Family Composition Effects on Gender Differentials in Nutrition and Immunization in Rural India

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The literature on gender discrimination over the last several decades has shown that gender differentials in child mortality exist in most regions of the developing world, with particularly severe excess female mortality in India and other parts of South and East Asia. However, discrimination against *surviving* girls tends to be presumed from the evidence about excess female child mortality and the nature of such discrimination has not received adequate attention. In addition, while recent literature in India and Bangladesh has found that mortality discrimination is "selective" (Das Gupta, 1987), and that girls born into households with many daughters fare particularly badly (Muhuri and Preston, 1991), there has been little research into the role of sibling sex composition in selective discriminatory practices affecting health status among living children. This lack of evidence persists despite the recognition that discrimination in aspects of child care such as morbidity, nutritional status or use of preventive and curative health care is likely to contribute to greater gender differentials in mortality (Miller, 1981; Arnold et al, 1996; Bardhan, 1974, 1982; Kishor, 1993, 1995; Kurz and Johnson-Welch, 1997; Waldron, 1987; Makinson, 1994; Obermeyer and Cardenas, 1997).

This paper addresses some of these gaps in the literature by examining gender differentials in immunization and severe stunting among surviving rural Indian children under the age of five years. In particular, I focus on the effects of the sex composition of surviving older siblings[1] on such gender differentials.

Background

A number of cross-national studies of child mortality levels and differentials in the 1970s and 1980s point to widespread excess female child mortality in many areas of the developing world (Tabutin and Willems, 1995; Arnold, 1992; Hill and Upchurch 1995). The Demographic and Health Surveys (DHS), conducted widely in developing countries, also point to widespread excess female child mortality. In one analysis, 18 of 26 DHS countries studied had higher girls' than boys' mortality for children ages 1-4 years (Arnold, 1992). Hill and Upchurch (1995), in a review of DHS data, found that girls in the developing world have higher mortality risks than boys for any given mortality level, compared to historical differentials in the now developed world at a comparable mortality level. There are distinct geographic patterns in the extent of discrimination against girl children, and studies point to the Middle East, North Africa, East Asia and South Asia as having the largest differentials (Arnold, 1992).

In India, excess female mortality and a preference for sons has existed for centuries. In the last one hundred years, population sex ratios from censuses have shown more or less a continuous increase (Visaria and Visaria, 1983, 1995), rising from 103 males per 100 females in 1901 to 107.6 males per 100 females in 1991 (Desai, 1994). The sex ratio for India in the 1991 census is the highest to be recorded in

[1] The terms "family composition" and "sex composition" are used interchangeably to refer to the sex composition of children in a household.

this century[2] and shows 32 million more men than women in the country as a whole (Visaria and Visaria, 1995).

Much of this excess female mortality occurs in childhood: the first reports of excess female child mortality noted by the British date back to 1789, to the open practice of female infanticide among the Rajkumar clan of Rajputs in Jaunpur district, eastern Uttar Pradesh (Miller, 1981). Indian censuses since 1872 have noted masculine sex ratios among children under the age of 12 years (Miller, 1981). There is also wide agreement that since the 1960s, excess female child mortality has remained high, if not increased, in many parts of the country (El-Badry, 1969; Miller, 1989; Desai, 1994; Bhat 1989; Parasuraman and Roy, 1991). Censuses from 1961 to 1991 show that sex ratios for children under 10 years have become more masculine all across India, including in the East, West and South (Miller, 1989; Das Gupta and Bhat, 1995). In 1992-93, child mortality — the risk of dying between the ages of one and five years — for girls in India as a whole at 42 per 1,000 was 43% higher than for boys at 29.4 per 1,000 (IIPS, 1995; p218). Studies from other South Asian countries provide evidence of discrimination against girl children in these countries as well (Bairagi, 1986; Caldwell and Caldwell, 1990; D'Souza and Chen, 1980; Koenig and D'Souza, 1986; Faisel et al., 1993).

Despite a large body of evidence on gender differentials in mortality, there remains little research on gender discrimination among *living* children. There are several reasons why we need a better understanding of the extent and nature of discriminatory behavior against living girls. First, gender discrimination that contributes to poorer health status for girls compared to boys is likely to be the main pathway for excess female child mortality and thus needs to be understood. Second, such discrimination, which may be manifested in non-fatal episodes of untreated or poorly treated disease or levels of malnutrition, can still produce increasing frailty among survivors and thus is an important child health issue in itself (Mosley and Chen, 1984; Mosley and Becker, 1991). Third, poor health for surviving girls in childhood has implications for their health in later life, including reproductive years, and may be perpetuated across generations (Merchant and Kurz, 1992).

That such discrimination occurs in India was recognized at least as early as the 1901 Census, which notes that "...there is no doubt that, as a rule, she [a girl] receives less attention than would be bestowed upon a son. She is less warmly clad,...she is probably not so well fed as a boy would be, and when ill, her parents are not likely to make the same strenuous efforts to ensure her recovery..." (1901 Census, quoted in Miller, 1981; p. 67). Recent evidence corroborates these findings.

Evidence of gender differentials in health care and nutrition comes from both northern and southern India. Studies in several states across the country have found that boys are much more likely

[2] This high sex ratio could partly be due to relatively greater under-enumeration of females in the 1991 census (Kapoor, 1991; IIPS, 1995). However, even the National Family Health Survey's slightly lower estimate of the sex ratio for 1992-93 at 106 males per 100 females is the third-highest recorded in this century, is close to that estimated by the 1991 Census, and indicates that the current sex ratio is high by historical standards (IIPS, 1995).

than girls to be taken to a health facility when sick (Ganatra, B. and S. Hirve, 1994; Kishor 1995; Visaria 1988; Ravindran, 1986; Caldwell, Reddy and Caldwell, 1982; Caldwell and Caldwell, 1990; Govindaswamy and Ramesh, 1996; Das Gupta, 1987). In a recent study, boys were almost 50 percent more likely than girls to be immunized in Uttar Pradesh in the North and about 20 percent more likely to be immunized than girls in Tamil Nadu in the South (Govindaswamy and Ramesh, 1996). In 1992-93, boys had higher immunization rates than girls in all but two states in India (Goa and Karnataka), though the extent of this differential varies by state (Kurz and Johnson-Welch, 1997). Similarly, girls are more likely to be malnourished than boys in both northern and southern states (Basu, 1989; Das Gupta, 1987; Pebley and Amin, 1991; Wadley, 1993; Sen and Sengupta, 1983; Caldwell and Caldwell, 1990; Arnold et. al., 1996). Research in Bangladesh and Pakistan shows similar discriminatory behavior against girl children in nutrition and health care (Chen, Huq and D' Souza, 1981; Koenig and D' Souza, 1987; Baqui et al, 1998; Faisel et al, 1993).

The strength and persistence of such discrimination against girls (and the concomitant preference for sons) in South Asia, particularly in India, stem from the perceived economic, social, and religious utility of sons compared to daughters.[3] Parents of girls are socially-bound to find suitable husbands for their daughters at an early age, often pay all marriage costs, and provide a dowry; social norms dictate that parents cannot expect much emotional or economic support from married daughters (Dyson and Moore, 1983; Arnold et al., 1996; Kishor, 1995; Mandelbaum, 1988). On the other hand, parents expect sons to provide financial and emotional care and regard them as a "social security" for old age; sons also perform important religious roles, ensure the continuation of the family lineage, and may be desired to increase a family's capacity to defend itself or to exercise power (Cain, Khanam and Nahar, 1979; Cain, 1988, 1993; Vlassof, 1990; Oldenburg, 1992; Dharmalingam, 1996). Finally, inheritance laws largely favor sons, and, in agricultural economies such as exist in South Asia, this bias is cited as an important contributor to a lower value accorded to daughters compared to sons (Agarwal, 1994; Singh, 1993). In India, these social, religious and economic rules are more strongly associated with kinship structures in North than in South India (Karve, 1965; Sopher, 1980; Dyson and Moore, 1983). However, recent studies, reviewed above, show some blurring of North-South distinctions in gender discrimination.

