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# Productivity, Health, and Inequality in the Intrahousehold Distribution of Food in Low-Income Countries

By MARK M. PITT, MARK R. ROSENZWEIG, AND MD. NAZMUL HASSAN\*

*A model is formulated incorporating linkages among nutrition, labor-market productivity, health heterogeneity, and the intrahousehold distribution of food and work activities in a subsistence economy. Empirical results, based on a sample of households from Bangladesh, indicate that, despite considerable intra-household disparities in calorie consumption, households are averse to inequality. Furthermore, consistent with the model, the results also indicate that both the higher level and greater variance in the calories consumed by men relative to women reflect in part the greater participation by men in activities in which productivity is sensitive to health status. (JEL 824, 122, 850).*

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 (Mark Gersovitz, 1983) and wage determination (Harvey Leibenstein, 1957; Joseph Stiglitz, 1976; Partha Dasgupta and Debraj Ray, 1984), for example, have demonstrated the possibility that behavior at low income levels may be quite distinct from that observed when income levels are well above those required for survival. Low-income societies are also characterized by an occupational distribution in which activities requiring high levels of energy expenditure predominate, and a number of recent studies have shown that health and food consumption directly affect productivity and wage rates in low-income environments (John Strauss, 1986; Anil Deolalikar, 1988; Jere Behrman and Deolalikar, 1989). In a subsistence regime, the allocation of

food is thus particularly important, and the measurement of the overall level of inequality in low-income countries must take into account how households in such environments distribute food among their individual members.

One salient aspect of the distribution of food 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.<sup>1</sup> One hypothesis that has been advanced is that gender-based nutrient inequality reflects disparities in labor-market opportunities between 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. Indeed, 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 (Pranab Bardhan, 1974; Rosenzweig and T. Paul Schultz, 1982). However, there is little direct evidence of a relationship between the actual intrahousehold distribution of food across individuals and labor-market activities; nor is there a clear theoretical linkage established between labor-market characteristics and patterns of intrahousehold

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<sup>1</sup>For an extensive review of the literature concerned with gender inequality and the intrahousehold distribution of food, see Behrman (1990).

TABLE 1—HOUSEHOLD DISTRIBUTIONS OF CALORIES BY AGE AND SEX IN BANGLADESH

Statistic	Age < 6			6 ≤ Age < 12			Age ≥ 12		
	Males	Females	$X^2(d.f.)$	Males	Females	$X^2(d.f.)$	Males	Females	$X^2(d.f.)$
Mean household calorie consumption	891	751	2.35 (217)	1,549	1,536	0.25 (220)	2,672	2,063	609.1 (465)
Mean household coefficient of variation	43.6	41.1	0.26 (38)	11.1	10.5	0.23 (29)	11.5	7.05	4.48 (143)

Source: Nutrition Survey of Bangladesh, 1981–2.

food allocation that may arise in low-income environments.<sup>2</sup>

Although attention has mainly focused on gender inequality in food allocation, if the relationship between healthiness and productivity differs across occupations and activities, the distribution of activities across individuals within gender classes should also be related to the intrahousehold distribution of foods. Table 1 presents the means of average household calorie consumption and the intrahousehold coefficient of variation in calorie consumption by age and sex for a probability sample of 345 households from 15 villages in Bangladesh.<sup>3</sup> These figures show that, while there is a large (30 percent) and statistically significant difference in the average number of calories allocated to men and women aged 12 and above, there is no difference between sexes 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 aver-

age 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, is 64 percent higher among males aged 12 and over than among females of the same age. Among children less than 12 years of age, however, inequality in calorie consumption among boys and girls is similar.<sup>4</sup>

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. The similarity in energy intensity and diversity of activities exhibited by girls 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. Furthermore, the large disparities in participation rates in high-energy-intensive activities between men and women aged 12 and over are consistent with the gender differences in the variability of calorie consumption depicted in Table 1 for that age group.

Tables 1 and 2 are suggestive of a direct linkage between the type of work activities

<sup>2</sup>One important study of the intrahousehold distribution of nutrients (Behrman, 1988) finds no apparent link between expected labor-market opportunities and sex disparities in nutrient consumption. However, this study only considers the allocation of foods among children less than 13 years of age, 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 modeled.

<sup>3</sup>Later in this paper, we describe 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 in Section III. The sex- and age-specific coefficients of variation are computed only for those households with two or more individuals in each group.

<sup>4</sup>Similar patterns characterize the Indian village data used by Behrman (1988). He shows that there are no sex differences in average nutrient allocations for children younger than 13, the subset of the population he studies. However, using the same data set, we find 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.

