Policy Research Working Paper 9390

On the Quantity and Quality of Girls

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WORLD BANK GROUP
Development Economics
Development Research Group
September 2020

Policy Research Working Paper 9390

Abstract

The introduction of prenatal sex-detection technologies in India has led to a phenomenal increase in abortion of female fetuses. This paper examines the impacts of this on girl relative to boy mortality rates after birth, using data from 1973–2005. The analysis finds a narrowing of the gender gap in under-5 mortality rates, in line with surviving girls being more wanted. The estimates show that for every three aborted girls, one additional girl survives to age five. Investigation of the mechanisms finds a narrowing of gender gaps in parental investments in children, moderation of son-biased fertility stopping, and shrinking of the gap between actual and desired fertility. Heterogeneity in fertility responses suggests a shift in the distribution of girls toward lower socioeconomic status families. The findings have implications not only for counts of missing girls, but also for the later life outcomes of girls.

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On the Quantity and Quality of Girls:

Fertility, Parental Investments, and Mortality*

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Note: The findings, interpretations, and conclusions expressed in this paper are entirely those of the authors. They do not necessarily represent the views of the World Bank and its affiliated organizations, or those of the Executive Directors of the World Bank or the governments they represent.

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.

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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 (1990a), Gruber et al. (1999), Donohue and Levitt (2001), Bitler and Zavodny (2002), 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)).[1] 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)).[2] Even in the absence of son preference, and even in environments in which contraception is available, parents do not always want the children that they conceive: e.g., 21 percent of all pregnancies in 2011 in the United States ended in abortion (Guttmacher Institute).[3]

When fetal sex determination was impossible, parents could adjust the gender composition of their children in two ways. First, by continuing childbearing until they achieved, for instance, the desired number of sons. Several studies document son-biased fertility stopping behavior,[4] which results in girls having more siblings than boys (Clark (2000), Bhalotra and van Soest (2008), Filmer et al. (2009), 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,

[1] Although ultrasound technology can be used as a diagnostic tool for various health conditions, for the purposes of this paper, we focus on its use for fetal sex detection.

[2] 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. Note that prenatal sex detection has been illegal in India in public facilities since 1976 and in private facilities since 1994.

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

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

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)).[5] In this 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 (we demonstrate rather than assume this); 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 also directly investigate, 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 purges unobservable trends that equally affect both boys and girls. Using three axes of variation allows us to control not only for fixed effects for birth-year, state, and birth-order but also for interactions between these variables, and for their interactions with indicators for female child and for sex of the firstborn child in the family. Our data span the 1973-2005 period.

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 percent after the introduction of ultrasound technology. These estimates are robust to conditioning upon mother fixed effects that control for woman-specific preferences that influence uptake of the new technology. We conduct a number of data and specification checks, including a check on whether gendered recall bias might drive our findings, which we conclude is unlikely.

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

Using different approaches, we find no evidence that our finding of post-ultrasound declines in excess female mortality (henceforth EFM, or the excess of female over male mortality) in families with firstborn girls was driven by changes in son preference specific to this group. Instead, the mechanism appears to be that parents use selective abortion to align the timing, number, and sex of births to their preferences. We substantiate this by analyzing four sets of additional outcomes—changes in parental investments, son-biased fertility stopping, birth spacing, and the gap between actual and desired fertility.[6]

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 significantly reduced post-ultrasound in firstborn-girl families, and this can explain 26 to 29 percent of the observed decline in under-5 excess female mortality.[7] Note that the critical investments relevant for early childhood mortality are not expenditures but non-monetary investments, such as breastfeeding and vaccination.

Analyzing fertility, we find that the gender gap in sibling size between firstborn-girl and firstborn-boy families (which, pre-ultrasound, was 0.14) declined by 40 to 50 percent. Since sex-selective abortion allows parents to avoid unwanted children, we tested whether availability of ultrasound drove actual fertility closer to desired fertility. Our estimates imply that undesired fertility, which was 0.115 births in firstborn-girl relative to firstborn-boy families in the pre-ultrasound years, declined by nearly 74 percent.

We also investigated birth spacing. Families with firstborn girls had shorter birth intervals in the pre-ultrasound period, consistent with their use of fertility continuation as a means of achieving sons (Bhalotra and van Soest (2008), Jayachandran and Kuziemko (2011), Milazzo (2018)). In line with sex-selective abortion being concentrated in firstborn girl families and with abortion mechanically lengthening birth spacing (because a conception occurs but no birth is observed), we find an increase in birth spacing in firstborn girl families relative to firstborn boy families in the post-ultrasound period relative to the pre-ultrasound period.

In sum, we observe a fairly dramatic post-ultrasound break in the trends in mortality and sibship size for girls relative to boys in firstborn-girl families (relative to firstborn-boy families). We also elucidate the mechanisms quite clearly by examining birth spacing, excess fertility, and parental investments in child health—all of which fit together to tell a coherent story. Nevertheless, the decline in postnatal death of girls only partially offsets the rise in female feticide. Our estimates

[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.

[7] 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.

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.

Investigation of heterogeneity in the impacts of availability of prenatal sex-detection technology on birth outcomes further elucidates mechanisms. We find that the decline in fertility in firstborn-girl families is, in general, larger among high socioeconomic status (SES) families, consistent with previous work that shows that sex-selective abortion in India is more prevalent among urban, educated, upper-caste, and wealthy women (Jha et al. (2006), Bhalotra and Cochrane (2010)). This is what we would expect given that families tend to substitute sex-selective abortion for son-biased fertility-continuation.

The estimated decline in EFM among women with firstborn daughters is, however, broadly similar across socioeconomic groups. This is not surprising given that baseline mortality is higher among low-SES families, making them more "treatable" in this dimension. Overall, our 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. 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, China decreased neonatal (but not post-neonatal) female mortality for higher-parity births conditional on SES, and that it reduced fertility at third and higher parity. They do not examine health investments.[8] 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).[9] This paper is also related to the literature showing that allowing parents to sift (through abortion), so that they retain only "wanted" children leads to higher average investments in children (see the references in paragraph 1 of the paper, and also see Grossman and Joyce (1990b) and Levine and Zimmerman (1996)).[10]

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

[9]In comparison, before the 1985 abortion legalization that Lin et al. (2014) examine, the average post-neonatal infant EFM in Taiwan, China in 1981 was -0.11 p.p., i.e., small, and biased in favor 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 is 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.

[10]The objective of this study is neither to condone nor to reject the practice of sex-selective abortion. Our

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.[11] 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 this puzzle. In Section 8 we delineate the difference in empirical strategy, highlighting that we are able to identify meaningful differences because we leverage relevant quasi-experimental variation between families (in the sex of the firstborn child) and over time (on account of policy-induced variation in access to ultrasound technology). 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 9).

The rest of the paper is organized as follows. Section 2 describes the Indian context. Sections 3 and 4 discuss the data and the empirical strategy. Section 5 presents results for EFM and Section 6 explores the underlying mechanisms. Section 7 presents estimates of the implied magnitude of substitution between postnatal discrimination and sex-selective abortions due to ultrasound access. Section 8 relates our paper to a previous literature and Section 9 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.[12] Abortion is legal if the pregnancy that it terminates endangers the 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

paper offers a deeper understanding of parental decision-making about fertility and child investments and presents a more complete picture of the effects of access to ultrasound technology in India.

[11] 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.

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

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.

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)).[13] The trend in ultrasound use (also in Figure 1) closely tracks the supply of ultrasound machines.[14] 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.[15] Additional amendments to the MTP Act in 2002 and 2003 increased public sector provision 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 among women's and hu-

[13] 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.

[14] 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.

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

man rights organizations.[16] 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.[17] 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) is 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 Data and Descriptive Statistics

To examine mortality and fertility, we use 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.[18] The sample comprises 503,316 births of 232,259 mothers that occurred in 1973-2005.[19]

[16]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)).

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

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

[19]Women who do not survive to the date of interview, for whatever reason, are not in the NFHS data, and nor are their children. Sex-selective abortion in the post-ultrasound period may elevate maternal mortality

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.[20] 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 to analyze effects on health investments.[21]

For the mortality and postnatal health specifications (presented in the next section), we pool all births of the surveyed women to create a child-level data set. The birth hazard sample is a retrospective mother-year panel that we create from women's retrospective birth histories; women enter this panel in the year of marriage and exit in the year of survey. The sample used for the fertility stock estimates simply pools all surveyed women to form a woman-level data set. Summary statistics for mortality, (actual) fertility, and child health investments are in Table 1. In Section 5.1.1, we examine the robustness of our results to data-related issues, including gendered recall bias. The maximum recall period is 20 years in our NFHS sample and is 26 years in our REDS sample.

4 Empirical Strategy

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 conception and investment decisions in a gendered manner, with implications for excess female mortality (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.[22] However, the wave of economic liberalization in India that led

risk, and more so among families with firstborn girls (who are more treated). However, maternal mortality is numerically small and the share of maternal mortality associated with abortion even smaller; so, any resulting bias is unlikely to drive our findings. We, nevertheless, considered the direction of any such bias. If, as observed in Jha et al. (2006) and Bhalotra and Cochrane (2010), high-SES women are more likely to commit sex-selection then, given that their children face lower mortality risks, our estimates of EFM decline will tend to be conservative. However, this downward bias may be somewhat offset if low-SES women, even with lower take up of abortion, are more likely to die of abortion.

[20] 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.

[21] REDS is restricted to rural women from 16 major states. 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. Since REDS data does not report the year of birth or death for children that did not survive until the year of survey, we cannot perform a bounding exercise for this bias. However, our results for postnatal health investments using NFHS data, that does not suffer from this issue, are broadly similar.

[22] 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

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 framework.

Previous work has shown that more educated and wealthier women are more likely to engage in sex-selective abortion in India (Jha et al. (2006), Bhalotra and Cochrane (2010)). We directly account for selection into uptake by estimating a specification that includes woman fixed effects. Even in the absence of woman fixed effects, by leveraging variation in firstborn child's sex, we obtain variation in the risk of sex-selection that is independent of socioeconomic status (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. Our assumption that the sex of the first child is randomly determined is supported by the data. We demonstrate this as follows. First, the top left graph in Figure 3 shows that the proportion of females among first births in India lies within the normal range that one would expect in the absence of manipulation (48.80 percent - 49.26 percent); and it shows no tendency to increase in the post-ultrasound period. This contrasts sharply with the time profile of the sex ratio for higher order births preceded by a girl.[23] Second, Table 2 shows that families with firstborn boys and firstborn girls are similar along a number of observables.[24]

Third, we combine these two stylized descriptive patterns in a regression framework—see Table A.3, where firstborn sex is the dependent variable. Column 1 shows that the share of firstborn girls (as opposed to boys) does not change post-ultrasound, conditional on state fixed effects, linear time trends, and indicators of SES. In column 2, the specification is enriched to allow interactions between every SES indicator and the post-ultrasound indicator. This is relevant because previous work shows that the proclivity to commit sex-selective abortion varies systematically with SES (Bhalotra and Cochrane (2010)). The interaction terms are jointly insignificant, indicating no

the costs of ultrasound and abortion were not prohibitive even for relatively poor households.

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

[24]Table A.4 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 (1995-1999) because we use the 1999 survey. The sample sizes in Table 2 are smaller than in other tables because there is only one first birth per woman, but multiple births or multiple years per woman in the other tables. The reason we have more firstborn-girl women in more recent years in the data is that we construct a retrospective woman-year panel from three cross-sections and firstborn-girl mothers have higher mortality as shown in Milazzo (2018), causing fewer of them to appear in the sample when we project backwards from cross-sections of surviving women.

significant change in the sex-composition of first births by socio-economic group after ultrasound technology became available.

Exogeneity of firstborn sex has also been previously defended (Das Gupta and Bhat (1997), Visaria (2005), Bhalotra and Cochrane (2010)). It also 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.5), 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 and daughters are 1.38 and 1.01, respectively.

Pre-trends not different by firstborn sex. In order to test for a significant difference in trends in under-5 mortality and health investments 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 the outcomes of interest 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 coefficients from these regressions are presented in Figure 4 which demonstrates no significant divergence between EFM, immunization, and breastfeeding in firstborn-boy and firstborn-girl families for pre-ultrasound cohorts for children of parity greater than one. The same is true for pre-trends estimated separately for the sub-samples of boys and girls (see Figure A.1).

Firstborn sex predicts sex-selection at higher parities. Previous studies, namely 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.[25] So, the interaction with first child's sex captures the differential incentives to sex-select among otherwise similar families.

4.1 Unconditional Trends in Outcomes

Our main findings are evident in the raw data. 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.

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

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, and not boys, in this group, narrowing the EFM gap among firstborn-girl families and, hence, overall. As we shall observe in Section 8, ignoring the distinction 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) and families with firstborn sons as the control group.

Table 1 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.

4.2 Regression Specifications

To capture the two trend breaks in availability of ultrasound, 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) confirmed 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.

4.2.1 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:[26]

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

The dependent variable, Y_i, is either a mortality indicator for child i or a measure of parental investments in children, including breastfeeding and immunization.[27] The variable $FirstbornGirl_j$ equals one if the first child of mother j is a girl, and is zero otherwise. The dummy variable $Female_i$ 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. We show results for under-5

[26]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.

