SELECTIVE GENDER DIFFERENCES IN CHILDHOOD NUTRITION AND IMMUNIZATION IN RURAL INDIA: THE ROLE OF SIBLINGS*

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This article examines the role of the sex composition of surviving older siblings on gender differences in childhood nutrition and immunization, using data from the National Family Health Survey, India (1992–1993). Logit and ordered logit models were used for severe stunting and immunization, respectively. The results show selective neglect of children with certain sex and birth-order combinations that operate differentially for girls and boys. Both girls and boys who were born after multiple same-sex siblings experience poor outcomes, suggesting that parents want some balance in sex composition. However, the preference for sons persists, and boys who were born after multiple daughters have the best possible outcomes.

Gender differences 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. Discrimination against *surviving* girls, however, tends to be presumed from the evidence about the excess mortality of girls, and the nature of such discrimination has not received adequate attention. In addition, while studies in India and Bangladesh have found that mortality discrimination is "selective" (Das Gupta 1987) and that girls who are born into households with many daughters fare particularly badly (Muhuri and Preston 1991), there has been little research on the role of the sex composition of siblings in selective discriminatory practices that affect the health of surviving children. This gap persists despite the recognition that discrimination in morbidity, nutritional status, or use of health care is likely to contribute to greater gender differences in mortality (Arnold, Choe, and Roy 1998; Bardhan 1974, 1982; Kishor 1993, 1995; Kurz and Johnson-Welch 1997; Makinson 1994; Miller 1981; Obermeyer and Cardenas 1997; Waldron 1987).

This article addresses some of these gaps in the literature by examining gender differences in immunization and severe stunting among surviving rural Indian children under age 5. In particular, I focus on the effects of the sex composition of surviving older siblings on such gender differences.

BACKGROUND

Excess female child mortality exists in many areas of the developing world (Arnold 1992, 1997; Hill and Upchurch 1995; Tabutin and Willems 1995). Arnold (1997) found that 27 out of 44 Demographic and Health Survey (DHS) countries had higher girls' than boys' mortality for children aged 1–4, although the average excess female mortality over all 44 countries was only 2%. There are distinct geographic patterns, and studies have pointed to the Middle East, North Africa, East Asia, and South Asia as having the largest gender differences (Arnold 1992, 1997).

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In India, excess female mortality and a preference for sons has existed for centuries. In the past 100 years, population sex ratios from censuses have shown more or less a continuous increase (Visaria and Visaria 1983, 1995; Visaria 1967, 1969), rising from 103 males per 100 females in 1901 to 107.2 males per 100 females in 2001 (Banthia 2001; Desai 1994; Visaria and Visaria 1995). Much of this excess female mortality occurs in childhood, and Indian censuses since 1872 have noted masculine sex ratios among children under age 12 (Miller 1981). From 1961 to 1991, sex ratios for children under age 10 became more masculine all across India (Bhat 1989; Das Gupta and Bhat 1997; Desai 1994; El-Badry 1969; Miller 1989; Parasuraman and Roy 1991). In 1992–1993, child mortality for girls in India as a whole, at 42.0 per 1,000, was 43% higher than for boys, at 29.4 per 1,000 (International Institute for Population Sciences, IIPS 1995:218). Studies from other South Asian countries have provided evidence of discrimination against girls as well (Bairagi 1986; Caldwell and Caldwell 1990; D'Souza and Chen 1980; Faisel, Ahmad, and Kundi 1993; Koenig and D'Souza 1986).

In contrast, there has been little research on gender differences in health among *surviving* girls and boys. This is an important issue because gender discrimination[1] that contributes to poorer health status for girls than for boys is likely to be the main pathway for excess female child mortality. Moreover, such discrimination, even when it is not fatal, can still produce greater frailty among survivors and thus is an important child health issue in itself (Mosley and Becker 1991; Mosley and Chen 1984). Finally, poor health has implications for surviving girls' health in 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 noted 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 1981:67).

Recent evidence has corroborated these findings. Studies in north and south India have found that boys are much more likely than girls to be taken to a health facility when sick (Caldwell, Reddy, and Caldwell 1982; Caldwell and Caldwell 1990; Das Gupta 1987; Ganatra and Hirve 1994; Govindaswamy and Ramesh 1996; Kishor 1995; Ravindran 1986; Visaria 1988). In 1992–1993, boys had higher immunization rates than did girls in all but two states in India (Goa and Karnataka), although the extent of this difference varied by state (Kurz and Johnson-Welch 1997). Similarly, girls are more likely to be malnourished than boys in both northern and southern states (Arnold et al. 1998; Basu 1989; Caldwell and Caldwell 1990; Das Gupta 1987; Pebley and Amin 1991; Sen and Sengupta 1983; Wadley 1993). Research in Bangladesh and Pakistan indicated similar discriminatory outcomes (Baqui et al. 1998; Chen, Huq, and D' Souza 1981; Faisel et al. 1993; Koenig and D'Souza 1986).

The persistence of such discrimination against girls in South Asia, particularly in India, stems from the perceived greater economic, social, and religious utility of sons than of daughters. Parents of girls are socially bound to find suitable husbands for them 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 (Arnold et al. 1998; Dyson and Moore 1983; Kishor 1995; Mandelbaum 1988). In contrast,

1. Such discrimination is likely to be a combination of "active" discrimination (such as a deliberate choice to provide health care to a sick boy but not to a sick girl), "passive" neglect (e.g., discovering that a girl is sick later than would be the case for a boy, simply because girls may be more ignored in day-to-day interactions than are boys), and what may be termed "selective favoritism" (choices made by resource-constrained families that favor those children that the family can ill afford to lose). This article does not analyze which of these dynamics is at play, and the terms *selective neglect* and *selective discrimination* are used to encompass any of these dynamics.

parents expect sons to provide financial and emotional care and regard them as a "social security" for old age; inheritance laws largely favor sons; and sons 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 (Agarwal 1994; Cain 1988, 1993; Cain, Khanam, and Nahar 1979; Dharmalingam 1996; Oldenburg 1992; Singh 1993; Vlassoff 1990). Although these social, religious and economic rules are more strongly associated with kinship structures in north than in south India (Dyson and Moore 1983; Karve 1965; Sopher 1980), the more recent studies reviewed earlier found some blurring of North–South distinctions in gender discrimination.