There is evidence to suggest that co-existing with son preference parents also have a preference for a balanced sex composition in the family, and that at least one daughter may be wanted. Research shows, moreover, that parents may selectively discriminate against children with certain sex-birth order characteristics in order to attain their desired family sex composition. In other words, not all girls may be

[3] Such discrimination is likely to be a combination of "active" discrimination (such as, for example, a deliberate choice to provide health care to a sick boy but not to a sick girl) and "passive" neglect (for example, discovering that a girl is sick later than would be the case for a boy, simply because girl children may be more ignored in day-to-day interactions than are boy children). This paper does not analyze which of these two types of discrimination is at play in the case of the two outcomes studied here, and the words "discrimination" and "neglect" are used interchangeably.

equally “unwanted” and equally discriminated against; by the same token, all boys may not be equally “wanted”. In traditional high fertility societies, selective discrimination may be the only way to achieve a desired family size and composition (Scrimshaw, 1978).

In a society with strong son preference, as is the case in much of India, the preferred family composition often includes two sons and at least one daughter, where the daughter is seen to provide important religious, social or emotional value (Williamson, 1976; Dharmalingam, 1996; Mutharayappa et al., 1997). Among Hindus in India, daughters are additionally important because giving away a girl in marriage (*kanyadaan*) is considered very meritorious (Miller, 1989). On the other hand, more than one or two daughters are usually not welcome. Thus girls born into a family that already has daughters are the ones most likely to be perceived as least valuable by the household and most likely to face selective discrimination.

Studies in northern India, Bangladesh and China show that girls with older sisters have significantly higher mortality risks than girls with only surviving older brothers, or girls with no surviving older siblings (Arnold et al, 1996; Das Gupta, 1987; Muhuri and Preston, 1991; Amin, 1990; Choe et al., 1995). Son preference can also result in girls with older brothers being neglected in households that favor existing male children over newborn females, or families that do not want daughters at all (Simmons et al., 1982). On the other hand, again due to son preference, any harmful effect of same-sex siblings for boys may be weaker than is the case for girls (Muhuri and Preston, 1991). In India, the continuation of the desire for one or two sons combined with a growing desire for a small family suggests that selective discrimination against higher-birth order girls is likely to continue (Das Gupta, 1989; Das Gupta and Bhat, 1995).

While research has documented the extent and nature of such selective gender discrimination in mortality differentials, there is less clear evidence on the pathways through which parents’ conscious or subconscious desires for a certain sex composition of children are implemented, that is, the way they may treat *surviving* girls and boys and what impact any selective discriminatory practices have on surviving children’s health status. This issue is the central focus of this paper.

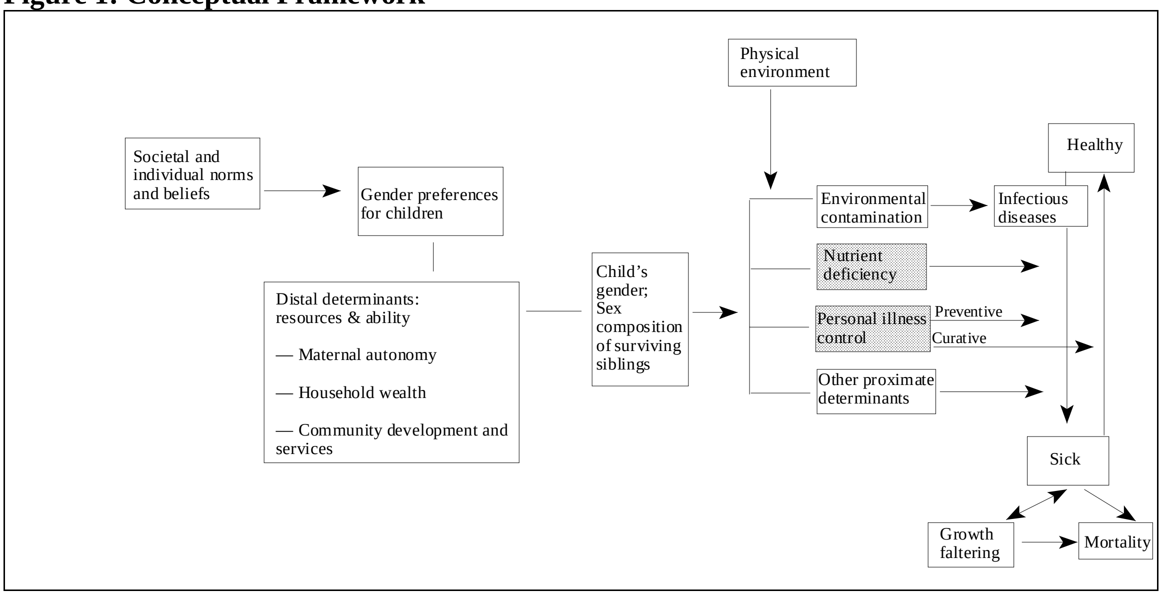
Framework and hypotheses

I model the effects of family sex composition on gender differentials in health status among surviving children using a modified proximate determinants framework (Mosley and Chen, 1984; Mosley, 1985). The proximate determinants model was originally developed to study factors affecting child mortality, and is based on the idea that all social and economic determinants of child mortality operate through a set of biological or proximate determinants to affect a child’s probability of survival. Thus, this model combines social, economic, medical and biological explanations of child mortality. Mosley and Chen (1984) group the proximate determinants into five categories, namely, maternal factors (mother’s age and parity); environmental contamination (routes of infection); nutrient deficiency; injury;

and personal illness control (preventive and curative care). All social and economic determinants of child mortality — the "distal" determinants — operate through these proximate determinants and are grouped by Mosley and Chen into individual-level, household-level and community-level variables.

My approach draws upon the proximate determinants framework inasmuch as I examine "health status", but with the health status of survivors — rather than mortality — as the outcome of interest. My model thus differs in that I analyze proximate determinants as outcomes themselves; gender differentials in the proximate determinants represent ways in which household attitudes against girls translate into worse health status for surviving girls and higher mortality rates for girls. My modified model is presented in Figure 1 below.

Figure 1: Conceptual Framework



The dependent variables analyzed in this study are in shaded boxes. The boxes and arrows to the right of the dependent variables illustrate the relationship of these variables to survival and frailty, closely following the Mosley-Chen (1984) framework. The theoretical integrity of the Mosley-Chen framework is, thus, preserved: these proximate determinants are conceptualized as the proximate pathways through which discriminatory health behavior contributes to mortality, as illustrated in the far right side of Figure 1. Thus, for instance, environmental contamination, nutrient deficiency and preventive personal illness control affect a child's movement from a healthy to a sick state, while curative personal illness control affects a child's movement from the sick to healthy states. A healthy child can fall sick due to infectious disease, malnutrition or lack of preventive health care; a sick child can either regain health (perhaps due to effective curative care), or survive at a more frail level than a healthy child, or die.[4]

[4] Testing the effects of gender differentials in the proximate determinants on excess female child mortality is beyond the scope of this study and will not be discussed here.

