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Reviewed work(s):

Source: *Journal of Political Economy*, Vol. 108, No. 4 (August 2000), pp. 688-727

Published by: [The University of Chicago Press](#)

Stable URL: <http://www.jstor.org/stable/10.1086/316098>

Accessed: 15/05/2012 04:40

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## **In Sickness and in Health: Risk Sharing within Households in Rural Ethiopia**

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Much of the literature on consumption smoothing and on risk sharing has focused on the ability of the household as a unit to protect its consumption. Little is known about the ability of individual members of the household to keep consumption smooth over time or relative to other members of the household. We use data on adult nutrition in Ethiopia to investigate whether individuals are able to smooth their consumption over time and within the household. We find that poorer households are not able to do so. Furthermore, poor southern households do not engage in complete risk sharing; women in these households bear the brunt of adverse shocks. This result implies that the collective model of household organization, which imposes Pareto efficiency on allocations, is rejected for these households. Finally, we obtain estimates of the relative Pareto weights in household allocation. We find that a wife's relative position is better if customary laws on settlements at divorce are favorable or if she comes from a relatively wealthy background and that poor southern women have lower Pareto weights in allocation.

This research uses data collected by the Department of Economics of Addis Ababa University and the Centre for the Study of African Economies, Oxford University. Data collection and research were financed by the Swedish International Development Cooperation Agency, the Economic and Social Research Council (United Kingdom), and the Fund for Scientific Research (Belgium). We would like to thank Harold Alderman, Bart Capéau, Stephen Nickell, Daphne Nicolitsas, Francis Teal, Dirk Van de gaer, seminar participants at Namur, Leuven, and Oxford, and a referee for their helpful comments. All errors remain our own.

[*Journal of Political Economy*, 2000, vol. 108, no. 4]  
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## I. Introduction

Households in rural Ethiopia, like farm households elsewhere in developing countries, live in a highly risky and volatile environment and suffer tremendous variation in incomes, mainly because of factors beyond their control. Keeping consumption smooth in the face of such fluctuations is difficult, and imperfect markets for credit and insurance make such smoothing doubly difficult for the poor. Much of the literature on saving and consumption smoothing has focused on the ability of the household as a unit to protect its consumption. Little is known about the ability of individual members of the household to keep consumption smooth over time or relative to other members of the household. We use panel data from rural Ethiopia, on individual nutritional status, to test whether individuals can smooth nutritional levels over time. Furthermore, we investigate whether households are able to share risk so that nutritional levels are smooth across members of the households. In doing so, we look at the factors determining the intrahousehold allocation of nutrition. We ask whether, in the face of fluctuating incomes, the burden of adjustment is borne by women, particularly in poorer households.

In examining this, we draw on several threads in the literature. The first refers to the ability of households to smooth consumption over a short horizon. There is evidence that households are able to do so, but the presence of liquidity constraints for some households qualifies these findings for it does limit the ability of such households to smooth their consumption effectively over time (e.g., Paxson 1992; Townsend 1994). The second thread is the ability of households to protect their consumption by relying on friends and the extended family, through village-level networks or government schemes, and thus to share the risk of volatile income with others. This is the cross-sectional counterpart to consumption smoothing across time. If risk is fully shared through market or nonmarket institutions, household consumption should not respond to idiosyncratic shocks in income. Research on the United States suggests that there is little evidence that such risk sharing takes place either at the national level or within extended families (Cochrane 1991; Altonji, Hayashi, and Kotlikoff 1992; Hayashi, Altonji, and Kotlikoff 1996). The evidence for developing countries favors at least partial risk sharing among households (Morduch 1991; Townsend 1993, 1994; Udry 1994). All these tests use data on the household to examine its behavior in the aggregate but clearly can be extended to examine risk sharing within the household.

There is substantial evidence that in some parts of the world, boys and men are favored over girls and women (see the review by Strauss

and Beegle [1995]). Most of these results are based on differences in levels of health outcomes or individual consumption of nutrients. While differences in average outcomes between the sexes do matter, both the variation in individual consumption and health over time and the differences in such variation between household members must be of interest. Behrman and Deolalikar (1990), using data on individual nutrient intakes from India, report that estimated price and wage elasticities of intakes are substantially and significantly higher for females than for males, suggesting that women and girls share a disproportionate burden of rising food prices. In a similar vein, there is some evidence that the consequences of shocks for men and women depend on whether the household is poor.<sup>1</sup>

Yet another thread is the extension of risk-sharing models to the examination of risk sharing within the household. Such an extension is a natural approach to examining intrahousehold allocation. In maximizing welfare for the household as a whole, household members will have to decide who gets what share of the total. The head of the household might take on the role of a social planner and allocate Pareto weights to each member, or some process of bargaining might take place that takes into account relative wealth (e.g., the amount of wealth brought in at marriage by each of the partners) and thus maximize the appropriately weighted sum of individual utilities. As Hayashi (1987) and Townsend (1994) point out, a natural test of altruism when household members are able to pool risk among themselves might be to test whether the weights are related to individual wealth. If they are not but the household is able to share risk, one might infer that household members are altruistic in the allocation of resources. Risk sharing within the household might be said to take place when individual shocks do not affect individual outcomes, when the household's resources are controlled for: the shock is shared according to some sharing rule. In tests of risk sharing, the weights would crop up as individual fixed effects. It might also be possible to find the underlying sharing rule by testing whether the weights are a function of individual wealth or endowment variables.

There has been considerable discussion of whether the characteristics of individuals or their sources of income determine their share in consumption (Chiappori 1992; Bourguignon et al. 1993; Browning et al. 1994). Chiappori develops models consistent both with the familiar dictatorial model of household allocation and with cooperative bargaining and suggests ways of discriminating between them. In essence, testing

<sup>1</sup> Rose (1994) finds that in rural India negative rainfall shocks are associated with higher mortality rates for boys and girls in landless households, but not in households with lots of land. However, Foster (1995) does not find any evidence of a sex bias in the evolution of child growth during and after the severe floods in Bangladesh in 1988.

the dictatorial model involves testing whether incomes are pooled among members. However, this is a test of a static model of resource allocation. A dynamic model of allocation of resources requires that even if incomes are not pooled, the shocks to income are "pooled," if risk sharing does take place within the household. In sum, the alternatives to the traditional model are consistent with the absence of income pooling, but if they are to remain Pareto-efficient allocations, these alternatives must allow for the pooling of shocks to income.

In this paper, we use anthropometric indicators to investigate these issues. Such indicators have substantial advantages. They are relatively easy to collect; are less prone to error than consumption, nutrient, or income data; and are data at the individual level rather than aggregated to the level of the household. We use such data to test whether individuals smooth consumption and explicitly test for differences between males and females. The focus on adults, however, complicates the problem since the effects of increased consumption on productivity and incomes must be explicitly accounted for. In fact, if resources are scarce and if returns to health vary by sex and age, we would expect households to allocate more health inputs to those members for whom the marginal product of health on income or wages is higher. This is the pure "life-boat" problem: poor households, which are liquidity constrained, might, in the face of a shock to their incomes, be forced to allocate limited resources toward those members who are more productive or more likely to survive.

The paper is organized as follows. In Section II, we present the survey evidence on adult nutritional status in rural Ethiopia. Section III develops a model of intertemporal resource allocation that explicitly accounts for the role of consumption in determining health, the influence of health on productivity, and the fact that health (or good health) is analogous to a durable good. In Section IV, we present the empirical specification and in Section V a discussion of the data, followed by a discussion of the results in Section VI. Section VII examines the results on the intrahousehold allocation and the determinants of the Pareto weights. Section VIII presents conclusions.

## **II. Adult Nutrition Using the Quetelet Index in Ethiopia**

The Quetelet or body mass index is now recognized as a reliable measure of the current nutritional status of adults (James, Ferro-Luzzi, and Waterlow 1988). It is measured as weight in kilograms divided by squared

height in meters.<sup>2</sup> It was proposed as a criterion for classifying chronic energy deficiency in adults and as supplementary to an indirect assessment using nutrient intake. Even though there is still substantial debate on the cutoff points for identifying the degree of malnutrition, extremely low values of the Quetelet index (below 17 or 18) have been associated with higher adult mortality (Waler 1984). However, the link between illness and the Quetelet index is less clear. For instance, Adams (1995) finds no significant correlation between the Quetelet index and the duration and incidence of adult morbidity across seasons. The index is also a recognized measure of the amount of energy stored in the body and is liable to some fluctuation over a short horizon.<sup>3</sup> It is likely, therefore, that the index is affected by changes in prices and income (Strauss and Thomas 1995*a*).

There is increasing evidence of a relationship between the Quetelet index and productivity, especially at lower levels of the index.<sup>4</sup> In this paper, we use the Quetelet index and its link to general health and productivity to examine the well-being of individuals over time. We use data from a panel of households and individuals from rural Ethiopia. The sample consists of 1,477 households in 15 areas of the country, with extensive health and socioeconomic data on over 9,000 adults and children, interviewed three times between 1994 and 1995. Details of the survey can be found in Dercon and Krishnan (1997).

Table 1 presents mean levels of the Quetelet index in rural Ethiopia for individuals aged 20 and over. Data are based on 2,343 individuals for whom complete information was available, excluding lactating and pregnant women. About a quarter of the adult population is malnourished, with little difference between men and women. About 7 percent of men and 9 percent of women display a Quetelet index below 17. Mean levels of the index are between 19 and 20, which is lower than the usual national averages in poor countries of between 21 and 23,

<sup>2</sup> The measure and its appealing characteristics were first described by the Belgian statistician, economist, and demographer Lambert Adolphe Jacques Quetelet (1796–1874) in his *Sur l'homme et le développement de ses facultés, ou Essai de physique sociale* (1835).

<sup>3</sup> It ought to be stressed that there is little information on morbidity risk of low levels of the Quetelet index in developing countries (Satyanarayana et al. 1991), but there appears to be general agreement that large fluctuations may pose unacceptable risks to health and have functional consequences.

<sup>4</sup> For an excellent discussion, see Dasgupta (1993, pp. 432–36). The Quetelet index has been found to be related to physical functioning or the ability to perform certain tasks in Indonesia (Strauss and Thomas 1995*a*) and to perform strenuous work in Bangladesh and Rwanda (Pitt, Rosenzweig, and Hassan 1990; Bhargava 1997). It is positively related to earnings in Bangladesh (Pitt et al. 1990) and to the productivity of women in Brazil (Strauss and Thomas 1995*a*).

