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INTRAHOUSEHOLD ALLOCATION OF NUTRIENTS IN RURAL INDIA: ARE BOYS FAVORED? DO PARENTS EXHIBIT INEQUALITY AVERSION?

By JERE R. BEHRMAN*

INTRAHOUSEHOLD allocations account for an important portion of total goods and services and can serve to modify substantially policy interventions and market outcomes. In almost every society, most of the nutrients for most children are obtained through intrahousehold food allocations. For South Asia there are a number of indices of health outcomes (e.g., mortality rates, weight for age) that suggest that the intrafamilial allocation of nutrients favors boys over girls, particularly in northern areas (Banerjee; Bardhan; Boserup; Cassen; El-Badry; Hammond; Harriss; Natarajan; Pakrasi and Halder; Preston and Weed; Ram Gupta; Rosenzweig and Schultz; Sen; Sen and Sengupta; Visaria; and Waldron).[1] A number of commentators on this phenomenon suggest that the pattern in these health indices reflects parental preference for sons. In contrast, Rosenzweig and Schultz argue that it is consistent with a parental maximizing response to differential labor-market returns to investments in boys and girls and some others—such as Sen and Sengupta—recognize that differential labor-market returns by sex may play an important role in nutrient allocations.

However existing studies do not examine the nature of parental preferences underlying the intrahousehold distribution of nutrients between sons and daughters.[2] This paper helps to fill this void in the literature by

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[1] There also is some suggestion of age and/or birth order discrimination, which I have studied in Behrman (1988a). Conceptually and empirically (if age and sex are not correlated) gender and age discrimination can be examined separately. I consider only gender in this paper.

[2] Behrman, Pollak and Taubman (1986) explore the nature of gender preferences in the allocation of schooling among siblings in the United States, including a discussion of the impact of endowment differentials versus unequal concern for schooling investment and expected adult earnings and marriage market outcomes in the single-input, single-output case.

using rich new Indian data to estimate an extension of the BPT (Behrman, Pollak and Taubman, 1982) intrahousehold allocation model, adapted to consider multiple health-related outcomes and multiple nutrient inputs.[3] These estimates permit answers to a number of important questions: Is there evidence of promale bias in the sense that parents weigh equal health-related outcomes more heavily for boys than for girls? Do parents allocate nutrients in a pure investment sense (i.e., maximizing total returns, independent of distribution) as Rosenzweig and Schultz seem to suggest, or is there a tradeoff between the total returns of such investments and their distribution among children? If there is such a tradeoff, does it imply compensation or reinforcement of differentials in genetic-related endowments (including gender)? Does the extent of promale bias depend on whether households hold land, on their caste, or on the education of the household head? Is seasonality critical, as authors in Chambers, Longhurst and Pacey, authors in Sahn and others have stressed, with different responses during the "lean" than during the (relatively) "surplus" season?

1. Multi-input, multi-output intrahousehold allocation model with parental preferences by gender of children

Within an optimizing model of parental allocation, systematic gender differences in human capital investments may originate in at least three conceptually different ways. First, parents may have no gender preferences, but just may respond to expected gender differentials in expected outcomes as is suggested above. Second, parents may have preferences that favor children of one gender in the sense that they value equal expected outcomes at equal cost more highly for that gender than for the other. Third, parents may have no gender preferences, but just may respond to systematic gender differences in prices of human-capital investments.[4] These distinctions are important. For example, how parents change their nutrient investments in children in response to changes in the expected health returns or expected wage differentials by gender may depend critically on the relative importance of the first versus the second of these possibilities.

General multi-input, multi-output intrahousehold allocation model

The BPT preference model of intrafamilial allocation provides a framework within which these distinctions can be made explicit. However this model is for the single-observed-input (i.e., schooling), single-observed-output (i.e., earnings) case. In the present context there are several

[3] BPT consider one output (the weighted sum of own earnings and expected spouse's earnings) and one input (schooling) for U.S. adult sibling samples.
[4] The third possibility is included for completeness. In the empirical case considered in this paper there is no reason to expect such differentials for the price of nutrients so, for simplicity, I abstract from such a possibility. If such differentials were thought to be important, however, it would be easy conceptually to introduce them.

observed nutrient inputs and several observed health-related outputs. Therefore this model is adapted and generalized here to the multi-nutrient input, multi-health indicator output case.

Parents are assumed to have a utility function (U^*) defined over a separable subutility function (U) and other outcomes of interest (Z^*). The separable subutility function is defined over I expected health-related outcomes for each of the J children (H_{ij}; $i = 1,\ldots, I$; $j = 1,\ldots, J$):[5,6]

$$U = U(H_{11},\ldots, H_{IJ}) \tag{1}$$

These preferences are maximized subject to two sets of constraints.

The first constraint is the partial nominal budget constraint that applies to household resources devoted to investment in the children (R). In anticipation of the empirical application in Section 2, I distinguish between investments in K observed nutrients (N_{kj}, $k = 1,\ldots, K$; $j = 1,\ldots, J$) and other health-related investments (X_j, $j = 1,\ldots, J$). I also assume that parents face fixed prices (P_{N_k}, P_X) for both observed nutrients and other health-related investments and that these are identical for all children:[7]

$$\sum_{j=1}^{J}\left(\sum_{k=1}^{K} P_{N_k}N_{kj} + P_X X_j\right) \leq R \tag{2}$$

The second set of constraints are the I expected health-related-outcome generating relations for each of the J children. For the jth child each of these I relations depends on the observed nutrients (N_{kj}), other purchased (X_j) inputs and endowments (E_j):[8]

$$H_{ij} = H_i(N_{1j},\ldots, N_{Kj}, X_j, E_j); \qquad i = 1,\ldots, I; \qquad j = 1,\ldots, J. \tag{3}$$

The endowments are Beckerian in the sense that they include both genetic and environmental factors that affect the health-related outcomes, but they are not purchased particularly for that reason (and perhaps not purchased at all). Critical for the present purpose are any differences in endowments that affect the level of the health-related outcomes systematically by gender, which could range from genetic gender differences in disease susceptibility

[5] For the short-run nutrient allocation problem under investigation, the number of children is taken as predetermined. Separability is required to explore the nature of preferences related to investments in children's health in the absence of data on other intergenerational transfers (e.g., dowries, land bequests).