While Figure 1 presents all the proximate determinants so that the model is complete, the analysis focuses on two groups of proximate determinants that are expected to reflect significant discrimination between male and female children, and for which data are available, namely nutrient deficiency and personal (preventive) illness control.

I use severe stunting as my measure of nutrient deficiency, while recognizing that severe stunting can reflect the effects of nutrient deficiency as well as recurrent, untreated infection. Nevertheless, severe stunting is suitable to measure gender discrimination as it is likely to reflect long-term neglect or recurrent and poorly-treated illness episodes and thus should, theoretically, capture any continuing pattern of selective neglect against unwanted (usually female) children. Immunization is used as a proxy for personal illness control, as a measure for the use of preventive health care. Immunization as a preventive care measure is likely to reflect parental motivation to ensure future well-being and health for the immunized child. Parents are more likely to be motivated to make the effort for wanted (first or only male) children rather than unwanted (higher-birth order female) children; thus, one may see significant patterns of selective neglect in immunization. It is possible, however, that if immunization is provided free in a community, financial and other costs may be low enough that parents get all children immunized; in this case, gender differentials in immunization may not be observed.

The boxes to the left of the dependent variables outline the individual, household, and community characteristics that are the distal determinants of gender discrimination in health status. I group these factors somewhat differently from the proximate determinants framework to reflect the characterization in the literature of key factors that influence son preference and gender differentials in health status.

In the modified model presented here, the extent of gender differentials and the effect of family composition on these gender differentials is determined by two sets of distal determinants. First, individual and social norms and beliefs determine parental gender preferences and thereby the desired family sex composition. Second, the extent to which these parental preferences are translated into different health outcomes for children of different gender and birth order depends on the maternal, household and community resources and ability available to parents. Thus, we can expect to see family composition effects on gender differentials where gender preferences exist *and* where there is a capacity to realize these preferences (see Pande, 1999 for a fuller theoretical treatment of this model).

This use of a modified Mosley-Chen framework is suitable for several reasons. First, the model clearly outlines the relationship of health status among surviving children to mortality, and, in the context of this paper, clarifies the theoretical relationship between gender differentials among surviving children and gender differentials in mortality. Second, the literature demonstrates that child health is strongly influenced not just by child-level factors such as gender and birth order, but also by a number of other contextual factors such as maternal education, household wealth and community characteristics (Caldwell, 1979; Bhuiya and Streatfield, 1992; Bhuiya et al, 1989; Cleland and van Ginneken, 1988; Desai and Alva, 1998; Hobcraft, 1993; Schultz, 1984; Sastry, 1996). Thus, to accurately assess the extent to which family composition and gender affect severe stunting and immunization among children, other related factors have to be accounted for. The proximate determinants model — originally and as modified here — provides a conceptual framework within which to situate these "distal" determinants. Estimating all the relationships hypothesized in the conceptual model is, however, beyond the scope of this paper; here, I estimate a reduced-form model examining the extent of selective gender discrimination after controlling for social norms, and maternal, household and community factors.

Given the use of selective neglect to achieve a balanced family composition, regardless of gender preference, I expect children whose surviving older siblings are all of the same sex (girls with only surviving older sisters and boys with only surviving older brothers) to have a higher probability of severe stunting and lower chances of being immunized than children with surviving siblings of the opposite sex (girls with only surviving older brothers and boys with only surviving older sisters). However, where there is strong son preference and household capacity to realize such a preference, boys with same-sex siblings are expected to do better than girls with same-sex siblings, and girls with only surviving older sisters are expected to have the worst health status of all combinations of siblings.

Description of the data

I use the rural, *de jure* sample of children from the 1992-93 India National Family Health Survey (NFHS) to test these hypotheses. This section describes sample sizes and characteristics, and the definitions of variables used in the analysis. For all outcome and explanatory variables, values of "don't know" or "missing" are coded as "missing" and are excluded from the analysis.

Outcome variables

Severe stunting is defined as a height-for-age measurement more than 3 standard deviations below the international reference population recommended by the WHO and found applicable to Indian children by the Nutrition Foundation of India (IIPS, 1995; p.281). While age data are remarkably complete and available for over 98 percent of children (IIPS, 1995; p.322-323), heights were not measured in some states due to a lack of measuring boards.[5] In the remaining states, height data are available for 81.5 percent of eligible children. Exploratory analysis shows that children with and without height measurements are similar across key background characteristics; in particular, the sex ratio is similar between those measured and not measured (Pande, 1999). The sample for severe stunting used in this study comprises 14,715 rural surviving children between 6-47 months of age who have height-for-age data. I chose a minimum of 6 months of age based on literature that has found that little gender discrimination occurs in very early childhood (see for example, Muhuri and Preston, 1991; Bhuiya and

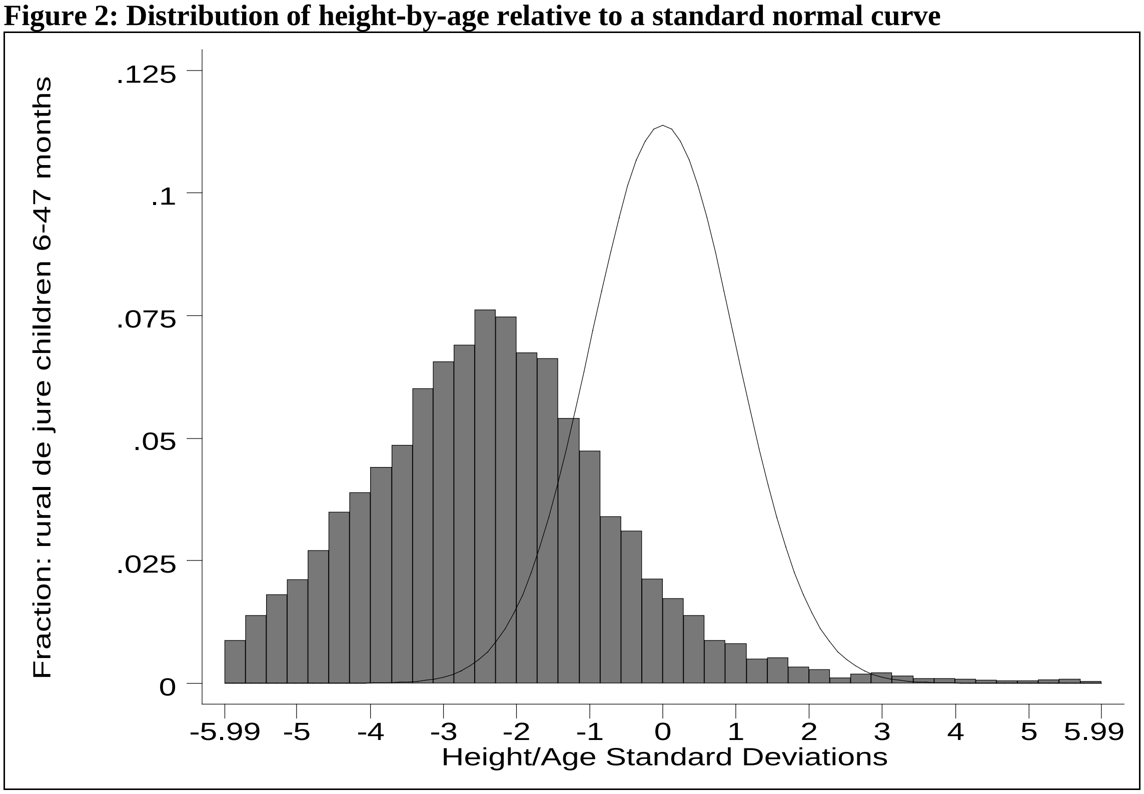
[5] Heights/lengths (supine or standing) were measured using an adjustable wooden measuring board. Height data are unavailable for Andhra Pradesh, Himachal Pradesh, Madhya Pradesh, Tamil Nadu and West Bengal. Exploratory analysis suggests that the lack of height measurements for these states is unlikely to bias the national average (IIPS, 1995; p.281).