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	70.5	69.1	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 ( <i>N</i> )	133	129	140	133	433	473
$X^2$ ( <i>d.f.</i> )	0.28 (1)		6.87 (3)		625.4 (4)	
$P^a$	0.600		0.076		0.0001	

<sup>a</sup>Significance level (probability).

and the intrahousehold distribution of food consumption, but they of course do not explain the diversity in activities within gender groups. In this paper, we examine the relationship between the household distribution of foods and labor-market activities in the context of a model incorporating (i) linkages among food consumption, health, and labor-market productivity and (ii) individual heterogeneity in inherent or “endowed” healthiness. The model takes as given differences in the opportunities for work activities by gender; conditional on the circumscribed activities of women, it yields implications for how the distribution of individual health endowments and nutrition-productivity linkages influence the distribution of food and energy expenditure (effort) across individual members of a household and provides a method for measuring gender-based discrimination by the household. Section I presents the model. Section II discusses the methodology used to compute individual endowments, and Section III 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 empirical results appear to be consistent with the hypothesis that work activity distributions substantially influence the intrahousehold distribution of food. In particular, the greater participation by men in

energy-intensive activities in which health status may importantly influence productivity is in part responsible for both the higher level of calories consumed by adult men and the greater variance in calories consumed among men compared to women. We are able to infer from our estimates, however, that households are averse to inequality in health outcomes, with men bearing slightly more of the “cost” of equalization than women as a consequence of their participation in activities requiring high energy levels.

### I. Theory

To analyze the relationships among the distribution of food, 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 food consumption  $c_i$  and by effort  $e_i$  expended in some work activity. In general, the effects of these variables on health may be nonmonotonic. However, we assume that in a subsistence economy food augments health, while effort decreases

health (stamina) such that

$$(1) \quad h_i^k = h^k(c_i, e_i, \mu_i)$$

$$\frac{\partial h_i^k}{\partial c_i} > 0 \quad \frac{\partial h_i^k}{\partial e_i} < 0$$

where  $\mu_i$  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 the following:<sup>5</sup>

$$(2) \quad w_i^k = w^k(e_i, h_i)$$

$$\frac{\partial w_i^k}{\partial e_i}, \frac{\partial w_i^k}{\partial h_i} > 0 \quad \frac{\partial^2 w_i^k}{\partial e_i \partial h_i} > 0.$$

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  $\mu_i$ , 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. 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.<sup>6</sup> Moreover, we allow the wage function to differ across classes of individuals, which may result from their

<sup>5</sup>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  $w = h(e + \gamma e^2)$ .

<sup>6</sup>We assume that work time is fixed (and set to unity) as is conventionally assumed in the nutrient-wage literature. It is possible to include home production activities in total work time, with (2) being replaced by a goods-production function, with no alteration in the basic implications of the model. Our data set contains no information on the amount of time allocated to any activity.

allocation to particular sets of activities. Thus, for example, in India, few women engage in plowing; 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 both women and men participate. Indeed, Behrman and Deolalikar (1989) and David Sahn and Harold Alderman (1988), based on data from India and Sri Lanka, respectively, found that health (measured as weight-for-height) and calorie consumption had significant positive effects on the wage rates of men but not women.

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

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

subject to

$$(4) \quad v + \sum_k \sum_i w_i^k - p \sum_k \sum_i c_i^k = 0$$

and functions (1) and (2), where  $v$  = nonearned 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

$$(5) \quad \left( \frac{\partial U}{\partial h_i^k} \right) \left( \frac{\partial h^k}{\partial c_i} \right) + \frac{\partial U}{\partial c_i^k}$$

$$= \lambda \left[ p - \left( \frac{\partial w^k}{\partial h_i^k} \right) \left( \frac{\partial h^k}{\partial c_i} \right) \right]$$

$$(6) \quad \left( \frac{\partial U}{\partial h_i^k} \right) \left( \frac{\partial h^k}{\partial e_i} \right) + \frac{\partial U}{\partial e_i^k}$$

$$= -\lambda \left[ \frac{\partial w^k}{\partial e_i} + \left( \frac{\partial w^k}{\partial h_i} \right) \left( \frac{\partial h^k}{\partial e_i} \right) \right]$$

where  $\lambda$  = 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 for which 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 nonmarket) time is the same for all individuals (there are few idle women in low-income countries), it is not market work time (or even the average wage rate) 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 that person's allocation of food and work effort are given by