[6] A number of observers have argued in favor of a bargaining approach rather than unified preferences for the household. While such arguments seem a priori plausible, existing studies do not persuasively distinguish between the two alternatives since they do not have data that plausibly reflects only the bargaining strengths of household members and not the opportunity cost of their time (i.e., wages, schooling, and earnings will not do for this purpose). See Behrman (1988b) for a review and more extensive discussion of these issues.

[7] For the empirical application below it seems plausible that parents face the same nutrient prices for different children. The prices of the other inputs may vary for different children, but these prices drop out of the analysis below so there is no loss of generality in assuming them to be the same across children in (2).

[8] At any point of time the totality of past inputs and endowments may affect the efficiency of the translation of nutrient intakes into health through, for example, infectious diseases.

to gender differences regarding the valuation of health-related outcomes (e.g., different wages by gender for health-related outcomes related to labor-market success).

Under the assumption that these functions have the desirable properties for an interior maximum, the K times J first-order conditions for parental nutrient investment of type k in the jth child can be derived. The K ratios of such expressions for the jth to the qth child are:

$$\left(\sum_{i=1}^{I} \frac{\partial U}{\partial H_{ij}} \frac{\partial H_{ij}}{\partial N_{kj}}\right) \Big/ \left(\sum_{i=1}^{I} \frac{\partial U}{\partial H_{iq}} \frac{\partial H_{iq}}{\partial N_{kq}}\right) = 1; \qquad k = 1, \ldots, K. \tag{4}$$

The ratio forms have the advantage of not including the unobservable family-specific Lagrangian multiplier, nor the often unobserved price of the kth nutrient. In the case of just one health-related outcome of interest (i.e., $I = 1$), these ratios can be rewritten as:

$$\frac{\partial U / \partial H_{ij}}{\partial U / \partial H_{iq}} = \frac{\partial H_{iq} / \partial N_{kq}}{\partial H_{ij} / \partial N_{kj}}; \qquad k = 1, \ldots, K. \tag{4A}$$

This special case makes clear the usual tangency condition for the maximum, e.g., the slope of the parental welfare function equals the slope of the health-related-outcome possibility frontier for the jth versus the qth child (point A in Fig. 1). Since the right side of this equation depends on the endowments of the jth versus the qth child, the health-related-outcome possibility frontier varies across different pairs of children in the same or in different households even if the partial nominal budget constraint in equation (2) and the functional forms for equations (1) and (3) are the

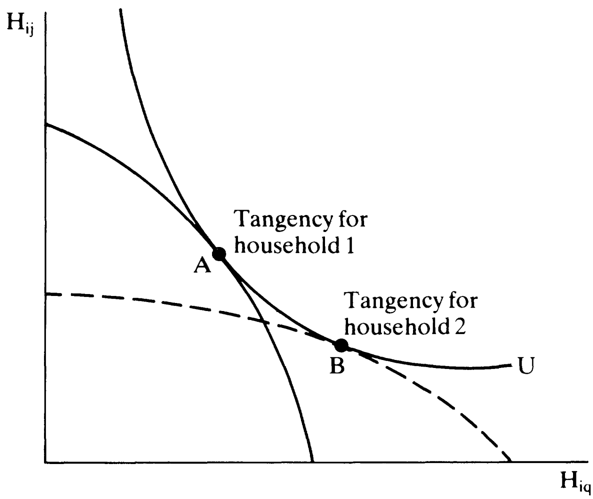

FIG. 1. Optimum intrahousehold allocation of human capital investments at point A for household 1 (solid lines). Relative endowments for child j lower in household 2 (dashed production possibility frontier) than in household 1.

same. This variation in the health-related-outcome possibility frontier across pairs of children permits identification of the slope of the welfare function (see BPT). Figure 1 also provides an illustration in which the endowments for child j versus q are lower in household 2 (dashed production possibility frontier) than in household 1. The slope of the welfare function is of great interest because it characterizes the productivity-equity tradeoff in parental preferences regarding the distribution of health-related outcomes among their children. Likewise any asymmetry of the curve around the 45° ray from the origin is of great interest because it reflects the extent of parental preference for children for a particular type, for example, identified by gender (see below).

Specific functional forms for the multi-input, multi-output BPT separable model

In order to estimate identifiable parameters in the parental subutility function, specific assumptions must be made about the functional forms in relations (1) and (3).

For the subutility function in relation (1) a CES form is used:

$$U = \left(\sum_{j=1}^{J} \sum_{i=1}^{I} a_{ij} H_{ij}^{c} \right)^{1/c} \tag{5}$$

This is a common functional form which allows the full range of parental inequality aversion with an one-parameter representation of that inequality aversion.[9] Parental inequality aversion is inversely associated with the parameter c. At one extreme, $c \rightarrow -\infty$ is the infinite inequality aversion case in which parents only value improved health for the worst-off child. At the other extreme $c \rightarrow 1$ is the zero inequality aversion case in which parents value all additional child health improvements equally, independent of the distribution of health-related outcomes among their J children. This extreme is equivalent to the widely-used standard pure investment model in which human-capital investments in children are made without reference to distribution among siblings (e.g., Becker and Tomes).

The parameters $(a_{11}, \ldots, a_{ij})$ are the weights that the parents place on each child's health-related outcomes. Equal concern about children is defined to be the case in which all a_{ij} are equal for a given outcome (say to a_i). If there is parental preference for children of one gender, there is unequal concern and a_{ij} is dependent on the gender of the jth child. If, say, a_{ij} is higher for boys than for girls, parents tradeoff positive health-related outcomes among their children, but they value equal health-related outcomes more for boys

[9] The generalized CES with displacement from the origin also was explored in this study. The estimates of the displacement parameter for this sample generally are not significantly nonzero and are not very robust to changes in specification. The estimates of the parameters of primary interest—inequality aversion and gender preference—however, are robust with the displacement parameters included or constrained to zero. Therefore in what follows focus is on the usual CES case with no displacement parameters.