Streatfield, 1992). I define severe stunting as a dichotomous variable (1=severely stunted, as defined above).

For immunization, I use a sample of 25,549 rural surviving children ages 12-59 months with data on immunization status. Children over the age of 12 months are chosen so as to allow every child in the analysis to have adequate exposure time for the full range of vaccinations, on the basis that, if the WHO schedule for immunizations is followed, then all children 12 months or older should be fully vaccinated or at least have had the chance to receive all of the vaccinations. Vaccination coverage is estimated by a combination of data from immunization cards and information from mother's recall where immunization cards are not available.[6] Only 21 percent of sample children had cards shown to the interviewer. Children of wealthier and more educated parents are both more likely to have an available card in the first place *and* are more likely to be immunized than children for whom only information from mother's recall is available (Pande, 1999). Consequently, multivariate analysis will control for whether a child's card was seen by the interviewer. Immunization is defined as an ordered categorical variable, coded as: none (no vaccinations by the time of the survey), some (at least one but not all vaccinations), and all (BCG, three doses each of DPT and oral polio vaccine, and measles vaccine).

Figure 2 shows the weighted distribution of height-by-age (standard deviations away from the

Figure 2: Distribution of height-by-age relative to a standard normal curve



[6] Studies have found that mothers report relatively accurately about their children's immunization, and that any bias in immunization rates calculated using both card and mothers' recall is likely to result in an under-estimate rather than an over-estimate of proportions immunized (Langsten and Hill, 1998).

reference population) among sample children ages 6-47 months, relative to a standard normal distribution, and provides evidence of the magnitude of moderate and severe stunting among rural Indian children.[7]

The data for both severe stunting and immunization paint a picture of poor health for the majority of rural Indian children at the time of the survey (table 1). Slightly over one-third (34.4 percent) of sample children were severely stunted. A low 30 percent of children between 12-60 months were fully immunized by the time of the survey, and close to 40 percent had not received any of the vaccinations.

TABLE 1: Severe stunting and immunization, total and by gender

	sev. stunt	No vacc.	BCG	DPT3	OPV3	Measles	All vacc.
Total	34.36	38.04	54.33	44.77	46.48	38.52	29.98
Girls	35.50	40.77	52.06	42.79	44.64	36.71	28.72
Boys	33.28	35.47	56.48	46.64	48.21	40.24	31.17
G.D.	6.25	13.00	-8.49	-9.00	-8.00	-9.62	-8.53

N=14715 for severe stunting; N=25549 for immunization

G.D.=gender differential, defined as percent difference between girls and boys

There is some evidence of gender differentials in these two health outcomes. Six percent more girls than boys are severely stunted. Gender differentials in immunization are larger than those in severe stunting, with 8-9 percent fewer girls than boys having received any or all vaccinations. The differential is largest among unvaccinated children, with 13 percent more girls than boys being totally unvaccinated.[8]

Explanatory variables

A number of explanatory variables are included as “distal” determinants to test this study’s hypotheses. The gender of the child is a key dichotomous variable, taking the value 1 if the child is female. Various combinations of the sex composition of older surviving siblings are used, namely, none, one surviving older brother and no surviving older sisters, one surviving older sister and no surviving older brothers, two or more surviving older sisters and no older brothers, two or more surviving older brothers and no older sisters, and a “mixed” category of surviving older brothers and sisters. I also control for child’s age and for birth order.

[7] Part of this pattern may reflect innate differences in the distribution of height-by-age among rural Indian children relative to the reference standard. However, a recent study by the Nutrition Foundation of India determined that the WHO standard is applicable to Indian children (IIPS, 1995); thus, at least part of this picture reflects actual patterns of height-by-age in the sample.

[8] As these cross-tabulations are weighted by NFHS sample weights, chi-square tests are not possible.

Several variables are included to control for maternal characteristics. I measure three levels of maternal education, whether or not a woman earns cash, a mother's access to media (defined as weekly access to radio or television or monthly access to cinema), and family structure (coded as 1 if a woman lives in an extended family, defined as including her husband, children and other relatives; coded as 0 if the woman lives in a nuclear family). The literature suggests that all these variables may be associated with gender differentials in child health, primarily as indicators of women's social and economic autonomy in the household.[9] Low women's autonomy can influence both gender preferences and discriminatory behavior against unwanted girl children (Dyson and Moore, 1983; Das Gupta, 1987; Mason, 1984, 1993; and others). In the Mosley-Chen framework, women's autonomy would correspond most closely to the individual-level distal determinant of decision-making. In the modified model used in this analysis, women's autonomy variables are conceptualized as household-level variables, based on the premise that in the Indian context maternal decision-making and autonomy play out at the level of the household and are not in the sole control of the mother (Dyson and Moore, 1983).

Household socioeconomic status may also be expected to be strongly associated with child health, and directly corresponds to the household-level distal determinants in the proximate determinants framework. The direction of the association of household wealth with gender differentials is, however, less clear (Bairagi, 1986; Bhuiya, Zimicki and D'Souza, 1986; Koenig and D'Souza, 1986; World Bank, 1991; Das Gupta, 1987; Muhuri and Preston, 1991).

The NFHS does not have income or consumption data that could be used to measure household socio-economic status. The survey does, however, ask questions about the ownership of a number of consumer durables, quality of housing, ownership of land and livestock, and source of drinking water and toilet. Studies have used a variety of ways to combine such data into one or more measures of household wealth and there is little consensus in the literature about the most appropriate or accurate way to do so; however, recent studies using different methods to construct measures of household wealth from indicators such as those found in the NFHS show that such data are a reasonable proxy for long-term household wealth (Montgomery et al., 1997; Filmer and Pritchett, 1998). In this study, I use information on household ownership of consumer durables to measure household socio-economic status.[10] The survey asked questions about a household's ownership of 11 different consumer and transport-related items, which were recorded as individual dichotomous variables. I combine these into a three-category variable coded 0 if the household owns no consumer or transportation goods, 1 if it owns at least one non-luxury

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[9] Mother's marital status, which may also be a measure of maternal autonomy, is not included in the analysis because of minimal variation in this variable: less than 2 percent of mothers of sample children are widowed, divorced or separated (202 mothers for the sample used in severe stunting and 396 for the immunization sample), while the rest are reported to be currently married.

[10] Previous multivariate analysis showed that variables measuring housing quality, ownership of land or cattle, or access to water and toilet facilities did not influence the results (Pande, 1999); these variables have not been included in this analysis.

item, and 2 if it owns at least one luxury item as well. Luxury items include refrigerator, car, television, video recorder; all other items are considered non-luxury variables.[11]

Community-level economic development — corresponding to the community-level distal determinants in the Mosley-Chen framework — may also be associated with child health, though the evidence is unclear, particularly as concerns gender differentials in child health (Jain, 1985; Kishor, 1993; Murthi et al., 1995). In this analysis, village-level economic development is proxied by variables measuring access to all-weather roads and electricity; village access to health services is also controlled for.

