Does Prenatal Sex Selection Improve Girls' Well-Being? Evidence from India[1]

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ABSTRACT

The paper studies the impact of prenatal sex selection on the well-being of girls by analyzing changes in children's nutritional status and mortality during the years since the diffusion of prenatal sex determination technologies in India. We use the ratio of male to female births in the year and state in which a child was born as a proxy for parental access to prenatal sex-selection. Using repeated cross-sections from a rich survey dataset, we show that high sex ratios at birth reflect the practice of sex-selective abortion. We then exploit the large regional and time variations in the incidence of prenatal sex selection to analyze whether changes in girls' outcomes relative to boys within states and over time are associated with changes in sex-ratios at birth. We find that an increase in the practice of prenatal sex selection appears to be associated with a reduction in the incidence of malnutrition among girls. The negative association is stronger for girls born in rural households and at higher birth parities. We find no evidence that prenatal sex selection leads to a selection of girls into families of higher SES, however we do find some evidence of a larger reduction in family size for girls relative to boys. We also find some suggestive evidence of better treatment of girls as reflected in breastfeeding duration. On the other hand, prenatal sex selection does not appear to be associated with a reduction in excess female child mortality.

Keywords: Son preference, prenatal sex selection, sex ratio at birth, gender discrimination, child health.

JEL codes: J13, J16, I1

I. INTRODUCTION

Son preference in India and other East Asian countries has been documented extensively. A large number of studies have shown that parental preference for boys is manifested in gender differences in the intrahousehold allocation of resources and medical care, which result in gender differentials in child outcomes such as nutrition, morbidity and mortality and, in extreme cases, female infanticide (see, for example, Arnold et al., 1998; Deaton, 1997; George, 1997; Miller, 1981; Kishor, 1993; Rose, 1999; Rosenzweig and Schultz, 1982; and Barcellos et al., 2010). The diffusion of prenatal sex determination technologies since the 1980s has provided parents with an alternative way to achieve the preferred sex composition of children. Indeed, starting from the late 1980s, there has been a steep increase in male to female ratios at birth, which has been attributed to the increasing practice of sex-selective abortion. A growing number of studies have examined the factors affecting the prevalence of prenatal sex selection and have assessed its effect on sex ratios at birth.[2] Little is known, however, about the impact of this phenomenon on the well-being of the girls who are born.

In this paper, we study the impact of pre-natal sex selection on the well-being of girls in India by analyzing changes in children's nutritional status and mortality during the period since the diffusion of prenatal sex determination technologies. Some researchers and policy makers have speculated that prenatal sex selection should reduce the number of unwanted girls and therefore improve the outcomes of girls who are born.[3] In particular, it has been suggested that prenatal sex selection may substitute for postnatal sex discrimination (see, for example, Goodkind, 1996 and Shepherd, 2008). Prenatal sex selection may also affect girls' well-being through a reduction in family size. Several studies have shown that in populations with strong son preference, girls have, on average, more siblings than boys since parents of girls continue to have additional children in order to attain the desired number of sons (see, for example, Yamaguchi, 1989; Arnold et al., 1998; and Jensen, 2003). Access to prenatal sex selection allows parents to rely on alternative family-building strategies that likely affect family size and birth spacing and therefore child outcomes.

India constitutes an interesting experimental case for studying the impact of prenatal sex selection due to its large regional variations in parental preference for sons and the striking differences in the use of prenatal sex-selection technologies over time and across regions. Using a rich data set on child outcomes and household characteristics, we apply a *triple-differences* approach to examine whether

[2] We review some of this literature in the next section.
[3] This is a similar rationale to that proposed by Donohue and Levitt (2001) and Ananat et al. (2009) in the context of abortion legalization in the US.

changes in the nutritional status and mortality rates of girls relative to boys within states and over time are systematically associated with changes in the prevalence of prenatal sex selection. In the absence of a direct measure of parental access to prenatal sex selection, we use the ratio of males to females at birth (MFR) in the child's state and birth cohort and provide evidence showing that increases in MFR at birth reflect the practice of prenatal sex selection starting from the late 1980s. We then analyze whether changes in the outcomes of girls versus boys within states and over time are associated with changes in MFR at birth. We also apply an alternative strategy and compare changes in outcomes of girls versus boys over time between regions in which the prevalence of prenatal sex selection is known to be increasing and regions in which it is rarely practiced.

The paper is related to a limited number of recent studies, most of them conducted concurrently with our own, that examine the effects of prenatal sex selection on girls' outcomes. Using data from the first two rounds of the Indian National Family Health Survey (NFHS), Shepherd (2008) examines whether there is differential change over time in the outcomes of girls versus boys between families with high versus low likelihood of using prenatal sex selection and finds inconclusive evidence of a substitution effect on mortality. Lin et al. (2009) examine the effects of an increase in prenatal sex selection in Taiwan on female mortality using variation in access generated by the legalization of abortion in 1985-6. Their findings suggest a positive association between prenatal sex selection and female survival. Almond et al. (2010) use variation in the diffusion of ultrasound technology across regions in China and find that prenatal sex determination is associated with an increase in girls' neonatal mortality, thus suggesting that prenatal investment in girls is reduced as a result of sex determination. They do not find any changes in postnatal investment in girls as manifested in breastfeeding duration and vaccination rates.[4]

Our paper adds to this limited literature in several aspects. As mentioned above, India's high degree of heterogeneity in female discrimination and the differences in the practice of prenatal sex selection across regions makes it an optimal case for studying the effects of prenatal sex selection. This is also due to India's large variation in family size across households, which provides an opportunity to examine the extent and impact of prenatal sex selection across different parities and study the impacts of prenatal sex selection on fertility decisions. A unique characteristic of our study is the availability of a rich data which enable us to examine the extent of prenatal sex selection and its impact on both mortality and

[4] Other recent related studies are Portner (2010), which looks at the determinants of sex-selective abortion in India though not their effect on outcomes, and Bhardadwaj and Nelson (2010) which examine gender differences in prenatal investment in countries with strong son preference, including India, although they do not estimate the impact of sex-selective abortion.

nutritional outcomes. In addition, we are able to explore various mechanisms underlying the relationship between prenatal sex selection and child outcomes. In comparison to Shepherd (2008), we expand the sample to include more recent data (the third round of NFHS), which allows us to verify the continuing upward trend in the practice of prenatal sex selection and to explore its impact on child outcomes. We also use a different empirical strategy to that used in Shepherd (2008) in order to test the substitution hypothesis. Rather than using the estimated propensity of prenatal sex selection at the household level, we exploit variation in MFR at birth across states and over time in order to proxy for the change in the incidence of prenatal sex selection. Essentially, we adopt both a continuous and a discrete version of a triple-differences approach. This approach has several appealing features as it allows us to control for regional variation in son preference by looking at changes within states over time. Moreover, we are also able to control for differential trends across states by examining changes in the relative outcomes of boys versus girls.

Our findings can be summarized as follows. (1) An increase in the practice of prenatal sex selection appears to be associated with a reduction in the prevalence of malnutrition among girls. (2) The negative association is stronger in some subpopulations in which discrimination against girls is usually more pronounced (for example in rural areas and among children of parity higher than one). (3) We find no evidence that prenatal sex selections lead to a selection of girls into families of higher SES since we do not observe a differential improvement in the family characteristics of girls versus boys. (4) However, we do find some evidence of a larger reduction in family size for girls versus boys and some suggestive evidence of better treatment of girls as measured by breastfeeding practices. In particular, we find some reduction in girls' disadvantage in breastfeeding duration in regions that experienced an increase in the incidence of prenatal sex selection. (5) On the other hand, prenatal sex selections do not appear to be associated with a reduction in girls' excess mortality.

The rest of the paper is organized as follows: In the next section, we review the literature and describe the institutional background of unbalanced sex ratios and prenatal sex selection in India. Section III discusses the conceptual framework for analyzing the effects of prenatal sex selections on child outcomes. Section IV describes the data. Section V lays out the empirical strategy and Section VI presents the results. Finally, Section VII concludes.

II. BACKGROUND AND INSTITUTIONAL FRAMEWORK

Imbalanced sex ratios have been documented in India as early as the 19th century and throughout the 20th century (see Visaria, 1971; Miller, 1981, 1984; Dyson and Moore, 1983; and Sen, 1990, who introduced the concept of "Missing Women"). Until the late 1970s, gender imbalances were mostly manifested at older ages and not particularly at birth and were attributed to excess female mortality due to maltreatment and neglect and in extreme cases to female infanticide (see, for example, Dreze and Sen, 1997 and Das Gupta, 1987).[5]

Several studies have documented sharp increases in male to female ratios at birth since the late 1980s, especially in northern and western states, which are regions historically known for strong sex discrimination and son preference (see, for example, Das Gupta and Bhat, 1997; Arnold et al., 2002; Bhat, 2002; Bhaskar and Gupta, 2007; and Retherford and Roy, 2003). These studies have shown that the increase in sex ratios at birth is concurrent with the diffusion of prenatal sex determination technologies, thus suggesting that sex-selective abortion is likely to be a major contributing factor.[6]

Sex-selective abortion requires two steps: sex determination of the fetus and an abortion. Abortion was legalized in India under the Medical Termination of Pregnancy Act (MTP) in 1972. Despite its legalization, most abortions are still practiced in unofficial and non-regulated facilities. According to Chhabra and Nuna (1993), the majority of the six to seven million abortions carried out annually take place in unofficial facilities. Based on a series of field surveys and qualitative studies conducted for the Abortion Assessment Project, Duggal and Ramachadran (2004) conclude that certified and legal abortion facilities account for only 24% of all private abortion facilities in the country. With regard to the motive for seeking an abortion, the authors find that only 31% of the abortions fell strictly within the limits established under the MTP Act; the rest were due to unwanted pregnancy (71%), economic reasons (7%), and the undesired gender of the fetus (13%).[7]

Sex determination during pregnancy was first made possible in the late 1970s by the use of amniocentesis (Jefferey et al., 1984) and became more accessible with the introduction of ultrasound

[5] Other possible explanations attributed to the lower number of females include census underenumeration of girls or different patterns of age misreporting by sex. While these phenomena might explain part of the gender imbalances, Visaria (1971), Miller (1981, 1984), and other researchers show that the main cause for imbalanced sex ratios until the late 1970s was sex differentials in mortality.

[6] Changes in the enumeration of girls or misreporting of age are unlikely to explain this upward trend since similar increases in sex ratios at birth are also observed among Indian populations living in the US, Canada, and the UK where birth registration is nearly complete and accurate (see Abrevaya, 2009; Almond and Edlund, 2008; Almond et al., 2009; and Dubuc and Coleman, 2007). We provide in the next section further evidence which suggests that sex-selective abortion is the main cause for the upward trend in sex ratios at birth.

[7] The total does not add up to 100% since it was a multiple response question.

technologies in the 1980s. In subsequent years, a continued decline in desired fertility coupled with a slower decline in the total number of desired sons increased the pressure to have sons at lower parities, thus raising the demand for sex-selective abortions (Das Gupta and Bhat, 1997). At the same time, economic development and trade liberalization accelerated the supply of the technology necessary to practice sex-selective abortion. As a result of the policy reforms of the 1990s, multinationals such as General Electric started manufacturing ultrasound machines in India, thus improving access to prenatal sex determination and increasing the practice of prenatal sex selection throughout the country (Geroge, 2006). The diffusion process took place from households of high socioeconomic status to those of low socioeconomic status and from urban to rural areas (Khanna, 1997).

In an attempt to eliminate the practice of sex-selective abortion and in response to increasing public pressure from several NGOs and women's organizations, the government of India passed the Prenatal Diagnostic Techniques Regulations and Misuse Act (PNDT Act) in 1994, thus making it illegal to use ultrasound or amniocentesis in order to determine the sex of a fetus. However, this legislation proved to be ineffective and the practice of sex-selective abortion continued to spread (see, for example, George, 2002 and Kishwar, 1995). In subsequent years, several private ultrasound clinics opened and mobile clinics and portable ultrasound machines made sex-selective abortion accessible in small towns and rural areas (Krugman, 1998) and most of the sex-selective abortions moved to underground clinics (Everett, 1998).

The publication of sex ratio figures from the 2001 Census revealed a further increase in MFR at ages 0-6, which led to increased public pressure on the government to enforce and expand the legal power of the PNDT Act. In 2001, the Centre for Enquiry into Health (CEHAT) and the Mahila Sarvangeen Utkarsh Mandal (MASUM) filed a public interest litigation against the Union of India and all the state governments for the non-implementation of the PNDT Act and for the inclusion of all emerging technologies that could be used for sex-selective abortion. As a result, the Indian Government decreed in 2001 that manufacturers of ultrasound equipment could sell the machines only to registered clinics, which must maintain strict records of their use. One of the amendments to the Act, which were approved by the Indian Parliament in 2003, was the inclusion of technologies that allow sex selection during the preconception and pre-implantation stage. Another important addition to the scope of the Act was the requirement on the part of clinics and doctors who provide ultrasound services to maintain written records that specify the specific reason why an ultrasound or amniocentesis test had been recommended.[8] Recent reports indicate some improvement in the enforcement of the Act. However, sex-selective abortion is still being practiced extensively and enforcement of the law appears to be difficult, if not impossible, to achieve (Subramanian and Selvaraj, 2009).[9] A recent study by Portner (2010) also shows that the enforcement of the PNDT Act has little impact.

III. CONCEPTUAL FRAMEWORK

There are various channels through which parental preference for sons can affect girls' outcomes in the absence of a technology that allows them to realize their preferences for sex composition. First, parents can allocate more resources to children that provide them with higher utility (i.e. boys) as long as boys and girls are substitutes in parents' utility function (Rosenzweig and Schultz, 1982).[10] A similar conclusion can be drawn from Becker and Tomes (1979, 1986) when capital markets are imperfect.

Parents can also attempt to adjust excess fertility and undesired sex composition of children through selective allocation of resources that results in the mortality of the less-desired children (Simmons et al., 1982). Alternatively, if parents continue to bear children until they achieve their desired number of sons, girls will have, on average, a larger number of siblings than boys (Yamaguchi, 1989). Therefore, even if parents treat boys and girls equally within a household, girls can be expected to grow up in larger families with fewer resources per child (Jensen, 2003). In addition, parental preference for boys is likely to affect birth spacing (Ben-Porath and Welch, 1976) and duration of breastfeeding (Jayachandran and Kuziemko, 2009) since parents may wish to have a child sooner following the birth of a daughter, especially if they are approaching the end of the mother's fecund years. Both factors are known to affect children's health.