Research has found that coexisting with the preference for sons is the desire for a balanced sex composition and that parents may desire at least one daughter. Parents may selectively discriminate against children with certain sex–birth-order characteristics to attain their desired sex composition or may selectively favor certain preferred children whom the household cannot afford to lose (Scrimshaw 1978). Where a strong preference for sons exists, as in much of India, the preferred sex composition often includes two sons and one daughter, and the daughter is seen to provide important religious, social, or emotional value (Dharmalingam 1996; Miller 1989; Mutharayappa et al. 1997; Williamson 1976). On the other hand, more than one or two daughters are usually not welcome, and girls who are born into a family that already has daughters are the most likely to be least valued and thus discriminated against by the household.

Studies in northern India, Bangladesh, and China have shown that girls with older sisters have significantly higher mortality risks than do girls with only surviving older brothers or girls with no surviving older siblings (Amin 1990; Arnold et al. 1998; Choe, Hongsheng, and Feng 1995; Das Gupta 1987; Muhuri and Preston 1991). The preference for sons can also result in girls with older brothers being neglected in households that favor existing sons over newborn daughters or families who do not want daughters at all (Simmons et al. 1982). On the other hand, the preference for sons means that any harmful effect of same-sex siblings for boys may be weaker than it is for girls (Muhuri and Preston 1991). In India, a continuing desire for 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 1997).

Although research has documented the extent and nature of such selective gender discrimination in mortality differences, there is less clear evidence about the nature of such discrimination among surviving children for whom such neglect is not yet fatal but may nonetheless contribute to a poorer health status. This is the central focus of this article.

FRAMEWORK AND HYPOTHESES

I modeled the effects of sex composition on gender differences in health status among surviving children using a modified proximate-determinants framework (Mosley 1985; Mosley and Chen 1984). In this framework, originally developed to study factors that affect child mortality, Mosley and Chen (1984) proposed five categories of biological and medical proximate determinants of child mortality, 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 "distal" determinants of child mortality operate through these proximate determinants and were grouped by Mosley and Chen into individual-level, household-level, and community-level variables.

My approach drew on the proximate-determinants framework inasmuch as I examined "health status," but with the health status of survivors, specifically, nutrient deficiency and personal (preventive) illness control—rather than mortality—as the outcome of interest. My model thus differed in that I analyzed proximate determinants as outcomes themselves. Gender differences in the proximate determinants represent ways in which household attitudes and behavior against girls translate into a worse health status for surviving girls than for boys. Figure 1 illustrates this modified model. The figure shows all the proximate determinants in the complete model; the determinants used as outcomes for this analysis 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, as in the original framework, although I did not analyze this relationship in my study.

I used severe stunting as my measure of nutrient deficiency. Although I recognize that severe stunting can also reflect the effects of recurrent, untreated infection, by its very definition, it is likely to be a more reliable marker of long-term nutritional neglect of unwanted children (usually girls) than are other nutrition indicators, which are more likely to be markers of acute—but not necessarily chronic or sustained—malnutrition. Immunization was used as a measure for personal illness control, specifically, preventive health care, and is indicative of parents' motivation to ensure a child's future well-being and health. Parents are likely to be motivated to care 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, low financial and other costs may encourage parents to immunize all their children, and hence gender differences may not be observed.

I grouped the distal determinants (the left side of Figure 1) differently from the proximate-determinants framework to reflect the characterization in the literature of key influences on the preference for sons and gender differences in health. Gender differences and the effect of sex composition on these differences are hypothesized to be determined by, first, individual and social norms and beliefs that influence parents' gender preferences and thereby the desired sex composition of children. Second, the extent to which parents' 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, sex-composition effects on gender differences may be seen when gender preferences exist *and* when there is a capacity to realize these preferences (for a fuller theoretical explanation, see Pande 1999).

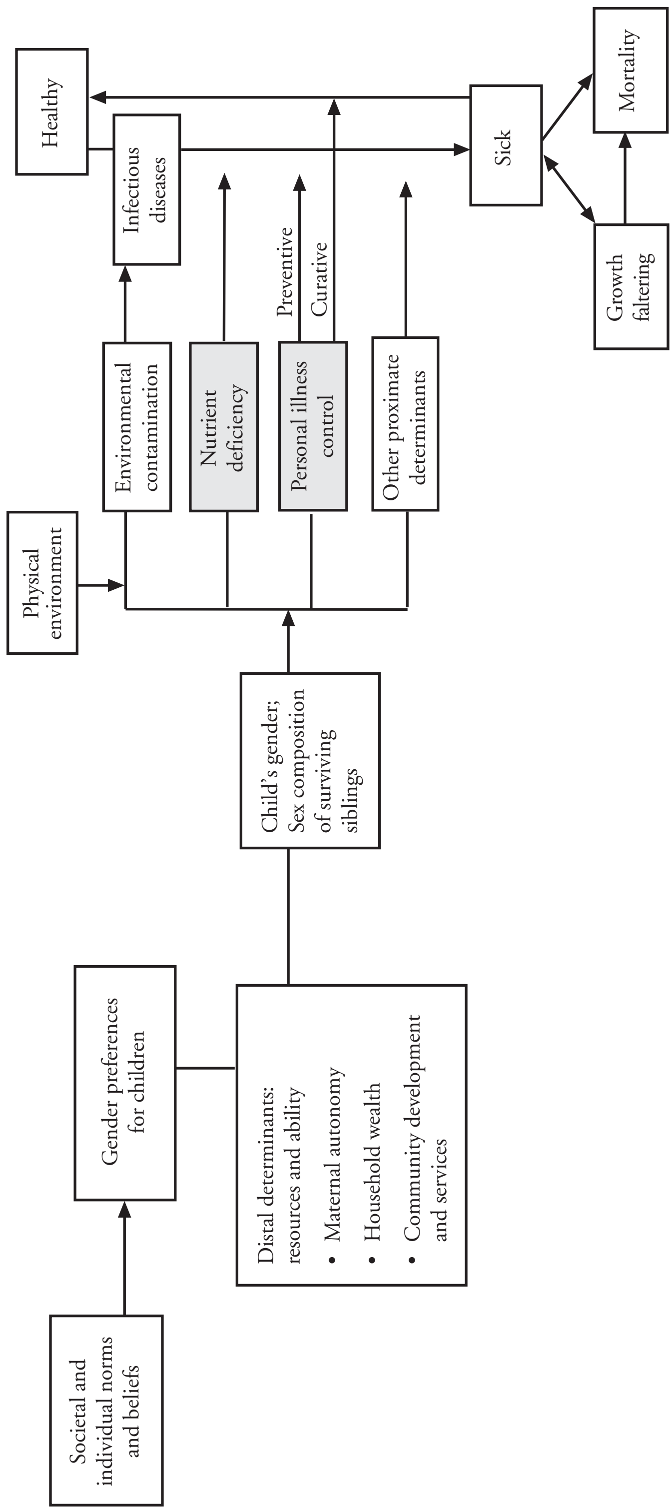
This modified Mosley-Chen framework specifies the theoretical relationship between gender differences among surviving children and gender differences in mortality. It also provides a conceptual framework that relates child-level determinants of child health to "distal" contextual factors, such as maternal education, household wealth, and community characteristics that are known to influence child health (Bhuiya and Streatfield 1992; Bhuiya, Wojtyniak, and Karim 1989; Caldwell 1979; Cleland and van Ginneken 1988; Desai and Alva 1998; Hobcraft 1993; Sastry 1996; Schultz 1984). Structural estimation of the full model is beyond the scope of this article; here, I estimated 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 sex composition, regardless of gender preference, I expected children whose surviving older siblings are all 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 a strong preference for sons and the household's capacity to realize such a preference, I expected boys with same-sex siblings to do better than girls with same-sex siblings and girls with only surviving older sisters to have the worst health status of all combinations of siblings.