A number of variables are included to control for social norms and cultural factors, at the individual and community levels. At the village level, I include female literacy and female labor force participation as measures of village social and economic norms that reflect the environment within which individual women's gender preferences and discriminatory behavior can occur (Mason, 1994). Marriage rules — kinship endogamy/exogamy (whether the community sanctions marrying blood relatives) and spatial endogamy/exogamy (whether marrying within the village is allowed) — are included as indicators of social norms reflecting women's position in marriage and family structure.[12] In India, exogamy (marrying outside of kin and/or community) is associated with kinship patterns that have lower women's autonomy and higher gender differentials compared to areas with endogamous marriage rules (Dyson and Moore, 1983; Kishor, 1991).

Household caste, religion, and region of residence are included as broad indicators of social and cultural norms not captured by other variables. Region is particularly important given the historically strong regional patterns of son preference and gender differentials in child mortality in India (Dyson and Moore, 1983; Kishor, 1993). States are grouped into regions based on the commonality of cultural, linguistic and kinship patterns as classified by Karve (1965).[13] As broad indicators that measure a variety of cultural and social influences, caste, religion and region can be difficult to interpret.

Table 2 below presents descriptive statistics for these explanatory variables. While the samples for severe stunting and immunization are slightly different from one another, sample distributions by

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[11] The survey asks about household ownership of sewing machine, clock or watch, sofa set, fan, radio or transistor, bicycle, motorcycle or scooter, refrigerator, television, video cassette recorder or player, and car. The division of these goods into "non-luxury" and "luxury" is based on the author's judgement and is admittedly somewhat arbitrary; however, bivariate and multivariate results suggest that the division is reasonable (Pande, 1999).

[12] Variables for village female literacy, female labor force participation and exogamy are created as weighted averages of the levels of these variables at the village and district levels, where the weights assigned to village (relative to the district) vary directly with the sample size of the village. See Pande (1999) for details.

[13] North: Bihar, Delhi, Haryana, Himachal Pradesh, Jammu, Punjab, Uttar Pradesh, West Bengal.
Central: Gujarat, Madhya Pradesh, Maharashtra, Orissa, Rajasthan
South: Andhra Pradesh, Goa, Karnataka, Kerala, Tamil Nadu
East: Assam, Arunachal Pradesh, Manipur, Meghalaya, Mizoram, Nagaland, Tripura

TABLE 2

Variable	Immunization Mean	s.d.	Min	Max	N	Severe stunting Mean	s.d.	Min	Max	N
Child's characteristics										
Female	0.49	0.50	0	1	25,549	0.49	0.50	0	1	14,715
Age in months	33.22	12.98	12	60	25,549	26.18	11.94	6	48	14,715
Birth Order	3.16	2.09	1	16	25,549	3.24	2.12	1	16	14,715
No surviving older siblings	0.27	0.45	0	1	25,549	0.26	0.44	0	1	14,715
1 surviving older sister	0.13	0.34	0	1	25,549	0.12	0.33	0	1	14,715
2+ surviving older sisters	0.08	0.27	0	1	25,549	0.08	0.27	0	1	14,715
1 surviving older brother	0.13	0.34	0	1	25,549	0.13	0.34	0	1	14,715
2+ surviving older brothers	0.07	0.25	0	1	25,549	0.07	0.25	0	1	14,715
Mixed	0.32	0.47	0	1	25,549	0.34	0.47	0	1	14,715
Mother's characteristics										
Age (in years)	26.71	5.89	14	49	25,549	26.64	5.92	14	49	14,715
Lives in extended family	0.59	0.49	0	1	25,546	0.63	0.48	0	1	14,714
Has regular access to media	0.38	0.49	0	1	25,540	0.36	0.48	0	1	14,715
Earns cash	0.20	0.40	0	1	25,500	0.16	0.36	0	1	14,693
Illiterate	0.72	0.45	0	1	25,549	0.71	0.45	0	1	14,715
Primary education	0.12	0.33	0	1	25,549	0.12	0.32	0	1	14,715
More than primary education	0.16	0.36	0	1	25,549	0.18	0.38	0	1	14,715
Household's characteristics										
Hindu	0.82	0.39	0	1	25,549	0.80	0.40	0	1	14,715
Muslim	0.13	0.34	0	1	25,549	0.13	0.34	0	1	14,715
Other	0.05	0.22	0	1	25,549	0.07	0.25	0	1	14,715
Household size	7.93	4.09	1	38	25,549	8.25	4.24	2	38	14,715
Owns no consumer goods	0.33	0.47	0	1	25,539	0.30	0.46	0	1	14,708
Owns non-luxury goods	0.55	0.50	0	1	25,539	0.56	0.50	0	1	14,708
Owns luxury goods	0.12	0.33	0	1	25,539	0.14	0.35	0	1	14,708
Village characteristics										
No health services	0.49	0.50	0	1	25,495	0.50	0.50	0	1	14,684
Government facility only	0.14	0.34	0	1	25,495	0.16	0.36	0	1	14,684
Hospital/clinic as well	0.38	0.48	0	1	25,495	0.34	0.47	0	1	14,684
All-weather roads	0.52	0.50	0	1	25,497	0.53	0.50	0	1	14,683
Electricity	0.77	0.42	0	1	25,284					
Proportion of female literacy	0.33	0.20	0	1	25,549	0.32	0.21	0	1	14,715
Proportion of women working	0.25	0.22	0	1	25,547	0.21	0.21	0	1	14,715
Proportion kinship endogamy	0.11	0.16	0	1	25,549	0.09	0.16	0	1	14,715
Proportion spatial endogamy	0.15	0.17	0	1	25,549	0.12	0.14	0	1	14,715
Region of residence										
North	0.47	0.50	0	1	25,549	0.54	0.50	0	1	14,715
Central	0.29	0.46	0	1	25,549	0.28	0.45	0	1	14,715
South	0.18	0.39	0	1	25,549	0.11	0.31	0	1	14,715
East	0.06	0.23	0	1	25,549	0.07	0.25	0	1	14,715

various background characteristics are largely similar. Forty-nine percent of the sample children are female, and about one-quarter of sample children have no surviving older siblings. A little over half the children live in extended families. Almost three-quarters (72 percent) of the mothers of sample children are illiterate, 20 percent or fewer earn cash, and fewer than 40 percent have regular access to media (television, radio or cinema). Half the children live in households that own only non-luxury goods, while a high one-third live in poor households that own no consumer goods. Almost half the sample children live in villages with no medical services, and in villages that are not accessible by an all-weather road.

Statistical methods

I use logit regression for analysis of severe stunting and ordered logit for immunization. While my unit of observation is the child, I include household and community data as explanatory variables. As a result, child-level observations are unlikely to be independent of each other, and are likely to be clustered at the household and community levels (Zenger, 1993; Liang and Zeger, 1993; Curtis et al., 1993). Previous analysis found clustering at the household level to be insignificant in this sample and thus it is not taken into account in this study (Pande, 1999). In the severe stunting logit model, I correct standard errors for clustering at the village level by correcting the variance-covariance matrix using a method developed by Huber and White and incorporated in STATA (Statacorp, 1997). Coefficient estimates are as estimated by conventional logit models. The ordered logit model presented here does not include such a correction; given the large sample size, inferences are not significantly affected for variables of interest (see Pande, 1999 for details).