Increased access to sex-selective abortion may improve the outcomes of girls who are born for a number of reasons. First, girls will be more likely to be born into families that have a lower disutility

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[8] For a description of the PNDT Act and its enforcement, see Mallik (2003) and George (2002).

[9] For the state of implementation and enforcement of the Act see the section on the PNDT Act in annual reports published by the Ministry of Health and Family Welfare at http://pndt.gov.in/index1.asp?linkid=15 and the report of the National Commission for Women, India at http://ncw.nic.in/Pdfreports/PC_PNDT_REPORT.pdf.

[10] Differences in parents' utility from boys and girls might be rooted in taste differences or in differences in the costs and benefits of raising boys versus girls (Ben-Porath and Welch, 1976). Arnold et al. (1998) provide a review of the literature and highlight three main channels that affect parental preferences for sons. The first is economic utility which is based mainly on differences in labor market productivity between boys and girls (either in the form of wages or agricultural productivity), security in the case of illness, support in old age and inheritance rules. A second channel is related to social utility, which provides families with more sons a higher status and additional sources of income in the form of dowry payments. A third channel is religious utility, which is derived from the performance by sons of important religious functions, especially upon the death of his parents. Girls, on the other hand, are seen as an economic drain on family resources because of the dowry system and the high cost of weddings.

from girls and therefore will suffer less from discrimination in resource allocation within the household. Second, since parents can achieve their preferred number of boys and girls without the need of having additional children, family size will not necessarily be larger for girls. Girls' outcomes might also improve as a result of selection into better households if parents who practice sex-selective abortion have different characteristics than parents who do not. For example, if poorer families are more likely to abort girls, then girls will be more likely to be born into better-off families and would therefore be expected to have better outcomes even if parental treatment has not changed.[11] On the other hand, prenatal sex determination might have a negative impact on girls' outcomes if parents reduce their prenatal investment in female fetuses in response to prenatal sex determination rather than having an abortion (see, for example, Bharadwaj and Nelson, 2010 and Almond et al., 2010).

IV. DATA

The data for our empirical analysis are taken from the National Family Health Survey (NFHS). The NFHS is a large-scale, multi-round survey conducted in a representative sample of households throughout India. Three rounds of the survey were conducted in 1992-3, 1998-9 and 2005-6.[12] Each round covered approximately 90,000 households which contained more than 500,000 individuals and was designed to provide state-level and national-level estimates. The survey includes detailed information on the demographic and socioeconomic background of the household members, as well as additional modules designed to investigate health, fertility and mortality.

We pooled the three survey rounds and selected only children who live with their biological mother and whose mother is between 15 and 49 years old in order to obtain a consistent sample across the three survey rounds. One major advantage of the dataset is that is has complete birth histories of a large number of women over a long period, making it feasible to estimate MFR at birth by state and cohort.

Table 1 reports summary statistics for male and female children included in our main samples (columns 1, 2, 4, and 5) and differences between the characteristics of girls and boys (columns 3 and 6). The nutritional status sample reported in columns 1-3 includes the last two children aged less than three of ever-married women sampled in one of the three NFHS rounds who have valid anthropometric data. The children sample reported in columns 4-6 includes all children of ever-married women born within

[11] In fact, it has been shown that sex-selective abortion is more prevalent among households with higher socioeconomic status (see, for example, Retherford and Roy, 2003) and therefore we do not expect to see an improvement in girls' outcomes that derives from the selection of girls into wealthier families.

[12] The NFHS was designed along the lines of the Demographic and Health Survey (DHS) that has been conducted in many developing countries since the 1980s.

the 10 years preceding each survey round. Household characteristics reported in the table are used as control variables in the empirical analysis.

Most children (about 75 percent) in both samples live in rural areas. About half of the children have mothers with no education and about 30 percent have fathers with no education. Mothers' age at first birth is relatively low at 19, on average. At the time of the survey, mothers in the nutritional analysis sample were 25 years old, on average, and they had an average of 3 children. Mothers in the children's sample are 29 years old on average at the time of the survey and have an average of about 4 children.

Columns 3 and 6 show that girls tend to be born into more disadvantaged families than boys. Their families have lower wealth levels, lower parental education and a lower degree of exposure to mass media. A possible explanation for differences in the family characteristics of girls versus boys is the practice of sex-selective abortion.[13] Another fact worth noting is that girls appear to have more siblings than boys in the children's sample, which is consistent with parental stopping rules in fertility behavior and son preference. There is no similar difference in the nutritional status sample, which is likely due to the fact that a large proportion of children in the sample come from households with incomplete fertility. It is important to bear these facts in mind when we later discuss possible mechanisms underlying the main results (in particular, the selection and family size channels).

V. EMPIRICAL STRATEGY

As described in the previous sections, the practice of sex-selective abortion has steadily increased in India during the last three decades though its incidence varies considerably across regions (Arnold et al., 2002 and George, 2002) and over time. We therefore exploit this variation in order to examine the effect of prenatal sex selection on the outcomes of girls.

One potential limitation of this analysis is that we do not directly observe the practice of sex-selective abortion. However, we do observe its consequences, primarily the abnormal sex ratio at birth. **Figure 1** shows the variation among states in male-female ratios (MFR) at birth as reported in census records from 1961 to 2001. Until the 1980s, MFR at birth did not exceed the normal ranges of 103-107 males per 100 females as found in various large-scale studies (for example, Visaria, 1971 and Jacobsen et al., 1999). Increases in MFR at birth become evident at the transition points of 1981, 1991 and 2001.

[13] In a different analysis (not shown here), we examined how family characteristics of boys versus girls differ over time by looking at differences between boys' and girls' households over the three survey rounds. The comparison suggests that the gaps between boys' and girls' family characteristics has widened over time. This is consistent with the increasing trend in the practice of sex-selective abortion. Nevertheless, since sex-selective abortion is also diffusing to rural and less disadvantaged areas, the gap may decline in the next few years.

Further evidence of the diffusion of sex-selective abortion is shown in **Figures 2a and 2b** which plot sex-ratios at birth by state stratified by urban and rural areas. The figures show that sex ratios at birth have also increased in rural areas though the trend has lagged somewhat behind that in urban areas. This is also noted by Retherford and Roy (2003) who examine sex ratios at birth over the first two rounds of the NFHS and report higher sex ratios in urban as opposed to rural areas in the first round but find no evidence of a rural/urban differential in the second round once socioeconomic variables are controlled for.

Interestingly, there are large variations in MFR (both in their levels and in their growth rates) even across those states that appear to have a strong preference for boys. To illustrate this point, Panel a of **Table 2** shows data on MFR at birth by state and selected indicators for son preference as reported by women in the first round of the NFHS (in 1992). The table reports MFR at birth by state for various census years (columns 1-5) and indicators of fertility and preference for number and sex composition of children based on tabulations from the first round of the NFHS (columns 6-9).[14] States are grouped by region.

The largest increases in sex ratios at birth are found in the northern states, which are characterized by a strong degree of son preference. In Punjab, for example, while MFR at birth was within the normal range between 1961 and 1981, it increased dramatically between 1981 and 2001 from 106 to 129. Western states also exhibit strong son preference and a large increase in sex ratios. For example, in Gujarat, MFR at birth remained at 103 between 1961 and 1981 but increased from 103 to 116 between 1981 and 2001. Both northern and western states appear to have strong son preference as manifested by the ideal sex ratio reported by mothers (1.46 and 1.33 as reported in column 7) and the proportion of mothers who desire a larger number of sons than daughters (0.48 and 0.42 as reported in column 9).

While strong son preference is found in states with upward trends in MFR, we also observe that in several states in the northeast, central and eastern regions with similar strong preferences for sons there was either no increase in MFR or only a mild one. In Madhya Pradesh, for example, mothers reported an ideal sex ratio of 1.44 but sex ratios at birth remained close to natural levels (MFR of 106 in 2001). Similarly, sex ratios in the northeastern states of Arunachal Pradesh, Assam and Manipur have not increased significantly despite the relatively high ideal MFR reported by mothers in these states.

[14] Women with living children were asked: "If you could go back to the time you did not have any children and could choose exactly the number of children to have in your whole life, how many would that be?" Women with no living children were asked, "If you could choose exactly the number of children to have in your whole life, how many would that be?" All women who gave a numerical response to the question on the ideal number of children were also asked how many of these children they would like to be boys, how many they would like to be girls and for how many the sex would not matter.

Southern states are usually characterized by a low degree of son preference and stable sex ratios at birth.

Overall, we conclude that while strong son preferences are found in states that exhibit increases in MFR at birth, this factor cannot alone explain state variation in MFR since there are several states with strong son preference that have not shown any significant increase in MFR. Panel b of **Table 2** presents several economic and demographic indicators by state for the years 1991-2. As can be clearly seen, northern and western states have a higher degree of development relative to states in the northeast, central and eastern regions. This is reflected by a higher wealth index, income per capita, share of households with electricity and degree of exposure to mass media (TV). On the other hand, development and income levels in many of the northern and western states are comparable to those in southern states where sex ratios have remained balanced. In terms of women's educational level and religion, it is hard to find a clear pattern that differentiates states with increases in MFR from the rest. For example, MFR has increased significantly in Punjab, which has a high share of Sikhs and also in Himachal Pradesh where the majority of the population is Hindu.

Overall, the picture obtained from **Table 2** suggests that the primary factors which distinguish between states with an increasing MFR are a strong preference for boys and a higher degree of development and modernization (in combination). Still, there are some exceptions, such at the state of Rajasthan which is poorer and less developed than other states with increasing MFR. There is also a clear geographical pattern that points to a higher incidence of sex-selective abortion in northern and western states. The fact that there are states with a strong preferences for boys and states with high levels of development that have not exhibited significant increases in MFR provide us with a heterogeneous group of states that are comparable to states with increasing MFR across different dimensions.

Figure 1 and Table 2 show a high variation in MFR across states and over time. But is the increase in MFR directly related to the practice of sex-selective abortion? Several studies have shown that increases in MFR at birth occur simultaneously with the diffusion of prenatal sex determination technologies (see for example, George, 2002). In order to rigorously examine this question, we test whether the propensity of giving birth to a boy is higher among families who presumably have a stronger desire for sons and whether this propensity changed when prenatal sex-determination became feasible. In **Table 3**, we report the likelihood of a male birth at parity *N* (two or three) as a function of the sex composition of the older siblings who were alive at the time of conception using a linear probability model. We examine two samples: children born between 1975 and 1989 and children born from 1990 onwards.

This split is meant to proxy for the availability of ultrasound technology. As seen in column 1, the probability of a male birth during the pre-ultrasound period did not vary significantly across households according to the sex composition of their previous children. In contrast, column 5 shows that during the post-ultrasound period this probability was significantly higher for households that had only girls. Estimates of the differential probabilities are virtually unchanged after adding controls for household characteristics (columns 2 and 6).

As noted earlier, sex determination technologies spread from urban to rural areas, thus enhancing access to prenatal sex selection. In columns 3-4 and 7-8 of **Table 3** we provide evidence on the incidence of sex-selective abortion stratified by urban and rural residence. The estimates suggest that the likelihood of a male birth increased both in rural and urban areas between the pre- and post-ultrasound periods among households who are more likely to desire a boy.

Evidence presented in **Table 3** shows that the likelihood of a male birth increased significantly in the 1990s among households who presumably had a stronger desire for a son. This suggests that increases in MFR at birth are likely to be a result of access to sex-selective abortion. In order to further examine the link between son preference and sex ratios at birth, additional information on ultrasound use and abortions is needed. Information on the incidence of abortions is difficult to obtain since a large share of abortions take place in unofficial and non-regulated facilities. In addition, abortion is usually misreported, especially if it is carried out for sex-selection purposes. We therefore examine patterns of ultrasound use.

The second and third round of the NFHS survey asked mothers whether they performed an ultrasound test for each of their births during the three or five years prior to the survey. While most ultrasound tests are performed as part of routine antenatal checkups, the association between ultrasound tests and sex ratios can provide suggestive evidence for the practice of sex-selective abortion. In columns 1 through 4 of **Table 4** we report the differential likelihood that a mother performed an ultrasound test during a pregnancy of parity N as a function of the sex composition of her N-1 previous children. As clearly seen in the table, mothers with no living sons are more likely to perform an ultrasound test in pregnancies at parities 2, 3, and 4. This increased likelihood is present in both urban and rural areas.

The link between ultrasound use and prenatal sex selection can be further examined by looking at the associations between a male birth and the sex composition of older children among mothers who reported conducting an ultrasound test during pregnancy. Estimates shown in columns 5 trough 8 suggest mothers who reported doing an ultrasound test during pregnancy have a significantly higher probability of giving birth to a boy if they have no older sons. A comparison between estimates in column 6 in Table 3 and column 6 in Table 4 shows that the increase in the likelihood of a male birth among mothers with no living sons is significantly higher in the sample of mothers who report doing an ultrasound test relative to the sample of mothers who gave birth after 1989. For example, the difference in the probability of a male birth at parity two between mothers with a girl at parity one and mothers with a boy at parity one is 7 percentage points in the sample of mothers who reported an ultrasound test at parity two. In contrast, the difference in probability for mothers who gave birth after 1989 is 2 percentage points.

In summary, the evidence presented in Tables 3 and 4 suggests that increases in MFR at birth are likely to be induced by the increasing practice of sex-selective abortion rather than by biological or environmental factors. We therefore use MFR at the state level as a proxy for the practice of sex-selective abortion.

How should we interpret changes in the incidence of sex-selective abortion (as proxied by MFR)? In principle, an increase in sex-selective abortion could be due to an increasing preference for sons or increased access to sex-determination technology.[15] Several studies have shown that despite the decline in the self-reported preference for boys over time, MFR at birth has increased (see, for example, Bhat and Zavier, 2003 and Retherford and Roy, 2003). This can be explained by the greater pressure to have a boy at lower parities due to the decline in ideal family size and the increased availability of prenatal sex-selection technologies. We therefore, interpret the increase in MFR as being due to these last two factors (i.e. an increase in access to the technology and a decline in the desired number of children) rather than an increase in discriminatory preferences against girls. Note that an increase in MFR that reflects an increase in the discriminatory preferences of parents would tend to rule out a finding that girls' well-being has improved. In that case, our estimates would provide a lower bound to the effects of sex-selective abortion.[16]

We exploit the variation in MFR across states and over time in order to examine the impact of sex-selective abortion on girls' outcomes. Specifically, we analyze whether changes in MFR within a state

[15] We see an increase in access as equivalent to a reduction in the cost of using the technology (which could be either a monetary cost or the psychic cost of a sex-selective abortion).