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

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

where

$$\Phi = \left( \frac{\partial w^k}{\partial h_i} \right) \left( \frac{\partial^2 h^k}{\partial c_i \partial c_i} \right) \times \left[ \frac{\partial^2 w^k}{\partial e_i \partial e_i} + \left( \frac{\partial^2 w}{\partial e_i \partial h_i} \right) \left( \frac{\partial h^k}{\partial e_i} \right)^2 + \left( \frac{\partial w^k}{\partial h_i} \right) \left( \frac{\partial^2 h^k}{\partial e_i \partial c_i} \right) \right] - \left[ \left( \frac{\partial^2 w^k}{\partial e_i \partial h_i} \right) \left( \frac{\partial h^k}{\partial c_i} \right) + \left( \frac{\partial w^k}{\partial h_i} \right) \left( \frac{\partial^2 h^k}{\partial e_i \partial c_i} \right) \right] \times \left[ \frac{\partial^2 w^k}{\partial e_i \partial h_i} + \left( \frac{\partial w^k}{\partial h_i} \right) \left( \frac{\partial^2 h^k}{\partial e_i \partial c_i} \right) \right] > 0.$$

As indicated by equation (8), under the income-maximization regime those individuals with greater endowments of health supply more effort, because health augments the labor-market returns to effort ( $\partial^2 w^k / \partial e_i \partial h_i > 0$ ). More-endowed individuals also receive more food because food increases health, which increases the returns to effort, and because effort depletes health status; increased food consumption both compensates for and enhances the return from 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 for which 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-con-

sumption slope in (7) depends positively 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

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

$$(10) \quad \frac{de_i^k}{d\mu_i^k} = \left[ \frac{\partial w^k}{\partial e_i} + \left( \frac{\partial w^k}{\partial h_i} \right) \left( \frac{\partial h^k}{\partial e_i} \right) \right] \\ \times \left( \frac{\partial h^k}{\partial e_i} \right)^{-1} \left[ - \left( \frac{\partial^2 h^k}{\partial e_i^2} \right) (S_{e_i e_i}) + \frac{de_i^k}{dv} \right] \\ + (S_{e_i e_i}) \left( \frac{\partial^2 w^k}{\partial e_i \partial h_i} \right) + \left( \frac{de_i^k}{dv} \right) \left( \frac{\partial w^k}{\partial h_i} \right)$$

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 (negative and positive, 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) arises from the welfare function in (3). This term indicates that the relationships among own endowments, food consumption, and effort depend on the relative magnitudes of substitution and income effects. If income effects are small, then in the absence of labor-market returns, higher-endowed individuals receive *less* 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 two 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 good affects health status and welfare. An alternative method of gauging compensation, and of more meaningfully assessing the differential treatment of different classes of individuals by the household, 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:

$$(11) \quad \frac{dh_i^k}{d\mu_i^k} = 1 + \left( \frac{\partial h_i^k}{\partial c_i^k} \right) \left( \frac{dc_i^k}{d\mu_i^k} \right) \\ + \left( \frac{\partial h_i^k}{\partial e_i^k} \right) \left( \frac{de_i^k}{d\mu_i^k} \right).$$

If the sum of the last two terms in (11) is negative (positive), then compensation (reinforcement) with respect to health occurs for group  $k$ ; reinforcement with respect to foods is thus not inconsistent with a household's aversion to inequality in health status. Expression (11) may differ across groups; intergroup differences in (11) thus are a measure of net discrimination across groups by the household with respect to health that incorporates both food and effort allocations.

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). It is straightforward to show that, although such effects cannot be signed in general, the cross effect of  $j$ 's endowment on  $i$ 's food consumption is more negative (given that the consumption of  $i$  and  $j$  are substitutes in the household welfare function) the stronger is the relationship between health and effort productivity for  $j$ .<sup>7</sup> Thus, the cross effect of a woman's endowment on a man's food consumption, given the gender differences in activities 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. 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.

## II. Estimating the Relationships between Endowments and Household Resource Allocations

To estimate the association among the endowments of members of a household, their food consumption, and their expendi-

<sup>7</sup>The 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

$$\begin{aligned} \frac{dc_i^k}{d\mu_j^l} = & \left[ p - \left( \frac{\partial w^l}{\partial h_j} \right) \left( \frac{\partial h^l}{\partial c_j} \right) \right] \left( \frac{\partial h^l}{\partial c_j} \right)^{-1} \\ & \times \left[ - \left( \frac{\partial^2 h^l}{\partial c_j \partial c_j} \right) (S_{c_i c_j}) + \frac{dc_i^l}{dv} \right] \\ & + (S_{c_i e_i}) \left( \frac{\partial^2 w^l}{\partial e_j \partial h_j} \right). \end{aligned}$$

The cross effect of  $j$ 's endowment on  $i$ 's food consumption is more negative (given that  $e_i$  and  $c_j$  are substitutes in the household welfare function) the stronger is the relationship between health and effort productivity for  $j$ .

ture of effort, we employ a 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. There are two problems with this "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  $\mu_i$  will be correlated in (1). One method of identifying the technology is to use 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.