The logit model estimates the conditional mean value of a dichotomous outcome (Kennedy, 1996; Hosmer and Lemeshow, 1989). The logit (L) is the log of the odds ratio and can be expressed as follows:

$$L = \ln\left[\frac{\Pr(Y=1)}{\Pr(Y=0)}\right] = x\beta$$

where β is a vector of parameters for x explanatory variables. The left-hand side of this equation represents the log odds of the outcome associated with the explanatory variable in question. The logit model used in this study for severe stunting can be described as follows:

$$\text{logit}(p_{ij}) = \alpha + G_{ij}\beta_g + S_{ij}\beta_s + G_{ij}S_{ij}\beta_{gs} + X\beta_x$$

where p_{ij} is the probability of being severely stunted for the *i*th child in the *j*th community. $G_{ij} = 1$ if *i*th child in *j*th village is female, and β_g is the main effect of gender and measures the gender differential in severe stunting after controlling for other variables in the model. S_{ij} is a vector of sex compositions of surviving older siblings for the *i*th child in the *j*th community, and β_s measures the effect of the sex composition of surviving older siblings on the probability of severe stunting for the index child, independent of other model variables. The fourth term on the right-hand side of the above equation represents an interaction of gender and sex composition. The coefficient on this term is used to test the central hypothesis of this paper, namely that the sex composition of surviving siblings affects the odds of severe stunting in gender-specific ways. *X* is a vector of other maternal, household and community level factors.

The ordered logit model is used when the outcome variable is categorized on an ordinal scale, ordered by some conceptual or subjective criteria (McCullagh and Nelder, 1989). It is the probability model used here for immunization, which is ordered from 0 (none) to 1 (any) to 2 (all vaccinations), as explained in a previous section. The ordered logit model can be expressed as:

$$\log\left[\frac{\gamma_j x}{1-\gamma_j x}\right] = \theta_j - \beta^T x,$$
$$j = 1,\ldots,k-1$$

for k categories of the response variable, where $\beta_j = \Pr(Y = \theta_j | x)$ is the cumulative probability up to and including the *j*th category, for a covariate vector x, and β_j is the cut-point for the *j*th category. Exponentiating both sides of the above equation gives the odds of falling into category *j* or lower versus falling into a category higher than *j*, with a given set of covariates. The odds ratio for a unit change in a particular covariate, say from $x=x_1$ to $x=x_2$, is given by:

$$\left[\frac{\gamma_j x_1 / (1-\gamma_j x_1)}{\gamma_j x_2 / (1-\gamma_j x_2)}\right] = \exp(-\beta_x (x_1 - x_2))$$

where β_x is the coefficient of interest (the right-hand side of this equation can be expanded to include the same vectors of independent variables as presented for the logit model earlier.) The *negative* sign on the coefficient means that a *higher* value of the variable *increases* the odds of a *lower* value of the outcome (less immunization). For example, a *negative* coefficient on the variable "gender" means that girls (gender=1) have *higher* odds for *less* immunization compared to boys.

Results

Figures 3 and 4 suggest that the patterns of gender differentials in severe stunting and immunization for the index child are associated with the sex composition of surviving older siblings in ways that are consistent with previous research on family composition and gender differentials in mortality (Das Gupta, 1987; Muhuri and Preston, 1991). Thus, all girls do not face the same level of discrimination: the first girl born after two or more boys may face less discrimination than a boy who has two or more older brothers, suggesting that parents may desire at least one daughter. On the other hand, girls born into a family that already has two or more surviving daughters and no surviving sons are among the most likely to be severely stunted (38 percent), and are less likely to be immunized than first daughters.


Figure 3: Proportion of girls and boys severely stunted by sex composition of surviving siblings
Proportion severely stunted
0.50
0.40
0.30
0.20
0.10
0.00
0.31
0.30
0.34
0.36
0.27
0.38
0.29
0.34
0.40
0.36
0.37
0.39
None
1sis
2+sis
1bro
2+bro
Mixed
Sex composition of surviving older siblings
Boys
Girls

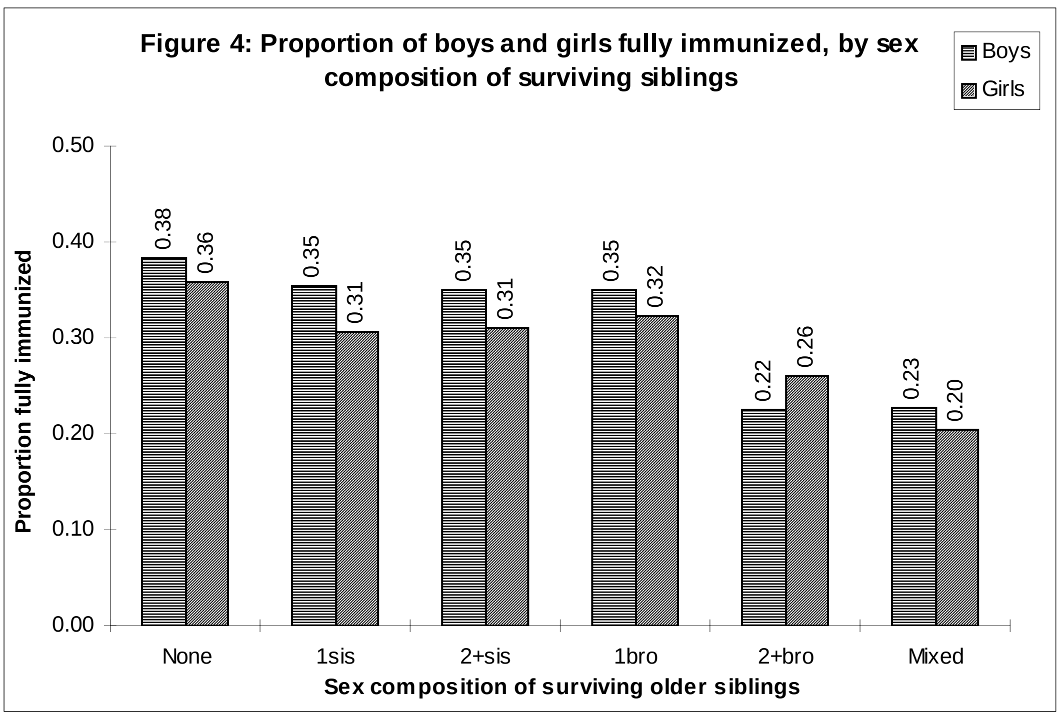
Figure 4: Proportion of boys and girls fully immunized, by sex composition of surviving siblings
Boys
Girls
Proportion fully immunized
0.50
0.40
0.30
0.20
0.10
0.00
0.38
0.36
0.35
0.31
0.35
0.31
0.35
0.32
0.22
0.26
0.23
0.20
None
1sis
2+sis
1bro
2+bro
Mixed
Sex composition of surviving older siblings

As with girls, some boys may be more “wanted” than others, and the first boy born after two or more girls fares particularly well. A boy with one older brother and no older sisters also has low risks of severe stunting and a high likelihood of being immunized; it is only after a family already has two surviving sons that a third son appears to be relatively neglected and faces the highest severe stunting risks and lowest immunization chances of all boys.

As discussed earlier in this paper, the effect of the sex composition of surviving older children on gender differentials in health status is influenced by social norms that form gender preferences for children, as well as by the household’s ability and access to resources to realize their preferences. While it is outside the scope of this paper to formally test the structural model implied by this conceptual framework, multivariate analysis results are presented that estimate a reduced-form model which controls for social norms and household and community factors that could influence gender differentials. These factors correspond to the “distal” determinants in the Mosley-Chen framework.