TABLE 1  
ADULT NUTRITIONAL STATUS IN RURAL ETHIOPIA, 1994–95

| Frequencies                         | Men  | Women |
|-------------------------------------|------|-------|
| Normal                              | 75.5 | 75.0  |
| Grade I chronic energy deficiency   | 17.5 | 16.0  |
| Grade II chronic energy deficiency  | 4.5  | 5.6   |
| Grade III chronic energy deficiency | 2.5  | 3.4   |
| Mean value of Quetelet index        | 19.9 | 19.5  |

SOURCE.—Pooled data from rounds 1–3 of the Ethiopian Rural Household Survey (ERHS), 1994–95. Adult population (above 20 years of age). There were 2,343 observations in each round. Grade I chronic energy deficiency refers to a Quetelet index between 17 and 18.4; grade II, between 16 and 17; grade III, less than 16; and normal, greater than 18.5 (see James et al. 1988). Sample excludes pregnant and lactating women.

although not particularly low in comparison to other rural areas in developing countries.<sup>5</sup>

We are interested in the fluctuations of the Quetelet index over a relatively short time horizon. The lowest level of the index as a percentage of the highest level is, on average, 91.6 percent for men and 90.7 percent for women.<sup>6</sup> These measures are somewhat lower than for other countries for which comparable data could be found, suggesting very large fluctuations over time in the Quetelet index, with women being adversely affected. Since these results imply a weight loss of, on average, 8–9 percent, this does not look like consumption smoothing.

The observed scale of adult malnutrition is likely to have functional consequences and to affect labor productivity. Data collected on physical functioning do show clear correlations between malnutrition and the ability to perform standard tasks. Table 2 shows the relationship between measured energy deficiency and an adult's ability to hoe a field for a morning. The correlations are striking. A third of those measured as grade II or grade III malnourished have some difficulty hoeing a field, compared to less than 15 percent of normally nourished adults. Almost a quarter of the adults who are grade III malnourished cannot hoe a field. Other activities such as carrying water for a short distance displayed similar correlations.

These productivity effects suggest that there are substantial costs to large fluctuations in nutritional status beyond the direct effects on well-

<sup>5</sup> Strauss and Thomas (1995*a*) report data on Brazil, Indonesia, and Côte d'Ivoire. Vosti and Witcover (1993) report means of the Quetelet index for a particular area in south central Ethiopia that are even lower than our measures: 18.4 for males and 19.1 for females. Gillespie and McNeill (1992) report means for a village in South India of 19.1 for both men and women. In the United Kingdom, mean values are close to 24, and in the United States the mean value is about 25.

<sup>6</sup> Data exclude pregnant and lactating women. Ferro-Luzzi et al. (1990) report an index of 96.5 percent for women in a relatively rich area in Ethiopia, Gillespie and McNeill (1992) report 97.2 and 98.8 percent for men and women in South India, and Dugdale and Payne (1987) report 93.4 percent for men in rural Burma.

TABLE 2  
FREQUENCY DISTRIBUTION OF FUNCTIONING BY LEVEL OF MALNUTRITION

|                                       | LEVEL OF MALNUTRITION |         |          |           |
|---------------------------------------|-----------------------|---------|----------|-----------|
|                                       | Normal                | Grade I | Grade II | Grade III |
| Can person hoe a field for a morning? |                       |         |          |           |
| Easily                                | 85.6                  | 76.9    | 62.5     | 62.6      |
| With a little difficulty              | 4.4                   | 6.8     | 8.3      | 8.2       |
| With a lot of difficulty              | 3.4                   | 5.3     | 14.8     | 6.8       |
| Not at all                            | 6.6                   | 10.9    | 14.4     | 22.4      |

SOURCE.—Rounds 2 and 3 of the ERHS. Questions were phrased with relevance to local units and necessarily rephrased to suit division of labor by sex.

being. Since labor requirements and returns fluctuate over time, these productivity effects create incentives to reduce food intakes in some seasons to be able to boost nutrition in peak periods. Similarly, they create incentives to skew allocation to the more productive. The high fluctuations observed in table 2 might well be compatible with optimizing behavior by the households.

Table 3 presents the average Quetelet index and the average of the minimum Quetelet index as a percentage of its maximum level, for men and women, by consumption per capita in quartiles, landholdings per capita in terciles, and geographical location. Average Quetelet index increases, whereas its variability falls with consumption levels per capita. Landholdings per capita show a similar pattern: more variability and a lower mean Quetelet index for households with smaller holdings. Finally, mean levels of the Quetelet index are lowest and its variability the highest in the South. In the central areas of the country and to a lesser extent in the North, men and women have similar levels of nutrition and suffer similar variability. In the South, the gap in the Quetelet index between men and women is larger and the fluctuations are much larger for women.

We investigate why we observe such large fluctuations and whether this is consistent with consumption smoothing at the individual, household, and village levels. We explore whether the higher fluctuation in the Quetelet index of the poor is a reflection of liquidity constraints and ask why women in the South suffer larger fluctuations. Finally, we look for an explanation for the apparent differences between men and women in the average allocations. Is this just a reflection of productivity differences or individual health endowments, or a consequence of a bias against women? We examine whether a bargaining view of intra-household allocation is consistent with the observed differences. In the next section, we describe our formal approach to these questions.

TABLE 3  
 QUETELET INDEX AND CONSUMPTION, LAND, AND GEOGRAPHICAL LOCATION

|                                     | AVERAGE QUETELET INDEX |       | MEAN OF MINIMUM AS PERCENTAGE OF MAXIMUM |       |
|-------------------------------------|------------------------|-------|--|-------|
|                                     | Men                    | Women | Men                                      | Women |
| By consumption per capita quartile: |                        |       |  |       |
| Poorest                             | 19.5                   | 19.0  | 91.1                                     | 90.3  |
| Lower middle                        | 19.8                   | 19.4  | 91.4                                     | 90.7  |
| Higher middle                       | 20.0                   | 19.8  | 91.5                                     | 90.6  |
| Least poor                          | 20.2                   | 19.9  | 92.2                                     | 91.3  |
| By landholdings per capita tercile: |                        |       |  |       |
| Lowest                              | 19.6                   | 19.2  | 91.5                                     | 90.3  |
| Middle                              | 19.9                   | 19.8  | 91.5                                     | 90.7  |
| Highest                             | 20.1                   | 19.7  | 91.7                                     | 91.3  |
| By region:                          |                        |       |  |       |
| North                               | 20.2                   | 19.8  | 91.3                                     | 90.7  |
| Central                             | 19.8                   | 19.8  | 92.2                                     | 92.1  |
| South                               | 19.6                   | 19.0  | 91.3                                     | 89.6  |

SOURCE.—Consumption from the first round of ERHS. Mean consumption per capita is 72 birr per month (about \$12). Landholdings are land suitable for cultivation, and its mean is about 0.3 hectare per capita.

### III. The Model

We begin by assuming that the household maximizes a “common preference” instantaneous utility function, postponing the discussion of alternative assumptions about preferences for later in this section. The household is assumed to derive utility from the nutritional status  $N$  of each member  $j$  of the household at each moment  $t$  in time. Nutrition is obtained from the consumption of food or nutrients  $C$  in the same period. Utility also depends on consumption expenditure on nonfood items, denoted by  $O$ , which is assumed to be a public good within the household and to be weakly separable from allocations to  $N$ . Let  $\beta$  be the discount rate ( $\beta < 1$ ) and let the household consist of  $J$  members. When uncertainty is ignored for the time being, the household maximizes the following intertemporally additive utility function at each point in time  $\tau$ :

$$\sum_{t=\tau}^T \beta^{t-\tau} \cdot U(N_1(C_{1t}), N_2(C_{2t}), \dots, N_j(C_{jt}), O_t), \quad (1)$$

in which  $U_N > 0$ ,  $U_{NN} < 0$ ,  $U_O > 0$ , and  $U_{OO} < 0$ .

Households maximize this utility function subject to a standard intertemporal full-income budget constraint, linking asset levels  $A$  in  $t + 1$  with assets, income  $Y$ , and total consumption of nutrients and other items in period  $t$ . There is no bequest motive and no debts in the final period  $T$ . Let  $L_{jt}$  be the total available labor of household member  $j$  at time  $t$ . Using nonfood prices as the numeraire and denoting  $r$  as the

return on asset holdings between periods and nutrient prices as  $P$ , we can write this equation as

$$A_{t+1} = (1 + r_{t+1}) \cdot \left( A_t - P_t \cdot \sum_{j=1}^J C_{jt} - O_t \right) + \sum_{j=1}^J Y_{jt+1}(L_{jt}). \quad (2)$$

For convenience we refer to these assets as financial assets, although in the Ethiopian context they could also be livestock or other commodities. Note that income  $Y$  earned at  $t + 1$  is a function of labor supply  $L$  in the previous period, as is usual in agricultural production. The income earned by each member is a function of individual labor supply and is pooled by the household. Each member's income depends on own productivity, on the division of labor across members, and on general labor market conditions. In the model we assume that leisure is exogenous; however, labor supply becomes endogenous through its link with nutrition and illness, hence affecting productivity. In particular, we assume that labor supply is the sum of total work time available  $T$  minus the time lost because of illness  $Z$ . Time available for work is made dependent on nutrition, as in the usual literature on nutrition-productivity links: the better-nourished people can endure longer hours and perform tasks in less time (Dasgupta 1993; Strauss and Thomas 1995*a*), and illness is negatively affected by nutritional status:

$$L_{jt} = T_j(N_{jt}) - Z_{jt}(N_{jt}), \quad (3)$$

with  $T_N > 0$ ,  $T_{NN} < 0$ ,  $Z_N < 0$ , and  $Z_{NN} > 0$ . The links between income, labor supply, and nutrition result in some efficient productivity level of nutrition intake.

Nutritional status must be considered as a stock or durable ("a store of energy"). The next constraint is the structural equation determining the stock of nutritional status in each period. Past levels of the stock of nutritional status will determine present levels, taking into account "depreciation": even at rest, the body uses energy. The function  $f$  in equation (4) below gives this depreciation function for each individual. In each period, nutritional status can be boosted by consumption ( $C$ ), the transformation of which is determined by the function  $m$ , which is increasing in  $C$ . Morbidity in each period ( $Z$ ) is likely to negatively affect this transformation, especially with diseases related to the functioning of the gastrointestinal tract. Finally, working, or expending energy on income-generating activities, will reduce nutritional status, and this loss needs to be taken into account as a function  $a$ , which is increasing in  $L$ . It is assumed that  $f_N$  is between zero and one: some but not total depreciation takes place, but consumption always implies some improvement in nutritional levels:

$$N_{jt} = f^j(N_{jt-1}) + m^j(C_{jt}, Z_{jt}) - a^j(L_{jt}), \quad (4)$$

with  $0 < f_N < 1$ ,  $m_c > 0$ ,  $m_z < 0$ , and  $m_{Lz} < 0$ .