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 (1983), 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 true (measured without error) endowment  $\mu_i^*$  is thus

$$(12) \quad \mu_i^* = H_i^* - C_i^* \Gamma \quad i = 1, \dots, n$$

where  $H_i^*$  and  $C_i^*$  are the (unobserved) true values of health and calorie consumption, respectively, and  $\Gamma$  is the calorie effect on health. Assume that the observed values  $H$  and  $C$  have measurement errors  $u_i$  and  $e_i$  with classical errors-in-variables proper-

ties; that is,

$$(13) \quad H_i = H_i^* + u_i$$

$$(14) \quad C_i = C_i^* + e_i$$

where  $E(H_i^*, u_i) = 0$ ,  $E(C_i^*, e_i) = 0$ , and  $E(u_i, e_i) = 0$ . Therefore, the estimated endowment  $\hat{\mu}_i$  is

$$(15) \quad \hat{\mu}_i = (H_i^* + u_i) - (C_i^* + e_i)\hat{\Gamma} \\ = \mu_i^* + u_i - e_i\hat{\Gamma}$$

where  $\hat{\Gamma}$  is the two-stage least-squares estimate of the calorie effect, and the endowment measurement error is  $\nu_i = u_i - e_i\hat{\Gamma}$ .

If there are no other observables or unobservables, the estimated (linear) calorie allocation equation is

$$(16) \quad C_i = a + b\hat{\mu}_i + e_i.$$

The least-squares estimator of  $b$ , if the true health endowments are nonstochastic and  $(1/n)\sum\mu_i^2$  converges as  $n \rightarrow \infty$  to a positive finite limit  $\sigma_{\mu\mu}$ , is

$$(17) \quad \text{plim}_{n \rightarrow \infty} \hat{b} = b \frac{\sigma_{\mu\mu}}{\sigma_{\mu\mu} + \sigma_{\nu\nu}} + \frac{\sigma_{e\nu}}{\sigma_{\mu\mu} + \sigma_{\nu\nu}}.$$

The first term in (17) corresponds to the classical error-in-variables bias. However, the second term appears because of the indirect estimation of the endowment from (12). When calorie consumption has a positive marginal product in the production of health ( $\hat{\Gamma} > 0$ ), then from (15),  $\sigma_{e\nu} < 0$ . Thus, if  $b$  is positive,  $\hat{b}$  will underestimate  $b$  unambiguously, but if the true endowment effect is negative, the sign of the bias is indeterminate. The (biased) least-squares estimator of the error-ridden endowment effect would tend to reject reinforcement with respect to calories if it in fact were true; errors in measurement in calories will make households appear to be more com-

pensatory with respect to calories than they really are.<sup>8</sup>

Consistent parameter estimates in the presence of errors in variables can be obtained by using instrumental-variables methods. 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. 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. Moreover, if there are available multiple indicators of health and thus 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.

### III. 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 1981-2 Nutrition Survey of Rural Bangladesh (Kamaluddin Ahmad and Hassan, 1986) provides information on individual-specific food consumption and anthro-

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

pometric measures of health along with other individual and household attributes for 385 households in 15 villages scattered throughout Bangladesh.<sup>9</sup> 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 Bangladesh data thus permit estimation of the health technology from the cross-sectional sample, as well as estimation of endowment responses purged of measurement error, based on the longitudinal component of the data set.<sup>10</sup>

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, who obtained measures of weight, height, skinfold thickness, and mid-arm circumference. Information was also obtained on the occupation of each household member, 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). The prices of a wide variety of foods sold in the village market were separately obtained in the survey, so that there is one price per commodity per village.

We use the information on weight-for-height to measure health, which is considered a good short-run measure of nutri-

tional 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). Calorie consumption, however, would appear to be a sufficient indicator of nutritional intake. The typical Bangladeshi diet is very simple; cereals account for 87 percent of calorie consumption, as well as 78, 82, 84, 70, and 82 percent of the consumption of protein, iron, thiamine, riboflavin, 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. Moreover, 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.

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.<sup>11</sup> 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 (well, pond, tube well, or river/canal). In

<sup>9</sup>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 those of ethnic Bengalis.