[16] On the other hand, it is also possible that discriminatory preferences against girls have actually declined as a result of an increase in MFR as parents internalize the general equilibrium effect of the increasing scarcity of girls. Therefore, the effect of MFR on girls' outcomes should be interpreted as the overall effect of sex-selective abortion (including any possible general equilibrium effect resulting from girls' scarcity). We further examine this issue in Section VI.1 below.

and over time are systematically associated with changes in the relative outcomes of females versus males. The main estimating equation is as follows:

$$(1)\, y_{ist} = \alpha_{s0} + \alpha_{s1} \, female_i + \delta_{t0} + \delta_{t1} \, female_i + x_i \, '\beta + \pi_{\,0} MFR_{\,st} + \pi_{\,1}\left(MFR_{\,st} * female_i \right) + \varepsilon_{ist}$$

where y_{ist} is the outcome of child *i* in state *s* born in year *t*, α_{s0} and α_{s1} are vectors of gender-specific state fixed effects, δ_{t0} and δ_{t1} are vectors of gender-specific year-of-birth fixed effects and x_i is a vector of individual characteristics that include indicators for twin birth, residence in an urban area, religion, mother's and father's level of education, mother's age (grouped), wealth quintiles, mass media exposure and mother's age at first birth.[17] The variable MFR_{st} is the Male-Female Ratio at birth for the cohort born in year *t* in state *s*.[18] In all the specifications, we use the national sampling weights provided in the surveys and cluster standard errors by state.

The parameter of interest is π_1, which captures the effect of prenatal sex selection on females' outcomes relative to males within the same state and birth cohort. This approach is essentially a continuous version of a *triple-difference* estimation strategy. This strategy has the advantage that it allows us to control for state-level fixed factors that differentially affect boys and girls (for example, the degree of discrimination against girls in a state). We can also control for state-time-varying factors that affect boys and girls similarly within each state and cohort and which might be correlated with changes in MFR over time (for example, improvement in access to health facilities). In addition, we control for differential trends in boys' and girls' outcomes at the national level in a very flexible way by including gender-specific year- of-birth fixed effects. Identification relies on the assumption that changes in MFR within a state over time are unrelated to other factors that could *differentially* affect male and female outcomes. We assess the plausibility of this assumption in the next subsection, where we also consider more general specifications that allow for gender-specific coefficients in all covariates and control for time-varying state characteristics, thus allowing them to have a differential effect by gender.

We first estimate equation (1) using a sample of children born at all parities. We also estimate the same equation while stratifying the sample by parity. Note that estimates based on a sample of children at all parities (as opposed to estimates that condition on parity) are less susceptible to selection bias

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[17] The index for media exposure is defined by the sum of indicators for exposure to TV, radio and newspapers or magazines. Each indicator receives a value of one if the mother reported exposure of at least once a week or almost every day. The wealth index is a constructed index provided in the NFHS data. The index is based on household assets and housing characteristics. The wealth index denotes the wealth quintile of the household relative to all households in the same survey round.

[18] We use a smoothed version of MFR which is computed as a 7-year moving average of the ratio of the number of male births to female births by year and state based on the pooled data of the three rounds of the NFHS survey.

generated by a decreasing trend in family size or to bias from a decline in family size generated by sex-selective abortion. Nonetheless, it is interesting to examine differential effects by parity given that the extent of sex-selective abortion varies by birth order and it is possible that girls are treated differently at different parities.[19]

We also employ an alternative strategy by estimating a discrete version of the triple differences model. As discussed above, northern and western states have a long tradition of strong preference for boys and have shown increasing use of prenatal sex-selection technologies. In contrast, the incidence of prenatal sex selection in other states has been relatively low.[20] We therefore analyze the differential change in girls' and boys' outcomes in northern and western states relative to other states over the three survey rounds by estimating the following equation:

$$(2)\ y_{is\tau} = \alpha_{s0} + \alpha_{s1}\, female_i + \delta_{\tau 0} + \delta_{\tau 1}\, female_i + x_i'\beta + \gamma_{\tau 0} Treated_s + \gamma_{\tau 1}(Treated_s * female_i) + \varepsilon_{is\tau}$$

where $y_{is\tau}$ is the outcome of child *i* in state *s* and in survey round τ , α_{s0} and α_{s1} are vectors of gender-specific state fixed effects, $\delta_{\tau 0}$ and $\delta_{\tau 1}$ are vectors of gender-specific survey-round fixed effects, x_i is a vector of individual characteristics that includes the same covariates as in equation (1) and "Treated" is an indicator that equals 1 if child *i* was born in a state with a high incidence of sex-selective abortion and 0 otherwise.

The parameters of interest are $\gamma_{\tau 1}\ (\tau = 1998; 2005)$ which denote the difference in outcomes of girls versus boys in 1998 and in 2005 (the years of the second and third rounds of the NFHS survey) relative to 1992 (the year of the first round of the NFHS) in states with an upward trend in sex-selective abortion relative to states in which sex-selective abortion is rare and has not increased over time.

The continuous version of our estimation strategy, in which MFR at birth is used as a regressor (equation 1), has the advantage of capturing the fact that the diffusion of pre-natal sex selection technologies is a gradual rather than a discrete change. On the other hand, the discrete version, in which states are divided into treated and comparison groups (equation 2) is more robust to possible measurement errors in MFR. The comparison of the results obtained from the two strategies provides a useful way of assessing the robustness of our findings.

[19] Kishor and Gupta (2009) show that MFR increased over time at parities two, three and four and, to a lesser extent, at parity one, while it seems to have remained relatively stable and at normal levels at parities higher than four. For evidence on differential treatment of girls by parity, see Mishra et al. (2004).
[20] In the next section we provide further evidence for sex-selective abortion in these states.

Note that we do not include household fixed effects in the estimating equations. This is due to several reasons: First, identification in a household fixed-effect model would rely on a comparison between siblings of the opposite sex. However, households with children of opposite sex cannot be viewed as a randomly selected sample since the sex composition of children is affected by parental stopping rules in fertility behavior and sex-selective abortion. In addition, given that our main sample includes children born within three years prior to the survey date, a household fixed-effects model would generate estimates for a selected sample of households (i.e. those which had two births within the last three years). Finally, as discussed in Section 3, parental preference for boys could result in gender differences in outcomes without necessarily implying differential treatment within a household. We are therefore interested in the well-being of girls relative to boys across all families and not necessarily within the same household.

VI. EMPIRICAL RESULTS

1. MAIN RESULTS ON NUTRITIONAL OUTCOMES

Gender discrimination within households can be analyzed across two dimensions: the allocation of inputs such as material resources, time, health care and feeding practices or children's outcomes such as nutrition, morbidity and mortality. We focus our analysis on children's nutrition and mortality since they embed information on various types of parental input. These outcomes are measured more easily than most others and are therefore less likely to suffer from measurement error or recall bias. We do not examine morbidity since it is more likely to suffer from bias due to gender differences in diagnosis and recall.[21]

The main analysis focuses on children's nutritional status as measured by anthropometric indicators based on height, weight and age. In particular, we consider three indicators of malnutrition: stunting, underweight and wasting. All three are defined based on z-scores, which are computed by subtracting the median and dividing by the standard deviation of a reference population of the same age and gender. Specifically, a child is considered stunted if his/her height-for-age z-score is at least 2 standard deviations below the median of the reference population. An underweight child has a weight-for-age z-score at least 2 standard deviations below the median and a wasted child has a weight-for-height z-score at least 2 standard deviations below the median.

[21] For example, Timaeous et al. (1998) report sex bias in reporting and recalling episodes of disease and sex differentials in childhood risks of illness and illnesses severity perception.

The three indicators capture malnutrition from different perspectives. Thus, stunting reflects long-term malnutrition or cumulative nutrition from conception and is also affected by recurrent or chronic illnesses. Wasting measures acute malnutrition and represents the failure to receive adequate nutrition in the period immediately preceding the survey and may be the result of inadequate food intake or a recent episode of illness leading to weight loss. An important feature of the wasting indicator is that it does not depend on the accuracy of age reporting. On the other hand, it is more sensitive to seasonal shocks. Underweight is a composite index of chronic or acute malnutrition. Note that z-scores are normalized by gender and age and therefore it is appropriate to examine the gender gap in these outcomes even if boys and girls follow different growth trajectories (unless the reference (well-fed) population is different).[22] Our analysis uses z-scores based on the US National Center for Health Statistics (NCHS) standard, which was the most commonly used measure until 2006.[23]

In order to maintain consistency across survey rounds, we limit the sample to children aged 0-35 months, born in the last two births to ever-married women aged 15-49.[24] In addition, we only include biological children of survey respondents since anthropometric data for other children were only collected in the third round of the survey.[25] About 18 percent of the children that satisfied the sample restrictions have missing values in at least one of the anthropometric indicators. Nevertheless, we do not find any significant gender differences in the likelihood of having a missing value in these indicators. Moreover, we do not find any associations between state variation in MFR and the likelihood of missing anthropometric data for girls or boys.

The literature has been inconclusive on whether there is a female disadvantage in nutritional outcomes of children in India.[26] The lack of evidence for female disadvantage in children's nutritional

[22] In an alternative specification, we included age-in-months and its interaction with *female* (instead of child's year-of-birth) in order to control for any differences in child outcomes that could result from differences in the age distribution of boys and girls. The results are virtually identical to those presented here.

[23] A new international reference population was published by the World Health Organization in 2006. The new standards are based on properly-fed children with no significant morbidity in Brazil, Ghana, India, Norway, Oman and the United States and use the breastfed child as the normative model for growth and development. As we show in the test of robustness subsection, our results are not sensitive to the change in z-scores.

[24] The main results remain similar when we remove the age restriction (results not reported here but available upon request).

[25] Height was not measured in the first round of the NFHS in Andhra Pradesh, Himachal Pradesh, Madhya Pradesh, Tamil Nadu and West Bengal because height measuring boards were not available at that time (IIPS, 1995). Given that our results for weight are highly consistent with the results for height, we believe that the lack of height data for these five states is unlikely to bias the main results.

[26] Mishra et al. (1999) find that stunting and underweight levels are similar for boys and girls while wasting appears to be somewhat higher among boys. Pande (2003) and Mishra et al. (2004) find different patterns across outcomes (stunting and underweight versus wasting) and different patterns by birth order and sibling sex composition. Barcellos et al. (2010) find that the gender gap in children's nutritional status is sensitive to the alternative

status, despite considerable evidence for son preference, is consistent with findings from several large-scale cross-country studies (see, for example, Hill and Upchurch, 1995; Arnold, 1992; Sommerfelt and Arnold, 1998; and Marcoux, 2002). One of the main challenges in testing for gender discrimination by comparing nutritional outcomes or mortality rates between boys and girls is that the outcomes are also affected by differences in biological processes, genetic endowments and illnesses that might differentially affect the genders. Rosenzweig and Schultz (1982) argue that since boys and girls may need different levels of inputs in order to achieve the same survival rates, it is not possible to infer discriminatory allocation of resources within the household by comparing gender differentials in survival rates. However, it is still possible to infer differences in resource allocation by comparing *changes* in the relative outcomes of boys and girls that result from changes in the economic environment. Following this approach, they find that survival rates of girls relative to boys are related to the variation in their expected earnings opportunities as adults. Using a similar approach, Rose (1999) finds that survival rates of girls relative to boys increase for cohorts that experienced favorable rainfall shocks during the first two years of life. Our triple-differences strategy follows the empirical approach used in these studies by enabling us to derive changes in resource allocation for boys and girls by comparing changes in relative outcomes across two different dimensions: time and region.

An additional methodological challenge in the analysis of gender discrimination is the availability of anthropometric data only for surviving children. Therefore, if sex-selective abortion has any impact on mortality, our analysis of nutrition will be based on a selective sample. For example, if increases in MFR are associated with a reduction in female child mortality, they might lead to an increase in the proportion of girls who are close to a survival threshold, thus attenuating the estimated effect of MFR on nutritional status. As shown below, the effects on mortality are negligible or not significant, which mitigates the concern about selectivity issues in the nutritional status results.

Table 5 reports the main results for children's nutritional status. As seen in column 2, a little less than half of the children are considered underweight or stunted (49 and 43 percent, respectively) and on average, girls are slightly more disadvantaged than boys. The gender gap for both measures is small

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reference charts used for the standardization of z-scores. A similar conclusion is reached by Tarozzi (2008). While evidence on gender differential in children's nutritional status is mixed, recent studies have found evidence of female disadvantage in longer term anthropometric outcomes. For example, Song and Burgard (2008) present evidence of female disadvantage in growth trajectories from early childhood through adolescence in China (a society with strong son preference) relative to Philippines (a society without a strong tradition of son preference) using longitudinal data. Marcoux (2002) reports evidence of a higher incidence of female malnutrition among adults in various countries and Deaton (2008) finds a higher degree of sexual dimorphism in height among adults living in Indian states with higher MFR.

(about 1.6 and 1.4 percentage points, respectively) though statistically significant. Wasting has a lower incidence (about 18 percent of the sample) and is slightly more prevalent among boys than girls. The gender gap favoring boys in this case is about 1.5 percentage points. Overall, and similar to previous studies, we do not find marked gender differences in the mean nutritional outcomes of children.

Columns 4-7 of **Table 5** report estimates for *MFR* and *female*MFR* from a linear probability model for the likelihood of being underweight, wasted, or stunted.[27] Estimates from a basic model that includes no covariates, except for a female dummy, and gender-specific state and year-of-birth fixed effects are reported in columns 4 and 5. Columns 6 and 7 report estimates for the full model specified in equation (1) which controls also for the household characteristics reported in Table 1. The key parameter of interest is the coefficient of the interaction term *female*MFR*, which is negative for all three outcomes and, except for stunting, is statistically significant (column 7). Note that the coefficients for the main effect of *MFR* are statistically insignificant (column 6), suggesting that changes in MFR are not associated with changes in the nutritional outcomes of boys. This last finding is important since it suggests that increases in MFR at the state level are unlikely to be associated with other changes in unobservables that affect children's nutritional status.

Taken together, the results suggest that girls' nutritional status improved relative to that of boys in regions where the incidence of sex-selective abortion (as proxied by MFR) increased. In other words, an increase in the practice of sex-selective abortion appears to be associated with a reduction in the incidence of malnutrition among surviving girls relative to boys.

Note that the key *female*MFR* coefficient reflects the change in girls' outcomes relative to those of boys that results from a one-unit increase in MFR. In order to put the estimated magnitude into perspective, the estimate reported in the first row of column 7 suggests that a 20-point increase in MFR (which is the increase observed in Punjab between the first and the third round of the NFHS) is associated with a 10 percentage point reduction in the proportion of girls who are underweight. Alternatively, we can say that an increase of one standard deviation in MFR (7 points) is associated with a 1.8 percentage point reduction in the proportion of girls who are wasted.

