percent of women aged 35 to 39 remain unmarried.

Second, parents are in a position to determine the order in which their children marry and they attempt to marry sons before daughters. The assumption that marriage patterns of family members are determined by the household as a whole is not uncommon (see Rosenzweig and Stark, 1989; for instance). In fact, around the time of the introduction of the Dowry

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<sup>8</sup>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?*

<sup>9</sup>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.

<sup>10</sup>This is modelled in a regression framework where the dependent variable takes the value 1 if the woman experiences another birth. The controls include the total number of children born, a time trend and mother level unobserved heterogeneity.

Prohibition Rules, most marriages in India were arranged (Dixon, 1971). The short birth intervals commonly observed in India further facilitate the marriage of sons before daughters. Figures from the NFHS-1 (1994) show that the median birth interval is two years. This is comparatively short compared to the standard deviation in the age at marriage, which is 3.5 years.

As pointed out by Vogl (2013), parents attempt to marry their children in *gender specific* birth order, which implies that marriages of girls (or boys) are influenced predominantly by the number of their older sisters (or brothers). Figure 3 shows the share of children married by age and gender of next youngest sibling. India is a virilocal society in which, after marriage, the bride moves to the groom and his family. Thus, the marriage of a woman are approximated by her leaving the household whereas the marriage of a man is self reported. The two figures suggest that women with younger sisters tend to marry earlier whereas men are less affected by the gender of their younger sibling. These patterns are compatible with parental incentives to marry sons before daughters.

To give more detail on marriage behaviour in India, this paper follows the methodology by Vogl (2013) and estimates the probability of a child marrying as a function of individual controls (like age<sup>11</sup>) and the gender of younger siblings. Columns (1) and (2) of table 1 employ the gender of the next child only and suggest that the birth of a younger sister (as compared to the birth of a younger brother) increases the probability of getting married by 3 percentage points for women and decreases the probability by 1 percentage points for men. Columns (3) to (6) control for the total number of younger siblings and include the number of younger sisters and brothers separately. The parameter estimates show that younger sisters increase and younger brothers decrease the marriage probability for women. The opposite is true for men. Like before, these parameter estimates are compatible with parental behaviour that endeavours to marry sons before daughters.

## 4 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.

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<sup>11</sup>The controls include age, birth order and state of residence fixed effects and parental education, caste and religion.

## 4.1 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. Do et al. (2013) and Tertilt (2005) consider the importance of marriage patterns; Roy (2011); Dalmia (2004); Deolalikar and Rao (1998) focus on the characteristics of the groom, Caldwell et al. (1983) on the ones of the bride. There is no consensus, however, on the precise monetary value of these transfers. Rao (1993, 2000) argues that dowries amount to up to 68 percent of assets before marriage. Other research, by contrast, has put forward much lower figures (Arunachalam and Logan, 2008; Anderson, 2007b; Edlund, 2000). In particular, Edlund (2006) distinguishes gross from net dowry payments and argues that the increase in net dowries has been negligible.

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). The survey is carried out in 6 different villages in the Indian state of Andhra Pradesh between the years 1975 and 1984. This survey contains self-reported information on, inter alia, age, marital status of all household members and inventory files for current physical stocks as well as on financial assets and liabilities such as bank accounts and dowries. Table 2 reports average dowries paid and received by households using ICRISAT data. The majority of households report to have either paid or received a dowry, 81 percent in column (1). Rows (b) and (c) report the dowries paid, (d) and (e) dowries received and (f) and (g) the net out payments per household.<sup>12</sup> 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.<sup>13</sup> A possible reason for this disparity that parents can only capture part of the dowry given to their sons. Dowries make up more than half of the household's non-durable consumption expenditures, see row (h), and 8 percent of the income of all household members combined see row (i).

## 4.2 Dowry Laws 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 (Dowry Prohibition Act, 1961).<sup>14</sup> Despite this legislation few dowry cases reached the courts and the practice of

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<sup>12</sup>These are defined as the dowries paid minus the dowries received by the household.

<sup>13</sup>All numbers are given in 1984 Rupees. Official GDP per capita in 1980 was Rupees 1,630.

<sup>14</sup>Dowry is 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

dowries persisted (see evidence in section 4.1). A common reason put forward for this is that the 1961 Act's provisions were not strong enough to implement successful prosecutions (Chowdhary, 1998).

In response to this, the government of India introduced the Dowry Prohibition Rules (1985),<sup>15</sup> 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 kept by the bride whereas the list containing presents to the groom is kept with the groom. These lists must be in writing and contain the approximate value of the present. 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.

To facilitate the implementation of the newly established rules, the amended act introduced the Dowry Prohibition Officers. The tasks of these public sector employees included the prevention of the taking and demanding of dowries and the collection of evidence necessary for the prosecution of persons committing offences under the Dowry Prohibition Act.<sup>16</sup> Legal research has pointed to a marked increase in dowry cases heard by courts in the mid 1980s (Menksi, 1998). Furthermore, India's high court took a much stricter approach to dowry offenders. Overall, the new rules were perceived by many as countering the prevalent attitude of patriarchal traditions that women were owned by men.

### 4.3 The Impact of the Dowry Prohibition Rules

Evidence from the Survey of Status of Women (2000) suggests that the Dowry Prohibition Rules had a marked effect on dowry transfers. The SWAF is part of a series implemented in India, Malaysia, Pakistan, the Philippines and Thailand. The questionnaire was carried out in 1993 and 1994 in four districts in the states of Tamil Nadu and Uttar Pradesh. It interviewed 1,600 women in total. The survey collected information on current health and different dimensions of female autonomy as well as retrospective information on marriages and dowry transfers.<sup>17</sup>

Figure 4 reports the differences in the conditional probability of a dowry being paid at 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"

<sup>15</sup>Amendment Act 63 of 1984 came into force on the 2.10.1985.

<sup>16</sup>Section 8B of the amended Dowry Prohibition Act.

<sup>17</sup>The data are publicly available at [swap.pop.upenn.edu/datasets](http://swap.pop.upenn.edu/datasets)

a woman's marriage between the years 1970 and 1994.<sup>18</sup> Whilst the conditional probability estimates of dowry payments in the years leading up to the policy are very similar, around 12% to 16% lower than the base years, the estimates drop to -30% in 1985. In subsequent years, the probability estimates oscillate between -30% and -40% before eventually dropping to -60% for the last year, 1992. This final drop may have been caused by the gradual introduction of the Dowry Prohibition Officers. As mentioned, the 1985 law stipulated the creation of these new public sector employees and it is likely that their impact was delayed. It is unlikely that this drop is the result of respondents misreporting their dowry transfers. As previously mentioned, the questions about dowries were asked to all women retrospectively in 1993 and 1994. Hence, whilst respondents may have had an incentive to misreport in general, it is unlikely that this incentive varied according to whether they married before or after the policy change.

Further descriptive evidence from the SWAF suggests that the change in policy affected parents' expectations of a dowry being paid. One question concerns mothers and the dowries they expect to pay. Figure 5 shows the percentage of mothers expecting to pay a dowry by the year of the last born child. The percentage of women intending to pay a dowry increases for birth years before the introduction of the policy. These individuals are more likely to have heard of the policy. For children born after the policy, by contrast, the fraction decreases. However, since the sample includes women that have not yet completed their childbearing years, this trend may reflect inexperience regarding dowry issues. Hence, the evidence is only to be seen as suggestive.