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

<sup>11</sup>A possibly superior procedure would have been to employ individual dummy variables for each of the 14 occupations provided in the data. Because we treat occupation as a choice variable, however, we would need more instruments than we have available to identify all of the individual activity effects on weight-for-height. The FAO/WHO categories provide a parsimonious way of representing occupations in terms of their consequences for short-term health. If the categorization is correct, our estimates are more efficient than those that would be obtained from the more agnostic specification, if it could be estimated.

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 are the household head's age and schooling level, household landholdings, and the village food prices interacted with household landholdings, the head's schooling and age, and the individual age and sex variables. The food prices are those for rice, wheat flour, potatoes, leafy vegetables, okra, green chilies, sugar and sweets, eggs, mustard oil, pulses, fish, milk, onions, garlic, and meat.<sup>12</sup>

Table 3 presents both (inconsistent) ordinary least squares (OLS) and consistent two-stage least squares (2SLS) estimates of the parameters of the Cobb-Douglas production function for weight-for-height. These estimates are obtained using the cross-sectional component of the data describing 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 to the consistent 2SLS estimates, indicating a possible strong relationship between activity choice and the health residual, containing the endowment. The 2SLS estimates 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 that of "exceptionally active" occupations.

We also tested whether calorie consumption was a sufficient statistic for nutrient consumption and whether the calorie elasticity differed between males and females. We could not reject the null hypothesis that

TABLE 3—EFFECTS OF CALORIE CONSUMPTION, ACTIVITY LEVEL, AND PREGNANCY STATUS ON WEIGHT-FOR-HEIGHT

Variable <sup>a</sup>	Ordinary least-squares estimates <sup>c</sup>	Two-Stage least-squares estimates <sup>c</sup>
Calorie consumption <sup>b</sup>	0.0295 (4.09)	0.136 (3.37)
Very active occupation <sup>b</sup>	0.0859 (5.34)	-0.0119 (0.23)
Exceptionally active occupation <sup>b</sup>	0.0668 (3.43)	-0.0817 (1.26)
Pregnant <sup>b</sup>	0.262 (7.69)	0.326 (1.34)
Lactating <sup>b</sup>	0.144 (9.28)	0.513 (4.65)
Age	0.284 (16.6)	0.0987 (1.90)
Age squared	-0.00456 (1.44)	0.0174 (2.37)
Sex (male = 1)	0.00196 (0.08)	-0.0578 (1.81)
Age × sex	0.0152 (1.74)	0.0687 (4.04)
Water drawn from tube well	-0.0478 (3.13)	-0.0406 (2.10)
Water drawn from well	-0.0720 (4.11)	-0.0693 (3.15)
Water drawn from pond	-0.0460 (2.30)	-0.0649 (2.55)
Constant	-2.56 (52.4)	-3.12 (13.9)
<i>N</i>	1,737	1,737
<i>R</i> <sup>2</sup>	0.775	—
<i>F</i>	395.1	—
<i>H</i> <sub>0</sub> : No influence of calcium, carotene, thiamine, and riboflavin consumption <sup>b</sup> ( <i>F</i> )	—	1.23
<i>H</i> <sub>0</sub> : No difference in effect of calorie consumption by sex ( <i>F</i> )	—	2.16

<sup>a</sup>All variables in logs, except sex, water sources, and activity level.

<sup>b</sup>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.

<sup>c</sup>Asymptotic *t* ratios in parentheses.

four additional nutrients found by James Ryan et al. (1984) to be potentially important determinants of short-run health in a rural area of India—calcium, carotene, thiamine, and riboflavin—do not influence

<sup>12</sup>Pitt (1983) shows that nutrient consumption is significantly responsive to food prices in Bangladesh.

weight-for-height in our sample. ( $F_{[4, 1724]} = 1.23$ ). 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, the difference between sexes in the production of weight-for-height is 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. Endowments and Calorie Consumption

Having obtained estimates of the health technology, we can compute health (weight-for-height) endowments for each individual based on actual calorie consumption and activity. In order to use the repeated-measure methodology to mitigate the effects of errors in measurement, we use the longitudinal component of the sample, which provides four rounds of data for 50 households in two of the villages (Jorbaria and Falshatia). 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  $\tau$  are thus constructed: the estimated endowments of the three health attributes averaged over the survey rounds in which the individual was present *excluding* period  $\tau$ .<sup>13</sup>

<sup>13</sup>Formally, the set of instruments  $Z_{i\tau}^j$  associated with the endowment of type  $j$  for individual  $i$  in period  $\tau$  is constructed as

$$Z_{i\tau}^j = \frac{1}{T_i - 1} \sum_{t \neq \tau} \hat{\mu}_{it}^j$$

where  $j$  = weight-for-height, skinfold thickness, or mid-arm circumference, and 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  $\tau$  are constructed as the group means of the individual-specific means  $Z_{i\tau}^j$ .