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

mortality and also for mortality in each year between birth and age 5. In 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.

The 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 age at birth, caste, religion, and residence in a rural area. We control for the main effects of $FirstbornGirl_j$ and $Female_i$ and fixed effects for state, birth year (or cohort, of the child), and birth order—importantly, we also include all pairwise interactions between these controls in our specification.

Thus, we allow birth cohort fixed effects to vary by firstborn sex ($\omega_t FirstbornGirl_j$), by child gender ($\sigma_t Female_i$), 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 ($\delta_s Female_i$, $\nu_s FirstbornGirl_j$, $\psi_b Female_i$, $\xi_b FirstbornGirl_j$), and state fixed effects to vary by birth order (η_{bs}) and birth year (ϕ_{st}).

Child gender-specific cohort fixed effects ($\sigma_t Female_i$) account for any nationwide changes that may influence gender gaps in the outcomes, including any decline in son preference or 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 ($\omega_t FirstbornGirl_j$) 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 ($\delta_s Female_i$) and sex of the firstborn child ($\nu_s FirstbornGirl_j$) 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 ($\psi_b Female_i$, $\xi_b FirstbornGirl_j$) given previous evidence that son preference varies with birth order (Jayachandran and Pande (2017)) 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., state-specific trends in son preference, 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; we have 25 states in our sample.[28]

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. In an extension, we also allow interactions of all covariates with child gender. 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 a variant of 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. Although we show results with and without these endogenous controls, our main specification excludes them. We also show the robustness of our estimates to the inclusion of Female x State x Year FE.

4.2.2 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.[29] 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.[30] 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 OLS regression:

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

The outcome variable, $Birth_{it}$, equals one if the mother gives birth in year t and is zero otherwise. $Post^1_t$, $Post^2_t$, and $FirstbornGirl_j$ 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 ($\nu_s FirstbornGirl_j$), and state-specific year fixed effects (θ_{st}).[31]

Additionally, we estimate the effects of ultrasound availability on the "stock" of children a

[28]Our results do not change if we instead use robust standard errors without clustering.

[29]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.

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

[31]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.

woman has at the time of the survey.[32] Specifically, we estimate the following specification for woman j in state s who has N_{jt} children in the year of survey, t:

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

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 $FirstbornGirl_j$ 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 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 ($\nu_s FirstbornGirl_j$), 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.[33]

5 Main Results: 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, left to right. Panel A presents estimates of specification (1). The coefficient on *Firstborn girl * Female* confirms the pattern shown in the summary statistics in Table 1: 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.[34]

The triple-interaction coefficients, *Firstborn girl * Female * Post1* and *Firstborn girl * Female*

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

[33] 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.

[34] 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.

* *Post2*, imply that this baseline EFM gap between firstborn-girl and firstborn-boy families was significantly reduced once ultrasound technology became available, narrowing by 60 percent. We can readily see this change in Panel B of Table 3 by replacing 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., $Post$.

The coefficients of interest are robust to the inclusion of interactions between covariates X_{ijt} and the $Female$ dummy in column (4), which allows the link between covariates and under-5 mortality risk to vary across gender. Moreover, our findings are robust to the inclusion of mother fixed effects in column (5), where the coefficients of interest are identified from two sources of variation—male vs female children within mother, and mothers who give birth pre vs post ultrasound, whether to same-sex or opposite-sex children. In Table A.21, we also show the robustness of our estimates to the inclusion of Female x State x Year FE.

The coefficients on *Firstborn girl * Post* are positive and significant and the coefficients on *Female* Post* are insignificant in Table 3. Since our story rests upon a negative and significant coefficient on the triple interaction term, *Firstborn girl * Post * Female*, these positive coefficients only underline the divergence in outcome trends post-ultrasound that we emphasize. For example, they show that within the group *Firstborn girl * Post*, mortality declined for female children relative to male children.

Just to document the underlying variation and make explicit the differences in the behaviors of families with firstborn boys vs girls, Table A.6 unpacks the triple difference. It presents a double-difference version of specification (1) for the samples of firstborn-girl and firstborn-boy families. The pattern is strikingly clear: there was no EFM in firstborn-boy families at baseline (i.e., pre-ultrasound), and no significant change post-ultrasound. In sharp contrast, firstborn-girl families had significantly positive EFM pre-ultrasound which declined significantly post-ultrasound.[35]

5.1 Extensions and Checks

Other mortality outcomes. Table A.7 shows results from regressions using death before age 1, 2, 3, and 4 as additional outcomes. The triple-interaction coefficients are always negative and increase in magnitude as we move from infant to under-5 mortality. We do not find any significant impact on gender differences in neonatal mortality (death in the first month of life), although the triple-interaction coefficients are consistently negative (see Table A.8). 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. Notice that any decline in neonatal EFM

[35]The impact on boys is captured by the coefficient on the terms $Post1$ and $Post2$ which show that, with increasing availability of ultrasound, boy mortality fell more rapidly in families with a firstborn-boy. However, recall that families with firstborn girls are the families where girls were more likely than boys to die after birth. This makes them "treatable" in the context of our paper. After the treatment, the pre-treatment gender gap narrows—girl mortality within these families falls more rapidly than boy mortality. But in the control group of families with firstborn boys, the reverse is the case.

(although insignificant) would bias 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, 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. 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)).

Results by birth order. Estimates by birth order are in Table A.9. 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] so there remains a girl-boy differential in under-5 mortality at higher orders.

Equalization of inheritance rights by gender. We also test if our results are biased by state-level legislative changes during our sample period that equalized inheritance of ancestral property between men and women. Recent work shows that this led to an increase in sex-selection (Bhalotra et al. (2018a)). We find that our results hold even if we restrict the sample to states that did not experience changes in inheritance rights, i.e., when we exclude Kerala, Andhra Pradesh, Tamil Nadu, Maharashtra, and Karnataka. These results are available on request.

Gendered recall bias. As we combine fertility histories of mothers interviewed in three DHS waves, recall of birth year varies within the sample but is, for some women, as long as 20 years. For the analysis in this paper, it is relevant to consider whether accuracy of recall of the year of birth or death of a child varies with the gender of the child. The firm gathering the NFHS data was alert to recall bias, and they report that the data were subject to several probes designed to address it. We, nevertheless, check for the presence of recall bias—if women are more likely to have forgotten girl deaths and if this error is increasing in distance of birth year from survey year, then this could, in principle, bias our findings. On the other hand, if families with firstborn girls had stronger gender bias in recall in the pre-ultrasound period then our EFM estimates will be conservative.

We examine gendered recall bias 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

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

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.2 shows that the reported sex ratio at birth does not vary by survey round.

We also check if the fraction female among deceased children varies by recall period for deaths within one month of birth and for deaths within five years of birth for firstborn children, always restricting the sample to allow full exposure to the allowed window of death. Figures A.3 and A.4 suggest that there is no significant gendered recall bias.[37]

We also previously showed in Table A.3 that the sex ratio of first births did not change over time as a function of maternal characteristics, including characteristics like education and religion that have been associated with son preference. If groups with high son preference are more likely to exhibit *gendered* recall of deceased children, this result eases concern that recall bias drives our conclusions.

Selection vs. changes in son preference. 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 already had a daughter at first birth). We have shown that the average girl who survives to birth is more likely to survive to age five in the post-ultrasound period. 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 arise from sex-selective abortion allowing selection, resulting in girls being increasingly born to parents who want them. We do not need to invoke a change in son preference. This is similar to discussions in the wider literature on abortion, which highlight that average outcomes for children improve when abortion is allowed because low-SES women are more likely to avail of abortion (Grossman and Joyce (1990b), Levine and Zimmerman (1996), Gruber et al. (1999)).[38]

Note that the estimated specifications allow for flexible trends in son preference insofar as they includes gender-specific birth year fixed effects. Son preference in India is centuries old and, while availability of ultrasound facilitated the exercise of son preference, there is no reason that it would change preferences. We, nevertheless, investigate this further—by controlling for preferences in different ways, and by modeling trends in preferences and allowing these to vary by firstborn sex.

We include controls for women's self-reported preferences (ideal number of children and ideal fraction of sons of children) in specification (1) and we include interactions of the preference variables with the $Female$ and $Firstborn\ girl$ indicators. Even if stated preferences are often measured

[37]The reason we restrict the sample to firstborn children is that we know that the sex ratio of first birth has not exhibited a trend. For higher-order births, any differences in the sex ratio at birth for deceased children by recall period will not only capture gender-bias in recall of child death (which is what we want to measure), but also trends in actual EFM and trends in the sex ratio of birth.

[38]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.

with error, it is notable that the coefficients of interest are not discernibly changed with the inclusion of these controls (Table A.10). In column (2) of Table A.10, we control for a different (also imperfect) measure of preferences that reflects actual choices rather than stated preferences —the state-year gender school 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 fraction of sons * Female* is positive and that of *Girl-boy enrollment rate 6-11* $*$ *Female* is negative, implying that EFM is higher in families with a stronger preference for sons. The coefficient on *Ideal number of children * 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).

We also estimated equations in which son preference (measured as the mother's reported ideal fraction of sons) is the dependent variable with a view to testing if it evolved differently between families with firstborn girls and firstborn boys; see Table A.11. This is not straightforward because we do not have longitudinal data on son preference. Instead, we have son preference recorded at the time of the survey in three cross-sections. The survey dates for the NFHS are 1991-92, 1998-99, and 2005-06, which are all post-ultrasound, but they span 15 years, allowing us to investigate preference decline during this period. For each wave, we keep women who had their first birth within a calendar year from their survey date, so that their preferences at the survey date reflect their preferences around the birth of their first child. We find a trend decline in son preference but no significant difference in the trend decline between firstborn-girl and firstborn-boy families.

As discussed earlier, Table A.3 suggests that there was no change in the composition of firstborn-girl families post-ultrasound. To the extent that son preference is systematically associated with observables, this contributes to showing that preferences did not change post-ultrasound in firstborn-girl relative to firstborn-boy families.[39]

Overall, using different approaches, we find no evidence that our finding of post-ultrasound declines in EFM in families with firstborn girls was driven by changes in son preference specific to this group. Instead, the mechanism appears to be that parents use selective abortion to align the timing, number, and sex of births to their preferences. We substantiate this by analyzing four additional sets of outcomes—parental investments, son-biased fertility stopping, the gap between actual and desired fertility, and birth spacing.

6 Mechanisms

In the following sub-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,

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

by virtue of being more wanted, post-ultrasound daughters in firstborn-girl families receive higher parental investments than girls born pre-ultrasound, narrowing investment gaps in these families between sons and daughters, relative to children in firstborn-boy families. 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.[40] A corollary of this is that actual fertility should move closer to desired fertility and we are able to test this as, unusually, we have individual data on desired fertility. Finally, we argue that if it is the availability of sex-selective abortion rather than something else that is driving our findings, then we should see an increase in birth spacing in the post-ultrasound period among families with firstborn girls relative to other families. We also investigate heterogeneity in treatment impacts by different indicators of SES with a view to further elucidating mechanisms.

6.1 Postnatal Health Investments

We posited above that declining under-5 mortality among girls signals increased parental investments in them. We directly test for this in Table 4. The outcomes are the number of months a child is breastfed;[41] 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.[42] 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 significantly declined during the post-ultrasound period in firstborn-girl families, although in some cells we lose significance for breastfeeding coefficients. The coefficients for medical expenditure during sickness are also positive, but imprecise, and only significant in one specification. 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, 19 percent and 7 to 10 percent (details in Online Appendix C).

For reasons discussed in Section 3, we also conducted this analysis using NFHS data. 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 and are available in Table A.12.

Our strictest specifications identify gender gaps and do not directly deliver an estimate of

[40] 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.

[41] Breastfeeding results are based on the last two surviving births of a mother and are based on the sample of children above age two to take into account censoring issues.

[42] 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.

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 significantly extended only for girls. These results are consistent with smaller sibship sizes in firstborn-girl families.[43]

6.2 Son-Biased Fertility Stopping

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 using an OLS specification. The coefficient of *Firstborn Girl* is positive and significant in both tables confirming that, pre-ultrasound, women whose first child was a girl were 3 p.p. more likely to give birth in a given year and had 0.189 more births than women with a firstborn son. Our estimates show that these differentials were significantly reduced once ultrasound technology became available. Estimates conditional upon mother fixed effects in column (6) of Table 5 are also similar. Column 3 in Table 6 implies that the pre-ultrasound gap in the number of births declined by 0.081 to 0.106, or by 40 to 50 percent. 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)). Controlling for mother's fertility preference and son preference, and their interactions with $Firstborn\ girl$ does not alter these effects in any way.[44] Table 7 shows that the relative fertility decline in firstborn-girl families is driven by a shift from having three or more children toward having two or fewer children.

Note that our fertility and EFM results remain similar if we restrict the NFHS sample to rural households from the states included in the REDS sample. Moreover, our results in Table 6 are robust to restricting the sample to women that have likely completed childbearing (i.e., women of age ≥ 33 or ≥ 37 at the time of survey), easing concerns about censoring, which may be more severe among firstborn-girl relative to firstborn-boy households in the post-period, driving our results. As a robustness check, we also estimate the effect on the hazard of birth using a Cox proportional hazard model. Table A.13 reports the hazard ratios from this exercise; our results remain the same.