Tables 3 and 4 present results from the multivariate models for severe stunting and immunization, respectively. The first model in each table presents the odds of the outcome without taking into account either the gender-specific effect of the sex composition of older siblings or maternal, household and community factors. Models 2 and 5 introduce an interaction between the child’s gender and family composition, and models 3 and 6 add other characteristics. In addition to the variables noted earlier, I control for mother’s age and household size. For the most part, similar models are estimated for the two outcomes to allow qualitative comparisons across outcomes; the results for immunization also control for whether a child had an immunization card available at the time of the survey.

The simplest models (models 1 and 4) show that girls are 10 percent more likely to be severely stunted (e^{β} =exp(0.098)=1.10, p=.007 in model 1); girls also have significantly higher odds of less immunization compared to boys (e^{β} =exp(-0.152)=0.86, p<.001 in model 4). Children with two or more surviving older sisters are less likely to be severely stunted than children with no surviving older siblings. On the whole, however, family composition does not appear to be a very strong determinant of severe stunting for all children combined. For immunization, having one older sister, two or more older sisters, or one brother increases the odds of a better immunization status relative to a child with no older surviving siblings.

Adding an interaction between gender and family composition in models 2 and 5 shows the significance of the difference in the impact of family composition on health status for girls and boys; this is true of both outcomes presented here, and remains significant even after controlling for other related factors in models 3 and 6.

The main effect for gender in these models is difficult to interpret due to multiple interactions. The coefficient on gender in the full models (models 3 and 6) suggests that, all else being equal, girls and boys no longer have significantly different odds of severe stunting. However, with all other variables at

TABLE 3. Multivariate results for severe stunting

Variable	MODEL 1 Coeff.	MODEL 1 s.e.	MODEL 1 P>|z|	MODEL 2 Coeff.	MODEL 2 s.e.	MODEL 2 P>|z|	MODEL 3 Coeff.	MODEL 3 s.e.	MODEL 3 P>|z|
Gender	0.098	0.036	0.007	0.008	0.072	0.908	-0.045	0.085	0.599
Child's age (6-11)									
12-23	1.005	0.072	0.000	1.004	0.072	0.000	1.037	0.076	0.000
24-35	1.103	0.072	0.000	1.103	0.072	0.000	1.197	0.079	0.000
36-47	1.408	0.072	0.000	1.410	0.072	0.000	1.489	0.082	0.000
Birth Order	0.091	0.014	0.000	0.091	0.014	0.000	0.070	0.019	0.000
Surviving older sibs (none)									
1 sister only	0.089	0.065	0.166	0.044	0.089	0.622	0.095	0.092	0.303
2+ sisters	-0.142	0.085	0.096	-0.383	0.115	0.001	-0.216	0.119	0.070
1 brother only	-0.021	0.064	0.737	-0.085	0.089	0.338	-0.064	0.097	0.513
2+ brothers	0.020	0.087	0.817	0.187	0.116	0.106	0.203	0.119	0.090
Mixed	0.017	0.074	0.818	-0.049	0.087	0.574	0.013	0.093	0.890
Gender*1sis				0.093	0.125	0.457	0.115	0.130	0.377
Gender*2+ sis				0.488	0.150	0.001	0.475	0.152	0.002
Gender*1bro				0.129	0.123	0.296	0.226	0.132	0.087
Gender*2+bro				-0.324	0.155	0.037	-0.296	0.155	0.056
Gender*mixed				0.135	0.094	0.154	0.168	0.098	0.087
Mother's education (none)									
Primary							-0.157	0.065	0.015
More than primary							-0.368	0.066	0.000
Lives in extended family							-0.047	0.049	0.338
Has access to media							-0.062	0.048	0.199
Mother earns cash							0.137	0.057	0.016
Mother's age							-0.028	0.005	0.000
Household size							0.013	0.006	0.032
Low caste/tribe (no)							-0.047	0.054	0.380
Religion (Hindu)									
Muslim							0.125	0.069	0.071
Other							-0.402	0.083	0.000
Owns goods (none)									
Non-luxury							-0.206	0.051	0.000
Luxury							-0.676	0.088	0.000
Roads in village (no)							0.007	0.057	0.897
Health facility in village (no)							-0.093	0.056	0.096
Female literacy							-0.635	0.155	0.000
Female labor force partn.							0.303	0.150	0.043
Territorial endogamy							0.918	0.140	0.000
Kinship endogamy							0.103	0.208	0.621
Region (North)									
Central							-0.482	0.092	0.000
South							-0.968	0.125	0.000
East							-0.463	0.110	0.000
Gender*Central							0.066	0.095	0.489
Gender*South							0.234	0.126	0.064
Gender*East							-0.092	0.118	0.439
Constant	-2.166	0.076	0.000	-2.123	0.082	0.000	-0.775	0.161	0.000
Sample size	14715			14715			14646		
Log likelihood statistic	-8861.246			-8851.05			-8358.155		

TABLE 4. Multivariate results for immunization

	MODEL 4			MODEL 5			MODEL 6		
Variable	Coeff.	s.e.	P>|z|	Coeff.	s.e.	P>|z|	Coeff.	s.e.	P>|z|
Card seen	1.842	0.031	0.000	1.842	0.031	0.000	1.698	0.034	0.000
Gender	-0.152	0.024	0.000	-0.146	0.045	0.001	-0.269	0.056	0.000
Child's age (12-23)									
24-35	0.192	0.032	0.000	0.191	0.032	0.000	0.160	0.034	0.000
36-47	0.165	0.032	0.000	0.165	0.032	0.000	0.113	0.034	0.001
48-60	0.173	0.038	0.000	0.173	0.038	0.000	0.076	0.042	0.066
Birth Order	-0.154	0.010	0.000	-0.154	0.010	0.000	-0.068	0.012	0.000
Surviving older sibs (none)									
1 sister only	0.076	0.042	0.072	0.108	0.058	0.062	0.048	0.061	0.428
2+ sisters	0.252	0.056	0.000	0.307	0.072	0.000	0.184	0.076	0.016
1 brother only	0.089	0.042	0.032	0.106	0.058	0.067	0.047	0.061	0.436
2+ brothers	-0.083	0.059	0.155	-0.197	0.077	0.011	-0.227	0.082	0.006
Mixed	-0.010	0.050	0.832	-0.012	0.058	0.833	-0.083	0.061	0.174
Gender*1sis				-0.066	0.081	0.417	-0.094	0.086	0.275
Gender*2+ sis				-0.119	0.097	0.219	-0.210	0.102	0.040
Gender*1bro				-0.033	0.080	0.678	-0.067	0.085	0.432
Gender*2+bro				0.235	0.104	0.024	0.280	0.110	0.011
Gender*mixed				0.003	0.063	0.960	-0.012	0.066	0.852
Mother's education (none)									
Primary							0.344	0.040	0.000
More than primary							0.638	0.042	0.000
Lives in extended family							0.081	0.032	0.011
Has access to media							0.326	0.031	0.000
Mother earns cash							0.046	0.035	0.189
Mother's age							0.011	0.003	0.002
Household size							-0.019	0.004	0.000
Low caste/tribe (no)							-0.160	0.031	0.000
Religion (Hindu)									
Muslim							-0.456	0.043	0.000
Other							0.021	0.052	0.691
Owns goods (none)									
Non-luxury							0.221	0.031	0.000
Luxury							0.696	0.051	0.000
Roads in village (no)							0.063	0.029	0.028
Health services (none)									
Government only							-0.090	0.042	0.031
Hospital/clinic as well							0.147	0.030	0.000
Electricity in village (no)							0.455	0.036	0.000
Female literacy							1.617	0.083	0.000
Female labor force partn.							0.764	0.075	0.000
Territorial endogamy							0.069	0.085	0.417
Kinship endogamy							1.121	0.106	0.000
Region (North)									
Central							0.017	0.046	0.705
South							-0.071	0.064	0.269
East							-1.369	0.070	0.000
Gender*Central							0.137	0.062	0.026
Gender*South							0.198	0.073	0.007
Gender*East							0.245	0.084	0.003
_cut1	-0.590	0.036		-0.587	0.040		1.230	0.096	
_cut2	0.902	0.036		0.905	0.041		2.984	0.098	
Sample size	25494			25494			25110		
Log likelihood statistic	-25412.8			-25408.2			-22576.9		

their reference value, there is still a significant female disadvantage in immunization status (e^{β} =exp(-0.269)=0.76, p<.001 in model 6).