The household welfare-maximization model, as noted, implies that the calories allocated to an individual in the household depend on that person's characteristics (age, sex, and endowment), the characteristics of all other household members, and household or village-specific characteristics such as health-program availability and food prices. With respect to village-level variables, because our longitudinal sample is taken from only two villages, a village dummy variable captures all village-specific determinants. To summarize parsimoniously the intrahousehold distribution of the exogenous characteristics of household members, we computed the household means of those variables, namely 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 treated as an endogenous variable because wages are assumed to depend on endowments, calorie allocations, and the level of effort.

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 the 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 evidently net reinforcement with respect to calories.<sup>14</sup> The uninstrumented family or cross-endow-

<sup>14</sup>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.

TABLE 4—TWO-STAGE GENERALIZED LEAST-SQUARES ESTIMATES:  
EFFECTS OF PERSONAL AND FAMILY CHARACTERISTICS ON THE  
ALLOCATION OF PERSONAL CALORIE CONSUMPTION

Variable <sup>a</sup>	Two-stage least-squares estimates <sup>b</sup>			
	All family members		By sex	
	No instruments for endowments	Instruments for endowments <sup>c</sup>	Males	Females
Individual weight/height endowment	-0.145 (1.98)	0.132 (1.33)	0.676 (4.14)	0.0662 (0.26)
Family endowment	-0.867 (1.01)	-1.15 (0.75)	—	—
Family endowment, males	—	—	-0.743 (1.72)	-0.414 (1.27)
Family endowment, females	—	—	-0.325 (1.10)	-0.0709 (0.20)
Family income <sup>c,d</sup>	0.0640 (2.04)	0.122 (3.95)	0.0839 (1.98)	0.0961 (2.07)
Age	1.35 (25.8)	1.35 (24.7)	1.44 (17.5)	1.33 (15.7)
Age squared	-0.199 (19.8)	-0.199 (19.0)	-0.200 (12.9)	-0.196 (11.6)
Sex (male = 1)	-0.0191 (0.24)	0.0492 (0.60)	—	—
Age × sex	0.0679 (2.63)	0.0801 (2.97)	—	—
Mean age of family members	-0.0711 (1.10)	-0.114 (1.66)	0.00318 (0.03)	-0.122 (1.44)
Variance of ages of family members	-0.0757 (1.39)	-0.115 (2.03)	-0.0837 (0.99)	-0.120 (1.61)
Proportion of family members male	-0.0330 (0.32)	-0.0588 (0.54)	0.00762 (0.04)	-0.0777 (0.55)
Water drawn from tube well	0.227 (3.03)	0.221 (2.81)	0.162 (1.58)	0.271 (3.29)
Jorbaria village	0.245 (3.36)	0.254 (3.34)	0.196 (1.95)	0.283 (3.54)
Constant	4.86 (18.3)	4.65 (16.7)	4.52 (10.3)	4.82 (13.2)
<i>N</i>	806	806	407	371
<i>X</i> <sup>2</sup> (no family error component) <sup>e</sup>	245.6	243.5	129.0	38.06
Share of family error component variance in total error variance	0.205	0.204	0.234	0.218

<sup>a</sup>All variables in logs, except sex, water source, location, and sex ratio.

<sup>b</sup>Asymptotic *t* ratios in parentheses.

<sup>c</sup>Instruments for income and endowments are: household landholdings and household head's schooling and age; and means of individual and family endowments for weight/height, skinfold thickness, and arm circumference calculated over all survey rounds, excluding the round from which observation is drawn.

<sup>d</sup>Endogenous variable.

<sup>e</sup>Lagrange multiplier (Breusch-Pagan test).

ment parameter is algebraically greater 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 in Table 4 also reveals that other parameters are biased substantially as well.

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 ( $t = 1.32$ ). However, the activity distributions reported in Table 2 indicate that there are important gender differences in activities, and thus, as our framework suggests, endowment effects may 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).<sup>15</sup> 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. These differences are consistent with the theory, given the lack of participation by women in energy-intensive activities and the findings in similar settings that health matters for the wages of men but not women.

Corresponding to the positive and significant 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. 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 the effect of an increase in a male's health endowment, if health status is less important for women in their

activities. 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, although the difference is not statistically significant because of the imprecision with which both endowment effects are measured.