DATA

I used the rural sample of the National Family Health Survey (NFHS), India, 1992–1993. The NFHS followed the format of the Demographic and Health Surveys (DHS),

Figure 1. Conceptual Framework



large-scale household surveys conducted in Asia, Africa and Latin America. It used three questionnaires: the Women's Questionnaire, which asked for detailed health and demographic information from married women in their reproductive years, including information about their children; the Household Questionnaire, which included demographic and socioeconomic information on respondents' households; and the Village Questionnaire, which had data on facilities in respondents' villages of residence. The survey was conducted separately in each state, with a uniform sampling design across states. A multistage, systematic, stratified sampling design was used, with the primary sampling units (usually, but not always, comprising a village in rural areas and an urban block in urban areas) selected systematically, with probability proportional to size. Households were then sampled using systematic sampling with equal probability, and all eligible women in each household were interviewed. National- and state-level sampling weights reflect the sampling design (IIPS 1995).

I used the sample of children born to de jure, rural, ever-married 13- to 49-year-old women and merged data across the Women's, Household, and Village Questionnaires. I used the rural sample because the Village Questionnaire, which allowed me to include community characteristics in the model, existed only for rural areas. Because about 73% of India's population is rural (IIPS 1995:xxxi), the results are still applicable to the majority of the population. I used the de jure sample (all residents), rather than the de facto sample (all those who slept in the sampled household the previous night), because detailed data on household structure and villages are available only for de jure women.

Dependent Variables

Severe stunting is defined as a height-for-age measurement more than 3 standard deviations below the international reference population that was recommended by the World Health Organization (WHO) and found to be applicable to Indian children by the Nutrition Foundation of India (IIPS 1995:281). My sample comprises 14,715 rural surviving children aged 6–47 months for whom height-for-age data were available. I chose a minimum of age 6 months because little gender discrimination seems to occur in infancy (Bhuiya and Streatfield 1992; Muhuri and Preston 1991). I measured severe stunting as a dichotomous variable (1 = severely stunted, as was just defined). Complete data on age were available for over 98% of the children (IIPS 1995:322–23), but some births that occurred 0–4 years before the survey tended to be reported erroneously as occurring 5–9 years before the survey (Mishra, Lahiri, and Luther 1999). Thus, estimates of severe stunting may be biased upward to the extent that this age exaggeration occurs. To the extent that such age exaggeration is greater among girls, the gender differences may be exaggerated. Heights were not measured in some states because of the lack of measuring boards, but other analyses suggested that the lack of height measurements for children in these states is unlikely to bias the national average (IIPS 1995:281; Mishra et al. 1999).[2] In the remaining states, data on heights were available for 81.5% of the eligible children. An exploratory analysis showed that children with and without height measurements were similar across key characteristics, including the sex ratio (Pande 1999).

For immunization, I used a sample of 25,549 rural surviving children aged 12–60 months. Children aged 12 months and older were chosen so that every child had adequate exposure time for the full range of vaccinations, as per the WHO schedule for immunizations.[3] The upper age limit was chosen to maximize the use of available data, given that the NFHS had data on immunizations for children up to 60 months old. Vaccination coverage

2. Data on heights were unavailable for Andhra Pradesh, Himachal Pradesh, Madhya Pradesh, Tamil Nadu, and West Bengal.

3. As per the WHO schedule, children should receive the BCG vaccine at birth or soon after; three doses each of DPT and polio vaccine at 6, 10, and 14 weeks of age; and the measles vaccine at 9 months or soon thereafter (WHO 1984).

was estimated by combining data from immunization cards and information from the mothers' recall when immunization cards were not available. Studies elsewhere have shown that mothers' recall of their children's immunizations is fairly accurate (Langsten and Hill 1998). Immunization was 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).

Explanatory Variables

Gender was a key dichotomous variable, coded 1 if the child was female. Various combinations of the sex composition of surviving older siblings were 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 controlled for the child's age and for birth order.

I measured three levels of maternal education, whether or not a woman earned cash, a mother's access to media (weekly access to radio or television or monthly access to movies), and family structure (1 if a woman lived in an extended family, defined as including her husband, children, and other relatives; 0 if the woman lived in a nuclear family). The literature suggests that all these variables may be associated with gender differences in child health, primarily as indicators of women's social and economic autonomy in the household. Women's low autonomy can influence both gender preferences and discriminatory behavior against unwanted daughters (e.g., Das Gupta 1987; Dyson and Moore 1983; Mason 1984, 1993). In the Mosley-Chen framework, women's autonomy would correspond to the individual-level distal determinant of decision making. In my modified model, women's autonomy was conceptualized as a household-level determinant because in the Indian context, maternal decision making and autonomy are often not in the sole control of the mother (Dyson and Moore 1983).

