On the Quantity and Quality of Girls: Fertility, Parental Investments, and Mortality*

S Anukriti — Boston College†

Sonia Bhalotra — University of Essex‡

Hiu Tam — University of Oxford§

November 2017

Abstract

The introduction of prenatal sex-detection technologies in India has led to a phenomenal increase in abortion of female fetuses. We examine fertility and investment responses to these technologies. We find a moderation of son-biased fertility stopping, erosion of gender gaps in parental investments in breastfeeding and immunization, and convergence in the under-5 mortality rates of boys and girls. For every three aborted girls, roughly one additional girl survives to age five. We also find a shift in the distribution of girls in favor of low-socioeconomic status families. Our findings have implications not only for counts of missing girls but also for the later life outcomes of girls conditioned by greater early life investments in them.

JEL Codes: I15, J13, J16

Keywords: abortion, child mortality, fertility, gender, health, India, missing girls, parental investments, prenatal sex detection, sex-selection, ultrasound

*We thank Prashant Bharadwaj, Abhishek Chakravarty, Ritam Chaurey, Diana Greene Foster, Liberated Gonzalez, Rachel Heath, Mahesh Karra, Melanie Khamis, Adriana Kugler, Annemie Maertens, Claudia Olivetti, Rohini Pande, Laura Schechter, Alessandro Tarozzi, Ushma Upadhyay, and Shing-Yi Wang, and participants at various seminars and conferences for their helpful comments and suggestions. We are grateful to Luojia Hu and Analia Schlosser for discussion of our related papers.

†Department of Economics and IZA. anukriti@bc.edu.

‡Department of Economics and ISER. srbhal@essex.ac.uk.

§Centre for Business Taxation, Saïd Business School. eddy.tam@sbs.ox.ac.uk.

1 Introduction

Innovations in birth control technology have had substantial socioeconomic impacts. The contraceptive pill, for instance, gave women unprecedented control over fertility, preventing unwanted births and allowing women to determine the timing of births, with dramatic consequences for their marriage and labor market choices (Goldin and Katz (2002), Bailey (2006)). The legislation of abortion has similarly empowered women by enabling them to eliminate unwanted births (Gruber et al. (1999)). In fact, birth control technology has had far-reaching implications not only for women but also for children to the extent that parental investments tend to be greater in children that are more "wanted" when they are born (Grossman and Joyce (1990), Gruber et al. (1999), Donohue and Levitt (2001), Charles and Stephens Jr. (2006), Donohue et al. (2009), Bailey et al. (2017)).

This paper investigates the impacts on child quantity and quality of another new technology that has altered the demographic landscape in countries where sons are valued more than daughters, namely, prenatal sex detection technology (henceforth, ultrasound; see Section 2). An ultrasound scan can reveal fetal sex quite reliably at as early as 12 weeks of gestation, enabling selective abortion of unwanted girls without, in principle, risking the mother's health (Epner et al. (1998)). We focus on India, where ultrasound was introduced in the mid-1980s, before which abortion had been legalized. The low cost and the non-invasive nature of ultrasound scans has led to their widespread use for fetal sex determination, resulting in a staggering rise in sex-selective abortion, equivalent to 6 percent of potential female births during 1995-2005 (Bhalotra and Cochrane (2010)).[1] Even in the absence of son preference, and even in environments in which contraception is available, parents do not always want the children they conceive: e.g., 21 percent of all pregnancies in 2011 in the United States ended in abortion (Guttmacher Institute).[2]

When fetal sex determination was impossible, parents could adjust the gender composition of their children in two ways. First, by continuing childbearing till they achieved, for instance, the desired number of sons. Several studies document son-biased fertility stopping behavior,[3] which results in girls having more siblings than boys ((Clark (2000), Bhalotra and van Soest (2008), Jensen (2012), Rosenblum (2013)). The quantity-quality trade-off, driven by the budget constraint, implies that, even if parents do not actively discriminate against daughters, a gender gap in outcomes will emerge at the aggregate level simply because girls, on average, grow up in families with fewer per capita resources. The second option is to subject girls to deliberate neglect, culminating in excess girl mortality during early childhood (Das Gupta (1987), Pitt and Rosenzweig (1990), Sen (1990), Rose (2000), Oster (2009), Bhalotra (2010), Jayachandran and Kuziemko (2011)).[4] In this

[1] Bhalotra and Cochrane (2010) estimate that 480,000 girls—greater than the number of girls born in the United Kingdom each year—were aborted per year in India during 1995-2005.

[2] https://www.guttmacher.org/fact-sheet/induced-abortion-united-states

[3] This refers to a higher likelihood of continued childbearing after a female relative to a male birth to achieve the desired number of sons.

[4] Families that practice son-biased stopping rules may, of course, also actively discriminate against their daughters.

paper, we test the hypothesis that the facility to detect and, subsequently, terminate unwanted female fetuses in the post-ultrasound era weakened both these channels. In other words, we expect ultrasound access to decrease fertility driven by son-biased stopping rules and to reduce postnatal discrimination against girls.

To estimate the causal impact of prenatal sex-selection technology on postnatal gender gaps, we combine supply-driven changes in ultrasound availability with plausibly exogenous family-level variation in the incentive to sex-select. We construct an indicator for cohorts born pre- versus post-ultrasound exploiting information on the first *imports* of ultrasound scanners after tariff reductions in the mid-1980s. We construct a second indicator for cohorts born after a major expansion in ultrasound availability associated with initiation of *local production* from the mid-1990s, driven by relaxation of industrial licensing regulations. Previous work shows that these supply-side changes resulted in a phenomenal rise in sex-selective abortion, and that family-level variation in the incentive to utilize ultrasound technology to conduct sex-selection is captured in an indicator for the sex of the firstborn child (Bhalotra and Cochrane (2010)). Our identification strategy exploits (a) that the sex of the firstborn child is quasi-random and (b) that sex-selective abortion among second- and higher-order births is concentrated in families with a firstborn daughter (Almond and Edlund (2008), Abrevaya (2009), Bhalotra and Cochrane (2010)). The identifying assumption, which we verify, is that, in the absence of ultrasound technology, trends in the outcomes would have been identical across firstborn-son and firstborn-daughter families. In contrast to earlier studies of abortion (Gruber et al. (1999), Pop-Eleches (2006), Ananat et al. (2009)), we analyze outcomes of girls *relative* to boys. Differencing by gender allows us to control for unobservable trends that equally affect both boys and girls.

We find that sex-selection technology led to substantial improvements in the relative survival of girls after birth. In particular, we estimate that the pre-ultrasound gender gap in mortality before age five (henceforth, under-5 mortality) for second- and higher-order births in firstborn-girl families relative to firstborn-boy families (equal to 2.77 percentage points (p.p.)) declined by 60 to 100 percent after the introduction of ultrasound technology. These estimates are robust to conditioning upon mother fixed effects.

We also investigate parental investments in girls relative to boys and changes in son-biased fertility stopping as potential mechanisms for the decline in the gender gap in mortality. Pre-ultrasound, in families with a firstborn girl, boys were breastfed for a longer duration and were more likely to be vaccinated. We find that gender gaps in vaccination and breastfeeding were virtually eliminated post-ultrasound in firstborn-girl families, and this can explain 27 to 31 percent of the observed decline in under-5 excess female mortality (EFM), which we define as the excess of female over male mortality.[5] We also find complete closure of the gender gap in sibling size between firstborn-girl and

[5]Note that vaccination and breastfeeding are only two of the many markers of parental investments. It is possible that parents also increased other investments in girls and that can further explain the observed EFM decline; however, we do not have data on other investments.

firstborn-boy families (which, pre-ultrasound, was 0.14). Since sex-selective abortion allows parents to avoid unwanted children, we also tested whether availability of ultrasound drove actual fertility closer to desired fertility. Our estimates imply a complete elimination of undesired fertility, which was 0.117 births in firstborn-girl relative to firstborn-boy families in the pre-ultrasound years. Essentially, in the post-ultrasound era, girls were more likely to be born into households that wanted them.

Although the standard measures of fertility preference and son preference in survey data have their limitations, we find that the results are robust to controlling for the mother's reported ideal fertility and ideal sex ratio of her births.[6]

In sum, we observe a fairly dramatic post-ultrasound break in the trends in mortality, parental investments, and sibship size for girls relative to boys in firstborn-girl families (relative to firstborn-boy families). Nevertheless, the decline in postnatal death of girls only partially offsets the rise in female feticide. Our estimates imply that 90,200 excess under-5 female child deaths were *indirectly* averted each year by ultrasound technology, but that the rise in the number of aborted girls was higher. We estimate that for every three girls that "went missing" before birth, only one girl survived after birth who otherwise would have died.

The absolute decline in EFM in firstborn-girl families is, in general, larger among low socioeconomic status (SES) families, while the decline in fertility among women with firstborn daughters is broadly similar across socioeconomic groups. Previous work shows that sex-selective abortion in India is more prevalent among educated and wealthy women (Bhalotra and Cochrane (2010)). Thus, responses to ultrasound technology in terms of sex-selective abortion, fertility, and postnatal investments in girls versus boys occur on different SES margins. These results suggest that, accounting for its influence on abortion, fertility, and mortality, ultrasound availability generated a shift in the distribution of girls in favor of low-SES families: more girls were aborted in high-SES families and more girls survived to age five in low-SES families. This implies that the average long-run outcomes for girls (driven by improved early-life investments) are unlikely to converge toward those for boys as much as within-SES-group and within-household differences.

While a number of studies have sought to identify how changes in abortion *law* in different countries have modified fertility and investment in children (Gruber et al. (1999), Pop-Eleches (2006), Ananat et al. (2009)), the evidence on how prenatal sex detection *technology* (which made sex-selective abortion feasible) modifies fertility and investment in girls is more limited. Lin et al. (2014) show that abortion legalization in Taiwan decreased neonatal (but not post-neonatal) female mortality for higher-parity births conditional on SES, and that it reduced fertility at third and

[6]We define fertility, as is usual, as live births. Note that, in general, increased opportunities for abortion may lead to lower or higher fertility (e.g., Ananat et al. (2009), Ananat and Hungerman (2012)). Abortion mechanically reduces the number of live births conditional upon the number of pregnancies; but the knowledge that a pregnancy can be aborted may stimulate more pregnancies and the net effect on fertility is ambiguous *a priori*. However, since ultrasound led to sex-selective abortion and not abortion *per se*, we can test an unambiguous prediction, which is that there is less *son-biased* fertility stopping post-ultrasound.

higher parity. They do not examine health investments.[7] Our finding of a post-ultrasound decline in under-5 EFM in India is empirically relevant, given how high it was pre-ultrasound, at 1.37 p.p. overall and 2.87 p.p. among children preceded by a firstborn girl (Table 1 and Table A.1).[8]

There are two studies similar to ours—one for India and one for China—that seek to estimate the impacts of ultrasound technology on postnatal outcomes, and neither finds any post-ultrasound narrowing of gender gaps in survival. Almond et al. (2010) find, in sharp contrast to us, that ultrasound access in China *increased* neonatal EFM, which suggests that parents consciously reduced *prenatal* inputs in girls.[9] They find no impact on gender gaps in post-neonatal mortality, and they do not examine fertility. In a study of India, Hu and Schlosser (2015) find a narrowing of gender gaps in malnutrition but, despite this, find no change in excess girl mortality. They state: "Particularly puzzling is why we find a differential improvement in female nutritional status but do not see any significant increase in female survival probabilities (*pp. 1257*)." Our findings appear to resolve the puzzle and in Section 7 we discuss the source of this difference. Our results contribute several new insights to current debates on biased population sex ratios and to research on abortion, fertility, and parental investments (more details in Section 8).

The rest of the paper is organized as follows. Section 2 describes the Indian context. Sections 3 and 4 discuss the empirical strategy and the data. Section 5 presents results and Section 6 presents estimates of the implied magnitude of substitution between postnatal discrimination and sex-selective abortions due to ultrasound access. Section 7 relates our paper to a previous literature and Section 8 concludes.

2 Context

While son preference has characterized parts of Indian society for centuries, the availability of affordable prenatal sex-diagnostic techniques combined with legal access to abortion is more recent. Abortion was legalized in India with the passage of the Medical Termination of Pregnancy (MTP) Act in 1971, effective in most states in 1972. The Act specifies the reasons for which an abortion can be legally performed and requires that it be performed by a registered medical practitioner in certified abortion facilities.[10] Abortion is legal if the pregnancy that it terminates endangers the

[7] Kalsi (2015) shows that low SES girls (but not boys) born at higher birth orders after the legalization of abortion in Taiwan were more likely to attend university.

[8] In comparison, before the 1985 abortion legalization that Lin et al. (2014) examine, the average post-neonatal infant EFM in Taiwan in 1981 was -0.11 p.p., i.e., small, and biased in favour of girls (Yang et al. (1996)). India and China differ in the age distribution of missing girls. In China, the imbalance in the sex ratio for under-5 children in primarily at birth (83 percent), while there is a more even spread across early childhood in India (Table A.2). In particular, the contribution of post-infancy EFM to the number of missing girls is 33 percent in India but only 3 percent in China.

[9] Bharadwaj and Lakdawala (2013) find that male fetuses receive higher prenatal investments in India too, suggesting that revelation of the sex of the child may lead to greater miscarriage of female fetuses or higher post-ultrasound neonatal mortality for girls that are not aborted. However, we find no significant impact of ultrasound access on neonatal mortality suggesting that this channel is not dominant.

[10] More information on the certification criteria is available in Stillman et al. (2014).

woman's life, causes grave injury to her physical or mental health, is a result of rape or contraceptive failure (the latter applies only to married women), or is likely to result in the birth of a child suffering from serious physical or mental abnormalities. Consent is not required from the woman's husband or from other family members; however, a guardian's consent is required if the woman seeking an abortion is either less than 18 years old or is mentally ill. The Act allows an unintended pregnancy to be terminated up to 20 weeks' gestation; however, if the pregnancy is beyond 12 weeks, approval is required from two medical practitioners (Arnold et al. (2002)). The stated purpose of the Act was to regulate and ensure access to safe abortion, although it has been argued that the political motivation was population control (Phadke (1997)).

Fetal sex determination first became possible in India with the advent of amniocentesis in the 1970s. This technology was introduced to detect genetic abnormalities but was soon being used to detect fetal sex. As early as 1976, the government banned the use of these tests for sex determination in government facilities (Arnold et al. (2002)). The private sector remained unregulated but widespread use was limited by the high direct cost and the invasiveness of amniocentesis. Fetal sex-selection only really became feasible after 1980—becoming evident at the population level after 1985 and widespread by 1995—with the arrival of ultrasound scanners. Ultrasound availability during the early diffusion period was driven by the liberalization of India's import sector. The first ultrasound scanner was imported in 1987 (Mahal et al. (2006)). Thereafter, the quantity of imports rapidly increased (Figure 1) as import duties on medical equipment were gradually lowered. The import tariff on medical devices declined from 40 to 60 percent in the 1980s to 25 percent in the late 1990s to 12.5 percent in 2003-04, and then to the currently uniform rate of 5 percent. Domestic production of ultrasound machines grew 15-fold between 1988 and 2003 (George (2006), Grover and Vijayvergia (2006)), following relaxation of industrial licensing regulations. The bottom graph in Figure 1 shows that, once domestic production began, it was orders of magnitude larger than imports.

Demand for ultrasound scans proliferated as a result of the technology being non-invasive and its wide affordability at about $10-$20 for a scan or an abortion (Arnold et al. (2002)).[11] The trend in ultrasound use (also in Figure 1) closely tracks the supply of ultrasound machines.[12] Additionally, Figure 2 shows that the officially reported number of abortions (that includes both sex-selective and other abortions) follows a similar trend and is positively correlated with self-reported ultrasound use during pregnancy. Clinics and portable facilities have mushroomed, advertising availability of ultrasound with slogans conveying that the cost of a scan is much lower than the future costs of dowry.[13] Additional amendments to the MTP Act in 2002 and 2003 increased public sector provision

[11] The costs cumulate if repeated scans and abortions are needed before a boy is conceived and vary with distance of the household from the clinic and with the safety of the procedures.

[12] We infer the trend in ultrasound use in Figure 1 by plotting the fraction of births in a year for which the mother reports getting an ultrasound test at some point during the pregnancy. However, this excludes pregnancies that resulted in abortion, and therefore underestimates the ultrasound usage rate, but is likely to reflect the trend accurately.

[13] Dowry is a ubiquitous feature of the Indian marriage market and payments from the bride's family to

and made abortion safer (Stillman et al. (2014)). Other things equal, this could have contributed to a further increase in feticide since 2002.

Since the late 1980s, sex-selection has become the dominant concern amongst women's and human rights organizations.[14] Their campaigns led to the central government passing the Prenatal Sex Diagnostic Techniques (Regulation and Prevention of Misuse) (PNDT) Act in 1994. This act was effective from January 1, 1996. The PNDT Act made it illegal to use prenatal sex-diagnostic techniques (like ultrasound) to reveal the sex of a fetus. Following the revelation in the 2001 Census of a continuing deterioration in the sex ratio, the PNDT Act was strengthened by a 2002 Amendment (effective 2003) incorporating a ban on advertising prenatal sex determination and increased penalties for violations.[15] It is widely believed that these regulations have made little difference (Visaria (2005)), although Nandi and Deolalikar (2013) find that they did have some impact. These bans are difficult to enforce because ultrasound (or alternatives like amniocentesis) are also used for medical purposes and in routine prenatal care, making it easy to cover up sex determination as a motive.

Before we move on to the empirical strategy, it is useful to also consider whether we expect an *underlying* trend in the sex ratio at birth, i.e., a trend in the absence of parental interventions, and in what direction. In general, the fetal environment has improved in India. The growth in income and the decline in poverty since the early 1980s has been widely documented; fertility decline set in from 1981 (Bhalotra and van Soest (2008)); and neonatal mortality rates have been decreasing. Maternal mortality is estimated to have declined (Bhat (2002)) and maternal age at birth has risen. These trends will all have led to improvements in fetal health. Indeed, the availability of ultrasound scans that monitor fetal development will also have contributed to better fetal health. Bio-medical studies show that fetal health improvements tend to favor boys (Trivers and Willard (1973), Almond and Edlund (2007)) and thus the underlying tendency, if any, would be for male child survival to increase over time. Our hypothesis, on the other hand, is that the availability of ultrasound scanners has improved female child survival relative to male children.

3 Empirical Methodology

The hypothesis we test in this paper is that the availability of prenatal sex-detection technology (in particular, ultrasound), which led to large-scale selective abortion of female fetuses, also modified

the groom's family at the time of marriage can amount to several multiples of annual household income (Anukriti et al. (2016)), and can motivate parents to eliminate female births (Bhalotra et al. (2016a)).

[14]Feminist and socialist groups in the United States and other richer countries have hotly defended a pro-choice stance against a pro-life stance on abortion. The focus of public discussion is on benefits for women rather than on benefits for children. For instance, `http://www.theguardian.com/commentisfree/2014/oct/14/abortion-right-to-privacy-women-right-to-equality` and `https://socialistworker.org/2013/11/01/abortion-every-womans-right`. Indian feminists, on the other hand, have been divided by the seeming contradiction of supporting a woman's right to abortion while opposing sex-selective abortion (Kumar (1983), Gangoli (1998)).