This equation is important for specifying the demand for nutrients in each period since it means that profitable arbitrage can take place over time, depending on the time path of prices for nutrients. In particular, the body can function as a store of wealth, as an asset alternative to financial assets. While financial assets have a well-defined opportunity cost in terms of the given interest rate  $r$ , for nutrition this price is not as easily defined and is in any case individual-specific. Since nutritional status is a durable, we shall have to define the opportunity cost of boosting nutritional status in one period relative to another period. This cost of increasing nutritional well-being today relative to tomorrow can be obtained by defining the "user cost" or the "rental price" associated with increasing the nutritional status in one period only, leaving it unaffected in other periods (Grossman 1972; Deaton and Muellbauer 1980). Equation (5) defines this user cost  $\Pi$  of increasing nutritional status of individual  $j$  in period  $t$  using equations (3) and (4):

$$\begin{aligned} \Pi_{jt} = & P_t \cdot \frac{1 + r_{t+1}}{m_{c_t}^j} - \frac{f_{N_t}^j}{m_{c_{t+1}}^j} \cdot P_{t+1} - \frac{\partial Y_{t+1}}{\partial L_{jt}} \cdot (T_{N_{jt}} - Z_{N_{jt}}) \\ & + a_{L_{jt}}^j \cdot (T_{N_{jt}} - Z_{N_{jt}}) \cdot P_t \cdot \frac{1 + r_{t+1}}{m_{c_t}^j}. \end{aligned} \quad (5)$$

The first term in (5) defines the cost of taking sufficient resources from the future and transforming it into nutritional outcomes via the consumption transformation function. However, since part of this stock increase persists into the future, the actual user cost in period  $t$  is lower by the second term, in which the value of what will be carried over to the future is given, when depreciation is called for. Furthermore, increased nutritional status now yields increased returns since more nutrition relaxes the time constraint (3).

For our purposes, this user cost is crucial because it may vary considerably across individuals. While consumption prices are constant across individuals within the same family or even village, individual characteristics could determine the efficiency of transforming consumption into nutrition and the rate of depreciation. Even more important, individual marginal returns from increasing nutrition may vary considerably, especially at low levels of nutrition, affecting nutritional allocation across individuals. To see this, consider the optimal allocation of nutritional status between two individuals  $i$  and  $j$  of the same household from optimizing (1) subject to the constraints

$$\frac{U_{N_{it}}}{U_{N_{jt}}} = \frac{\Pi_{it}}{\Pi_{jt}}. \quad (6)$$

Equation (6) states that in each period households will allocate consumption such that the marginal rate of substitution of nutritional status between members  $i$  and  $j$  is equal to the ratio of marginal prices of increasing nutritional status for each individual. The actual allocation among members of the household will depend on both preferences and the price of nutrition. Even if the same weight is attached to the utility of each member, the allocation within the household will not be equal if there are differences in the productivity of different members.<sup>7</sup> Finally, (6) also implies that relative nutritional status will not necessarily remain the same over time if there are fluctuations or cycles in marginal productivities of different household members. Even if the household has some aversion to inequality and even if, as in (1), the household prefers smooth nutrition to fluctuations for its members, the household may choose to switch resources to particular members in particular periods.

We can make this more explicit by writing the instantaneous utility function in (1) as in Samuelson's (1956) consensus model, mirroring a social welfare function with Pareto weights:

$$\sum_{j=1}^J \theta_j \cdot U_j(N_{jt}, O_t), \quad (7)$$

in which  $\theta$  is the Pareto weight given to individual  $j$ , which sum up to one for all  $j$ . With such preferences, the optimal allocation between any members  $i$  and  $j$  of the same household (eq. [6]) can be written as

$$\frac{U_{N_{it}}}{U_{N_{jt}}} = \frac{\theta_j}{\theta_i} \cdot \frac{\Pi_{it}}{\Pi_{jt}}. \quad (8)$$

This formulation can be used to investigate the nature of the relationship between members of the same household since (7) describes what Chiappori (1988, 1992) refers to as the "collective" model, which nests in its specification both a common preference model as in (1) and a model in which individuals with well-defined utility functions interact in a collective decision-making process to achieve a Pareto-efficient allocation. The outcome can always be written as though a social welfare function is maximized subject to a pooled resource. In general, the weights could depend on prices, incomes, preferences, and other var-

<sup>7</sup> Formally, if indifference curves are linear, so that the sum of utilities is maximized irrespective of the distribution, substantial inequality may arise because of these price differences. Only if indifference curves are L-shaped will price differences not matter. Behrman and Deolalikar (1990, 1993) refer to this as the degree of inequality aversion within the household, with L-shaped indifference curves implying maximum concern.

variables, including variables that do not affect the total budget constraint or individual preferences but still affect the decision process or arise from cooperative bargaining within the household.

Typically, as in McElroy and Horney (1981), a Nash bargaining model is assumed, and threat points associated with the default outcome of divorce are supposed to affect the equilibrium outcome between spouses. The weights in the allocation within the household will then depend on variables describing this outside option. Lundberg and Pollak (1993) have argued that the relevant threat points for the Nash bargaining solution are those implied by an "uncooperative marriage" in which spouses may revert to a traditional sexual division of labor or, as Bergstrom (1996) puts it, "burnt toast and harsh words."<sup>8</sup>

Chiappori (1988, 1992) and Browning et al. (1994) have derived a series of testable restrictions from the general collective model (i.e., without imposing restrictions on the type of underlying decision process), which can be used on cross-section data (for a review of the empirical evidence, see Lundberg and Pollak [1996]).<sup>9</sup> Most of the tests on the collective model are derived from static and deterministic models. In such models, no distinction needs to be made between persistent and transient changes in variables such as income. Nevertheless, whether someone's increase in current income is persistent or not is likely to change the resulting relative allocation considerably. In developing countries, with substantial risks in incomes, not making the distinction between permanent changes in incomes and transitory shocks may blur the analysis. In fact, the implications of the collective model go even further: when Pareto efficiency is imposed on all allocations, this requires all gains from trade between the individuals to be exploited, for otherwise one individual could obtain a higher utility without affecting the utility of the other individuals. Pareto efficiency requires all individual-specific shocks in income (or in other variables) to be insured by the other members of the households. *Collective models are consistent with the absence of income pooling but require income shock pooling: they view*

<sup>8</sup> An alternative solution to Nash bargaining is to use the Kalai-Smorodinsky solution, which is responsive to the best possible outcome for each member within the marriage (Dasgupta 1993). In most empirical applications, if a bargaining model is appealed to, the results tend to be interpreted as reflecting the Nash solution, but it is not obvious that this is the best interpretation.

<sup>9</sup> Despite the impressive number of testable restrictions that can be obtained from this literature, the actual tests are not without their problems. Heterogeneity in productivity is one example. Behrman (1993) suggests that differences in labor and unearned incomes for, say, females in cross-section data may be closely linked to productivity differences: more productive women not only have higher incomes from past and present labor, but also have a positive impact on their children's health. The apparent rejection of income pooling is then just a rejection of a too restrictive common preference model, in which heterogeneities are not appropriately modeled. Equation (8) also shows that such productivity and "collective weights" effects may not be easily identified in a cross section.

the household as a risk-sharing institution. This also suggests that a more appropriate test of the collective model may be a test of endowment pooling rather than of income pooling.

To see this formally, we derive the perfect risk-sharing solution (Cochrane 1991; Townsend 1994; Udry 1994). Assume that uncertainty is summarized by a state variable  $s$ , which can take on  $S$  possible values with probability  $p^s$  in each period. Uncertainty affects individual income outcomes. Assume further that we can write expected intertemporal utility of the household by combining (1) and (7) as

$$\sum_{j=1}^J \sum_{t=\tau}^T \beta^{t-\tau} \cdot \sum_{s=1}^S \theta_j \cdot p^s U_j(N_{jt}^s, O_t^s). \quad (9)$$

The household can then be viewed as though a social planner has given each individual  $j$  a weight  $\theta_j$  according to which consumption can be allocated in each period and in each state of the world. Assume further that assets are accumulated only by the household.<sup>10</sup> Uncertainty is assumed to be limited to individual incomes. Under these circumstances, the aggregate endowment within the household is still defined as in (2), albeit with income realizations being state-dependent. In each state and in each period a household-level multiplier can be defined related to equation (2) that is the same for each member of the household. The Pareto-optimal allocation of nutrition across individuals for each set of states of the world  $s$  and  $s'$  and at each  $t$  will be defined as

$$\frac{U_{N_{it}^s} / \Pi_{it}^s}{U_{N_{jt}^s} / \Pi_{jt}^s} = \frac{\theta_j}{\theta_i} = \frac{U_{N_{it}^{s'}} / \Pi_{it}^{s'}}{U_{N_{jt}^{s'}} / \Pi_{jt}^{s'}}. \quad (10)$$

Equation (10) shows that the optimal Pareto-efficient allocation will be independent of the state of the world and will be the same in each period: full insurance is obtained.<sup>11</sup> This outcome can be thought of as being achieved through a series of state-contingent transfers in each state of the world; transactions agreed between the individuals ex ante, with a veil being drawn over the precise way in which this might be done.

The household appears to be an ideal place for full risk sharing to take place, and indeed many have argued that if risk sharing is ever likely to exist, it is among members of the same family in the form of

<sup>10</sup> This assumption could be relaxed to allow for assets to be kept at the level of the individual, if one assumes that they cannot be traded outside the household. If they can be used within the household, then individual positions in each period will be dependent not just on current income outcomes but also on past income realizations and savings.

<sup>11</sup> From (5) it can be seen that under uncertainty user costs may vary across states and time as well. It is assumed that once the state of the world is known, the value of the user cost at  $t$  is uniquely defined. In a full-information equilibrium within the households, the values of these virtual prices in each state of the world are known by each member.

an informal insurance arrangement (Hayashi et al. 1996). In fact, risk-sharing opportunities provide powerful incentives for entering into marriage contracts: the problems caused by asymmetric information are likely to be limited, and it is likely that members care about each other, which will reduce incentives to cease participation in agreements and renege on commitments.<sup>12</sup>

The disadvantage, of course, is that incomes within households are likely to covary. Rural villages may be suitable settings for further risk sharing to take place: information flows relatively freely, enforcement may be easily obtained, and incomes are not perfectly covariate. Full risk sharing can then be defined as condition (10) being valid for all households in the village. Even beyond the village some risk sharing may be expected between family members, and evidence has been found to support this (Lucas and Stark 1985; Ravallion and Dearden 1988; Rosenzweig and Stark 1989). Nevertheless, full risk sharing requires high information flows and the absence of enforcement problems, as well as limited covariance in shocks among the members. Substantial risks are unlikely to be insurable in these villages. Credit or quasi credit may take on the role of insurance, but it suffers from many of the same problems.