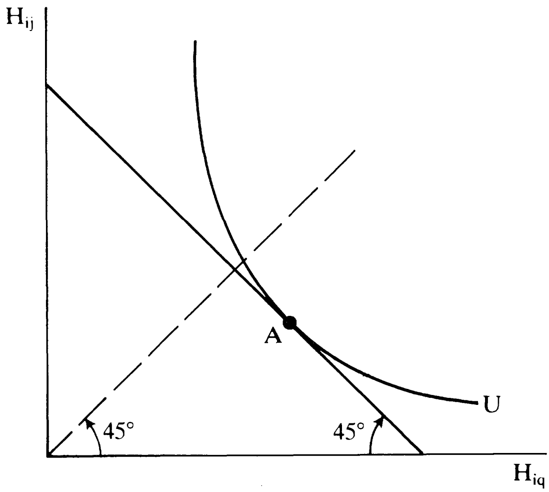

FIG. 2. Unequal concern favoring child q.

than for girls. Below I discuss more explicitly the impact of unequal concern on parental allocations of human capital and on relative health-related outcomes. Figure 1 assumes equal concern between the jth and qth children so the isoutility curves are symmetrical around a 45° ray from the origin; Fig. 2 in contrast exhibits unequal concern with parental preferences favoring child q.

The health-related-outcome production functions are assumed to be quasi-Cobb–Douglas:

$$H_{ij} = f_i(E_j) X_j^{\beta_i} \prod_{k=1}^{K} N_{kj}^{\alpha_{ki}}; \qquad i = 1, \ldots, I; \qquad j = 1, \ldots, J. \tag{6}$$

This is a widely-used and tractable form with diminishing marginal returns to nutrients (assuming $\alpha_{ki} < 1$ for $i = 1, \ldots, I$ and $k = 1, \ldots, K$).

Substitution of the first derivatives of (5) and (6) into (4) and additional definitions for the η's gives a set of K relations:

$$\eta_0 \frac{\sum_{i=1}^{I} \eta_{ijk} H_{ij}^{c}}{\sum_{i=1}^{I} \eta_{iqk} H_{iq}^{c}} = \frac{N_{kq}}{N_{kj}}; \qquad k = 1, \ldots, K, \tag{7}$$

where η_0 is the common gender-dependent portion of the ratio of a_{ij}/a_{iq} across I health-related outcomes, and η_{ijk} (and η_{iqk}) are the products of the rest of a_{ij} (a_{iq}) and α_{ki}. Under the assumption that any parental preferences by gender of a child shifts all of the preference weights for that child by the same proportion,[10] $\eta_0 = (1 + \varepsilon D_j)/(1 + \varepsilon D_q)$ reflects any parental gender

[10] That is, $a_{ij} = a'_{ij}(1 + \varepsilon D_j)$, $a_{iq} = a'_{iq}(1 + \varepsilon D_q)$.

preference regarding the children with D_j (D_q) one if the jth (qth) child is male and zero if that child is female. Then $\eta_{ijk} = a'_{ij}\alpha_{ki}$ and $\eta_{iqk} = a'_{iq}\alpha_{ki}$.[11]

With a stochastic specification and data on nutrient inputs, health-related outputs and gender for siblings, estimates can be obtained of c, $\eta_0 = a_{ij}/a_{iq}$, and η_{ik}, $i = 1,\dots, I$. Though endowments other than gender play an important role in identifying the curvature of the parental subutility function, one striking feature of relation (7) is that such endowments do not appear explicitly—which is convenient since most such endowments generally are not observed in social science data sets.

Although the parental preference weights (a_{ij}, a_{iq}) cannot be identified separately, the dependence of their ratio on the genders of the jth and qth children can be explored under the above assumptions. Thus whether parental preferences exhibit this kind of gender favoritism can be identified from estimation of relation (7)—conditional on the specific assumptions about functional forms.

Implications of gender differentials in endowments on intrahousehold nutrient and health outcome allocations

One perhaps surprising feature of relation (7) is that the gender endowment differentials that affect health-related outcomes do not appear explicitly, though they are in the background since they work through the H_{ij}'s. This does *not* mean that gender endowment differentials leave intrafamilial allocations unaffected. To the contrary, they shift the health-related-outcome possibility frontier around just like other endowments.

Figure 3 provides an illustration of this effect in the simple case in which there is only one health-related outcome and in which parents have equal concern independent of the gender of their children. The solid health-related-outcome possibility frontier is one for which both children are males; it is tangent to the solid isoutility curve at point A. Consider what happens if child j is replaced by an otherwise identical female child, there is a gender differential in returns to nutrient investments that favors males, and total family resources allocated to health-related investment in the children (R) are constant. The result is a health-related-outcome possibility frontier everywhere interior to the previous one except on the H_{iq} axis, as is indicated by the dashed line. The boy–girl health-related outcome possibility frontier is tangent to the dashed parental isoutility curve at B. At point B the expected health-related outcome is less for both the jth and qth children due to the change in gender of the jth child from male to female,

[11] Only these products can be identified, not the non-gender-dependent component of the preference weights (a'_{ij}, a'_{iq}), nor the elasticities of the ith health-related output with respect to the k nutrient input (α_{ki}). Moreover, the η_{ijk}'s are identified only in relative terms since each could be divided by the same arbitrary nonzero factor without affecting (7) because it would cancel out between the numerator and the denominator. Therefore in the estimates below the normalization that $\sum_{i=1}^{I} \eta_{ijk} = \sum_{i=1}^{I} \eta_{ik}$ (since the person index can be supressed) = one is utilized.

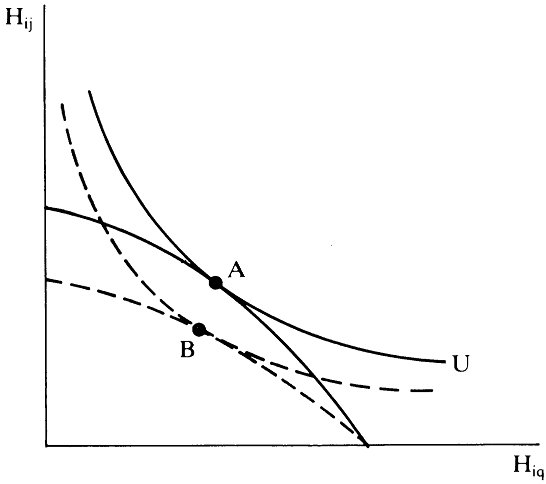

FIG. 3. Impact of comparative static change in gender of child j from male to female given endowment differentials favoring males.

given gender differentials in returns to nutrient investments. But, except for the extreme inequality aversion case in which the parental isoutility line is L-shaped along the 45° ray, the decline is greater for the jth child because parents respond to the less favorable returns to this child if it is a girl as compared to if it were a boy.