The parameter estimates reported in Table 4 may specify cross effects in an imperfect manner by assuming that they can be represented by the means (and higher moments) of household distributions, although adding second- and third-order moments did not significantly improve the fit of these equations. A household fixed-effects estimator, however, provides an estimate of own endowment effects that requires no assumptions about the parameterization of the household variables, because the full set of cross terms and household-specific regressors are impounded in the fixed effect. 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, by gender, of the effects of personal characteristics on individual calorie consumption. The set of instruments for the own endowment measure remains the same. The estimation procedure takes into account the sample property that individuals appear more than once. The null hypothesis of no individual-specific error components is indeed rejected by the Breusch-Pagan (Lagrange multiplier) test statistic in each equation estimated.<sup>16</sup>

<sup>15</sup>The reduction in total sample size that occurs when the sample is stratified by sex results from the necessity of using only those households that have both males and females in order to estimate gender-specific cross and own endowment effects.

<sup>16</sup>Note that only household (and not individual) random effects were specified in estimating the calorie-allocation equations of Table 4. The Breusch-Pagan (Trevor S. Breusch and Adrian Pagan, 1980) statistics of Table 5 confirm the importance of individual effects even when controlling for household fixed effects. The parameter estimates of Table 4 are nonetheless consistent, but standard errors are underestimated by about 10 percent, based on our experience in obtaining the estimates reported in Table 5.

TABLE 5—FIXED-EFFECTS TWO-STAGE GENERALIZED LEAST SQUARES:  
EFFECTS OF PERSONAL CHARACTERISTICS ON  
INDIVIDUAL CALORIE CONSUMPTION

Variable <sup>a</sup>	Two-stage least-squares estimates <sup>b</sup>			
	Males		Females	
	Endowment effects constant	Endowment effects vary with age	Endowment effects constant	Endowment effects vary with age
Own endowment <sup>c</sup>	0.447 (3.58)	—	-0.0278 (0.15)	—
Age < 6 <sup>c</sup>	—	-0.435 (1.35)	—	-0.314 (0.46)
6 ≤ age < 12 <sup>c</sup>	—	0.923 (2.29)	—	1.86 (2.13)
Age ≥ 12 <sup>c</sup>	—	1.21 (2.69)	—	0.0894 (0.13)
Age	1.44 (22.9)	1.31 (14.9)	1.34 (18.1)	1.35 (17.9)
Age squared	-0.201 (16.7)	-0.170 (9.16)	-0.199 (13.4)	-0.206 (13.7)
<i>N</i>	429	429	371	371
<i>X</i> <sup>2</sup> (no individual error components)	46.5	48.35	32.36	26.17
Individual error variance/total error variance	0.287	0.300	0.258	0.282

<sup>a</sup>All variables in logs.

<sup>b</sup>Asymptotic *t* ratios in parentheses.

<sup>c</sup>Instrumental variables used are means of individual and family endowments for weight/height, skinfold thickness, and arm circumference calculated over all survey rounds, excluding the round from which observation is drawn.

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 is only  $-0.028$  ( $t = -0.15$ ) for females.

In columns 2 and 4 of Table 5, we report estimates obtained using specifications of the calorie equation in which sex-specific own endowment effects are allowed to vary across the three age groups that appear from Table 2 to be related to the differentiation of activity patterns. 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, in this case solely via the allocation of foods.

Male and female children aged 6–12 years evidently engage in a more diverse set of activities ranked by energy intensity, and the own endowment parameters exhibit reinforcement and are statistically significant for both males and females. A 10-percent increase in the health endowment of a 6–12-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 activities, categorized by energy intensity, displayed 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 and are concentrated in less energy-intensive activities. Reflecting this, the estimated own endowment elasticity of calorie consumption for adult males is positive, statistically significant, and the largest of all groups (1.21), while that for adult females is close to zero (0.09).

If  $\beta_k$  is the estimated own endowment effect for age-sex group  $k$ , then the variability in consumption in group  $k$  depends on  $\beta_k$  and on the group's dispersion in endowments, as  $\text{Var}(c^k) = \beta_k^2 \text{Var}(\mu^k)$ . Based on our estimates of the endowments, we cannot reject the hypothesis that all within-group endowment variances are equal. Our estimates of  $\beta_k$  values thus imply that household allocation rules and the variation in the effects of health on productivity across activities are in part responsible for the higher variability in intrahousehold calorie allocations among males relative to females for individuals aged 12 and over. These factors also contribute 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 than 6.<sup>17</sup>