## 5 Empirical Strategy

This paper exploits the exogenous decrease in expected dowries resulting from the Dowry Prohibition Rules outlined above to test implications (i) and (ii) empirically.

### 5.1 Empirical Framework

This paper investigates the probability that a woman gives birth after a given number of children. This probability is estimated as a function of individual controls and the gender composition of her offspring. The study constructs complete, retrospective birth histories. Thus, 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

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<sup>18</sup>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.

for her entering motherhood. The following linear probability model is estimated:

$$y_{ij} = \gamma_g g_{ij} + x'_{ij}\beta + \alpha_s + \nu_{ij} + u_{ij} \quad (8)$$

where  $y_{ij}$  is equal to 1 if woman  $i$  at birth order  $j$  gives birth to a child,  $g_{ij}$  is the number girls alive at birth order  $j$ ,  $x_{ij}$  a vector of exogenous covariates (among them the age at birth, which acts as a time trend),  $\alpha_s$  state (or mother) and  $\nu_s$  birth order specific fixed effects and  $u_{ij}$  a woman and birth order specific error term. Compared to parity progression (Arokiasamy, 2002) or count models (Wang and Famoye, 1997) this specification has two advantages. First, it models the woman's complete birth history and second it allows the gender composition of the woman's offspring to vary after every birth.<sup>19</sup>

The fertility decision outlined in equation 5 states that at every birth order the number of boys decreases and the number of girls increases the benefit to a further birth. Due to collinearity only one of these two variables can be included in the specification. The latter is chosen for two reasons. First, the data suggest a consistently positive correlation between the number of girls and fertility decisions. The number of boys, by contrast, has an ambiguous effect on fertility (see figure 2). Second, the evidence in Appendix A shows that the relation between daughters and fertility rates is linear and robust to the number of brothers as well as to other gender compositions of siblings. Note that  $g_{ij}$  is a measure of the number of girls alive as opposed to the total number of female births experienced by the woman.<sup>20</sup> A test of the hypothesis  $\gamma_g = 0$ , therefore, will investigate implication (i) of the theoretical model.

The coefficient  $\gamma_g$  estimated in equation 8 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, the paper includes dummy variables for specific gender compositions of children

$$y_{ij}^* = D_g + D_{bb} + D_{bg} + D_{gg} + D_{bbb} + D_{bbg} + D_{bgg} + D_{ggg} + x'_{ij}\beta + \alpha_s + \nu_{ij} + u_{ij} \quad (9)$$

where  $D_g$  is specific to birth order one and takes the value 1 if the firstborn is a girl and  $D_{bb}$ ,  $D_{bg}$  and  $D_{gg}$  apply to birth order two and take the value 1 if the two first children are two boys, a boy and a girl and two girls respectively. Finally,  $D_{bbb}$ ,  $D_{bbg}$ ,  $D_{bgg}$  and  $D_{ggg}$  are specific to birth order three and take the value 1 if the first three children are three boys, two boys and a girl, two girls and a boy and three boys respectively. Note that the order of

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<sup>19</sup>Arulampalam and Bhalotra (2006) employ a similar specification in the context of infant mortality.

<sup>20</sup>The specification also includes a dummy for the death of one of the children in the family.



children is not considered here. The estimation of equation 9 only considers children born at birth order four or less.

## 5.2 Difference-in-differences specification

Implication (ii) of the theoretical model states that a decrease in  $d_e$  attenuates the positive correlation between the number of girls and fertility (see equation 5). In the reduced form equation, this correlation is denoted by  $\gamma_g$  (see equation 8). Thus, we would expect the introduction of the Dowry Prohibition Rules to decrease the parameter estimate of  $\gamma_g$ . The paper investigates this hypothesis by estimating a difference-in-differences model, which also allows for a change in the slope parameter  $\gamma_g$

$$y_{ij}^* = (T + P_{1985-2000} + P_{1985-2000}T)(1 + g_{ij}) + g_{ij}\gamma_g + x_{ij}\beta + \alpha_s + \nu_{ij} + u_{ij} \quad (10)$$

where  $T$  is a dummy variable for the child's mother belonging to the treatment group and  $P_{1985-2000}$  a dummy for child  $i$  being born after the introduction of the policy; this encompasses births between the years 1985 and 2000. A test for the parameter on the triple interaction  $P_{1985-2000}Tg_{ij}$  being negative estimates whether the policy attenuated the correlation between the number of girls and fertility. Further, a test of the hypothesis that the coefficient estimate on the post-treatment interaction  $P_{1985-2000}T$  is different from zero investigates whether the policy increased fertility independently of its effect via  $\gamma_g$ .<sup>21</sup>

Because the policy changes outlined in section 4.2 were introduced in the whole of India simultaneously, geographical variation cannot be employed to identify the effect of the policy. Instead, this paper uses three different variations to distinguish treatment and control groups. The first and preferred specification exploits the fact that exposure to the Dowry Prohibition Rules varied with the gender composition of children at the time of the reform. Recall that, in the theoretical model, the expected dowry ( $d_e$ ) influences fertility choices by determining the expected change in dowry payments from another birth ( $E_D$ ). This term is defined as  $E_D = d_e\delta^{b+1}(1 - \delta^g)$  and increases in the number of girls,  $g$ . The impact of the decrease in  $d_e$  resulting from the introduction of the Dowry Prohibition Rules will thus be stronger for families with more girls. For example, for households with no daughters  $E_D = 0$ . Consequently, these families are not affected by the Dowry Prohibition Rules. This paper focuses on the first child only and uses the gender of the firstborn to distinguish treated from untreated individuals. Parents with firstborn daughters are the treatment group. House-

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<sup>21</sup>The model also allows for a change in the intercept to avoid imposing assumptions on the effect of the Dowry Prohibition Rules on fertility.

holds, whose firstborn is male, are the control group. The exogeneity of the firstborn’s gender is investigated in section B.

To illustrate how exposure to the reform varies with the gender of the firstborn child, consider two families with  $k$  children. Family  $T$  belongs to the treatment group and has  $g_T$  girls and  $b_T$  boys. Family  $C$  belongs to the control group and has  $g_C$  girls and  $b_C$  boys. Further, assume that (with the exception of the first child) the gender compositions of children in families  $T$  and  $C$  are identical, which implies that  $g_T = g_C + 1$  and  $b_T = b_C - 1$ . The expected change in dowry transfers following a subsequent birth are, respectively

$$E_D(T) = d_e \delta^{b_T+1} (1 - \delta^{g_T}) \quad (11)$$

$$E_D(C) = d_e \delta^{b_C+1} (1 - \delta^{g_C}) \quad (12)$$

subtracting 12 from 11 yields

$$d_e \delta^{b_T+1} (1 - \delta) \quad (13)$$

which denotes the difference in expected dowry transfers resulting from another child between the treatment and control groups (holding everything else constant). Term 13 implies that a decrease in  $d_e$  will lead to a larger decrease in  $E_D$  for the treatment than for the control group. We would thus expect  $\gamma_g$  to decrease more for families with a firstborn daughter. Term 13 also implies that, before the reform, the expected gains from another birth are larger for the treatment than for the control group. The difference-in-differences specification can test for this. Note that the employment of the gender of the firstborn would induce a strong collinearity between the treatment dummy  $T$  and one of the measurements of the gender composition, the number of girls  $g_{ij}$ . To avoid this, the estimations for this treatment group re-define  $g_{ij}$  and only indicate the number of girls born at birth order 2 or above. Hence, the first birth is not used when constructing  $g_{ij}$ .