These notions can be illustrated algebraically for the multi-input, one-outcome case. In this case, the first-order conditions can be solved to obtain reduced-form expressions for the nutrition investments and the health-related outcomes for the jth child relative to the qth child:

$$\frac{N_{kj}}{N_{kq}} = \left(\frac{a_j}{a_q}\right)^{1/(1-\delta c)} W^{c/(1-\delta c)} \tag{8}$$

$$\frac{H_j}{H_q} = \left(\frac{a_j}{a_q}\right)^{\delta/(1-\delta c)} W^{1/(1-\delta c)} \tag{9}$$

where

$$\delta = \sum_{k=1}^{K} \alpha_k + \beta \quad \text{and} \quad W = f(E_j)/f(E_q)$$

and the i subscripts are supressed since there is only one health-related output. These expressions indicate directly how the relative endowments (W) affect intrahoushold nutrient investments and intrahousehold health-related outcomes. The comparative static exercise of changing the jth child's gender from male to female lowers W, ceteris paribus, if investment returns are greater for males.

The sign and magnitude of the exponents of W are critical. Under the assumptions that endowments have an impact on the health-related outcome and that the health-related-outcome production function has nonincreasing returns to scale, the sum of the health-related outcome elasticities with respect to the nutrient and other health-related investments is less than one ($\delta < 1$). This limit on δ together with the limit on the range of c to assure a maximum (i.e., $c \leqslant 1$) assures that the demoninators in the exponents for both expressions are strictly positive. Other properties of these exponents depend critically on the extent of parental inequality aversion (c).

As $c \rightarrow -\infty$ parents are concerned more and more single-mindedly with equality. At the limit they focus all of their nutrient and other health-related investments in the child for whom the endowments, including those related to gender, otherwise would result in the lower health-related outcome until the point that health-related outcome equality is achieved. Changing the gender of the jth child from male to female in the infinite inequality aversion case has no effect on the distribution of parental resources between the jth and qth child if the sign of $(W-1)$ is not changed. However if W changes from above one to below one because of sufficiently strong endowment differentials favoring males, there is an accompanying radical shift from investing only in the qth child (son) to investing only in the jth child (daughter).

Between the infinite inequality aversion extreme and the Cobb–Douglas case ($-\infty < c < 0$), parental inequality concern is strong enough so that the child less favored by endowments receives greater nutrient investments than does the other child. This partially compensates for the relative health-related-outcome impact of the factors in W, but does not completely offset it. The change in the gender of the jth child from male to female causes a relative shift in nutrients from the qth child, but there is a switch from the qth child receiving more nutrients to the jth child receiving more nutrients only if W changes from below to above one.

In the Cobb–Douglas case ($c = 0$), parents just balance off equity versus productivity concerns in the sense that nutrients are distributed independently of the value of W (e.g., the exponent in relation (8) is zero), and relative health-related outcomes are proportional to W((e.g., the exponent in relation (9) is one). In this case a change in the gender of the jth child from male to female leaves intrafamilial nutrient allocation unaffected, but alters the relative health-related outcome unfavorably for the jth child in exact proportion to W.

If parental concern about productivity outweighs their concern about equity on the investment side of the Cobb–Douglas case (including the pure-investment extreme at which there is no inequality aversion), nutrient investments result in greater relative inequality in health-related outcomes than would result alone from the gender endowment differentials that are incorporated into W. That is, the parents are interested sufficiently strongly in productivity that they respond to gender endowment differentials by investing so as to reinforce the inequality of health-related outcomes. If the gender of child j is changed from male to female so W falls, the health-related outcome of the W qth to the jth child increases. Whether that results in more (less) health-related outcome inequality than in the case in which the jth child is a son depends on whether W moves closer to or further from one as a result.

Because on the pure-investment side of the Cobb–Douglas case parents act so that health-related outcomes are distributed more unequally than they would be due to W alone, however, does *not* necessarily mean that they distribute nutrients and other health-related investments in greater quantities to the child favored by gender endowments differentials. Between the Cobb–Douglas and pure-investment cases there is a value of $c = \bar{c}$, at which level these human-capital investments are distributed exactly proportional to W:

$$\bar{c} = 1/(1 + \delta) \tag{10}$$

Given that $0 < \delta < 1$, $\bar{c}$ can be bounded more narrowly to be between 0.5 and 1, with a value closer to one the smaller is δ. If $0 < c < \bar{c}$, nutrients are distributed more equally than the impact of the gender endowment differentials in W (e.g., the exponent in relation (10) is below one), even though health-related outcomes are distributed more unequally. If $c > \bar{c}$, both nutrients and health-related outcomes are distributed more unequally than is the impact of W. The impact on the distribution of nutrients of a change in the jth child's gender from male to female obviously depends on whether c is above, below, or just at $\bar{c}$. If c is below $\bar{c}$, there is a relative shift in resources towards the jth child (and vice versa if $c > \bar{c}$, with no impact at $c = \bar{c}$).

This discussion has focused on the multi-input but only one-outcome case because the relations are clearer in this case than if there are the sums over outputs in the numerators and demonimators in the left side of (7) as in the multi-output case. But the basic thrust of these comments, though not the exact details, carries over to the multi-output case.

Implications of parental gender preferences

Figures 1 and 2 suggest that the effect of parental unequal concern favoring one child is to shift nutrient and other health-related investments from the other child to that one (except in the case in which parental preferences are characterized by infinite inequality aversion). The shift in nutrient and other health-related investments, of course, results in a shift in the relative health-related outcome towards that child, which is larger the more productive are nutrients and other health-related investments (i.e., the larger is δ).

The dependence of parental allocation of nutrients on whether the parental preference weights (a's) on health-related outcomes favor one gender is apparent from relation (7). The allocation of nutrients is directly related to the ratio of the gender-dependent component of these weights (η_0) in the multi-output model, though whether this relation is less than, exactly, or more than proportional is not obvious because of the induced changes in H_{ij} and H_{iq}.

