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A prominent if not distinguishing feature of low-income countries that has been incorporated into many models of behavior in such settings is the proximity of average income-levels to subsistence. Models of savings behavior (Gersovitz, 1983) and wage determination (Leibenstein, 1957; Stiglitz, 1976), for example, have demonstrated the possiblity that at low income levels behavior may be quite distinct from that observed where income levels are well-above those required for survival. In a subsistance regime, the allocation of foods is thus particularly important and there have been a large number of recent empirical studies concerned with how households in low-income environments distribute foods among their members.

One salient aspect of the distribution of foods in low-income settings that has caught the attention of many social scientists is the disparity in nutrients received by women compared to men, particularly in South and West Asian societies.[1] One hypothesis that has been advanced is that gender-based nutrient inequality reflects disparities in labor market opportunities across men and women in these settings, with the pecuniary returns to a household from the allocation of food to women being less than those for men. While some empirical studies have shown the existence of a relationship between sex differences in infant mortality rates and differences in labor market participation rates between men and women (Bardhan, 1974; Rosenzweig and Schultz, 1982), there is little evidence on how the actual intrahousehold distribution of food across individuals is related to labor market characteristics nor a clear theoretical linkage established between market opportunities and patterns of intrahousehold food allocations.

One important study of the intrahousehold distribution of nutrients (Behrman, 1988b) finds no apparent link between expected labor market opportunities and sex disparities in nutrient consumption. However, this study only looks at the allocation of foods among children less than 13, a large proportion of whom do not participate in the labor market and, perhaps more importantly, among whom there may be little differentiation with respect to work activities. For this group, the link between the labor market and food consumption can only be indirect and is in any case not explicitly modelled.

In this paper we examine the relationship between the household distribution of foods and labor market opportunities in the context of a model incorporating (i) linkages between food consumption, health and labor market productivity and (ii) individual heterogeneity in inherent healthiness. A number of recent studies have shown in the context of low-income environments that health and food consumption affect productivity and wage rates (Strauss, 1986, Behrman and Deolaliker, 1987). While these studies have not linked their findings to household behavior, if the relationship between health and productivity differs across occupations and activities, between rickshaw pulling and mat weaving, say, and occupational participation is constrained for women, then the disparities in occupations and activities across men and women may be reflected in their food consumption.[2]

While attention has mainly focussed on gender inequality in food allocations, if the type of work activity matters for food consumption, the distribution of activities across individuals within gender class should be related to the intrahousehold distribution foods. Indeed, both average sex differences in food consumption and within-gender inequality in food consumption appear to vary over the life-cycle. Table 1 presents the means of average household calorie consumption and the intrahousehold coefficient

Table 1
Household Distibutions of Calories by Age and Sex in Bangladesh

	Age<6		6≤Age<12		Age≥12	
	Males	Females	Males	Females	Males	Females
Mean household calorie consumption	891	751	1549	1536	2672	2063
H_o: No sex differential in means (d.f.)	2.35 (217)		0.25 (220)		609.1 (465)	
Mean household coefficient of variation	43.6	41.1	11.1	10.5	11.5	7.05
H_o: No sex differential in CVs (d.f.)	0.26 (38)		0.23 (29)		4.48 (143)	

Source: Nutrition Survey of Bangladesh, 1981-82.

of variation in calorie consumption by age and sex for a probability sample of 345 households from 15 villages in Bangladesh.[3] The sex and age-specific coefficients of variation are computed only for those households with two or more individuals in each group. These figures show that while there is a large (30 percent) and statistically significant difference in the average calories allocated to men and women aged 12 and above, there is no difference by sex in mean calories consumed by children ages 7 through 11. For children aged 6 and below, boys on average receive more calories than girls, however. Gender differences in average calorie consumption are thus highly age-dependent. Table 1 also shows, more interestingly, that mean within-household inequality in food consumption, measured by the coefficient of variation, among males aged 12 and over is also significantly higher than that for females of the same age -- the mean intrahousehold coefficient of variation in food consumption is 64 percent higher among males than among females. Among children less than 12 years of age, inequality in calorie consumption among boys and girls is similar, however.[4]

Table 2 displays the distribution of activities ranked by their energy requirements, within the same sex and age groups. These figures demonstrate that stratification by activities also varies by age and sex and in large part parallels what is observed in Table 1 for calorie consumption. Not surprisingly, young children below age six are very rarely engaged in activities requiring even light energy, and never in any of the three more energy-intensive occupational categories. However, a substantial number of older children aged 6 to 12 years are engaged in energy-using activities, and the share of girls engaged in activities requiring energy at moderate levels or above actually exceeds that of boys (5.3 percent versus 0.7 percent). The similarity in energy intensity and diversity of activities exhibited by girls

Table 2
Percentage Activity Distribution by Energy Requirements, Age and Sex

Energy Requirement	Age<6		6≤Age<12		Age≥12	
	Males	Females	Males	Females	Males	Females
Insignificant	98.7	99.3	65.9	73.7	26.8	20.6
Light	1.3	0.7	28.8	25.6	22.6	8.5
Moderate	0	0	0	4.5	2.82	68.2
Very high	0	0	0.7	0.8	31.9	1.2
Exceptionally high	0	0	0	0	15.9	1.5
Sample size	133	129	140	133	433	473
H_o: No sex differential χ^2(d.f.)	0.28(1) (.600)*		6.87(3) (.076)		625.4(4) (.0001)	

*Significance level (probability).

and boys in the below-six and 6-to-12 age groups mirror the similarity in the mean and variability in calorie consumption among boys and girls in those age-groups exhibited in Table 1. And the large disparities in participation rates in very and exceptionally high-energy intensive activities across men and women aged 12 and over are consistent with the calorie distribution differences depicted in Table 1 for that age group.

Tables 1 and 2 are suggestive of a direct linkage between work activities and the intrahousehold distribution of food consumption, but they of course do not explain the diversity in activities within gender groups. In section 1, we set out the model and derive implications for how the distribution of health endowments influence the distribution of food and activities across members of a household as conditioned by nutrition-productivity linkages. Section 2 discusses the methodology used to compute individual endowments and section 3 reports estimates, based on a sample of households from 15 villages in Bangladesh, of the effects of food consumption and activities on weight-for-height, the effects of health endowments on calorie consumption by sex and age and the effects of endowments on activity choice and income. The results indicate that calorie consumption affects weight-for-height equally for males and females. However, higher endowed men aged 12 and over receive higher levels of food in terms of calories and are more likely to participate in energy-intensive activities while higher-endowed women do not, contributing to a greater variance in calories consumed among men than among women in this age group, among whom males and females have similar endowment variances. These results are shown to be consistent with a model in which labor market considerations, and in particular the greater participation by men in activities in which health status importantly influences productivity, substantially influence the intrahousehold distribution of food. Households also display some degree of aversion to inequality in health outcomes, however, with men bearing slightly more of the "cost" of equalization than women as a consequence of their allocation to activities requiring high energy levels.