[15]More details on the PNDT Act are available in Retherford and Roy (2003) and Visaria (2005).

conception and investment decisions in a gendered manner, with implications for EFM.

As discussed in Section 1, we exploit exogenous variation in the supply of ultrasound scanners. The first supply shock was the first imports of ultrasound scanners in the early stages of India's import liberalization (tariff reduction), and the second was initiation of local production following industrial delicensing policies.[16] However, the wave of economic liberalization in India that led to reduced import tariffs and relaxed licensing of domestic production was potentially correlated with other trends, for instance, with greater exposure of women to Western media, and with rising incomes for large sections of the population. We therefore interact birth cohort variation in exposure to the new technology with the sex of the firstborn child of the mother, a previously well-tested proxy for the proclivity to commit sex-selection. Since we are interested primarily in gender gaps in parental investments, survival, and sibship size, we use the sex of the child as a third interaction in a triple differences-in-differences (DDD) regression specification.

The identification challenge is that families that take up prenatal ultrasound scans and obtain sex-selective abortion may be different from families that do not abort girls even when they can. For instance, we know that more educated and wealthier women are more likely to engage in sex-selective abortion in India (Bhalotra and Cochrane (2010)). As a result, post-ultrasound, the average girl is more likely to be born into a low-SES family and girls' outcomes will be worse even if parental investments are unchanged.[17] By leveraging variation in firstborn child's sex, we obtain variation in the risk of sex-selection that is independent of SES. It is therefore key to our identification strategy to demonstrate that (a) the sex of the firstborn child is quasi-random, (b) pre-ultrasound trends in outcomes did not differ by child gender, and (c) firstborn child sex predicts the proclivity to commit sex-selection (conditional on underlying preferences). We now demonstrate that these three conditions hold.

Testing exogeneity of firstborn sex. The assumption that the sex of the first child is randomly determined is supported by the data.[18] We demonstrate this in three different ways. The top left graph in Figure 3 shows that the proportion of females among first births in India lies within the normal range (48.8 percent - 49.26 percent) during our sample period, and shows no tendency to increase over time.[19] Table 2 shows that families with firstborn boys and firstborn girls are similar along a number of observables (nonetheless, we also investigate a mother fixed effects specification to account for any unobservable differences).[20] Third, we regress firstborn sex on indicators for

[16]We do not use measures of state-specific penetration or adoption as they would be endogenous; see Jayachandran et al. (2010) for a similar argument pertaining to the introduction of antibiotics in the United States. Bhalotra and Cochrane (2010) document that access to ultrasound in India was widespread and that the costs of ultrasound and abortion were not prohibitive even for relatively poor households.

[17]Since we find that girls' outcomes improve, failure to control for SES would lead us to under-estimate impacts. We do control for various SES characteristics, including education and wealth, but here our point is to illustrate the nature of the potential identification problem.

[18]We describe the data in more detail in Section 4.

[19]Figure 3 reproduces Figures 1-4 from Bhalotra and Cochrane (2010).

[20]Table A.5 reports the corresponding summary statistics for the REDS sample. The definition of pre-ultrasound and early diffusion period is the same as in Table 2 but the late diffusion period is shorter

post-ultrasound cohorts and various controls (state fixed-effects, state-specific linear time trends, and a set of SES variables), and find that it is not significantly predicted by these variables (see Table A.3). Exogeneity of firstborn sex has also been previously defended (Das Gupta and Bhat (1997), Visaria (2005), Bhalotra and Cochrane (2010)) and lines up with recent survey data that suggest that parents do not always prefer having a son over a daughter. Jayachandran (2016) finds that although the vast majority of families want to have a son if they can only have one child, at a family size of two they prefer having one daughter and one son over having two sons. As desired and actual fertility in India are well above one (Table A.4), it is reasonable to assume that parents are not averse to having one girl, despite a strong desire for at least one boy. In our sample, on average, desired fertility is 2.7; the desired number of sons is 1.38; and the desired number of daughters is 1.01.

Pre-trends not different by firstborn sex. In order to test for a significant difference in under-5 mortality trends for firstborn-boy and firstborn-girl families in the pre-ultrasound period, we restrict the sample to the pre-ultrasound period, 1973-1984, and regress an indicator for under-5 mortality on the full set of interactions between indicators for firstborn girl (G_j), female birth (F_i), and year of birth, with and without fixed effects for birth order and state and examine the coefficients of the triple-interactions. The two graphs in Figure 4 demonstrate no significant divergence between EFM in firstborn-boy and firstborn-girl families for pre-ultrasound cohorts.

Firstborn sex predicts sex-selection at higher parities. Previous studies, such as Bhalotra and Cochrane (2010) and Rosenblum (2013), have established that parents randomly exposed to a "firstborn-girl treatment" are more likely to practice sex-selection at higher-parity births, consistent with a documented desire for at least one son. Figure 3 clearly depicts this pattern: after ultrasound technology became available, second, third, and fourth births became increasingly male but *only* for families without a son.[21] So, the interaction with first child's sex captures the differential incentives to sex-select among otherwise similar families.

Unconditional trends in outcomes. Table 1 shows that in the pre-ultrasound era, 1973-1984, averaging across families with a firstborn girl, girls were 2.87 p.p. more likely than boys to suffer under-5 mortality. After 1985, following introduction of ultrasound, the gender gap in mortality declined for this group, to 1.59 p.p. during 1985-1994 and to 1.09 p.p. during 1995-2005. In comparison, EFM in families with a firstborn son was close to zero pre-ultrasound (indeed, it was negative at -0.08, vastly different from 2.87 p.p.). Moreover, there is no evidence that post-ultrasound mortality reductions favored girls in this group. This is compelling evidence showing that (a) the baseline problem of EFM was entirely restricted to families with firstborn girls and (b) post-ultrasound reductions in mortality favored girls in this group, narrowing the EFM gap among firstborn-girl families and hence overall. As we shall observe in Section 7, ignoring the distinction

(1995-1999) because we use the 1999 survey.

[21] The graphs look similar if we compare the proportion of females among third and fourth births by firstborn sex instead.

between families by the sex of the firstborn child makes it much harder to identify the variation of interest. Henceforth, we think of children in families with firstborn daughters as the "treated" group (i.e., treated by availability of ultrasound).

Table 1 also shows that the same broad patterns hold for parental investments and fertility. Thus, the sex of the firstborn child is a predictor of gender gaps not only in mortality but also in child investments and fertility.

3.1 Regression Specifications

To capture the time variation in ultrasound availability, we split our sample into three broad time-periods, defining 1973-1984 as the pre-ultrasound period, 1985-1994 as the early diffusion period, and 1995-2005 as the late diffusion period when ultrasound supply and use became widespread. Using non-parametric plots and flexible parametric specifications, Bhalotra and Cochrane (2010) confirm that the trend in the sex ratio at birth breaks in line with the availability of ultrasound. Recall that 1985 and 1995 mark, respectively, the first imports and the initiation of local production of ultrasound scanners (see Figure 1). We nevertheless confirm that our results are similar if we vary the precise thresholds used to define the three time periods.

3.1.1 Mortality and Health Investments

For child i of birth order b born to mother j in year t and state s, we estimate the following ordinary least-squares (OLS) specification:[22]

$$\begin{aligned} Y_{ibjst} =& \alpha + \beta_1 G_j * F_i * Post_t^1 + \beta_2 G_j * F_i * Post_t^2 \\ &+ \gamma G_j * F_i + \omega_t G_j + \sigma_t F_i \\ &+ \mathbf{X'_{ijt}} \tau + \delta_s F_i + \nu_s G_j + \psi_b F_i + \xi_b G_j \\ &+ \rho_{bt} + \eta_{bs} + \phi_{st} + \epsilon_{ibjst} \end{aligned} \tag{1}$$

The dependent variable, Y_{ibjst}, is either a mortality indicator for child i or a measure of parental investments in children, including breastfeeding and immunization.[23] The indicator variable G_j equals one if the first child of mother j is a girl. The variable F_i equals one if child i is female. $Post_t^1$ indicates that t belongs to the early diffusion period (1985-1994) and $Post_t^2$ indicates that t belongs to the late diffusion period (1995-2005). Attached to the two triple interaction terms are the coefficients of interest, β_1 and β_2. For regressions where mortality before x months or years is the outcome, we exclude children that are less than x months or years old from the sample to allow each child "full exposure" to the risk of mortality.

A vector of socioeconomic and demographic characteristics, $\mathbf{X_{ijt}}$, comprises indicators for household wealth quintiles, educational attainment of child's parents, mother's birth cohort, mother's

[22]The variable state refers to the mother's state of residence at the time of survey and may differ from the child's state of birth. Restricting the sample to women who have not migrated between their first birth and the survey date does not substantively change the estimates.

[23]More details on the variables used in the regression analysis are available in Appendix B.

age at birth, caste, religion, and residence in a rural area. We also control for the main effects of G_j and F_i and fixed effects for state, birth year (or cohort, of the child), and birth order. We allow birth cohort fixed effects to vary by firstborn sex ($\omega_t G_j$), by child gender ($\sigma_t F_i$), and by birth order (ρ_{bt}), which provides a flexible suite of controls for possibly omitted trends. We also allow state and birth order fixed effects to vary by firstborn sex and by child gender ($\delta_s F_i$, $\nu_s G_j$, $\psi_b F_i$, $\xi_b G_j$), and state fixed effects to vary by birth order (η_{bs}) and birth year (ϕ_{st}).

Child gender-specific cohort fixed effects ($\sigma_t F_i$) account for any nationwide changes that may influence gender gaps in the outcomes, including improvements in maternal health or prenatal care which we expect to benefit male fetuses more than female fetuses given the evidence on greater sensitivity of males to prenatal inputs (Low (2000)). They also account for any trends associated with modernization. Cohort fixed effects varying by sex of the firstborn child in the family ($\omega_t G_j$) control for nationwide trends that may have differentially affected mortality of children in firstborn-girl versus firstborn-boy families. For instance, the Trivers-Willard hypothesis (Trivers and Willard (1973)) implies that firstborn-boy families are more often of higher SES than firstborn-girl families, and it is plausible that trends in the outcomes differ by SES and that the SES-observables we control for do not capture every relevant expression of SES.

Allowing the state fixed effects to vary with both child gender ($\delta_s F_i$) and sex of the firstborn child ($\nu_s G_j$) allows state-level time-invariant factors, such as soil quality (Carranza (2015)), to have gender-specific effects and ensures that we absorb any cross-sectional heterogeneity that may be correlated with firstborn sex. We interact indicators for child gender and sex of the firstborn with birth order ($\psi_b F_i$, $\xi_b G_j$) given previous evidence that son preference varies with birth order and that sex of the firstborn child influences the exercise of son preference. Lastly, ϕ_{st}, η_{bs}, and ρ_{bt} control non-parametrically for, respectively, state-specific time effects (e.g., differential growth rates of state GDP or availability of abortion and other health services), state-specific birth order effects, and birth order specific time effects.

Specification (1) is estimated for second- and higher-order births; the "control" group thus comprises pre-ultrasound births and second- and higher-order births to mothers whose firstborn is a boy. The coefficient γ measures the difference in EFM or the gender gap in health investments between the treatment and control groups during the pre-ultrasound period. The coefficients β_1 and β_2 capture how these gaps evolved over the early and late diffusion periods relative to the pre-ultrasound period. Standard errors are clustered by state.

As discussed, we do not observe large differences in the socioeconomic characteristics of firstborn-boy and firstborn-girl families (Table 2), so selection on firstborn sex is of limited concern. We nevertheless condition upon SES (education, wealth, caste, rurality); and we effectively difference between firstborn-girl and firstborn-boy families. To ensure that there remain no unobservables that are potentially correlated with firstborn sex and that influenced girl outcomes differently from boy outcomes in the post- relative to the pre-ultrasound era, we introduce mother fixed-effects in specification (1). While recognizing that the self-reported fertility preference variables are potentially endogenous (being asked during or after the fertility process), we also consider sensitivity of the estimates to conditioning upon mother's stated desired fertility and desired sex composition of children.

3.1.2 Fertility

We examine the impact of prenatal sex-detection technology on gender gaps in fertility in two ways. First, we test if ultrasound altered the male-bias in the hazard of birth in a given year for firstborn-girl versus firstborn-boy mothers.[24] For this specification, we utilize a retrospective mother-year panel in which a woman enters the panel in her year of marriage and exits in the year of survey.[25] For mother i from state s of age a in year t, who has given birth to $b-1$ children by t and whose last birth took place r years ago, we estimate the following logistic regression:

$$\begin{aligned} Birth_{it} =& \alpha + \beta_1 G_i * Post_t^1 + \beta_2 G_i * Post_t^2 + \gamma G_i + \omega_t \\ &+ \mathbf{X_i}'\tau + \phi_a + \psi_b + \sigma_r + \delta_s + \nu_s G_i + \theta_{st} + \epsilon_{it} \end{aligned} \tag{2}$$

The outcome variable, $Birth_{it}$, equals one if the mother gives birth in year t and is zero otherwise. $Post_t^1$, $Post_t^2$, and G_j are defined as earlier. The vector $\mathbf{X_i}$ comprises indicators for household wealth quintiles, educational attainment of the mother and her husband, caste, religion, residence in a rural area, and mother's year of birth. We include fixed effects for year (ω_t), state (δ_s), mother's age (ϕ_a), parity (ψ_b), and years since last birth (σ_r), state-specific firstborn-girl fixed effects ($\nu_s G_j$), and state-specific year fixed effects (θ_{st}).[26]

Additionally, we estimate the effects of ultrasound availability on the "stock" of children a woman has at the time of survey.[27] Specifically, we estimate the following specification for woman j in state s who has N_{jt} children in the year of survey, t:

$$\begin{aligned} N_{jt} =& \alpha + \beta_1 G_j * Post_t^1 + \beta_2 G_j * Post_t^2 + \gamma G_j + \sigma Post_t^1 + \psi Post_t^2 \\ &+ \mathbf{X_j}'\tau + \delta_s + \nu_s G_j + \theta_s Post_t^1 + \omega_s Post_t^2 + \epsilon_{jt} \end{aligned} \tag{3}$$

We restrict the sample to mothers who either were always exposed or never exposed to ultrasound for the year-span of their births. In other words, we retain women who had all their births strictly within one of the three time-periods—pre-ultrasound, early diffusion, or late diffusion. $Post_t^1$ and $Post_t^2$ indicate that a woman began and completed childbearing respectively during 1985-1995 and after 1995. The variable G_j is, as before, an indicator for the firstborn being a girl. The vector $\mathbf{X_j}$ comprises indicators for household wealth quintiles, educational attainment of the woman and

[24] Since our empirical strategy relies on the sex of the first birth, the sample excludes the 11 percent of women in the data who had never given birth by the time of the survey.

[25] Out of wedlock birth in India is rare (0.06 percent in 2005 National Family Health Survey).

[26] We also modify this specification by including mother fixed effects to test if ultrasound availability delays time to the next birth for a given mother, conditional upon the time since last birth, in the post-ultrasound (relative to the pre-ultrasound) period and whether this delay is on average greater in firstborn-girl (relative to firstborn-boy) families. The results from this specification are available upon request.

[27] Like specification (2), here too we exclude women who had never given birth by the time of the survey.

her husband, caste, religion, residence in a rural area, woman's birth year, and woman's age at the time of survey. We include fixed effects for the woman's birth year as fertility is right-censored for some women. Moreover, we include state fixed effects (δ_s), state-specific firstborn-girl fixed effects ($\nu_s G_j$), and allow the effects of the post-ultrasound indicators to vary by state ($\theta_s Post_t^1$ and $\omega_s Post_t^2$). The coefficient γ provides an indication of the extent to which the pre-ultrasound period was characterized by son-biased fertility stopping. The coefficients β_1 and β_2 test our hypothesis that, after fetal sex determination became feasible, there was less son-bias in fertility decisions; i.e., the fertility difference between families with a firstborn girl and families with a firstborn boy narrowed.[28]

4 Data and Descriptive Statistics

The mortality and fertility equations are estimated using three pooled rounds of the National Family Health Survey (NFHS) conducted in 1992-93, 1998-99, and 2005-06. These nationwide, repeated, cross-sectional surveys are representative at the state level and report complete birth histories for all interviewed women, including children's month, year, and order of birth, mother's age at birth, and age at death of deceased children.[29] The sample comprises 503,316 births of 232,259 mothers that occurred in 1973-2005.

Although the NFHS contain fairly rich data on investments in children, these questions are asked only for children born in a few years preceding each survey and there are no pre-ultrasound data on investments.[30] For this reason, we utilize the 1999 round of the Rural Economic and Demographic Survey (REDS) that reports data on health investments for all children alive at the time of survey.[31] To the extent that deceased children are likely to have received lower health investments than surviving children and (as we show) excess girl mortality was higher during the pre-ultrasound period than in the post-ultrasound years, the exclusion of deceased children will tend to bias the estimated effects downward. In this sense, the estimates we present are conservative. We nevertheless also present estimates using the NFHS data, which is not biased by the exclusion of deceased children.

For the mortality and postnatal health specifications, we pool all births of the surveyed women to create a child-level dataset. The birth hazard sample is a retrospective mother-year panel that we

[28] In principle, the excluded mothers, whose fertility spans more than one period should be similar to the included mothers. We checked for balance and found that, in fact, the excluded mothers are, on average, of lower SES and are older. However, we always control for SES characteristics and flexibly control for age, and, within the sample, our estimates identify differences by firstborn sex. In any case, we also present fertility results separately for each SES group (within our sample) and, as we show later, for most SES indicators, the differences in the coefficients of interest are small. Lastly, our results are robust to the inclusion of fixed effects for the years of first and last birth for a mother and fixed effects for their interactions with each other, with state fixed effects, and with the firstborn sex indicator.

[29] Since the state of Sikkim changed its border during the period of analysis, we exclude it from our sample.

[30] NFHS-1, 2, and 3 collected health investments for, respectively, the last three children born after January 1988, the last two children born after January 1995, and all children born after January 2001.

[31] REDS is restricted to rural women from 16 major states.

create from women's retrospective birth histories. The sample used for the fertility stock estimates simply pools all surveyed women to form a woman-level dataset. Summary statistics for mortality and (actual) fertility are in Table 1.

Measurement error due to imperfect recall. One may be concerned that a 20-year recall period introduces measurement error in mother reports of the gender of births. For instance, women may omit or forget to mention children that died many years ago and, if such misreporting is correlated with child gender and increasing in distance of birth year from survey year, then it could, in principle, bias our findings. The firm gathering the NFHS data was alert to this potential problem, and the data have been subject to several probes designed to assess this. Nevertheless, since recall error is likely to increase with distance from the survey, we examined the extent of this issue by exploiting the fact that a range of birth years is available in more than one NFHS round and that distance between a given year and the survey date varies by round. For instance, we may expect greater reporting bias for the year 1988 in the survey conducted in 2005-06 than in the survey conducted in 1992-93. Figure A.1 shows that the reported sex ratio at birth does not vary by survey round.