From a more intuitive point of view, a firstborn girl mechanically shifts the gender composition of children within a family towards girls. See Appendix B for evidence.<sup>22</sup> A predominantly female gender composition, in turn, translates into high dowry payments net of dowry receipts. This is confirmed by descriptive evidence from the ICRISAT. Columns (2) and (3) of Table 2 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 place, 74 percent.<sup>23</sup> A decrease in expected dowry

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<sup>22</sup>The magnitude of the estimate is in line with what reported by Rosenblum (2013a).

<sup>23</sup>These figures consider the overall gender composition of the household and not just the gender compo-

payments will thus affect these families more than families with a similar number of sons and daughters. The gender of the firstborn child has previously been employed by other studies to proxy fertility behaviour or preferences for sons (Rosenblum, 2013a; Bhalotra and Cochrane, 2010; Jensen, 2003).

The second and third sources of variation exploit the heterogeneity in dowry payments before the introduction of the Dowry Prohibition Rules. The second specification distinguishes patrilineal and matrilineal states. In the former, the ancestral descent is traced through the father, in the latter through the mother. In the matrilineal states chosen in this paper women enjoy, on average, a higher status.<sup>24</sup> Additionally, kinship systems in these states appear to show larger interactions between married women and their families of birth. As a consequence, parents receive on average more support from their married daughters (See Das Gupta et al., 2003), which may counteract the dowry paid at her wedding. Patrilineal states, by contrast, have exhibited larger gender gaps in a number of human development indicators.<sup>25</sup> Further, qualitative research has pointed out that these societies are characterised by higher dowry payments. The empirical specification thus defines patrilineal states as the treatment group.<sup>26</sup>

Finally, this study compares lower and higher caste households across India. Dowries weigh heavier on the budget of lower caste household for two reasons. First, lower caste status is seen as an unattractive feature in a bride and lower caste brides have an incentive to marry "upwards" in terms of caste to gain status. The dowry payment may serve as compensation for this. Second, lower caste households exhibit lower socio-economic outcomes. The combination of these two factors lead parents belonging to lower castes to pay higher dowries, *ceteris paribus*. In fact recent qualitative work points out that lower caste members started paying disproportionately high dowries in an effort to copy the higher castes (Srinivas, 1997). Evidence from the ICRISAT confirms this. Columns (4) and (5) of table 2 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 influence the inflow of dowry payments, lower caste households show considerably higher out payments of dowries. This 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.

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sition of the offspring, and can thus only be seen as suggestive evidence.

<sup>24</sup>Matrilineal states are Kerala, Meghalaya and West Bengal.

<sup>25</sup>Patrilineal states are Bihar, Delhi, Haryana, Himachal Pradesh, Madhya Pradesh, Punjab and Uttar Pradesh.

<sup>26</sup>States that are not classified as either matri- or patrilineal are dropped from the analysis.

## 6 Results

### 6.1 Gender Composition of Children and Fertility Choices

Implications (i) and (ii) of the theoretical model are tested using information on completed reproductive histories of women in India born between 1953 and 1970.

Table 3 contains results from the reduced form fertility equation 8. The dependent variable,  $y_{ij}$ , takes the value 1 if woman  $i$  at birth order  $j$  gave birth to another child. The gender composition of children is measured via  $g_{ij}$  and the more detailed dummy variables outlined in section 5.1. The parameter estimates suggest a strong correlation between the number of girls alive at every birth order and fertility choices (coefficient  $\gamma_g$ ); at every birth order the birth (and survival) of a girl is associated with a 1 to 2 percentage point increase in the probability of the woman giving birth to another child. Thus, substituting a son for a daughter increases the probability of a further birth by around 2 percentage points. If parents were indifferent with regards to their children's gender, the parameter estimate would be zero. The size of the coefficient estimated here, by contrast, is considerable. The increase in fertility resulting from the female gender of a child is comparable to a decrease in 1.5 years of maternal education. The estimates are robust to the inclusion of family-specific characteristics, see columns (1) to (3). The positive sign of  $\gamma_g$  lends support to implication (i).

Column (4) reports the parameter estimates for the specification outlined in equation 9. At birth order one, the birth of a girl increases the probability of a further birth by almost 2 percentage points compared to the base category (women with a firstborn male). The positive correlation between the number of girls and fertility is also apparent at birth order two. The parameter estimates on  $D_{bb}$ ,  $D_{bg}$  and  $D_{gg}$  are strictly increasing in the number of girls. At birth order three, gender compositions with more girls and boys (rows  $i$  and  $j$ ) increase fertility whereas more boys than girls (rows  $g$  and  $h$ ) decrease birth rates. The parameter estimates of equation 9 can be used to calculate the probability that a woman with average characteristics has another birth at each birth order. These are reported in column (5). In some cases the gender composition of children has a stronger effect than its total number. A woman with three daughters, for instance, shows a higher estimated probability to have another child compared to a woman with only one daughter.

## 6.2 Dowries and Fertility Decisions

### 6.2.1 Difference-in-differences estimates

Table 4 reports the difference-in-differences results using the gender of the firstborn child to distinguish treatment and control group. This is the preferred specification. Columns (1) to (3) use  $g_{ij}$  to measure the gender composition of children. Rows (a) to (c) report the parameter estimates for the coefficients on the variables  $D_{1985-2000}$ ,  $D_T$  and  $D_{1985-2000}D_T$  in equation 10. They estimate the effect of the policy on the intercept of the reduced form fertility equation. The coefficient on the post-treatment interaction (row c) suggests that the policy increased the chances of a further birth between 2 and 3 percentage points.

Row (d) reports the parameter estimates on  $\gamma_{ij}$ . This corresponds to the coefficient  $\gamma_g$  for the control group before the treatment. The results show that before the policy change the control group exhibited a weakly positive correlation between  $g_{ij}$  and  $y_{ij}^*$ .<sup>27</sup> The parameter estimates for the number of girls alive are around 0.01. Row (e) reports the coefficients on the number of girls interacted with the post dummy ( $D_{1985-2000}g_{ij}$ ). This parameter estimates how  $\gamma_g$  changed after the policy for the control group. The estimates show a small strengthening of this correlation for the control group after 1985. Row (f) reports the parameter estimates on the interaction between the number of girls and the treatment dummy ( $D_Tg_{ij}$ ), which estimates the difference in  $\gamma_g$  between the treatment and control group before the introduction of the policy. As shown in equation 13, the theoretical model predicts this coefficient to be positive. 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 subsample 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 8).