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

The NFHS did not have income or consumption data, but it did have data on household asset ownership, which I used to measure household socioeconomic status. Recent studies have shown that such data are a reasonable proxy for long-term household wealth (Filmer and Pritchett 2001; Montgomery et al. 2000). Data on ownership of 11 consumer and transport-related items were used to create a three-category variable, coded 0 if the household owned no consumer or transportation goods, 1 if it owned at least one nonluxury item, and 2 if it owned at least one luxury item as well. Luxury items included a refrigerator, car, television, and video recorder; all other items were considered nonluxury items.[4] I also controlled for mother's age and household size.

Community-level economic development—the community-level distal determinants in the Mosley-Chen framework—may also be associated with child health, although the evidence is unclear, particularly for gender differences (Jain 1985; Kishor 1993; Murthi, Guio, and Dreze 1995). I used measures of access to all-weather roads and electricity as proxies for village-level economic development; I also controlled for village access to health services.

4. Nonluxury items included a sewing machine, clock or watch, sofa set, fan, radio or transistor, bicycle, motorcycle, or scooter. The division of these goods into "nonluxury" and "luxury" was based on my judgment and was admittedly somewhat arbitrary; however, bivariate and multivariate results suggested that the division is reasonable.

To measure social norms and cultural factors at the village level, I used female literacy and female labor-force participation as measures of the immediate social and economic environment in which individual women's gender preferences and discriminatory behavior occur (Mason 1994). I used endogamous/exogamous marriage rules as indicators of social norms reflecting women's position in marriage and the family structure. Kinship endogamy (marrying blood relatives) was directly measured; spatial endogamy (marrying within the village) was measured by whether the woman currently resided in her village of birth. In India, exogamy is associated with kinship patterns that include women's lesser autonomy and higher gender differences than in areas with endogamous marriage rules (Dyson and Moore 1983; Kishor 1993). These village-level social and cultural variables were weighted (see the appendix for the details).

Household caste, religion, and region of residence were included as indicators of social and cultural norms that were not captured by other variables. Region was particularly important, given the historically strong regional patterns of preferences for sons and gender differences in child mortality in India (Dyson and Moore 1983; Kishor 1993). States were grouped into regions on the basis of the commonality of cultural, linguistic, and kinship patterns, as classified by Karve (1965).[5] The cultural and social influence of caste, religion, and region are complex and multifaceted, so the associated coefficients can be difficult to interpret.

STATISTICAL METHODS

I used logit regression to analyze severe stunting and ordered logit for immunization. Because I included household and community data as explanatory variables, child-level observations are unlikely to be independent of one another and are likely to be clustered at the household and community levels (Curtis, Diamond, and McDonald 1993; Liang and Zeger 1993; Zenger 1993). Exploratory analysis found clustering at the household level to be insignificant, and thus it is not included in the results presented here. In the severe-stunting logit model, I corrected 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 were estimated by conventional logit models. The ordered logit model for immunization did not include such a correction, but a previous analysis suggested that the inferences were not significantly affected for the variables of interest (Pande 1999).

The logit model estimates the conditional mean value of a dichotomous outcome (Hosmer and Lemeshow 1989; Kennedy 1996). The logit model used in this study for severe stunting can be described as

$$\text{logit}(p_{ij}) = \alpha + G_{ij}\beta_g + S_{ij}\beta_s + G_{ij}S_{ij}\beta_{gs} + X\beta_x, \qquad (1)$$

where p_{ij} is the probability of being severely stunted for the ith child in the jth community. $G_{ij} = 1$ if the ith child in the jth village is female, and β_g is the main effect of gender and measures the gender difference in severe stunting after other variables in the model are controlled. $\mathbf{S_{ij}}$ is a vector of the sex composition of surviving older siblings for the ith child in the jth 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 variables in the model. The fourth term on the right-hand side of Eq. (1) represents an interaction of gender and sex composition. The coefficient on this term was used to test the central hypothesis of this article, namely, that the sex composition of surviving

5. *North* includes Bihar, Delhi, Haryana, Himachal Pradesh, Jammu, Punjab, Uttar Pradesh, and West Bengal; *Central* includes Gujarat, Madhya Pradesh, Maharashtra, Orissa, and Rajasthan; *South* includes Andhra Pradesh, Goa, Karnataka, Kerala, and Tamil Nadu; and *East* includes Assam, Arunachal Pradesh, Manipur, Meghalaya, Mizoram, Nagaland, and Tripura.

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 previously. The ordered logit model can be expressed as

$$\log[\gamma_j x / 1 - \gamma_j x] = \theta_j - \beta^T x, \text{ where } j = 1,\ldots,k - 1 \quad (2)$$

for k categories of the response variable, where $\gamma_j = \Pr(Y \leq \theta_j | x)$ is the cumulative probability up to and including j, for a covariate vector **x**, and θ_j is the cut-point for the jth category. Exponentiating both sides of the 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

$$[\gamma_j x_1 / (1 - \gamma_j x_1)] / [\gamma_j x_2 / (1 - \gamma_j x_2)] = \exp(-\beta_x(x_1 - x_2)), \quad (3)$$

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

RESULTS

Univariate and Bivariate Analyses

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 (see Table 1). Slightly over one-third (34.4%) of the children in the sample were severely stunted. Only 30% of the children aged 12–60 months were fully immunized, and close to 40% had not received any of the vaccinations.

There is some evidence of gender differences in nutrition, and 6% more girls than boys were severely stunted. In an all-India analysis of NFHS data, Mishra et al. (1999) found no gender discrimination in moderate and severe stunting combined. The rural

Table 1. Severe Stunting and Immunization, Total and by Gender of the Child, Rural India: 1992–1993

	Outcome Measure						
Sample	Severe Stunting	No Immunizations	BCG	DPT3	OPV3	Measles	Fully Immunized
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
GD[a]	6.25	13.00	–8.49	–9.00	–8.00	–9.62	–8.53

Notes: The total sample size is 14,715 for severe stunting and 25,549 for immunization. Because NFHS sample weights were used, chi-square tests were not possible.