Previous studies of gender discrimination show differential patterns for urban and rural populations and usually find a higher incidence of discrimination against girls in rural areas (see, for example, Rosenzweig and Shultz, 1982; Simmons, 1982; Subramaniam and Deaton, 1991; and Deaton, 1997). Following this approach, we also look at the differential effects of sex-selective abortion on girls'

[27] Marginal effects from logit models provide similar results.

outcomes in rural and urban areas.[28] We stratify the sample by place of residence (urban/rural) and then estimate equation (1) separately for the two samples.[29]

Estimates for the rural/urban stratification are reported in Panels A and B of **Table 6.** As expected, column 5 shows that children are more likely to be malnourished in rural as opposed to urban areas. Interestingly, there are no marked differences in the gender gap in nutritional status (column 3). Estimates for *MFR* and its interaction with *female* as reported in columns 4 and 5 show that improvements in girls' nutritional status relative to that of boys associated with increases in MFR are substantially larger in rural areas than in urban areas. These results suggest that the improvement of girls' nutritional status following the appearance of sex-selective abortion was primarily concentrated in rural areas.[30] The larger improvement in girls' outcomes in rural areas is consistent with the literature in showing that gender discrimination in resource allocation is more prevalent among more constrained families.

We also stratify the analysis according to parity since our earlier results show that sex-selective abortions are more likely to occur at higher parities. Panels C and D of **Table 6** report estimates of the effect of *MFR* and its interaction with *female* for the sample of children born at parity one and the sample of children born at parity two and above. The results provide some evidence for the stronger effect of MFR in improving girls' nutritional status at parities higher than one (in particular with respect to wasting). This is consistent with the fact that sex-selective abortion is more prevalent at parities higher than one and therefore is likely to reduce the proportion of unwanted girls at these parities.

Robustness Checks

We performed additional tests for the robustness of the main results reported in Table 5. The results of these tests are reported in Appendix **Table A1**. In order to facilitate comparison, we reproduce the estimates obtained for our main specification in columns 1 and 2 of the table. Overall, the additional estimates are qualitatively similar across various models and specifications.

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[28] Note that our earlier analysis showed that while sex-selective abortion was more prevalent in urban areas, there is an increasing trend in the use of prenatal sex-selection in rural regions.

[29] We measure MFR at the state level since there are insufficient observations in the NFHS data to compute separate sex ratios for the urban and rural samples within each state for each birth cohort. Evidence from census data suggests a high correlation between MFR in rural and urban areas within each state. For example, the correlation is about 0.7 for the 1991 census and 0.9 for the 2001 census.

[30] Note also that since most (about 75%) of the population lives in rural areas in a majority of states, our estimates for the rural sample are more precise thanks to the larger sample size and better proxy for MFR which uses state-level data.

First, we assess the validity of our main identifying assumption that there are no changes in unobserved factors associated with MFR that could affect girls' outcomes. Note that this concern is partially addressed by the triple-difference strategy which controls for state-level time-varying factors that affect boys and girls similarly.[31] Still, there may be state-level time-varying factors that differentially affect boys and girls. For example, it might be the case that increases in a state's MFR are related to more rapid economic development and modernization, which in turn may affect girls' and boys' health outcomes differentially (for example, by providing better health care access or by means of a reduction in discrimination against girls). Therefore, we estimate expanded versions of equation (1) in which household and time-varying state characteristics and their interactions with gender are also controlled for. Specifically, we estimate three models with alternative sets of additional controls. In the first set, we add interactions between household covariates and a female dummy to the basic model. In the second and third specifications, we include, in addition to gender-specific controls for household characteristics, two alternative sets of state-level time-varying covariates interacted with gender. The first set of state variables is compiled from the NFHS data and includes state means of the wealth index and the proportion of households in urban areas, proportion with electricity access, and proportion with access to TV (watch TV at least once a week). The alternative set contains state information on GDP growth (compiled from reports of India's central bank -- the Reserve Bank of India), average wage income, the rates of poverty, literacy and employment and the share of employment in agriculture (from the Indian Socio-Economic Survey downloaded from IPUMS International).

The results for the three expanded models are reported in columns 3-6 of **Table A1,** and, in general, exhibit similar patterns. For example, the coefficients of female*MFR in the underweight regression for the three expanded models are -0.514 (s.e.=0.126), -0.433 (s.e.=0.115) and -0.344 (0.169), respectively. Overall, the estimates confirm that our main results are largely robust to other time-varying changes at the state level that could differentially affect boys' and girls' outcomes.

We also tested the sensitivity of our results to the specific definition of the anthropometric indicators. As noted above, in 2006 the World Health Organization published new growth standards for assessing children's nutritional status. Therefore, we re-defined the anthropometric indicators using the new WHO tables and re-estimated the main specification. As a result, the proportion of underweight children slightly decreases when using the new standards (0.440 versus 0.485) while the proportion of

[31] We also estimated an alternative specification in which we control for state fixed effects interacted with year of birth (instead of including MFR). The coefficients of the interaction between *female* and *MFR* were virtually unchanged.

children who are wasted or stunted slightly increases (0.221 versus 0.179 and 0.499 versus 0.433, respectively). More importantly, we observe that the gender gap in nutritional status is reversed with girls being slightly less likely to be underweight, wasted or stunted relative to boys. This last finding emphasizes the importance of examining changes rather than levels when comparing between the nutrition or health status of boys versus girls. Columns 9 and 10 in **Table A1** show that, despite the nontrivial changes in the definition of outcomes, the coefficient of *MFR* and its interaction with *female* remain virtually unchanged.

We also assessed the sensitivity of our main results to the specific definition of the main explanatory variable. Note that our basic specification includes MFR in a linear fashion although it may be the case that the relationships between MFR and the outcomes of interest are nonlinear. We therefore re-estimated equation (1) after replacing MFR with its natural logarithm. The estimates reported in columns 1 and 2 of the lower panel of **Table A1** show that the estimated effects when MFR is replaced with its logarithm are of comparable magnitude. For example, the coefficient on female*ln(MFR) in the underweight regression is -0.609 (s.e.=0.147) which implies that a 20-point increase in MFR from 1.05 to 1.25 (or 17%) is associated with a decrease of 10 percentage points (17*0.609) in the proportion of underweight girls, which is similar to the results reported in Table 5.

We also experimented with different ways of smoothing MFR. For example, we considered both a shorter and a longer window for computing the moving average. The results, which are reported in columns 3-6 in the lower panel of the table, are basically similar to those obtained from our main specification. For example, the coefficients on *female*MFR* in the underweight regressions are -0.426 and -0.591 when MFR is computed using a 5-year and 9-year MA, respectively (as compared to -0.536 in Table 5).

Finally, we also looked at the sensitivity of our results to the weighting scheme of the data by re-estimating our model using unweighted data (columns 7 and 8 in the lower panel) and using state weights (columns 9 and 10 in the lower panel). All the estimates are quite similar to the main results.

Regional Patterns

In this section, we apply an alternative estimation strategy which involves a discrete version of our triple difference-in-differences model. The basic idea is to abstract from the specific cohort measure of MFR and examine changes in the outcomes of females versus males over the three survey rounds in states that have experienced an upward trend in the practice of sex-selective abortion relative to states where the practice of sex-selective abortion is relatively rare.

As described in Section V, states with an upward trend in sex ratios at birth are clustered in the northern and western regions. We therefore identify the following eight states as *"treated"*: Gujarat, Haryana, Himachal Pradesh, Maharashtra, Punjab, Rajasthan, Jammu and Kashmir and Delhi. This list coincides with the classification of Bhat (2002) who examines the dynamics of sex ratios in India and adds the states of Jammu and Kashmir and Rajasthan to the classification proposed by Retherford and Roy (2003) (who identify those states that have a high incidence of sex-selective abortion by examining trends in sex ratios at birth and differential trends by parity and sex composition of previous children).[32]

Appendix Table A2 presents data for the incidence of sex-selective abortion in treated states and in all other states in the sample (18 in total). This is done by computing the differential probability of a male birth at parities 2 or 3 as a function of the sex composition of previous children in the pre- and post-ultrasound periods while applying the same strategy used for Table 3. The results show that, prior to the 1990s, the likelihood of a male birth at parity 2 in treated states is not associated with the gender of the older sibling. At parity 3, we already observe some positive association for households with two girls in treated states although the estimate is relatively small (2.7 percentage points). The differential probability of a male birth among households with one or two older girls increases considerably during the 1990s in treated states and is highly significant. Estimates for parity two show that households in treated states that have a girl at parity one, are almost four percentage points more likely to have a boy at parity two relative to households that have a boy at parity one. For parity three, we find that the likelihood of having a boy is 7.9 percentage points higher among households with two girls and 5.3 points higher among households with a girl and a boy relative to households with two boys.

In sharp contrast to the pattern observed for treated states, we find no statistically significant differences in the likelihood of a male birth at parity two or three according to the sex composition of previous children among households in the other states during the post-ultrasound period.

Overall, the evidence reported in Appendix Table A2 points to an increasing trend in the use of sex-selective abortion in the treated states as opposed to all remaining states, in which the practice of sex-selective abortion appears to be less common.[33]

Table 7 reports the results of the estimation of equation (2) which compares changes in the nutritional status of girls versus that of boys in treated versus other states over the three survey rounds. The parameters of interest are the coefficients on the triple interaction terms, *female*round2*treated*

[32] Retherford and Roy (2003) focused on 17 major Indian states and did not examine patterns of sex-selective abortion in Jammu and Kashmir.

[33] We do not claim that sex-selective abortion is not practiced at all in other states but rather that its effects are expected to be smaller relative to the treated group.

and *female*round3*treated*. Both coefficients are negative for the three nutritional status indicators and their magnitude is larger (in absolute terms) in the third round relative to the second round (see columns 5 and 7), which is consistent with the upward trend in MFR over the three survey rounds. For example, estimates for underweight (first row) suggest that girls' likelihood of being underweight decreased about 4 percentage points more in states with a higher incidence of sex-selective abortion relative to other states between the first and the second round of the NFHS. The reduction observed in the third round relative to the first is about 6 percentage points. Note that sex ratios at birth increased in treated states from 1.093 to 1.151 between the first and the third round, while in other states, they increased only slightly from 1.064 to 1.067.

Similar to the results reported in Table 5, the coefficients for the two interaction terms, *round2*treated* and *round3*treated*, are small, not significant and have inconsistent signs over the survey rounds and outcomes, thus suggesting that there were no major changes over time in the nutritional status of boys in treated relative to other states.

Taken together, our results suggest that girls' nutritional status improved to a greater extent than that of boys in states with a high incidence of sex-selective abortion.

2. MECHANISMS

The relative improvement in girls' nutritional status may have resulted from improved parental treatment of girls who are born (for example, substitution of prenatal discrimination for postnatal discrimination). It could also be the result of selection, such that girls are born into families with better endowments. An additional channel that could explain the results is family size, which may have declined due to lessened reliance on stopping rules in fertility behavior. In an attempt to explore the mechanisms for the relative improvement of girls' nutritional status, we performed some additional analyses.

2.a. Are girls increasingly born into better-endowed families?

In order to examine the selection channel, we examine whether family characteristics of girls (such as parental education, mother's age at first birth and household wealth status) improved relative to boys in states where MFR has increased. This is done by estimating a triple-differences equation in which we regress each of the family characteristics on *MFR* and *female*MFR* in a model that controls for state fixed effects, survey-year fixed effects and their interactions with a female indicator.

Table 8 reports estimates for the main MFR effect and the interaction between *MFR* and *female*. The table also reports gender differences in household characteristics (column 3). Estimates of the main MFR effect reported in column 4 (rows 1-6) suggest that regions with increasing MFR experienced improvement in some family characteristics, in particular, an increase in the level of parental education and mother's age at first birth and a decline in the likelihood of living in a rural area. This is consistent with the fact that the practice of sex-selective abortion is related to economic development and urbanization. On the other hand, there is no association between MFR and maternal age or the household wealth index.

Estimates for the interaction between *MFR* and *female* reported in column 5 (rows 1-6) are all small, not significant and have inconsistent signs across the different variables. Overall, these results suggest that there is no differential improvement in household characteristics among families with girls relative to families with boys and therefore the improvement in girls' nutritional status relative to that of boys is unlikely to be explained by the fact that girls are born into *"better"* families.

The positive association between MFR and some household characteristics raises some concerns regarding our main results for girls' nutritional outcomes if an improvement in household characteristics has a differential effect on the nutritional or health status of boys and girls. Nevertheless, as reported in Table A1 and discussed in the previous section, our main results were virtually identical to those obtained from an enriched model in which we allow for a differential effect of household characteristics by gender. This suggests that the improvement in parental education or other household characteristics did not have a differential effect on the nutritional status of girls relative to boys.

2.b. Are girls increasingly being born into smaller families?

In this subsection will examine the family size channel. As noted above, as a direct consequence of parental stopping rules in fertility behavior, girls are more likely to be born in larger families. With access to sex-selective abortion, parents of girls would not necessarily need to have more children in order to attain the desired sex composition. As a result, girls might fare better due to an increase in family resources per child. We therefore examine whether family size has differentially changed for girls relative to boys in regions with increasing MFR. To this end, we regress family size on *MFR* and *female*MFR*. The model also controls for state fixed effects, year fixed effects and their interactions with a female indicator.

The last row of Table 8 reports the coefficients of *MFR* and *MFR* interacted with *female*. The estimate reported in column 4 indicates that MFR is negatively associated with family size which may be due to various factors. One possibility is reverse causation such that a desire to have fewer children may increase the demand for sons at lower parities, thus increasing the incidence of sex-selective abortion. A second possible channel may be related to unobserved factors that affect both MFR and fertility. For example, economic development is generally associated with a reduction in fertility and an increase in access to ultrasound technology. A third channel may arise due to a causal link from access to sex-selective abortion to family size. Parents with access to sex-selective abortion are more likely to attain their preferred sex composition of children without the need of having additional children.[34] The first two channels should have a similar effect on boys and girls while the last should show a differential effect by gender with larger reductions for girls. As seen in column 5, the estimate for the interaction between *female* and *MFR* is indeed negative (-0.923) and marginally significant (s.e.=0.534) suggesting that reductions in family size were larger for girls relative to boys in regions with upward trends in the incidence of sex-selective abortion.

2.c. Are girls increasingly receiving better treatment from parents?