Row (g) reports the estimates for the triple interaction between the post and treatment dummies and the number of girls ( $D_{1985-2000}D_Tg_{ij}$ ). This coefficient is the estimated effect of the Dowry Prohibition Rules on  $\gamma_g$  and is the coefficient of interest. The estimates show that the correlation decreased by between 3 and 4 percentage points, which lends support to prediction (ii), 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) can be used to calculate that, for the control group, the coefficient  $\gamma_g$  increased from 0.005 to 0.019. For the treatment group, by contrast, it decreased from 0.06 to 0.031.

Column (4) reports the coefficient estimates on the triple interaction between the post and

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<sup>27</sup>Recall that  $g_{ij}$  in this specification only considers girls born at birth order 2 or above.

treatment dummies on the one hand and the dummies for the gender compositions outlined in equation 9 on the other.<sup>28</sup> The figures suggest that the decrease in the correlation between the gender composition of children and fertility was strongest at birth orders three and four. At birth order three the coefficient for couples with one boy and one girl decreased by 4 percentage points. Similarly, at birth order 4, the coefficient for families with two girls and one boy decreased by 6 percentage points. The correlation for birth order 2, by contrast, increased as a result of the policy.

Table 5 reports the difference-in-differences estimates employing two alternative specifications to distinguish the treatment and control group. Columns (1), (2) and (3) compare patri-lineal (the treatment group) and matri-lineal states (the control group). The results are similar to the ones outlined in the main specification. The Dowry Prohibition Rules are estimated to increase the intercept of the fertility equation by between 6 and 8 percentage points. These magnitudes exceed the ones of table 4. For the control group the number of girls appears to have a weaker correlation to fertility choices. This has intuitive appeal; matri-lineal states place high importance on daughters and have been documented to pay lower dowries. It is, therefore, unlikely that these states exhibit strong son preferences in fertility behaviour. The estimates in row (d) show that also for this specification the policy decreased the correlation between the number of girls and fertility by around 3 percentage points. Columns (4), (5) and (6) report the parameter estimates comparing families of low castes (the treatment group) with high caste individuals (the control group). Akin to before, the estimates are in line with the ones reported in table 4 and the ones reported in columns (1) to (3). For this specification, the reform is estimated to have decreased the correlation between girls and fertility by around 2 percentage points.

### 6.2.2 Treatment Heterogeneity

This section employs the specification that distinguishes treatment and control via the gender of the firstborn child to investigate how the effect of the policy changed with the characteristics of the mother. Columns (1) and (2) of table 6 show the impact of the Dowry Prohibition Rules distinguishing by the religion of the mother. The Dowry Prohibition Rules excluded dowers or mahr, which are both given from the husband to the wife under the Sharia personal law. Because dowers and mahrs act as brideprices and alleviate the financial pressure of dowries, one would expect that the reform had a smaller impact on Muslims than Hindus. The results confirm this. For Muslims, the policy decreased  $\gamma_g$  by less than half (2 percentage points) compared to Hindus (5 percentage points). Columns (3) and (4) distinguish by patri-

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<sup>28</sup>Recall that the gender of the first child is not used in the construction of these variables. The estimates, therefore, apply to birth order two and above.

and matri-lineal states. As mentioned before, dowries are more pronounced in patri-lineal states and the results show that the effect of the policy in these states is more pronounced (5 percentage points) compared to matri-lineal states (2 percentage points). Columns (5) and (6), moreover, suggest a clear positive correlation between the impact of the Dowry Prohibition Rules and the mother’s education. The slope parameter for the former group decreases by 5 percentage points whereas for the latter it is only attenuated by 4 percentage points. 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. Columns (7) to (10) employ two different variables that have previously been used as proxies for female autonomy: age and age difference at marriage (Abadian, 1996). The results confirm that more autonomous women responded more strongly to the Dowry Prohibition Rules.

## 7 Identification Concerns

The key identifying assumption of the difference-in-differences estimator is that the time trend in reproductive behaviour would have been the same for treatment and control group in the absence of the reform. This sub-section assesses the plausibility of this assumption in the present context.

### 7.1 Confounding Factors

A first concern is the presence of confounding factors. Children in the post period are born between 1985 and 2000. If, in this time interval, anything caused the fertility behaviour of treatment and control groups to be different, its effect would erroneously be attributed to the Dowry Prohibition Rules. A first example of a confounding factor is the introduction of policies that affected the economic value of girls, such as the Hindu Succession Act. This legislation facilitated inheritance by daughters and may thus have affected reproductive choices of parents. It has been analysed recently (Deininger et al., 2013; Rosenblum, 2013b; Roy, 2011).<sup>29</sup> Indeed, Roy (2011) finds that the Act had a pronounced impact on dowry payments. A second confounding factor is the phenomenon of dowry inflation. If dowries exhibit an increasing trend over time (as pointed out by Rao, 2000; 1993), the decrease in the probability of dowry transfers analysed here is unlikely to translate into a substantial change

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<sup>29</sup>Kerala changed the law in 1976, Andhra Pradesh in 1986, Tamil Nadu in 1989, Maharashtra and Karnataka in 1994.

in  $d_e$ . A common strategy for addressing these concerns is to include state specific trends in the estimation in equation 10. Column (1) of table 7 shows that the parameter estimates are robust to the addition of state specific trends; the difference-in-differences parameter is -0.038.

The paper also estimates a truncated version of equation 10 as

$$y_{ij}^* = (P_{1985-a} + T + P_{1985-a}T)(1 + g_{ij}) + n_{ij}\gamma_n + g_{ij}\gamma_g + x_{ij}\beta + \alpha_i + u_{ij} \quad (14)$$

where  $a = 1995, 1990, 1986, 1985$ . In practice, only children born between 1985 and the cut-offs defined in  $a$  are used to identify the effect of the Dowry Prohibition Rules. Children born after the cut-off are dropped from the analysis. This shorter time frame significantly decreases the importance of confounding factors. It is, for instance, unlikely that other policies affected treatment and control groups differently between 1985 and 1986. Furthermore, dowry inflation is unlikely to have changed substantively in those years. Columns (2) to (5) of table 7 report the parameter estimates, which suggest that the results are stable to the changes in the post period. The effect of the policy on the slope parameter remains negative. For the estimates in row (c) in columns (2) to (5) the magnitudes of the effect of policy remain similar to the previous specification, with magnitudes between 2 and 4 percentage points.

## 7.2 Sex Selective Abortions

A second concern is the introduction of prenatal sex determination techniques around the time of the Dowry Prohibition Rules. As pointed out by Bhalotra and Cochrane (2010), families with a firstborn daughter have a stronger incentive to abort female fetuses at birth order two or above, which would violate the common time trend assumption.

The paper addresses this concern in three ways. First, the specification outlined in equation 14 helps us understand the importance of sex selective abortions. Previous work has documented that the practice of sex selective abortions increased in the 1990s (see Arnold et al, 2002b, for instance). A treatment effect estimated for children born in 1985 (and 1986) only is thus significantly less likely to be biased by sex selective abortions. Hence, the parameter estimates reported in table 7 can be seen as first evidence against the influence of sex selective abortions.