Other main effects for all sample children in both models 3 and 6 are, by and large, consistent with the literature. Thus, for example, children with educated mothers and children from wealthier households are less likely to be severely stunted and have higher odds of a better immunization status than other children; female literacy also has a strong protective role at the community level. Interestingly, children in extended families are less likely to be severely stunted and have higher odds of better immunization than children in nuclear families; results are significant for immunization even after controlling for household size, wealth and other factors. This could reflect a beneficial impact for any child of the presence of multiple caretakers in an extended family. In the case of immunization, compared to no health services, the presence of a government health facility appears to lower the odds of better immunization for children in that village while a hospital/clinic is associated with higher odds for better immunization. While these results may reflect lack of confidence in government health facilities, they could also reflect greater community wealth that is associated with both higher immunization and the presence of a hospital or clinic, and that may not be adequately controlled for by the village economic development variables used here.

The interactions between gender and family composition provide the test for the key hypothesis of this paper. Coefficients for key statistically significant interactions between gender and family composition from models 3 and 6 are presented in table 5 as odds ratios, with 95 percent confidence intervals (obtained from the variance-covariance matrix). The results in table 5 support the hypothesis that family composition influences health status, and does so differently for girls and boys, in ways that suggest that parents want some balance in the sex composition of their children.

The results below show that, despite son preference, and consistent with the bivariate analysis, certain boys are neglected. A boy with two or more older brothers has significantly lower odds of a better immunization status than the first son (OR=0.80); a boy with at least two older brothers is also 22 percent more likely to be severely stunted than the first son (borderline not statistically significant).

Table 5: Odds ratios for selected interactions between gender and family composition

	Immunization					
	2 or more older surviving sisters			2 or more older surviving brothers		
	OR	lower CI	upper CI	OR	lower CI	upper CI
Boys (ref: none)	1.20	1.04	1.40	0.80	0.68	0.94
Girls (ref: none)	0.97	1.14	0.83	1.05	1.24	0.89
Girls (ref: boys)	0.62	0.75	0.51	1.01	1.24	0.83
	Severe stunting					

	2 or more older surviving sisters			2 or more older surviving brothers		
	OR	lower CI	upper CI	OR	lower CI	upper CI
Boys (ref: none)	0.81	0.64	1.02	1.22	0.97	1.55
Girls (ref: none)	1.30	1.63	1.03	0.91	1.14	0.73
Girls (ref: boys)	1.61	2.01	1.29	0.74	0.92	0.60

On the other hand, boys born into households that have only girls have significantly better health outcomes than the reference category. In particular, boys born into a household with no other boys and with two or more surviving girls appear to be the most "wanted". Such boys have the highest odds of better immunization (OR=1.20), and among the lowest odds of severe stunting (OR=0.81, marginally non-significant), of children with any combination of surviving older siblings.

Similarly, all girls do not face equal discrimination. Girls with multiple surviving sisters face more neglect than other children, a result consistent with Das Gupta's (1987) findings about mortality differentials in the Punjab in India. Compared to girls with no surviving older siblings, a third or higher birth-order daughter is 30 percent more likely to be severely stunted; immunization odds are not significantly different, however, for first and subsequent daughters. The gender differential is also significant. A girl born into a household with two or more surviving girls is 61 percent more likely to be severely stunted compared to a boy born after two or more surviving girls, and has significantly lower odds of a better immunization status compared to a boy with two or more older sisters (OR=0.62).

Parents' desire for a balanced sex composition does not appear to provide the same degree of protection to girls with opposite-sex siblings as it does for boys with opposite-sex siblings, reflecting the strength of son preference in this population. A girl born into a household that has only boys (at least two) does not have significantly different odds for severe stunting and immunization than a girl with no surviving older siblings. The gender differential does, however, present mild evidence that a girl born into a household with two or more surviving boys and no other girls may be preferred to a boy born into such a household: girls with two or more surviving brothers are about 25 percent less likely to be severely stunted than boys with two or more surviving brothers (OR=0.74).

Most of the other background maternal, household and other factors included in the full models (models 3 and 6 on tables 2 and 3 respectively) did not have statistically significant gender-specific effects on the odds of severe stunting and immunization and these interactions were dropped from the final models shown here. The interaction with region, however, was significant. In particular, girls in the South are less likely to be severely stunted (OR=exp(-0.97+0.23)=0.48) and have higher odds of better immunization (OR=exp(-0.071+0.20)=1.14) than girls in the North, consistent with the literature on higher son preference and greater discrimination against girls in northern than in southern India (Dyson and Moore, 1983; Arnold et al, 1996; Govindaswamy and Ramesh, 1996).

Discussion

Despite long-standing persistence of gender differentials in mortality on the Indian sub-continent, patterns of gender discrimination in the proximate determinants of child mortality have not received adequate attention in the literature. This paper focused on this gap by examining gender differentials in severe stunting and immunization, and in particular, the effects of family sex composition on gender discrimination for surviving girls and boys for these two outcomes.

The overall results show that severe stunting and lack of immunization are serious health problems among rural Indian children. In addition, both severe stunting and immunization appear to be important mechanisms of gender differentials in health status among surviving children.

It also appears to be the case, as hypothesized, that not all girls and boys are treated equally and there is evidence of patterns of selective neglect in the case of severe stunting and immunization that are consistent with the literature on mortality differentials (Muhuri and Preston, 1991; Das Gupta, 1987), and that persist even after maternal, household and community factors are taken into account. Thus, both girls and boys with only surviving siblings of the opposite sex fare better than children with no surviving older siblings. Conversely, children with two or more surviving same sex siblings are worse off in terms of these two health outcomes. However, the strength of son preference and the low value of girls is still evident, and the harmful effect of having surviving older siblings of the same sex alone is harsher for girls than it is for boys, while the protective effect of having only opposite-sex surviving older siblings is weaker for girls than is the case for boys.

The strength of these results for immunization have particularly sobering policy implications for programs such as the global Expanded Program on Immunization (EPI), which has been in place in India for many decades. Given that immunization is provided free of cost, and is provided through a mass program that should theoretically be accessible to all eligible children, the female disadvantage such as we see here is somewhat surprising. These results reflect the fact that there are costs to immunization other than monetary costs, for example opportunity costs of time taken to vaccinate a child or to take care of side-effects, that may contribute to a female disadvantage. The analysis suggests that, where immunization is seen as a life-saving or positive health practice, mothers do in fact use immunization opportunities to improve the odds of good health primarily for preferred children.

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