[a]GD is the gender difference, defined as the percentage difference between girls and boys.

Table 2. Description of Explanatory Variables for Analysis of Immunization and Severe Stunting in Rural India: 1992–1993

	Immunization					Severe Stunting				
Variable	Mean	*SD*	Min.	Max.	*N*	Mean	*SD*	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
Surviving Older Siblings										
None	0.27	0.45	0	1	25,549	0.26	0.44	0	1	14,715
1 older sister	0.13	0.34	0	1	25,549	0.12	0.33	0	1	14,715
2+ older sisters	0.08	0.27	0	1	25,549	0.08	0.27	0	1	14,715
1 older brother	0.13	0.34	0	1	25,549	0.13	0.34	0	1	14,715
2+ 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

(continued)

gender difference in severe stunting alone found here suggests that gender differences in nutrition are apparent in only the most sustained neglect, as is found in severe stunting. Gender differences in immunization were larger, with 8%–9% fewer girls than boys having received any or all vaccinations. The largest difference was among unvaccinated children, with 13% more girls than boys being totally unvaccinated.

Table 2 presents descriptive statistics for the explanatory variables. Sample distributions are similar for the two outcomes: 49% of the children in the sample were female, and about one-quarter had no surviving older siblings. A little over half the children lived in extended families. In addition, 72% of the mothers of the children in the sample were illiterate, 20% or fewer earned cash, and fewer than 40% had regular access to media. Half the children lived in households that owned only nonluxury goods, while one-third lived in households that owned no consumer goods. Almost half the children lived in villages with no medical services and no all-weather road.

Figures 2 and 3 suggest that the patterns of gender differences 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 research on sex composition and gender differences in mortality (Das Gupta 1987; Muhuri and Preston 1991). All girls do not face

(Table 2, continued)

	Immunization					Severe Stunting				
Variable	Mean	*SD*	Min.	Max.	*N*	Mean	*SD*	Min.	Max.	*N*
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 nonluxury 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
PHC/subcenter 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

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. On the other hand, girls who were 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%) and are less likely to be immunized than are first daughters.

As with girls, some boys may be more "wanted" than others, and the first boy who is 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 the third son appears to be relatively neglected and faces the highest risk of severe stunting and lowest chance of immunization of all boys.

Figure 2. Severe Stunting, by Gender and Sex Composition of Surviving Older Siblings, Children Aged 6–47 Months in Rural India: 1992–1993

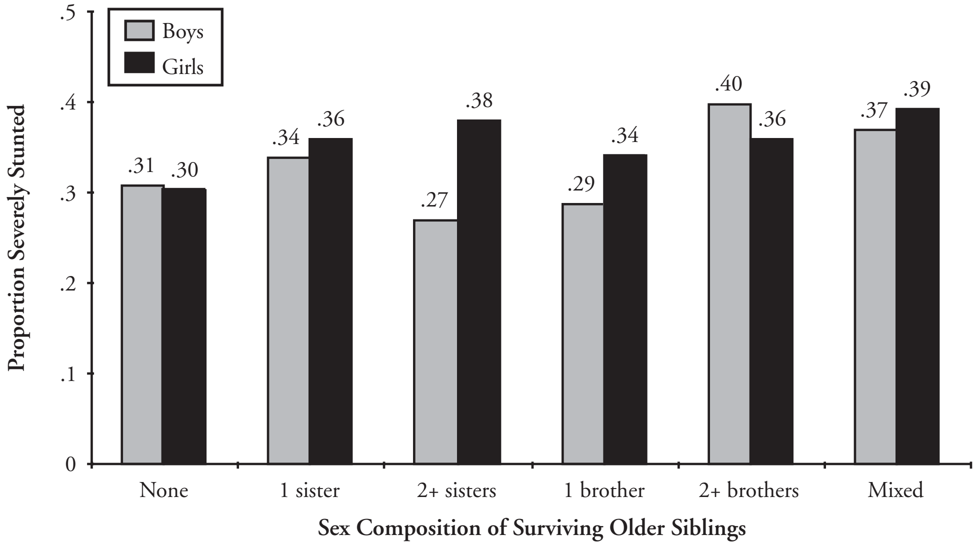


Figure 3. Full Immunization, by Gender and Sex Composition of Surviving Older Siblings, Children Aged 12–60 Months in Rural India: 1992–1993

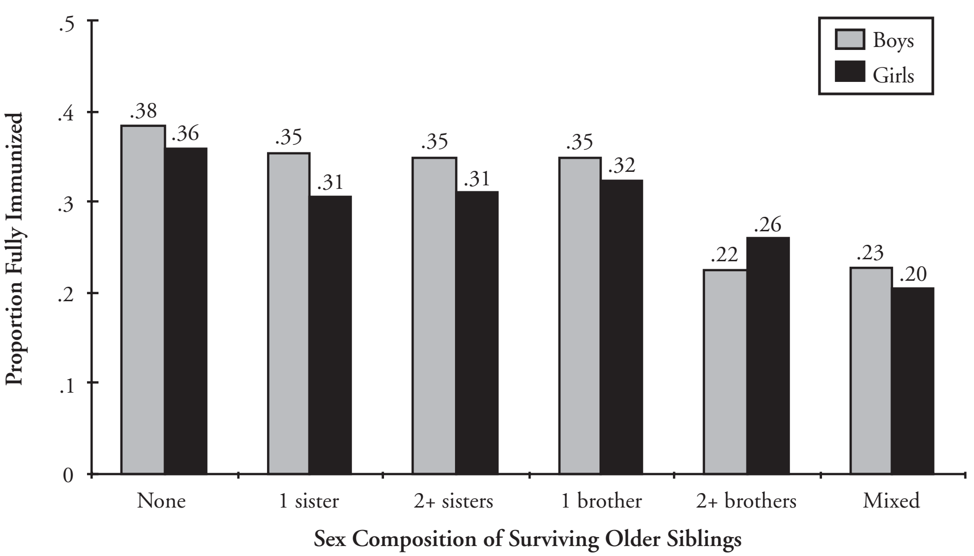


Table 3. Logit Model for Severe Stunting Among Children Aged 6–47 Months in Rural India: 1992–1993