Maternal survival selection. The NFHS gathers birth (and child mortality) history data from women in their reproductive years. It naturally excludes women who have died, for example, from maternal (birth-related) mortality. Sex-selective abortion in India has been illegal since 1996 (see Section 2), but illegal abortion is conducted on a large scale and is often unsafe, elevating maternal mortality risk. Thus, if maternal mortality scales with the practice of (selective) abortion, the fact that we never observe women who die before the survey means that our estimates will be under-estimates. We must also consider any bias in our estimates that arises because the unsafe abortions and the underlying maternal mortality risks are concentrated among women of low SES. Below, we will show that the post-ultrasound decline in EFM is larger among low SES women (who exhibit higher pre-ultrasound EFM). So if maternal mortality removes a disproportionate share of low SES women, our results under-estimate the average impacts of ultrasound availability on EFM.

5 Results

5.1 Excess Female Mortality

Under-5 Mortality. In Table 3, we present estimates of the impacts of the introduction and diffusion of ultrasound technology on under-5 mortality. We add controls as we move across columns, with columns (3) and (4) being the richest specifications without and with mother fixed effects, respectively. Panel A presents the results for specification (1). The coefficient of *Firstborn girl * Female* confirms that, during the pre-ultrasound period, girls were significantly more likely to die before age 5 (by 2.768 p.p. in column (3)) among children preceded by a firstborn sister relative to a firstborn brother.[32] The triple-interaction coefficients, *Firstborn girl * Female * Post1* and

[32] In contrast, under-5 mortality for boys in the pre-ultrasound period was smaller in firstborn-girl families. It was 12.90 percent in firstborn-boy families and 10.49 percent in firstborn-girl families.

*Firstborn girl * Female * Post2* together imply that *all* of the baseline EFM gap between firstborn-girl and firstborn-boy families (i.e., 2.768 p.p.) was eliminated once ultrasound technology became available.[33] In Panel B of Table 3, we replace the two post-ultrasound indicators that indicate the start of ultrasound imports ($Post1$) and the beginning of domestic production ($Post2$) with one dummy variable that captures the divide between the pre- and the post-ultrasound regime, i.e., $Post$. Like Panel A, we find that ultrasound availability led to a significant decline in the EFM gap between firstborn-girl and firstborn-boy families across all columns, closing the pre-ultrasound gap by 60 percent. The coefficients of interest are robust to the inclusion of mother fixed effects in column (4). These coefficients are identified using a sample in which women have children of mixed gender at second or higher parities.[34] Note that the inclusion of mother fixed effects does not imply that we have removed the influence of all compositional change.

Neonatal Mortality. We do not find any significant impact on gender differences in neonatal mortality, although the triple-interaction coefficients are consistently negative (see Table A.6). Thus, we can decisively reject an *increase* in relative girl neonatal mortality, which could arise if, having detected child sex, parents made smaller fetal investments in girls. In fact, the decline in neonatal EFM (although insignificant) biases us against finding a decline in under-5 EFM because, in the post-ultrasound cohorts, the marginal survivors of the neonatal period are negatively selected (i.e., in the pre-ultrasound period they would have succumbed to neonatal mortality, but now they "just survive").

Larger reductions in under-5 than in neonatal mortality are consistent with the increases in postnatal investments that we document below, given that neonatal survival is less dependent upon postnatal investments and is more closely linked to maternal health and delivery conditions. Our findings are also congruous with previous work; for instance, Almond et al. (2006) show that hospital de-segregation after the Civil Rights Act led to a narrowing of the racial gap in post-neonatal but not in neonatal mortality.[35] The significance of our results is enhanced by the fact that reductions in post-neonatal mortality also improve later life circumstances, predicting adult height, a marker of health (Bozzoli et al. (2009)) and cognitive performance (Chay et al. (2009)).

Extensions. Estimates by birth order are in Table A.7. The only statistically significant reductions in under-5 EFM are among second births; although there are sizable but less precisely estimated reductions among third births too. This may reflect the common finding that parents are particularly averse to having more than two girls (which applies at order three and above),[36]

[33] To determine what percentage of the pre-ultrasound EFM gap by firstborn sex is eliminated post-ultrasound, we compare the sum of the two triple interaction coefficients with the coefficient of *Firstborn girl * Female*.

[34] Balance tests show that this sample has higher fertility so this is just an additional specification test.

[35] Similarly, The Million Deaths Study (2010) in India shows that only 3.2 percent of neonatal deaths were caused by diarrhea—a function of clean water and nutrition—in contrast to 22.2 percent of post-neonatal deaths.

[36] Almond and Edlund (2008) show that Indian, Chinese, and Korean families with no previous sons exhibit male-biased sex ratios at third parity but not before in the 2000 US Census and Bhalotra and Cochrane

so there remains a girl-boy differential in under-5 mortality at higher orders.

We may be concerned that the sharp changes in ultrasound availability occasioned by economic liberalization policies in India coincided with changes in women's education and job opportunities and, hence, with changes in preferences over fertility and the sex composition of births. Although there is no evidence of sharp (as opposed to smooth) changes in women's achievements and opportunities, we investigate this concern by including controls for women's self-reported preferences (ideal number of children and ideal sex ratio of children) in specification (1). We also interact the preference variables with the *Female* dummy and the *Firstborn girl* dummy. Although these variables are often measured with error, it is notable that the coefficients of interest are not discernibly changed by these additions (Table A.8). In column (2) of Table A.8, we control for another measure of gender biased preferences—the state-year gender enrollment ratio at ages 6-11 and 11-14—and, again, the results are stable. The coefficients of these additional terms are interesting in their own right, and are consistent with our expectations. The coefficients of *Ideal Sex Ratio * Female* and *Enrollment Gender Gap * Female* are positive, implying that EFM is higher in families with a stronger preference of sons. The coefficient on *Ideal Fertility * Female* is negative, implying that, conditional on son preference, a declining trend in fertility leads to higher EFM, a finding that is consistent with Jayachandran (2016).

Mechanisms. As discussed in Section 1, availability of ultrasound scans led to an increase in abortion of female fetuses by parents who did not want a girl (and the data suggest that a greater proportion of such parents were those who, by a random act of nature, had daughters at first birth). We have shown that the marginal girl who survives to birth post-ultrasound is also more likely to survive to age five. This is because she is born into a family that wants her; since families that do not want girls now tend to abort them. In this sense, the estimated effects on EFM are driven by the composition of parents who have girls. Put differently, ultrasound availability does not affect parents' preferences but, rather, how girls are matched to parents.[37] Since the data contain self-reported son preference, we investigated this and found no post-ultrasound change in son preference in first-girl relative to first-boy families (see Table A.9).[38]

In the following sections, we investigate proximate mechanisms, i.e., the behaviors that operate in the treated (firstborn-girl) families that have daughters post-ultrasound. The hypothesis is that, by virtue of being more wanted, post-ultrasound daughters in firstborn-daughter families receive higher parental investments than girls born pre-ultrasound, narrowing investment gaps in these

(2010)) show that the male-bias in the sex ratio at birth in India is increasing in birth order.

[37] Note that in the mother fixed effects specification, differences in under-5 survival between pre- and post-ultrasound daughters are still driven by parents desiring the later born daughter (which they chose to have) more than the first (which they could not choose) on average.

[38] As an indirect test of whether preferences changed, we considered exploiting the random variation in firstborn sex to look at whether EFM among first born children changed post-ultrasound but this is not a clean test because the outcomes of firstborn children in first-boy versus first-girl families may have changed post-ultrasound because of spillovers to firstborns, for example, from fertility decline which relaxes resources constraints.

families between sons and daughters. Similarly, parents who achieve their desired sex-mix of births using sex-selective abortion in the post-ultrasound period have no need to use son-biased fertility stopping.[39]

5.2 Postnatal Health Investments

We posited above that declining under-5 mortality among girls signals increased parental investments. We directly test for this in Table 4. The outcomes are the number of months a child is breastfed; a dummy variable indicating that the child has received at least one vaccine; and medical expenditure (in Rupees) on the child in the year prior to the survey.[40] Pre-ultrasound, in families with a firstborn girl, boys were breastfed for a longer duration and were more likely to be vaccinated and to receive expenditure during illness. The estimates show that the gender gaps in breastfeeding and vaccination were virtually eliminated post-ultrasound in firstborn-girl families, although in some cells we lose significance for breastfeeding coefficients due to insufficient precision. The coefficients for medical expenditure during sickness are also positive but imprecise and only significant in one column. A previous literature estimates that breastfeeding differences explain about 9 percent of the gender gap in post-neonatal child mortality in India (Jayachandran and Kuziemko (2011)) and that sex differences in vaccinations explain between 20 - 30 percent (Oster (2009)). We estimate that the contributions of breastfeeding and vaccination to the ultrasound-led decline in EFM are, respectively, 20 percent and 7 to 11 percent (details in Appendix C).

For reasons discussed in Section 4, we also present results using the NFHS data in Table A.10. The NFHS contains information not only on immunization and breastfeeding, but also reports the number of antenatal checks during pregnancy. The results are broadly similar.[41] Estimates by birth order (available upon request) are also congruous with the birth order specific results for survival. We also found a tendency for ultrasound access to decrease pre-existing gender gaps in children's school enrollment (using NFHS data and specification (1)), although the relevant coefficient is not statistically significant.[42]

[39]Of course, only a fraction of families with firstborn daughters do not want subsequent girls and only a fraction of those will commit abortion. The contention is that, after ultrasound becomes available, a larger fraction of those families that do have daughters want them.

[40]Medical expenditure is conditional upon illness and includes doctor's fees, medicines, and costs of special diets during the illness. The specifications are similar to those for mortality except that we drop the urban indicator (since REDS covers only rural households) and the wealth quintiles (not reported in REDS), and, since we have a smaller sample, we drop ρ_{bt}, $\xi_b G_j$, and $\nu_s G_j$ and replace ϕ_{st} with state-specific linear time trends.

[41]The estimates of gender gaps during the early diffusion period in Table A.10 are similar to the corresponding numbers (i.e., sums of the coefficients in the first two rows) in Table 4. The coefficients in the second row of Table A.10 capture the differences in gender gaps during the two post-ultrasound periods, and are also similar to the corresponding estimates (i.e., differences of the coefficients in the second and third rows) in Table 4. For instance, column (2) of Table A.10 implies that the gender gap in the likelihood of receiving at least one vaccine was 2.5 p.p. lower in the late diffusion period relative to the early diffusion period, while the corresponding magnitude is 1.18 p.p. in column (6) in Table 4.

[42]Public school education is free in India, and thus parents may not discriminate in terms of sending their children to school unless they have a demand for child labor. Gender bias may manifests in the quality of

Our strictest specifications identify gender gaps and do not directly deliver an estimate of the absolute gains or losses for boys in firstborn-girl families (which would require examining the coefficients of $Firstborn\ girl$ x $Post$ indicators) because we include among controls $Firstborn\ girl$ x $Year$ fixed effects ($\omega_t G_j$). Replacing $\omega_t G_j$ with $Firstborn\ girl$ x $Post$ indicators in specification (1) shows that, in firstborn-girl families, not only girls but boys too received greater immunization post-ultrasound (though the magnitude of gains is smaller for boys than for girls), while breastfeeding duration was only significantly extended for girls. These results are consistent with smaller sibship sizes in firstborn-girl families.[43]

5.3 Fertility

For reasons detailed earlier, in Tables 5 and 6 we investigate if son-biased fertility stopping behavior changed subsequent to the availability of sex-selection technology. The coefficient of *Firstborn Girl* is positive and significant in both tables confirming that, pre-ultrasound, women whose first child was a girl were 15.5 p.p. more likely to give birth in a given year and had 0.155 more births than women with a firstborn son. Our estimates show that these differentials were eliminated once ultrasound technology became available.[44] Note that the coefficients in Tables 5 are coefficients of the index function; column (2) implies that the gap in the predicted yearly probability of birth between firstborn-girl and firstborn-boy families declined from 0.033 in the pre-ultrasound period to -0.004 in the late diffusion period ($Post2$). Controlling for mother's fertility preference and son preference does not significantly alter these effects.[45]

Since sex-selective abortion allows parents to avoid unwanted children, we also test whether availability of ultrasound drove actual fertility closer to desired fertility. Our estimates in Table 6 imply a complete elimination of undesired fertility, which the first row shows was 0.117 births in first-girl relative to first-boy families. To put this in a wider perspective, the presence of HIV reduces the average number of births a woman has during her life-cycle by 0.15 (Shapira (2013)). The coefficients of interest (those on the triple-interaction terms) are similar for actual fertility (in column 2) and for actual minus desired fertility (in column 3). This again confirms that the decline in actual fertility in first-girl families that we document is not driven by a decline in desired fertility.

The coefficients of the stated preference terms reveal that actual fertility is, as we may expect, increasing in desired fertility and in the desired ratio of sons to daughters. However, excess fertility is decreasing in the ideal ratio of sons to daughters, suggesting that desired fertility rises more

schooling (private versus public schools (Azam and Kingdon (2013)) and enrollment in higher education; however, we do not have adequate data on these dimensions.

[43]These results are available upon request. We find no significant change in boy survival alongside increases in girl survival in firstborn-girl families, however.

[44]Estimates conditional upon mother fixed effects are not statistically significantly different, and are available on request.

[45]As a robustness check, like for EFM, we also ran a specification where $Firstborn\ girl$ * $Post1$ and $Firstborn\ girl$ * $Post2$ are interacted with the preference variables; the results remain the same.

steeply with son preference than actual fertility.[46] Again, this makes sense, given that the desire to have sons leads to the widely documented phenomenon of son-biased fertility stopping (which is evident from the baseline statistics in the first row). Since actual fertility is not fully in the control of parents, it will tend to rise less than proportionately with desired fertility. Table 7 shows that the relative fertility decline in firstborn-girl families is driven by a shift from having four or more children toward having two or three children.

5.4 Heterogeneity

Although stated son preference is weaker among urban, literate, and wealthy women (Figure A.2), they exhibit higher rates of prenatal sex selection (Figure A.3). This pattern is consistent with their reporting lower desired fertility (Figure A.4), with educated individuals being more likely to adopt a new technology (Lleras-Muney and Lichtenberg (2005)), and with their being more efficacious in achieving their targets (Rosenzweig and Schultz (1989)). Also, wealth may matter at the margin for affordability of ultrasound scans and (safe) abortion, especially if a woman engages in multiple events. So, if there were a strict substitution of prenatal for postnatal girl mortality, we may expect the reductions in mortality and fertility that we document in this paper to be concentrated among educated and wealthy mothers. However, it is possible that these responses occur at different margins.

We examine if our results differ by mother's educational attainment (illiterate versus literate), household wealth (bottom 40 versus top 20 percent of the national wealth distribution), mother's employment status (paid employment versus rest),[47] household caste (scheduled caste (SC) versus other),[48] and rural versus urban residence. Tables 8 and 9 respectively present estimates for under-5 EFM and gender gaps in sibling size.[49] In each regression, we continue to control for all SES variables, except the one being used to examine heterogeneity. The tables show the baseline gender gaps in firstborn-girl families (first rows) and test to what extent they narrowed during the post-ultrasound period (second and third rows).

EFM by SES. In the pre-ultrasound era, the general pattern was that, among firstborn-girl families, excess girl mortality was greater in low SES groups (illiterate, poor, unemployed, rural); the estimated post-ultrasound EFM decline was also in general steeper in these low-SES households. The only exception is in the case of caste where upper caste (i.e., non-SC) families had higher EFM

[46] Regressions of actual fertility and ideal fertility on the ideal sex ratio variable confirm this. However, note that unobserved shocks that increase, say the (measured) desired number of boys, would drive up both the ideal sex ratio and ideal fertility creating a positive upward bias on the latter coefficient.

[47] The results are robust to using alternative comparisons, including employed (paid or unpaid) versus unemployed. Women's labor force participation is hockey-stick shaped in India, being more common among the poor (Das and Desai (2003)) than among the middle-classes.

[48] We pool high castes and other backward classes (OBC) because the first survey round does not distinguish them. Since OBC are better-off than SC households, the categories we use preserve the caste hierarchy in India. We also pool scheduled tribes (ST) with the higher caste group based upon finding that they take similar coefficients when included as a separate category.

[49] Table A.11 presents the heterogeneity results for effects on the number of children.

than lower castes in the pre-ultrasound era and experienced steeper EFM declines post-ultrasound. This is consistent with previous evidence that the higher castes in India have higher son preference (Srinivas (1962)). Notice that our finding that EFM generally declined more in low-SES households is consistent with larger increases in investment in girls in these households and also with a given change in investment having larger survival impacts in low-SES households where other causes of child mortality, such as infection rates, are higher.

Fertility decline by SES. Baseline gender gaps in sibling size among firstborn-girl families were larger among literate, rich, urban, and non-SC women as well as for unemployed women.[50] Like EFM, the decline in son-biased fertility stopping is greater among household-types that had larger pre-ultrasound gender gaps.

SES samples by caste. We also split each SES-subsample by caste.[51] Unlike other dimensions of SES, caste is exogenous in that an individual is born into a caste and remains in it. The caste hierarchy has been preserved by the low prevalence of inter-caste marriages.[52] The upper-castes have historically laid greater emphasis on ritual purity and adherence to religious texts, which often compromises the position of women (Das Gupta et al. (2003), Das Gupta (2010)). In accordance with this, pre-ultrasound excesses in mortality and family size in girl-led families and post-ultrasound declines were, on average, larger in higher caste households. Interacting caste with other indicators of SES shows that at the low-end of the SES distribution, low castes are more gender-equal but at the high-end, low caste behavior is similar to that of high-castes. This is consistent with the process of *Sanskritization*, wherein lower castes emulate the upper castes in seeking upward mobility (Srinivas (1962)).

6 Estimates of Substitution

To assess the extent of substitution of postnatal for prenatal discrimination after the introduction of ultrasound technology, we use our estimates to compute the number of female child deaths that have been averted and compare them with the number of girls who are missing, both due to prenatal sex-detection. These calculations are described in Appendix D. We find that for every girl that survived due to ultrasound technology, three girls were aborted before birth. These estimates take into account endogenous changes in fertility. Lower average fertility implies that the share of all births that are lower parity is increasing and since sex ratios are closer to the biological norm at lower parities (and consistent with the norm for first births), this will contribute (*through a compositional effect*) to the average sex ratio being less male-biased than otherwise.

For firstborn girl families, we calculate the proportion of discriminated births for which parents substituted postnatal discrimination with prenatal discrimination as the decline in the number of

[50]In India, on average, low-SES women are more likely to be employed, driven to work by poverty.

[51]We include households of all religions and use the self-reported caste of the household for our analysis while using religion as a control variable. These results are available upon request.

[52]According to the 2005 India Human Development Survey, only 4.4 percent of women were married to a spouse from a different caste.

girls missing due to postnatal EFM (= -90, 200) divided by the total number of missing girls due to postnatal EFM in the pre-ultrasound period (= 120, 983). This calculation implies that, for nearly 75 percent of the births preceded by a firstborn girl, parents who were practicing postnatal discrimination in the pre-ultrasound period switched to prenatal discrimination after ultrasound became available. The percentage of switchers is much larger than the estimates in Lin et al. (2014) who find that 4 percent of parents of second-parity births and 8.5 percent of parents of third- and higher-parity births made the switch in Taiwan.

7 Discussion: Relation to Recent Work

A recent study by Hu and Schlosser (2015) (henceforth HS), written in parallel with ours, addresses a similar question using similar data sources. It is therefore important that we defend the value added of this study and, importantly, explain the source of the difference in our findings. Although, like us, HS find evidence of increased parental investment in girls in the post-ultrasound period,[53] they find no change in EFM, and they refer to this as a puzzle. Whether or not girl mortality has declined in the post-ultrasound era is a question of considerable import. Unabated sex-selection biases the (cohort) population sex ratio in favor of men; and a vast literature suggests that it influences crime, marriage quality, and domestic violence.