The increased resources devoted to a favored child due to parental gender preference result in increased relative health-related outcomes for that child, ceteris paribus, in the multi-output model, though the general result cannot be expressed in a very illuminating simple algebraic manner. For the special case of the multi-input, single outcome model in relations (8) and (9), however, the exact impact of gender preferences in the a's is immediately clear. Relative nutrient and other health-related investments are changed directly, but less than proportionately to the ratio of these weights (i.e., $1/(1-\delta c) < 1$). Relative health-related outcomes change less than proportionately if $\delta - 1 < c$ (so $\delta < (1-\delta c) < 1$) and vice versa.

Taste heterogeneity, health-related outcome possibility frontier heterogeneity and simultaneity biases

When the parents choose an allocation of nutrients between two children, they elect both the ratio of nutrients and the ratio of expected health-related outcomes. This joint decision causes simultaneity bias if ordinary-least-squares estimates are made of (7) with an additive stochastic term on the left side to reflect random measurement error and differences between expected and realized health outcomes. Unfortunately instruments are not available in the data set used for this study (nor likely to be available in most other data sets) to permit use of some simultaneous estimator since useful instruments must differ among the children in each household. However some insight is possible about the nature of the bias. In the one-outcome case this simultaneity results in an upward bias in the estimate of c (i.e., a bias towards the pure investment case), as is demonstrated in Behrman and Taubman.[12] While an extension to the multioutput case is messy, it appears that the bias also is upward in this case.

[12] Suppose there is heterogeneity in tastes, with the average curve yielding A as a tangency in Fig. 1, but with a tangency for a particular household at A', directly below A, at which point the jth sib in that household has less nutrition and health than has the jth sib in an average family. Now consider the bias obtained in an OLS nonlinear regression with the nutrition ratio as the dependent variable. Denote the natural logarithm of the ratio by small letters. Then the equation to be considered is $n_j = ch_j + w_j$, which is simply (7) rewritten for the single input, single outcome case. The expected value of the least square estimate is given by

$$\text{plim}\,(\hat{c}) = \frac{\text{plim}\sum n_j h_j}{\text{plim}\sum h^2}\frac{\text{plim}\sum (ch_j + w_j)h_j}{\text{plim}\sum h_j^2} = \text{plim}\,(c) + \frac{\text{plim}\sum w_j h_j}{\text{plim}\sum h_j^2}$$

Now w_j can be written as $u_j + v_j$, where v_j consists of measurement error arising, for example, from differences between expected health and the health realization and where u_j represents the error arising from taste heterogeneity. The standard assumption is that v_j and h_j are uncorrelated. From (6), each unit decrease in n_j caused by heterogeneity results in an α reduction in h. Thus $\text{plim}\sum w_j h_j = \text{plim}\sum \alpha h_j^2$. Therefore, $\text{plim}\,(\hat{c})$ equals $c + \alpha$ and is biased up since $0 \leq \alpha \leq 1$. If, alternatively, there is error in the health outcome possibility frontier, points such as A and A' are observed, and the same analysis goes through.

2. Data and estimates

Data

The households in the sample are the 240 households in the ICRISAT VLS (International Crops Research Institute for the Semi-Arid Tropics Village Level Studies) Panel Data Set. This panel was established in 1975 by selecting two typical villages with regard to crop patterns, land-use, irrigation, etc. in each of three agroclimatic regions of semi-arid-tropical (SAT) India (two villages in Andhra Pradesh and four in Maharashtra). Within each village a random sample of 40 households (stratified by primarily labor income and three sizes of cultivating groups) was selected for the panel.

Between September 1976 and January 1978, in collaboration with the National Institute of Nutrition (NIN) and the College of Home Science of Andhra Pradesh Agricultural University (APAU), home-science graduates in nutrition collected individual 24-hour recall dietary information four times at intervals of 3–4 months (but not on fast nor festival days). The intakes of 11 nutrients of these individual diets then were calculated on the bases of conversion factors in Gopalan *et al.* and subsequently converted to percentages of age-specific Indian Recommended Daily Allowances (RDA) for moderate activity levels from the same source. At the same time medical doctors collected anthropometric data on weight, arm circumference, triceps skinfold thickness and height, which subsequently were standardized for weight and height by age-specific standards for higher income groups in Hyderabad as determined by NIN (and which are at about the 50th percentile of the widely-used Harvard standards) and for weight-for-height, arm circumference and triceps skinfold thickness by international standards in Jelliffe. (For further details see Ryan, Bidinger, Rao and Pushpamma, hereafter RBRP.)

The standardized nutrient intake data and the standardized anthropometric data for children under 13 years of age from households in which more than one child resided are the critical data for this study. To explore whether there are seasonal differences, the RBRP distinction is maintained between the lean (or relatively tight or depressed with respect to food supplies) and the surplus (or relatively abundant) season. To lessen random measurement error, mean values for each individual across rounds in a given season are used. Because of the fairly short-run nature of the allocation process under investigation with the focus on seasonal differences, the relatively long-run health index of height is not used in the basic estimates. Therefore I begin below with four standardized anthropometric indices of health: weight-for-height, arm circumference, weight and triceps skinfold thickness. Because RBRP conclude that the critical nutrients for this population are five—calories (or energy), β-carotene, riboflavin, vitamin C (ascorbic acid), and calcium—these nutrients are considered in the present study.

The standardized anthropometric measures (Table I) indicate that most of

TABLE I

Summary Statistics for Distributions of Standardized Anthropometric Health Indicators and Standardized Nutrient Intakes for Lean and Surplus Seasons for Boys and Girls in South Indian Rural Households, 1976–8