6.3 Actual vs Desired Fertility

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 large reductions in undesired fertility, which the first row shows was 0.115 births in firstborn-girl relative to firstborn-boy families. The coefficients of interest (those on the triple-interaction terms) are similar for actual fertility and for excess fertility (actual minus desired fertility). This again

[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]Using a logit specification in Table 5 yields very similar results.

confirms that the decline in actual fertility in firstborn-girl families that we document is not driven by a decline in desired fertility (i.e., preferences).

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 steeply with son preference than actual fertility.[45] 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.

6.4 Birth Spacing

We also study birth spacing between the first two births as an outcome. We focus on the sample of women who have two children to avoid selection in terms of fertility preference as most couples in India have at least two children. We expect two mechanisms to be at play here: a behavioral mechanism that distinguishes firstborn-girl from firstborn-boy families, and a mechanical effect which is strictly active only in the post-ultrasound period. The first is that firstborn-girl mothers choose shorter birth spacing than firstborn-boy mothers, driven by their desire to conceive a son—one may think of this as an intensive margin behavior corresponding to their fertility-continuation behavior (Bhalotra and van Soest (2008), Jayachandran and Kuziemko (2011), Milazzo (2018)). The second is the mechanical effect associated with abortions being unobserved in the fertility history data. The observed interval between births will tend to be longer when an abortion takes place since there is a conception but then no birth. This implies that firstborn-girl households will tend to have longer birth spacing post- relative to pre-ultrasound since (sex-selective) abortions are predominantly in these households. In sum, other things equal, we expect birth spacing to be shorter in firstborn-girl than in firstborn-boy families during the pre-ultrasound period (due to the first mechanism), and this difference should narrow post-ultrasound (due to the second mechanism). This is exactly what we find.

Table A.14 compares birth spacing between the first two births for firstborn-boy and firstborn-girl families before and after ultrasound access. In the pre-ultrasound era, firstborn-girl mothers gave birth to their second child about one month sooner than firstborn-boy mothers, in line with the literature. This gap significantly narrows in the post-ultrasound period, consistent with greater use of sex-selective abortions by firstborn-girl mothers.

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

6.5 Heterogeneity

Although stated son preference is weaker among urban, literate, and wealthy women, they exhibit higher rates of prenatal sex selection (see Table 9 in Bhalotra and Cochrane (2010)). This pattern is consistent with their reporting lower desired fertility, 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 for different women, or that other factors (fecundity or causes of mortality) play a role too. We investigate heterogeneity in treatment impacts with a view to further elucidate mechanisms.

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),[46] household caste (scheduled caste (SC) versus other),[47] rural versus urban residence, and religion. Tables 8 and 9 respectively present estimates for gender gaps in sibling size and under-5 EFM.[48] 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). We also show tests in the tables of the significance of the difference between the coefficients in each group-pair.

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.[49] The decline in son-biased fertility stopping is significantly greater in high-SES households that had larger pre-ultrasound gender gaps. Thus is consistent with greater use of sex-selective abortions by high-SES families (Jha et al. (2006), Bhalotra and Cochrane (2010)).

EFM decline by SES. In the pre-ultrasound era, excess girl mortality among firstborn-girl families was greater in low-SES groups (illiterate, poor, unemployed, rural) with the exception that there was more EFM in upper-caste (non-SC) families, consistent with the stronger association of sons with status in these families (Srinivas (1962)). In general, we find no significant differences in

[46]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. So results for education may not line up with results for the other markers of SES, such as, wealth and caste.

[47]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.

[48]Table A.15 presents the heterogeneity results for effects on the number of children.

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

post-ultrasound declines in EFM across SES groups. The magnitudes of the absolute declines are larger in low-SES households where pre-ultrasound rates were higher, but these differences are not statistically significant, and the proportional declines are similar. This is not surprising because even if high-SES households were more affected by the introduction of prenatal sex detection because they showed a greater inclination to substitute sex-selective abortion for fertility-continuation, low-SES households are more "treatable" in any attempt to reduce mortality because it is easier to bring mortality down from initially higher levels.[50] A smaller change in investment can have larger survival impacts in low-SES households as they are more vulnerable to other causes of child mortality, such as infection.

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)).

Heterogeneity by religion. We examine heterogeneity by religion across the two largest religious groups in India, Hindus and Muslims, because they are known to differ in attitudes to abortion. Previous work shows that Muslims, who are more averse to abortion due to religious reasons, are more likely to attain their desired number of sons through son-biased stopping rules, whereas Hindus are more likely to practice sex-selective abortions to achieve the same end (Almond and Edlund (2008), Bhalotra and Cochrane (2010), Bhalotra et al. (2018b)). In line with this, Tables A.17, A.18, and A.19 show that our results on fertility, postnatal health investments, and mortality are driven by Hindus rather than Muslims. In the pre-ultrasound period, Muslim women with a firstborn girl had significantly more children and had significant under-5 EFM relative to those with a firstborn boy as displayed by the coefficients of *Firstgirl * Female* in Table A.17 and in Panel B of Table A.19. In the post-ultrasound era, when some Hindus switched from fertility continuation or postnatal neglect to sex-selective abortion, Muslims continued, as in the pre-ultrasound

[50]The one case where group differences in EFM decline are significantly different is in rural relative to urban households. Table 9 shows that absolute declines are significantly larger in rural families but, in fact, the proportional declines are smaller. This is in line with greater prenatal sex-selection in urban areas offsetting the role of higher baseline mortality in rural areas.

[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 in Table A.16.

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

era, to practice gender-biased fertility stopping.

Heterogeneity by region. We also examine if the effects differ across states that have historically been documented to exhibit varying degrees of son preference. We split the states in our sample into two groups consisting of states with above and below median sex ratios at birth to capture the variation in son preference. Consistent with expectation, we find that both EFM and fertility decline in the post-ultrasound period are stronger in states where son preference is more deeply rooted, as reflected in their above-median sex ratios at birth. See Tables A.20 and A.21.

7 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 Online 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 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, China.[53]

8 Discussion: Relation to Recent Work

A recent study by Hu and Schlosser (2015) (henceforth HS) addresses a similar question using similar data sources but finds no drop in EFM following availability of ultrasound (even though they do, like us, find evidence of increased parental investment in girls in the post-ultrasound period). In this section, we elucidate the value added of this study and, importantly, explain the source of the difference in our findings. Whether or not girl mortality has declined in the post-ultrasound

[53] As Lin et al. (2014) point out, their estimates are likely to understate the effect relative to places like India where son preference is stronger and the health care environment is poorer, making child mortality more sensitive to parental inputs. Note that, since mortality is an extreme case indicator, only parents that changed investments in children beyond a certain threshold (that averts child death) are, by definition, counted as "switchers" in both papers, and this threshold will be higher in Taiwan, China.

era is a question of considerable importance (see next section). There is a growing awareness in the field of economics of the importance of replication (Kane (1984), Card and Krueger (1995), Hammermesh (2007), Anderson and Kichkha (2017)), and that this may entail changes in design (Rosenblum (2001)).

Our investigations suggest that an important source of the difference is that we leverage the systematic pre- versus post-ultrasound difference between the behavior of families with firstborn girls versus firstborn boys, whereas HS pool these families and lose that identifying variation. In their specification, the independent variable is a 7-year moving average of the state-year average of the male-female sex ratio at birth, MFR_{st}. The summary statistics in Table 1 and estimates of a double-difference model displayed in Table A.6 show that pre-ultrasound EFM and the post-ultrasound decline in EFM were both concentrated in firstborn-girl families. The pre-ultrasound 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 is likely to veil relevant changes. The MFR_{st} variable in HS essentially pools a treated group with an untreated group that is roughly as large. It is useful to observe here that a number of other studies resemble ours in using quasi-experimental variation in the sex of the first born child and showing that it predicts son-preferring behaviors (e.g., Dahl and Moretti (2008), Rosenblum (2013), Milazzo (2018), Bhalotra et al. (2018a)).

To demonstrate the empirical relevance of this difference, we start with the HS specification[54] and make one change—we calculate the state-year averaged female-male sex ratio at birth separately for firstborn-girl (FMR_{st}^{FG}) and firstborn-boy families (FMR_{st}^{FB}), and then compute the difference between the two, calling it the "*excess* female-male sex ratio at birth" ($EFMR_{st} = FMR_{st}^{FG} - FMR_{st}^{FB}$). We then replace FMR_{st} in equation (4) with $EFMR_{st}$ as follows:

$$Y_{ist} = \alpha_{s0} + \alpha_{s1} female_i + \delta_{t0} + \delta_{t1} female_i + x_i^{'}\beta + \pi_0 EFMR_{st} + \pi_1 (EFMR_{st} * female_i) + \epsilon_{ist} \quad (5)$$

Table A.24 presents the results for specification (5). Columns (1)-(5) use the same sample as HS. Columns (6)-(8) expand the sample to include children who were born within 20 years of interview to make the estimates comparable to ours. This simple change to the HS specification uncovers a significant narrowing of the girl-boy difference in mortality, in line with the findings in our paper.[55]

[54]In our paper, unlike HS, we define *mortality* rather than survival as the outcome and we define the sex ratio at birth as the share of *females* at birth instead of the share of males at birth. To ensure comparability with our estimates, we rewrite HS' specification as follows:

$$Y_{ist} = \alpha_{s0} + \alpha_{s1} female_i + \delta_{t0} + \delta_{t1} female_i + x_i^{'}\beta + \pi_0 FMR_{st} + \pi_1 (FMR_{st} * female_i) + \epsilon_{ist} \quad (4)$$

where Y_{ist} is a mortality indicator for child i.

[55]We find similar results when we merge our $EFMR_{st}$ variable with the HS data published on the EJ webpage, and use their program to estimate their specification except for replacing MFR_{st} with $EFMR_{st}$. This exercise eliminates all differences in sample construction and the estimation method, so the only remaining difference is that we replace MFR with EFMR.

Another important difference between our specification and that of HS is that we focus on second- and higher-order births while HS include children of all parities in their regression specification (equation (1) above). Since the sex ratio of first births was not impacted by ultrasound access (we demonstrated this earlier), this dilutes the relevant variation in MFR_{st}. The estimated impacts in Table A.24 become larger when we exclude first births from the sample.

Third, aggregate state-year variation in the sex ratio at birth (the key independent variable in HS) is not a clean measure of availability, as it combines plausibly exogenous availability of the new technology (supply) with state-year varying uptake (demand). In other words, the sex ratio at birth is likely to be jointly determined with the outcomes.[56] Since we model EFM by firstborn sex (which (i) is quasi-random and (ii) varies at the mother level), we are able to, more effectively than HS, difference out a host of these omitted variables by controlling for Female x State x Year FE, Female x Year FE, Female x State FE, Female x Birth order FE, for instance. Table A.21 shows that our EFM results are robust to the inclusion of Female x State x Year FE—this is a check that HS cannot do given their identification strategy. We also introduce mother fixed effects in a variation of the specification, to account for selective uptake.

Fourth, we use quasi-experimental variation in the availability of ultrasound scanners determined by policy changes governing imports and industrial licensing in India. The MFR_{st} variable used by HS does not adequately capture structural breaks in the availability of sex-selection technology in India—we capture these by dividing our data into the pre-ultrasound period, the early diffusion period, and the late diffusion period through two indicators ($Post_t^1$ and $Post_t^2$). We demonstrate the empirical relevance of this by starting with our specification and replacing $Post_t^1$ and $Post_t^2$ in a simplified version of equation (1) with the average male-female sex ratio at birth, creating one average for each of the three periods of ultrasound availability, $\overline{MFR_{st}}$. Therefore, $\overline{MFR_{st}}$ takes three distinct values. We then estimate the following specification:

$$
\begin{aligned}
Y_i =& \alpha + \beta FirstbornGirl_j * Female_i * \overline{MFR_{st}} \\
&+ \gamma FirstbornGirl_j * Female_i + \omega_t FirstbornGirl_j + \sigma_t Female_i \\
&+ \mathbf{X}'_{\mathbf{ijt}}\tau + \delta_s Female_i + \nu_s FirstbornGirl_j + \psi_b Female_i + \xi_b FirstbornGirl_j \\
&+ \rho_{bt} + \eta_{bs} + \phi_{st} + \epsilon_i
\end{aligned}
\tag{6}
$$

The only difference between equations (1) and (6) is that we have replaced the two triple-interaction terms ($FirstbornGirl_j * Female_i * Post_t^1$ and $FirstbornGirl_j * Female_i * Post_t^2$) in (1) with one term, $FirstbornGirl_j * Female_i * \overline{MFR_{st}}$ in (6).

[56] 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.

Columns (1)-(3) in Panel B of Table A.25 present estimates of equation (6) for our sample (i.e., births that occurred within 20 years of interview). The triple-interaction term is now statistically insignificant. In columns (4)-(6) we include only births within 10 years of interview to make our sample comparable to that in HS, and we find similarly statistically insignificant results. For comparison, in Panel A we report the estimates from our main specification, i.e., equation (1).