1. Theory

To analyze the relationship among the distribution of foods, health and labor market activities we set out a framework describing the allocation of food and the choice of labor market "effort" across heterogeneous individuals residing in integrated household units, defined by common objective functions. For simplicity we assume that there is only one food, or nutrient. The model can be readily extended to incorporate multiple foods and nutrients with no alteration in its basic implications. The health status h_i^k of an individual i among a class of individuals k, is assumed to be influenced by his/her food consumption c_i and by his/her effort e_i expended in some work activity, as described by (1). We assume that food augments health while effort decreases health such that,

$$(1) \qquad h_i^k = h^k(c_i, e_i, \mu_i) \qquad \partial h^k/\partial c_i > 0, \quad \partial h^k/\partial e_i < 0,$$

where μ_i is the endowed health of an individual, that component of health influenced by neither consumption nor effort.

Effort is rewarded in the labor market, with the returns to effort increasing with health status. The wage rate, w_i^k for an individual i in the class of individuals k is given by

$$(2) \qquad w_i^k = w^k(e_i, h_i) \qquad \partial w^k/\partial e_i,\ \partial w^k/\partial h_i > 0, \quad \partial^2 w^k/\partial e_i \partial h_i > 0.$$[5]

Individuals are assigned to classes (age, sex) by the characteristics of the health and effort wage functions so that every member of each class has the same h and w functions; individuals are individually differentiated by their health endowments μ_i, which are known to all family members. Expressions (1) and (2) capture the essential assumption of the nutrition-wage literature, that food consumption augments labor market productivity, presumably via health status (as found by Deolalikar (1988) for India). However, while the nutrition-based efficiency wage literature assumes a purely technological relationship between effort and health (or food consumption), here both food consumption and labor effort are choice variables.[6] Moreover, we allow the wage function to differ across classes of individuals, which may result from their allocation to particular sets of activities. Thus, for example, in India few women engage in ploughing; in Bangladesh no women are observed pulling rickshaws. The relationship between health and the returns to effort are likely to be quite different in those activities in which women and men participate.

The allocations of food and work effort across individuals in a household unit are determined as the solution to the maximization problem:

$$\max_{c_i^k, e_i^k} \quad U(h_1^k,\ldots,h_{n_k}^k,\ c_1^k,\ldots,c_{n_k}^k,\ e_1^k,\ldots,e_{n_k}^k) \qquad k = 1,\ldots,m \tag{3}$$

subject to

$$v + \sum_k \sum_i w_i^k - p \sum_k \sum_i c_i^k = 0 \tag{4}$$

and functions (1) and (2), where v = non-earned income and p is the price of the food good. In the household welfare function (3), it is assumed that increases in both health status and food consumption augment utility while increases in work effort lower utility.

The necessary first-order conditions for the allocation or assignment of food and work effort to individual i of class k are:

$$\frac{\partial U}{\partial h_i^k}\frac{\partial h^k}{\partial c_i} + \frac{\partial U}{\partial c_i^k} = \lambda \left[p - \frac{\partial w^k}{\partial h_i^k}\frac{\partial h^k}{\partial c_i} \right] \tag{5}$$

$$\frac{\partial U}{\partial h_i^k}\frac{\partial h^k}{\partial e_i} + \frac{\partial U}{\partial e_i^k} = -\lambda \left[\frac{\partial w^k}{\partial e_i} + \frac{\partial w^k}{\partial h_i}\frac{\partial h^k}{\partial e_i} \right], \tag{6}$$

where λ = marginal utility of income. Condition (5) states that the marginal cost of allocating an additional unit of food to person i is lower the greater the extent to which health augments work efficiency. Thus if the members of class ℓ participate in activities where the market returns to health are greater compared to those activities in which members of class k participate (who are otherwise identical), then on average class ℓ individuals will receive higher allocations of food than will class k individuals. Since we assume for simplicity that work (whether market or non-market)time is the same for all individuals (there are few idle women in low-income countries), it is not market work time (or even average wage rates) that matters for food allocation (as in Rosenzweig and Schultz, 1982), but the type of activity engaged in, as defined by the wage-effort-health association.

Within a class, the distribution of food and work effort across individuals will depend on the distribution of endowments. To highlight the roles of both health in the labor market and household preferences in influencing these distributions, assume that the endowment is additive in (1). Thus differences in endowments do not influence the health returns to food consumption. Consider first a model, nested in (3), in which household income is maximized. The maximand is the left hand side of expression (4), and the necessary first-order conditions are given by (5) and (6) with the left hand side of each expression replaced by zero. In the income-maximizing model the relationships between the endowment of an individual i in class k and his/her allocation of food and work effort are given by

$$\text{(7)} \qquad \frac{dc_i^k}{d\mu_i^k} = \left[\frac{\partial^2 w^k}{\partial e_i \partial h_i} \frac{\partial h^k}{\partial c_i} + \frac{\partial w^k}{\partial h_i} \frac{\partial^2 h^k}{\partial e_i \partial c_i} \right] \Phi^{-1} \frac{\partial^2 w^k}{\partial e_i \partial h_i} > 0$$

$$\text{(8)} \qquad \frac{de_i^k}{d\mu_i^k} = - \left[\frac{\partial w^k}{\partial h_i} \frac{\partial^2 h^k}{\partial c_i \partial c_i} \right] \Phi^{-1} \frac{\partial^2 w^k}{\partial e_i \partial h_i} > 0,$$

$$\text{where } \Phi = \left(\frac{\partial w^k}{\partial h_i} \frac{\partial^2 h^k}{\partial c_i \partial c_i} \right) \left[\frac{\partial^2 w^k}{\partial e_i \partial e_i} + \frac{\partial^2 w}{\partial e_i \partial h_i} \left(\frac{\partial h^k}{\partial e_i} \right)^2 + \frac{\partial w^k}{\partial h_i} \frac{\partial^2 h^k}{\partial e_i \partial c_i} \right]$$

$$- \left[\frac{\partial^2 w^k}{\partial e_i \partial h_i} \frac{\partial h^k}{\partial c_i} + \frac{\partial w^k}{\partial h_i} \frac{\partial^2 h^k}{\partial e_i \partial c_i} \right] \cdot \left[\frac{\partial^2 w^k}{\partial e_i \partial h_i} + \frac{\partial w^k}{\partial h_i} \frac{\partial^2 h^k}{\partial e_i \partial c_i} \right] > 0.$$

Under the income-maximization regime, those individuals with greater endowments of health supply more effort (participate in effort-intensive activities) since their labor market returns are higher but also receive more food, as long as the health returns to food consumption are not (entirely) diminished by increased effort. The greater allocation of food occurs because effort depletes health status; increased food consumption compensates for increased effort. Thus, those individuals exerting greater effort or in effort-intensive activities will also be consuming more food. Moreover, those classes of individuals in activities where the returns to work effort are more sensitive to health status will be characterized by greater differences in food consumption (and effort) compared to an otherwise identical class of individuals with the same distribution of endowments. This is because the magnitude of the (positive) endowment-food consumption slope depends positively in (7) on the degree to which health augments market returns to effort.