In the rest of this section, we discuss the source of the difference in our findings. First, HS restrict their sample to children born within 3 to 10 years preceding the survey (see Table 10 in HS), so the earliest year of birth in their sample is either 1989 or 1982, which implies either no ultrasound cohorts, or three. In contrast our pre-ultrasound data span 1973-1984. We demonstrate that this matters. If we expand the HS sample to include children born within 20 years of survey (the criterion that we use) but use their specification (so ignoring the variation in EFM by firstborn sex that our specifications exploit), we recover significant declines in neonatal, infant, and under-5 mortality (see columns (4)-(6) of Table 10). The same coefficients become insignificant when only births within 10 years of birth are included (see Columns (1)-(3) of Table 10 that replicate the findings in HS).

Second, HS do not distinguish between firstborn-girl and firstborn-boy families. The summary statistics in this paper show that, in the pre-ultrasound period, EFM was concentrated in firstborn-girl families. The girl-boy differential in under-5 mortality was 2.87 p.p. for births preceded by a firstborn girl in contrast to -0.08 p.p. for births preceded by a firstborn boy. This is an enormous difference, so a specification forcing equal coefficients for the two groups may veil relevant changes. Columns (7)-(9) of Table 10 demonstrate the importance of this restriction in generating small and insignificant changes in EFM in HS. If we incorporate variation by firstborn sex in HS' specification (without changing their sample otherwise), we find a significant decline in under-5 EFM.

Third, HS model EFM and differential investments in girls as a function of fairly aggregate

[53] They use NFHS data for their analysis, which, as discussed before, is not ideal for examining the impacts on parental investments.

state-year variation in the sex ratio at birth. However, the sex ratio at birth is likely to be jointly determined with the outcomes.[54] Since we model EFM by firstborn sex (which varies at the household level), we difference out a host of state-year omitted variables. More importantly, we use quasi-experimental variation in the availability of ultrasound scanners determined by policy changes governing imports and industrial licensing in India.

8 Conclusion

As ultrasound technology became increasingly available, the global annual number of sex-selective abortions increased from nearly zero in the late 1970s to 1.6 million per year in 2005-2010 (Bongaarts and Guilmoto (2015)), with India and China being the biggest contributors. The stark growth in female feticide has garnered a lot of attention from academics, policymakers, and popular media.

Moral arguments can be made both in favor of parents' right to choose the sex of their offspring as well as against selective abortion of girls (Kumar (1983)). Abstracting from these ethical dilemmas, there are several reasons why a significantly male-biased sex ratio at birth is undesirable. The resulting scarcity of women on the marriage market can substantially increase the number of unmarried and childless men,[55] who may face destitution in old age since children through marriage are the most important source of support for the elderly in countries like India that lack institutional social security (Das Gupta et al. (2010)). Rising sex ratios can lead to increased trafficking of women,[56] higher prevalence of sexually-transmitted diseases (Ebenstein and Sharygin (2009)), and more crime (Edlund et al. (2007), Drèze and Khera (2000), Amaral and Bhalotra (2017)). Sex-selection may also result in girls being consistently born to lower-status parents, thereby relegating women to lower social strata (Edlund (1999), Bhalotra and Cochrane (2010)). While most public attention has focused upon the increasing deficit of girl children, it has also been noted that a large share of sex-selective abortions in India are conducted in unsafe environments. Complications due to unsafe abortion account for an estimated 9 percent of all maternal deaths in India (Stillman et al. (2014)).[57] On the other hand, a shortage of women on the marriage market may increase their bargaining power and welfare.[58] It has also been argued that sex-selective abortions might be

[54] Once fetal sex detection is feasible, parents simultaneously decide whether to conceive, whether to use prenatal sex diagnosis, whether to abort if the fetus is a girl, and how much to invest in male versus female births that are taken to term. The joint outcomes, thus, are fertility, the sex ratio at birth (which is more male if there is more girl abortion), post-birth investments in girls relative to boys, and girl relative to boy survival. HS impose the fairly strong identifying assumption that changes in state-year sex ratio at birth are uncorrelated with unobserved factors that could differentially affect male and female outcomes.

[55] Bhaskar (2011) estimates that one in five boys born in recent cohorts in China will be unable to find female partners.

[56] Recent evidence shows that a shortage of women in north Indian states has led to the import of brides from other poorer states in India (Kaur (2004), Ahlawat (2009)).

[57] The maternal mortality ratio in India was 178 maternal deaths per 100,000 live births in 2010-12.

[58] See Chiappori et al. (2002) and related papers for the large literature on household bargaining in developed countries. **?** shows that a relative scarcity of women in Haryana has increased their bargaining power on the marriage market and they are able to secure improved sanitation facilities at home as a result.

preferable to infanticide or postnatal discrimination (Goodkind (1996)).

Our analysis shows that the increase in sex-selective abortions fueled by ultrasound technology substantially decreased postnatal gender discrimination against girl children in India. Relative to available studies, we contribute new evidence and present a more comprehensive analysis. We find that sex-selection eliminated gender gaps in post-neonatal child mortality, postnatal health investments, and sibling size among second- and higher-parity births in households with a firstborn daughter relative to households with a firstborn son. So, although fewer girls were born, those that survived to birth were treated more equally, were more likely to survive to age five, and received higher investments during childhood, we can project that they are more likely to do well as adults in terms of cognitive attainment, income, longevity (Bhalotra and Venkataramani (2013), Currie and Rossin-Slater (2015), Bhalotra et al. (2016b)), and outcomes of their offspring (Currie and Moretti (2007), Almond and Currie (2011), Bhalotra and Rawlings (2011)). Narrowing of gender gaps in human capital also tends to be associated with higher growth rates and social change (Klasen (2002), Lagerlöf (2003)). The fertility decline (concentrated in firstborn-girl families) we observe not only benefits girls through increased resources per capita, it is also potentially beneficial for the health of mothers, which is depleted by the high levels of fertility motivated by the desire to bear sons (Milazzo (2014)). More generally, fertility decline in developing countries has been shown to be associated with economic growth, human capital accumulation, and women's empowerment (Joshi and Schultz (2007), Rosenzweig and Zhang (2009), Miller (2010), Ashraf et al. (2013)).

However, our evidence suggests that surviving girls in the post-ultrasound regime are more likely to be in low-SES households[59] and for every additional girl that survived after birth, three girls were aborted.

References

Abrevaya, J. (2009): "Are There Missing Girls in the United States? Evidence from Birth Data," *American Economic Journal: Applied Economics*, 1, 1–34.

Ahlawat, N. (2009): "Missing Brides in Rural Haryana: A Study of Adverse Sex Ratio, Poverty and Addiction," *Social Change*, 39, 46–63.

Almond, D., K. Y. Chay, and M. Greenstone (2006): "Civil Rights, the War on Poverty, and Black-White Convergence in Infant Mortality in the Rural South and Mississippi," *MIT Department of Economics Working Paper No. 07-04*.

Almond, D. and J. Currie (2011): "Killing Me Softly: The Fetal Origins Hypothesis," *Journal of Economic Perspectives*, 25, 153–72.

Almond, D. and L. Edlund (2007): "Trivers-Willard at Birth and One Year: Evidence from US Natality Data 1983-2001," *Proceedings of the Royal Society B*, 274, 2491–2496.

——— (2008): "Son Biased Sex Ratios in the US 2000 Census," *Proceedings of the National Academy of Sciences of the United States of America*, 105, 5681–5682.

[59] This has implications for marriage and violence, for instance, because of marital hypergamy (Edlund (1999)).

Almond, D., H. Li, and L. Meng (2010): "Son Preference and Early Childhood Investments in China," *Working Paper*.

Amaral, S. and S. Bhalotra (2017): "Population Sex Ratios and Crime Against Women: Long Run Effects of Sex Selection in India," Mimeo, University of Essex.

Ananat, E. O., J. Gruber, P. B. Levine, and D. Staiger (2009): "Abortion and Selection," *Review of Economics and Statistics*, 91, 124–136.

Ananat, E. O. and D. M. Hungerman (2012): "The Power of the Pill for the Next Generation: Oral Contraception's Effects on Fertility, Abortion, and Maternal and Child Characteristics," *Review of Economics and Statistics*, 94, 37–51.

Anderson, S. and D. Ray (2010): "Missing Women: Age and Disease," *Review of Economic Studies*, 77, 1262–1300.

Anukriti, S., S. Kwon, and N. Prakash (2016): "Dowry: Household Responses to Expected Marriage Payments," *mimeo*.

Arnold, F., S. Kishor, and T. Roy (2002): "Sex-selective Abortions in India," *Population and Development Review*, 28, 759–785.

Ashraf, Q. H., D. N. Weil, and J. Wilde (2013): "The Effect of Fertility Reduction on Economic Growth," *Population and development review*, 39, 97–130.

Azam, M. and G. G. Kingdon (2013): "Are girls the fairer sex in India? Revisiting intra-household allocation of education expenditure," *World Development*, 42, 143–164.

Bailey, M. (2006): "More Power to the Pill: The Impact of Contraceptive Freedom on Women's Life Cycle Labor Supply," *Quarterly Journal of Economics*, 121, 289–320.

Bailey, M. J., O. Malkova, and Z. M. McLaren (2017): "Does Parents' Access to Family Planning Increase Children's Opportunities? Evidence from the War on Poverty and the Early Years of Title X Does Parents' Access to Family Planning Increase Children's Opportunities? Evidence from the War on Poverty and the Early Years of Title X," *NBER Working Paper 23971*.

Bhalotra, S. (2010): "Fatal Fluctuations: Cyclicality in Infant Mortality in India," *Journal of Development Economics*, 93, 7–19.

Bhalotra, S., A. Chakravarty, and S. Gulesci (2016a): "The Price of Gold: Dowry and Death in India," *IZA Discussion Paper 9679*.

Bhalotra, S. and T. Cochrane (2010): "Where Have All the Young Girls Gone? Identification of Sex Selection in India," *IZA Discussion Paper No. 5381*.

Bhalotra, S., M. Karlsson, and T. Nilsson (2016b): "Infant Health and Longevity: Evidence from a Historical Intervention in Sweden," *Journal of European Economic Association*, Forthcoming.

Bhalotra, S. and S. Rawlings (2011): "Intergenerational Persistence in Health in Developing Countries: the Penalty of Gender Inequality?" *Journal of Public Economics*, 95, 286–299.

Bhalotra, S. and A. van Soest (2008): "Birth-Spacing, Fertility and Neonatal Mortality in India: Dynamics, Frailty, and Fecundity," *Journal of Econometrics*, 143, 274–290.

Bhalotra, S. and A. Venkataramani (2013): "Cognitive Development and Infectious Disease: Gender Differences in Investments and Outcomes," *IZA Discussion Paper 7833*.

Bharadwaj, P. and L. K. Lakdawala (2013): "Discrimination Begins in the Womb: Evidence of Sex-Selective Prenatal Investments," *Journal of Human Resources*, 48, 71–113.

Bhaskar, V. (2011): “Sex Selection and Gender Balance,” *American Economic Journal: Microeconomics*, 3, 214–244.

Bhat, P. (2002): “Maternal Mortality in India: An Update,” *Studies in Family Planning*, 33, 227–236.

Bongaarts, J. and C. Guilmoto (2015): “How Many More Missing Women? Excess Female Mortality and Prenatal Sex Selection, 1970–2050,” *Population and Development Review*, 41, 241–269.

Bozzoli, C., A. Deaton, and C. Quintana-Domeque (2009): “Adult Height and Childhood Disease,” *Demography*, 46, 647–669.

Carranza, E. (2015): “Soil Endowments, Female Labor Force Participation, and the Demographic Deficit of Women in India,” *American Economic Journal: Applied Economics*, 6, 1997–225.

Charles, K. K. and M. Stephens Jr. (2006): “Abortion Legalization and Adolescent Substance Abuse,” *Journal of Law and Economics*, 49, 481–505.

Chay, K. Y., J. Guryan, and B. Mazumder (2009): “Birth Cohort and the Black-White Achievement Gap: The Roles of Access and Health Soon After Birth,” *NBER Working Paper 15078*.

Chiappori, P.-A., B. Fortin, and G. Lacroix (2002): “Household Labor Supply, Sharing Rule and the Marriage Market,” *Journal of Political Economy*, 110, 37–72.

Clark, S. (2000): “Son Preference and Sex Composition of Children: Evidence from India,” *Demography*, 37, 95–108.

Currie, J. and E. Moretti (2007): “Biology as Destiny? Short and Long-Run Determinants of Intergenerational Transmission of Birth Weight,” *Journal of Labor Economics*, 25, 231–264.

Currie, J. and M. Rossin-Slater (2015): “Early-life Origins of Lifecycle Well-being: Research and Policy Implications,” *Journal of Policy Analysis and Management*, 34, 208–242.

Das, M. B. and S. Desai (2003): “Why are Educated Women Less Likely to be Employed in India? Testing Competing Hypotheses,” *Social Protection Discussion Paper Series 0313, World Bank*.

Das Gupta, M. (1987): “Selective Discrimination Against Female Children in Rural Punjab, India,” *Population and Development Review*, 13, 77–100.

——— (2010): “Family Systems, Political Systems, and Asia’s ‘Missing Girls’: The Construction of Son Preference and Its Unraveling,” *Asian Population Studies*, 6, 123–152.

Das Gupta, M. and P. Bhat (1997): “Fertility Decline and Increased Manifestation of Sex Bias in India,” *Population Studies*, 51, 307–315.

Das Gupta, M., A. Ebenstein, and E. J. Sharygin (2010): “China’s Marriage Market and Upcoming Challenges for Elderly Men,” *World Bank Policy Research Working Paper 5351.*

Das Gupta, M., J. Zhenghua, L. Bohua, X. Zhenming, W. Chung, and B. Hwa-Ok (2003): “Why is Son Preference So Persistent in East and South Asia? A Cross-country Study of China, India and the Republic of Korea,” *Journal of Development Studies*, 40, 153–187.

Donohue, J. J. I., J. Grogger, and S. D. Levitt (2009): “The Impact of Legalized Abortion on Teen Childbearing,” *American Law and Economics Review*, 11, 24–46.

Donohue, J. J. I. and S. D. Levitt (2001): “The Impact of Legalized Abortion on Crime,” *Quarterly Journal of Economics*, 141, 379–420.

Drèze, J. and R. Khera (2000): “Crime, Gender, and Society in India: Insights from Homicide Data,” *Population and Development Review*, 26, 335–352.

Ebenstein, A. Y. and E. J. Sharygin (2009): "The Consequences of the Missing Girls of China," *The World Bank Economic Review*, 23, 399–425.

Edlund, L. (1999): "Son Preference, Sex Ratios and Marriage Patterns," *Journal of Political Economy*, 107, 1275–1304.

Edlund, L., H. Li, J. Yi, and J. Zhang (2007): "Sex Ratios and Crime: Evidence from China's One Child Policy," *IZA Discussion Paper 3214*.

Epner, J. E. G., H. S. Jonas, and D. L. Seckinger (1998): "Late-term Abortion," *Journal of American Medical Association*, 280, 724–729.

Gangoli, G. (1998): "Reproduction, Abortion and Women's Health," *Social Scientist*, 26, 83–105.

George, S. (2006): "Sex Ratio in India, Response to Jha, Prabhat, Rajesh Kumar, Neeraj Dhingra," *The Lancet*, 367, 1725.

Goldin, C. and L. Katz (2002): "The Power of the Pill: Oral Contraceptives and Women's Career and Marriage Decisions," *Journal of Political Economy*, 110, 730–770.

Goodkind, D. (1996): "On Substituting Sex Preference Strategies in East Asia: Does Prenatal Sex Selection Reduce Postnatal Discrimination?" *Population and Development Review*, 22, 111–125.

Grossman, M. and J. Joyce, Theodore (1990): "Unobservables, Pregnancy Resolutions, and Birth Weight Production Functions in New York City," *Journal of Political Economy*, 98, 983–1007.

Grover, A. and R. Vijayvergia (2006): "Sex Ratio in India, Response to Jha, Prabhat, Rajesh Kumar, Neeraj Dhingra," *The Lancet*, 367, 1726.

Gruber, J., P. B. Levine, and D. Staiger (1999): "Abortion Legalization and Child Living Circumstances: Who is the Marginal Child?" *Quarterly Journal of Economics*, 114, 263–292.

Hu, L. and A. Schlosser (2015): "Prenatal Sex Determination and Girls' Well-Being: Evidence from India," *Economic Journal*, 125, 1227–1261.

Jayachandran, S. (2016): "Fertility Decline and Missing Women," *American Economic Journal: Applied Economics (forthcoming)*.

Jayachandran, S. and I. Kuziemko (2011): "Why Do Mothers Breastfeed Girls Less Than Boys: Evidence and Implications for Child Health in India," *Quarterly Journal of Economics*, 126, 1485–1538.

Jayachandran, S., A. Lleras-Muney, and K. Smith (2010): "Modern Medicine and the 20th-Century Decline in Mortality: Evidence on the Impact of Sulfa Drugs," *American Economic Journal: Applied Economics*, 2, 118–146.

Jensen, R. (2012): "Another Mouth to Feed? The Effects of Fertility on Girls' Malnutrition," *CESifo Economic Studies*, 58, 322–47.

Joshi, S. and T. P. Schultz (2007): "Family Planning as an Investment in Development: Evaluation of a Program's Consequences in Matlab, Bangladesh," *Yale University Economic Growth Center Discussion Paper*.

Kalsi, P. (2015): "Abortion Legalization, Sex Selection, and Female University Enrollment in Taiwan," *Economic Development and Cultural Change*, 64, 163–185.

Kaur, R. (2004): "Across-Region Marriages: Poverty, Female Migration and the Sex Ratio," *Economic and Political Weekly*, 39, 2595–2603.

Klasen, S. (2002): "Low Schooling for Girls, Slower Growth for All? Cross-Country Evidence on the Effect of Gender Inequality in Education on Economic Development," *World Bank Economic Review*, 16, 345–373.

Kumar, D. (1983): "Male Utopias or Nightmares?" *Economic and Political Weekly*, 18, 61–64.

Lagerlöf, Nils-Petterf, N.-P. (2003): "Gender Equality and Long-Run Growth," *Journal of Economic Growth*, 8, 403–426.

Lin, M.-J., J.-T. Liu, and N. Qian (2014): "More Missing Women, Fewer Dying Girls: The Impact of Abortion on Sex Ratios at Birth and Excess Female Mortality in Taiwan," *Journal of European Economic Association*, 12.

Lleras-Muney, A. and F. Lichtenberg (2005): "The Effect Of Education On Medical Technology Adoption: Are The More Educated More Likely To Use New Drugs?" *Annales d'Economie et Statistique*.

Low, B. S. (2000): *Why Sex Matters: A Darwinian Look at Human Behavior*, Princeton, NJ: Princeton University Press.

Mahal, A., A. Varshney, and S. Taman (2006): "Diffusion of Diagnostic Medical Devices and Policy Implications for India," *International Journal of Technology Assessment in Health Care*, 22, 184–190.