Another distinction between our paper and HS is that we examine fertility more closely, looking at son-biased fertility stopping, the gap between actual and desired fertility, and birth spacing as outcomes. We find lower rates of son-biased fertility stopping in firstborn-girl families, a narrowing of the desired vs. actual fertility gap, and that the pre-ultrasound pattern of shorter birth spacing in firstborn-girl relative to firstborn-boy families is significantly moderated once ultrasound technology became available. These results are of substantive importance since children often receive lower investments in larger families (Becker and Lewis (1974), Mogstad and Wiswall (2016)); unwanted fertility imposes costs on women and children; and short birth spacing has been shown to be associated with maternal health deterioration (see discussion in Milazzo (2018)).

9 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,[57] 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,[58] 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)). Grosjean and Khattar (2018) show that male-biased sex ratios can lead to more conservative cultural attitudes, labor supply decisions, and occupational choices in the long-term. 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

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

[58] 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)).

account for an estimated 9 percent of all maternal deaths in India (Stillman et al. (2014)).[59] On the other hand, a shortage of women on the marriage market may increase their bargaining power and welfare.[60] It has also been argued that sex-selective abortions might be 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[61] and for every additional girl that survived after birth, three girls were aborted.

The objective of this study is neither to condone nor to reject the practice of sex-selective abortion. Our paper offers a deeper understanding of parental decision-making about fertility and child investments and presents a more complete picture of the effects of access to ultrasound technology in India. Lastly, we note that the findings of this study are based on data until 2005 and may or may not apply to more recent years.

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

[60] See Chiappori et al. (2002) and related papers for the large literature on household bargaining in developed countries. Stopnitzky (2017) 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.

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

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. 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.

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, R. G. and A. Kichkha (2017): "Replication, Meta-analysis, and Research Synthesis in Economics," *American Economic Review*, 107, 56–59.

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.

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*.

Becker, G. S. and H. G. Lewis (1974): "The Interaction between Quantity and Quality of Children," *Economics of the Family: Marriage, Children, and Human Capital*, 81–90.

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

Bhalotra, S., R. Brule, and S. Roy (2018a): "Women's Inheritance Rights Reform and the Preference for Sons in India," *Journal of Development Economics*, Forthcoming.

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

Bhalotra, S., I. Clots-Figueras, and L. Iyer (2018b): "Religion and Abortion: The Role of Politician Identity," *IZA Discussion Paper 11292*.

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.

Bitler, M. and M. Zavodny (2002): "Child Abuse and Abortion Availability," *American Economic Review Papers and Proceedings*, 92.

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.

Card, D. and A. Krueger (1995): "Time-Series Minimum-Wage Studies: A Meta-analysis," *American Economic Review, P&P*, 85, 238–243.

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.

Dahl, G. B. and E. Moretti (2008): "The Demand for Sons," *Review of Economic Studies*, 75, 1085–1120.

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.

Filmer, D., J. Friedman, and N. Schady (2009): "Development, Modernization, and Childbearing: The Role of Family Sex Composition," *World Bank Economic Review*, 23.

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.

Grosjean, P. and R. Khattar (2018): "It's Raining Men! Hallelujah? The Long-Run Consequences of Male-Biased Sex Ratios," *Review of Economic Studies*, Forthcoming.

Grossman, M. and J. Joyce, Theodore (1990a): "Pregnancy Wantedness and the Early Initiation of Prenatal Care," *Demography*, 27, 1–17.

——— (1990b): "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.

Hammermesh, D. S. (2007): "Replication in economics," *Canadian Journal of Economics*, 40.

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.

Jayachandran, S. and R. Pande (2017): "Why Are Indian Children So Short? The Role of Birth Order and Son Preference," *American Economic Review*, 107, 2600–2629.

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

Jha, P., R. Kumar, P. Vasa, N. Dhingra, Neeraj, D. Thiruchelvam, and R. Moineddin (2006): "Low Male-to-female Sex Ratio of Children Born in India: National Survey of 1.1 Million Households‘," *Lancet*, 367, 211–218.

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.

Kane, E. J. (1984): "Why journal editors should encourage the replication of applied econometric research," *Quarterly Journal of Business and Economics*, 23, 3–8.

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.

Levine, P. B., T.-A. B. and D. J. Zimmerman (1996): "The effect of Medicaid abortion funding restrictions on abortions, pregnancies and births," *Journal of Health Economics*, 15, 555–578.

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*.

——— (2018): "Why Are Adult Women Missing? Son Preference and Maternal Survival in India," *Journal of Development Economics*, 134, 467–484.

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

Mogstad, M. and M. Wiswall (2016): "Testing the Quantity-Quality Model of Fertility: Linearity, Marginal Effects, and Total Effects," *Quantitative Economics*, 7, 157–192.

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.

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.

Rosenblum, P. (2001): "Replicating Effects and Biases," *The American Statistician*, 55, 223–227.

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.

Stopnitzky, Y. (2017): "No toilet no bride? Intrahousehold bargaining in male-skewed marriage markets in India," *Journal of Development Economics*, 127, 269–282.

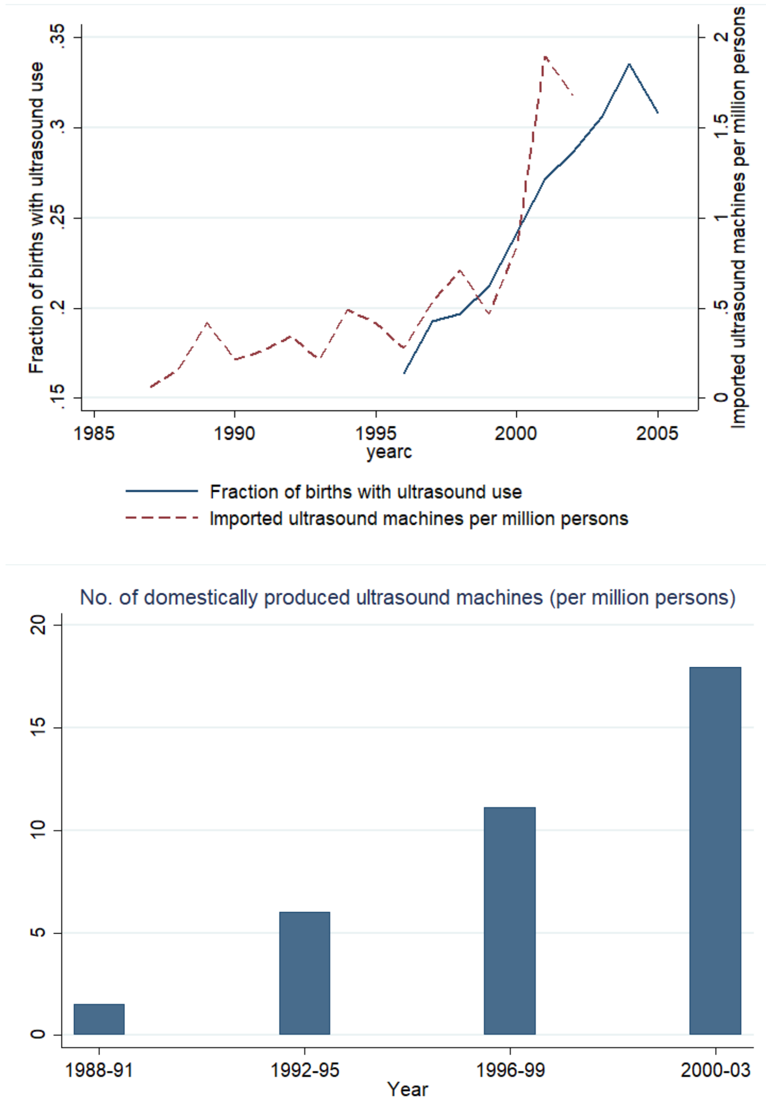
Trivers, R. L. and D. E. Willard (1973): "Natural selection of parental ability to vary the sex ratio of offspring," *Science*, 179, 90–92.

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.

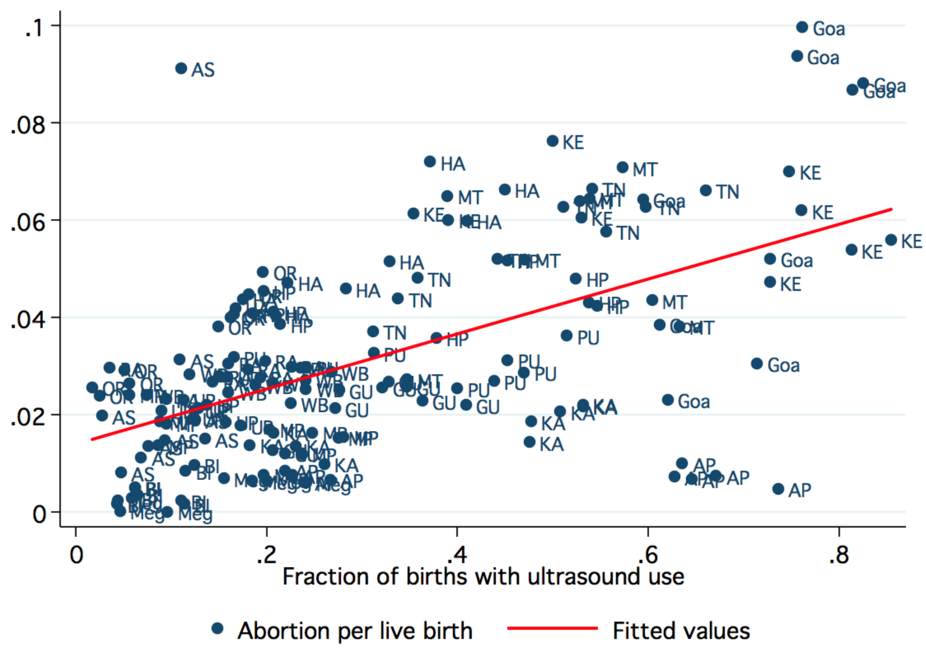
Figures

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 per million persons, 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 per million persons. Data source: George (2006).

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

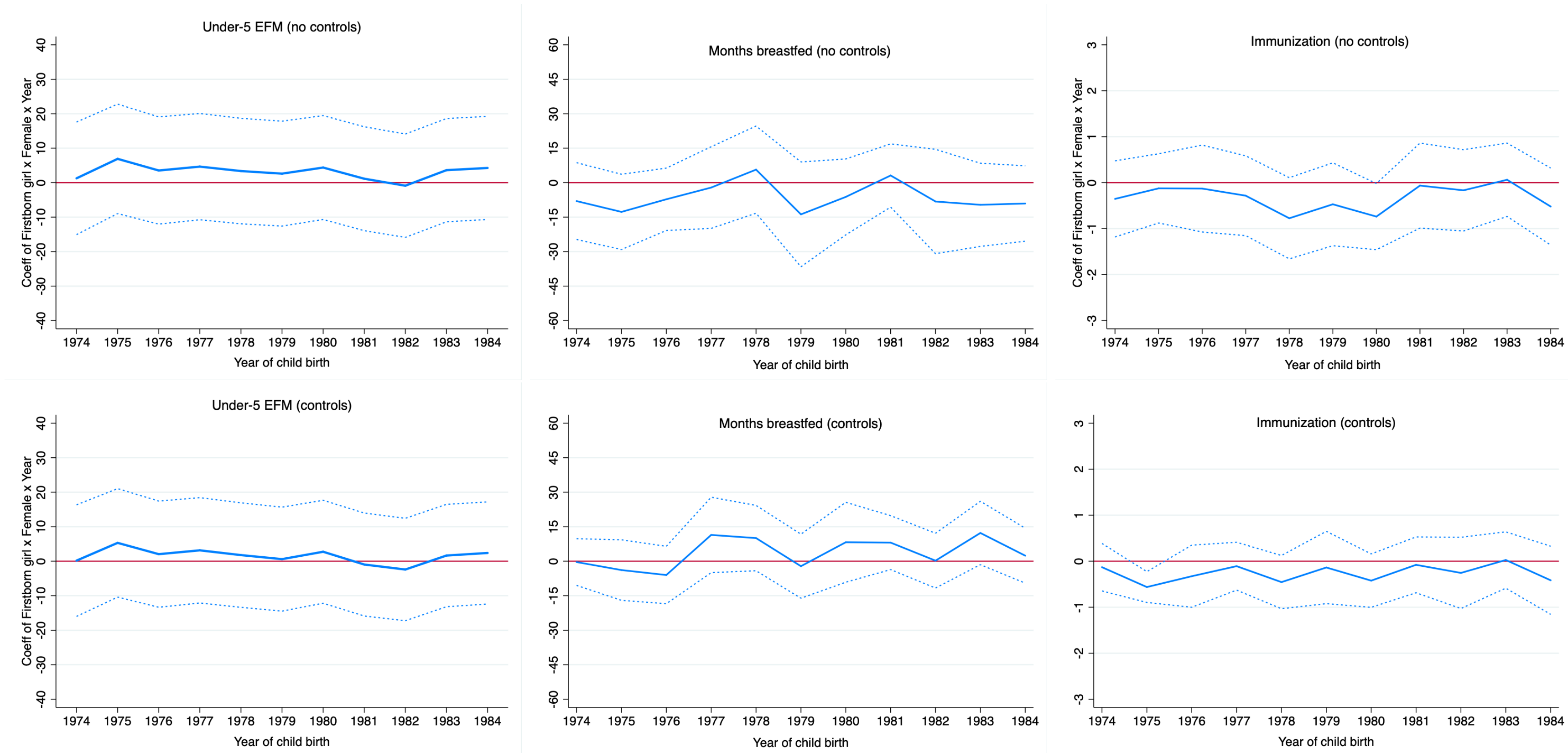


NOTES: This graph plots the state-year variation in the officially reported number of abortions and the fraction of births in a state-year for which the mother reports getting an ultrasound test at some point during the pregnancy (other details in Figure 1 notes). Data for the officially reported number 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. Data: NFHS.