In the utility maximization model (3), the relationships among own endowments, food consumption and work effort are given by:

$$
\begin{aligned}
\frac{dc_i^k}{d\mu_i^k} = {} & \left[p - \frac{\partial w^k}{\partial h_i^k}\frac{\partial h^k}{\partial c_i}\right]\left(\frac{\partial h^k}{\partial c_i}\right)^{-1}\left[-\frac{\partial^2 h^k}{\partial c_i \partial c_i}(S_{c_i c_i}) + \frac{dc_i^k}{dv}\right] \\
& - (S_{c_i e_i})\frac{\partial^2 w^k}{\partial e_i \partial h_i} + \frac{dc_i^k}{dv}\frac{\partial w^k}{\partial h_i}
\end{aligned}
\tag{9}
$$

$$
\begin{aligned}
\frac{de_i^k}{d\mu_i^k} = {} & \left[\frac{\partial w^k}{\partial e_i} + \frac{\partial w^k}{\partial h_1}\frac{\partial h^k}{\partial e_i}\right]\left(\frac{\partial h^k}{\partial e_i}\right)^{-1}\left[-\frac{\partial^2 h^k}{\partial e_i \partial e_i}(S_{e_i e_i}) + \frac{de_i^k}{dv}\right] \\
& + (S_{e_i e_i})\frac{\partial^2 w^k}{\partial e_i \partial h_i} + \frac{de_i^k}{dv}\frac{\partial w^k}{\partial h_i},
\end{aligned}
\tag{10}
$$

where dc_i^k/dv and de_i^k/dv are income effects on food and effort, $S_{c_i c_i}$ and $S_{e_i e_i}$ are the Hicks-Slutsky compensated own substitution effects (positive and negative respectively), and $S_{e_i c_i}$ is the Hicks-Slutsky cross compensated substitution effect, which is negative if effort (a "bad") and food consumption are substitutes. The first of the three right-hand side terms in (9) and (10) arise from the welfare function in (3). In the absence of labor market health returns, this term indicates that the relationships between own endowments, food consumption and effort depend on the relative magnitudes of substitution and income effects. If income effects are small then higher-endowed individuals receive <u>less</u> food and provide more labor market effort. Some of their higher health is thus taxed away via both the food and effort allocations; low endowment individuals are "compensated" for their low endowments by higher food and lower effort allocations.

The last three right-hand side terms in (9) and (10) arise because of the health-effort interaction in the labor market. Both of these terms are positive in the food allocation equation (9), given that food is a normal good. Thus, the association between own endowments and food consumption will be algebraically higher the more strongly health augments the returns to effort. If women are barred (or refrain) from participating in activities in which health status strongly affects productivity, then compensation (reinforcement) with respect to food is more (less) likely than among men.

We note that defining compensation with respect to the sign of the relationship between own endowments and an individual-specific input, such as foods, can be misleading when more than one allocated item affects health status and welfare. In the profit-maximizing regime, we saw that the more-endowed household members received more food (reinforcement) but also expended more effort. Indeed, in some sense the increased food consumption provided to the more endowed fuels their higher effort. An alternative method of gauging compensation is to examine the net change in health status associated with a change in endowment. This is given by (11), in the additive endowment case:

$$\text{(11)} \qquad \frac{dh_i^k}{d\mu_i^k} = 1 + \frac{\partial h_i^k}{\partial c_i^k}\frac{dc_i^k}{d\mu_i^k} + \frac{\partial h^k}{\partial e_i^k} + \frac{de_i^k}{d\mu_i^k} .$$

If the sum of the last two terms in (11) is negative (positive) then overall compensation (reinforcement) with respect to health occurs; reinforcement with respect to foods is thus not inconsistent with a household's aversion to inequality in health status. If (11) is equal to one, the household exhibits overall neutrality with respect to its joint assignments of individual-specific food and effort levels.

While it is clear that the signs of the own endowment effects on food and effort do not necessarily distinguish between the income and welfare maximizing models, the existence of cross endowment effects can only arise when household welfare is being maximized (and the welfare function is not linear in its arguments). A change in the endowment of individual j in class ℓ on the food consumption of individual i in class k, in the welfare maximization model, is given by:

$$(12)\quad \frac{dc_i^k}{d\mu_j^\ell} = \left[p - \frac{\partial w^\ell}{\partial h_j}\frac{\partial h^\ell}{\partial c_j}\right]\left(\frac{\partial h^\ell}{\partial c_j}\right)\left[-\frac{\partial^2 h^\ell}{\partial c_j \partial c_j}(S_{c_i c_j}) + \frac{dc_i^\ell}{dv}\right] - (S_{c_j e_j})\frac{\partial^2 w^\ell}{\partial e_j \partial h_j}$$

Here, the cross effect of j's endowment on i's food consumption is more negative (given that c_i^k and c_j^d are substitutes) the stronger is the relationship between health and effort productivity for j. Thus, the cross effect of a woman's endowment on a man's food consumption, given restrictions on women's activities as exhibited in many South Asian societies, will be algebraically greater than will the effect of a change in a man's endowment on the woman's food allocation, while own endowment effects will be algebraically greater for males. Of course, the net effects of endowments on allocations will in the general case also depend on whether or not there are differential returns to health associated with food consumption and effort in (1). Knowledge about both the health technology and the role of health in augmenting productivity is thus critical for understanding the determinants of the allocation of foods and effort levels.

2. Estimating the Relationships between Endowments and Houshold Resource Allocations.

To estimate the association between the endowments of members of a household, their food consumption and their expenditure of effort we employ the method, first used in Rosenzweig and Schultz (1983), in which the health technology (1) is estimated directly and, based on the technology parameter estimates and the actual resources consumed or expended by each individual, individual-specific endowments are computed. It is possible to learn about some aspects of intrahousehold behavior without estimating the underlying technology of health production, with suitable choices of functional forms for both the technology (1) and the household welfare function (3), as in Behrman et al (1982) and Behrman (1988). Estimation of the technology and endowments, however, permits more agnosticism regarding preferences because endowment effects on allocations, as in relations (9) through (11), can be robustly estimated with approximations to household decision rules because endowments, if properly estimated, should be orthogonal to all of the exogenous determinants of household resource allocations.

There are two principal problems with the "residual" endowment method. First, if endowments, which are not directly observed by the researcher, influence resource allocations, consistent estimates of the household production technology cannot be obtained using least squares; that is, c_i, e_i and the unobserved μ_i will be correlated in (1).[7] One method of identifying the technology is to make use of instruments. In this case, food prices, labor market variables reflecting labor demand and exogenous components of income determine resource allocations but do not directly affect health status, given food and activity levels. Identification is facilitated if foods are broken down into a reduced set of basic elements (nutrients) that are presumed to influence health, since the number of foods consumed, and thus food prices, is likely to exceed the number of nutrients that enter into (1).

A second less well-recognized problem with extracting estimates of endowments from estimates of the technology, which arises even when the technology is estimated consistently, is that the derived endowments will be measured with systematic error. Contrary to the assertions in Rosenzweig and Schultz (1986), the measurement errors in estimated endowments are not likely to be random because the technology inputs, in this case individual-specific levels of nutrients, are unlikely to be measured without error. Endowment effects estimated by least squares are thus unlikely to be consistent and the biases cannot necessarily be signed a priori.