Hence, for much risk, self-insurance—through the accumulation and depletion of assets—may be the only alternative for the household. Let us assume that assets are pooled as in equation (2) and that risk is shared within the household. Assume further that transfers or gifts may be given by other households but that they are received by the household and not by individual members. This implies that in each period each member of the household faces the same household budget constraint, which means that resources are allocated according to preferences, weights, and user costs, as in (2). Finally, assume that credit is unobtainable, so that an additional constraint,  $A_t \geq 0$ , has to be added. The Euler equation describing the optimal path over time for the nutritional status of each individual  $j$  in the household can be defined in the usual

<sup>12</sup> Enforcement problems after each realization of the state of the world imply, however, that the derivation of the Pareto weights determining each individual's outcome from a one-period game-theoretic model is not possible. Nevertheless, such arrangements can come about from repeated interaction, and several contributions have determined or characterized constrained Pareto-efficient equilibrium informal insurance contracts using a repeated game structure (Coate and Ravallion 1993; Thomas and Worrall 1994). These arrangements deviate from full risk sharing and therefore from Pareto efficiency. Rejecting full risk sharing does not exclude the existence of some risk sharing within the household; it does reject, however, the assumption of Pareto efficiency in the collective approach to modeling the household. In empirical tests, full risk sharing within (extended) families has been rejected in the United States by Hayashi et al. (1996). Pareto efficiency itself has been tested in the context of cross-section data using tests derived from the collective model. Bourguignon et al. (1993), Browning et al. (1994), and Thomas and Chen (1994) do not find evidence to reject Pareto efficiency. However, Udry (1996) rejects Pareto efficiency in production within households in Burkina Faso.

way (as in Deaton [1991]; for details, see Dercon and Krishnan [1997]). If households are unconstrained, the standard result of intertemporal allocation is obtained, in which marginal utilities are equated across time, appropriately discounted, and possible price changes over time are controlled for. However, if households face binding liquidity constraints in period  $t$ , the household will consume all assets. Individual marginal utility will then be equal to the transformation into nutrition of the amount of total household resources given to the individual. The presence of liquidity constraints may then cause substantial fluctuations in the nutrition of households and their members.<sup>13</sup>

However, the tendency to link the absence of smooth consumption over time with the existence of liquidity constraints may be misleading. If food prices are systematically lower in certain seasons, then fluctuations in consumption and therefore in nutritional status may be optimal. Furthermore, different returns to labor in peak and slack seasons are likely on top of fluctuations in prices. If depreciation is limited, if returns to other liquid assets are low, and if food prices show very substantial fluctuations, then it may pay the household to use the body as a store of wealth: feast when prices are low and fast when prices are high. Peak labor seasons usually coincide with high food prices since substantial labor is needed in periods before the next harvest. While this will make boosting consumption in slack seasons more costly, with very high food prices in peak labor seasons, it may make consumption intakes even higher in slack seasons, resulting in even more, but optimal, variability in household consumption over the seasons.<sup>14</sup> The optimality of this strategy is a sign of the imperfections of the asset and food markets: if anything, it shows that the returns to assets and the return to stocks in body weight are not integrated and arbitrage is profitable. Given the aversion to fluctuations in nutrition by individuals, welfare improvements could be obtained by asset and food markets that function better.

#### IV. Empirical Model and Econometric Specification

Our empirical approach focuses on the behavior of nutritional status for individuals across households and over time. Risk sharing requires that allocations are unaffected by individual-specific shocks, except for their effect on the household budget constraint. If we find evidence

<sup>13</sup> There is extensive evidence on the use of savings to smooth consumption in developing countries, even if less than perfect smoothing results (Paxson 1992, 1993; Rosenzweig and Wolpin 1993; Fafchamps, Udry, and Czukas 1998). Evidence also suggests that poorer households are less able to smooth consumption, suggesting differential liquidity constraints between poorer and richer households (Morduch 1991; Townsend 1994).

<sup>14</sup> A model with storage losses gives similar results. Dugdale and Payne (1987) discuss this using data from Gambia and Burma. Fluctuations in consumption intakes and nutritional outcomes over the seasons are optimal in the circumstances.

against it, we have also found evidence against a “collective” specification of the household allocation process. We also try to identify whether households smooth consumption either through risk sharing with other households or through self-insurance at the household level. To investigate this, we look at shocks that do not affect household permanent income and investigate whether these shocks appear to be insured by the households or within the community. We make a distinction between poorer and richer households to proxy for liquidity constraints. We use the analysis to make a direct link to intrahousehold allocation and investigate the nature of the arrangements within the household by obtaining the determinants of the “sharing rule.” Focusing specifically on married or cohabiting couples, we shall proceed in the next section to investigate which factors determine the respective weight in the allocations of husbands and wives: assets, altruism, or user cost-related efficiencies.

Suppose that the utility function of each member of a household is additive separable in nonfood consumption over time. Assume for simplicity that the risk aversion coefficient is the same within a household (but not necessarily so between households). Let  $N_{jt}$  denote the nutritional status or the Quetelet index of person  $j$  in household  $J$  at time  $t$ . Suppose that marginal utility can be defined as  $N^{-\rho}$ , with  $\rho$  the coefficient of relative risk aversion. An alternative (structural) way of writing the first-order conditions of our model is to equate the weighted marginal utilities for each individual to the value of the Lagrange multiplier on the life cycle asset constraint of household  $J$  at  $t$ ,  $\lambda_{jt}$ , which is constant for members of the same household. If risk is shared within the household and equation (10) is used, the optimal solution for each household member in each period can be written as

$$\frac{\theta_j \cdot U'(N_{jt})}{\Pi_{jt}} = \lambda_{jt} = \frac{\theta_j \cdot N_{jt}^{-\rho}}{\Pi_{jt}}. \quad (11)$$

First-order conditions linking periods  $t$  and  $t - 1$  can then be obtained by dividing a first-order condition in period  $t$  by the condition in period  $t - 1$ , so that the Pareto weights drop out. Rewriting the expression further by taking logs, and adding an error term  $\epsilon_{jt}$  (assumed to be identically distributed), we obtain

$$\ln \frac{N_{jt}}{N_{j,t-1}} = \frac{1}{\rho} \ln \frac{\Pi_{j,t-1}}{\Pi_{jt}} + \frac{1}{\rho} \ln \frac{\lambda_{j,t-1}}{\lambda_{jt}} + \epsilon_{jt}. \quad (12)$$

If households do not face credit constraints, then the shadow price of the household budget constraint will be equalized over time, up to the rate of time preference and the interest rate. If interest rates and time preference are equal and if user costs remain constant over time, then

we obtain the familiar martingale property of the logarithm of durable stocks under permanent income (Mankiw 1982). Shocks to income would affect the Lagrangean multipliers only to the extent that they shift the life cycle asset constraint. However, if credit constraints bind, then the intertemporal relationship between the multipliers breaks down and the relevant household-level budget constraint is determined by current asset levels.<sup>15</sup> Shocks to household income would have a larger effect on the evolution of nutritional status over time. To the extent that we can distinguish liquidity-constrained households from others, this observation lies at the basis of our test of consumption smoothing across households. To implement the tests, income shocks are identified. Village-level shocks are treated differently than household-level shocks, for the former cannot be insured within the village. As in Udry (1994), to measure shocks to income we focus on the causes of income shocks rather than on the actual changes in income.

A test for risk sharing within the household also requires us to identify idiosyncratic shocks to income earned by individuals within the household. When the effect on total household life cycle wealth or, for constrained households, on current wealth levels, under full risk sharing is controlled for, individual shocks should have no effect on the path of nutritional status. It is difficult to identify such shocks in agricultural households, where, as in Ethiopia, tasks related to the agricultural cycle are carried out together, on a joint family farm.<sup>16</sup> Many of the risks faced by the individual members—such as risks related to farming—are unlikely to be insurable within the family since they will be perfectly covariate. Substantial time is likely to be spent on household work and in the production of other household goods, and idiosyncratic shocks to the returns to this effort may not be easily measurable.

To get around these problems, we assume that all time actually spent has an individual-specific return  $w_{j,t+1}$ , which is risky, but this risk is perfectly covariate across members of the same family  $j$ . As before, nutritional levels will affect the actual number of hours or days of work

<sup>15</sup> Equation (12) is invalid since an additional multiplier  $\mu_{jt}$  needs to be added, related to the nonnegativity constraint on assets. Therefore, the first-order condition can be written as

$$\frac{\theta_j \cdot U(N_{jt})}{\pi_{jt}} = \lambda_{jt} + \mu_{jt} = \frac{\theta_j \cdot N_{jt}^{\rho}}{\pi_{jt}}.$$

Equation (12) can be amended with the prediction that changes in the shadow price of credit constraints would affect the path on nutritional status. If income changes are not persistent, then the effect of changes in current asset positions can produce relatively large effects on  $\mu_{jt}$  compared to  $\lambda_{jt}$ .

<sup>16</sup> If household members each owned separate plots, as in parts of West Africa (see, e.g., Udry 1996), then one might be able to derive direct measures of idiosyncratic income shocks.

that the individual member can commit to earning this return. Idiosyncratic income shocks will be defined as being related to time lost because of illness shocks. Formally, we define

$$Y_{jt} = \varphi_{jt} \cdot w_{jt} \cdot L_{jt-1}, \quad (13)$$

where  $w_j$  is the return to labor time for individual  $j$ ;  $\varphi_{jt}$  is the multiplicative risk to returns to labor, which is the same for all members  $j$  of household  $j$ ; and  $E_{t-1}(\varphi_{jt}) = 1$ . Using equation (3) and defining  $Z_{jt-1}$  as the individual-specific time lost because of illness, we can define idiosyncratic income shocks as

$$\frac{\partial Y_{jt}}{\partial Z_{jt-1}} = -\varphi_{jt} \cdot w_{jt}. \quad (14)$$

Working time lost because of individual illness shocks will reduce household earnings by a multiple of (14). However, if the household operates as a full risk-sharing institution, then this should not affect the allocation of nutrition to each household member. Consequently, terms measuring individual illness shocks in the period preceding the measurement of nutritional status ought not to be significant, when the effect on household-level earnings in equation (13) is controlled for.

However, this test lacks power to distinguish risk sharing: it is correct only if illness shocks operate solely through the budget constraint. First, illness may not only affect the ability to earn income but could also have a direct effect on nutritional status. For example, illnesses such as gastrointestinal ailments could result in direct weight loss from an inability to absorb nutrients.<sup>17</sup> Second, illness may be endogenous to nutritional status. In (3), earnings at  $t + 1$  are affected by illness at  $t$ , which in turn are affected by nutritional status. As in (4), both problems will cause the user cost of nutritional status to change. To address the first problem, we use the details on the type or symptoms of the illness episodes reported to exclude those illnesses that are likely to cause direct weight loss. The second problem is all too common when health data are used and is discussed further in Section VI. Besides these two problems, we must acknowledge that the entire testing procedure is dependent on the assumption that illness does not affect the marginal utility of nutrition. In (1), illness itself was not an argument in the utility function; in general, separability of illness and nutrition would be required for the test to stand up. Otherwise, illness shocks would affect nutrition even under full insurance. Assuming separability is a strong assumption,

<sup>17</sup> Lusky et al. (1996) use data on 17-year-old Israeli males to examine the relationship between morbidity and extreme values of the Quetelet index. They find that functional disorders associated with underweight are bronchial and lung conditions, intestinal conditions, and emotional neuroses. They point out that underweight may be a consequence of some disorders rather than a cause.

which is not easily tested: however, in discussing the results, we examine whether they are likely to be caused by nonseparability rather than reflect risk sharing.<sup>18</sup>