Figure 4: Test of differential pre-trends in outcomes by firstborn sex



NOTES: This figure tests for parallel trends in under-5 EFM, months breastfed, and receiving at least one vaccine by firstborn sex during the pre-ultrasound period (1973-1984) for births of parity > 1. 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 the outcome variable on the full set of interactions between indicators for *First girl*, *Female*, and *Year*, without other controls in the top row and with fixed effects for birth order and state in the bottom row. The omitted year is 1973. Data: NFHS and REDS.

Tables

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

	Firstborn boy families			Firstborn 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	1.58
N	30,066	28,333		31,382	28,412	
Late diffusion period: 1995-2005	7.71	7.80	0.09	7.14	8.22	1.09
N	10,314	9,709		11,543	9,921	
B. Received at least one vaccine						
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. Months breastfed						
Pre-ultrasound: 1973-1984	19.05	18.63	-0.42	19.15	18.90	-0.25
N	1,537	1,124		619	401	
Early diffusion period: 1985-1994	18.32	18.35	0.03	18.50	18.10	-0.40
N	2,197	1,758		1,105	734	
Late diffusion period: 1995-1999	12.78	12.52	-0.26	13.71	12.95	-0.76
N	1,453	1,302		673	532	
D. Medicines/special food during illness last year (Rs.)						
Pre-ultrasound: 1973-1984	153.44	174.34	20.90	265.96	165.22	-100.74
N	645	399		396	245	
Early diffusion period: 1985-1994	162.15	149.74	-12.41	159.56	143.26	-16.30
N	1,816	1,651		1,407	1,250	
Late diffusion period: 1995-1999	191.71	162.80	-28.91	214.36	148.81	-65.55
N	769	733		683	553	
E. Total number of children						
Pre-ultrasound: 1973-1984		2.91			3.12	
N		14,095			10,690	
Early diffusion period: 1985-1994		2.13			2.22	
N		25,857			22,169	
Late diffusion period: 1995-2005		1.85			1.96	
N		23,480			22,372	

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. Panel E shows the total number of children per mother for the sample of mothers whose first and last birth took place within each of the three periods. Data: NFHS in Panels A and E. REDS in Panels B, C, D.

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

Outcome: Death before age 5	(1)	(2)	(3)	(4)	(5)
Panel A:					
*Firstborn girl * Female*	2.957***	2.821***	2.768***	2.785***	2.563***
	(0.607)	(0.598)	(0.588)	(0.593)	(0.588)
*Firstborn girl * Female * Post1*	-1.646**	-1.527*	-1.525*	-1.536*	-1.602
	(0.773)	(0.776)	(0.768)	(0.774)	(0.974)
*Firstborn girl * Female * Post2*	-1.998**	-1.953**	-1.996**	-2.022**	-1.841**
	(0.744)	(0.749)	(0.744)	(0.745)	(0.734)
*Firstborn girl * Post1*	1.422**	1.455**			
	(0.576)	(0.538)			
*Firstborn girl * Post2*	1.785***	1.793***			
	(0.636)	(0.623)			
*Female * Post1*	0.303	0.225			
	(0.478)	(0.489)			
*Female * Post2*	0.141	0.0209			
	(0.578)	(0.585)			
Panel B:					
*Firstborn girl * Female*	2.957***	2.821***	2.775***	2.784***	2.564***
	(0.607)	(0.598)	(0.586)	(0.593)	(0.587)
*Firstborn girl * Female * Post*	-1.741**	-1.642**	-1.647**	-1.662**	-1.659*
	(0.711)	(0.710)	(0.695)	(0.702)	(0.828)
*Firstborn girl * Post*	1.519**	1.546***			
	(0.560)	(0.526)			
*Female * Post*	0.261	0.173			
	(0.445)	(0.455)			
N	227,129	227,129	227,129	227,129	227,129
Baseline mean	0.124	0.124	0.124	0.124	0.124
X_{ijt}		x	x	x	x
Firstborn 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 Birth year FE			x	x	x
Birth order x State FE			x	x	x
State x Birth year FE			x	x	x
Firstborn girl x State FE			x	x	x
Firstborn girl x Birth order FE			x	x	x
Female x X_{ijt}				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 OLS 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)-(4). The vector X_{ijt} comprises mother's age at birth and, except in column (5), 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. Data: NFHS. *** 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*	-5.116**	-4.585**	-3.625**	-0.130**	-0.093**	-0.069	-28.473	-45.830	-95.486**
	[1.861]	[1.820]	[1.536]	[0.051]	[0.039]	[0.048]	[42.037]	[37.982]	[43.982]
*First girl * Female * Post1*	4.499**	3.957*	1.859	0.113**	0.099*	0.076*	17.708	35.897	73.112
	[2.011]	[2.128]	[2.207]	[0.049]	[0.048]	[0.042]	[64.694]	[48.992]	[58.856]
*First girl * Female * Post2*	5.672*	4.277	2.331	0.094*	0.073*	0.118**	67.488	103.445*	154.470**
	[2.712]	[2.746]	[2.396]	[0.051]	[0.035]	[0.044]	[44.108]	[52.554]	[69.308]
*Female * Post1*	-2.374*	-1.201		-0.005	0.001		9.401	12.427	
	[1.321]	[0.999]		[0.056]	[0.045]		[21.840]	[23.550]	
*Female * Post2*	-1.730	-0.798		-0.029	-0.015		-38.946	-58.308	
	[1.733]	[1.708]		[0.038]	[0.030]		[31.540]	[36.770]	
*First girl * Post1*	-1.142	-0.634		0.052*	0.050		-30.552	-21.272	
	[1.224]	[1.425]		[0.026]	[0.035]		[52.685]	[45.289]	
*First girl * Post2*	0.061	0.506		0.039	0.059		-54.144	-58.160	
	[1.871]	[1.746]		[0.039]	[0.049]		[59.750]	[53.773]	
N		10,965			20,562			10,293	
Baseline mean		18.931			0.704			186.426	
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 OLS 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 and restrict the sample to children above age 2. 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%.

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

	(1)	(2)	(3)	(4)	(5)	(6)
Firstborn girl	0.0352***	0.0357***	0.0301***	0.0314***	-0.00967*	0.0455***
	(0.00226)	(0.00215)	(0.00509)	(0.00512)	(0.00495)	(0.00997)
*Firstborn girl * Post1*	-0.0197***	-0.0213***	-0.0217***	-0.0231***	-0.0224***	-0.0213***
	(0.00157)	(0.00165)	(0.00154)	(0.00180)	(0.00183)	(0.00310)
*Firstborn girl * Post2*	-0.0372***	-0.0400***	-0.0400***	-0.0418***	-0.0401***	-0.0427***
	(0.00249)	(0.00248)	(0.00199)	(0.00225)	(0.00222)	(0.00361)
Ideal fraction of sons				0.0126***	-0.0143***	
				(0.00349)	(0.00332)	
Ideal no. of children				0.0273***	0.0275***	
				(0.00152)	(0.00162)	
*Firstborn girl * Ideal fraction of sons*					0.0766***	
					(0.00480)	
*Firstborn girl * Ideal no. of children*					-0.000471	
					(0.000968)	
N	2,455,633	2,455,633	2,455,633	2,276,264	2,276,264	2,276,264
X_i	x	x	x	x	x	
Year FE	x	x	x	x	x	x
State FE	x	x	x	x	x	
Age FE	x	x	x	x	x	x
Parity FE	x	x	x	x	x	x
State x Year FE	x	x	x	x	x	x
Years since last birth FE		x	x	x	x	x
Firstborn girl x State FE			x	x	x	
Mother FE			x	x	x	x

NOTES: Coefficients from specification (2) estimated using OLS regression on the 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. Data: NFHS. *** 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)	(4)	(5)
Firstborn girl	0.155***	0.141***	0.189***	0.117***	0.115**
	(0.012)	(0.015)	(0.031)	(0.018)	(0.043)
*Firstborn girl * Post1*	-0.088***	-0.079***	-0.081***	-0.085***	-0.085***
	(0.016)	(0.019)	(0.019)	(0.024)	(0.024)
*Firstborn girl * Post2*	-0.112***	-0.100***	-0.106***	-0.093***	-0.093***
	(0.018)	(0.023)	(0.024)	(0.025)	(0.026)
Ideal no. of children		0.315***	0.314***		
		(0.018)	(0.023)		
Ideal fraction of sons		0.052***	0.073***	-0.345***	-0.346***
		(0.018)	(0.025)	(0.024)	(0.032)
*Firstborn girl * Ideal no. of children*			0.002		
			(0.016)		
*Firstborn girl * Ideal fraction of sons*			-0.044		0.002
			(0.028)		(0.028)
N	118,663	88,475	88,475	88,475	88,475
Baseline mean	3.001	3.001	3.001	0.451	0.451

NOTES: Estimates of specification (3) estimated using OLS regression. The dependent variable in columns (1)-(3) is the number of births at the time of interview and in columns (4)-(5) 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. Data: NFHS. *** 1%, ** 5%, * 10%.

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

	Number of children				
	≥ 2	≥ 3	≥ 4	≥ 5	≥ 6
	(1)	(2)	(3)	(4)	(5)
Firstborn girl	-0.012**	0.054***	0.069***	0.034***	0.009***
	(0.005)	(0.005)	(0.006)	(0.004)	(0.002)
*Firstborn girl * Post1*	0.010	-0.011	-0.042***	-0.034***	-0.009**
	(0.007)	(0.007)	(0.007)	(0.005)	(0.003)
*Firstborn girl * Post2*	0.017**	-0.023***	-0.058***	-0.036***	-0.009***
	(0.007)	(0.008)	(0.009)	(0.005)	(0.003)
N	118,663	118,663	118,663	118,663	118,663
Baseline mean	0.886	0.615	0.315	0.127	0.042

NOTES: This table presents estimates from specification (3) using indicators for the mother having various number of children at the time of survey in columns (1)-(5). Data: NFHS. *** 1%, ** 5%, * 10%.

Table 8: Probability of birth: 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)	(1)-(2)	(3)	(4)	(3)-(4)	(5)	(6)	(5)-(6)
Firstborn girl	0.0159***	0.0511***	-0.0352***	0.0215***	0.0533***	-0.0318***	0.0333***	0.0278***	0.00549
	(0.00418)	(0.00431)	(0.00319)	(0.00642)	(0.00410)	(0.00603)	(0.00606)	(0.00460)	(0.00352)
*Firstborn girl * Post1*	-0.0150***	-0.0277***	0.0128***	-0.0201***	-0.0223***	0.00216	-0.0262***	-0.0157***	-0.0105***
	(0.00182)	(0.00263)	(0.00303)	(0.00114)	(0.00345)	(0.00359)	(0.00180)	(0.00248)	(0.00325)
*Firstborn girl * Post2*	-0.0268***	-0.0506***	0.0237***	-0.0332***	-0.0451***	0.0119**	-0.0454***	-0.0326***	-0.0128***
	(0.00251)	(0.00295)	(0.00382)	(0.00302)	(0.00351)	(0.00541)	(0.00233)	(0.00287)	(0.00351)
N	1,351,795	1,103,838		972,072	600,212		1,494,683	959,110	
Baseline mean	0.289	0.279		0.288	0.263		0.287	0.283	
	Caste			**Rurality**					
	SC	General/OBC/ST		Rural	Urban				
	(7)	(8)	(7)-(8)	(9)	(10)	(9)-(10)			
Firstborn girl	0.0204**	0.0324***	-0.0119***	0.0216***	0.0491***	-0.0275***			
	(0.00762)	(0.00475)	(0.00375)	(0.00571)	(0.00450)	(0.00456)			
*Firstborn girl * Post1*	-0.0136***	-0.0232***	0.00962**	-0.0198***	-0.0244***	0.00464			
	(0.00417)	(0.00153)	(0.00402)	(0.00174)	(0.00266)	(0.00303)			
*Firstborn girl * Post2*	-0.0336***	-0.0415***	0.00784	-0.0356***	-0.0472***	0.0116***			
	(0.00632)	(0.00185)	(0.00603)	(0.00273)	(0.00277)	(0.00397)			
N	374,747	2,080,886		1,598,757	856,876				
Baseline mean	0.295	0.284		0.289	0.278				

NOTES: Estimates of specification (2) in the text. The dependent variable is an indicator for birth to a given mother in a given year. Each column within a panel is a separate OLS regression. 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. The third column in every panel shows a test of the significance of the difference between the coefficients for the sub-groups shown in the first two columns. Data: NFHS. *** 1%, ** 5%, * 10%.