To see the measurement error problem, and one solution, assume for simplicity that the health production function contains only one nutrient (calories). The endowment estimate is

$$(13) \qquad \hat{\mu}_i = H_i - C_i\hat{\Gamma}, \quad i=1,\ldots,n,$$

where $\hat{\Gamma}$ is the two-stage least squares estimate of the calorie effect of health. The estimated endowment $\hat{\mu}_i$ contains the true health endowment plus pure measurement error ν_i

$$(14) \qquad \hat{\mu}_i = \mu_i + \nu_i, \quad i=1,\ldots,n,$$

where $E(\mu_i \nu_i) = 0$. In the classical errors-in-variables case, if the noisy measure of health endowment were the only noisy regressor in a calorie demand equation so that

$$(15) \qquad C_i = a + b\hat{\mu}_i + e_i, \quad i=1,\ldots,n,$$

where e_i represents the error in measuring calorie consumption and all other regressors are suppressed, $E(e_i \nu_i) = 0$, the least squares estimator of b will be biased towards zero. As is well known, if the true health endowments are nonstochastic and $(1/n)\ \Sigma\mu_i^2$ converges as $n\to\infty$ to a positive finite limit $\sigma_{\mu\mu}$, then

$$(16) \qquad \operatorname*{plim}_{n\to\infty} \hat{b} = b \frac{\sigma_{\mu\mu}}{\sigma_{\mu\mu}+\sigma_{\nu\nu}} .^{8}$$

In computing the health endowment from estimates of (1), however, the measurement errors associated with the health endowment and calorie consumption of the nutrient will be negatively correlated ($\sigma_{e\nu} < 0$). This is made clear by writing the true (measured without error) endowment as

$$(17) \qquad \mu_i^* = H_i^* - C_i\Gamma,$$

where H_i^* and C_i^* are the (unobserved) true values of health and calorie consumption respectively. The observed values H and C have measurement errors u_i and e_i

$$(18) \qquad H_i = H_i^* + u_i$$

$$(19) \qquad C_i = C_i^* + e_i.$$

Therefore, the estimated endowment $\hat{\mu}_i$ is

$$(20) \qquad \hat{\mu}_i = (H_i^* + u_i) - (C_i^* + e_i)\hat{\Gamma}$$

$$= \mu_i + u_i - e_i\hat{\Gamma}$$

and the endowment measurement error is $\nu_i = u_i - e_i\hat{\Gamma}$.

If calorie consumption has a positive marginal product in the production of health ($\hat{\Gamma} > 0$), as is likely in low-income environments, the estimated endowment $\hat{\mu}_i$ will be negatively correlated with the calorie measurement error. With $\sigma_{e\nu} < 0$, the least squares estimator of b, will have a component biasing it towards zero plus a term biasing it negatively even if the (unobserved) true endowment and true calorie consumption are not correlated with any of the measurement errors:

$$(21) \qquad \underset{n\to\infty}{\text{plim}} \ \hat{b} = b \frac{\sigma_{\mu\mu}}{\sigma_{\mu\mu}+\sigma_{\nu\nu}} + \frac{\sigma_{e\nu}}{\sigma_{\mu\mu}+\sigma_{\nu\nu}} .$$

If b is positive, $\hat{b}$ is less than b unambiguously, but if b is negative, the sign of the bias is indeterminate. Thus, the biased least squares estimator would tend to reject reinforcement (conpensation bias) with respect to calories if it in fact were true.

Consistent parameter estimates in the presence of errors in variables can be obtained by using instrumental variables methods. The problem, of course, is finding a set of instruments that are correlated with the health endowment but not with the error ν_i. As discussed, even in the absence of measurement error, consistent estimation of the health production function requires (two-stage least squares) instrumental variable estimation as a result of the correlation between calorie consumption and the health endowment. The food prices and health programs that act as identifying instruments in correcting for this heterogeneity problem also serve to correct for errors-in-variables bias in the production function. However, since the exogenous variables are by assumption uncorrelated with the true endowment, these variables cannot correct the errors-in-variables problem contaminating the estimated endowment effects on calorie allocations.

Repeated observations on the measured-with-error variable or the availability of different but related indicators of the phenomena to be measured (for example, multiple proxy variables) are potential sources of instruments. The use of repeated observations to obtain consistent estimates of parameters in the presence of errors in variables has a long history in the statistics literature (Tukey, 1951). If individual-specific food intake and anthropometric measures of health were measured at more than one point in time, then even if all measurements of (calorie) consumption are made with error, all that is required for consistent instrumental variables estimation is that the period-specific errors be uncorrelated across time periods.

An estimation procedure for the pure repeated observations case has been set out by Tukey (1951). His approach estimates coefficients from a one-criterion analysis of variance, that is, as ratios of within-set and between-set variances and covariances. Our problem differs from the pure repeated observation case considered by Tukey, not only because the two measurement errors e_i and ν_i are not independent, but because errors are not likely to be independent across a single observational strata. That is, for any observational round $E(e_i e_j) \neq 0$ for $i \neq j$. This arises because the units of observation are individuals, groups of whom make up a much smaller number of households, and household unobservables are likely to be important. Moreover, if there are available multiple indicators of endowments in addition to repeated measures, they can be used as well, as long as the measurement errors in each endowment type are uncorrelated across time periods (noncontemporaneously) with the health endowment measurement error.

3. Empirical Results

a. Data and the Estimation of the Health Technology

As the previous discussion has made clear, to obtain direct estimates of endowment effects requires data that not only provide individual-specific information on health and consumption but contain i) sufficient cross-sectional variation in exogenous variables needed as instruments for estimation of the health technology and (ii) repeated observations on individuals to purge estimated endowments of measurement errors. The 1981-82 Nutrition Survey of Rural Bangladesh (Ahmad and Hassan, 1983) provides information on individual-specific food consumption and anthropometric measures of health along with other individual and household attributes for 385 households in 15 villages scattered throughout Bangladesh.[9] Intra-household food consumption information was collected once for 25 (out of 50 sampled) households in each of 12 randomly-selected villages and in 35 (out of 70 sampled) households in an industrial town. In addition, the same information was collected at four separate times within a year for 25 (out of 50 sampled) households in each of two of the remaining villages. These Bangledesh data thus permit estimation of the health technology from the full sample, as well as of endowment responses purged of measurement error, based on the longitudinal component of the data set.[10]

The intra-household dietary information was collected by specially trained female dietary investigators who measured dietary intake by weighing each individual's intake in the home over a 24 hour period. All individuals covered by the dietary survey were also examined by a clinician and measures of weight, height, skinfold thickness and mid-arm circumference were taken. Special effort was made to return to a household if a member was absent when visited by the clinical team. Information was also obtained on age, sex, education and relationship to the head of each household member, as well as whether female household members were pregnant or lactating. The occupation of each member was recorded and the energy intensity of his or her activity was coded using guidelines established by the Food and Agriculture Organization and the World Health Organization (see Appendix A). Also available are data on household income, land ownership, and source of drinking water. The prices of a wide variety of foods sold in the village market were obtained from the Agricultural Extension Service for both the date of the survey and prior periods.