To derive a testable equation from (12), we consider three factors determining the path of the user cost. The first two parts describe the path of prices: time-varying but constant for a community ( $P$ , e.g., prices, returns to assets) or for similar individuals ( $W$ , e.g., wages, relevant for  $j$  and similar for individuals with characteristics  $a_j$ ). We also assume that this time path is a function of lagged nutritional status. The final part includes individual-specific variables  $v_{jt}$  to capture differences in physiology. The time path of the user cost is defined additively as<sup>19</sup>

$$\ln \frac{\Pi_{jt-1}}{\Pi_{jt}} = \alpha \cdot \ln \frac{P_{t-1}}{P_t} + \omega \cdot \ln \frac{W_t^{a_j}}{W_{t+1}^{a_j}} + \eta \cdot \ln N_{jt} + \ln \frac{v_{jt-1}}{v_{jt}}. \quad (15)$$

Higher prices at  $t-1$  relative to  $t$  would increase user costs at  $t-1$  compared to  $t$ , whereas higher wages at  $t$  compared to  $t+1$  would reduce user costs at  $t-1$  compared to  $t$ . The final term depends on individual physiological characteristics.<sup>20</sup> We assume that this term (times  $1/\rho$ ) is a fixed effect  $\pi_j$ . From (15), (12) becomes

$$\begin{aligned} \ln \frac{N_{jt}}{N_{jt-1}} &= \pi_j + \frac{\eta}{\rho} \cdot \ln N_{jt-1} + \frac{\alpha}{\rho} \cdot \ln \frac{P_{t-1}}{P_t} + \frac{\omega}{\rho} \cdot \ln \frac{W_t^{a_j}}{W_{t+1}^{a_j}} \\ &\quad + \beta \cdot S_{ct} + \gamma \cdot S_{jt} + \delta \cdot Z_{jt-1} + \epsilon_{jt}, \end{aligned} \quad (16)$$

where  $\epsilon_{jt} \sim \text{ID}(0, \sigma_\epsilon^2)$ ;  $E[\epsilon_{jt}] = E[\epsilon_{js}] = 0$  for  $t \neq s$  (the residuals are not serially correlated); and  $S_{ct}$  and  $S_{jt}$  are insurable shocks affecting income occurring at the community and household levels, respectively,

<sup>18</sup> Another issue is nonseparability between illness and nutrition effects in the marginal return to nutrition. In that case, an illness shock would also affect the user cost of nutrition. This makes the risk-sharing test formulation not valid; i.e., the assumptions underlying (13) and (14) are too restrictive and the test invalid. In Sec. VI, when discussing the results, we shall briefly address this issue as well.

<sup>19</sup> We assume that lagged nutritional status enters the Euler equation on a purely physiological basis: the costs of adapting to new information prices or incomes, which affect the relative costs of boosting nutritional status, are a function of lagged nutritional status. To derive this from the optimization model, further simplifying assumptions will need to be made regarding eq. (4). In the literature on durables, lagged dependent variables have also been obtained by imposing utility costs of adjustment (Bernanke 1984) or other fixed costs of adjustment specifications (Bertola and Caballero 1990). It is unlikely that adjustment costs in utility or costs of a pure physiological adjustment can be distinguished in empirical work.

<sup>20</sup> We therefore assume that each individual has a specific path of nutritional status. Individual heterogeneity is a major finding in most health analysis and a potential source of serious misspecification (Behrman and Deolalikar 1990; Strauss and Thomas 1995a). Strauss and Thomas (1995b) comment on the problems related to ignoring heterogeneity in the Euler equations estimated by Foster (1995). Behrman, Deolalikar, and Lavy (1995) argue strongly for the need to control for heterogeneity in dynamic models of nutritional status.

in the period between  $t - 1$  and  $t$ . The coefficients  $\beta$  and  $\gamma$  are expected to be nonnegative. A test for consumption smoothing involves testing whether  $\beta$  and  $\gamma$  are equal to zero. Risk sharing within the household would imply that  $\delta$  in (16) is zero once the household-level effect of the shock (see eq. [14]) is controlled for.

Owing to the presence of a lagged endogenous variable,  $\ln N_{jt-1}$ , (16) cannot be consistently estimated using generalized least squares on the equation in first differences. To estimate (16) we use the optimal generalized method of moments (GMM) estimator derived by Arellano and Bond (1991) and estimated it in first differences. We use  $\ln N_{jt-2}$  and a set of exogenous and predetermined variables  $Z$ , such as family and parental background, and serious health problems in the past to instrument for the lagged endogenous variable. The set of instruments pertains to periods  $t - 3$  and earlier. The two-step Arellano and Bond estimator, which corrects for possible heteroscedasticity in  $\epsilon_{jt}$ , is used to obtain consistent estimates. We report the Sargan test for overidentifying restrictions to test the null hypothesis of the validity of the instruments.

Since data were collected on nutritional status for all members of the household, a further test for risk sharing can be implemented using (10). Consider two members of the same household,  $i$  and  $j$ , facing the same household budget constraint so that  $\lambda_{jt}$  in (12) is equal for both. Time-variant household-level variables and shocks are differenced out by taking ratios between nutritional status for  $j$  and  $i$  to yield equation (17). Note that  $u_{ijt}$  is an independently distributed random error:

$$\ln \frac{N_{jt}}{N_{it}} = \frac{1}{\rho} \ln \frac{\Pi_{it}}{\Pi_{jt}} + \frac{1}{\rho} \ln \frac{\theta_j}{\theta_i} + u_{ijt} \quad (17)$$

By first-differencing this equation, we can eliminate the relative Pareto weights. Using (15) and adding an overidentifying variable related to illness shocks at  $t - 1$ , we get

$$\begin{aligned} \ln \frac{N_{jt}}{N_{j,t-1}} - \ln \frac{N_{it}}{N_{i,t-1}} &= (\pi_j - \pi_i) + \eta \cdot (\ln N_{j,t-1} - \ln N_{i,t-1}) \\ &+ \frac{1}{\rho} \cdot \left( \ln \frac{W_t^{a_j}}{W_{t+1}^{a_j}} - \ln \frac{W_t^{a_i}}{W_{t+1}^{a_i}} \right) \\ &+ \delta_j \cdot Z_{j,t-1} - \delta_i \cdot Z_{i,t-1} + u_{ijt}, \end{aligned} \quad (18)$$

where  $\pi_i$  and  $\pi_j$  are defined as before. Significance of  $\delta_i$  or  $\delta_j$  would signify the absence of risk sharing. Since, by definition, this coefficient ought to be equal to those of the same variables in (16), this constitutes a test of robustness for the risk-sharing results.

Differencing equation (18) and using the GMM estimator as described

above obtains consistent estimates of the parameters but loses an important variable of interest: the ratio between the Pareto weights (“the sharing rule”), which determines the intrahousehold allocation of nutrition between  $i$  and  $j$ . We can retrieve these weights by using our regression results either on the intertemporal or on the intrahousehold allocation (eqq. [16] and [18]). Details on the derivation are discussed in the Appendix. The factors determining the relative weights of household members are investigated by running a regression of individual endowments on this fixed effect.

## V. The Empirical Specification

To estimate equations (16) and (18), we construct a series of variables measuring individual-, household-, and village-level shocks in incomes and prices. The variables used are summarized in table 4. First, we define variables reflecting the prices and wages relevant for the evolution of user costs of nutrition, as in (16). Since complete data on wages were not available, we construct proxies to describe prices and wages, and we assume that returns on assets  $r$  are relatively constant over the seasons.<sup>21</sup> In (16), since earnings from labor expended today are obtained only in the next period, higher future wages relative to given current levels increase current nutritional levels compared to past levels. In highly seasonal agriculture, peak labor periods are those in which good nutrition would have high (future) returns. Consequently, we expect that if a survey round coincided with a local peak labor period relative to the previous observation of the Quetelet index, it would have positive effects on the difference in the index between periods. We introduce sex-specific interaction terms since returns to labor may be different for men and women. Lower food prices in  $t$  relative to  $t + 1$  make the user cost of nutrition at  $t$  lower relative to  $t + 1$ . Part of the seasonal changes in food prices are predictable, with prices lower in postharvest periods. Consequently, we introduce dummies indicating whether the measurement of the Quetelet index was taken in the postharvest period, when prices were low. The postharvest period was defined as within four months of the start of the harvest (for details, see Dercon and Krishnan [1997]).

Rainfall data for each village, measured as rainfall in the cropping season related to the last harvest relevant for the survey period and expressed as a percentage of mean rainfall using historical data, are also introduced. Rainfall data for the nearest local weather station were

<sup>21</sup> An alternative specification using site-specific dummies was also tried, but it did not change the econometric results substantially. Since using proxies allows a clear interpretation of the results, only these findings are reported.

TABLE 4  
 VARIABLES USED IN THE REGRESSION OF CHANGES IN THE LOG OF THE QUETELET INDEX

| Variable                          | Definition   |
|-----------------------------------|--|
| Community-Level Variables         |  |
| Rainfall                          | Change in rainfall as a percentage of mean rainfall in periods relevant for harvest before $t$ and $t-1$   |
| Rainfall $\times$ road            | Change in rainfall multiplied by an indicator of whether the village has an all-weather road   |
| Postharvest                       | Difference in dummies each defined as one if measurement of the Quetelet index was taken in the postharvest period (defined as within four months of the start of harvest)   |
| Peak labor                        | Difference in dummies each defined as one if measurement of the Quetelet index was taken during a peak period for male labor, such as plowing and harvesting periods   |
| Household-Level Shocks            |  |
| Crop shock                        | Index of the absence of farm-level problems, such as plant diseases, flooding, insects, animal trampling, etc. The higher the better. Used as differences of indices   |
| Rain shock                        | Farm-specific index of rainfall experiences on the farm, such as whether plowing occurred too early or too late for the rains, whether it rained when harvesting, etc. The higher the better. Used as differences of indices |
| No oxen                           | Whether farmer could not obtain oxen at the right time for plowing. One if constrained. Used as a difference of dummies  |
| No labor                          | Whether labor could not be hired when needed on the farm. One if constrained. Used as a difference of dummies  |
| No off-farm                       | Whether no off-farm wage labor could be found when wanted in the period between observations. One if constrained   |
| Livestock shock                   | Farm-specific index of problems with livestock related to access to grazing land and to livestock diseases. The higher the index, the worse the outcome. Used as differences   |
| Livestock loss                    | Value of the loss of livestock due to death or missing in the period between $t-1$ and $t$   |
| Family Labor Supply               |  |
| Male/female adult illness days    | Number of days in the last 28 days preceding the survey that male or female adults in the households were too ill to work, as a percentage of male adults in the family. The higher, the worse                               |
| Number of dead male/female adults | Number of male or female adults who died between rounds $t$ and $t-1$  |

| Individual-Level Shocks        |   |
|--------------------------------|---|
| Idiosyncratic illness shock    | Unpredicted shock in working days lost because of illness in last 28 days preceding round $t$ . Residuals obtained from a fixed-effect regression using previous illness experience, lagged Quetelet index, and controls for pregnancy and breast feeding, individual and household characteristics, and time-varying village-level dummies |
| Time-Varying Control Variables |   |
| Breastfeeding/pregnancy        | One if woman is breast feeding or pregnant. Used as change in dummies   |
| Lagged log Quetelet index      | Predicted lagged dependent variable. Identifying instruments are family background variables, individual-specific health history variables, including the number of days of serious illness in the last five years as instruments, and lagged level of ln Quetelet index  |
| Interaction Terms              |   |
| Male/female                    | Sex-specific effects  |
| South                          | One if village is in the Southern Ethiopian People's Administrative Region, i.e., southern Ethiopia   |
| Lowland                        | One if household owns less than the median land per capita in the village of residence (i.e., half the households in each village will get the value one)   |

used to construct the series. Positive shocks in rainfall are likely to affect food prices negatively, proxying the unpredictable part of price movements. They should also be seen as a proxy for positive community-level income shocks. We interact rainfall with an indicator of whether the area has an all-weather road. In a country with a poor road infrastructure, the presence of a good road may have different effects on prices and incomes.