The second strategy compares the sex ratios at birth for the treatment and control group before and after the introduction of the policy. If parents respond to the Dowry Prohibition Rules by aborting female fetuses, we would expect the sex ratio at birth to decrease for



the treatment group in the years after the policy. The sex ratio for the control group, by contrast, should remain the same. Figure 6 plots the percentage of girls born in the years before and after the introduction of the Dowry Prohibition Rules.<sup>30</sup> In both time periods, the differences between treatment and control group do not appear to be significant. In fact, the only statistically significant differences between the two samples are in the years 1983 and 1989. Note that, as shown in table 7 the results are robust to using children born before 1989.

Third, this paper estimates a model similar to the one employed by Bhalotra and Cochrane (2010) to test the hypothesis that parents adopted sex selective abortions after the year 1985. The paper estimates the probability that a child born to woman  $i$  at birth order  $j$  is female ( $w_{ij}$ ) as a function of individual characteristics, the gender of the oldest sibling and a dummy for whether the index child was born after the introduction of the Dowry Prohibition Rules.

$$w_{ij} = P_{1985-2000} + T + P_{1985-2000}T + x_{ij}\beta + f_s + f_st + u_{ij} \quad (15)$$

where  $P_{1985-2000}$  is an indicator variable for the child being born after the introduction of the policy in 1985 (and up until the year 2000), and  $T$  a dummy variable for the child's oldest sibling being female. The specification also includes state fixed effects and state specific trends. A test on the interaction  $TP_{1985-2000}$  investigates whether the Dowry Prohibition Rules had a significant impact on the probability of a girl being born.

Table 8 reports the parameter estimates of equation 15.<sup>31</sup> Overall, the estimates suggest that the policy had no significant impact on the probability of a female birth. The value of the post-treatment interaction coefficient with and without state specific trends (columns 1 and 3) are not significant. Furthermore, the sizes of the parameter estimates are very small. They suggest that the change in the law decreased the chances of a female birth between 0.4 and 0.7 percentage points. The specification also divides the post-period into two (columns 2 and 4). Here again, no significant effect is found. One possible reason for the difference between the estimates presented here and the ones of Bhalotra and Cochrane (2010) may be due to the different samples used. Whilst the authors employ births to all women in India, this paper only uses women who had come to the end of their reproductive years and who are born between the years 1953 and 1970.

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<sup>30</sup>Because the treatment is defined along the lines of the firstborn child (the exogeneity of which has already been argued for in table 10) this panel only considers children of birth order 2 or above.

<sup>31</sup>The dependent variable,  $w_{ij}$ , takes the value 100 if woman  $i$  at birth order  $j$  gave birth to a girl.

### 7.3 Robustness Checks

To address the concern that the 1985 dummy is correlated with changes in fertility that are independent of the Dowry Prohibition Rules, this paper carries out a number of falsification tests. Columns (1) and (2) of table 9 estimate equation 10 adding, separately, two placebo treatments, one for the years 1976-84 and one for the years 1981-84. In practice, the specification estimated here employs two post-treatment periods. The first is the placebo and includes the years 1976 to 1984 (or 1981 to 1984, depending on the specification). The second comprises births between 1985 and 2000. By including the original post period (the second, going from 1985 to 2000) the paper simultaneously investigates whether the original model is robust to the inclusion of pre-treatment trends. The parameter estimates suggest that the placebo treatment had a negligible effect on fertility behaviour. Although the post-treatment interaction is positive and significant, the size of the coefficients (around 1.3 percentage points) are smaller compared to the post 1985 estimates (around 4 percentage points). Furthermore, the difference-in-differences estimator for the change in  $\gamma_g$  using the placebo time periods are very close to zero. The estimator employing the post 1985 time period, by contrast, remains negative with a very similar magnitude to the main specification, a decrease of around 4 percentage points.

Another way of checking the robustness of the specification is to employ women 15 years older than the estimation sample. These individuals had come to the end of their reproductive years by the time the Dowry Prohibition Rules were introduced. The specification carries out two placebo treatments for these individuals in the years 1961 and 1971 (which correspond to the years 1975 and 1985 for the estimation sample). Columns (3) and (4) find no significant effects of the falsification tests on reproductive behaviour.

A commonly employed method to investigate the plausibility of the common time trends assumption is to investigate the behaviour of cohorts too old to be affected by the policy change. Figure 7 explores the fertility behaviour of women born between the years 1942 and 1953. The time trends in fertility behaviour for these cohorts look parallel. Although not conclusive, this can be seen as suggestive evidence in favour of the common time trend assumption.

## 8 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

worthy of a short discussion. 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 boys are more valuable than girls. The results put forward here, by contrast, argue that a large part of this behaviour can be 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.

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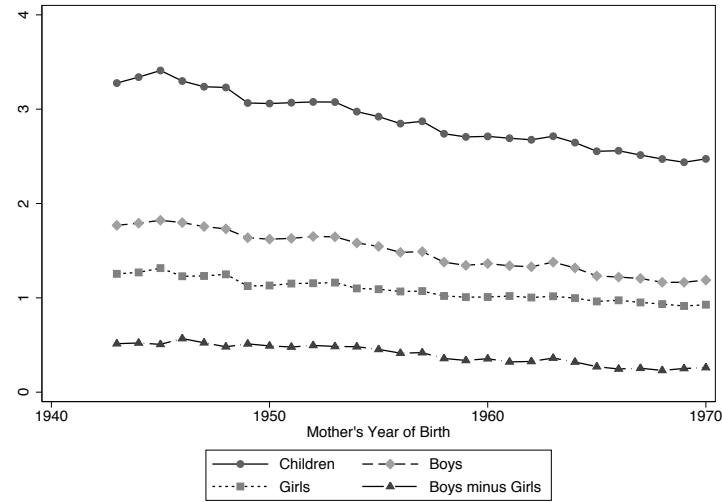
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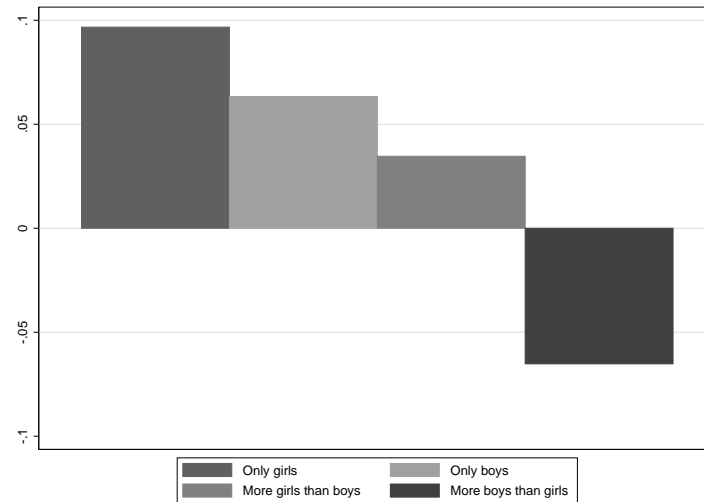
# Figures