We use the information on weight-for-height to measure health, which is considered a good short-run measure of nutritional status that will be sensitive to daily food consumption and activity levels. To estimate the health technology (1), food consumption was converted into nutrient intakes using conversion factors specific to Bangladeshi foods (Institute of Nutrition and Food Science, 1980). The typical Bangladeshi diet is very simple; an unusually large share of total nutrition is derived from the cereals rice and wheat. Cereals accounted for 87 percent of calorie consumption, as well as 78, 82, 84, 70 and 82 precent of the consumption of protein, iron, thiamine and niacin, respectively. As the consumption of each nutrient is a linear function of all foods consumed, the large share of consumption derived from just one food group makes the set of observed nutrient intakes nearly perfectly collinear. The consumption of most nutrients are thus nearly proportional to the consumption of any one of them. This collinearity suggests that the full set of individual nutrient coefficients would be imprecisely estimated in a health production function no matter how health was measured. In any case we would expect that weight-for-height, as an indicator of short-run health, should not respond substantially if at all to the intake of any nutrient except calories. Daily changes in weight reflect the difference between calories consumed and calories expended. Thus, given the diet composition we would expect that the addition of nutrients other than calories will not significantly improve the fit of a Bangladesh weight-for-height production function having calories as an input, as we test below.

To reflect calorie outflow we add to individual-specific nutrient consumption in the weight-for-height production function two dummy variables reflecting participation in occupations categorized as "very active" or "exceptionally active" based on the occupational data. We add as well dummy variables indicating whether a woman was pregnant or lactating at the time of the survey. Exogenous regressors included are age, age squared, sex, the interaction of sex and age, and a set of dummy variables indicating the source of the household's drinking water as between well, pond, tubewell or river/canal. In accord with the model we treat nutrients, activities, pregnancy and lactation as endogenous variables and estimate the production function using two-stage least squares. Identifying instruments include the household head's age and schooling level, household landholdings and the village food prices interacted with individual age and sex variables, household landholdings and the head's schooling and age. The food prices include rice, wheat flour, potatoes, other roots, leafy vegetables, okra, green chillies, other nonleafy vegetables, sugar and sweets, eggs, mustard oil, pulses, tumaric, fish, milk, onions, garlic, other spices and meat.

Table 3 presents both (inconsistent) ordinary least squares (OLS) and consistent two-stage least squares (2SLS) estimates of the parameters of the Cobb-Douglas weight-for-height production function. These estimates are obtained using the full set of 15 villages (with one round from each of the two multiple-round villages). The calorie elasticity is seriously underestimated by OLS, although it is positive and statistically significant using either procedure. Moreover, the OLS estimates of the effect of the energy intensity of effort on weight-for-height are of the opposite sign of the consistent 2SLS estimates. The 2SLS estimates thus indicate that increased calorie consumption significantly increases weight-for-height and demonstrate that participation in exceptionally active occupations tends to deplete weight-for-height, although the estimated activity coefficient has a relatively large standard error. The less active occupations categorized as "very active" have an estimated coefficient only one-eighth of "exceptionally

Table 3
Effects of Calorie Consumption, Activity Level and Pregnancy Status on Weight-for-Height[a]

	Ordinary Least Squares Estimates	Two-Stage Least Squares Estimates
Calorie consumption[b]	.0295	.136
	(4.09)[c]	(3.37)
Very active occupation[b]	.0859	-.0119
	(5.34)	(0.23)
Exceptionally active occupation[b]	.0668	-.0817
	(3.43)	(1.26)
Pregnant[b]	.262	.326
	(7.69)	(1.34)
Lactating[b]	.144	.513
	(9.28)	(4.65)
Age	.284	.0987
	(16.6)	(1.90)
Age squared	-.00456	.0174
	(1.44)	(2.37)
Sex (male=1)	.00196	-.0578
	(0.08)	(1.81)
Age · sex	.0152	.0687
	(1.74)	(4.04)
Water drawn from tubewell	-.0478	-.0406
	(3.13)	(2.10)
Water drawn from well	-.0720	-.0693
	(4.11)	(3.15)
Water drawn from pond	-.0460	-.0649
	(2.30)	(2.55)
Constant	-2.56	-3.12
	(52.4)	(13.9)
R^2	.775	-
F	395.1	-
n	1737	1737
H_o: No influence of calcium, carotene, thiamine and riboflavin consumption[b] (F)	-	1.23
H_o: No difference in effect of calorie consumption by sex (F)	-	2.16

a. All variables in logs, except sex, water sources and activity level.
b. Endogenous variable: Instruments include household head's age and schooling level, landholdings and prices of all foods consumed interacted with individual age and sex variables, land and head's schooling and age.
c. Asymptotic t-ratios in parentheses.

active" occupations.

We also tested whether calorie consumption was a sufficient statistic for nutrient consumption and whether the calorie elasticity differed across males and females. The null hypothesis that four additional nutrients found by Ryan <u>et al</u> (1984) to be potentially important determinants of short-run health in a rural area of India -- calcium, carotene, thiamine and riboflavin -- do not influence weight-for height in our sample could not be rejected ($F(4,1724) = 1.23$). The null hypothesis that the calorie output elasticity is the same for men and women also could not be rejected ($F(1,1724) = 2.16$). Thus, gender differences in the production of weight-for-height are sufficiently well specified as an age-dependent intercept shift. Except for the first 2.5 years of life, Bangladeshi males are predicted to have greater weight-for-height than females having identical levels of inputs.

b. <u>Endowments and Calorie Consumption</u>

Having obtained estimates of the health technology, we can compute health (weight-height) endowments for each individual based on his/her actual calorie consumption and activity. However, to estimate the effects of endowments on calorie allocations requires instruments, as discussed above, given the likelihood that individual calorie consumption is measured with error. In order to use the repeated measure methodology we use the longitudinal sample from the two villages, which provides four rounds of data for 50 households in two villages (Jorbaria and Falshattia). We also use as instruments estimated endowments of mid-arm circumference and skinfold thickness derived from production functions, estimated by two-stage least squares, containing the same regressors and instruments as the weight-for-height production function. Three instruments for an individual's weight-for-height endowment in a period τ are thus constructed -- the estimated endowments of the three health attributes averaged over the survey rounds in which the individual was present <u>excluding</u> period τ. Formally, the j set of instruments $Z_{i\tau}$ associated with the endowment of individual i in period τ are constructed as

$$Z^j_{i\tau} = (1/(T_i-1))\Sigma_{t\neq\tau}\hat{\epsilon}_{i,t}\ , \quad j\text{=weight-for-height, skinfold thickness, mid-arm circumference,}$$

where T_i is the number of repeated measures (rounds) available for person i. Instruments for the mean weight-for-height endowments of groups (classes) of family members in period τ are constructed as the group means of the individual specific means $Z^j_{i\tau}$.

The household welfare-maximization model, as noted, implies that the calories allocated to an individual in the household depend on his/her characteristics -- age, sex and endowment -- and also the characteristics of all other household members, plus household or village-specific characteristics such as health program availability and food prices. With respect to village-level variables, since our longitudinal sample is taken from only two villages, a village dummy variable captures all village-specific determinants. However, a difficulty with estimating a calorie allocation equation from a sample of individuals who are members of relatively large households is the large set of parameters that needs to be estimated. If there are m classes of individuals, T household characteristics (including an intercept) and n household members, then $m\cdot n$ + T parameters need to be estimated. Moreover, the size of households vary in the sample and households of differing size and technology classes (age and sex) will belong to different choice regimes -- that is, parameters of the calorie allocation equation are specific to a technology class configuration. Samples of households having common demographic attributes will be small and are likely to impart selection bias to the results.