	Boys			*Girls*		
Season and Variables	*Mean*	*Standard Deviation*	*Standard Error of Mean*	*Mean*	*Standard Deviation*	*Standard Error of Mean*
LEAN SEASON						
Anthropometric Measures						
Weight-for-Height	88	7.5	0.6	88	8.5	0.8
Arm Circumference	82	7.1	0.6	83	8.0	0.7
Weight	75	11	0.8	78	13	1.0
Triceps Skinfold Thickness	48	25	2.0	46	22	2.0
Nutrient Intakes						
Calories	89	27	2.2	90	32	2.9
β-Carotene	29	34	2.7	25	17	1.6
Riboflavin	54	34	2.8	55	31	2.8
Vitamin C	22	21	1.7	23	27	2.4
Calcium	48	29	2.4	50	34	3.0
SURPLUS SEASON						
Anthropometric Measures						
Weight-for-Height	89	7.5	0.6	88	8.5	0.8
Arm Circumference	82	7.6	0.6	82	8.4	0.8
Weight	77	11	0.9	78	13	1.0
Triceps Skinfold Thickness	54	22	1.9	50	23	2.0
Nutrient Intakes						
Calories	90	27	2.3	85	29	2.6
β-Carotene	23	17	1.4	22	18	1.6
Riboflavin	69	56	4.7	61	44	4.0
Vitamin C	21	20	1.7	25	27	2.5
Calcium	47	26	2.1	52	46	4.1

the children in the sample are below international standards for both seasons, particularly for triceps skinfold thickness. This also is the only health indicator for which the surplus-season mean exceeds significantly the lean-season one, both for boys and girls.[13] The mean for boys exceeds that for girls significantly only for triceps skinfold thickness in the surplus season, and the mean for girls exceeds that for boys significantly only for weight in the lean season. In terms of the means for anthropometric health indicators, thus, there is some limited evidence of better health-related outcomes for the surplus season than for the lean season, but not particularly for boys versus girls.

[13] I use "significant" to refer to the standard 5% level unless otherwise qualified.

The standardized nutrient input measures suggest diets significantly below the standards for both boys and girls in both seasons, particularly for β-carotene and vitamin C. The only means that differ significantly between the seasons are for riboflavin (higher in the surplus season) and β-carotene (higher in the lean season). None of the means for the standardized nutrient input measures differ significantly between boys and girls. On the basis of these means alone, thus, one might be tempted to conclude that there is no gender discrimination and very little difference between the seasons. However there still may be gender preferences that underlie these allocations, within the framework of the model of the previous section, if they are offset more or less by gender endowment differentials (see relations 8 and 9).

Basic multi-health indicatior, multi-nutrient input intrahousehold allocation estimates

For these estimates the health-dependent outcomes that are posited to enter into the parental subutility function are the four anthropometric measures. That is, the parents are posited to be interested directly in the health of their individual children, as indicated by a weighted sum of these anthropometric measures. This interest may be due to an altruistic interest in the children themselves or due to the perception that the children's health is related to some output of longer-run interest to the parents, such as the children's adult labor market capacities (more on this below). The nutrient inputs are the five emphasized by RBRP.

Table II gives weighted seemingly unrelated regression (SUR) system estimates of the basic model, with separate estimates for 390 pairs of children for the lean season and 379 pairs of children for the surplus season.[14] The estimates vary significantly between the two seasons, thus supporting emphasis by authors in Chambers, Longhurst and Pacey and in Sahn on the critical nature of seasons in poor agrarian areas.[15]

The lean-season estimates are in the top half of the table. These estimates have a number of interesting features. First, the $\hat{\eta}_{ik}$'s suggest that weight is the most important health indicator, with significant positive effects for parental preferences pertaining to the distribution of calories, β-carotene, riboflavin and calcium. Weight-for-height has significant positive effects for calories and vitamin C,[16] and arm circumference is significant for vitamin

[14] The weights are used to assure that each household is weighted equally even though different households have different numbers of pairs of children. Single-child households, of course, are excluded from the analysis. The number of pairs of children differ between the lean and surplus seasons because a few children were not present in both seasons.

[15] What appears to be different parental utility parameters between the seasons may reflect the existence of a more general utility function that encompasses both sets of seasonal estimates. However exporations undertaken with generalizations of the CES utility function did not lead to a single empirical estimate that encompasses both seasons without seasonal differences in the parameters.

[16] The puzzling significantly negative effect for β-carotene also should be noted.

TABLE II

System Estimates of Basic Multi-Health Indicator, Multi-Nutrient Input Intrahousehold Allocation Model with Male Preference for Children in Rural Semi-Arid Tropical India 1976–78.[a]

LEAN SEASON:	η_{ik}			Residual Correlations				
	Weight-for-Height	*Arm Circumference*	*Weight*	*Calories*	*β-Carotene*	*Riboflavin*	*Vitamin C*	*Calcium*
Calories	0.67(3.1)	−0.07(0.2)	0.58(2.2)	1.00				
β-Carotene	−1.24(2.3)	−0.56(0.5)	3.17(3.5)	0.48	1.00			
Riboflavin	0.50(1.6)	−0.43(1.0)	1.10(2.6)	0.63	0.31	1.00		
Vitamin C	1.28(2.6)	1.23(2.9)	−0.78(1.8)	0.39	0.39	0.25	1.00	
Calcium	−0.24(0.6)	0.21(0.5)	0.88(2.1)	0.30	0.54	0.28	0.15	1.00
Inequality Aversion (c)		0.47(3.4)	SSE	4.6	32.4	12.2	12.7	24.7
Gender Preference (ε)		0.046(2.7)	R^2	0.89	0.62	0.75	0.64	0.63
SEE = 1918, MSE = 4.92								

SURPLUS SEASON:	η_{ik}			Residual Correlations				
	Weight-for-Height	*Arm Circumference*	*Weight*	*Calories*	*β-Carotene*	*Riboflavin*	*Vitamin C*	*Calcium*
Calories	0.92(9.5)	0.03(0.4)	0.04(0.5)	1.00				
β-Carotene	0.70(1.5)	−2.22(4.9)	1.73(2.8)	0.00	1.00			
Riboflavin	0.75(6.3)	−0.16(1.4)	0.37(3.5)	0.77	0.40	1.00		
Vitamin C	0.31(2.0)	0.05(0.4)	0.61(4.5)	0.52	0.36	0.69	1.00	
Calcium	0.76(6.4)	0.25(2.3)	−0.04(0.4)	0.24	0.41	0.45	0.15	1.00
Inequality Aversion (c)		−1.74(14.1)	SSE	11.0	263	13.8	15.5	20.7
Gender Preference (ε)		0.032(1.0)	R^2	0.69	0.38	0.62	0.53	0.67
SEE = 1788, MSE = 4.72								

[a] Absolute t values are in parentheses to the right of point estimates.

C.[17] Second, the SSEs and R^2s suggest that the model is most consistent ($R^2 = 0.89$) with the distribution of calories, the most important nutrient, and least consistent with the distribution of β-carotene ($R^2 = 0.62$). Third, the correlations among all of the residuals are positive, ranging from 0.15 between vitamin C and calcium to 0.63 between calories and riboflavin, thus suggesting the existence of some systematic factor(s) affecting allocation of nutrients among children beyond those incorporated into the model.