Figure 1: Ideal Number of Children, Boys and Girls



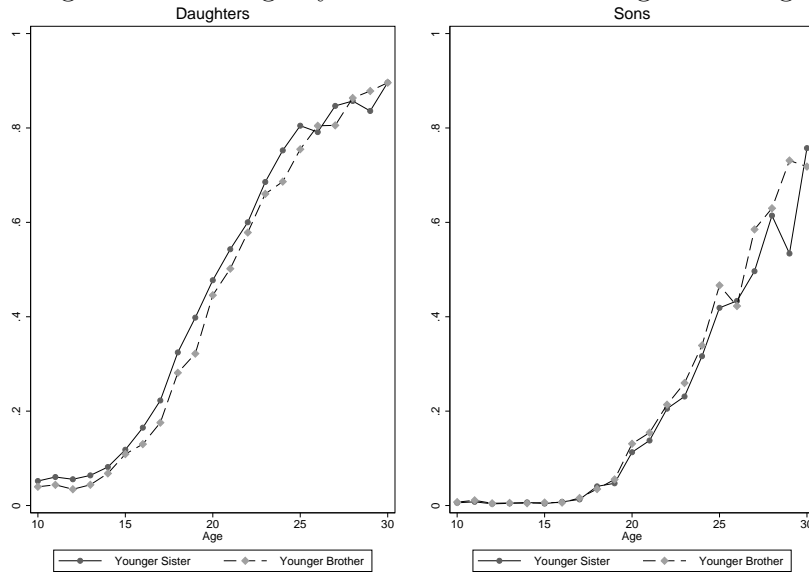
**Notes:** Sample consists of women born between 1942 and 1970, aged 36 to 49 years at interview and drawn from NFHS-1, NFHS-2 and NFHS-3; own calculations.

Figure 2: Conditional Differences in Probability of another Birth



**Notes:** Parameter estimates reported are from linear probability model; dependent variable takes value 1 if woman gives birth to a child; 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; base category is women with equal number of boys or girls; mother level fixed effects included.

Figure 3: Marriage by Gender of Next Youngest Sibling



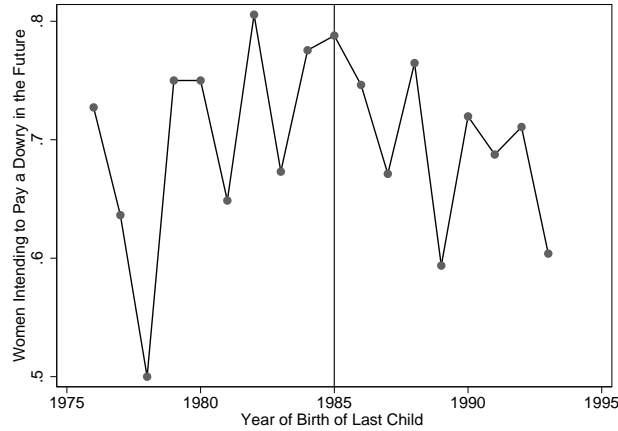
**Notes:** Sample consists of boys and girls aged 10 to 30 with at least one younger sibling; drawn from NFHS-3.

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



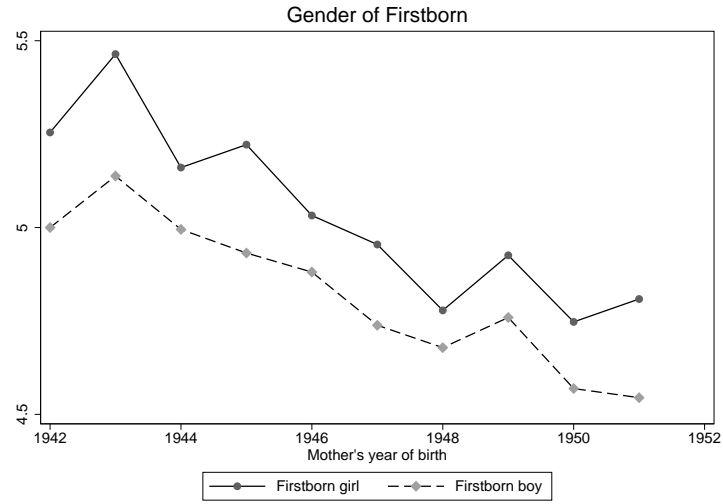
Notes: sample consists of women interviewed in 1993 drawn from the SWAF; own calculations.

Figure 6: 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 7: 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: Estimates of determinants of marriage

	(1)	(2)	(3)	(4)	(5)	(6)
	Dependent variable: Probability of marriage					
	Daughters	Sons	Daughters	Sons	Daughters	Sons
(a) <b>Younger sister</b>	0.028*** (0.004)	-0.009 ** (0.003)				
(b) <b>Number of younger sisters</b>			0.008 ** (0.003)	-0.013*** (0.004)		
(c) <b>Number of younger brothers</b>					-0.010*** (0.003)	0.004 (0.003)
<b>Children</b>	37,431	28,650	37,431	28,650	37,431	28,650
<b>R-Squared</b>	0.345	0.298	0.350	0.301	0.350	0.300

**Notes:** Parameter estimates reported are from linear probability model; for daughters, dependent variable takes value 1 if daughter leaves the house; for sons, dependent variable takes value 1 if son reports to be married; sample consists of children aged 10 to 30 years at interview and drawn from NFHS-3; child level covariates include age and birth order fixed effects; mother level covariates include education, religion, caste; father level covariates include education; standard errors are reported in parentheses and are clustered at the state level; \*\*\*, \*\*, and \* indicate significance at the 1%, 5% and 10% levels.

Table 2: Dowries paid in rural India 1975 - 1984

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

	(1)	(2)	(3)	(4)	(5)
Dependent variable: Probability of further birth					
	Parameter estimates				Predicted Probability
(a) <b>Number of girls alive</b>	0.018*** (0.004)	0.021*** (0.004)	0.011*** (0.004)		
<b><u>Birth order 1</u></b>					
(b) <b>Boy</b>				–	77%
(c) <b>Girl</b>				0.016*** (0.002)	78%
<b><u>Birth order 2</u></b>					
(d) <b>Boy, boy</b>				0.018*** (0.004)	72%
(e) <b>Boy, girl</b>				0.050*** (0.003)	75%
(f) <b>Girl, girl</b>				0.130*** (0.004)	83%
<b><u>Birth order 3</u></b>					
(g) <b>Boy, boy, boy</b>				–0.061*** (0.008)	64%
(h) <b>Boy, boy, girl</b>				–0.114*** (0.005)	58%
(i) <b>Boy, girl, girl</b>				0.042*** (0.005)	74%
(j) <b>Girl, girl, girl</b>				0.163*** (0.006)	86%
<b>Parental controls</b>	no	yes	no	no	no
<b>Birth order fixed effect</b>	yes	yes	yes	yes	yes
<b>Mother fixed effect</b>	no	no	yes	yes	yes
<b>State fixed effect</b>	yes	yes	no	no	no
<b>Observations</b>	333,115	326,323	355,203	208,756	208,756
<b>Mothers</b>	67,247	67,683	72,247	47,518	47,518
<b>R-Squared</b>	0.250	0.265	0.387	0.320	0.320