One solution to this problem is to assume that the effect on the calories consumed by member 1 of household ℓ due to changes in the characteristics of household members 2 through n_ℓ of household ℓ are captured by moments of the distributions of the characteristics of household members and also do not directly depend on household size n_ℓ. The first two columns of Table 4 present parameter estimates of (log) calorie allocation equations in which the characteristics of other household members are summarized by household means. Individual determinants of calories consumed include own age, age squared, sex and endowment, and hence the variables reflecting the cross effects of other household members include mean age, mean age squared (variance of ages), proportion of household members male and the mean of the household's endowments. Household-specific variables include water sources and family income. Family income is clearly endogenous if wages depend on endowments, calorie allocations and the level of effort. Identifying instruments for family income include household landholdings, household head's schooling and head's age.

The first column of Table 4 provides two-stage generalized least squares estimates of the logarithmic calorie allocation equation, estimated with the full sample of individuals from the two villages, but in which instruments are not used for the endowment variables. The second column provides parameter estimates that use instruments for the endowment variables. As predicted, the uninstrumented coefficient estimate for own endowment is algebraically less than the (positively-signed) instrumented own endowment coefficient estimate -- indeed, it is of opposite sign, indicating compensation when there is net reinforcement with respect to calories.[12] The uninstrumented family or cross endowment parameter is algebraically greater

Table 4
Two-Stage Generalized Least Squares Estimates: Effects of Personal and Family Characteristics on the Allocation of Personal Calorie Consumption[a]

Variable	All Family Members: No Instruments for Endowments	All Family Members: Instruments for Endowments	Males	Females
Individual weight /height endowment	-.145 (1.98)[c]	.132 (1.33)	.676 (4.14)	.0662 (0.26)
Family endowment	-.867 (1.01)	-1.15 (0.75)	-	-
Family endowment males	-	-	-.743 (1.72)	-.414 (1.27)
Family endowment females	-	-	-.325 (1.10)	-.0709 (0.20)
Family income[b]	.0640 (2.04)	.122 (3.95)	.0839 (1.98)	.0961 (2.07)
Age	1.35 (25.8)	1.35 (24.7)	1.44 (17.5)	1.33 (15.7)
Age squared	-.199 (19.8)	-.199 (19.0)	-.200 (12.9)	-.196 (11.6)
Sex (male=1)	-.0191 (0.24)	.0492 (0.60)	-	-
Age · sex	.0679 (2.63)	.0801 (2.97)	-	-
Mean age of family members	-.0711 (1.10)	-.114 (1.66)	.00318 (0.03)	-.122 (1.44)
Variance of ages of family members	-.0757 (1.39)	-.115 (2.03)	-.0837 (0.99)	-.120 (1.61)
Proportion of family members male	-.0330 (0.32)	-.0588 (0.54)	.00762 (0.04)	-.0777 (0.55)
Water drawn from tube well	.227 (3.03)	.221 (2.81)	.162 (1.58)	.271 (3.29)
Jobaria village	.245 (3.36)	.254 (3.34)	.196 (1.95)	.283 (3.54)
Constant	4.86 (18.3)	4.65 (16.7)	4.52 (10.3)	4.82 (13.2)
χ^2 (no family error component)	245.6	243.5	129.0	38.06
$\sigma_\mu^2/(\sigma_\epsilon^2+\sigma_\mu^2)$	.205	.204	.234	.218
n	806	806	407	371

a. All variables in logs, except sex, water source, location and sex ratio.
b. Endogenous variable. Instruments for income and endowments are: household landholdings, household head's schooling, head's age; mean of individual and family endowments for weight/height, skinfold thickness, and arm circumference from all survey rounds excluding the round from which observation is drawn.
c. Asymptotic t-ratios in parentheses.

than the consistent estimate, but as the consistent estimate is negative, the sign of the bias could not be predicted unambiguously a priori. A comparison of the first two columns also reveals that other parameters are biased substantially as well. The consistent family income coefficient is twice that estimated without instrumenting endowments, although both estimates treat income as endogenous. The coefficients on the mean and variance of ages also differ substantially by estimation method.

The consistent estimate of the own endowment effect in column 2 of Table 4 suggests that there is reinforcement with respect to calories, although the coefficient is not statistically different from zero at standard levels of significance ($t = 1.32$). Income effects on calorie consumption are positive and significant. Own and cross age effects are also significant as is an age-sex interaction. The latter's positive sign suggests that calories are increasing with age for males. Adding second and third moments of the family endowment, age and sex variables to the specification did not significantly improve the fit.

The activity distributions reported in Table 2 indicate that there are important gender differences in labor market participation and thus, as our framework suggests, endowment effects should differ by gender. In columns 3 and 4 of Table 4 we provide two-stage generalized least squares estimates of calorie allocation equations stratified by sex (with all endowments instrumented). As the theory predicts, the own endowment effect for males, the group with heavier participation in activities intensive in effort and in which shorter-run health status is likely to be an important determinant of labor market returns, is algebraically greater than that for females, whose participation in highly effort-intensive activities is limited. The estimates indicate that a 10 percent increase in a male's endowment increases his calorie allocation by 6.8 percent; the own endowment effect for females is one tenth that of males.

Corresponding to the positive own endowment effect for males, the cross effect of the endowment of other males in the household is negative. These results thus reject the pure income-maximizing model, since if households allocate calories and effort so as to maximize income, all cross effects will be zero. Moreover, if households maximize welfare rather than income, we have seen that the cross endowment effect of those in the labor market (males) should be of opposite sign to the male own endowment effect, as we have found. The theory also predicts that if there is calorie reinforcement, the effect of an increase in a female's health endowment on the calories allocated to others in the household should be less in absolute value than an increase in a male's health endowment if health status is less important for women in their activities. Changes in a female's health endowment presumably does not induce as large a return to effort (because of evident limits to occupational choice) as do changes in a male's health endowment. In both the male and female calorie allocation equations of Table 4, columns 3 and 4, the effect of the mean endowment of females on calorie consumption is indeed considerably less in absolute value than the effect of the mean endowment of males.

The parameter estimates reported in Table 4 may specify cross effects in an imperfect manner by assuming they can be represented by the means (and higher moments) of household distributions. A method of estimating own endowment effects that requires no assumptions about the parameterization of the household variables employs a household fixed effects estimator, in which the full set of cross terms and household-specific regressors are impounded in the fixed effect. Estimation is accomplished with a sample transformed by differencing individual observations from the mean values of the household to which they belong. Although this approach is more likely to avoid specification error it, of course, prevents identification of the parameters associated with family endowments and other household-specific regressors.

Table 5 reports household fixed effects two-stage generalized least squares estimates of the effects of personal characteristics on individual calorie consumption. The sample has been stratified by sex since the theory predicts (and our previous results confirm) that the calorie response of males to the endowments, ages and gender characteristics of other family members will differ from the calorie responses of females, given the stratification of activities by gender. Two-stage instrumental variables estimation is still required because the own endowment remains a regressor. The set of instruments for the own endowment measure remains the same.