### C. *Endowments, Family Income, and Activity Participation: Household Discrimination*

Although 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 individuals with higher endowments are more likely to choose activities with a greater

energy intensity of effort, as implied by the theory. Moreover, with estimates of endowment effects on energy intensity, we can use (11) to compute how the household on net responds to differences in endowments for each gender. While the Bangladesh data do not provide information on individual-specific wage rates or earnings, we can test whether households with higher average health endowments among adult males have higher incomes, given land resources and schooling. Table 6 (column 1) provides estimates of the determinants of (log) per capita income. These 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 (column 2) also reports maximum-likelihood (ML) instrumental-variable estimates of a probit activity-choice equation for individuals aged 12–60.<sup>18</sup> The dichotomous dependent variable in this equation has the value of 1 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) and 0 otherwise. 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. The coefficient on sex (male = 1) is positive and statistically significant, reflecting the differences between sexes in the diversity of occupations, given endowments. Thus, the results reported in Table 6 confirm that there is a pecuniary return to health and effort, that adult males with

<sup>17</sup>To test for 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.

<sup>18</sup>The likelihood maximized is given in Richard J. Smith and Richard W. Blundell (1986).

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	Per capita income (two-stage least-squares) <sup>a</sup>	Exceptionally active occupation (full-information ML IV probit) <sup>a</sup>
Own endowment <sup>b</sup>	—	13.9 (1.64)
Family endowment males ≥ 12 years old <sup>b</sup>	2.38 (2.86) <sup>b</sup>	−16.74 (2.29)
Family endowment females ≥ 12 years old <sup>b</sup>	0.378 (0.75)	−3.67 (1.25)
Age	—	1.28 (1.06)
Sex	—	6.92 (2.36)
Landholding	0.0200 (0.64)	−0.0219 (2.46)
Household head's schooling	0.109 (1.80)	−1.09 (1.54)
Mean age of family members	0.0444 (0.14)	−1.58 (1.18)
Variance of ages of family members	0.591 (1.91)	−5.55 (2.50)
Proportion of family members male	0.566 (1.09)	−4.11 (1.82)
Jorbaria village	−0.199 (1.30)	−5.98 (2.19)
Constant	4.23 (4.11)	4.95 (1.17)
<i>N</i>	45	153
<i>F</i>	3.73	—
$\chi^2_{[12]}$	—	76.2

<sup>a</sup>Asymptotic *t* ratios in parentheses.

<sup>b</sup>Instrumented.

higher endowments are more likely to undertake exceptionally energy-intensive work, and that adult female health endowments are relatively unimportant in determining activity choices or household income compared to adult male endowments.

Finally, the net effect of a change in own endowment on own health [eq. (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. 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.<sup>19</sup>

<sup>19</sup>The elasticity is given by

$$d \ln h / d \ln \mu = 1 + (\partial \ln h / \partial \ln c)(d \ln c / d \ln \mu) \\ + (\partial \ln h / \partial e)(de / d \ln \mu).$$

The health elasticity of calories, the first parenthetical term in the elasticity expression, is 0.136 for both males and females (Table 1). The elasticity of calorie consumption with respect to own health endowment ( $d \ln c / d \ln \mu$ ) is 1.21 for adult males and 0.089 for

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 vs. 3 percent).

#### IV. 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 individual heterogeneity in exogenous healthiness and differences in labor-market returns to health and effort across groups of individuals. The empirical analysis was applied to individual and household-level data from Bangladesh, a country that exhibits large differences in calorie consumption and in the energy intensity of activity by age and gender. Our results 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 6–12) 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 adult women in Bangladesh, appear to account for part of the disparities in the average consumption of nutrients across adult men and women and to contribute to the greater variability among men in nutrient consumption.

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 likelihood of adult males with higher endowments to undertake exceptionally energy-intensive work resulted in a "tax" on adult male endowments that exceeded that of adult females (12 percent vs. 3 percent), signaling some discrimination against males by the household.

Our evidence that disparities by gender in food consumption in a low-income society like Bangladesh reflect the gender differentiation in the energy intensities of activities 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 an increase in the occupational diversity of women is likely to reduce (calorie) consumption inequality between the sexes, it will increase inequality among 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. However, to the extent that economic development is characterized by a transformation of work activities to those in which linkages between food consumption and productivity are weak, overall inequality in food consumption may be attenuated for all groups as incomes rise.

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adult females (Table 5). The effect of participation in an exceptionally active occupation on (log) health ( $\partial \ln h / \partial e$ ) is  $-0.082$  for both males and females (Table 3). The estimated effect of (log) own health endowment on the probability of participating in an exceptionally active occupation is 13.9 (Table 6) multiplied by the standard normal density function evaluated at the value of the underlying latent activity variables for adult males and females, which are 0.247 and 0.038, respectively.

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