Of primary interest are the estimates of parental inequality aversion (c) and of gender preference (ε). Parental inequality aversion is estimated to be 0.47, significantly above the Cobb–Douglas value of zero and significantly below the pure-investment value of one. While this value indicates some parental concern with distribution, such concern is quite limited in comparison with previous estimates for the United States of about 0.10 in BPT.[18] It suggests that in the lean season these rural Indian households allocate nutrients to their children in a manner more approximating a pure investment model with much more limited distributional concern than has been observed for schooling allocations in the United States. However this c probably does not exceed $\bar{c}$ in relation (10); therefore only health-related outcomes—but not also nutrients—are distributed more unequally than is the impact of endowments in W.

Male-gender preference, finally, is estimated to be significantly positive at 0.046. That is, during the lean season parents weigh a given health-related outcome for boys almost 5% more heavily than the identical health-related outcome for girls. This result suggests that when food supplies are very tight these parents exhibit male preference, and do not just respond to higher marginal returns to nutrient investments in males. This issue is explored further below.

The surplus-season estimates are in the bottom half of Table II. As for the lean season, the surplus-season estimates are most consistent with the allocation of calories ($R^2 = 0.69$) and least consistent with the allocation of β-carotene ($R^2 = 0.38$) and there is considerable positive collinearity among the disturbance terms. But there are three important differences between the lean- and surplus-season estimates. First, the surplus-season estimates suggest that weight-for-height is a more critical indicator of health-outcomes than is weight. Second, parental inequality aversion in the surplus season with $\hat{c}$ equal to -1.74 is significantly stronger than for the lean season and for the values for the United States in BPT; a value of c below the Cobb–Douglas value of zero implies that in the surplus season parents distribute nutrients to favor the less well endowed children sufficiently so

[17] The weights for the fourth anthropometric measure, tricep skinfold thickness, are defined as one minus the other three weights (see the last sentence in note 11). Therefore they could be positive only for calcium.

[18] This comparison must be qualified because, as noted above, $\hat{c}$ may be biased towards the pure-investment case and the estimates in BPT also may have this bias, but not necessarily of the same magnitude.

that the relative health outcomes are more equal than the impact of the relative endowments (i.e., compensatory behavior).[19] Third, the point estimate of gender preference is less than two thirds as large for the surplus season as for the lean season and is not significantly nonzero even at the 10% level.

Thus the picture emerges of these households exhibiting promale preferences and following closer to a pure investment strategy with relatively limited concern for equity during the lean season when food supplies are relatively limited even though there is not significant promale preference and there is considerable inequality aversion (probably more than estimated for the United States) when food is relatively abundant in the surplus season.

Multi-nutrient input, multi-health output intrahousehold model with expected adult labor-market outcomes included

As noted above, parents may be interested in their children's health in part because that health relates to adult labor market productivity (and perhaps the sharing of the fruits of that productivity with the parents). Rosenzweig and Schultz, for example, emphasize this possibility in their study of gender differences in Indian child mortality. Since adult labor market productivity is posited to depend on adult health which is associated with child health, in fact, it is possible that the estimates in Table II are only reflecting this relation and what appears to be male preference in the lean season only reflects greater returns to male than to female health.

To explore this possibility, I posit that additional health-related outcomes are the expected probability of adult labor force employment and the expected adult wage of a current child, both conditional on current health characteristics. To construct these expected adult labor market outcomes, substantial stability over time in the labor market relations and in the individual child's health status as he or she grows into adulthood both are assumed. Regarding the former, as Rosenzweig and Schultz argue in a similar context, such an assumption is relatively palatable in a society that is relatively stable as has been the SAT part of rural India.[20] Regarding the latter, there plausibly is a link between childhood and adult health status.

To be more explicit, first probits for labor force employment probabilities and semilog wage functions dependent on health for men and women in the same villages where the children in the sample live were estimated. Then these results were used to estimate expected adult labor force employment probabilities and wages for current boys and girls, conditional on their expected adult health status being perfectly correlated with their current

[19] The qualification of the previous note again applies.
[20] In fact the assumption of such stability seems more legitimate for the SAT areas under study here than for the broader geographical area considered by Rosenzweig and Schultz since SAT India has been relatively unaffected by the Green Revolution and other changes that have had substantial labor market effects in other parts of India.

childhood health status. These estimated labor force employment probabilities and wages then are used as additional health-related outcomes in the nutrient allocation model.[21]

What happens if the specification in Table II is expanded to include as additional outcomes these expected adult labor market outcomes? Do the constructed labor market outcome variables dominate over the direct health indicators in Table II? Of course such a test is limited because of considerable multicollinearity and the assumptions necessary for the constructed labor market employment probabilities and wages to represent the relevant expected adult outcomes. Nevertheless, the results may be informative. They suggest that adding the expected labor market outputs to the direct anthropometric indicators of health-related outcomes used in Table II does not improve the empirical power of the model to explain actual nutrient allocations.[22] For neither the lean nor the surplus season, in fact, are any of the coefficients of the constructed labor market variables significantly positive in the expanded formulation. If the labor market outcomes alone are included in the model, moreover, the estimates are not as consistent with the empirical patterns as are the estimates in Table II. Subject to the qualifications about the labor market outcome measures already made, such patterns suggests that the estimates in Table II are not dominated by estimates which focus on the health returns in the form of expected adult labor market outcomes. Therefore the lean-season male preference does not seem to reflect primarily differential expected labor-market returns to health by gender, which contrasts with the Rosenzweig and Schultz claim about the importance of such labor-market returns in understanding the distribution of nutrients between girls and boys.[23]

Sensitivity of parental gender preference to landholding class, caste, education of the household head

The results to this point suggest that, at least during the lean season, south Indian rural households exhibit male preference in their allocation of nutrients among their children. However there are frequent conjectures that the extent of gender preference may vary with household characteristics.

[21] The labor force employment probability and wage relations posit that each depends on a number of standardized anthropometric health indicators with a quadratic in age to control for life-cycle paths. Separate regressions were estimated for men and for women since Ryan and Ghodake report significant differences in wage functions by gender for the same villages. The anthropometric measures included are the four used in Table III plus standardized height since height is posited to reflect long-run health status. The estimates differ significantly between the two sexes. The results reported in the text are not sensitive to variations in the specification that were explored.