As mentioned above, an improvement in girls' nutritional status may be a direct consequence of better care and treatment in the family since girls are more likely to be wanted when prenatal sex-selection is available. Parental treatment is difficult to assess since we do not directly observe household investment. As a partial solution, we use breastfeeding practices as a proxy. Medical and public health research has found that breastfeeding has large benefits for children's health, especially in environments with poor sanitary conditions (see, for example, The World Health Organization, 2000). The WHO recommends that breastfeeding in developing countries should continue, in combination with supplementary foods, up to the child's second birthday or beyond (WHO, 1991).

Although breastfeeding is nearly universal in India, girls tend to be breastfed for a shorter period than boys. This in shown in **Figure 3** which plots Kaplan-Meyer survival functions for breastfeeding duration for boys and girls. As can be seen, the likelihood of being breastfed does not differ by gender during the first year of life but girls are less likely to continue being breastfed after 12 months of age.

[34] On the other hand, the ability to choose a child's gender increases the expected utility of children and might therefore increase the likelihood of having additional children. Ben-Porath and Welch (1976) also suggest that if the net cost of raising boys is lower than the cost of raising girls, family size might increase if parents are able to choose the gender of their children. Furthermore, even in the absence of cost differences, family size might also increase if boys yield higher utility than girls and there is substitution between utility from children and ordinary consumption or between utility from boys and utility from girls. As a result, family size might increase or decrease after the appearance of sex-selective abortion. Nevertheless, differences in family size by gender are expected to narrow.

Parents with son preference may breastfeed their sons for a longer period than their daughters either actively through discriminatory treatment or passively through stopping rules in fertility behavior (Jayachandran and Kuziemko, 2010).[35]

We examine the effect of sex-selective abortion on gender disparity in breastfeeding by estimating equation (1) with the dependent variables being indicators of whether the child was breastfed for at least 12, 18 or 24 months. These indicators were chosen in order to account for heaping of observations at six–month intervals (either due to rounding error in duration reporting or actual propensity to breastfeed up to a focal point). In order to take into account the possibility of right-censoring in duration, the indicators are defined conditional on children being at least 12, 18 or 24 months old, respectively, at the time of the survey.

We estimate the model using our main sample (the last two children born within 3 years prior to each survey round) as well as for samples stratified by rural/urban residence and parity (i.e., the same samples used for the nutritional analysis). Results are reported in **Table 9.** Consistent with the pattern shown in Figure 3, the differences in outcome means reported in column 3 show a disadvantage for girls in breastfeeding duration that widens with age. For example, girls are one percentage point less likely than boys to be breastfed for at least 12 months, about 4 percentage points less likely to be breastfed for at least 18 months and almost 6 percentage points less likely to be breastfed for at least 24 months. Girls' disadvantage in breastfeeding duration is larger in rural areas and in parities higher than one (Panels B and E).

Estimates for the interaction between *MFR* and *female* suggest a larger reduction in girls' breastfeeding disadvantage at long durations (24 months or longer) in states with higher increases in MFR. The differential effect is stronger in rural areas and at higher parities. These results are broadly consistent with the findings from Tables 5 and 6, which suggests that breastfeeding practices may be another channel (in addition to family size) through which sex-selective abortion reduces gender disparity in nutritional outcomes among girls. Consistent with the results for nutritional status, we find a larger improvement in rural areas and at parities larger than one.

2.d. Are girls benefiting from declining son preference?

Girls may be treated better if sex-selection technology allows parents to substitute prenatal discrimination for postnatal discrimination. It is also possible that girls are treated better when their

[35] Jayachandran and Kuziemko (2010) consider a dynamic model for the breastfeeding decision by parents with son preference and use NFHS data to empirically test some of the model's predictions. They do not, however, examine the effect of sex-selective abortion on breastfeeding duration.

parents internalize the change in the value of women as they become a scarce commodity, thus reducing their preference for sons. In order to test this channel, we examine whether increases in MFR are associated with a decline in the preference for boys at the state level. We focus on two measures of son preference: the ratio of the ideal number of sons to the ideal number of daughters and the proportion of women who want more sons than daughters. Note that we cannot establish a causal relationship between sex ratios and son preference since the link between the two can go in both directions. In addition, there may be other unobserved factors that can affect both simultaneously. Nevertheless, it may still be of interest to examine the association between MFR and son preference.

We collapse the data by state and survey year, thus limiting the sample to mothers whose youngest child was born within 3 years prior to the survey and estimate models in which the indicators of son preference are regressed on MFR. The results are reported in Appendix **Table A3**. Simple bivariate regressions show no association between son preference and MFR (column 2). When controlling for state fixed effects (column 3), we observe that states with increasing MFR are associated with declining son preference. Note, however, that these associations may be due to common factors related to economic development and modernization which affect both parental preferences and MFR and do not necessarily imply a causal link between MFR and son preference. In fact, when we add controls for household characteristics by state and survey year, the associations between changes in son preference and changes in MFR fade away (column 4).[36] Overall, while it is hard to provide compelling evidence for the link between changes in MFR and changes in son preference, the observed associations between these two factors do not seem to point to a larger decline in son preference in states that have experienced an increase in MFR.[37]

3. MORTALITY

Gender preferences and discrimination may also affect child mortality. Indeed, a large literature has documented excess female mortality which becomes evident after the age of one month.

In this section, we examine the effect of sex-selective abortion on gender disparity in mortality. We focus on the following set of binary outcomes: neonatal mortality (death under 1 month of age), post-neonatal mortality (death between 1 month and 12 months of age, conditional on surviving up to 1 month), child mortality (death between 1 and 5 years of age, conditional on surviving up to 1 year) and

[36] The controls include the state-year means of the following background characteristics: index of media exposure, religion, mother's education, age at first birth and indicators for urban/rural residence and wealth.

[37] In fact, Anderson (2003) shows that in caste societies, economic modernization might actually lead to dowry inflation, thus reducing (rather than increasing) the economic value of girls.

under-five mortality (death before 5 years of age).[38] For each outcome, we estimate a linear probability model with a specification similar to that of equation (1).[39] We consider two samples: the first includes the last two children born within 3 years prior to each survey and parallels the sample selection criteria used for the nutritional outcomes and the second includes all children born within 10 years prior to each survey.

The results are reported in **Table 10**. Consistent with the findings of previous studies, we find excess female mortality after the age of one month. Nevertheless, we do not find a larger reduction in female disadvantage in child mortality in states with a high incidence of sex-selective abortion. The coefficients on the key interaction term *MFR*female* (reported in columns 5 and 8) are roughly zero in all outcomes and samples and none of them is statistically significant.

The results for child mortality are consistent with Shepherd (2008) which reports inconclusive evidence of sex-selective abortion on child mortality with results varying by specific mortality measures and time.[40] On the other hand, the results are somewhat at odds with our previous results on nutritional outcomes, family size and breastfeeding duration. Particularly puzzling is why we find an improvement in female nutritional status but do not see any reduction in excess female mortality. One possible explanation is that families which substitute between prenatal and postnatal female discrimination are not the same families characterized by excess female child mortality. A second possible explanation is that differential recall (by gender) of deaths and measurement error may be biasing the estimated MFR effects towards zero. The problem may be more severe with longer recall periods as is the case for the second sample which includes all children born within 10 years prior to each survey. By comparison, the analysis of nutrition outcomes of living children (born within 3 years prior to the survey) is based on anthropometric measures observed at the time of the survey and is thus less likely to be affected by these biases. Indeed, the estimates for the interaction between *MFR* and *female* for the mortality outcomes using the sample of younger children are negative although not precise enough to be statistically significant.

A third possible explanation is that while the nutritional status of surviving girls has improved the magnitude of the improvement is still not large enough to reduce the likelihood of death for marginal girls. Medical research suggests that the leading contributors to child mortality are respiratory ailments

[38] In order to deal with possible right-censoring in duration, the indicators are only defined for children who would be "old enough" (for example 1 month, 1 year and 5 years old, respectively) at the time of the survey.
[39] Results from logit models (not reported here) provide similar findings.
[40] We did some further analysis by stratifying the sample according to place of residence (rural/urban) and parity, but did not find any consistent evidence for reductions in female mortality.

and infectious and gastro-intestinal diseases. Malnutrition, while often underlying and exacerbating these diseases, is not by itself a fatal factor, except in severe or extreme cases.[41] It is also possible that other types of parental investment, such as preventative care (for example vaccinations), have a more important impact on reducing diseases and thus mortality.

VII. CONCLUSION

In this paper, we study the impact of prenatal sex selection on the well-being of girls in India. To the extent that prenatal sex selection constitutes a substitute for postnatal gender discrimination, the situation of girls born after sex determination technologies became available might improve. Prenatal sex selection may also affect girls' relative well-being through a differential reduction in family size or a selection of girls born into families with higher SES.

We explore these issues using data from the National Family Household Survey. We proxy parental access to prenatal sex selection using the ratio of male to female births in the year and state in which the child was born and provide evidence showing that starting in the 1990s high sex ratios at birth reflect the practice of sex-selective abortion. We then analyze whether changes in girls' outcomes relative to boys within states and over time are associated with changes in sex ratios at birth.

We find that an increase in the practice of prenatal sex selection appears to be associated with a reduction in the incidence of malnutrition among surviving girls. This negative association is stronger for girls born in rural households and at higher birth parities. We find no evidence that prenatal sex selection leads to selection of girls into families of higher SES. We do find some evidence of a larger reduction in family size for girls than for boys and we also find some suggestive evidence of better treatment of girls as reflected in breastfeeding duration. On the other hand, prenatal sex selections do not appear to be associated with a reduction in excess female child mortality.

[41] Our indicators of malnutrition (i.e. anthropometric measures that are two standard deviations below the median of the reference population) are often thought to reflect mild cases. Indeed, when examining comparable samples (children aged 0-35 months), we find only a moderate correlation between the malnutrition indicators and mortality at the state level in each of the survey rounds. Our results are consistent with those of Hill and Upchurch (1995) who report a lack of association between female mortality disadvantage and nutritional disadvantage across several developing countries.

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Table 1. Summary Statistics

	Nutritional status sample			Children's sample		
	Girls	Boys	Difference	Girls	Boys	Difference
	(1)	(2)	(3)	(4)	(5)	(6)
Urban	0.236	0.237	-0.002	0.232	0.235	-0.003
			(0.003)			(0.002)
Index of mass media exposure	0.799	0.824	-0.025	0.738	0.755	-0.017
			(0.010)			(0.005)
Wealth index	2.82	2.86	-0.041	2.74	2.77	-0.030
			(0.009)			(0.009)
Mother's Age	25.2	25.3	-0.098	28.7	28.7	-0.047
			(0.024)			(0.019)
Mother's age at 1st birth	19.0	19.0	0.000	18.7	18.7	-0.023
			(0.025)			(0.013)
Mother's education						
No education	0.541	0.538	0.004	0.613	0.607	0.006
			(0.004)			(0.003)
Primary school	0.158	0.148	0.009	0.146	0.147	-0.001
			(0.003)			(0.002)
Secondary school	0.247	0.256	-0.010	0.200	0.203	-0.004
			(0.005)			(0.002)
Higher	0.053	0.057	-0.004	0.040	0.042	-0.002
			(0.002)			(0.001)
Missing	0.001	0.001	0.000	0.001	0.001	0.000
			(0.000)			(0.000)
Father's education						
No education	0.300	0.291	0.009	0.341	0.339	0.002
			(0.003)			(0.002)
Primary school	0.193	0.189	0.004	0.205	0.201	0.003
			(0.005)			(0.003)
Secondary school	0.381	0.389	-0.008	0.346	0.350	-0.004
			(0.006)			(0.002)
Higher	0.119	0.124	-0.005	0.101	0.103	-0.002
			(0.003)			(0.002)
Missing	0.006	0.007	-0.001	0.006	0.007	0.000
			(0.001)			(0.000)
Religion						
Hindu	0.792	0.792	0.001	0.791	0.794	-0.003
			(0.002)			(0.002)
Muslim	0.158	0.156	0.002	0.161	0.156	0.005
			(0.003)			(0.001)
Other religion	0.049	0.052	-0.003	0.05	0.05	-0.002
			(0.002)			(0.002)
Missing	0.001	0.001	0.000	0.0	0.0	0.000
			(0.000)			(0.000)
Number of children in the family	2.92	2.93	-0.007	3.9	3.7	0.125
			(0.009)			(0.015)
Sample size	36,940	39,560	76,500	172,472	185,337	357,809

Notes: The table reports summary statistics for boys and girls (cols. 1,2,4, and 5) included in the analysis samples and differences between the characteristics of girls and boys (cols. 3 and 6). Standard errors of the differences clustered at the state level are reported in parenthesis. The samples pool rounds 1, 2, and 3 of the NFHS. The nutritional status sample reported in columns 1-3 includes the last two children under three years of age of ever married women with valid anthropometric data. The children sample reported in columns 4-6 includes all children born within the last 10 years preceding the survey date of ever married women. Observations are weighted using national-level weights.