Milazzo, A. (2014): "Why Are Adult Women Missing? Son Preference and Maternal Survival in India," *World Bank Policy Research Working Paper 6802*.

Miller, G. (2010): "Contraception as Development? New Evidence from Family Planning in Colombia," *The Economic Journal*, 120, 709–736.

Nandi, A. and A. B. Deolalikar (2013): "Does a legal ban on sex-selective abortions improve child sex ratios? Evidence from a policy change in India," *Journal of Development Economics*, 103, 216–228.

Oster, E. (2009): "Proximate Sources of Population Sex Imbalances in India," *Demography*, 46.

Phadke, S. (1997): "Pro Choice or Population Control: A Study of the Medical Termination of Pregnancy Act, Government of India 1971," *Re/Productions*.

Pitt, M. and M. R. Rosenzweig (1990): "Estimating the Intrahousehold Incidence of Illness: Child Health and Gender-inequality in the Allocation of Time," *International Economic Review*, 31, 969–989.

Pop-Eleches, C. (2006): "The Impact of a Change in Abortion Regime on Socio-Economic Outcomes of Children: Evidence from Romania," *Journal of Political Economy*, 114, 744–773.

Retherford, R. D. and V. Mishra (2001): "An Evaluation of Recent Estimates of fertility Trends in India," *National Family Heath Survey Subject Reports*.

Retherford, R. D. and T. Roy (2003): "Factors Affecting Sex-Selective Abortion In India," *National Family Heath Survey Bulletin*, 17.

Rose, E. (2000): "Gender Bias, Credit Constraints and Time Allocation in Rural India," *Economic Journal*, 110, 738–758.

Rosenblum, D. (2013): "The effect of fertility decisions on excess female mortality in India," *Journal of Population Economics*, 26, 147–180.

Rosenzweig, M. and T. P. Schultz (1989): "Schooling, Information and Nonmarket Productivity: Contraceptive Use and Its Effectiveness," *International Economic Review*, 30, 457–477.

Rosenzweig, M. R. and J. Zhang (2009): "Do Population Control Policies Induce More Human Capital Investment? Twins, Birth Weight and China's "One-Child" policy," *The Review of Economic Studies*, 76, 1149–1174.

Sen, A. (1990): "More Than 100 Million Women Are Missing," *New York Review of Books*, 61–66.

Shapira, G. (2013): "How subjective beliefs about HIV infection affect life-cycle fertility : Evidence from rural Malawi," *Journal of Human Resources (forthcoming)*.

Srinivas, M. (1962): *Caste in Modern India: And Other Essays*, Asia Publishing House, Bombay.

Stillman, M., J. J. Frost, S. Singh, A. M. Moore, and S. Kalyanwala (2014): "Abortion in India: A Literature Review," New York: Guttmacher Institute.

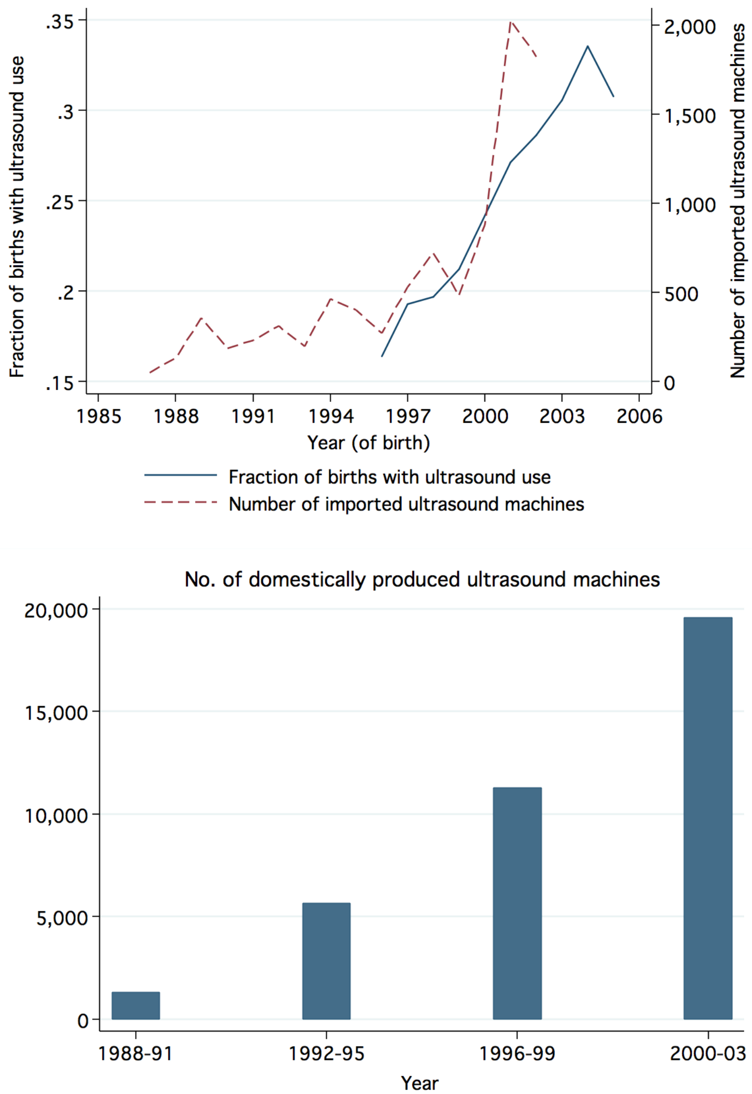
Trivers, R. L. and D. E. Willard (1973): "Natural selection of parental ability to vary the sex ratio of offspring," *Science*, 179, 90–92.

Visaria, L. (2005): "Female Deficit in India: Role of Prevention of Sex Selective Abortion Act," *mimeo*.

Yang, W.-S., H. H. Knöbel, and C.-J. C. Chen (1996): "Gender differences in postneonatal infant mortality in Taiwan," *Social Science and Medicine*, 43, 1461–1465.

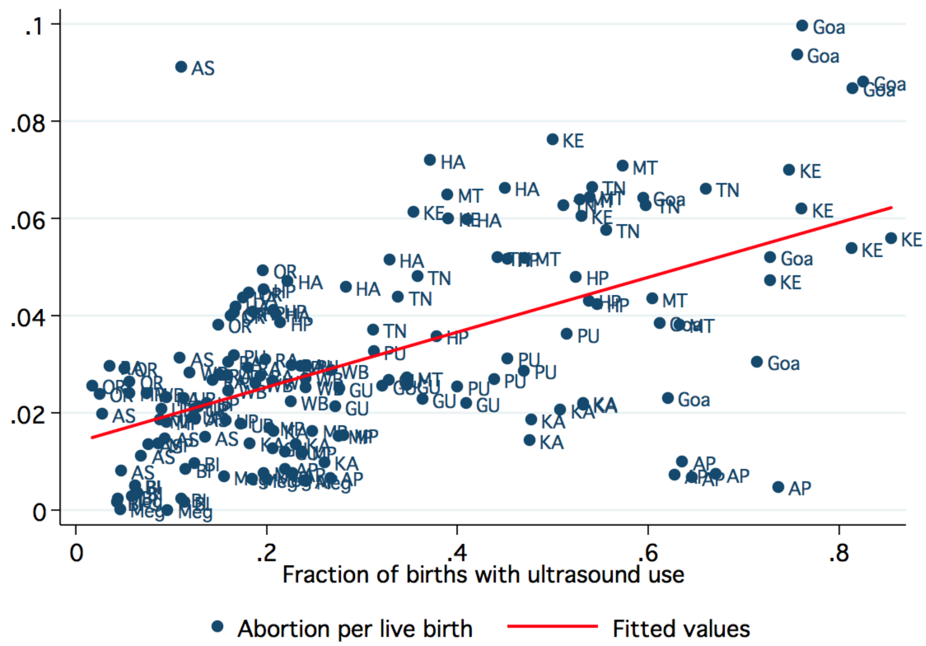
Figures

Figure 1: Ultrasound use by mothers and supply of ultrasound scanners in India



NOTES: (1) The solid line in the top graph plots the fraction of births in a year for which the mother reports getting an ultrasound test at some point during the pregnancy (the denominator equals the number of births with a non-missing response on ultrasound use). The relevant question was not asked in NFHS-1 but in NFHS-2 and NFHS-3, data on ultrasound use was collected for births since January 1995 and January 2001, respectively. The years 1995 and 2000 have been dropped due to extremely small sample sizes. (2) The dashed line in the top graph plots the number of ultrasound scanners imported at the national level, the first records of which appear in the import data in 1987; indeed, there was no category coding these scanners before then (Source: Mahal et al. (2006)). (3) The bars in the bottom graph plot the number of ultrasound machines produced domestically in India. Data source: George (2006).

Figure 2: Relationship between number of abortions and ultrasound use by mothers

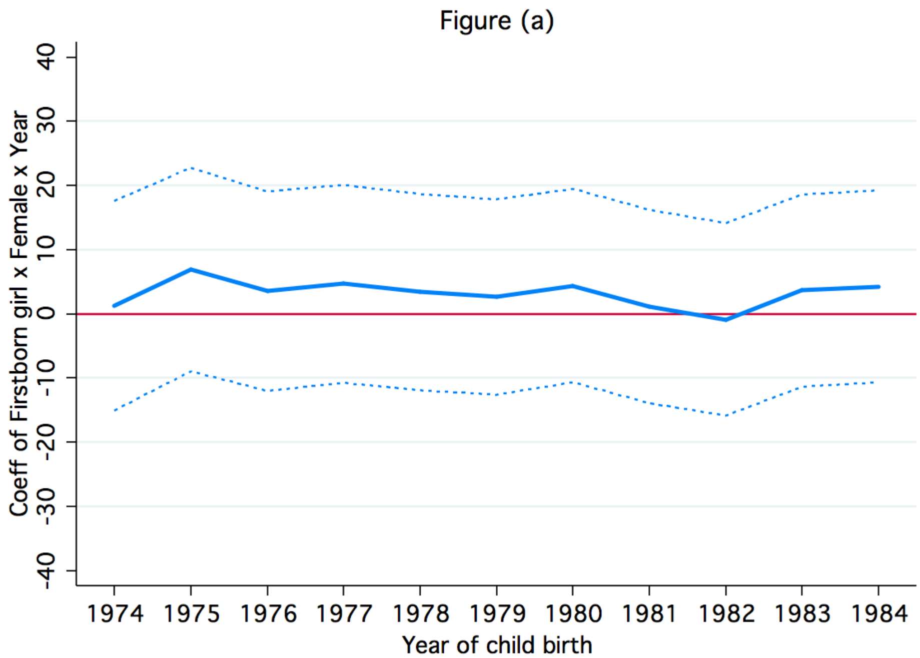


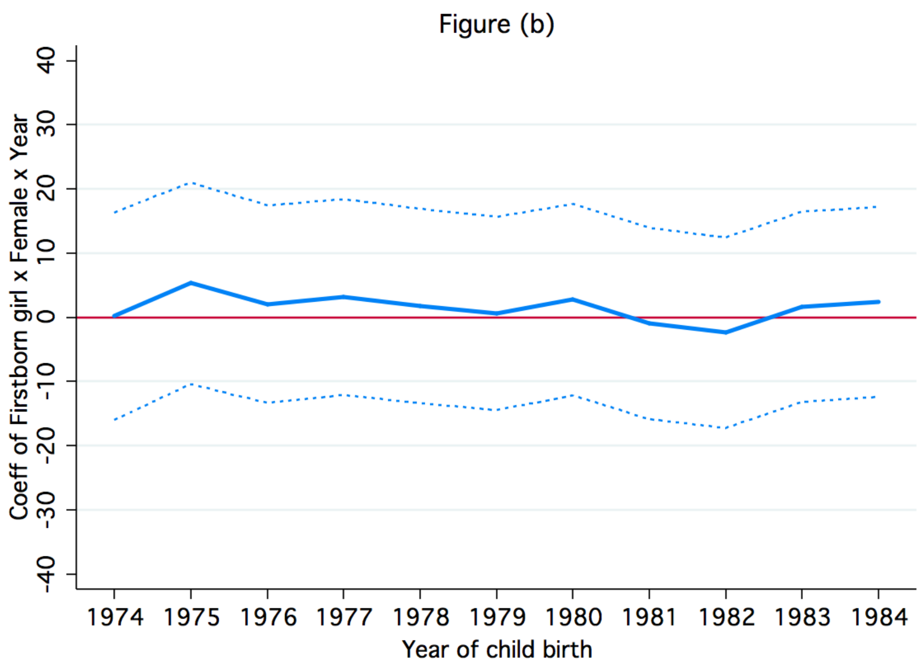
NOTES: This graph plots the state-year variation in the officially reported number of abortions and the fraction of births in a state-year for which the mother reports getting an ultrasound test at some point during the pregnancy (other details in Figure 1 notes). Data for the officially reported number of abortions comes from `http://www.johnstonsarchive.net/policy/abortion/india/ab-indias.html`.

Figure 3: Trends in proportion of females at birth by birth order and sex composition of older siblings

NOTES: These plots are derived from Bhalotra and Cochrane (2010). Panel A shows the evolution of percent female among first births over time. In panels B, C, and D the trend in percentage of births that are female is plotted respectively for second, third, and fourth births separately for families that have at least one son and families with no sons at the time of the respective birth. In all cases, the y-axis shows the 5-year moving average of percentage of births that are female. The figures show that, despite ultrasound availability, the sex ratio of first births has remained normal. It also shows that the sex ratio at birth in families without sons diverges from the sex ratio in families with sons *after* the introduction of ultrasound.

Figure 4: Test of differential pre-trends in excess female under-5 mortality by firstborn sex





NOTES: This figure tests for parallel trends in excess female under-5 mortality by firstborn sex during the pre-ultrasound period (1973-1984). Excess female under-5 mortality equals the percentage of female births that die minus the percentage of male births that die before age five. The graphs plot the coefficients of the *First girl* x *Female* x *Year* indicator variables (and the 95% confidence intervals) from the regression of under-5 mortality on the full set of interactions between indicators for *First girl*, *Female*, and *Year*, without other controls in Figure (a) and with fixed effects for birth order and state in Figure (b). The omitted year is 1973.

Tables

Table 1: Unadjusted sample means by child gender, firstborn sex and pre-post ultrasound (%)

	Firstborn boy families			Firsborn girl families		
	(1)	(2)	(3)	(4)	(5)	(6)
	Male	Female	(2)-(1)	Male	Female	(5)-(4)
A. Under-5 mortality						
Pre-ultrasound: 1973-1984	12.90	12.82	-0.08	10.49	13.36	2.87
N	17,769	16,520		17,252	15,908	
Early diffusion period: 1985-1994	10.04	10.26	0.22	9.04	10.63	1.59
N	40,940	38,766		43,267	39,252	
Late diffusion period: 1995-2005	7.71	7.80	0.09	7.14	8.23	1.09
N	24,147	22,907		27,447	23,897	
B. Immunized						
Pre-ultrasound: 1973-1984	0.707	0.708	0.001	0.720	0.679	-0.041
N	1,861	1,624		1,044	916	
Early diffusion period: 1985-1994	0.859	0.858	-0.001	0.898	0.886	-0.012
N	1,882	1,723		1,438	1,298	
Late diffusion period: 1995-1999	0.924	0.910	-0.014	0.938	0.942	0.004
N	792	754		696	568	
C. Probability of birth in a year						
Pre-ultrasound: 1973-1984		0.29			0.28	
N		408,857			233,131	
Early diffusion period: 1985-1994		0.26			0.23	
N		682,371			451,674	
Late diffusion period: 1995-2005		0.23			0.18	
N		405,628			305,998	

NOTES: This table shows pre- versus post-ultrasound trends in unadjusted (raw) outcomes among second- and higher-order children, by firstborn sex and child gender. EFM denotes excess female mortality. Panel A shows a striking decline in female relative to male under-5 mortality in firstborn-girl families alongside no decline in firstborn-boy families. Panel B shows exactly that pattern for immunization. Panel C shows fertility declined between the pre- and post-ultrasound period in all families but more rapidly in families with firstborn-girls.

Table 2: Test of balance in samples by firstborn sex

	1973-1984		1985-1994		1995-2005		All years
	FB	FG	FB	FG	FB	FG	FB-FG
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Rural	0.68	0.68	0.64	0.65	0.59	0.58	-0.0004
Hindu	0.78	0.78	0.76	0.76	0.73	0.74	-0.0009
Muslim	0.11	0.11	0.12	0.12	0.13	0.13	-0.0001
SC	0.14	0.14	0.15	0.16	0.16	0.17	-0.0030*
ST	0.12	0.13	0.13	0.13	0.14	0.13	-0.0002
Mother's Education							
No education	0.59	0.58	0.48	0.49	0.32	0.32	0.002
Incomplete secondary	0.35	0.36	0.40	0.40	0.48	0.48	-0.002
Secondary or higher	0.06	0.07	0.11	0.11	0.20	0.20	0.00009
Father's Education							
No education	0.32	0.31	0.26	0.27	0.18	0.18	-0.002
Incomplete secondary	0.50	0.50	0.52	0.52	0.58	0.58	0.0005
Secondary or higher	0.18	0.19	0.22	0.21	0.24	0.24	0.001
Mother's birth cohort							
1942-1960	0.52	0.52	0.04	0.04	0.00	0.00	0.002
1961-1970	0.48	0.48	0.56	0.56	0.07	0.06	0.001
1971-1987	0.00	0.00	0.41	0.40	0.93	0.93	-0.003
Mother's age at birth							
12-15	0.14	0.13	0.11	0.11	0.05	0.06	0.001
16-18	0.35	0.35	0.32	0.32	0.26	0.26	0.002
19-24	0.44	0.44	0.47	0.48	0.53	0.53	-0.002
25-30	0.07	0.07	0.09	0.09	0.14	0.13	-0.001
31-49	0.01	0.01	0.01	0.01	0.02	0.02	0.0002
Household wealth							
2nd quintile	0.15	0.14	0.16	0.17	0.16	0.16	-0.00003
3rd quintile	0.17	0.17	0.17	0.17	0.18	0.18	-0.001
4th quintile	0.21	0.21	0.23	0.22	0.23	0.23	0.001
Richest quintile	0.23	0.23	0.24	0.23	0.27	0.27	0.003
N (1st births)	26,064	23,905	41,430	38,712	22,108	21,074	173,293

NOTES: This table compares the socioeconomic characteristics of firstborn-boy (FB) and firstborn-girl (FG) families during the pre-ultrasound period and during the two post-ultrasound periods in the NFHS sample. The sample is restricted to first births as only these are quasi-random. SC and ST denote Scheduled Castes and Scheduled Tribes, respectively. Column (7) shows the differences in sample means among FB and FG families for the entire sample. The larger sample size of FB relative to FG families reflects the fact that the sex ratio at birth is naturally above one. *** 1%, ** 5%, * 10%

Table 3: Excess female under-5 mortality as a function of ultrasound availability and firstborn sex

Dep var: Death before age 5	(1)	(2)	(3)	(4)
Panel A:				
*Firstborn girl * Female*	2.957***	2.821***	2.768***	2.563***
	(0.607)	(0.598)	(0.588)	(0.588)
*Firstborn girl * Female * Post1*	-1.646**	-1.527*	-1.525*	-1.602
	(0.773)	(0.776)	(0.768)	(0.974)
*Firstborn girl * Female * Post2*	-1.998**	-1.953**	-1.996**	-1.841**
	(0.744)	(0.749)	(0.744)	(0.734)
Panel B:				
*Firstborn girl * Female*	2.957***	2.821***	2.775***	2.564***
	(0.607)	(0.598)	(0.586)	(0.587)
*Firstborn girl * Female * Post*	-1.741**	-1.642**	-1.647**	-1.659*
	(0.711)	(0.710)	(0.695)	(0.828)
N	227,129			
Baseline mean	12.38			
*Female * Post1* and *Female * Post2*	x	x		
*Firstborn Girl * Post1* and *Firstborn Girl * Post2*	x	x		
X_{ijt}		x	x	x
Firstborn Girl x Birth year FE			x	x
Female x Birth year FE			x	x
Female x State FE			x	x
Female x Birth order FE			x	x
Birth order x Birth year FE			x	x
Birth order x State FE			x	x
State x Birth year FE			x	x
Firstborn girl x State FE			x	x
Firstborn girl x Birth order FE			x	x
Mother FE				x

NOTES: Sample of second- and higher-order births. In Panel B, instead of splitting the post-ultrasound period into two sub-periods, we only use one indicator, *Post*, to divide the sample into pre- and post-ultrasound years. Each column is a separate regression. The outcome is an indicator for death before age 5. We drop children that are less than 5 years old to allow each child in the sample full exposure to the risk of under-5 mortality. We always control for *Female* and fixed effects (FE) for birth year and birth order and for *Firstborn girl* and state FE in columns (1)-(3). The vector X_{ijt} comprises mother's age at birth and, except in column (4), household wealth quintiles, caste, religion, residence in a rural area, educational attainment of child's parents, and mother's birth cohort. Standard errors in parentheses are clustered by state. Baseline mean refers to the average likelihood of under-5 mortality in the pre-ultrasound period. *** 1%, ** 5%, * 10%.