Next, household-level income shock variables were constructed as an index of self-reported adverse occurrences affecting crops and livestock production. Almost all households were involved in crop production, and many owned livestock. During each interview, data were collected on events in the past cropping season and the relevant harvest. The first index measures whether common problems such as pests, flooding, insects, and animal trampling of crops affected crops. The second measures the farm-specific experience related to rainfall during different parts of the season. Since oxen are very important in many farming systems in Ethiopia, a third index measures whether there was a problem in obtaining oxen in time for plowing. A fourth index measures problems in hiring labor. The last two indices are proxies for the degree of imperfection in oxen and agricultural labor markets and the effect on the farmer. We also constructed an index of problems associated with

livestock diseases and sufficient access to grazing land. To measure shocks, we use the first difference of each index in each period.<sup>22</sup> We also include the value of livestock that died or got lost in the previous period to proxy losses in assets.

We also include variables measuring illness shocks to household labor supply. We used the total number of days a person was not able to work because of illness, as reported by adults in the household and divided by the number of adults, as a proxy for the reduction in household income from illness. The measures were constructed separately for men and women. We also include the number of male and female adults who died between rounds.

To measure individual illness shocks in the period preceding measurement of the Quetelet index, we use working days lost because of illness in the previous month, as discussed before. There are several problems with using this variable.<sup>23</sup> Illness is assumed to work through the budget constraint and not via the user cost. One problem is that during illness the same food intake might result in a lower weight gain (if the body becomes less efficient in transforming food into nutrition). To handle this problem, we investigated the nature of the illness symptoms reported. Just under 30 percent of adults in the sample reported at least one episode of illness in any of the rounds used in this paper. The most common symptoms reported by those ill were severe headaches (20 percent of cases) and fever (19 percent). Symptoms such as diarrhea, vomiting, abdominal pains, and other symptoms related to gastrointestinal illnesses were reported in 15 percent of the cases. General weakness and fainting (9 percent) and respiratory problems (8 percent) are the next most important. Most of these reported episodes of illness do not appear to be directly related to weight loss with the clear exception of gastrointestinal diseases. Less than 7 percent of the sample was affected by these symptoms in one or more rounds of the survey. Since it is unlikely that this effect can be controlled for, we decided to exclude individuals suffering from symptoms of gastrointestinal disease. A second problem relates to the predictability of illness. A correct measure of an idiosyncratic shock in a risk-sharing test ought

<sup>22</sup> In a highly risky agricultural environment with imperfect markets, it would be incorrect to consider the absence of problems as a norm. During the piloting of the survey modules, it was found that farmers could not define problems relative to a norm or the usual outcome. Differences also have the advantage that we can allow shocks for those without livestock to be put at zero.

<sup>23</sup> One problem is the self-reporting bias often observed with illness in the poor, with more educated or richer people reporting more episodes of illness (Behrman and Deolalikar 1988). Using days unable to work reduces some of these problems (Schultz and Tansel 1997). In our data set, regressions on working days lost show predictable correlations with the loss of fewer days of work due to illness by educated and richer people than by poor and uneducated people, in contrast to regressions on reported illness days.

to exclude predictable episodes of illness. Days unable to work at  $t$  were predicted using the lagged Quetelet index in both  $t - 1$  and  $t - 2$ , individual and family background variables, and time-varying community-level dummies, controlling for health infrastructure and local seasonal factors. The panel allowed us to do a fixed-effects regression; therefore, we could control for health endowments and other fixed factors. In short, information available to the household and individual at  $t - 1$  was used to derive unpredicted illness as the deviation of actual days lost from days predicted.<sup>24</sup> This does not necessarily solve a third problem: that even unpredictable illness shocks are endogenous to the path of nutritional status. As was mentioned before, nutrition may affect the probability and length of illness and therefore affect the user cost of nutrition, as in (3) and (4). Simply using instruments to control for endogeneity is not possible: if the instruments are information available to the household and the individual before the illness strikes, then this information could have been used to predict illness anyway. However, unpredicted illness shocks, which are measured before  $t$  but after  $t - 1$ , have been purged of the effects of the Quetelet index at  $t - 1$  via the prediction model of illness shocks. Plots of unpredicted illness shocks against the lagged Quetelet index reveal no discernible pattern. While in itself this is no proof, it seems that endogeneity is unlikely to be at the core of the results. In fact, the pattern of the results obtained below further strengthens this argument.

All the household- and community-level variables describing income shocks were interacted with whether the household was in the lower half of the per capita distribution of land in each village. Land is an important asset in rural Ethiopia, closely correlated with other measures of wealth. However, by law it cannot be sold or bought, so it is not a variable that the household can control. The hypothesis is that landholdings may be a good proxy for liquidity constraints. The distribution of land per village is used to allow for the fact that the returns to land are dependent on local agroclimatic circumstances.

To test whether there is risk sharing within families, the individual illness shocks were interacted with sex. Table 3 suggested that fluctuations were not just larger for women, but especially so in the South and in poorer households. Hence, the idiosyncratic shock variables are interacted with a dummy for southern villages and for individuals in households with small landholdings. Other control variables include changes in the breast feeding or pregnancy status of women.

Equation (16) was estimated for all adults aged 20 and over for whom we have complete information in all rounds. We excluded those suf-

<sup>24</sup> A similar procedure was used to construct unpredicted illness shocks at the household level, but this did not affect the results.

fering gastrointestinal-related illnesses in any round of the survey and restricted the sample to those households consisting of at least one male and one female to avoid drawing conclusions on intrahousehold allocations in households with adults of only one sex. This left 1,787 individuals. Equation (16) was also estimated on a smaller sample of heads and spouses, leaving 816 individuals.<sup>25</sup> This sample was also used to estimate (18), providing a consistency test as well as a means of retrieving the intrahousehold weights between husbands and wives. The advantage of restricting the sample to couples is that issues related to marriage arrangements can be explored.

## VI. Estimation Results

Table 5 gives the results of the estimation of the intertemporal path of nutritional status. The first set of results pertains to the entire sample, and the second set pertains to only married or cohabiting couples. The dependent variable is the difference in the logarithm of the Quetelet index at  $t$  and  $t - 1$ . The model was estimated using naive instrumental variable estimation and proxies for the lagged value of the Quetelet index (in first differences, to remove the fixed effects) and also using the GMM estimator developed by Arellano and Bond (1991). The GMM estimator produces more efficient estimates, but the point estimates do not differ significantly between the naive instrumental variable estimator and the two-step GMM estimator. Only the results for the GMM estimates are reported here.

The regression results on both the full sample and the married couples show jointly significant estimates, and the Sargan test on the validity of the instruments used for the lagged dependent variable is not rejected. Pregnancy has its expected effect on weight gain, whereas breast feeding is not significant. The lagged dependent variable is highly significant, rejecting a pure random walk specification.<sup>26</sup> Its coefficient

<sup>25</sup> Polygamous households were excluded, but polygamy is not very common so that few observations are lost through this restriction. No detailed information on the nature of polygamous relationships was collected, so little would be gained from including them, except for possible complications in the interpretation of the findings.

<sup>26</sup> Measurement error in the Quetelet index could pose a problem in obtaining consistent estimates. However, it is likely that measurement errors are uncorrelated over time: scales were randomly distributed across and within villages and taken back to the capital, Addis Ababa, at the end of each round of data collection. Measurement errors are also likely to be household-specific since each household member was weighed at the same site, with the same scale. If so, this allows a test of the robustness of coefficient estimates by comparing the estimates of eq. (18) (on differences across couples, where identical measurement errors are netted out) with those of eq. (16), for individuals. The coefficients on the lagged Quetelet index in (16) are  $-.87$  compared with  $-.80$  in eq. (18). In short, if measurement error were a problem, we would expect the coefficient on the lagged Quetelet index in eq. (16) to be biased toward zero compared to the estimates obtained by differencing between husbands and wives. Since there is little difference in the coefficients of both the lagged Quetelet index and the individual-specific shocks, it might be concluded that household-specific measurement error is unlikely to be an issue.

TABLE 5  
GMM ESTIMATION OF INTERTEMPORAL PATH  
Dependent Variable: Change in Logarithm of Quetelet Index at  $t$  and  $t - 1$