Table 2a. Male Female Ratios and Fertility Preferences by State

	Male-Female Ratio (MFR) at age 0					Fertility preferences			
	1961 (1)	1971 (2)	1981 (3)	1991 (4)	2001 (5)	Number of children (6)	Ideal number of children (7)	Ideal MFR (8)	Wants more sons than daughters (9)
North	**103**	**103**	**105**	**111**	**117**	**2.46**	**2.75**	**1.46**	**0.49**
Delhi	104	105	105	110	117	2.36	2.52	1.25	0.30
Haryana		104	108	115	124	2.45	2.56	1.41	0.45
Himachal Pradesh	102	104	103	108	115	2.29	2.36	1.30	0.37
Jammu & Kashmir	102	103	107	N/A	114	2.58	2.77	1.48	0.49
Punjab	103	105	106	117	129	2.46	2.57	1.46	0.48
Rajasthan	103	102	104	108	112	2.49	3.02	1.55	0.58
West	**103**	**103**	**104**	**108**	**113**	**2.23**	**2.56**	**1.29**	**0.38**
Gujarat	103	103	103	109	116	2.24	2.60	1.33	0.42
Maharashtra	103	103	105	107	111	2.22	2.54	1.27	0.36
Northeast	**98**	**102**	**102**	**104**	**104**	**2.73**	**3.33**	**1.33**	**0.40**
Arunachal Pradesh	N/A	109	100	101	103	2.55	4.67	1.41	0.43
Assam	98	101	N/A	105	105	2.74	3.17	1.38	0.44
Manipur	102	94	101	102	106	2.89	3.74	1.36	0.43
Meghalaya	N/A	106	100	101	104	2.78	4.62	1.01	0.14
Mizoram	N/A	102	N/A	99	100	2.66	4.29	1.18	0.33
Nagaland	64	101	103	102	102	2.99	4.03	1.12	0.28
Tripura	99	106	106	103	105	2.43	2.57	1.28	0.33
Sikkim	95	88	101	105	106	2.32	2.23	1.13	0.22
Central	**100**	**102**	**104**	**107**	**110**	**2.47**	**3.28**	**1.52**	**0.55**
Madhya Pradesh	101	99	101	104	106	2.30	3.12	1.44	0.52
Uttar Pradesh	100	104	105	109	112	2.55	3.36	1.55	0.57
East	**99**	**100**	**103**	**106**	**106**	**2.29**	**3.03**	**1.41**	**0.45**
Bihar	101	102	104	108	107	2.38	3.40	1.56	0.56
Orissa	97	98	102	103	106	2.23	3.01	1.36	0.45
West Bengal	99	98	103	104	104	2.19	2.58	1.25	0.31
South	**100**	**99**	**102**	**104**	**105**	**2.08**	**2.48**	**1.17**	**0.23**
Andhra Pradesh	99	98	101	103	104	1.99	2.75	1.25	0.33
Goa	105	105	105	104	106	2.34	2.69	1.20	0.28
Karnataka	101	101	102	104	106	2.30	2.53	1.20	0.27
Kerala	101	99	102	104	103	2.07	2.62	1.12	0.18
Tamil Nadu	99	99	101	103	105	2.00	2.08	1.07	0.11

Notes: Columns 1-5 report male-female ratios (MFR) at age zero by state for various census years. Columns 6-9 report indicators for fertility, desired fertility, and son preferences based on mothers' reports from the first round of the NFHS. Tabulations for Sikkim are based on the second round of the NFHS as Sikkim was not sampled in the first round. Observations are weighted using state-level weights.

Table 2b. State Characteristics

		Household characteristics			Religion			Mother's charateristics		
	Real per capita income 1991 (1990 rupees) (1)	Urban (2)	Wealth Index (3)	HH with electricity (4)	Hindu (5)	Muslim (6)	Other (7)	Avg. years of schooling (8)	TV exposure (9)	Iliterate (10)
North		**0.29**	**3.55**	**0.75**	**0.79**	**0.05**	**0.16**	**2.84**	**0.40**	**0.66**
Delhi	10,177	0.92	4.79	0.96	0.82	0.10	0.08	6.35	0.83	0.37
Haryana	7,502	0.26	3.90	0.86	0.89	0.04	0.07	3.01	0.49	0.64
Himachal Pradesh	4,790	0.10	3.62	0.92	0.97	0.01	0.02	3.62	0.47	0.50
Jammu & Kashmir	3,872	0.18	3.74	0.88	0.77	0.17	0.06	3.91	0.50	0.57
Punjab	8,373	0.28	4.26	0.94	0.38	0.01	0.61	3.88	0.57	0.53
Rajasthan	4,113	0.20	2.79	0.54	0.92	0.06	0.02	1.36	0.18	0.82
West		**0.39**	**3.56**	**0.77**	**0.81**	**0.11**	**0.08**	**3.85**	**0.44**	**0.52**
Gujarat	5,687	0.35	3.60	0.78	0.89	0.09	0.02	3.61	0.39	0.55
Maharashtra	7,316	0.42	3.54	0.76	0.76	0.13	0.11	3.97	0.47	0.50
Northeast	N/A	**0.16**	**2.69**	**0.31**	**0.61**	**0.21**	**0.18**	**3.13**	**0.22**	**0.55**
Arunachal Pradesh		0.15	3.17	0.62	0.35	0.01	0.64	2.25	0.29	0.70
Assam	4,014	0.12	2.44	0.20	0.67	0.28	0.04	2.80	0.18	0.59
Manipur	3,893	0.32	3.55	0.64	0.62	0.06	0.31	4.44	0.38	0.48
Meghalaya	N/A	0.19	3.10	0.43	0.09	0.02	0.89	3.26	0.24	0.51
Mizoram	N/A	0.49	3.82	0.76	0.02	0.00	0.98	5.69	0.25	0.08
Nagaland	N/A	0.21	3.64	0.78	0.05	0.01	0.94	4.11	0.23	0.43
Tripura	3,420	0.20	2.96	0.47	0.87	0.08	0.05	4.01	0.34	0.42
Sikkim	N/A	0.14	3.73	0.80	0.60	0.01	0.38	3.72	0.56	0.49
Central		**0.21**	**2.69**	**0.44**	**0.86**	**0.12**	**0.01**	**2.01**	**0.21**	**0.75**
Madhya Pradesh	4,149	0.22	2.85	0.65	0.93	0.05	0.02	1.98	0.27	0.74
Uttar Pradesh	3,516	0.20	2.62	0.34	0.83	0.16	0.01	2.03	0.19	0.76
East		**0.19**	**2.46**	**0.24**	**0.83**	**0.15**	**0.02**	**2.40**	**0.21**	**0.66**
Bihar	2,655	0.15	2.32	0.17	0.82	0.16	0.02	1.78	0.13	0.78
Orissa	3,077	0.15	2.42	0.29	0.97	0.01	0.02	2.16	0.16	0.67
West Bengal	4,753	0.27	2.67	0.30	0.76	0.22	0.02	3.30	0.33	0.51
South		**0.31**	**3.39**	**0.65**	**0.82**	**0.11**	**0.07**	**3.72**	**0.43**	**0.54**
Andhra Pradesh	4,728	0.26	3.20	0.65	0.88	0.08	0.04	2.48	0.39	0.69
Goa	N/A	0.50	4.32	0.92	0.67	0.05	0.27	5.38	0.71	0.34
Karnataka	4,696	0.33	3.27	0.66	0.86	0.11	0.03	3.13	0.40	0.61
Kerala	2,418	0.28	3.89	0.61	0.54	0.26	0.19	6.76	0.42	0.16
Tamil Nadu	5,047	0.35	3.42	0.66	0.88	0.06	0.06	4.07	0.50	0.50

Notes: The table reports selected economic and demograpchic characteristics by state. Data on per capita income comes from Cashin and Sahay (1996). Figures reported in columns 2-10 are based on tabulations from the first round of the NFHS. Tabulations for Sikkim are based on the second round of the NFHS. Observations based on NFHS data are weighted using state-level weights.

Table 3. Differential Probability of a Male Birth at Parity N as a Function of the Sex Composition of Older Siblings

	Born between 1975-1989				Born after 1989			
	Full Sample		Rural	Urban	Full Sample		Rural	Urban
	No controls (1)	Full controls (2)	Full controls (3)	Full controls (4)	No controls (5)	Full controls (6)	Full controls (7)	Full controls (8)
A. Parity 2 (omitted category=Boy)								
Girl	-0.004	-0.004	-0.005	0.000	0.020	0.020	0.016	0.031
	(0.003)	(0.003)	(0.004)	(0.013)	(0.007)	(0.007)	(0.008)	(0.010)
Sample Size	*50,175*	*50,175*	*34,624*	*15,551*	*80,424*	*80,424*	*51,350*	*29,074*
B. Parity 3 (omitted category=Boy-Boy)								
Girl-Girl	-0.006	-0.007	-0.009	0.000	0.032	0.030	0.018	0.067
	(0.010)	(0.009)	(0.008)	(0.024)	(0.010)	(0.010)	(0.008)	(0.014)
Girl-Boy	0.009	0.009	0.007	0.016	0.019	0.019	0.021	0.016
	(0.007)	(0.007)	(0.010)	(0.019)	(0.009)	(0.009)	(0.009)	(0.014)
Boy-Girl	-0.004	-0.004	-0.008	0.009	0.005	0.005	-0.003	0.028
	(0.010)	(0.010)	(0.007)	(0.030)	(0.008)	(0.008)	(0.010)	(0.012)
Sample Size	*39,042*	*39,042*	*27,963*	*11,079*	*55,289*	*55,289*	*34,849*	*15,895*

Notes: The table reports the differential probability of a male birth at parity 2 (panel A) and parity 3 (panel B) as a function of the sex composition of older siblings. The samples include all children born in the 15 years prior to each survey date. Estimates reported in columns 1-4 are for children born between 1975 and 1989. Estimates reported in columns 5-8 are for children born in 1990 or afterwards. Columns 1, 2, 5, and 6 report estimates for the full sample. Columns 3, 4, 7, and 8 report estimates from samples stratified by rural/urban residency. Regression estimates reported in columns 2-4 and 6-8 are from models that control also for twin status, mother's age, mother's education, mother's age a first birth, indicators for mother's religion, father's education, mother's mass media exposure, wealth, and rural/urban status. Observations are weighted using national-level weights. Standard errors clustered at the state level are reported in parenthesis. Sample sizes are reported in *italics* .

Table 4. Sex-Ratios and Ultrasound Use

	Ultrasound Test				Pregnancy outcome = boy conditional on doing ultrasound test			
	Full Sample		Rural	Urban	Full Sample		Rural	Urban
	No controls (1)	Full controls (2)	Full controls (3)	Full controls (4)	No controls (5)	Full controls (6)	Full controls (7)	Full controls (8)
A. Parity 2 (omitted category=1 son)								
No sons	0.022	0.020	0.023	0.016	0.071	0.070	0.064	0.081
	(0.007)	(0.006)	(0.005)	(0.012)	(0.016)	(0.016)	(0.020)	(0.022)
Sample Size	*20,265*	*20,265*	*12,104*	*8,161*	*6,225*	*6,225*	*2,323*	*3,902*
B. Parity 3 (omitted category=2 sons)								
No sons	0.102	0.087	0.075	0.128	0.103	0.101	0.161	0.063
	(0.022)	(0.021)	(0.022)	(0.024)	(0.028)	(0.029)	(0.056)	(0.040)
1 son	0.027	0.021	0.010	0.059	-0.007	-0.008	0.087	-0.083
	(0.011)	(0.012)	(0.014)	(0.028)	(0.040)	(0.043)	(0.076)	(0.040)
Sample Size	*11,777*	*11,777*	*7,822*	*3,955*	*2,398*	*2,398*	*991*	*1,407*
C. Parity 4 (omitted category= 3 sons)								
No sons	0.080	0.060	0.077	0.003	0.219	0.134	0.083	0.291
	(0.015)	(0.015)	(0.020)	(0.051)	(0.130)	(0.133)	(0.241)	(0.102)
1 son	0.021	0.015	0.028	-0.026	0.102	0.039	0.060	0.057
	(0.009)	(0.009)	(0.013)	(0.041)	(0.114)	(0.120)	(0.217)	(0.131)
2 sons	-0.013	-0.004	0.014	-0.068	0.101	0.048	0.043	0.091
	(0.013)	(0.012)	(0.013)	(0.037)	(0.099)	(0.106)	(0.249)	(0.110)
Sample Size	*6,843*	*6,843*	*4,800*	*2,043*	*938*	*938*	*423*	*515*

Notes: Columns 1 through 4 report the differential likelihood that a mother performs an ultrasound test during pregnancy as a function of the sex composition of her older children. Columns 5 through 8 report the differential likelihood of a male birth as a function of the sex composition of her older children among mothers who performed an ultrasound test during that pregnancy. Panels A, B, and C report estimates for pregnancies/births outcomes at parity 2, 3, and 4 respectively. Columns 1,2,5, and 6 report estimates for the full sample of mothers. Columns 3, 4, 7, and 8 report estimates for samples stratified by rural/urban residency. Regression estimates reported in columns 2-4 and 6-8 are from models that control also for twin status, mother's age, mother's education, mother's age a first birth, indicators for mother's religion, father's education, mother's mass media exposure, wealth, and rural/urban status. Observations are weighted using national-level weights. Standard errors clustered at the state level are reported in parenthesis. Sample sizes are reported in italics.

Table 5. Effects on Nutritional Status of Children

				Basic specification		Individual controls	
Outcome	Sample size (1)	Outcome mean (2)	Females -Males (3)	MFR (4)	Female *x* MFR (5)	MFR (6)	Female *x* MFR (7)
Underweight	76,314	0.485	0.016	-0.162	-0.535	-0.095	-0.536
			(0.003)	(0.336)	(0.160)	(0.318)	(0.132)
Wasted	69,784	0.179	-0.015	-0.050	-0.247	-0.029	-0.250
			(0.004)	(0.284)	(0.109)	(0.275)	(0.104)
Stunted	69,571	0.433	0.014	0.316	-0.287	0.386	-0.293
			(0.004)	(0.230)	(0.180)	(0.236)	(0.164)

Notes: The table reports the association between MFR at birth in the state of residence and nutritional status of children. Column 2 reports the outcome means and column 3 reports the female-male differential in nutritional status. Columns 4 and 5 report regression estimates for MFR and MFR interacted with a female dummy from a linear probability model that includes gender specific state and year of birth fixed effects. Columns 6 and 7 report regression estimates from a model that controls also for the covariates specified in table 3. The sample includes children aged 0 to 35 months born in the last two births of ever married women sampled in rounds 1-3 of the NFHS surveys. Height measures were not taken in round 1 for the following 5 states: Andhra Pradesh, Himachal Pradesh, Madhya Pradesh, Tamil Nadu and West Bengal. Observations are weighted using national-level weights. Standard errors clustered at the state level are reported in parenthesis.

Table 6. Effects on Nutritional Status of Children by Place of Residence and Parity

Outcome	Sample size (1)	Outcome mean (2)	Females -Males (3)	MFR (4)	Female *x* MFR (5)
A. Rural					
Underweight	53,158	0.513	0.014	-0.123	-0.547
			(0.004)	(0.299)	(0.182)
Wasted	48,191	0.185	-0.018	0.046	-0.301
			(0.004)	(0.289)	(0.126)
Stunted	48,037	0.458	0.012	0.345	-0.360
			(0.005)	(0.252)	(0.193)
B. Urban					
Underweight	23,156	0.395	0.014	0.055	-0.407
			(0.008)	(0.409)	(0.185)
Wasted	21,593	0.158	-0.007	-0.212	-0.147
			(0.005)	(0.290)	(0.222)
Stunted	21,534	0.351	0.017	0.335	0.312
			(0.009)	(0.235)	(0.274)
C. Parity 1					
Underweight	22,820	0.428	0.006	0.027	-0.565
			(0.007)	(0.280)	(0.186)
Wasted	20,892	0.160	-0.018	0.231	-0.137
			(0.008)	(0.237)	(0.203)
Stunted	20,868	0.379	0.004	0.508	-0.285
			(0.008)	(0.241)	(0.196)
D. Parity>1					
Underweight	53,494	0.508	0.020	-0.137	-0.542
			(0.005)	(0.334)	(0.144)
Wasted	48,892	0.186	-0.013	-0.111	-0.317
			(0.004)	(0.304)	(0.178)
Stunted	48,703	0.454	0.019	0.345	-0.325
			(0.004)	(0.245)	(0.199)

Notes: The table reports the association between MFR at birth in the state of residence and nutritional status of children for samples stratified by place of residence (Panels A and B) and parity (Panels C and D). Columns 4 and 5 report regression estimates from a model that controls also for the covariates specified in table 3. The sample includes children aged 0 to 35 months born in the last two births of ever married women sampled in rounds 1-3 of the NFHS surveys. Observations are weighted using national-level weights. Standard errors clustered at the state level are reported in parenthesis.