Table 4: Postnatal health investments as a function of ultrasound availability and firstborn sex

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	Months breastfed			Received $\geq$ one vaccine			Medicines/special food during illness last year (Rs.)		
First girl * *Female*	-5.128**	-4.462**	-3.932**	-0.130**	-0.093**	-0.069	-21.695	-15.745	-73.239*
	(1.875)	(1.817)	(1.725)	(0.051)	(0.039)	(0.048)	(35.677)	(37.157)	(38.891)
First girl * *Female* * *Post1*	4.688**	3.980*	2.368	0.113**	0.099*	0.076*	19.842	13.687	63.299
	(2.013)	(2.029)	(2.121)	(0.049)	(0.048)	(0.042)	(45.916)	(45.364)	(53.199)
First girl * *Female* * *Post2*	4.174	3.599	1.432	0.094*	0.073*	0.118**	36.163	32.147	92.722*
	(3.137)	(2.538)	(2.511)	(0.051)	(0.035)	(0.044)	(37.880)	(40.721)	(44.802)
N		13,085			20,562			15,157	
Baseline mean		0.167			-0.035			-108.717	
*Female***Post1* & *Female***Post2*	x	x		x	x		x	x	
*1st Girl***Post1* & *1st Girl***Post2*	x	x		x	x		x	x	
X_{ijt}		x	x		x	x		x	x
1st Girl x Birth year FE			x			x			x
Female x Birth year FE			x			x			x
Female x State FE			x			x			x
Female x Birth order FE			x			x			x
Birth order x State FE			x			x			x
State-specific time trends			x			x			x
Firstborn girl x State FE			x			x			x
Firstborn girl x Birth order FE			x			x			x

NOTES: This table reports investment effects for children of second- and higher-order birth order using the REDS sample. Each column is from a separate regression. We control for *Female*, *Firstborn girl*, and fixed effects for birth year, birth order, and state in all columns. The vector X_{ijt} comprises mother's age at birth, caste, religion, educational attainment of child's parents, and mother's birth cohort. Breastfeeding results are based on the last two surviving births of a mother. Vaccination and health expenditure results are based on all surviving children of a mother. Standard errors in parentheses are clustered by state. Baseline mean refers to the average of the relevant outcome variable in the pre-ultrasound period. *** 1%, ** 5%, * 10%. Further investment results using the NFHS data are in Table A.10.

Table 5: Fertility: Hazard of birth as a function of ultrasound availability and firstborn sex

	(1)	(2)	(3)	(4)
Firstborn girl	0.173***	0.185***	0.155***	0.168***
	(0.011)	(0.012)	(0.038)	(0.038)
*Firstborn girl * Post1*	-0.073***	-0.084***	-0.092***	-0.102***
	(0.006)	(0.007)	(0.009)	(0.010)
*Firstborn girl * Post2*	-0.190***	-0.211***	-0.220***	-0.233***
	(0.016)	(0.018)	(0.017)	(0.017)
Ideal sex ratio				0.076***
				(0.022)
Ideal no. of children				0.161***
				(0.010)
N	2,455,630	2,455,553	2,455,553	2,276,192
Baseline mean	0.285			
X_i	x	x	x	x
Year FE	x	x	x	x
State FE	x	x	x	x
Age FE	x	x	x	x
Parity FE	x	x	x	x
State x Year FE	x	x	x	x
Years since last birth FE		x	x	x
Firstborn girl x State FE			x	x

NOTES: Coefficients (and not odds ratios) from specification (2) estimated using a logistic regression on a mother-year sample. The dependent variable is an indicator for birth to a given mother in a given year. Sample includes all mothers who have ever given birth, for each year from their year of marriage to the year of interview. Standard errors in parentheses are clustered by state. Baseline mean is the mean probability of birth in a given year during the pre-ultrasound period. *** 1%, ** 5%, * 10%.

Table 6: Fertility: Number of children as a function of ultrasound availability and firstborn sex

	Number of births		Excess Fertility
	(1)	(2)	(3)
Firstborn girl	0.155***	0.141***	0.117***
	(0.012)	(0.015)	(0.018)
*Firstborn girl * Post1*	-0.088***	-0.079***	-0.085***
	(0.016)	(0.019)	(0.024)
*Firstborn girl * Post2*	-0.112***	-0.100***	-0.093***
	(0.018)	(0.023)	(0.025)
Ideal no. of children		0.315***	
		(0.018)	
Ideal sex ratio		0.052***	-0.345***
		(0.018)	(0.024)
N	118,663	88,475	88,475
Baseline mean	3.001	3.001	0.451

NOTES: Estimates of specification (3) estimated using OLS. The dependent variable in columns (1)-(2) is the number of births at the time of interview and in column (3) is excess fertility which equals number of births minus ideal number of children. Baseline means are average of the outcome variable in each column for mothers who had both their first and last birth within 1973-1984. Standard errors in parentheses are clustered by state. *** 1%, ** 5%, * 10%.

Table 7: Fertility: Investigating the margin of response by number of children

	Number of children			
	$= 1$	$= 2$	$= 3$	≥ 4
	(1)	(2)	(3)	(4)
Firstborn girl	0.012**	-0.067***	-0.015*	0.069***
	(0.005)	(0.006)	(0.007)	(0.006)
*Firstborn girl * Post1*	-0.010	0.021**	0.030***	-0.042***
	(0.007)	(0.008)	(0.009)	(0.007)
*Firstborn girl * Post2*	-0.017**	0.040***	0.035***	-0.058***
	(0.007)	(0.009)	(0.010)	(0.009)
N	118663	118663	118663	118663
Baseline mean	0.114	0.271	0.300	0.315

NOTES: This table presents estimates from specification (3) using indicators for the mother having, respectively, 1, 2, 3, and ≥ 4 children at the time of survey in columns (1)-(4). Standard errors in parentheses are clustered by state. Baseline means are average of the outcome variable in each column for mothers who had all their births within 1973-1984. *** 1%, ** 5%, * 10%.

Table 8: Under-5 mortality: Heterogeneity by socioeconomic status

	Mother's Education		**Wealth**		**Mother's Employment**	
	Illiterate	Literate	Bottom 40%	Top 20%	Paid employment = 0	Paid employment = 1
	(1)	(2)	(3)	(4)	(5)	(6)
*Firstborn girl * Female*	3.455***	1.679**	3.279***	1.022	2.993***	2.392***
	(0.830)	(0.774)	(1.143)	(0.752)	(0.753)	(0.725)
*Firstborn girl * Female * Post1*	-1.986*	-0.789	-1.622	0.386	-1.824*	-1.039
	(1.077)	(0.992)	(1.542)	(1.090)	(0.971)	(1.026)
*Firstborn girl * Female * Post2*	-2.320*	-1.232	-2.177	0.500	-2.463**	-1.360
	(1.244)	(1.031)	(1.723)	(1.506)	(1.113)	(0.870)
N	137,681	89,448	96,762	41,745	131,325	95,657
Baseline mean	15.41	7.12	16.91	6.04	11.69	13.41

	Caste		**Rurality**	
	SC	General/OBC/ST	Rural	Urban
	(1)	(2)	(3)	(4)
*Firstborn girl * Female*	1.707	2.816***	3.718***	0.315
	(1.547)	(0.555)	(0.776)	(0.946)
*Firstborn girl * Female * Post1*	-0.863	-1.525*	-2.522**	1.086
	(2.033)	(0.774)	(1.055)	(1.035)
*Firstborn girl * Female * Post2*	-0.683	-2.100**	-3.179***	0.833
	(2.483)	(0.787)	(1.072)	(1.313)
N	37,912	189,217	156,311	70,818
Baseline mean	15.90	11.80	14.15	8.15

NOTES: This tables reports estimates from the specification in column 3 of Table 3 for various sub-samples. Each column within a panel is a separate regression. SC, ST, OBC, and General respectively denote scheduled castes, scheduled tribes, other backward classes, and upper castes. The wealth categories are based on the national household wealth distribution. Standard errors in parentheses are clustered by state. Baseline mean refers to under-5 mortality for children born in the pre-ultrasound period, by mothers with the specific socioeconomic status. *** 1%, ** 5%, * 10%.

Table 9: Birth hazard: Heterogeneity by socioeconomic status

	Mother's Education		**Wealth**		**Mother's Employment**	
	Illiterate	Literate	Bottom 40%	Top 20%	Paid employment = 0	Paid employment = 1
	(1)	(2)	(3)	(4)	(5)	(6)
Firstborn girl	0.060**	0.324***	0.091**	0.388***	0.177***	0.140***
	(0.029)	(0.037)	(0.043)	(0.036)	(0.046)	(0.033)
*Firstborn girl * Post1*	-0.059***	-0.133***	-0.093***	-0.114***	-0.114***	-0.062***
	(0.009)	(0.015)	(0.006)	(0.021)	(0.012)	(0.013)
*Firstborn girl * Post2*	-0.127***	-0.312***	-0.172***	-0.317***	-0.253***	-0.174***
	(0.016)	(0.017)	(0.021)	(0.020)	(0.022)	(0.020)
N	1,351,731	1,102,413	971,697	599,531	1,494,545	958,583
Baseline mean	0.289	0.279	0.288	0.263	0.287	0.283

	Caste		**Rurality**	
	SC	General/OBC/ST	Rural	Urban
	(1)	(2)	(3)	(4)
Firstborn girl	0.078	0.172***	0.092**	0.310***
	(0.048)	(0.037)	(0.040)	(0.038)
*Firstborn girl * Post1*	-0.040**	-0.101***	-0.085***	-0.112***
	(0.020)	(0.009)	(0.010)	(0.017)
*Firstborn girl * Post2*	-0.160***	-0.234***	-0.184***	-0.296***
	(0.038)	(0.016)	(0.020)	(0.023)
N	374,215	2,080,707	1,598,030	856,144
Baseline mean	0.295	0.284	0.289	0.278

NOTES: Estimates of equation (2) in the text. The dependent variable is an indicator for birth to a given mother in a given year. We use a mother-year data set for all mothers who have ever given birth, for each year from their year of marriage to the year of interview. Standard errors in parentheses are clustered by state. Baseline mean is the mean probability of birth in a given year during the pre-ultrasound period, by mothers with the specific socioeconomic status. *** 1%, ** 5%, * 10%.

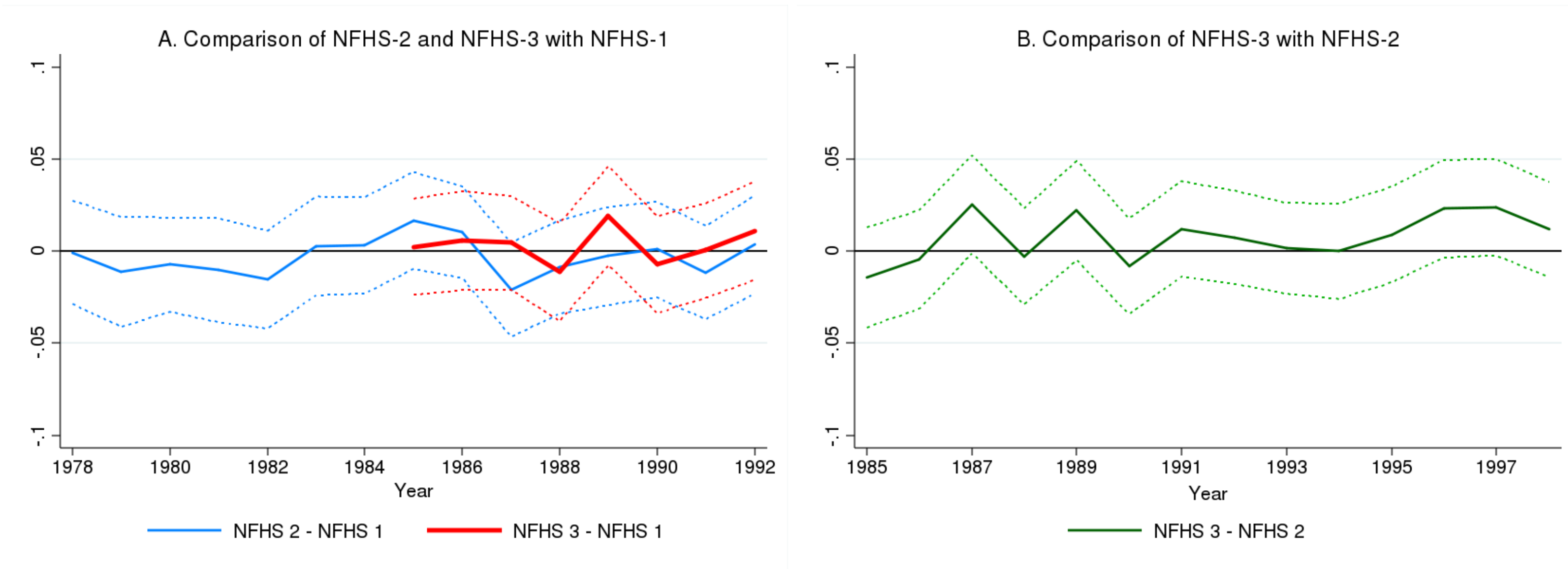
Table 10: Comparison with Hu and Schlosser (2015)

	Births within 10 years of survey [HS sample]			Births within 20 years of survey [Our sample]			Births within 10 years of survey [HS sample]		
	Neonatal	Infant	Under 5	Neonatal	Infant	Under 5	Neonatal	Infant	Under 5
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
MFR_{st}	2.827	3.474	-6.660	3.276	5.941**	7.251**	1.818	2.655	-13.949
	(2.898)	(3.810)	(7.344)	(2.132)	(2.658)	(3.236)	(3.544)	(4.554)	(9.127)
$Female * MFR_{st}$	0.047	-1.474	0.552	-3.555*	-7.323***	-11.985***	3.418	2.673	12.205**
	(1.874)	(2.781)	(4.728)	(1.974)	(2.438)	(3.630)	(3.227)	(4.194)	(5.288)
$Female * MFR_{st} * Firstborn\ girl$							-10.127	-12.527	-34.025**
							(6.493)	(9.804)	(15.927)
N	297,519	272,115	156,208	500,798	475,394	359,487	297,519	272,115	156,208

NOTES: Columns (1)-(3) replicate the findings in Table 10 of Hu and Schlosser (2015) (HS). Columns (4)-(6) show how HS' findings change when their sample is expanded to include births within 20 years of survey, and columns (7)-(9) show what happens when firstborn sex is used as a third interaction while maintaining their sample restriction of births in the last 10 years. MFR_{st} is the male-female ratio at birth in state s and year t. In columns (1)-(6), we use Hu and Schlosser (2015)'s specification, except that we exclude twins from our sample and do not control for mother's mass media exposure. In the last three columns, we include the main effects, double interactions, and triple interactions of $Female$, MFR_{st}, and $Firstborn\ girl$ in addition to the same SES controls as columns (1)-(6). *** 1%, ** 5%, * 10%.