[22] These results are summarized verbally here, but are not presented in detail in order to conserve space.

[23] Since Rosenzweig and Schultz focus on mortality rates rather than indicators of health-related outcomes for surviving children, they may be correct to emphasize the importance of expected labor market outcomes for mortality if the determinants of survival differ from the determinants of health conditional on survival.

For example, Sen and Sengupta suggest that the extent of male preference may depend upon whether or not the household holds land.

To explore such possibilities the model in relation (7) is now extended to incorporate the possibility that the extent of gender preference in the lean season[24] depends on various household characteristics:

$$\varepsilon = \varepsilon_0 + \varepsilon_1 E + \varepsilon_2 L + \varepsilon_3 C \tag{11}$$

Three household characteristics are considered: education of the household head (E), whether the household holds land (L) above some minimum level, and its relative caste (C) rank (in one of the seven rank-order caste groupings defined by an anthropologist familiar with the sample villages[25]) of the household within the village in which it resides.

In most respects these extended estimates in Table III are similar to those in Table II: the coefficients of the η_{ik}'s suggest the same general patterns (though the standard errors are larger), the correlations of the residuals are quite similar, and the inequality-aversion parameter (c) is not significantly different from that in Table II. The estimated lean-season dependencies of male preference on the education of the household head and on relative caste rank are significant. For the caste variable the estimates suggest that gender preference is inversely associated with caste rank. These caste effects are fairly substantial, with a difference in the weights of 0.22 to 0.32 between the highest and lowest castes in a village. In combination with the constant component of gender preference (and assuming no education for the household head), they suggest that parental gender preference favors boys only for the lower three or four caste groups (out of seven)—and girls may be favored by the highest ranked castes. The estimates also suggest that households with more-educated heads favor boys slightly more in the lean season, a result that I find somewhat surprising.

3. Summary

The extended household allocation model developed in this paper to examine parental gender preference and the special data utilized in this study have permitted a more systematic exploration than here-to-fore undertaken of critical parameters pertaining to parental inequality aversion and gender preference in the allocation of nutrients among children in rural Indian households. Of course the results have to be qualified because they are conditional on particular functional forms and a particular data set.[26] Subject to this caveat, the estimates have some interesting aspects.

[24] Only the lean season estimates are presented to conserve space since only for the lean season is gender preference estimated to be significant.

[25] Details of the caste structures in the sample villages and the caste rankings are discussed in Doherty.

[26] The general pattern of preference parameters estimates of particular interest—inequality aversion (c) and promale preferences (ε)—are robust to use of a generalized CES preferred function (note 6) and variations in the set of indicators for health-related outcomes.

TABLE III

System Estimates of Basic Multi-Health Indicator, Multi-Nutrient Input Intrahousehold Allocation Model with Male Preference Dependent on Education of Household Head and Caste Rank, Lean Season in Rural Semi-Arid Tropical India, 1976–78.[a]

LEAN SEASON:	η_{ik}		*Residual Correlations*					
	Weight-for-Height	*Arm Circumference*	*Weight*	*Calories*	*β-Carotene*	*Riboflavin*	*Vitamin C*	*Calcium*
Calories	0.80(2.3)	−0.33(0.6)	0.90(1.8)	1.0				
β-Carotene	−1.87(1.8)	−1.18(1.2)	4.57(2.3)	0.47	1.0			
Riboflavin	0.64(1.3)	−0.87(1.1)	1.62(2.0)	0.59	0.29	1.0		
Vitamin C	1.56(1.9)	1.43(2.2)	−0.93(1.4)	0.36	0.37	0.21	1.0	
Calcium	−0.48(0.7)	0.20(0.3)	1.17(1.6)	0.29	0.54	0.26	0.14	1.0
Inequality Aversion (c)		0.34(2.3)	SSE	4.0	29.8	11.0	12.0	24.5
Gender Preference (ε_0)		−0.17(2.5)	R^2	0.90	0.64	0.76	0.64	0.62
Gender Preference * Education (ε_1)		0.025(3.7)						
Gender Preference * Land Holder (ε_2)		0.058(1.2)						
Gender Preference * Caste Rank (ε_3)		0.035(3.3)						
SEE = 1782, MSE = 4.89								

[a] Absolute t values are in parentheses to the right of point estimates.

They suggest that the preference parameters governing allocation of nutrients among children differ significantly across seasons, thus supporting the emphasis of authors in Chambers, Longhurst and Pacey, in Sahn, and others on the importance of seasonal differences in poor rural societies. As a result analysis based on the surplus season alone may be misleading regarding, for example, the impact on less-well endowed children of increments in household food or other resources.

There also apparently is a promale bias in parental preferences on the order of magnitude of five percent in the lean season, which probably is not just a response to differential expected labor market returns to nutrient investments in boys versus those in girls. This promale bias is not associated with land ownership as Sen and Sengupta have conjectured, but does seem to be related to the caste rank in a manner such that lower-ranked castes exhibit more male preference and to the education of the household head in a manner which increases the extent of promale bias as education increases. Of course what I identify as a preference bias might reflect differential returns by gender in some other unobserved outcomes, such as performance of rites associated with death. With more information, such possibilities could be explored in future research.[27] But the exploration in this paper of whether control for expected labor-market outcomes dominates apparent gender preferences is a significant advance over the previous literature.

Finally, the extent of parental inequality aversion in the lean season is less than that observed in the United States (but vice versa for the surplus season),[28] indicating behavior more like the pure-investment case in rural India when food is most limited. However, there is evidence that parents do consider equity as well as productivity, so pure investment models may be misleading. Nevertheless, for the lean season when food is scarcest the combination of limited inequality aversion and promale preferences—particularly for the lowest ranked castes—may leave those children who are less-well endowed, especially if they are low-caste females, close to or even below the margin for survival.

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[27] There may be a danger in the extreme, however, of definining unobserved outcomes so that preferences cannot differ by gender. For example to say that the returns are higher to investing in boys than in girls purely because parents gain more pleasure from boys than from girls and therefore there are no preference biases would not seem to me to be very useful definitions of outcomes and preferences.

[28] Subject to the qualifications about the relative magnitudes of biases in notes 18 and 19.

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