| VARIABLES                                   | FULL SAMPLE (1,787 Adults) |            | COUPLES ONLY (816 Adults) |            |
|---|----------------------------|------------|---------------------------|------------|
|   | Coefficient                | tValue     | Coefficient               | tValue     |
| Constant                                    | -.0066                     | -1.81*     | -.0046                    | -.95       |
| Log Quetelet index, $t - 1$                 | -.8925                     | -14.51***  | -.8720                    | -9.53***   |
| Breast feeding, $t, t - 1$                  | .0004                      | .08        | .0080                     | 1.09       |
| Pregnancy, $t, t - 1$                       | .0187                      | 3.11***    | .0233                     | 3.05**     |
| Community-Level Effects                     |                            |            |                           |            |
| Peak male, $t, t - 1$                       | .0039                      | 1.70*      | .0057                     | 1.79*      |
| Peak female, $t, t - 1$                     | .0050                      | 2.10**     | .0060                     | 1.75*      |
| Highland postharvest, $t, t - 1$            | .0015                      | .69        | .0049                     | 1.34       |
| Lowland postharvest, $t, t - 1$             | .0049                      | 2.45***    | .0081                     | 2.48**     |
| Highland rainfall, $t, t - 1$               | -.0010                     | -.09       | -.0186                    | -1.01      |
| Lowland rainfall, $t, t - 1$                | -.0075                     | -.65       | -.0164                    | -1.02      |
| Highland rainfall $\times$ road, $t, t - 1$ | .0235                      | 1.28       | .0307                     | 1.20       |
| Lowland rainfall $\times$ road, $t, t - 1$  | .0227                      | 1.19       | .0263                     | .86        |
| Household-Level Effects                     |                            |            |                           |            |
| Highland rain shock, $t, t - 1$             | -.0012                     | -.14       | .0020                     | .17        |
| Lowland rain shock, $t, t - 1$              | .0089                      | 2.01**     | .0082                     | 1.26       |
| Highland crop shock, $t, t - 1$             | .0239                      | 2.23**     | .0100                     | .61        |
| Lowland crop shock, $t, t - 1$              | -.0235                     | -2.04**    | -.0237                    | -1.26      |
| Highland no oxen, $t, t - 1$                | -.0015                     | -.28       | .0091                     | 1.07       |
| Lowland no oxen, $t, t - 1$                 | -.0104                     | -1.76*     | -.0110                    | -1.14      |
| Highland no labor, $t, t - 1$               | .0119                      | 1.69*      | .0090                     | .64        |
| Lowland no labor, $t, t - 1$                | .0020                      | .49        | .0011                     | .22        |
| Highland livestock shock, $t, t - 1$        | -.0095                     | -1.55      | -.0052                    | -.55       |
| Lowland livestock shock, $t, t - 1$         | -.0057                     | -1.00      | -.0026                    | -.25       |
| Highland livestock loss, $t, t - 1$         | -.0039                     | -.71       | -.0209                    | -.91       |
| Lowland livestock loss, $t, t - 1$          | -.0123                     | -.23       | -.0143                    | -.74       |
| Highland no off-farm, $t, t - 1$            | -.0027                     | -.25       | -.0043                    | -.77       |
| Lowland no off-farm, $t, t - 1$             | -.0008                     | -.14       | .0026                     | .51        |
| Labor Supply Shocks                         |                            |            |                           |            |
| Highland male adults died, $t, t - 1$       | -.0162                     | -.84       |                           |            |
| Lowland male adults died, $t, t - 1$        | -.0964                     | -1.19      | -.0313                    | -.26       |
| Highland female adults died, $t, t - 1$     | .0155                      | .70        |                           |            |
| Lowland female adults died, $t, t - 1$      | .0034                      | .22        |                           |            |
| Highland male adults ill days, $t, t - 1$   | -.0008                     | -1.00      | -.0013                    | -1.47      |
| Lowland male adults ill days, $t, t - 1$    | -.0005                     | -.58       | -.0005                    | -.43       |
| Highland female adults ill days, $t, t - 1$ | -.0010                     | -.99       | -.0003                    | -.27       |
| Lowland female adults ill days, $t, t - 1$  | .0012                      | 1.56       | .0011                     | .77        |
| Idiosyncratic Individual Shocks             |                            |            |                           |            |
| Highland, South, male ill, $t, t - 1$       | -.0010                     | -.95       | -.0003                    | -.23       |
| Lowland, South, male ill, $t, t - 1$        | .0001                      | .04        | -.0018                    | -.56       |
| Highland, South, female ill, $t, t - 1$     | -.0022                     | -1.17      | -.0003                    | -.23       |
| Lowland, South, female ill, $t, t - 1$      | -.0042                     | -5.90***   | -.0030                    | -3.34***   |
| Highland, not South, male ill, $t, t - 1$   | .0013                      | 1.17       | -.0002                    | -.17       |
| Lowland, not South, male ill, $t, t - 1$    | -.0016                     | -1.35      | -.0013                    | -.69       |
| Highland, not South, female ill, $t, t - 1$ | .0004                      | .32        | .0013                     | .77        |
| Lowland, not South, female ill, $t, t - 1$  | -.0007                     | -.81       | -.0001                    | -.51       |
| Wald joint significance                     | 748.23                     | $p = .000$ | 404.60                    | $p = .000$ |
| Sargan test                                 | 2.05                       | $p = .842$ | 6.40                      | $p = .269$ |

\* Significant at the 10 percent level.  
\*\* Significant at the 5 percent level.  
\*\*\* Significant at the 1 percent level.

suggests a positive correlation over time in the path of nutritional status. If this reflects user costs in period  $t - 1$  relative to  $t$ , the negative coefficient in the regression suggests that user costs are higher at  $t$  for a higher level of the Quetelet index obtained at  $t - 1$ . This is consistent with expectations: boosting nutritional status will be more costly at higher levels of the Quetelet index.<sup>27</sup>

The change in the dummy describing whether the current period is a peak period for male labor needs has its expected positive and significant effect on nutritional status. The effect is valid for both males and females, suggesting that peak periods are not relevant just for men. Households appear, therefore, to be adjusting their nutrition to cash in on higher returns to labor in particular periods. Seasonal differences in productivity are responsible in part for observed fluctuations. The change in dummies related to the postharvest period also has the expected sign, suggesting that lower food prices in the postharvest period will induce increases in nutritional status, as part of arbitrage in user costs of the seasons. However, the effect is significant only for households with small landholdings: "feast now, fast later" is a strategy used by poorer households. Since the ability to store energy in the body is not linked to wealth, this effect suggests that access to good food storage facilities or to savings instruments may be costly.<sup>28</sup> Rainfall shocks appear to have little effect, although this may be a reflection of the fact that rainfall was good between 1993 and 1995 in most areas. Rainfall in areas with an all-weather road appears to be more favorable to nutritional status than rainfall in areas with a poor infrastructure.

Although not all shocks at the household level have the expected sign (in particular the index on crop shocks for poor households), we find significant (and sensible) effects on the farm-specific rain variable and on the availability of oxen for households with small landholdings. Since the farming system is entirely rain fed and in most cases is dependent on plowing, this indicates that farmers with little land cannot insure themselves or insure within the community. It does reject the presence of risk sharing within the village for poorer households. However, the evidence is far less clear in the regression focusing only on couples, where the presence of risk sharing or self-insurance within the village cannot be rejected, at least for households in which husbands and wives are present. Since the sample is much smaller in this case, it

<sup>27</sup> As was argued before, the formulation of the empirical model is also consistent with a model of utility costs of adjustment in a durable goods model, as in Bernanke (1984). The estimated coefficient on the lagged dependent variable is equivalent to a rate of stock adjustment to the desired level of .89 in Bernanke's model, suggesting low adjustment costs.

<sup>28</sup> The combined effects of postharvest and peak period variables could explain a more than 1 percent decline in nutritional status over about six months (i.e., between rounds of the survey).

is not clear whether this result has much significance. Although the effects on male illness and deaths of males at the household level have the expected signs (i.e., they reduce nutritional status consistent with a loss of income), they are not significant.

In the results on the overidentifying restrictions to test full risk sharing within the household, one result stands out: we find a very significant negative effect for women in poor southern households in both regressions. In all the other households (i.e., all the households in the northern and central areas and rich southern households) the effects are insignificant, suggesting that risk is being fully shared within the household. Southern households with little land do not share the risk of illness in women, and illness in men does not affect their nutritional status beyond an effect via household income. The very high significance of the effect is striking and is consistent with the findings in the descriptive statistics: higher fluctuations in the Quetelet index for women in southern households with little land. The coefficients suggest that an extra day of unexpected illness for these women would reduce the Quetelet index in the next period by 0.3–0.42 percent. Since the average number of days ill and unable to work is about 5.5 days per month for women reporting illness, this suggests a loss of 1.6–2.3 percent of body mass index due to the lack of risk sharing when ill for women in poor households in the South. The result also implies the absence of Pareto efficiency in the intrahousehold allocation. A priori, this does not exclude the possibility of constrained Pareto efficiency, resulting in partial risk sharing.<sup>29</sup>

Equation (18) was also estimated to provide a robustness test for the coefficients obtained. If they are not robust, then this may well suggest misspecification of the household- and community-level variables in the intertemporal regression. The robustness regression was highly significant and the instruments found valid (the Sargan test could not be rejected with  $p = .877$ ). The coefficients significant previously, the control variables, the lagged dependent variable, and the individual shock variables were again significant with similar values. The coefficient on the lagged Quetelet index was now slightly higher at  $-0.7960$  and on individual illness shocks for women in the South in households with little land was somewhat lower at  $-0.0048$ . This provides a powerful test for the robustness of the result that poor southern households do not fully share risk.<sup>30</sup>

<sup>29</sup> Since the result is asymmetric (men appear to be insured, but not women), it would reflect a set of sustainability constraints in which women fully insure men, but not vice versa.

<sup>30</sup> The coefficient for men in poor southern households was also significantly negative in the robustness test, suggesting that, contrary to the intertemporal results, they are also not fully insured. The main result, that poor southern households do not act as a risk-sharing institution, is reinforced by this finding.

The fact that the results in table 5 show significant effects only for members of poor households in the South suggests that nonseparability is unlikely to be the main cause of these results. Indeed, if nonseparable preferences were to drive the results, then it would have been unlikely that there would have been a systematic difference in the effects for poor households in a particular area, compared to other households—as though poor households would have a different preference structure related to illness and nutrition. Furthermore, a similar reasoning suggests that endogeneity—that is, that nutrition shocks and health shocks are simultaneously determined—is not driving our results. This explanation does not appear consistent with a different effect for poor households in the South compared to poor households in the North.<sup>31</sup> A related problem, nonseparability of the effects of illness and nutrition in the returns to labor time, is also unlikely to be behind the results since it would require important differences in the production technology between rich and poor households in each area, for which there is no evidence.<sup>32</sup>

## VII. Nutritional Allocation within the Marriage

The regressions allowed us to calculate the Pareto weights in the allocation of nutrition using the approach outlined in Section V. The estimates are an increasing function of the ratio of husband's weight in the allocation of nutrition relative to the wife's weight: they are obtained as the log of the relative weights multiplied by the discount rate common to the household. Further regression analysis on these weights allows us to explore the factors determining the nutritional allocation rules within the marriage. Using the information available within the data set, we try to test whether threat points outside the marriage influence allocations within the household. We also look for evidence that systematic productivity differences rather than bargaining between partners determine allocations within a marriage.

<sup>31</sup> In their test of household risk sharing against illness shocks, Gertler and Gruber (1997) make similar observations.

<sup>32</sup> There are differences in both the farming system and the household production technology between the North and the South. The North is cereal-oriented, with a main diet based on teff (an indigenous cereal crop) and barley, whereas the South has more permanent crops, with ensete (false banana) a main source of calories. In both parts of the country, agricultural and household work is definitely strenuous and time-consuming; it is hard to argue that the effect of illness shocks on the marginal return to nutrition and therefore on the user cost of nutrition would be fundamentally different between these areas. Also, there is no evidence of systematic technology differences in home and farm production, which could lead us to expect differential effects of illness on the user cost for some households and not for others. Lack of detailed time use data to determine shadow returns to different activities for different individuals does not allow us to explore this issue further.

Before we test these propositions, a brief digression on the institution of marriage by ethnic group and region is warranted, for it is quite different than in other parts of Africa. The main ethnic groups in our sample are the Amhara-Tigre (who inhabit the northern highlands), the Oromo (from the central areas), and four groups in the South: the Kembata, the Gurage, the Wolayata, and the Gedeo. Among the Amhara-Tigre (and, to a great extent, the Oromo), the basic unit of organization is the nuclear family. The first marriage is usually arranged by the parents, and once agreed, the two families make contributions of cattle and other goods to the new household. The gifts exchanged between the families are carefully balanced, and marriages usually occur between families of equal standing. A formal marriage contract is drawn up, and property division at divorce is equitable.<sup>33</sup> Most studies of the Amhara-Tigre stress that the wife is very much her husband's equal within the household.<sup>34</sup> The customs among the Oromo are not dissimilar, particularly where there has been close association with the Amhara. However, among the Muslim Oromo, customs do deviate from those of the mainly Orthodox Christian Amhara. Divorce is less common, and husbands have the advantage in initiating divorce. Divorce settlements are often less generous toward the wife, and the custom of bride-price is more prevalent at marriage.