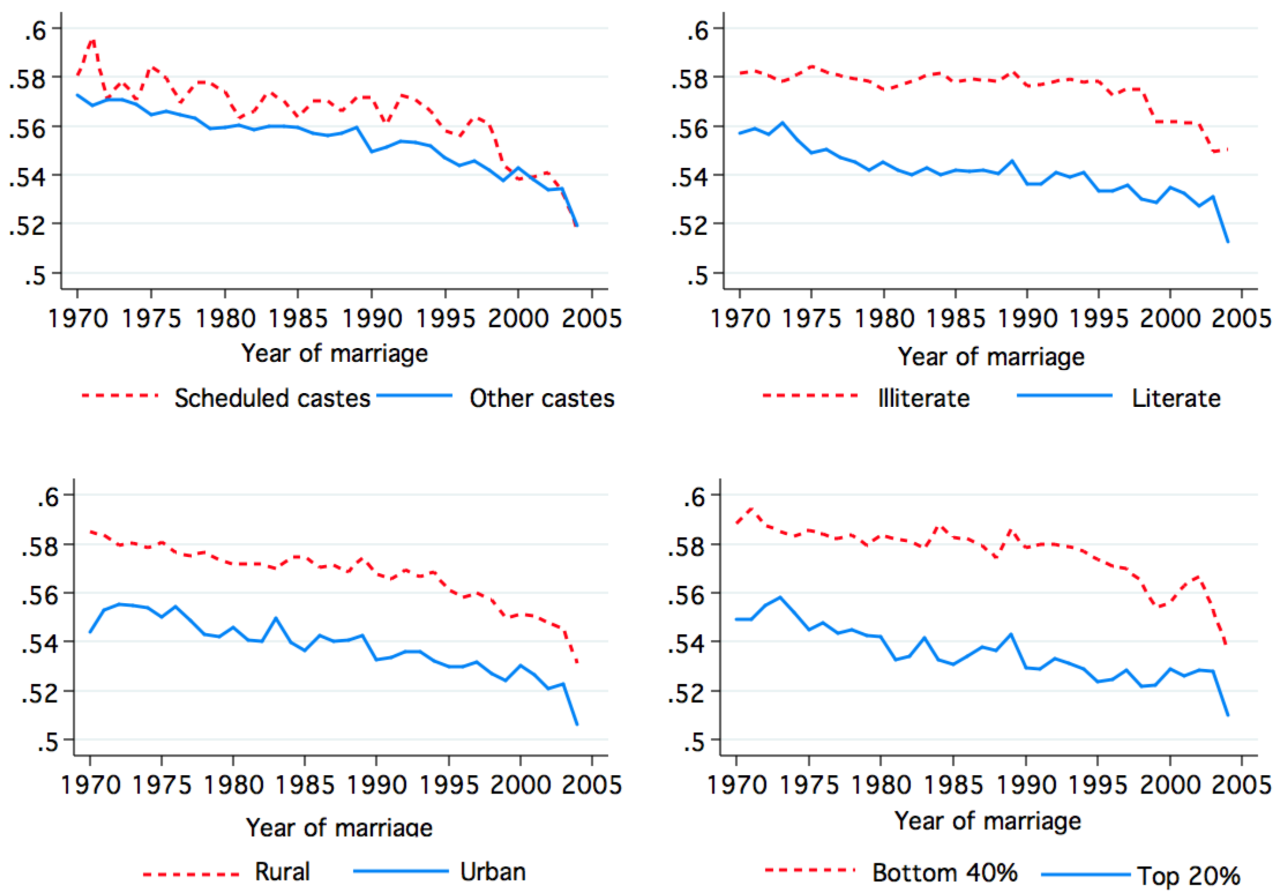
A Additional Figures and Tables

Figure A.1: Testing for recall bias in the reported sex ratio at first birth



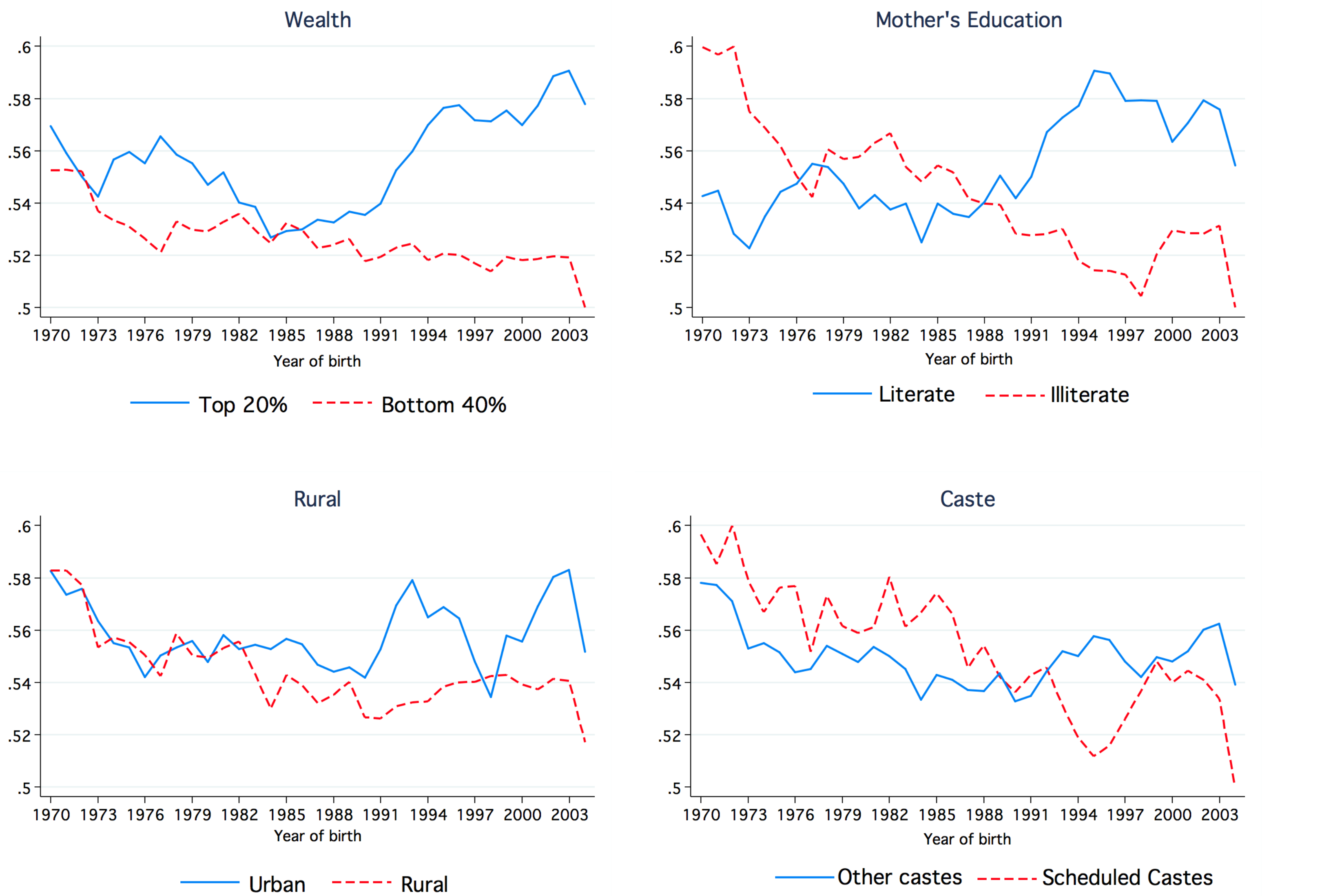
NOTES: This graph checks if there is any systematic recall bias in reports of first child's gender by comparing how the probability that a first birth is reported as female differs across rounds for the same year of birth. In panel A, the thinner solid blue (thicker solid red) line shows how the reported probability that a first birth is female differs in NFHS-2 (NFHS-3) as compared to NFHS-1. In panel B, a similar comparison is made between NFHS-2 and NFHS-3. The dotted lines denote the 95% confidence intervals.

Figure A.2: Trends in ideal fraction of sons



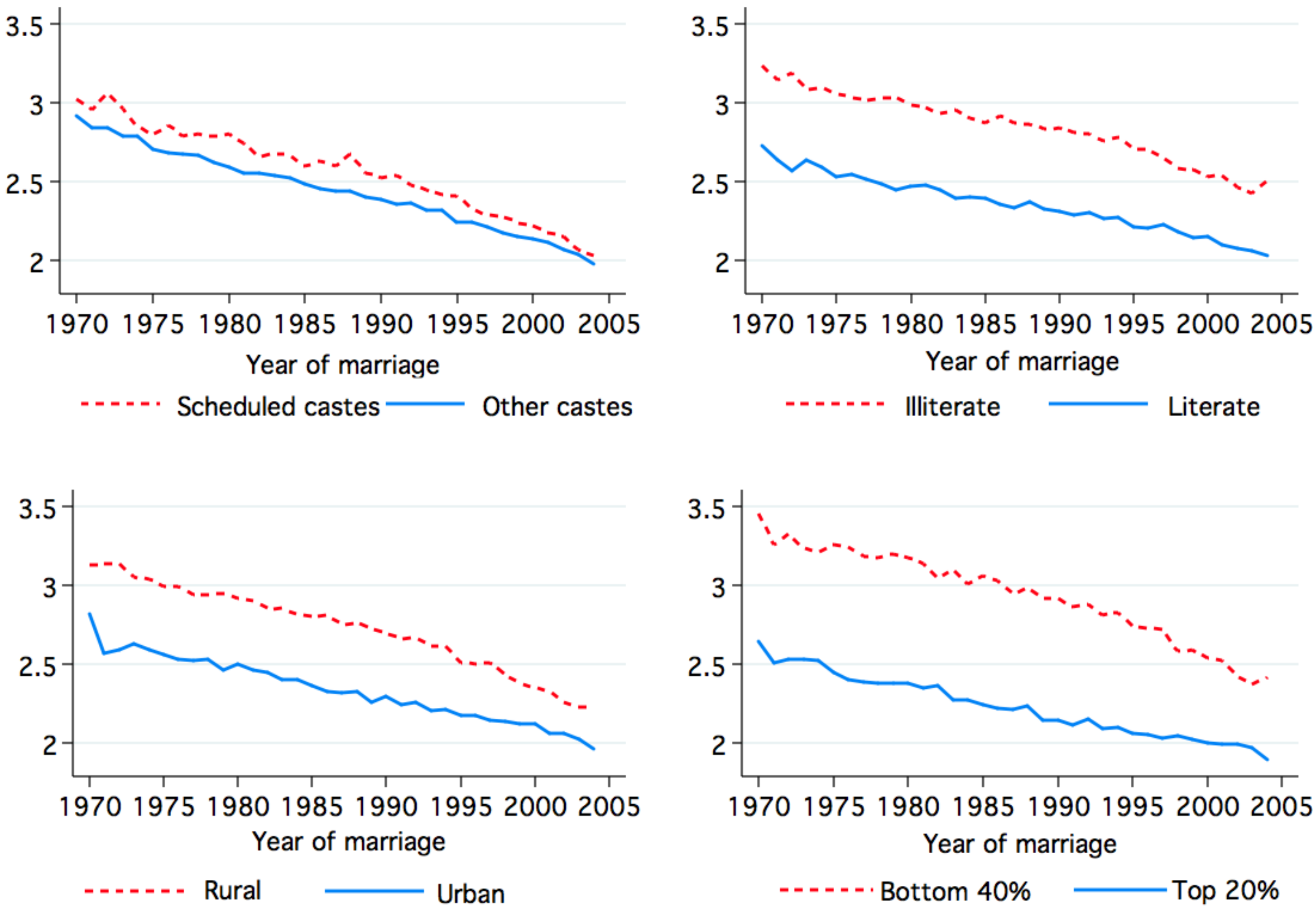
NOTES: The above graphs plot the trend in the 5-year moving average of the reported ideal fraction of sons by year of first marriage, separately for various socioeconomic groups, for mothers who had at least one child at the time of survey. Ideal fraction of sons = $\frac{ideal_{boys}+(0.5*ideal_{either})}{ideal_{kids}}$ if $ideal_{kids} > 0$, where $ideal_{kids}$ is the ideal number of total children; $ideal_{boys}$ is the ideal number of boys; and $ideal_{either}$ is the number of children that parents are indifferent over the gender of, as reported by a woman.

Figure A.3: Trends in actual fraction of sons



NOTES: The graph plots the trend in the 5-year moving average of actual fraction of male births among all births in a year, separately for various socioeconomic groups.

Figure A.4: Trends in ideal number of children



NOTES: The above graphs plot the trend in average reported ideal number of children by the year of first marriage for mothers who had at least one child at the time of survey, separately for various socioeconomic groups.

Table A.1: Under-5 mortality rates by time-period; pooling firstborn-girl and firstborn-boy families

	Male (1)	Female (2)	EFM (3)=(2)-(1)
Pre-ultrasound: 1973-1984	11.72	13.09	1.37
N	35,021	32,428	
Post-ultrasound: 1985-1994	9.53	10.44	0.91
N	84,207	78,018	
Post-ultrasound: 1995-2005	7.41	8.01	0.60
N	51,594	46,804	

NOTES: This table reports the percentage of second- and higher-order children, by child's gender, who suffered from mortality within 5 years of birth, over the three time-periods in our sample. Column (3) reports the difference between the mortality numbers in columns (1) and (2).

Table A.2: Age distribution of missing girls in 2000

	No. missing (in 000s)		% of all missing women		% of under-5 missing girls	
Age group	India (1)	China (2)	India (3)	China (4)	India (5)	China (6)
At birth	184	644	11%	37%	37%	83%
0-1	146	109	9%	6%	30%	14%
1-4	164	23	10%	1%	33%	3%
< 5	494	776	29%	45%		
5-14	93	2	5%	0%		
≥ 15	1125	947	66%	55%		
Total	1712	1727				

NOTES: This table is based on Anderson and Ray (2010) and reports the number of missing girls for various age groups (columns (1) and (2)), missing girls as a percentage of all missing women (columns (3) and (4)), and missing girls across age groups as a percentage of all under-5 missing girls (columns (5) and (6)), separately for India and China. It shows that India and China differ in the age distribution of missing girls. In China, the imbalance in the sex ratio for under-5 children in primarily at birth (83 percent), while there is a more even spread across early childhood in India. In particular, the contribution of post-infancy EFM to the number of missing girls is 33 percent in India but only 3 percent in China

Table A.3: Sex ratio at first birth

Dep Var: Child is female	
$Post1$	0.00496
	(0.00500)
$Post2$	0.00901
	(0.00829)
N	173,293

NOTES: This table tests if ultrasound access affected sex ratio at first birth. The outcome variable is an indicator of birth being female. The sample is restricted to first births and we control for state fixed effects, state-specific linear time trends, and a set of socioeconomic controls (the same as in specification (1)). Standard errors in parentheses are clustered by state. *** 1%, ** 5%, * 10%.

Table A.4: Actual and ideal fertility and fraction of sons; pooling all families

	(1) Actual	(2) Ideal	(3) Actual - Ideal
1. Pre-ultrasound: 1973-1984			
Fertility	2.96	2.21	0.75
Fraction of Sons	0.58	0.57	0.01
N		22,256	
2. Post-ultrasound: 1985-1994			
Fertility	2.16	2.17	-0.01
Fraction of Sons	0.55	0.56	-0.01
N		44,804	
3. Post-ultrasound: 1995-2005			
Fertility	1.91	1.93	-0.03
Fraction of Sons	0.53	0.55	-0.02
N		43,182	

NOTES: This table reports the actual and ideal fertility and fraction of sons as reported by the mother, and the gaps between the two, over the three time-periods in our sample. Ideal fraction of sons = $\frac{ideal_{boys}+(0.5*ideal_{either})}{ideal_{kids}}$ if $ideal_{kids} > 0$, where $ideal_{kids}$ is the ideal number of total children; $ideal_{boys}$ is the ideal number of boys; and $ideal_{either}$ is the number of children that parents are indifferent over the gender of, as reported by a woman.

Table A.5: Balance Tests: Sample means by firstborn sex, REDS

	1973-1984		1985-1994		1995-1999	
	FB	FG	FB	FG	FB	FG
	(1)	(2)	(3)	(4)	(5)	(6)
SC	0.13	0.15	0.14	0.14	0.14	0.15
ST	0.07	0.07	0.07	0.07	0.08	0.07
OBC	0.35	0.35	0.34	0.37	0.33	0.35
Hindu	0.87	0.90	0.87	0.90	0.90	0.89
Muslim	0.08	0.06	0.08	0.06	0.05	0.06
Sikh	0.03	0.03	0.04	0.03	0.04	0.02
Christian	0.01	0.01	0.01	0.01	0.01	0.02
Mother's age at birth	20.59	20.61	20.74	20.84	21.40	21.63
Mother's years of education	2.74	3.06	3.84	3.73	5.37	5.52
Father's years of education	5.54	6.11	6.54	6.44	7.79	8.01
N (1st births)	1,374	899	1,599	1,216	758	646

NOTES: This table compares the socioeconomic characteristics of firstborn-boy (FB) and firstborn-girl (FG) families during the pre-ultrasound period and during the two post-ultrasound periods in the REDS sample. SC, ST, and OBC denote Scheduled Castes, Scheduled Tribes, and Other Backward Classes, respectively. The sample is restricted to first children who were alive at the time of survey.

Table A.6: Neonatal EFM as a function of ultrasound availability and firstborn sex

Dep var: Death within 1 month of birth	(1)	(2)	(3)	(4)
*Firstborn girl * Female*	1.299***	1.247***	1.190***	0.985*
	(0.393)	(0.389)	(0.386)	(0.492)
*Firstborn girl * Female * Post1*	-0.620	-0.573	-0.564	-0.266
	(0.458)	(0.460)	(0.457)	(0.553)
*Firstborn girl * Female * Post2*	-0.600	-0.591	-0.570	-0.448
	(0.403)	(0.398)	(0.385)	(0.442)
N	328,072			
*Female * Post1* and *Female * Post2*	x	x		
*Firstborn Girl * Post1* and *Firstborn Girl * Post2*	x	x		
X_{ijt}		x	x	x
Firstborn Girl x Birth year FE			x	x
Female x Birth year FE			x	x
Female x State FE			x	x
Female x Birth order FE			x	x
Birth order x Birth year FE			x	x
Birth order x State FE			x	x
State x Birth year FE			x	x
Firstborn girl x State FE			x	x
Firstborn girl x Birth order FE			x	x
Mother FE				x

NOTES: This table reports estimates of equation (1) in the text for neonatal mortality of second and higher order births. Neonatal mortality is mortality in the first month of life. Each column is a separate regression. The outcome is an indicator for death within one month of birth. We drop children that are less than 1 month old to allow each child in the sample "full exposure" to the risk of neonatal mortality. We always control for *Female* and fixed effects (FE) for birth year and birth order and for *Firstborn girl* and state FE in columns (1)-(3). The vector X_{ijt} comprises mother's age at birth and, except in column (4), household wealth quintiles, caste, religion, residence in a rural area, educational attainment of child's parents, and mother's birth cohort. Standard errors in parentheses are clustered by state. *** 1%, ** 5%, * 10%.

Table A.7: Under-5 mortality: Heterogeneity by birth order

Birth order	2nd	3rd	4th	5th	6th	7th
	(1)	(2)	(3)	(4)	(5)	(6)
*Firstborn girl * Female*	3.484***	2.971**	1.123	-0.0783	-3.635	20.28
	(0.692)	(1.111)	(1.666)	(3.108)	(3.826)	(13.75)
*Firstborn girl * Female * Post1*	-2.847**	-0.881	1.019	3.061	-0.405	-22.43
	(1.034)	(1.047)	(1.602)	(4.228)	(4.271)	(18.44)
*Firstborn girl * Female * Post2*	-2.093**	-2.463	0.0841	1.044	2.977	-25.86
	(0.789)	(1.523)	(2.335)	(4.091)	(4.331)	(16.45)
N	101,529	64,383	34,318	16,384	6,885	2,539
Baseline mean	11.96	11.89	13.09	14.78	15.53	23.01

NOTES: Estimates corresponding to the specification in column (3) of Table 3 estimated separately for various birth orders. Each column is a separate regression. The outcome measures mortality as % of births that do not survive. Baseline mean refers to the average likelihood of under-5 mortality in the pre-ultrasound period. Standard errors in parentheses are clustered by state. *** 1%, ** 5%, * 10%.

Table A.8: Under-5 mortality: Controls for mother's preferences and gender gaps in school enrollment

Dep Var: Death before age 5	(1)	(2)
*Firstborn girl * Female*	0.793	2.807***
	(1.367)	(0.697)
*Firstborn girl * Female * Post1*	-0.747	-1.374
	(1.009)	(0.932)
*Firstborn girl * Female * Post2*	-1.877**	-2.050**
	(0.687)	(0.823)
Ideal number of children × Female	-0.652***	
	(0.223)	
Ideal sex ratio × Female	2.256***	
	(0.479)	
Ideal number of children × Firstborn girl	-0.414*	
	(0.226)	
Ideal sex ratio × Firstborn girl	0.0418	
	(0.330)	
Ideal number of children × Female × Firstborn girl	0.413	
	(0.334)	
Ideal sex ratio × Female × Firstborn girl	0.572	
	(0.637)	
*Girl-boy enrollment rate 6-11 * Female*		0.606***
		(0.0969)
*Girl-boy enrollment rate 14-17 * Female*		1.484
		(1.094)
N	182,287	179,799

NOTES: Estimates of specification (1) with additional controls for son preference. Ideal fraction of sons = $\frac{ideal_{boys}+(0.5*ideal_{either})}{ideal_{kids}}$ if $ideal_{kids} > 0$, where $ideal_{kids}$ is the ideal number of total children; $ideal_{boys}$ is the ideal number of boys; and $ideal_{either}$ is the number of children that parents are indifferent over the gender of, as reported by a woman. Standard errors in parentheses are clustered by state. *** 1%, ** 5%, * 10%.