While the fixed effects estimator purges all household unobservables from the error term of the calorie allocation equation, the errors are not likely to be independent across observations because individuals appear in the sample more than once and individual-specific unobservables may be important. Therefore, the estimator used is a fixed effects two-stage generalized least squares estimator. The null hypothesis of no individual error components is indeed rejected by Breusch-Pagan (Lagrange multiplier) test statistics in each equation estimated.[13]

Columns 1 and 3 of Table 5 report the within-household, gender-specific (logarithmic) calorie allocation equations having own endowment, own age and own age squared as regressors. The parameter estimates diverge very little from those reported in Table 4. The elasticity of calorie consumption with respect to own health endowment is 0.447 ($t = 3.58$) for males, indicating reinforcement, and only -0.028 ($t=-0.15$) for females.

Table 5

Fixed Effects Two-Stage Generalized Least Squares: Effects of Personal Characteristics on Individual Calorie Consumption[a]

Variable	Males		Females	
Own endowment[b]	.447 (3.58)[c]	-	-.0278 (0.15)	-
Endowment·(age<6)[b]	-	-.435 (1.35)	-	-.314 (0.46)
Endowment·(6≤age<12)[b]	-	.923 (2.29)	-	1.86 (2.13)
Endowment·(age≥12)[b]	-	1.21 (2.69)	-	.0894 (0.13)
Age	1.44 (22.9)	1.31 (14.9)	1.34 (18.1)	1.35 (17.9)
Age squared	-.201 (16.7)	-.170 (9.16)	-.199 (13.4)	-.206 (13.7)
H: No individual error components (χ^2)	46.51	48.35	32.36	26.17
$\sigma_\mu^2/(\sigma_\mu^2 + \sigma_\epsilon^2)$	.287	.300	.258	.282
n	429	429	371	371

a. All variables in logs.
b. Instrumental variables used; see Table 4.
c. Asymptotic t-ratios in parentheses.

The activity patterns by age and sex exhibited in Table 2 suggests that the distributions of both males and females among occupations (as reflected by their energy intensity) varies with age. In columns 2 and 4 of Table 5 we report estimates obtained using specifications of the calorie equation in which own endowment effects are allowed to vary across the three age groups that appear to be related to the differentiation of activity patterns within the samples of males and females. The pattern of estimated own endowment effects matches up well with the pattern of activities presented in Table 2. Both male and female young children (aged less than six years) have the (algebraically) smallest own endowment effects -- there is no labor market return to higher endowments for these family members and thus calorie compensation dominates -- part of the better health derived from a higher endowment is "taxed" away by the household solely via the allocation of foods.

Male and female children aged 6 to 12 years evidently engage in a more diverse set of activities by energy intensity and the own endowment parameters for this age group reflect this. The own endowment coefficients exhibit reinforcement and are statistically significant for both males and females. A 10 percent increase in the health endowment of a 6 to 12 year old child increases calorie consumption by 9.2 percent if the child is male and 18.6 percent if the child is female. The higher rate of reinforcement for girls in this age group is consistent with their greater diversity of activity by energy intensity displayed in Table 2.

The estimated effects of own health endowment on calorie consumption for adults (ages 12 and above) is also consistent with the patterns presented in Table 2. Adult males exhibit the greatest diversity of activity choice ranked by energy-intensity among all age/sex groups, while adult females have very limited diversity. Reflecting this ability to alter the intensity of effort, the own endowment elasticity of calorie consumption for adult males is large (1.21), positive and statistically significant while for adult females it close to zero (0.09).

Our estimates of differences in calorie responses to endowments by age and sex, associated with differences in activity distributions, would also appear to explain the patterns in intrahousehold calorie variability by age and sex exhibited in Table 1 since $\mathrm{var}(c^k)$ depends on $\beta_k^2\mathrm{var}(\mu^k)$, where β_k is the estimated own endowment effect for age-sex group k, as we cannot reject the hypothesis that all endowment variances are equal; indeed, the endowment variance for adult females is slightly higher than that of males. Our estimates of the β_k thus imply that the variation in the effects of health on productivity across activities is in part responsible for the higher variability in intrahousehold calorie allocations among males relative to females for persons aged 12 and over, and contributes to the variability among girls and boys for those household members aged between 6 and 12 years, with no effect among boys and girls less then 6.[14]

c. Endowments, Family Income, and Activity Participation

While the results thus far are consistent with there being a return to health in the labor market, it remains to demonstrate with these data that income is positively associated with health endowments and that those with higher endowments are more likely to choose activities with a greater energy-intensity of effort, as implied by the theory. While the Bangladesh data do not provide information on individual-specific wage rates or earnings, we can test if households with higher average health endowments among adult males, given land resources and schooling, have higher incomes. Table 6, column 1 provides estimates of the determinants of (log) per-capita income. These

Table 6

Determinants of the Log of Per-Capita Household Income and Probability of Participating in an "Exceptionally Active" Occupation Among Persons Aged 12-60 years

Variable/ Estimation Procedure	Per-Capita Income Two-stage Least Squares	Exceptionally Active Occupation Full-information ML IV Probit
Own endowment[a]	-	13.9 (1.64)
Family endowment-males≥12[a]	2.38 (2.86)[b]	-16.74 (2.29)
Family endowment-females≥12[a]	0.378 (0.75)	-3.67 (1.25)
Age	-	1.28 (1.06)
Sex	-	6.92 (2.36)
Land	.0200 (0.64)	-.0219 (2.46)
Head's schooling	.109 (1.80)	-1.09 (1.54)
Mean age of family members	.0444 (0.14)	-1.58 (1.18)
Variance of ages of family members	.591 (1.91)	-5.55 (2.50)
Proportion of family members male	.566 (1.09)	-4.11 (1.82)
Jorbaria village	-.199 (1.30)	-5.98 (2.19)
Constant	4.23 (4.11)	4.95 (1.17)
n	45	153
F	3.73	-
$\chi^2(12)$	-	76.2

a. Instrumented.
b. Asymptotic t-ratios in parentheses.

show that income is positively and significantly associated with the average endowments of males older than 12 years of age, but, as expected, the adult female endowment elasticity of income is only one-sixth as large as the male endowment elasticity and not statistically different from zero.

Table 6 also reports maximum likelihood instrumental variable estimates of a probit activity choice equation for individuals aged 12 to 60 in column two. The zero-one dichotomous dependent variable in this equation has the value of one if an adult is engaged in an exceptionally active occupation -- the only activity category that substantially reduced weight-for-height in the estimated production function (Table 3).[15] Here, own endowment has a positive and statistically significant (at the 10 percent level) effect on the probability of participating in an exceptionally active activity. In addition, consistent with the calorie allocation estimates of Table 4, the male family endowment has a large negative influence on this probability, five times larger than the influence of the female family endowment -- given one's own endowment, an increase in that of other males in the family significantly reduces the likelihood of being in an energy-intensive activity. The coefficient on sex (male=1) is positive and statistically significant, reflecting the importance of gender in circumscribing occupational choice, given endowments. Indeed, only two females in the sample were engaged in an exceptionally active occupation. Thus, the results reported in Table 6 confirm that there is a pecuniary return to health and effort, that adult males with higher endowments are more likely to undertake exceptionally energy intensive work, and that adult female health endowments are relatively unimportant in determining activty choices or household income compared to adult male endowments.