	Model 1			Model 2			Model 3		
Variable	Coeff.	*SE*	*P* > |*z*|	Coeff.	*SE*	*P* > |*z*|	Coeff.	*SE*	*P* > |*z*|
Intercept	–2.166	0.076	.000	–2.123	0.082	.000	–0.775	0.161	.000
Gender	0.098	0.036	.007	0.008	0.072	.908	–0.045	0.085	.599
Child's Age (6–11)									
12–23 months	1.005	0.072	.000	1.004	0.072	.000	1.037	0.076	.000
24–35 months	1.103	0.072	.000	1.103	0.072	.000	1.197	0.079	.000
36–47 months	1.408	0.072	.000	1.410	0.072	.000	1.489	0.082	.000
Birth Order	0.091	0.014	.000	0.091	0.014	.000	0.070	0.019	.000
Surviving Older Siblings (none)									
1 sister only	0.089	0.065	.166	0.044	0.089	.622	0.095	0.092	.303
2+ sisters	–0.142	0.085	.096	–0.383	0.115	.001	–0.216	0.119	.070
1 brother only	–0.021	0.064	.737	–0.085	0.089	.338	–0.064	0.097	.513
2+ brothers	0.020	0.087	.817	0.187	0.116	.106	0.203	0.119	.090
Mixed	0.017	0.074	.818	–0.049	0.087	.574	0.013	0.093	.890
Gender × 1 sister				0.093	0.125	.457	0.115	0.130	.377
Gender × 2+ sisters				0.488	0.150	.001	0.475	0.152	.002
Gender × 1 brother				0.129	0.123	.296	0.226	0.132	.087
Gender × 2+ brothers				–0.324	0.155	.037	–0.296	0.155	.056
Gender × mixed				0.135	0.094	.154	0.168	0.098	.087

(continued)

Multivariate Analysis

As I discussed earlier, the effect of the sex composition of surviving older children on gender differences 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. Although it is outside the scope of this article to test formally the structural model implied by this conceptual framework, I present the results of a multivariate analysis for a reduced-form model that controlled for social norms and household and community factors that could influence gender differences. To examine whether the effects of the gender and sex composition of children varied by region or other key characteristics, I ran initial multivariate analyses that included separate models for boys and girls and by region. Most two-way and three-way interactions of other variables with gender and sex composition suggested by these models were not statistically significant, and thus single models with interactions between gender and sex composition and between gender and region were used for both outcomes. Only the final models are presented here.

Tables 3 and 4 present the results of the final multivariate models for severe stunting and immunization, respectively. The first model in each table presents the odds of the

(Table 3, continued)

	Model 1			Model 2			Model 3		
Variable	Coeff.	*SE*	*P* > |*z*|	Coeff.	*SE*	*P* > |*z*|	Coeff.	*SE*	*P* > |*z*|
Mother's Education (none)									
Primary							–0.157	0.065	.015
More than primary							–0.368	0.066	.000
Lives in Extended Family							–0.047	0.049	.338
Has Access to Media							–0.062	0.048	.199
Mother Earns Cash							0.137	0.057	.016
Mother's Age							–0.028	0.005	.000
Household Size							0.013	0.006	.032
Low Caste/Tribe							–0.047	0.054	.380
Religion (Hindu)									
Muslim							0.125	0.069	.071
Other							–0.402	0.083	.000
Owns Goods (none)									
Owns nonluxury goods							–0.206	0.051	.000
Owns luxury goods							–0.676	0.088	.000
Roads in Village							0.007	0.057	.897
Health Facility in Village							–0.093	0.056	.096
Female Literacy							–0.635	0.155	.000
Female Labor-Force Participation							0.303	0.150	.043
Territorial Endogamy							0.918	0.140	.000
Kinship Endogamy							0.103	0.208	.621
Region (North)									
Central							–0.482	0.092	.000
South							–0.968	0.125	.000
East							–0.463	0.110	.000
Gender × Central							0.066	0.095	.489
Gender × South							0.234	0.126	.064
Gender × East							–0.092	0.118	.439
N		14,715			14,715			14,646	
Log-Likelihood Statistic	–8,861			–8,851.1			–8,358.2		

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 sex composition, and Models 3 and 6 add other characteristics. Similar models are estimated for the two outcomes to allow qualitative comparisons across outcomes.

The simplest models (Models 1 and 4) show that girls are 10% more likely to be severely stunted (e^{β} = exp(0.098) = 1.10, p = .007 in Model 1) and have significantly

Table 4. Ordered Logit Model for Immunization of Children Aged 12–60 Months in Rural India: 1992–1993

	Model 4			Model 5			Model 6		
Variable	Coeff.	*SE*	$P > \|z\|$	Coeff.	*SE*	$P > \|z\|$	Coeff.	*SE*	$P > \|z\|$
Gender	–0.180	0.023	.000	–0.154	0.044	.000	–0.293	0.055	.000
Child's Age (12–23)									
24–35 months	0.009	0.031	.780	0.009	0.031	.777	–0.014	0.033	.663
36–47 months	–0.169	0.031	.000	–0.169	0.031	.000	–0.189	0.033	.000
48–60 months	–0.213	0.036	.000	–0.213	0.036	.000	–0.284	0.040	.000
Birth Order	–0.172	0.010	.000	–0.172	0.010	.000	–0.073	0.012	.000
Surviving Older Siblings (none)									
1 sister only	0.066	0.041	.106	0.131	0.055	.018	0.071	0.059	.232
2+ sisters	0.278	0.054	.000	0.353	0.070	.000	0.214	0.074	.004
1 brother only	0.088	0.040	.028	0.087	0.055	.116	0.034	0.059	.560
2+ brothers	–0.117	0.057	.040	–0.236	0.075	.002	–0.260	0.080	.001
Mixed	–0.046	0.048	.344	–0.026	0.056	.644	–0.095	0.060	.112
Gender × 1 sister				–0.136	0.078	.083	–0.158	0.084	.059
Gender × 2+ sisters				–0.163	0.094	.082	–0.249	0.100	.012
Gender × 1 brother				0.001	0.077	.985	–0.059	0.082	.476
Gender × 2+ brothers				0.247	0.101	.014	0.296	0.108	.006
Gender × mixed				–0.041	0.060	.497	–0.054	0.064	.405
Mother's Education (none)									
Primary							0.399	0.039	.000
More than primary							0.689	0.041	.000
Lives in Extended Family							0.123	0.031	.000
Has Access to Media							0.340	0.030	.000
Mother Earns Cash							0.020	0.034	.563