Table A.9: Mother's reported son preference as a function of ultrasound availability

	Ideal Fraction of Sons
*Firstborn girl * Female*	0.011***
	(0.003)
*Firstborn girl * Female * Post1*	-0.002
	(0.002)
*Firstborn girl * Female * Post2*	0.001
	(0.003)
N	471,559

NOTES: Estimates of specification (1) where dependent variable is Ideal fraction of sons = $\frac{ideal_{boys}+(0.5*ideal_{either})}{ideal_{kids}}$ if $ideal_{kids} > 0$, where $ideal_{kids}$ is the ideal number of total children; $ideal_{boys}$ is the ideal number of boys; and $ideal_{either}$ is the number of children that parents are indifferent over the gender of, as reported by a woman. *** 1%, ** 5%, * 10%.

Table A.10: Health investments: A different data set (NFHS)

	# Antenatal Checks	At least 1 vaccination	# Months breastfed
	(1)	(2)	(3)
*Firsborn girl * Female*	-0.041	-0.029**	-0.422**
	(0.056)	(0.012)	(0.155)
*Firsborn girl * Female * Post2*	0.002	0.025**	0.129
	(0.060)	(0.012)	(0.258)
N	92,525	79,809	83,480

NOTES: NFHS data. Investments are only queried for cohorts born in a few years before each survey so there are no pre-ultrasound cohorts. The comparison here is therefore across the two post-ultrasound periods. The estimates are from the following specification for child i of birth order b born to mother j in year t: $I_{ibjt} = \alpha + \beta G_j * F_i * Post_t^2 + \gamma G_j * F_i + \omega_t G_j + \sigma_t F_i + \psi_b F_i + X_{ijt}^{'}\tau + \delta_s F_i + \rho_{bt} + \eta_{bs} + \phi_{st} + \epsilon_{ibjt}$. Among children who were at least 12 months old at the time of survey, we define a child to be fully immunized if he or she had received the eight most common vaccines by that time. Standard errors in parentheses are clustered by state. *** 1%, ** 5%, * 10%.

Table A.11: Fertility: Heterogeneity by socioeconomic status using the cross-sectional approach

	Mother's Education		**Wealth**		**Mother's Employment**	
	Illiterate	Literate	Bottom 40%	Top 20%	Paid employment = 0	Paid employment = 1
	(1)	(2)	(3)	(4)	(5)	(6)
Firstborn girl	0.113***	0.198***	0.126***	0.183***	0.172***	0.142***
	(0.020)	(0.018)	(0.031)	(0.019)	(0.014)	(0.021)
*Firstborn girl * Post1*	-0.063*	-0.110***	-0.060	-0.099***	-0.114***	-0.054*
	(0.031)	(0.023)	(0.042)	(0.029)	(0.018)	(0.030)
*Firstborn girl * Post2*	-0.059**	-0.148***	-0.069*	-0.121***	-0.132***	-0.088***
	(0.028)	(0.024)	(0.036)	(0.027)	(0.021)	(0.029)
N	46,597	72,066	36,831	36,352	78,717	39,840
Baseline mean	3.230	2.812	3.101	2.769	3.026	2.964

	Caste		**Rurality**	
	SC	General/OBC/ST	Rural	Urban
	(1)	(2)	(3)	(4)
Firstborn girl	0.146***	0.154***	0.134***	0.190***
	(0.039)	(0.012)	(0.017)	(0.015)
*Firstborn girl * Post1*	-0.014	-0.098***	-0.095***	-0.079***
	(0.056)	(0.016)	(0.023)	(0.023)
*Firstborn girl * Post2*	-0.032	-0.124***	-0.108***	-0.122***
	(0.041)	(0.018)	(0.026)	(0.018)
N	17,043	101,620	71,155	47,508
Baseline mean	3.243	2.966	3.109	2.850

NOTES: Estimates of specification (3). The dependent variable is the number of births at the time of interview. Standard errors in parentheses are clustered by state. Baseline mean refers to the mean fertility of mother at time of interview, with the specific socioeconomic status who had their first and last birth in the pre-ultrasound period 1972-1984. *** 1%, ** 5%, * 10%.

B Variable Descriptions

- Excess Female Mortality: Female mortality - Male mortality
- Neonatal mortality: Death within one month of birth
- Under-5 mortality: Death before age five
- $Post_t^1$: indicator variable for $t \in 1985 - 1994$
- $Post_t^2$: indicator variable for $t \in 1995 - 2005$
- F_i: child i is female
- G_j: first child of mother j is female
- Allopathic treatment: an indicator variable that equals 1 if a sick child received medical help from an allopathic doctor during the past year, and 0 otherwise
- Exp on education: amount (in Rupees) spent on fees, books, uniform, hostel pocket money, transportation, and private coaching during last year
- Doctors' fees: amount (in Rupees) spent on doctors' fees last year
- Medicine and special food: amount (in Rupees) spent on medicine and special food last year
- Medical exp: amount (in Rupees) spent on doctors' fees, medicine, and special food last year
- Ideal sex ratio: Ideal number of sons/ Ideal number of daughters (as reported by the mother)
- Ideal number of children: Mother's self-reported ideal number of children of any sex
- Education categories: no education, incomplete secondary education, and secondary or higher education
- Categories for mother's birth cohort: 1942-1960, 1961-1970, and 1971-1987
- Categories for Mother's age at birth: 12-15 years, 16-18 years, 19-24 years, 25-30 years, and 31-49 years
- Caste categories: Scheduled Caste (SC), Scheduled Tribe (ST), and Others
- Religion categories: Hindu, Muslim, and Others

C Contributions of Breastfeeding and Vaccination

Our estimates suggest that ultrasound access reduced under-5 mortality by 1.841 p.p. (column (4) in Table 3) and increased the likelihood of being breastfed for at least 24 months by 27.5 p.p. (result available upon request). Since there was no significant effect on breastfeeding during the first year of birth, we assume that the 27.5 p.p. increase took place between 12 and 24 months from birth. According to the World Health Organization (2000), breastfeeding between the ages one and two decreases mortality by 50 percent relative to no breastfeeding. Applying this factor to the share of children who are being breastfed and the mortality rate during 12-24 months,[60] the implied mortality rate for breastfed children is 1.35 percent[61] and is 2.7 percent (1.35 * 2 = 2.7) for non-breastfed children in the 12-24 months range. This implies that not being breastfed during the 12-24 months age range increases the risk of mortality by 1.35 p.p. (2.7 - 1.35 = 1.35). If breastfeeding disparities (during 12-24 months) were the only cause of under-5 mortality differences by gender, the EFM decline due to improvements in the breastfeeding gender gap would be 0.16 p.p. (0.275 * 1.35 = 0.371). Thus breastfeeding explains about 20 percent (0.371/1.841 = 0.202) of the estimated EFM decline.

Moreover, we find that ultrasound availability increased the probability of a child receiving at least one vaccination by 0.076 to 0.118 (triple-interaction coefficients in column (6) of Table 4). The average number of vaccinations (conditional on receiving at least one vaccination) for girls preceded by a firstborn girl during the early diffusion period is 6.64 in NFHS data. Thus the estimated effects for at least one vaccination translate into an average increase in the number of vaccinations of 0.505 (6.64 * 0.076 = 0.505) to 0.784 (6.64 * 0.118 = 0.784). Oster (2009) suggests that each vaccination reduces mortality during ages 1 to 4 by 0.26 p.p.. Thus the implied effect on EFM through vaccination is 0.131 p.p. (0.26 * 0.505 = 0.131) to 0.204 p.p. (0.26 * 0.784 = 0.204), which translates into 7 percent (0.131/1.841 = 0.071) to 11 percent (0.204/1.841 = 0.111) of the decline in EFM.

Note that the mortality measure is 12-36 months in Jayachandran and Kuziemko (2011) and is 1-4 years in Oster (2009). However, any exogenous change in these mortality measures would generate an almost one-to-one change in the mortality measure we use, i.e., death before age five.

[60] We assume that the share of children who are being breastfed and the mortality rate during 12-24 months are the same as those for 12-36 months used in Jayachandran and Kuziemko (2011).

[61] Solving $0.481x + 2(1 - 0.481)x = 2.05$ yields $x = 1.35$, where 0.481 is the fraction of children aged 12 to 36 months that are being breastfed in the sample analyzed by Jayachandran and Kuziemko (2011).

D Substitution of Prenatal for Postnatal Girl Mortality

Here we use our regression estimates to calculate the magnitude of substitution from postnatal EFM to prenatal sex-selection as the ratio of the number of girls selectively aborted and the number of girls who survived due to ultrasound (and would have otherwise died postnatally). Let:

- N: annual number of births in India
- N_{FG}: annual number of births in India that are preceded by a firstborn girl
- M_{FG}: fraction of births that are male among N_{FG}
- F_{FG}: fraction of births that are female among N_{FG}
- Δ: the pre-post (counterfactual) difference

Then, the number of "missing girls" each year, i.e., the difference between the expected number of female births (given the observed number of male births and the natural sex ratio at birth) and the observed number of female births is given by:

$$\underbrace{\frac{0.49}{0.51}(N_{FG} * M_{FG})}_{\text{Expected \#female births given the observed \#male births}} - \underbrace{N_{FG} * F_{FG}}_{\text{Observed \#female births}} = N_{FG}\left(\frac{0.49}{0.51} - \frac{F_{FG}}{0.51}\right)$$

The change in the annual number of missing girls due to ultrasound access is then calculated as the difference between the number of missing girls during the pre- and the post-ultrasound periods:

$$\begin{aligned}
&N_{FG,post}\left(\frac{0.49}{0.51} - \frac{F_{FG,post}}{0.51}\right) - N_{FG,pre}\left(\frac{0.49}{0.51} - \frac{F_{FG,pre}}{0.51}\right)\\
=&N_{FG,post} * \frac{0.49}{0.51} - N_{FG,post} * \frac{F_{FG,post}}{0.51} - N_{FG,pre} * \frac{0.49}{0.51} + N_{FG,pre} * \frac{F_{FG,pre}}{0.51}\\
=&\Delta N_{FG} * \frac{0.49}{0.51} - N_{FG,post} * \frac{F_{FG,post}}{0.51} + N_{FG,pre} * \frac{F_{FG,pre}}{0.51}
\end{aligned}$$

(Adding and subtracting $N_{FG,post} * \frac{F_{FG,pre}}{0.51}$)

$$\begin{aligned}
=&\Delta N_{FG} * \frac{0.49}{0.51} - N_{FG,post} * \frac{F_{FG,post}}{0.51} + N_{FG,post} * \frac{F_{FG,pre}}{0.51} - N_{FG,post} * \frac{F_{FG,pre}}{0.51} + N_{FG,pre} * \frac{F_{FG,pre}}{0.51}\\
=&\Delta N_{FG} * \frac{0.49}{0.51} + N_{FG,post} * \Delta F_{FG} * \frac{-1}{0.51} - \Delta N_{FG} * \frac{F_{FG,pre}}{0.51}\\
=&\underbrace{\Delta N_{FG} * \left(\frac{0.49}{0.51} - \frac{F_{FG,pre}}{0.51}\right)}_{\text{(conception effect)}} + \underbrace{N_{FG,post} * \Delta F_{FG} * \frac{-1}{0.51}}_{\text{(sex-selective abortions effect)}}
\end{aligned}$$

The change in the number of missing girls due to ultrasound can be decomposed into a "conception effect" and a "sex-selective abortion effect," as defined above. The sex-selective abortion effect refers to the fact that a smaller *fraction* of *post-ultrasound* births preceded by a firstborn girl are now female, where as the conception effect is driven by the change in the number of births that are preceded by a firstborn girl.

The number of excess postnatal female deaths in a year equals:

$$N_{FG} * F_{FG} * EFM_{FG}$$

where EFM_{FG} refers to the difference between the probability of death by age 5 among children preceded by a firstborn girl and a firstborn boy.
The change in the number of excess postnatal female deaths can thus be written as:

$$\begin{aligned} & N_{FG,post} * F_{FG,post} * EFM_{FG,post} - N_{FG,pre} * F_{FG,pre} * EFM_{FG,pre} \\ = & \underbrace{N_{FG,post} * F_{FG,post} * (\Delta EFM_{FG})}_{\text{(behavioral effect)}} + \underbrace{N_{FG,post} * (\Delta F_{FG}) * EFM_{FG,pre}}_{\text{(mechanical sex-selection effect)}} + \underbrace{(\Delta N_{FG}) * F_{FG,pre} * EFM_{FG,pre}}_{\text{(conception effect)}} \end{aligned}$$

The change in the number of excess postnatal female deaths can also be decomposed into three components: the "behavioral effect" refers to the EFM decline due to, say, better postnatal health investments in girls in the post-ultrasound period; the "mechanical sex-selection effect" is driven by the decrease in the *fraction* of *post-ultrasound* births that are preceded by a firstborn girl; and the "conception effect" that reflects the change in the *number* of births that are preceded by a firstborn girl itself.

The table below uses the aforementioned formulae to calculate the number of births preceded by a firstborn girl during the pre-ultrasound (in column (1)), during the late diffusion period (in column (2)), and the change in the number of missing girls as the difference between columns (1) and (2) in column (3). EFM_{FG} in column (1) is the coefficient of *Firstborn girl * Female* and in column (2) is the coefficient of *Firstborn girl * Female * Post2* from column (4) in panel B of Table 3. During the pre-ultrasound period, the fraction of females in births preceded by a firstborn girl was 47.9 percent. Ultrasound access decreased this fraction by 1.8 p.p..[62] We calculate $\frac{N_{FG}}{N}$ from the entire sample since the pre-ultrasound sample of births is likely to be underreported due to recall bias. The number of births in column (1), N, is the 1995 figure from UN Statistics Division. The number of mothers, m, is obtained as $\frac{N}{\text{General Fertility Rate}}$ using the general fertility rate estimated from NFHS-2 by Retherford and Mishra (2001) as 131.53 per 1000 women. ΔN_{FG} is estimated to

[62] This number is derived from a regression specification similar to the one estimated by Bhalotra and Cochrane (2010).

be equal to -0.0063 as is obtained from the fertility equation using the conversion method described below.

Thus, the number of "missing girls" increased by 305,496 per year due to ultrasound access. The estimated decline in EFM, on the other hand, implies that the annual number of postnatal female deaths by age 5 fell by 60,879. The ratio of increase in sex-selective abortions and the decrease in EFM is 5.02, i.e., for every 5.02 girls aborted, one girl survived due to access to ultrasound technology.

Table: Decomposition and Simulation

	Pre (Pre-ultrasound) (1)	Post (Late diffusion) (2)	Δ (Post - Pre) (3)
A. Regression Estimates			
EFM_{FG}	0.0277	0.0077	-0.020
F_{FG}	0.4790	0.4605	-0.0185
$\frac{N_{FG}}{N}$	0.3340		
B. Decomposition			
N	27,300,000		
m (#mothers)	$\approx$ 207,557,211		
m_{FG} (#mothers with a firstborn girl)	$\approx$ 0.3340*m		
N_{FG}	0.3340*(N) = 9,118,200	0.3340*(N)-0.0063*m_{FG} = 8,681,458	-0.0063*m_{FG} = -436,742
Δ number of excess postnatal female deaths:			
(1) $N_{FG,post} * F_{FG,post} * (\Delta EFM_{FG})$		$\approx$ -79,956	
(2) $N_{FG,post} * (\Delta F_{FG}) * EFM_{FG,pre}$		$\approx$ -4,449	
(3) $(\Delta N_{FG}) * F_{FG,pre} * EFM_{FG,pre}$		$\approx$ -5,795	
(4) = (1)+(2)+(3)		= -90,200	
Δ number of missing girls:			
(5) $\Delta N_{FG} * (\frac{0.49}{0.51} - \frac{F_{FG,pre}}{0.51})$		$\approx$ -9,420	
(6) $N_{FG,post} * \Delta F_{FG} * \frac{-1}{0.51}$		$\approx$ 314,916	
(7) = (5)+(6)		= 305,496	
Δ number of missing girls / Δ EFM:			
(8) = (7)/(4)		= -3.39	

– Conversion of treatment effect from the fertility specification (2) to expected number of children born in a year.

Let N to be the number of total children born in a single year. Then

$$N = N_1 + N_{FG} + N_{FB}$$

where N_1 is the number of firstborn children and N_{FB} and N_{FG} are the number of second or higher order births respectively preceded by a firstborn girl and a firstborn boy.

- m_{FG}: number of women in the 15-49 age-group with a firstborn girl
- $S_{a,FG}$ and $S_{a,FB}$: share of women aged a who have a firstborn girl and firstborn boy, respectively
- $P_{a,FG}$ and $P_{a,FB}$: probability that a women who has a firstborn girl (firstborn boy) and who is of age a in a given year gives birth in that year

$$\implies N_{FG} = (\textstyle\sum_{a=15}^{a=49} P_{a,FG} S_{a,FG}) * m_{FG}$$

$$\begin{aligned}
&\Delta N_{FG} = N_{FG,post} - N_{FG,pre} \\
=&(\sum_{a=15}^{a=49} P_{a,FG,post} S_{a,FG,post} - \sum_{a=15}^{a=49} P_{a,FG,pre} S_{a,FG,pre}) * m_{FG} \\
=&(\sum_{a=15}^{a=49} (P_{a,FG,post} - P_{a,FG,pre}) S_{a,FG}) * m_{FG}
\end{aligned}$$

assuming that $S_{a,FG,pre} = S_{a,FG,post} = S_{a,FG}$. This assumption is reasonable since ultrasound access had no effect on the decision to have a first birth or its sex ratio.

The treatment-on-the-treated effect we estimated in equation (2) equals

$$\begin{aligned}
&E[N|post, FG] - E[N|pre, FG] \\
=&\sum_{a=15}^{a=49} (P_{a,FG,post} - P_{a,FG,pre}) * S_{49,FG} \\
+&\sum_{a=15}^{a=48} (P_{a,FG,post} - P_{a,FG,pre}) * S_{48,FG} \\
+&\sum_{a=15}^{a=47} (P_{a,FG,post} - P_{a,FG,pre}) * S_{47,FG} \\
+&...
\end{aligned}$$

If we assume uniform impact of ultrasound access on the probability of birth across age-groups, i.e., $P_{a,FG,post} - P_{a,FG,pre} = P_{FG,post} - P_{FG,pre}$ and assuming $S_{a,FG} = S_{FG}$:

$$E[N|post, FG] - E[N|pre, FG] = S_{FG} * (P_{FG,post} - P_{FG,pre}) * (1 + 2 + 3 + ... + 35)$$

$$
\begin{aligned}
&\sum_{a=15}^{a=49}(P_{a,FG,post} - P_{a,FG,pre})S_{a,FG} \\
=&S_{FG} * (P_{FG,post} - P_{FG,pre}) * 35 \\
=&\frac{35}{(1 + ... + 35)} * (E[N|post, FG] - E[N|pre, FG]) \\
=&\frac{35}{630} * (E[N|post, FG] - E[N|pre, FG]) \\
\approx&0.056 * (E[N|post, FG] - E[N|pre, FG]) \\
\approx&0.056 * (-0.112) = -0.0063
\end{aligned}
$$