Table 9: Under-5 mortality: Heterogeneity by socioeconomic status

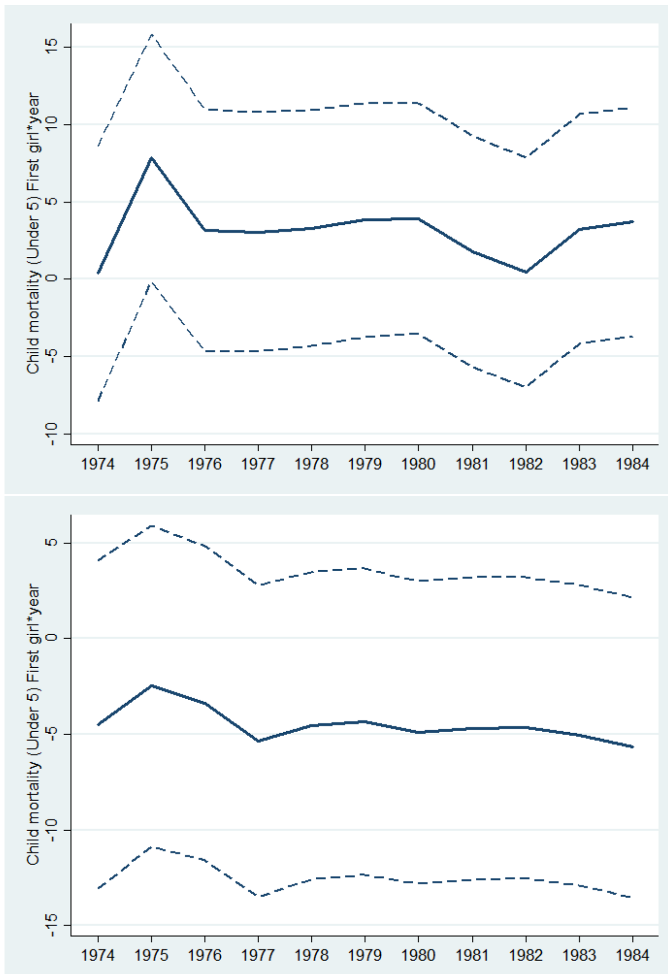
	Mother's Education			**Wealth**			**Mother's Employment**		
	Illiterate	Literate		Bottom 40%	Top 20%		Paid employment = 0	Paid employment = 1	
	(1)	(2)	(1)-(2)	(3)	(4)	(3)-(4)	(5)	(6)	(5)-(6)
*Firstborn girl * Female*	3.455***	1.679**	1.777	3.279***	1.022	-2.258	2.993***	2.392***	-0.6011
	(0.830)	(0.774)	(1.175)	(1.143)	(0.752)	(1.449)	(0.753)	(0.725)	(0.908)
*Firstborn girl * Female * Post*1	-1.986*	-0.789	-1.198	-1.622	0.386	2.008	-1.824*	-1.039	0.7851
	(1.077)	(0.992)	(1.435)	(1.542)	(1.090)	(2.047)	(0.971)	(1.026)	(1.279)
*Firstborn girl * Female * Post*2	-2.320*	-1.232	-1.089	-2.177	0.500	2.676	-2.463**	-1.360	1.103
	(1.244)	(1.031)	(1.845)	(1.723)	(1.506)	(2.981)	(1.113)	(0.870)	(1.396)
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				
	(7)	(8)	(7)-(8)	(9)	(10)	(9)-(10)			
*Firstborn girl * Female*	1.707	2.816***	-1.109	3.718***	0.315	3.403***			
	(1.547)	(0.555)	(1.439)	(0.776)	(0.946)	(1.267)			
*Firstborn girl * Female * Post*1	-0.863	-1.525*	0.6613	-2.522**	1.086	-3.608**			
	(2.033)	(0.774)	(1.983)	(1.055)	(1.035)	(1.522)			
*Firstborn girl * Female* Post*2	-0.683	-2.100**	1.417	-3.179***	0.833	-4.012**			
	(2.483)	(0.787)	(2.619)	(1.072)	(1.313)	(1.936)			
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 OLS 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. The third column in every panel shows a test of the significance of the difference between the coefficients for the sub-groups shown in the first two columns. Data: NFHS. *** 1%, ** 5%, * 10%.

ONLINE APPENDICES

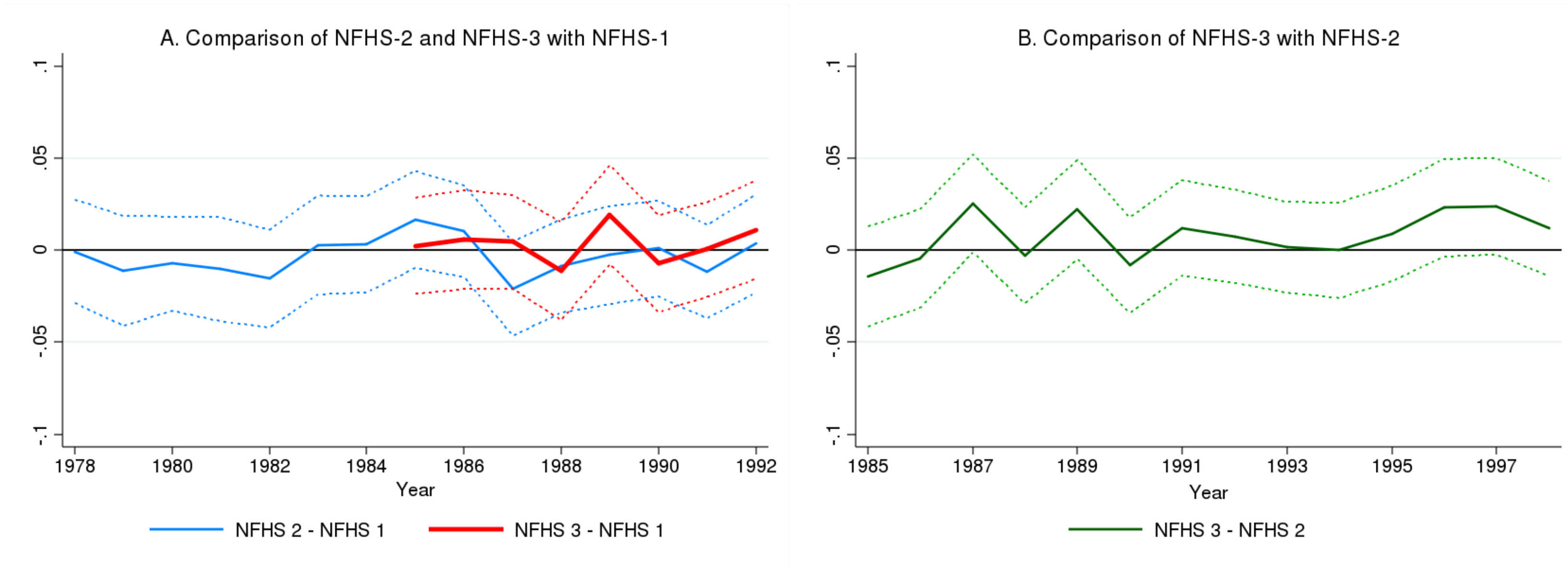
A Additional Figures and Tables

Figure A.1: Test of differential pre-trends in under-5 mortality by firstborn sex separately for boys and girls



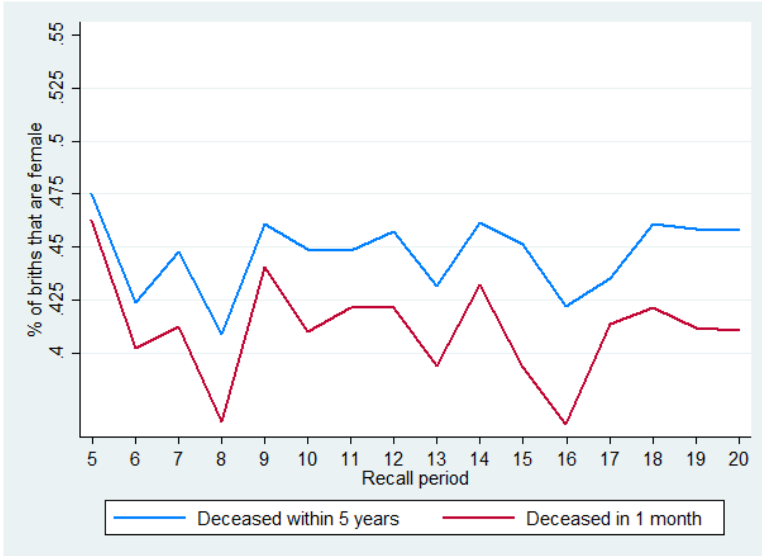
NOTES: This figure tests for parallel trends in under-5 mortality by firstborn sex during the pre-ultrasound period (1973-1984) for male and female births of parity > 1. The graphs plot the coefficients of the *First girl* x *Year* indicator variables (and the 95% confidence intervals) from the regression of under-5 mortality on the full set of interactions between indicators for *First girl* and *Year*, without other controls, for girls in the top figure and for boys in the bottom figure. The omitted year is 1973. Data: NFHS.

Figure A.2: Testing for recall bias in the reported sex ratio at first birth



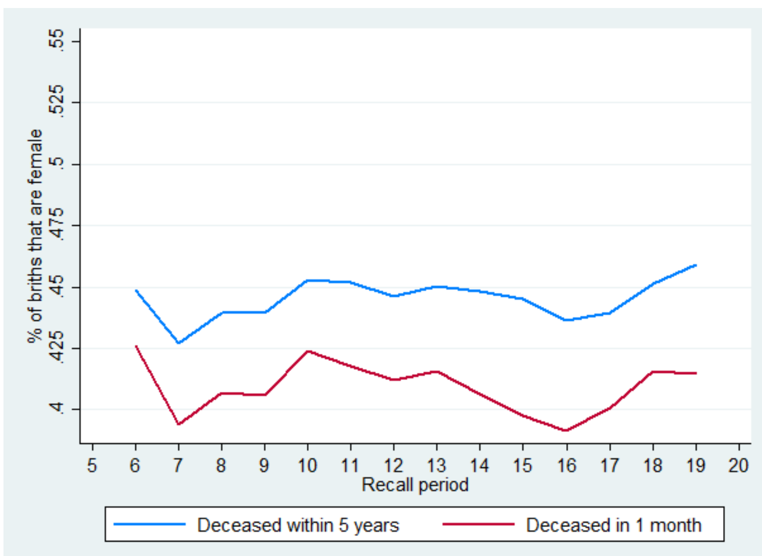
NOTES: This graph checks if there is any systematic recall bias in reports of first child's gender by comparing how the probability that a first birth is reported as female differs across rounds for the same year of birth. In panel A, the thinner solid blue (thicker solid red) line shows how the reported 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.3: Sex ratio of deceased firstborn children by recall period



NOTES: The figure plots the proportion female among firstborn children deceased within 5 years (or 1 month) of birth by recall period defined as the number of years between birth and interview. Data: NFHS.

Figure A.4: Sex ratio of deceased firstborn children by recall period, 3-year moving average



NOTES: The figure plots the 3-year moving average of the proportion female among firstborn children deceased within 5 years (or 1 month) of birth by recall period defined as the number of years between birth and interview. Data: NFHS.

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). Data: NFHS.

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 of first birth and changes in characteristics of FG vs FB families over time

Dep Var: First child is female	(1)	(2)
Post	0.00324	0.00993
	(0.00435)	(0.0147)
SC**Post*		0.00749
		(0.00802)
Educated (Mother)**Post*		-0.00747
		(0.00914)
Educated (Father)**Post*		-0.00525
		(0.00727)
Muslim**Post*		-0.00946
		(0.0175)
Hindu**Post*		-0.00278
		(0.0136)
Rural**Post*		-0.000712
		(0.00710)
High wealth (Top 60%)**Post*		0.00417
		(0.00603)
N	173,293	173,293
State FE	x	x
Households controls	x	x
Linear time trend	x	x
HH controls*Post		x
F-stat for joint significance of HH controls*Post		F(7,24) = 1.98

NOTES: This table tests if ultrasound access affected the sex ratio at first birth. The outcome variable is an indicator for first 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 that included in the main specification for health outcomes). In column (2), we control for a set of interactions between household controls with *Post* (indicating $t \geq 1985$). The table reports the F-statistic from a test for the joint-significance of the set of interaction terms in column (2). Standard errors in parentheses are clustered by state. Data: NFHS. *** 1%, ** 5%, * 10%.