(continued)

higher odds of less immunization than boys (e^{β} = exp(–0.180) = 0.83, $p < .001$ in Model 4). Children with two or more surviving older sisters are less likely to be severely stunted than are children with no surviving older siblings. On the whole, however, sex composition does not appear to be a strong determinant of severe stunting for all children combined. For immunization, having two or more older sisters or one brother increases the odds of a better immunization status relative to a child with no surviving older siblings.

The addition of an interaction between gender and sex composition in Models 2 and 5 shows the significance of the difference in the impact of sex composition on health

(Table 4, continued)

	Model 4			Model 5			Model 6		
Variable	Coeff.	*SE*	*P* > |*z*|	Coeff.	*SE*	*P* > |*z*|	Coeff.	*SE*	*P* > |*z*|
Mother's Age							0.012	0.003	.000
Household Size							–0.026	0.004	.000
Low Caste/Tribe							–0.184	0.030	.000
Religion (Hindu)									
Muslim							–0.486	0.042	.000
Other							–0.117	0.050	.018
Owns Goods (none)									
Owns nonluxury goods							0.245	0.031	.000
Owns luxury goods							0.740	0.050	.000
Roads in Village							0.095	0.028	.001
Health Services (none)									
PHC/subcenter only							–0.087	0.041	.032
Hospital/clinic as well							0.134	0.029	.000
Electricity in Village							0.433	0.035	.000
Female Literacy							2.042	0.080	.000
Female Labor-Force Participation							0.682	0.074	.000
Territorial Endogamy							–0.029	0.082	.724
Kinship Endogamy							1.072	0.103	.000
Region (North)									
Central							–0.035	0.045	.436
South							–0.079	0.062	.203
East							–1.202	0.069	.000
Gender × Central							0.162	0.060	.007
Gender × South							0.229	0.071	.001
Gender × East							0.245	0.082	.003
Cut-point 1	–1.199	0.033		–1.187	0.038		0.870	0.093	
Cut-point 2	0.116	0.033		0.129	0.037		2.471	0.094	
N		25,549			25,549			25,163	
Log-Likelihood Statistic		–27,375.3			–27,367.7			–23,940.3	

status for girls and boys; this is true of both outcomes presented here and remains significant after other related factors are controlled in Models 3 and 6.

The main effect for gender in these models is difficult to interpret because of multiple interactions. The coefficient on gender in the full model for severe stunting (Model 3) suggests that all else being equal, girls and boys no longer have significantly different odds of severe stunting. However, with all other variables at their reference value, there is still a significant female disadvantage in immunization status.

The interactions between gender and sex composition provide the test for the key hypothesis of this article. The interactions of gender with sex composition are jointly

Table 5. Odds Ratios for Selected Interactions Between Gender and Sex Composition in Childhood Immunization Status, Rural India: 1992–1993

	1 Surviving Older Sister			2 or More Surviving Older Sisters			2 or More Surviving Older Brothers		
Category	Odds Ratio	Lower CI[a]	Upper CI[a]	Odds Ratio	Lower CI[a]	Upper CI[a]	Odds Ratio	Lower CI[a]	Upper CI[a]
Boys (ref. = no siblings)	1.07	0.96	1.20	1.24	1.07	1.43	0.77	0.66	0.90
Girls (ref. = no siblings)	0.92	0.81	1.03	0.97	0.83	1.13	1.04	0.88	1.22
Girls (ref. = boys)	0.64	0.55	0.74	0.58	0.49	0.70	1.00	0.82	1.22

[a]CI = Confidence interval.

significant (not shown), suggesting that older-sibling composition in general plays a key role in determining gender-specific risks of severe stunting and immunization. In addition, certain sex compositions are individually significant when interacted with gender. Coefficients for key statistically significant interactions between gender and sex composition for immunization (Model 6) and severe stunting (Model 3) are presented in Tables 5 and 6, respectively, as odds ratios, with 95% confidence intervals. Tables 5 and 6 support the hypothesis that sex 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.

Despite the preference for sons, 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 does the first son (OR = 0.77); a boy with at least two older brothers is also 22% more likely to be severely stunted than is the first son, although this result is marginally nonsignificant (see Tables 5 and 6, respectively).

On the other hand, boys who are born into households that have only girls have significantly better health outcomes than does the reference category. In particular, boys who are 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 = 1.24) and among the lowest odds of severe stunting (OR = 0.81, marginally nonsignificant) of children with any combination of surviving older siblings (Tables 5 and 6, respectively).

Table 6. Odds Ratios for Selected Interactions Between Gender and Sex Composition in Severe Stunting Among Children, Rural India: 1992–1993

	2 or More Surviving Older Sisters			2 or More Surviving Older Brothers			Older Surviving Sisters and Brothers		
Category	Odds Ratio	Lower CI[a]	Upper CI[a]	Odds Ratio	Lower CI[a]	Upper CI[a]	Odds Ratio	Lower CI[a]	Upper CI[a]
Boys (ref. = no siblings)	0.81	0.64	1.02	1.22	0.97	1.55	1.01	0.84	1.22
Girls (ref. = no siblings)	1.30	1.03	1.63	0.91	0.73	1.14	1.20	1.00	1.43
Girls (ref. = boys)	1.61	1.29	2.01	0.74	0.60	0.92	1.13	0.98	1.31

[a]CI = Confidence interval.