Finally, the net effect of a change in own endowment on own health (equation (11)) can be calculated from the estimates of the health technology in Table 3 and the estimated endowment effects on calories and activities in Tables 5 and 6. Calculating the elasticity of health with respect to own endowment is straightforward as a result of the logarithmic functional forms chosen for all estimated relationships. For both adult males and females (aged 12 years and above) our estimates indicate that, in addition to its direct effect on health, an increase in endowment tends to increase health by increasing calorie consumption and to reduce health by inducing greater intensity of effort. The latter indirect effect dominates the former for both sexes. The elasticity of own health with respect to own endowment is 0.88 for adult males and 0.97 for adult females. Bangladesh households thus exhibit compensatory behavior with respect to health. Moreover, as the difference between the endowment elasticity and unity can be thought of as a "tax" levied by the household on the exogenous health of its members, our estimates indicate that the exogenous health of adult males is taxed at a higher rate than the exogenous health of adult females (12 percent versus 3 percent).

4. Conclusion

In this paper we have examined the determinants of calorie consumption and activity choices from the perspective of a model of intrahousehold allocation that incorporates the effects of differences in exogenous healthiness, health technology and labor market returns to health and effort on the allocation of resources to members. The empirical analysis was applied to unique individual and household-level data from Bangladesh, a country that exhibits large differences in calorie consumption and the energy-intensity of activity by age and gender. Our results, obtained using an estimation procedure that takes into account the endogeneity of health and the likelihood of error in the measurement of both calorie intake and health endowments, reveal that energy-intensive effort tends to reduce health as measured by weight-for-height, that there is a pecuniary return to health and effort, and that there is substantial calorie reinforcement for those classes of individuals best able to alter the energy-intensity of effort. In particular, adult males (aged 12 years and above) and male and female children (aged 6 to 12) were found to receive calorie reinforcement with respect to their health endowments. These classes of individuals were also those exhibiting the most diverse activity choices ranked by energy intensity. Thus, linkages between health levels and productivity, combined with the circumscribed activities of women in Bangladesh, appear to account for part for the disparities in the average consumption of nutrients across adult men and women and contribute to a greater variability across men in nutrient consumption. Furthermore, it was found that adult males with higher health endowments were more likely to undertake exceptionally energy intensive work.

Our results also reject the income maximizing model of the household in favor of a model in which households exhibit some aversion to inequality. Indeed, even though the rate of calorie reinforcement for adult males was quite high (1.21 in elasticity) and almost zero for adult females, the greater response of adult male effort to endowments resulted in a "tax" on adult male endowment that exceeded that of adult females (12 percent versus 3 percent).

Our evidence that disparities by gender in consumption in a low-income society like Bangladesh, marked by evident occupational stratification by gender, reflect labor market activities and are not only the result of gender bias in food allocation suggests that increases in labor force opportunities for women, _ceteris paribus_, will likely increase the calories allocated to women. However, as we have shown, the health and welfare benefits of such an increase in calorie consumption by women will be tempered by the increased level of energy-intensive activity associated with greater calorie consumption. Furthermore, while the lifting of occupational constraints on women may likely reduce (calorie) consumption inequality between the sexes, it will increase inequality within the class of adult women and thus may increase overall inequality in consumption and health. The increase in inequality among women will reflect the increased importance of the distribution of endowments in determining the distribution of calories when there is a greater return to effort and health in the labor market.

Footnotes

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1. For an extensive review of the literature concerned with gender inequality and the intrahousehold distribution of food, see Behrman (1988a).

2. Behrman and Deolalikar (1987) find that in their sample of Indian households, health (measured by weight for height) had a significant effect on the wage rate of males but not on the wage of females. They conjecture that this disparity in health effects on labor productivity results from the division of labor and might influence the allocation of nutrients. Sahn and Alderman (1988) also find that calorie consumption affected the wage rate of rural males but not females in rural Sri Lanka.

3. We describe below the characteristics of this data set. Calorie consumption in Bangladesh is a good indicator of overall nutrient consumption, given the simplicity of the Bangladeshi diet, as discussed below.

4. Similar patterns characterize the Indian village data used by Behrman (1988b). He shows that there are sex differences in mean nutrient allocations for children below 13, the sample population he studies. However, we find using the same data set, that for individuals aged 13 and above, mean calorie consumption is 12 percent higher for males. The variance in consumption among males is 15.6 percent higher than it is among females in that age group. Both of these differences are statistically significant.

5. We also assume that the marginal product of health vanishes if there is no effort so that the second derivative of health in the wage function is

zero. If (2) is quadratic, for example, then $w = h(e + \gamma e^2)$.

6. We assume that work time is fixed (and set to unity) as is conventially assumed in the nutrient-wage literature. It is possible to include in work time home production activities, with (2) being replaced by a goods production function with no alteration in the basic implications of the model.

7. The model estimated by Behrman (1988b) also requires a solution to the simultaneity problem that arises in the presence of the unobservable endowments. In that study, however, this problem is recognized but not dealt with.

8. The parameters associated with the other regressors measured without error are also biased; the sign of their bias can be determined from the variance-covariance matrix of the observations (Levi, 1973).

9. The data from one additional village of hill tribes -- who are not racially or ethnically related to Bengalis -- were not used in our analysis as their dietary and other behaviors are considered too unlike that of ethnic Bengalis.

10. Hassan (1984) has compared the nutritional information in the survey with that collected in prior nutrition surveys in Banlgadesh (and East Pakistan), to draw inferences concerning trends in Bangladeshi health and food consumption.

11. Pitt (1983) shows that nutrient consumption is significantly responsive to food prices in Bangladesh.

12. The coefficient on own endowment in these logarithmic calorie allocation equations should be interpreted as the elasticity of own health with respect to own endowment conditional on mean family endowments remaining fixed. This elasticity then corresponds to the experiment in

which a transfer of endowment occurs within the household that leaves mean endowments unchanged. This same interpretation also applies to the own age and sex coefficients.

13. Note that only household and not individual random effects were specified in estimating the calorie allocation equations of Table 4. The Breusch-Pagan statistics of Table 5 confirm the importance of individual effects even when controlling for household random effects. The parameter estimates of Table 4 are nonetheless consistent but standard errors are underestimated by about 10 percent based on our experience with obtaining the estimates reported in Table 5.

14. To test if there was seasonality in endowment effects we tested whether endowment responses varied with income by interacting the endowment and age variables with household income using the household fixed effects procedure. We could not reject the hypothesis that endowment and age effects were independent of income levels.

15. The likelihood maximized is given in Smith and Blundell (1986).

Table A

Energy Requirement Levels by Occupation: FAO/WHO Standards[a]

Activity Level	Occupation[b]
Light active	Businessman, teacher, student, doctor, unemployed, retired.
Moderately active	Skilled laborer, housewife.
Very active	Farmer, house servant, fisherman.
Exceptionally active	Agricultural laborer, non-agricultural unskilled laborer, rickshaw puller.

a. FAO/WHO Ad Hoc Expert Committee, Energy and Protein Requirements, WHO Technical Report Series N-522, WHO, Geneva, 1973.

b. Occupational titles from Ahmed, K. and Hassan, N., Nutrition Survey of Rural Bangladesh, Institute of Nutrition and Food Science, University of Dhaka, Bangladesh, May 1986.

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