Table A.4: Balance Tests: Sample means by firstborn sex, REDS

	1973-1984		1985-1994		1995-1999		All years	1973-1984	1985-1994	1995-1999
	FB	FG	FB	FG	FB	FG	FB-FG	FB-FG	FB-FG	FB-FG
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
SC	0.13	0.15	0.14	0.14	0.14	0.15	-0.008	-0.02	0.0001	-0.004
ST	0.07	0.07	0.07	0.07	0.08	0.07	0.003	0.0009	0.002	0.009
OBC	0.35	0.35	0.34	0.37	0.33	0.35	-0.02	-0.004	-0.03*	-0.02
Hindu	0.87	0.90	0.87	0.90	0.90	0.89	-0.02***	-0.03**	-0.03***	0.01
Muslim	0.08	0.06	0.08	0.06	0.05	0.06	0.02**	0.02*	0.02	-0.02*
Sikh	0.03	0.03	0.04	0.03	0.04	0.02	0.008*	0.004	0.008	0.02
Christian	0.01	0.01	0.01	0.01	0.01	0.02	-0.005	0.007	-0.0003	-0.01**
Mother's age at birth	20.59	20.61	20.74	20.84	21.40	21.63	-0.13	-0.03	-0.10	-0.23
Mother's years of education	2.74	3.06	3.84	3.73	5.37	5.52	-0.18*	-0.32*	0.11	-0.15
Father's years of education	5.54	6.11	6.54	6.44	7.79	8.01	-0.27**	-0.57***	0.10	-0.22
N (1st births)	1,374	899	1,599	1,216	758	646	6,492	2,273	2,815	1,404

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.5: Actual and ideal fertility and fraction of sons; pooling all families

	(1) Actual	(2) Ideal	(3) Actual - Ideal
1. Pre-ultrasound: 1973-1984			
Fertility	2.98	2.54	0.44
Fraction of Sons	0.58	0.57	0.01
N		23,414	
2. Post-ultrasound: 1985-1994			
Fertility	2.17	2.47	-0.30
Fraction of Sons	0.55	0.56	-0.01
N		45,920	
3. Post-ultrasound: 1995-2005			
Fertility	1.90	2.31	-0.41
Fraction of Sons	0.53	0.55	-0.02
N		44,813	

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. Data: NFHS.

Table A.6: Excess female under-5 mortality: double-difference models by firstborn sex

	Outcome: Death before age 5					
	Firstborn girl family			Firstborn boy family		
	(1)	(2)	(3)	(4)	(5)	(6)
Female	2.929***	2.893***	2.822***	-0.0338	-0.0609	0.00292
	(0.694)	(0.691)	(0.680)	(0.618)	(0.603)	(0.614)
Female × *Post1*	-1.362**	-1.343**	-1.305**	0.282	0.307	0.225
	(0.510)	(0.513)	(0.498)	(0.493)	(0.480)	(0.489)
Female × *Post2*	-1.882***	-1.853***	-1.928***	0.0976	0.148	0.0199
	(0.461)	(0.466)	(0.458)	(0.577)	(0.581)	(0.587)
Post1	-1.897***			-3.304***		
	(0.499)			(0.665)		
Post2	-3.967***			-5.738***		
	(0.796)			(1.135)		
N	114,418	114,418	114,418	112,711	112,711	112,711
X_{ijt}			x			x
Birth order FE	x	x	x	x	x	x
State FE	x	x	x	x	x	x
Birth year FE		x	x		x	x

NOTES: Sample of second- and higher-order births. 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. The vector X_{ijt} comprises mother's age at birth, 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: Excess female mortality across child ages

	Outcome: Death before				
	1 year	2 years	3 years	4 years	5 years
	(1)	(2)	(3)	(4)	(5)
*Firstborn girl * Female*	1.829***	2.330***	2.546***	2.781***	2.768***
	(0.476)	(0.578)	(0.581)	(0.610)	(0.588)
*Firstborn girl * Female * Post1*	-0.691	-0.827	-0.968	-1.212	-1.525*
	(0.607)	(0.731)	(0.723)	(0.726)	(0.768)
*Firstborn girl * Female * Post2*	-0.550	-1.095	-1.521*	-1.650**	-1.996**
	(0.424)	(0.710)	(0.750)	(0.712)	(0.744)
N	309,860	289,705	269,495	248,807	227,129
Baseline mean	9.65	11.07	11.75	12.12	12.38
X_{ijt}	x	x	x	x	x
Firstborn Girl x Birth year FE	x	x	x	x	x
Female x Birth year FE	x	x	x	x	x
Female x State FE	x	x	x	x	x
Female x Birth order FE	x	x	x	x	x
Birth order x Birth year FE	x	x	x	x	x
Birth order x State FE	x	x	x	x	x
State x Birth year FE	x	x	x	x	x
Firstborn girl x State FE	x	x	x	x	x
Firstborn girl x Birth order FE	x	x	x	x	x

NOTES: Sample of second- and higher-order births. Each column is a separate regression. The outcome is an indicator for death before age 1, 2, 3, 4, 5 for column (1), (2), (3), (4), and (5) respectively. For each column, we drop children that do not have full exposure to the risk of mortality. We always control for *Female* and fixed effects (FE) for birth year and birth order and for *Firstborn girl* and state FE. The vector X_{ijt} comprises mother's age at birth, 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. Data: NFHS. *** 1%, ** 5%, * 10%.

Table A.8: 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. Data: NFHS. *** 1%, ** 5%, * 10%.

Table A.9: 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

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. Data: NFHS. *** 1%, ** 5%, * 10%.

Table A.10: Under-5 mortality: Controls for mother's preferences and gender gaps in enrollment

Dep Var: Death before age 5	(1)	(2)
*Firstborn girl * Female*	0.793	2.899**
	(1.367)	(1.101)
*Firstborn girl * Female * Post*1	-0.747	-1.360
	(1.009)	(0.899)
*Firstborn girl * Female * Post*2	-1.877**	-2.028**
	(0.687)	(0.839)
Ideal number of children $\times$ *Female*	-0.652***	
	(0.223)	
Ideal sex ratio $\times$ *Female*	2.256***	
	(0.479)	
Ideal number of children $\times$ *Firstborn girl*	-0.414*	
	(0.226)	
Ideal sex ratio $\times$ *Firstborn girl*	0.0418	
	(0.330)	
Ideal number of children $\times$ *Female* $\times$ *Firstborn girl*	0.413	
	(0.334)	
Ideal sex ratio $\times$ *Female* $\times$ *Firstborn girl*	0.572	
	(0.637)	
*Girl-boy enrollment rate 6-11 * Female*		-0.795
		(2.514)
*Girl-boy enrollment rate 6-11 * Female * Firstborn girl*		0.327
		(1.454)
*Girl-boy enrollment rate 14-17 * Female*		0.145
		(0.226)
*Girl-boy enrollment rate 14-17 * Female * Firstborn girl*		-0.384
		(0.251)
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. Data: NFHS. *** 1%, ** 5%, * 10%.

Table A.11: Trends in son preference by firstborn sex

	Outcome variable: Ideal fraction of sons	
A. By survey round	(1)	(2)
Round 2	-0.0169***	-0.0134**
	(0.00586)	(0.00635)
Round 2 × Firstborn girl	0.00374	0.00407
	(0.00720)	(0.00733)
Round 3	-0.0441***	-0.0381***
	(0.00763)	(0.00771)
Round 3 × Firstborn girl	0.00770	0.00881
	(0.00774)	(0.00756)
B. By ultrasound availability	(1)	(2)
Post2	-0.0288***	-0.0226***
	(0.00560)	(0.00617)
Post2 × Firstborn girl	0.00510	0.00605
	(0.00638)	(0.00641)
N	10,160	10,160
Firstborn girl	x	x
State FE	x	x
Households controls		x

NOTES: The data pools the three waves of the NFHS, conducted in 1991/2, 1998/9 and 2005/6. For each survey round, we keep women who had their first birth within a calendar year from their survey data, so that their preferences at the survey date reflect their preferences around the birth of their first child. We then compare mother's reported son preference between firstborn-girl and firstborn-boy families and across the 3 survey rounds, that span about 15 years. In Panel A, we compare rounds 2 and 3 with round 1; the omitted group is the first NFHS round. In Panel B, we compare the two rounds (NFHS-2,3) in the late diffusion period of ultrasound availability with one round (NFHS-1) in the early diffusion period. Standard errors are clustered by state. *** 1%, ** 5%, * 10%.

Table A.12: Parental health investments, NFHS data

	Months breastfed (1)	Received $\geq$ 1 vaccine (2)	No. of antenatal checks (3)
*Firstborn girl * Female*	-0.373**	-0.028*	-0.076
	(0.172)	(0.014)	(0.061)
*Firstborn girl * Female * Post2*	0.056	0.022	0.061
	(0.291)	(0.015)	(0.071)
N	59,654	56,430	67,001
X_{ijt}	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: NFHS data. Sample include second and higher order births. 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. The sample in column (1) is restricted to children below age three and in column (2) to children of age $\geq$ 1. Standard errors in parentheses are clustered by state. *** 1%, ** 5%, * 10%.

Table A.13: Hazard of birth: Cox proportional-hazards model

Hazard Ratios	(1)	(2)	(3)
Firstborn girl	1.099***	1.093***	1.097***
	(0.00874)	(0.0148)	(0.0148)
*Firstborn girl * Post1*	0.975***	0.972***	0.968***
	(0.00523)	(0.00569)	(0.00674)
*Firstborn girl * Post2*	0.919***	0.917***	0.909***
	(0.0107)	(0.0105)	(0.00984)
Ideal fraction of sons			1.054***
			(0.0119)
Ideal no. of children			1.118***
			(0.00887)
N	2,286,065	2,286,065	2,114,681
X_i	x	x	x
Year FE	x	x	x
State FE	x	x	x
Age FE	x	x	x
Parity FE	x	x	x
State x Year FE	x	x	x
Years since last birth FE	x	x	x
Firstborn girl x State FE		x	x

NOTES: Hazard ratios from specification (2) estimated using a Cox proportional-hazards model on the 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. State-year fixed effects of the model are included by stratifying the Cox PH model by state-year. Data: NFHS. *** 1%, ** 5%, * 10%.

Table A.14: Birth spacing: Months between first and second birth

	No. of months
Firstborn girl	-1.094*
	(0.596)
*Firstborn girl * Post1*	1.342*
	(0.757)
*Firstborn girl * Post2*	1.901**
	(0.726)
N	41,329
X_i	x
Firstborn Girl x State FE	x
Year of first birth x State FE	x

NOTES: OLS regression. Sample include mothers who had exactly 2 births. The dependent variable is the number of months between first and second births. Standard errors in parentheses are clustered by state. Data: NFHS. *** 1%, ** 5%, * 10%.

Table A.15: 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)	(1)-(2)	(3)	(4)	(3)-(4)	(5)	(6)	(5)-(6)
Firstborn girl	0.113***	0.198***	-0.085***	0.126***	0.183***	-0.056	0.172***	0.142***	0.030
	(0.020)	(0.018)	(0.032)	(0.031)	(0.019)	(0.039)	(0.014)	(0.021)	(0.025)
*Firstborn girl * Post1*	-0.063*	-0.110***	0.047	-0.060	-0.099***	0.039	-0.114***	-0.054*	-0.060*
	(0.031)	(0.023)	(0.045)	(0.042)	(0.029)	(0.056)	(0.018)	(0.030)	(0.036)
*Firstborn girl * Post2*	-0.059**	-0.148***	0.089**	-0.069*	-0.121***	0.052	-0.132***	-0.088***	-0.044
	(0.028)	(0.024)	(0.040)	(0.036)	(0.027)	(0.048)	(0.021)	(0.029)	(0.035)
N	46,597	72,066		36,831	36,352		78,717	39840	
Baseline mean	3.230	2.812		3.101	2.769		3.026	2.964	

	Caste			**Rurality**		
	SC	General/OBC/ST		Rural	Urban	
	(7)	(8)	(7)-(8)	(9)	(10)	(9)-(10)
Firstborn girl	0.146***	0.154***	-0.008	0.134***	0.190***	-0.056**
	(0.039)	(0.012)	(0.039)	(0.017)	(0.015)	(0.023)
*Firstborn girl * Post1*	-0.014	-0.098***	0.084	-0.095***	-0.079***	-0.017
	(0.056)	(0.016)	(0.056)	(0.023)	(0.023)	(0.034)
*Firstborn girl * Post2*	-0.032	-0.124***	0.092**	-0.108***	-0.122***	0.015
	(0.041)	(0.018)	(0.041)	(0.026)	(0.018)	(0.030)
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. The third column in every panel shows a test of the significance of the difference between the coefficients for the sub-groups shown in the first two columns. Data: NFHS. *** 1%, ** 5%, * 10%.