Similarly, all girls do not face equal discrimination. Girls with multiple surviving sisters face more neglect than do other children, a result consistent with Das Gupta's (1987) findings about mortality differences in the Punjab in India. Compared with girls with no surviving older siblings, a third or higher-birth-order daughter is 30% more likely to be severely stunted (see Table 6); the odds of immunization are not significantly different, however, for first and subsequent daughters (see Table 5). Moreover, girls who have a combination of older brothers and sisters are also more likely to be severely stunted than are girls with no older siblings; this result is marginally statistically significant for severe stunting (see Table 6) and not significant for immunization. The gender difference, however, is statistically significant for both outcomes. A girl who is born into a household with two or more surviving girls is 61% more likely to be severely stunted than is a boy who is born after two or more surviving girls (see Table 6). For immunization (see Table 5), having any older sisters at all is likely to be deleterious for girls compared with boys, and girls with one or with two or more surviving older sisters have significantly lower odds of a better immunization status compared with boys with similar numbers of older sisters (OR = 0.64 for one older sister and OR = 0.58 for two or more older sisters).

Parents' desire for a balanced sex composition does not appear to provide the same degree of protection for girls with opposite-sex siblings as it does for boys with opposite-sex siblings, reflecting the strength of the preference for sons in this population. A girl who is born into a household that has only boys (at least two) does not have significantly different odds for immunization and severe stunting than does a girl with no surviving older siblings (see Tables 5 and 6, respectively). The gender difference does, however, present mild evidence that a girl who is born into a household with two or more surviving boys and no other girls may be preferred to a boy who is born into such a household. Table 6 shows that girls with two or more surviving brothers are about 25% less likely to be severely stunted than are boys with two or more surviving brothers (OR = 0.74).

The interactions of gender and region show that 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(–0.079 + 0.23) = 1.36) than girls in the north. This result is consistent with the literature, which showed a higher preference for sons and greater discrimination against girls in northern than in southern India (Arnold et al. 1998; Dyson and Moore 1983; Govindaswamy and Ramesh 1996).

Other results for all the children in both Models 3 and 6 are, by and large, consistent with the literature. 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. The odds of immunization decrease with age, perhaps reflecting that if a child does not get immunized close to the recommended age, he or she may miss a window of opportunity. Vaccination campaigns that occurred in India in this period may, for instance, have specifically targeted younger children. However, because age can be a proxy for many unobserved factors, this result is difficult to interpret.

It is interesting that children in extended families are less likely to be severely stunted and have higher odds of better immunization than do children in nuclear families; the results remained significant for immunization after household size, wealth, and other factors were controlled. This finding could reflect a beneficial impact for any child of the presence of multiple caretakers in an extended family. In the case of immunization, compared with no health services, the presence of only a governmental primary health center (PHC) or subcenter appears to lower the odds of better immunization for children in that village, whereas the additional presence of a public or private hospital or clinic is associated with higher odds of better immunization. These results may reflect that parents prefer to get their children immunized in clinics or hospitals—whether public or private—rather than at the more local levels of government health facilities, such as PHCs or subcenters. However, they could also reflect unobserved heterogeneity in the communities, not adequately controlled for by the village development variables, and associated with both higher immunization and the presence of a hospital or clinic in the village. Women's employment, at the individual and village levels, increases the odds of severe stunting. This finding is consistent with the findings of other studies that suggested that compared with nonworking mothers' children, working mothers' children may face poor health and higher mortality risks, regardless of gender, because working mothers may have less time for child care (Basu and Basu 1991; Desai and Jain 1994; Leslie 1989).

DISCUSSION

Despite the long-standing persistence of gender differences in mortality on the Indian subcontinent, patterns of gender discrimination in the proximate determinants of child mortality—in other words, in the health status of surviving children—have not received adequate attention in the literature. This article focused on this gap by examining gender differences in severe stunting and immunization, particularly the effects of sex composition on gender discrimination for surviving girls and boys for these two outcomes. The overall results show that severe stunting and the 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 differences in health status among surviving children.

It also appears that, as hypothesized, 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 differences 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 do 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. The strength of the preference for sons and the low value of girls are evident in that 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 it is for boys.

The strength of these results for immunization has particularly sobering policy implications for programs such as the Expanded Program on Immunization, 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 seen here is somewhat surprising. These results indicate 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, when immunization is seen as a life-saving or positive health practice, mothers do use opportunities for immunization to improve the odds of good health, primarily for preferred children.

That parents may discriminate selectively on the basis of gender *and* on where a particular son or daughter "fits" into the overall sex composition of the family strongly suggests that parental and societal norms about the values of girls relative to boys and about a desirable family sex composition are key to explaining why certain children fare worse than their siblings and why girls with older sisters fare particularly badly. While such an analysis is beyond the scope of this article, understanding better the process of how social norms around gender are formed, how they change, and who is more likely or less likely to adapt to these norms in terms of gender and sex composition differences in child health are crucial next steps for appropriate policy intervention in this domain.

APPENDIX

I created weighted continuous variables to measure the proportion of village-level female literacy, female employment, exogamy, and endogamy to adjust for the fact that in some villages, the village sample was too small to allow a statistically meaningful analysis. In such cases, I included district-level aggregation, using weights to reflect the size of the village-level sample relative to the district-level sample for each unit of analysis (the village). The weights were created such that as the village-level sample increased, the village-level value was assigned a higher weight relative to the district-level value for that village. This weight allowed me to create village-level variables even for small villages (where the district-level value from the larger district-level sample was weighted higher), yet at the same time to give more weight to the village-level when the sample size was appropriately large, thus keeping to the spirit of the conceptual framework that emphasizes the immediate—or village—environment.

I used a basic exponential weight chosen to bound the proportion between 0 and 1. The weight is $w_v = 1 - \exp(-\lambda n_v)$, where w_v is the weight for village v, n_v is the number of sample women in village v, and λ is a scale parameter chosen on the basis of the number of women at which λ approaches 1. Then, for example, the variable for village kinship endogamy is defined as $K = w_v \times K_v + (1 - w_v) \times K_d$, where K is the weighted village-level kinship endogamy variable, K_v is the proportion of endogamy in the village, and K_d is the proportion of endogamy in the district. The larger the number of sample women in the village, the greater the weight accorded to the village-level endogamy variable, and vice versa. I set $\lambda = .05$, based on the fact that the weight with this value reaches 1 as the number of sample women in the village approaches 100, a sufficient size to allow the variable to be based on the village (rather than district) level of this measure. In constructing these variables, I excluded the index woman.

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