**Notes:** Parameter estimates reported are from linear probability model; dependent variable takes value 1 if woman gives birth to a child; 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; standard errors are reported in parentheses and are clustered at the state or 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)
	Dependent variable: Probability of further birth			
(a) Post 1985	-0.105*** (0.014)	-0.089*** (0.016)	-0.097*** (0.005)	-0.145*** (0.008)
(b) Firstborn is female	0.013*** (0.002)	0.014*** (0.002)		
(c) Firstborn is female * post 1985	0.020*** (0.005)	0.020*** (0.005)	0.033*** (0.006)	0.031*** (0.008)
(d) Number of girls alive	0.011 * * (0.004)	0.013*** (0.004)	0.005 (0.003)	
<b>Interactions of number of girls alive with</b>				
(e) Post 1985	0.015*** (0.004)	0.012 * * (0.005)	0.014*** (0.003)	
(f) Firstborn is female	0.036*** (0.004)	0.037*** (0.005)	0.055*** (0.003)	
(g) Firstborn is female * post 1985	-0.034*** (0.004)	-0.033*** (0.004)	-0.043*** (0.004)	
<b>Interactions of firstborn is female * post 1985 with</b>				
<u>Birth order 2</u>				
(h) Boy				-
(i) Girl				0.036*** (0.014)
<u>Birth order 3</u>				
(j) Boy, boy				-0.025 (0.020)
(k) Boy, girl				-0.036 * * (0.015)
(l) Girl, girl				-0.008 (0.018)
<u>Birth order 4</u>				
(m) Boy, boy, boy				-0.045 (0.037)
(n) Boy, boy, girl				-0.036 (0.023)
(o) Boy, girl, girl				-0.059*** (0.022)
(p) Girl, girl, girl				-0.016 (0.030)
Parental controls	no	yes	no	no
Birth order fixed effect	yes	yes	yes	yes
Mother fixed effect	no	no	yes	yes
State fixed effect	yes	yes	no	no
Observations	239,778	235,734	250,339	227,154
Mothers	44,018	43,261	47,518	47,518
R-Squared	0.256	0.267	0.391	0.353

**Notes:** Parameter estimates reported are from linear probability model; dependent variable takes value 1 if woman gives birth to a child; 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; standard errors are reported in parentheses and are clustered at the state or mother level; \*\*\*, \*\*, and \* indicate significance at the 1%, 5% and 10% levels.



Table 5: Difference-in-Differences Estimates: Alternative Specifications

	(1)	(2)	(3)	(4)	(5)	(6)
	Dependent variable: Probability of further birth					
(a) <b>Patrilineal States * post85</b>	0.056* (0.025)	0.055 ** (0.023)	0.078*** (0.011)			
(b) <b>Lower caste * post 1985</b>				0.026*** (0.009)	0.028*** (0.009)	0.052*** (0.007)
(c) <b>Number of girls alive</b>	0.001 (0.020)	0.003 (0.022)	-0.022*** (0.005)	0.041*** (0.005)	0.046*** (0.005)	0.054*** (0.002)
<u>Interactions of number of girls alive with</u>						
(d) <b>Patrilineal States * post85</b>	-0.026*** (0.007)	-0.024 ** (0.010)	-0.030*** (0.006)			
(e) <b>Lower caste * post 1985</b>				-0.017*** (0.003)	-0.015*** (0.003)	-0.022*** (0.004)
<b>Parental controls</b>	no	yes	no	no	yes	no
<b>Birth order fixed effect</b>	yes	yes	yes	yes	yes	yes
<b>Mother fixed effect</b>	no	no	yes	no	no	yes
<b>State fixed effect</b>	yes	yes	no	yes	yes	no
<b>Observations</b>	165,074	162,564	165,074	326,882	326,323	348,165
<b>Mothers</b>	31,433	30,546	31,433	65,803	65,232	70,803
<b>R-Squared</b>	0.246	0.261	0.379	0.255	0.267	0.382

**Notes:** Parameter estimates reported are from linear probability model; dependent variable takes value 1 if woman gives birth to a child; 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; standard errors are reported in parentheses and are clustered at the state or mother level; \*\*\*, \*\*, and \* indicate significance at the 1%, 5% and 10% levels.

Table 6: Treatment Heterogeneity

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	Hindu	Muslim	Patri-lineal States	Matri-lineal States	Primary education	No Primary education	Above av. age at marriage	Below av. age at marriage	Below av. age difference	Above av. age difference
(a) Post 1985 * firstborn is female	0.036*** (0.007)	0.019 (0.015)	0.037*** (0.010)	-0.035 (0.022)	0.033*** (0.011)	0.041*** (0.008)	0.035*** (0.009)	0.036*** (0.009)	0.038*** (0.009)	0.027*** (0.009)
(b) Number of girls alive	0.004 (0.003)	0.003 (0.006)	0.012*** (0.004)	-0.031*** (0.010)	0.013 * * (0.005)	0.012*** (0.003)	0.002 (0.005)	0.009*** (0.004)	0.011*** (0.004)	-0.001 (0.004)
Interactions of Number girls alive with										
(c) Post 1985 * firstborn is female	-0.046*** (0.004)	-0.019 * * (0.009)	-0.046*** (0.006)	-0.020 (0.015)	-0.054*** (0.007)	-0.037*** (0.005)	-0.063*** (0.007)	-0.035*** (0.005)	-0.052*** (0.006)	-0.037*** (0.006)
Parental controls	no	no	no	no	no	no	no	no	no	no
Birth order fixed effect	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
Mother fixed effect	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
State fixed effect	no	no	no	no	no	no	no	no	no	no
Observations	215,787	34,456	99,644	23,194	98,295	151,940	110,224	140,115	118,824	131,515
Mothers	41,941	5,560	17,193	4,942	21,360	26,137	22,233	25,285		
R-Squared	0.404	0.370	0.378	0.410	0.456	0.383	0.412	0.377	0.384	0.382

**Notes:** Parameter estimates reported are from linear probability model; dependent variable takes value 1 if woman gives birth to a child; 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; standard errors are reported in parentheses and are clustered at the state or mother level; \*\*\*, \*\*, and \* indicate significance at the 1%, 5% and 10% levels.

Table 7: Difference in Differences Estimates: Confounding Factors

	(1)	(2)	(3)	(4)	(5)
	Dependent variable: Probability of further birth				
<b>Sample: Children born between</b>	1970-2000	1970-1995	1970-1990	1970-1986	1970-1985
(a) <b>Firstborn is female * Post 1985</b>	0.029*** (0.006)	0.029*** (0.007)	0.029*** (0.007)	0.036*** (0.007)	0.049*** (0.012)
(b) <b>Number of girls alive</b>	0.013 ** (0.003)	0.003 (0.003)	-0.007 ** (0.003)	-0.000 (0.003)	-0.003 (0.003)
<b>Interactions of number of girls alive with</b>					
(c) <b>Firstborn is female * post 1985</b>	-0.038*** (0.004)	-0.040*** (0.004)	-0.034*** (0.005)	-0.034*** (0.007)	-0.018 ** (0.009)
<b>Parental controls</b>	no	yes	no	no	no
<b>Birth order fixed effect</b>	yes	yes	yes	yes	yes
<b>Mother fixed effect</b>	no	no	yes	yes	yes
<b>State fixed effect</b>	yes	yes	no	no	no
<b>Observations</b>	239,778	245,222	228,695	198,924	187,813
<b>Mothers</b>	45,517	47,518	47,518	47,518	47,518
<b>R-Squared</b>	0.410	0.361	0.314	0.258	0.240

**Notes:** Parameter estimates reported are from linear probability model; dependent variable takes value 1 if woman gives birth to a child; 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; standard errors are reported in parentheses and are clustered at the state or mother level; \*\*\*, \*\*, and \* indicate significance at the 1%, 5% and 10% levels.