Table A.16: Under-5 mortality: Heterogeneity by caste x socioeconomic status

	Illiterate		Literate		Botton 40% wealth		Top 20% wealth	
	SC	Non-SC	SC	Non-SC	SC	Non-SC	SC	Non-SC
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
*Firstborn girl * Female*	2.277	3.762***	-0.688	1.797**	2.477	3.551***	5.626**	0.970
	(1.852)	(0.761)	(3.227)	(0.724)	(2.483)	(1.008)	(2.700)	(0.796)
*Firstborn girl * Female * Post1*	-1.285	-2.325*	0.550	-0.711	-1.756	-1.845	-3.897	0.651
	(2.390)	(1.137)	(3.614)	(0.948)	(2.922)	(1.431)	(3.067)	(1.244)
*Firstborn girl * Female * Post2*	-0.201	-2.776**	-0.232	-1.091	-3.172	-2.069	-6.795	1.023
	(3.080)	(1.251)	(3.726)	(1.031)	(3.591)	(1.873)	(6.203)	(1.576)
	Unemployed		**Employed**		**Rural**		**Urban**	
	SC	Non-SC	SC	Non-SC	SC	Non-SC	SC	Non-SC
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
*Firstborn girl * Female*	2.045	3.056***	2.025	2.394***	2.297	3.757***	-1.757	0.493
	(2.033)	(0.724)	(2.073)	(0.702)	(2.139)	(0.693)	(2.285)	(1.015)
*Firstborn girl * Female * Post1*	-0.506	-1.924*	-1.952	-0.804	-1.388	-2.524**	2.899	0.969
	(1.890)	(0.957)	(3.093)	(1.036)	(2.675)	(0.997)	(3.093)	(1.210)
*Firstborn girl * Female * Post2*	-0.864	-2.716**	-1.894	-1.198	-0.889	-3.411***	1.681	0.907
	(2.638)	(1.050)	(3.267)	(0.973)	(2.879)	(1.161)	(2.993)	(1.388)

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 OLS regression. SC denotes scheduled 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. The third column in every panel shows a test of the significance of the difference between the coefficients for the sub-groups shown in the first two columns. Data: NFHS. *** 1%, ** 5%, * 10%.

Table A.17: Number of births: Hindus vs Muslims

	Hindus (1)	Muslims (2)
Firstborn girl	0.131***	0.161***
	(0.016)	(0.056)
*Firstborn girl * Post1*	-0.077***	-0.054
	(0.019)	(0.058)
*Firstborn girl * Post2*	-0.093***	-0.046
	(0.026)	(0.065)
Ideal no. of children	0.347***	0.260***
	(0.017)	(0.032)
Ideal fraction of sons	0.039**	0.011
	(0.018)	(0.026)
N	67,433	9,038
Baseline mean	2.997	3.200

NOTES: Estimates of specification (3) estimated using OLS. The dependent variable is the number of births at the time of interview. 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. Data: NFHS. *** 1%, ** 5%, * 10%.

Table A.18: Postnatal health investments: Hindus vs Muslims, REDS

	Months breastfed				Received >= one vaccine				Medicines/ special food during illness			
	Hindus	Hindus	Muslims	Muslims	Hindus	Hindus	Muslims	Muslims	Hindus	Hindus	Muslims	Muslims
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
First girl * Female	-4.0527*	-3.2497*	-0.2781	-1.8091	-0.0764*	-0.0445	-0.0032	0.058	-65.5785*	-115.2283**	120.6338	61.8803
	[1.9879]	[1.5970]	[5.6447]	[3.4082]	[0.0401]	[0.0486]	[0.1394]	[0.1030]	[35.5459]	[41.1597]	[83.6157]	[144.8619]
First girl * Female * Post1	3.0801	2.13	-6.0742	-5.9526	0.1022**	0.0773*	-0.0486	-0.1137	66.1963	107.7535*	-102.2985	-58.7846
	[2.5654]	[2.5635]	[5.4057]	[3.4693]	[0.0458]	[0.0374]	[0.1049]	[0.1274]	[48.7842]	[61.5769]	[85.5359]	[125.8920]
First girl * Female * Post2	2.9154	1.7382	-2.0404	-0.4341	0.0663**	0.0862**	0.0033	-0.0115	117.3814**	172.0607**	-60.7643	-43.5918
	[2.5750]	[2.2636]	[6.4090]	[8.0712]	[0.0296]	[0.0344]	[0.1024]	[0.1322]	[55.0821]	[66.3537]	[101.5542]	[165.4993]
N	9,726	9,726	734	734	18,138	18,138	1,619	1,619	9,079	9,079	849	849
*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	x	x	x	x	x	x
1st Girl x Birth year FE		x		x		x		x		x		x
Female x Birth year FE		x		x		x		x		x		x
Female x State FE		x		x		x		x		x		x
Female x Birth order FE		x		x		x		x		x		x
Birth order x State FE		x		x		x		x		x		x
State-specific time trends		x		x		x		x		x		x
Firstborn girl x State FE		x		x		x		x		x		x
Firstborn girl x Birth order FE		x		x		x		x		x		x

NOTES: This table reports investment effects for children of second- and higher-order birth order using the REDS sample, separately for Hindus and Muslims. 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, educational attainment of child's parents, and mother's birth cohort. Breastfeeding results are based on the last two surviving births of a mother and restrict the sample to children above age 2. Vaccination and health expenditure results are based on all surviving children of a mother. Standard errors in parentheses are clustered by state. *** 1%, ** 5%, * 10%.

Table A.19: EFM: Hindus vs Muslims

	Hindus (1)	Muslims (2)
A. Death before 1 month		
*Firstborn girl * Female*	1.581***	-0.254
	(0.428)	(0.887)
*Firstborn girl * Female * Post*	-0.833*	0.709
	(0.436)	(1.027)
N	241,234	47,775
B. Death before age 5		
*Firstborn girl * Female*	2.860***	2.636**
	(0.676)	(1.055)
*Firstborn girl * Female * Post*	-1.751**	-0.563
	(0.766)	(1.209)
N	168,193	31,938
X_{ijt}	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

NOTES: Sample of second- and higher-order births. Each column is a separate regression, with the same specification to column (3) in Table 3. We always control for *Female* and fixed effects (FE) for birth year and birth order and for *Firstborn girl* and state FE. The vector X_{ijt} comprises mother's age at birth, household wealth quintiles, caste, residence in a rural area, educational attainment of child's parents, and mother's birth cohort. Standard errors in parentheses are clustered by state. Data: NFHS. *** 1%, ** 5%, * 10%.

Table A.20: Under-5 EFM, across states

	Outcome: Death before age 5	
	Above median SRB	Below median SRB
	(1)	(2)
*Firstborn girl*Female*	3.174***	2.050**
	(0.794)	(0.830)
*Firstborn girl*Female*Post1*	-1.916*	-0.836
	(1.047)	(1.035)
*Firstborn girl*Female*Post2*	-2.317**	-1.493
	(1.021)	(0.975)
N	149,140	77,989
Mean state-level SRB	1.132	1.043

NOTES: The table present estimates from the same specification as column (3) of Table 3. Columns (1) and (2) include states that have male-female sex ratio at birth (SRB) above and below median among all states, respectively. SRB is calculated using population counts of total number of children at age 0 in the state from the 2011 Census of India. Data: NFHS.

Table A.21: Number of births, across states

	Outcome: Number of births	
	Above median SRB	Below median SRB
	(1)	(2)
*Firstborn girl*Female*	0.199***	0.137***
	(0.016)	(0.019)
*Firstborn girl*Female*Post1*	-0.104***	-0.068**
	(0.021)	(0.025)
*Firstborn girl*Female*Post2*	0.129***	-0.092***
	(0.022)	(0.027)
N	69,254	49,409
Mean state-level SRB	1.132	1.043

NOTES: The table present estimates from the same specification as column (1) of Table 6. Columns (1) and (2) include states that have male-female sex ratio at birth (SRB) above and below median among all states, respectively. SRB is calculated using population counts of total number of children at age 0 in the state from the 2011 Census of India. Data: NFHS.

Table A.22: Under-5 EFM with different recall periods

	Recall period from interview			
	10 years	12 years	15 years	20 years
Outcome: Death before age 5	(1)	(2)	(3)	(4)
*First girl*Female*	1.737*	2.172**	2.715***	2.768***
	(0.897)	(0.780)	(0.705)	(0.588)
*First girl*Female*Post1*	-0.364	-0.964	-1.516	-1.525*
	(0.992)	(0.976)	(0.919)	(0.768)
*First girl*Female*Post2*	-0.954	-1.380	-1.944**	-1.996**
	(1.098)	(0.998)	(0.878)	(0.744)
N	134,986	169,344	206,830	227,129
X_{ijt}	x	x	x	x
Firstborn Girl x Birth year FE	x	x	x	x
Female x Birth year FE	x	x	x	x
Female x State FE	x	x	x	x
Female x Birth order FE	x	x	x	x
Birth order x Birth year FE	x	x	x	x
Birth order x State FE	x	x	x	x
State x Birth year FE	x	x	x	x
Firstborn girl x State FE	x	x	x	x
Firstborn girl x Birth order FE	x	x	x	x

NOTES: Sample of second- and higher-order births. Each column is a separate regression. Column (1)-(4) include samples with recall period within 10,12, 15, and 20 years. Specification same as column (3) of Table 3. 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. The vector X_{ijt} comprises mother's age at birth, 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. Data: NFHS. *** 1%, ** 5%, * 10%.

Table A.23: Under-5 EFM: Controlling for State x Year x Female FE

Outcome: Death before age 5	(1)	(2)
*Firstborn girl * Female*	2.780***	2.796***
	(0.575)	(0.580)
*Firstborn girl * Female*Post1*	-1.566**	-1.576*
	(0.758)	(0.764)
*Firstborn girl * Female*Post2*	-1.976**	-2.004**
	(0.740)	(0.740)
N	227,129	227,129
X_{ijt}	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
Female x X_{ijt}		x
Firstborn girl x State x *Female* FE	x	x

NOTES: This table checks the robustness of our findings to the inclusion of State x Year x Female FE. Columns (1) and (2) correspond, respectively, to columns (3) and (4) in Table 3. Standard errors in parentheses are clustered by state. Data: NFHS. *** 1%, ** 5%, * 10%.

Table A.24: Comparison with Hu and Schlosser (2015)

	Children born within:							
	36 months before survey		120 months before survey			20 years before survey		
Died within:	1 month	12 months	1 month	12 months	5 years	1 month	12 months	5 years
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$EFMR_{st}$	0.003	0.003	-0.004	-0.008**	-0.014**	-0.003	-0.005	-0.004
	(0.004)	(0.005)	(0.002)	(0.003)	(0.006)	(0.003)	(0.003)	(0.005)
$Female$ x $EFMR_{st}$	-0.001	0.005	0.007**	0.015***	0.020**	0.004	0.008*	0.007
	(0.005)	(0.008)	(0.003)	(0.004)	(0.007)	(0.003)	(0.004)	(0.005)
N	83,961	58,596	318,555	293,190	173,146	550,075	524,710	404,666

NOTES: This table presents estimates of equation (5). The outcome is an indicator of mortality. $EFMR_{st}$ is computed as the difference of the female-male sex ratio at birth between firstborn girl and firstborn boy families, averaged for each state and year cell, and is computed using our version of the NFHS data. The specification includes household characteristics, state-female fixed effects and year-female fixed effects. Twin indicator and mass media exposure index are not included in the controls, and regressions are unweighted. Standard errors clustered at the state level are reported in parentheses. *** 1%, ** 5%, * 10%.

Table A.25: Replacing post indicators in our specification with $\overline{MFR_{st}}$

	Children born within:					
	20 years before survey			10 years before survey		
Died within:	1 month	12 months	5 years	1 month	12 months	5 years
	(1)	(2)	(3)	(4)	(5)	(6)
A. Our estimates						
$Firstgirl$ x $Female$ x $Post1$	-0.00564	-0.00691	-0.0153*	-0.0115	-0.0205*	-0.0254**
	(0.00457)	(0.00607)	(0.00768)	(0.00769)	(0.0103)	(0.0107)
$Firstgirl$ x $Female$ x $Post2$	-0.00570	-0.00550	-0.0200**	-0.0111	-0.0194*	-0.0301**
	(0.00385)	(0.00424)	(0.00744)	(0.00823)	(0.0103)	(0.0122)
N	328,072	309,860	227,129	212,385	194,173	111,442
B. Replacing $Post1$ and $Post2$ with $\overline{MFR_{st}}$						
$Firstgirl$ x $Female$ x $\overline{MFR_{st}}$	-0.154	-0.018	-0.966	-0.092	-0.120	-0.964
	(0.371)	(0.503)	(0.751)	(0.506)	(0.744)	(0.848)
N	328,072	309,860	227,129	212,385	194,173	111,442

NOTES: This table presents estimates of equation (1) in Panel A and of equation (6) in Panel B. The outcome is an indicator of mortality. The purpose of this table is to show how the estimates change if we start with our specification and replace the post-ultrasound indicators with the average MFR used in HS. Standard errors clustered at the state level are reported in parentheses. *** 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
- Expenditure 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 expenditure: amount (in Rupees) spent on doctors' fees, medicine, and special food last year
- Ideal fraction of sons: (Ideal no. of boys + 0.5* Ideal no. of either sex) / (Ideal no. of boys + Ideal no. of girls + Ideal no. of either sex) (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 1.996 p.p. (column (3) 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,[62] the implied mortality rate for breastfed children is 1.35 percent[63] 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 19 percent (0.371/1.996 = 0.186) 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 (0.131/1.996 = 0.066) to 10 percent (0.204/1.996 = 0.102) 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.

[62] 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).

[63] 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:

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

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:

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

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

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

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..[64] 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

[64]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 $S_{a,FG} = S_{FG}$:

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

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