Table 7. Effects on Nutritional Status of Children by Region and Survey Round

Outcome	Sample size (1)	Outcome mean (2)	Females -Males (3)	round 2 x treated (4)	female x round 2 x treated (5)	round 3 x treated (6)	female x round 3 x treated (7)
Underweight	76,314	0.485	0.014	0.034	-0.043	0.000	-0.064
			(0.003)	(0.040)	(0.019)	(0.040)	(0.025)
Wasted	69,784	0.179	-0.015	0.004	-0.022	-0.009	-0.056
			(0.003)	(0.023)	(0.017)	(0.033)	(0.016)
Stunted	69,571	0.433	0.013	0.037	-0.018	0.032	-0.034
			(0.004)	(0.032)	(0.020)	(0.040)	(0.024)

Notes: Columns 4-7 report estimates from a triple-differences model that compares nutritional outcomes of girls versus boys in treated versus comparison states over the three survey rounds. The omitted category is survey round 1. The treated group includes the following states: Gujarat, Haryana, Himachal Pradesh, Maharashtra, Punjab, Jammu and Kashmir, Rajasthan, and Delhi. The models control for state fixed effects and survey round indicators interacted with gender. In addition the model controls for twin status, mother's age, mother's age a first birth, and indicators for mother's religion, mother's education, father's education, mother's mass media exposure index, wealth index, and rural/urban status. Observations are weighted using national-level weights. Standard errors clustered at the state level are reported in parenthesis.

Table 8. Family Characteristics of Girls vs. Boys

Outcome	Sample size (1)	Outcome mean (2)	Females -Males (3)	MFR (4)	Female *x* MFR (5)
Mother's age	76,500	25.3	-0.098	2.952	-1.817
			(0.024)	(2.268)	(1.703)
Mother's age at 1st birth	76,500	19.0	0.000	4.863	0.785
			(0.025)	(1.704)	(0.873)
Mother's educ	76,394	3.62	-0.118	6.489	0.243
			(0.042)	(1.769)	(1.419)
Father's educ	76,032	6.15	-0.137	4.538	0.682
			(0.051)	(1.215)	(2.275)
Wealth index	76,500	2.84	-0.041	-0.268	-0.039
			(0.009)	(0.623)	(0.522)
Rural	76,500	0.763	0.002	-0.300	0.158
			(0.003)	(0.121)	(0.159)
Number of children	76,500	2.92	-0.007	-1.421	-0.923
			(0.009)	(0.672)	(0.534)

Notes: The table reports the association between MFR at birth in the state of residence and household characteristics. Column 2 reports variable means and column 3 reports the female-male differential. Columns 4 and 5 report regression estimates for MFR and MFR interacted with a female dummy from a model that includes gender specific state and year of birth fixed effects. The sample includes children aged 0 to 35 months born in the last two births of ever married women sampled in rounds 1-3 of the NFHS surveys. Observations are weighted using national-level weights. Standard errors clustered at the state level are reported in parenthesis.

Table 9. Breastfeeding Duration

Outcome Breastfed for:	Sample size (1)	Outcome mean (2)	Females -Males (3)	MFR (4)	Female *x* MFR (5)
		A. Full Sample			
At least 12 months	59,670	0.881	-0.011	0.093	-0.058
(age>=12 months)			(0.004)	(0.132)	(0.133)
At least 18 months	43,295	0.707	-0.038	0.128	-0.206
(age>=18 months)			(0.007)	(0.190)	(0.230)
At least 24 months	29,537	0.544	-0.056	0.037	0.391
(age>=24 months)			(0.007)	(0.186)	(0.155)
		B. Rural			
At least 12 months	41,183	0.906	-0.013	0.001	-0.033
(age>=12 months)			(0.004)	(0.129)	(0.115)
At least 18 months	29,624	0.744	-0.045	0.109	-0.276
(age>=18 months)			(0.006)	(0.160)	(0.208)
At least 24 months	20,265	0.583	-0.069	-0.068	0.507
(age>=24 months)			(0.008)	(0.228)	(0.228)
		C. Urban			
At least 12 months	18,487	0.805	-0.009	0.129	-0.040
(age>=12 months)			(0.008)	(0.294)	(0.289)
At least 18 months	13,671	0.594	-0.020	0.054	0.005
(age>=18 months)			(0.013)	(0.382)	(0.394)
At least 24 months	9,272	0.421	-0.018	0.373	-0.358
(age>=24 months)			(0.010)	(0.402)	(0.542)
		D. Parity 1			
At least 12 months	17,789	0.829	-0.007	0.217	0.036
(age>=12 months)			(0.008)	(0.182)	(0.148)
At least 18 months	12,913	0.618	-0.015	0.159	-0.001
(age>=18 months)			(0.008)	(0.208)	(0.463)
At least 24 months	8,650	0.438	-0.017	0.370	0.162
(age>=24 months)			(0.013)	(0.237)	(0.267)
		E. Parity>1			
At least 12 months	41,881	0.903	-0.013	0.039	-0.126
(age>=12 months)			(0.006)	(0.167)	(0.176)
At least 18 months	30,382	0.743	-0.047	0.127	-0.295
(age>=18 months)			(0.010)		
At least 24 months	20,887	0.585	-0.070	-0.056	0.492
(age>=24 months)			(0.009)	(0.208)	(0.232)

Notes: The table reports associations between MFR and indicators for breastfeeding duration. Panel A reports estimates for the full sample. Panels B and C report estimates for samples stratified by rural/urban residence. Panels D and E report estimates for samples stratified by parity. All samples include children born within the last 35 months prior to survey date. The minimum age restriction used in each of the regressions is specified in the first column of the table. All models include gender specific state and year of birth fixed effects and control for the set of covariates specified in Table 3. In addition, the models control for age in months and age in months interacted with a female dummy. Observations are weighted using national-level weights. Standard errors clustered at the state level are reported in parenthesis.

Table 10. Effects on Mortality

		Children Born within 36 months preceding survey					Children Born within 120 months preceding survey				
Outcome	Sample (1)	Sample size (2)	Outcome mean (3)	Females -Males (4)	MFR (5)	Female *x* MFR (6)	Sample size (7)	Outcome mean (8)	Females -Males (9)	MFR (10)	Female *x* MFR (11)
Neonatal mortality	Age ≥	98,922	0.042	-0.007	0.010	-0.024	356,361	0.048	-0.007	0.001	-0.019
(death between 0-29 days)	1 month			(0.002)	(0.031)	(0.047)			(0.001)	(0.031)	(0.034)
Post-neonatal mortality	Age ≥	63,961	0.023	0.002	-0.004	-0.049	309,689	0.028	0.003	0.000	0.006
(Death between 1 month-12 months)	12 months			(0.001)	(0.033)	(0.061)			(0.001)	(0.030)	(0.028)
Child Mortality	Age ≥						174,978	0.029	0.010	0.067	0.023
(Death between age 1 and before age 5)	60 months								(0.003)	(0.034)	(0.040)
Under 5 mortality	Age ≥						188,924	0.108	0.004	0.062	0.031
(Death before fifth birthday)	60 months								(0.005)	(0.080)	(0.067)

Notes: The table reports associations between MFR and various indicators of mortality. The sample for columns 2-6 is the same sample used for the analysis on nutritional status of children (see e.g. Table 5) and includes all children aged 0 through 35 months born in the last two births to ever married women sampled in rounds 1-3 of the NFHS surveys. The sample for columns 7-11 includes all children born in the last ten years prior to survey date to ever married women sampled in rounds 1-3 of the NFHS surveys. The minimum age restriction used for each outcome is specified in column 1. Columns 3 and 8 report the outcome means and columns 4 and 9 report the female-male differential in the outcome variables. Columns 5 and 6 and columns 10 and 11 report regression estimates for MFR and MFR interacted with a female dummy from a linear probability model that includes gender specific state and year of birth fixed effects and control also for the covariates specified in table 3. Observations are weighted using national-level weights. Standard errors clustered at the state level are reported in parenthesis.