Table 8: Difference in Differences Estimates: Sex Selective Abortions

	(1)	(2)	(3)	(4)
Dependent variable: Probability of female birth				
(a) <b>Firstborn is female</b>	-0.261 (0.350)	-0.261 (0.350)	-0.357 (0.311)	-0.358 (0.311)
(b) <b>Firstborn is female * Post(1985-2000)</b>	-0.686 (0.508)		-0.422 (0.522)	
(c) <b>Firstborn is female * Post(1985-1994)</b>		-0.708 (0.522)		-0.481 (0.536)
(d) <b>Firstborn is female * Post(1995-2000)</b>		-0.476 (1.242)		0.172 (1.250)
<b>Birth order fixed effect</b>	yes	yes	yes	yes
<b>Mother level controls</b>	no	yes	no	yes
<b>State order fixed effect</b>	yes	no	no	no
<b>State specific trends</b>	no	yes	no	yes
<b>Observations</b>	155,303	155,303	146,120	146,120
<b>R-Squared</b>	0.001	0.001	0.001	0.001

**Notes:** Parameter estimates reported are from linear probability model; dependent variable takes value 1 if woman gives birth to a child; 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; standard errors are reported in parentheses and are clustered at the state or mother level; \*\*\*, \*\*, and \* indicate significance at the 1%, 5% and 10% levels.

Table 9: Robustness Checks

	(1)	(2)	(3)	(4)
	Placebo: 1976-84	Placebo: 1981-84	Placebo: 1971-84	Placebo: 1961-84
Dependent variable: Probability of further birth				
	Sample: Women born 1953 - 1970		Sample: Women born 1942 - 1948	
(a) <b>Placebo * firstborn is female</b>	0.013*** (0.003)	0.014*** (0.004)	-0.008 (0.013)	0.007 (0.006)
(b) <b>Treatment * firstborn is female</b>	0.040*** (0.007)	0.039*** (0.007)		
(c) <b>Number of girls alive</b>	0.021*** (0.007)	0.029*** (0.004)	0.013*** (0.006)	-0.018*** (0.005)
Interactions of number girls alive with:				
(d) <b>Placebo * firstborn is female</b>	-0.006 (0.010)	-0.003 (0.006)	0.000 (0.008)	-0.003 (0.008)
(e) <b>Treatment * firstborn is female</b>	-0.045*** (0.012)	-0.041*** (0.008)		
<b>Observations</b>	250,339	250,339	41, 646	41, 646
<b>Mothers</b>	47,518	47,518	6,713	6,713
<b>R-Squared</b>	0.391	0.392	0.372	0.372

**Notes:** Parameter estimates reported are from linear probability model; dependent variable takes value 1 if woman gives birth to a child; 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; standard errors are reported in parentheses and are clustered at the state or mother level; \*\*\*, \*\*, and \* indicate significance at the 1%, 5% and 10% levels.

# A Robustness of Functional Form of Section 5.1

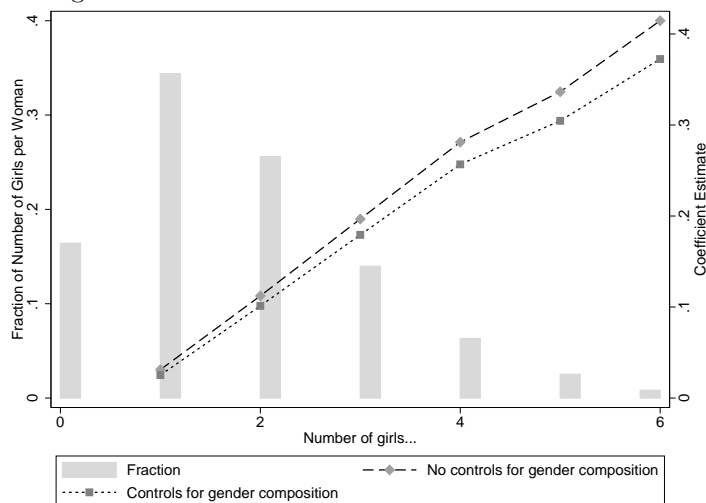
A number of concerns are connected with the modelling of the interaction between the number of daughters and the probability of the woman experiencing a further birth. First, the specification in equation 8 assumes a linear relationship between  $g_{ij}$  and  $y_{ij}^*$ . To investigate this linearity assumption, equation 8 is re-estimated using a dummy for every girl born as

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

where  $D_k$  is an indicator variable taking the value 1 if the number of surviving girls equals to  $k$ . The parameter estimates of  $\psi_k$  along with a histogram for the fraction of the number of girls alive in each family in the sample are shown in Figure A1. 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.<sup>32</sup>

Second, the specification in equation 8 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, the equation above 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 A1 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.

Figure A1: Coefficient Estimates for Number of Girls



**Notes:** Parameter estimates reported are from linear probability model; dependent variable takes value 1 if woman gives birth to a child; 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.

<sup>32</sup>The percentage of women with more than 6 girls alive is around 1%.

## B The Exogeneity of the Firstborn Child

The random assignment of treatment status in the first specification rests on the assumption that parents do not sex selectively abort at birth order one. A number of studies have already pointed out that sex ratios at first births in India lie within normal limits (Rosenblum, 2013a; Bhalotra and Cochrane, 2010; Poertner, 2010; Retherford and Roy, 2003). Descriptive evidence from the estimation sample lends further support to the exogeneity of the first born child's sex. Table 10 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 according to the gender of their firstborn child. The differences appear negligible.<sup>33</sup> Column (5) of Table 10 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. Just as above, the parameter estimates do not appear very large, 0.007 for the former and 0.001 for the latter. The low R-squared also points to the exogeneity of the firstborn's gender. A possible reason for these findings is that ultrasound technology was not widely available before 1985. This notion will be explored further in section 7.3.

Table 10: 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*** (0.002)	50.0%	51.1%	**	0.007** (0.003)
(e) Mother is muslim	0.012** (0.005)	11.3%	10.9%		-0.009 (0.005)
(f) Mother belongs to lower caste	0.002 (0.003)	15.6%	15.6%		0.002 (0.005)
(g) Age at first birth	0.000 (0.000)	20.4 years	20.5 years	***	0.001** (0.001)
(h) Father has primary education	0.004** (0.002)	47.4%	47.5%		-0.000 (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; standard errors are reported in parentheses and are clustered at the state level; \*\*\*, \*\*, and \* indicate significance at the 1%, 5% and 10% levels.

<sup>33</sup>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.