Figure 1: Male to Female Ratio at Birth

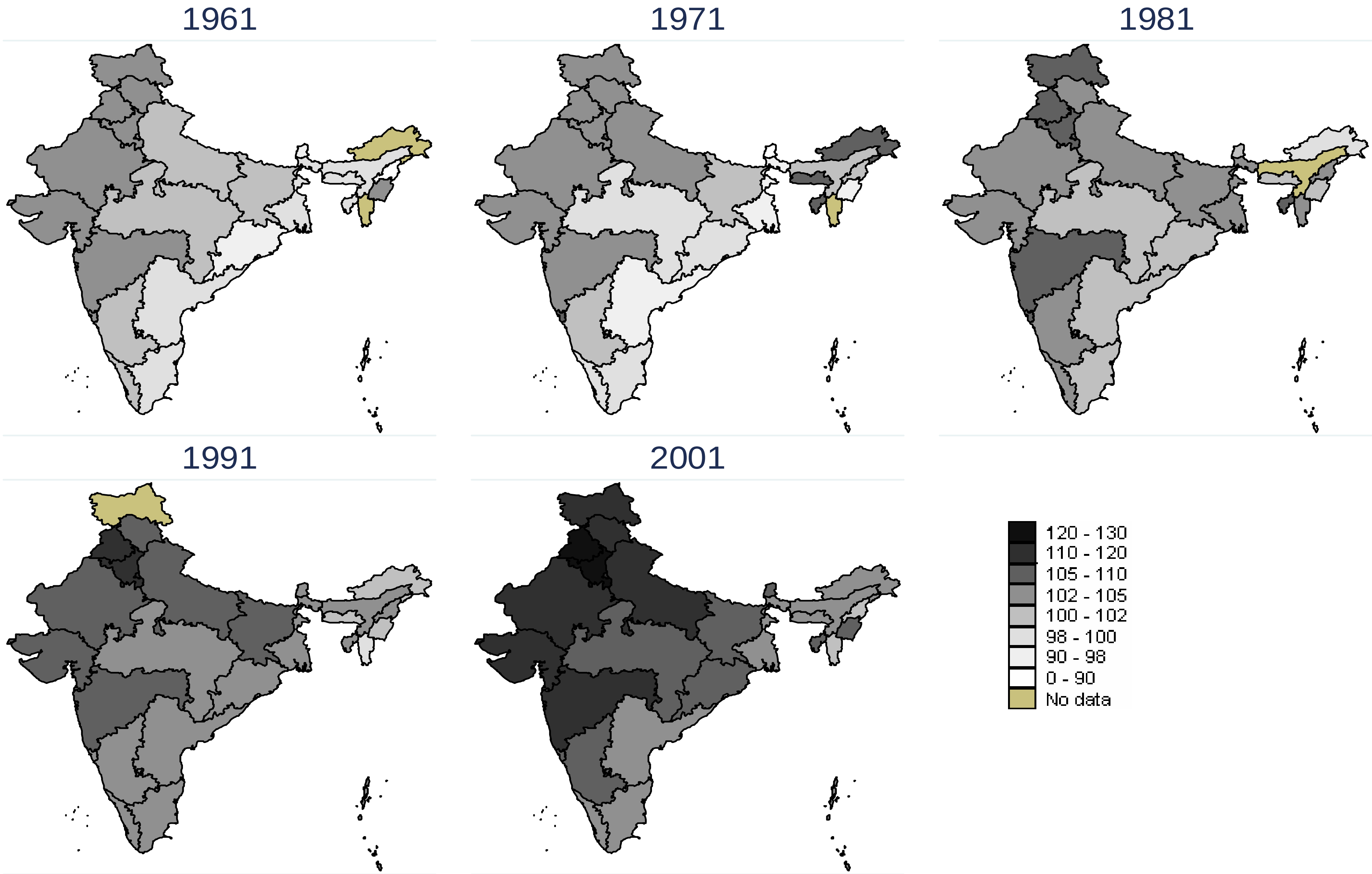

Figure 2a. Male to Female Ratio at Age 0 for the Urban Population

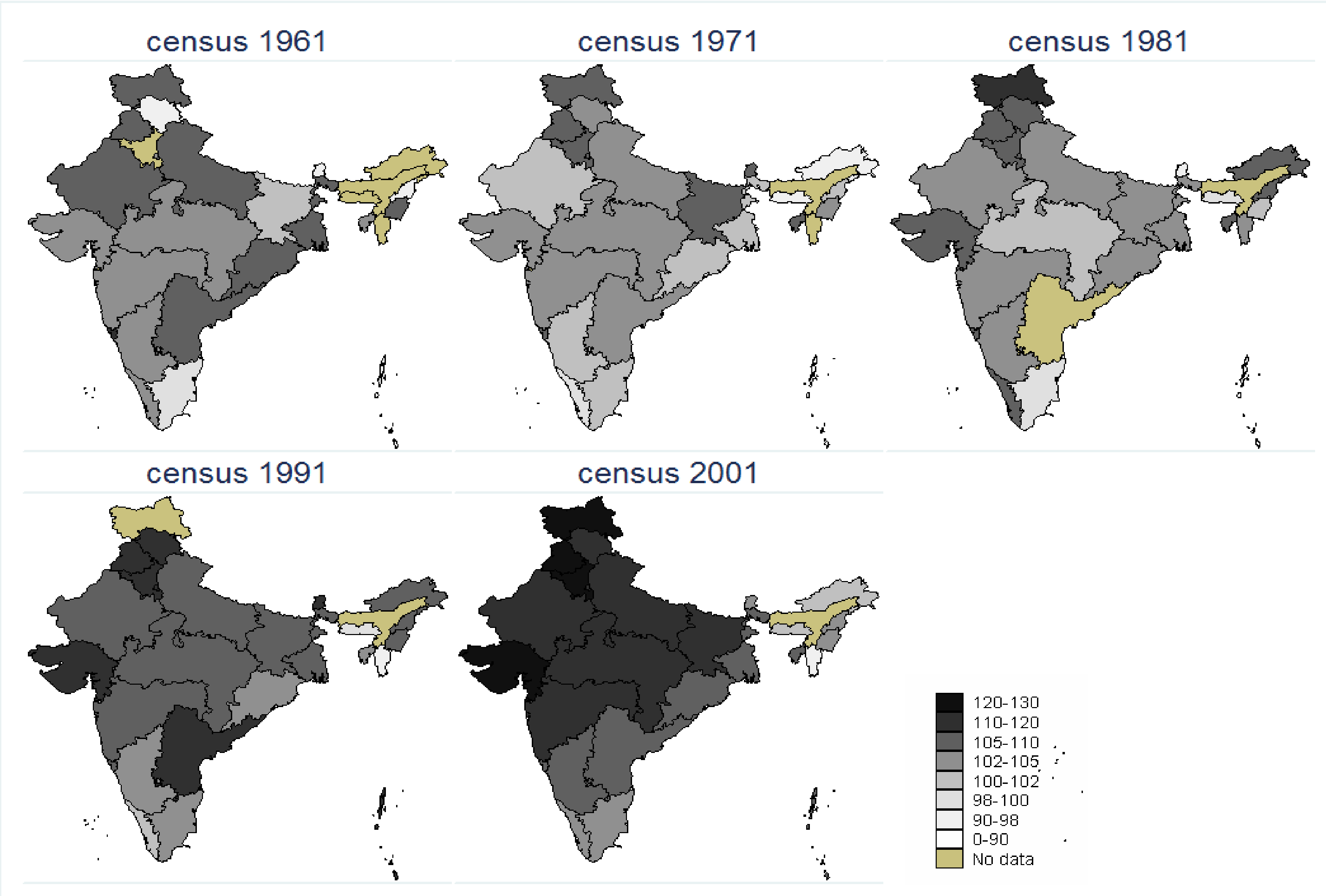

Figure 2b. Male to Female Ratio at Age 0 for the Rural Population

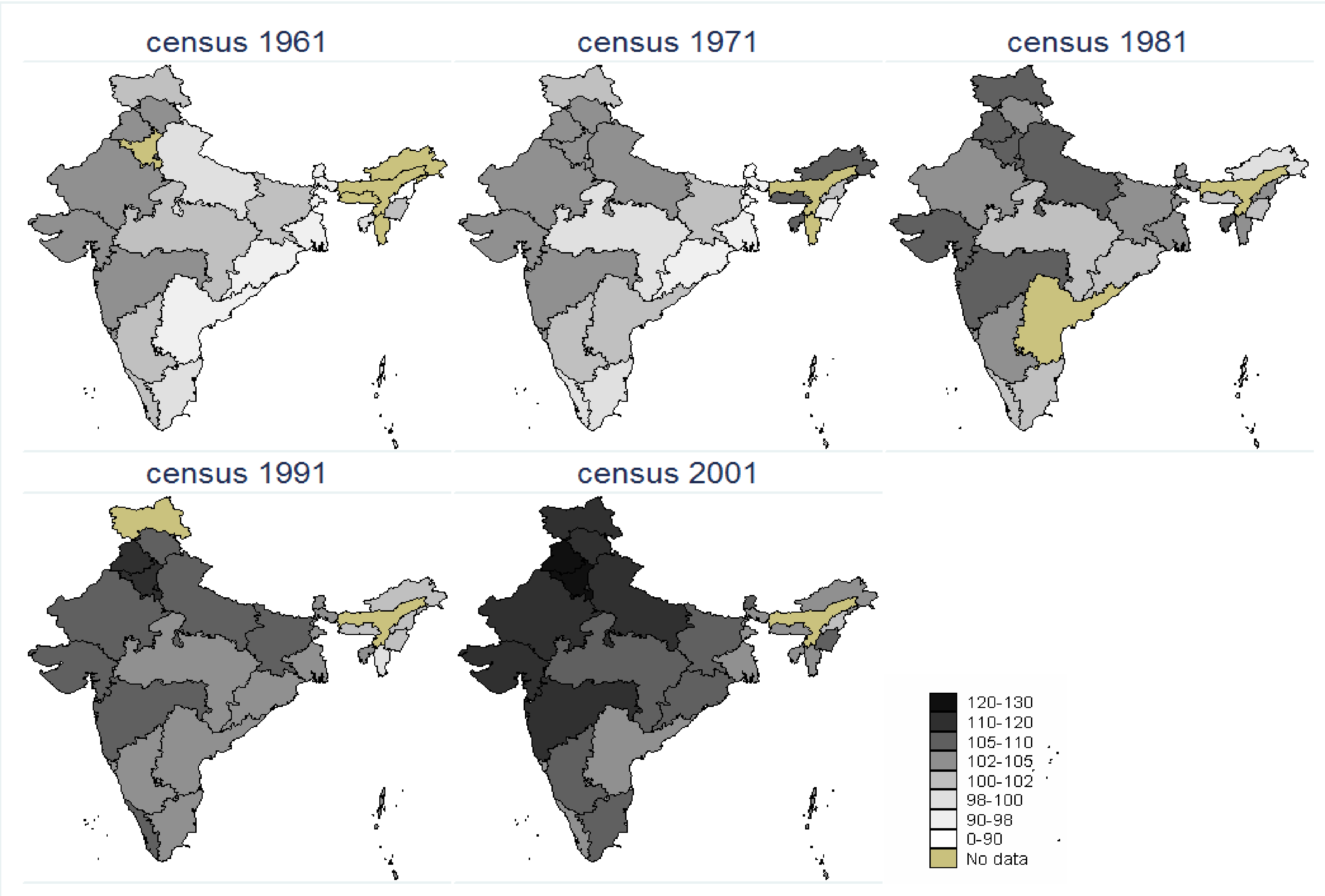

Figure 3: Proportion Breastfed for at least x months

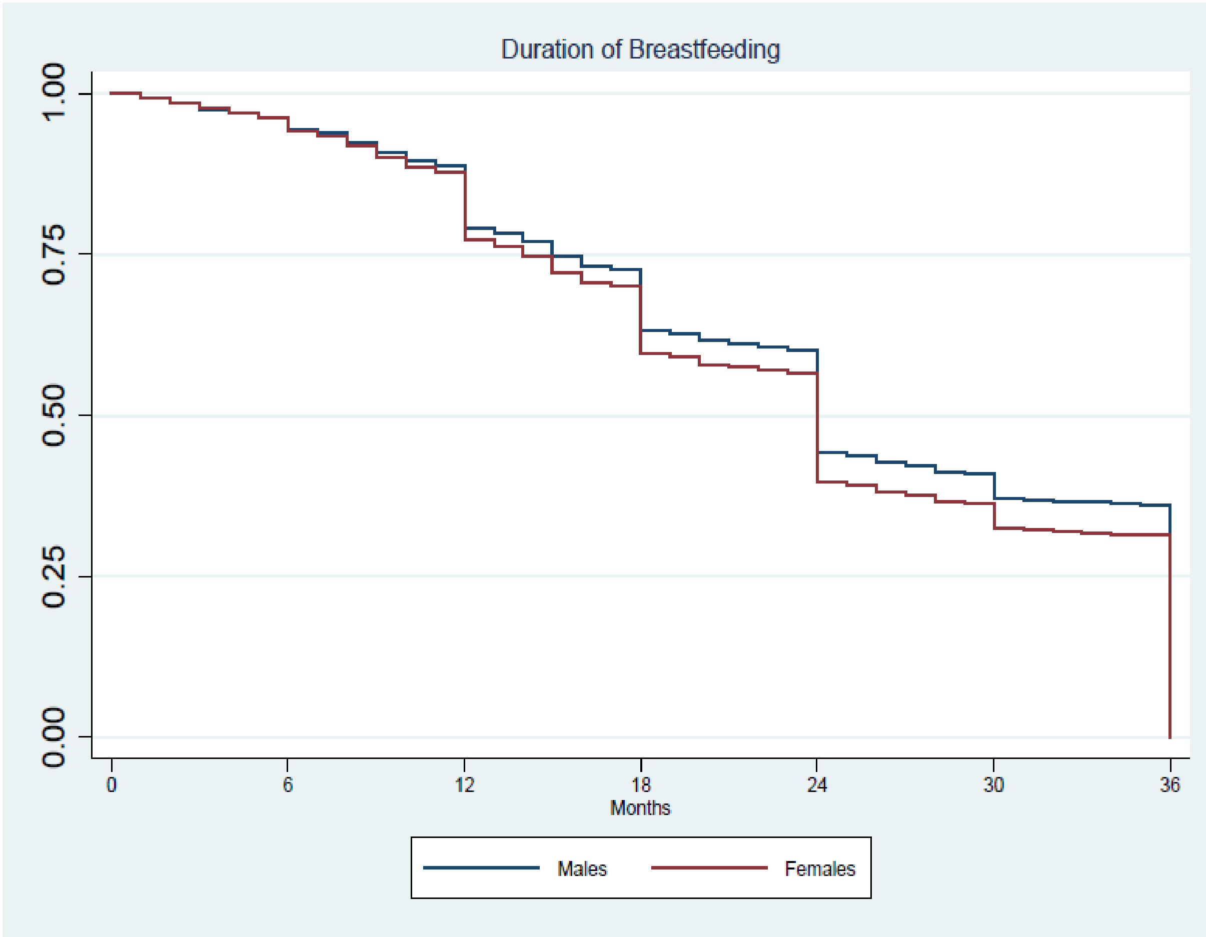

Table A1. Robustness checks

	Main results		Household covariates interacted with gender		State varying controls 1 + HH covariates interacted with gender		State varying controls 2 + HH covariates interacted with gender		New-zscores	
	MFR	Female *x* MFR	MFR	Female *x* MFR	MFR	Female *x* MFR	MFR	Female *x* MFR	MFR	Female *x* MFR
Outcome	(1)	(2)	(3)	(4)	(7)	(8)	(5)	(6)	(9)	(10)
Underweight	-0.095	-0.536	-0.107	-0.514	-0.091	-0.433	0.186	-0.344	-0.024	-0.543
	(0.318)	(0.132)	(0.319)	(0.126)	(0.254)	(0.115)	(0.350)	(0.169)	(0.258)	(0.137)
Wasted	-0.029	-0.250	-0.036	-0.236	0.140	-0.135	0.159	-0.268	0.069	-0.234
	(0.275)	(0.104)	(0.278)	(0.105)	(0.181)	(0.114)	(0.243)	(0.141)	(0.250)	(0.082)
Stunted	0.386	-0.293	0.379	-0.276	0.415	-0.217	0.193	-0.058	0.249	-0.191
	(0.236)	(0.164)	(0.239)	(0.164)	(0.175)	(0.112)	(0.177)	(0.100)	(0.200)	(0.119)

	Log MFR		5-year span for MFR		9-year span for MFR		Unweighted regression		State weights	
	MFR	Female *x* MFR	MFR	Female *x* MFR	MFR	Female *x* MFR	MFR	Female *x* MFR	MFR	Female *x* MFR
Outcome	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Underweight	-0.099	-0.609	-0.103	-0.426	0.130	-0.591	-0.002	-0.447	-0.002	-0.441
	(0.344)	(0.147)	(0.269)	(0.131)	(0.386)	(0.159)			(0.238)	(0.105)
Wasted	-0.029	-0.295	-0.015	-0.184	0.080	-0.316	0.150	-0.287	0.135	-0.284
	(0.302)	(0.105)	(0.199)	(0.077)	(0.295)	(0.138)	(0.200)	(0.096)	(0.201)	(0.089)
Stunted	0.463	-0.330	0.175	-0.222	0.540	-0.236	0.177	-0.224	0.191	-0.244
	(0.254)	(0.188)	(0.205)	(0.161)	(0.308)	(0.193)	(0.133)	(0.102)	(0.139)	(0.097)

Notes: The table reports estimates from various robustness checks. See section VI of the paper for a detailed explanation of each test.

Table A2. Differential Probability of a Male Birth at Parity N as a Function of Sex Composition of Previous Children:
Northern and Western States vs. Other States

	Born between 1975-1989		Born after 1989	
	Treated States	Other States	Treated States	Other States
	(1)	(2)	(3)	(4)
A. Parity 2 (omitted category=Boy)				
Girl	0.004	-0.007	0.038	0.013
	(0.010)	(0.003)	(0.008)	(0.009)
Sample Size	*16,697*	*33,478*	*24,287*	*56,137*
B. Parity 3 (omitted category=Boy-Boy)				
Girl-Girl	0.027	-0.020	0.079	0.012
	(0.006)	(0.010)	(0.024)	(0.007)
Girl-Boy	0.003	0.011	0.053	0.007
	(0.010)	(0.009)	(0.019)	(0.009)
Boy-Girl	0.017	-0.012	0.015	0.001
	(0.014)	(0.010)	(0.014)	(0.011)
Sample Size	*12,905*	*26,137*	*16,543*	*35,739*

Notes: The table reports the differential probability of a male birth at parity 2 (panel A) and parity 3 (panel B) as a function of the sex composition of previous children. The table reports estimates for the subsample of treated states (columns 1 and 3) and all other states (columns 2 and 4). The sample includes all women aged 15-49 surveyed in rounds 1-3 of the NFHS. Estimates reported in columns 1 and 2 are for children born between 1975 and 1989. Estimates reported in columns 3 and 4 are for children born in 1990 or afterwards. Regression estimates come from models that control also for twin status, mother's age, mother's education, mother's age a first birth, indicators for mother's religion, father's education, mother's mass media exposure, wealth, and rural/urban status. Observations are weighted using national-level weights. Standard errors clustered at the state level are reported in parenthesis. Sample sizes are reported in italics.

Table A3: Relationship Between Son Preference and MFR at the State Level

Outcome	Outcome mean (1)	Bivariate regression (2)	Adding state fixed effects (3)	Adding state fixed effects and mother's background characteristics (4)
Ideal MFR	1.249	-0.044	-0.555	0.014
		(0.319)	(0.333)	(0.213)
Wants more sons than daughters	0.285	-0.075	-0.524	0.057
		(0.235)	(0.249)	(0.181)

Note: The table reports associations between indicators for mother's son preferences and state MFR by survey round. The sample includes mothers whose youngest child was born within last 3 years prior to the survey date. Data are collapsed by state and round (No. obs =77) . Estimates reported in column 2 come from a simple bivariate regression. Estimates reported in column 3 control for state fixed effects. Estimates reported in column 4 come from regressions that control for state fixed effects and state-year means of the following variables: indicators for mass media exposure, religion, mother's education, mother's age at first birth , urban status, and wealth.