The customs of the southern groups in our sample deviate from the northern norms and vary considerably by ethnic group and area. There are pockets of strong evangelical Christianity in some areas, in contrast to a mixture of animist and Muslim beliefs elsewhere in the region. However, there appear to be some similarities in social norms in marriage and at divorce. The strong patrilineal links seem "to make the marital relationship a discordant one" (Kaplan et al. 1971, p. 147). Women cannot sue for divorce, but men may do so at whim.<sup>35</sup> Divorce is rare, and wives get no share of assets on divorce.

This anthropological evidence suggests strong differences between

<sup>33</sup> Levine (1965) provides a fascinating discussion: "Within the nuclear family itself, the ties that bind are limited by certain fundamental safeguards of private property... The land which each partner brings to the marriage or later acquires through bequests remains his own until death. If they should separate, each takes his own land, and the goods acquired by the household during their marriage are divided equally between them. There is also the provision of 'gil,' a quantity of property that either or both of the spouses may designate, prior to getting married, as strictly private property and indivisible in the case of divorce. In the event that a nuclear family enters a savings association or equb, husband and wife join as separate members" (p. 258).

<sup>34</sup> "Transactions appear to be constructed to facilitate divorce and the redivision of wealth between husband and wife. Theoretically, they are in an equal position" (Pankhurst 1992, p. 109).

<sup>35</sup> Kaplan et al. (1971) go further to explain that "a man may strike his wife at the least provocation; he may ignore her and devote his entire attention to a second wife; he can send her home to her parents and still demand that she be faithful sexually" (p. 147).

TABLE 6  
VARIABLES USED FOR DETERMINING INTRAHOUSEHOLD ALLOCATION

| Variable             | Definition  |
|----------------------|---|
| Age gap              | Age difference between husband and wife (years)   |
| Age husband          | Age of the husband (years)  |
| Years residence gap  | Difference in the number of years' residence in the village by husband and wife                                 |
| Schooling difference | Difference in dummies measuring whether husband and wife had completed primary schooling                        |
| Duration marriage    | How long they have been married   |
| Husband richer       | Wife claiming that husband's household richer than hers at the time of marriage                                 |
| Husband same         | Wife claiming that husband's household of similar wealth to hers at the time of marriage                        |
| Joint goods          | Value in thousands of birr of goods given to the couple as wedding gifts to help them start up household        |
| Bride-price          | Value in thousands of birr of goods given by the husband to the bride's family at marriage                      |
| Divorce settlements  | One if wife stated that she could take assets with her if she was divorced                                      |
| Divorce rules        | Whether it is customary to share assets equally at time of divorce (one if yes; no if husband takes everything) |
| Land                 | The value of land owned by the household in hectares per capita   |
| South                | One if living in South  |
| South $\times$ land  | Land interacted with South  |
| Distance town        | Distance to nearest town in kilometers  |

different parts of the country on the rights to divorce and its consequences for women. These divorce rules result in a very low divorce rate in the South compared to the North. For instance, 10 percent of the northern households in the sample were headed by divorced or separated women, compared to less than 1 percent of households in the South. We also found that divorced or separated female heads of households in the North were not worse off in terms of nutritional status: their average Quetelet index across rounds was 19.86 compared to 19.82 for married women. This was in contrast to widowed female heads of households in the South, whose average Quetelet index was only 19.42.<sup>36</sup>

All this provides some support for a bargaining view with divorce as the threat point. To explore this, a regression was estimated on the relative weights. The variables used in the determinants of the intra-household allocation are defined in table 6. Variables included are characteristics of husband and wife, related to age, education, and length of residence. Other variables describe their marriage: how long they have been married, whether the wife thinks that she married someone from the same wealth background or a richer man, the value of the

<sup>36</sup> Since very few females were heads of households in the South, equivalent means could not be calculated.

bride-price paid, whether the wife can take valuables with her if she were to divorce, and the value of goods given to the couple at the time of their marriage. Using data collected in the anthropological survey, we also include a dummy describing the customary arrangement related to divorce: whether the assets are shared equally or given to the husband if the couple were to split up. Finally, we include a measure of the wealth of the household (land per capita) and the distance of the village to the nearest town.

Some of these variables can be interpreted as reflecting outside options on divorce, such as the rules governing divorce and the claims on the assets. A higher value of joint goods received at the time of marriage may improve the position of the wife, traditionally the weaker partner, since she can claim half of these assets. The wealth of the husband's family would similarly affect the husband's exit options and therefore the bargaining outcome; education gaps and levels could be interpreted in a similar way. Other variables, such as the duration of the marriage, the age gap between spouses, and the difference in the years of residence in the village, are not readily susceptible to an interpretation of being related to outside options. A spouse who is considerably older may be able to rely on a higher allocation sustained by social norms, and a longer marriage may sustain a more extreme form of "uncooperative marriage" as in Lundberg and Pollak (1993). Distance to the nearest town could affect the sustainability of social norms, but it could also affect the outside options available.

Total landholdings owned by the household are also not easily interpreted as reflecting relative bargaining power. Land claims may traditionally have been more biased toward males, but at present, landholdings are not privately owned and tend to be allocated by the peasant association. Traditional rules on the division of assets at divorce may be followed *de facto*, although *de jure*, land should be allocated by the peasant association using criteria such as household size. If divorce rules are followed for land as well, then we may expect effects in line with the North-South division as for other assets. To account for this, we interact landholdings with the dummy variable indicating a southern household. Data were available only on the first marriage, so we had to restrict ourselves to husbands and wives each in their first marriage, limiting the sample to 190 couples. Table 7 shows the results.

Despite the relatively small sample, some interesting results emerge. First, variables describing the relative outside options at divorce appear to matter. If customary rules demand an equal division of assets and higher joint gifts were given at the time of marriage, then the woman's position is improved, although this effect is significant only at 20 percent. Marrying a husband from the same or from a wealthier background is clearly disadvantageous for the woman. A large age gap between hus-

TABLE 7  
OLS ESTIMATION  
Dependent Variable: Household Fixed Effect in (18), i.e. the  
Predicted Log of the Ratio of the Pareto Weights of Husband over  
Wife, Times  $1/\rho$ , Using Regression on Couples (190 Couples)

| Variable                         | Coefficient | t-Value |
|----------------------------------|-------------|---------|
| Age gap                          | .005        | 1.55    |
| Age husband                      | -.002       | -.78    |
| Difference in years of residence | .001        | 1.12    |
| Schooling difference             | .016        | .66     |
| Duration of marriage             | .002        | .69     |
| Husband richer                   | .060        | 2.00**  |
| Husband same                     | .070        | 2.49**  |
| Value of joint gifts on marriage | -.002       | -1.30   |
| Bride-price                      | .001        | .65     |
| Divorce settlements expected     | -.001       | -.01    |
| Customary divorce rules          | -.037       | -1.54*  |
| Land                             | .017        | .61     |
| South                            | .014        | .24     |
| South $\times$ land              | -.255       | -1.71*  |
| Distance to town                 | .001        | .10     |
| Constant                         | -.037       | -.42    |
| Joint $F$ -test                  |             | 1.58*   |

\* Significant at the 10 percent level.

\*\* Significant at the 5 percent level.

band and wife worsens the wife's allocation. We also find that while landholdings are entirely insignificant in the North, they are significant in the South. There is a surprisingly large, positive effect on allocations to southern wives of increased landholdings. If landholdings matter as a proxy for individual wealth, then it should improve the man's position since the division of assets is strongly in favor of the husband in the South. An alternative interpretation is that it is related to productivity effects: if landholdings are small, the returns to additional labor are likely to be low. With nutrition-productivity links, this would also mean that the returns to additional nutrition would be low. Since in the South (but not in the North) the husband is the main decision maker, he may not find it in his (or the family's) interest to increase the family's labor supply by improving his wife's nutrition. With increasing landholdings, the marginal returns to labor would increase, and if they, as one would expect, increase over some range at a decreasing rate, then it is in the family's interest to increase the nutrition of the wife as well. A related interpretation, a more general "lifeboat" effect, is also consistent with the evidence: poor households have to skew their consumption to some of the more productive individuals, resulting in systematic biases against women. Whatever the rationale, the result is a sex bias closely correlated with wealth and landholdings, and not just preferences and bargaining power, particularly in the South.

### VIII. Conclusions

In poor countries, where consumption expenditures are made almost entirely on food, testing for consumption smoothing over the seasons must take into account seasonal differences in prices, work loads, and returns to labor. We use data on adult nutrition to test smoothing within and across households. Using the Quetelet index, we find that outcomes in Ethiopia, as in many other rural areas in the world, vary a great deal. Fluctuations are larger for women, for individuals in poorer households, and for those in the South of the country.

We model nutrition as a durable good and define its user cost. We use this to look at the intertemporal and intrahousehold allocation of nutrition. We find that nutrition responds to prices and differences in returns to labor. We also find some evidence that poor households are affected by household-specific shocks in agriculture, suggesting the presence of liquidity constraints. Households with more land appear to be better able to self-insure or insure themselves within the community.

Our model also allowed us to test for risk sharing within the household. Using unpredicted illness shocks as a measure of an individual idiosyncratic shock, we find that in most households full risk sharing of illness shocks takes place. However, in poor southern households, households do not pool the illness shocks to women. Since risk sharing can be viewed as a test for Pareto efficiency of allocations, this contradicts the basic premise of the usual models of intrahousehold allocation.

Estimations of intertemporal and intrahousehold regressions were used to retrieve the implicit weights of the allocation of nutrition within families. We find that differences in the ages of husband and wife mattered for allocations, as did the relative wealth of the husband and the customary rules on divorce settlement. However, the wealth of the household—measured by its landholding—had a very large positive effect on the wife's allocation in the South, suggesting that productivity-related effects rather than bargaining may be at the root of the relative bias against women in the South.

### Appendix

#### Derivation of the Pareto Weights

We wish to obtain from equation (17) an estimate of the Pareto weights in the allocation  $(1/\rho) \cdot \ln(\theta_j/\theta_i)$ . In general, in any fixed-effects regression model, we can derive an estimate of the fixed effect as the average error term (Baltagi 1995, p. 27). Suppose that we have  $T$  observations on  $\ln(N_{jt}/N_{it})$ . In (17), this would mean that our estimate for the Pareto weights is equal to

$$\frac{1}{T} \cdot \sum_{i=1}^T \left( \ln \frac{N_{it}}{N_{it}} - \frac{1}{\rho} \ln \frac{\Pi_{it}}{\Pi_{jt}} \right). \quad (\text{A1})$$

This requires information about  $(1/\rho) \cdot \ln(\Pi_{it}/\Pi_{jt})$ . The presence of health fixed effects in the intertemporal model (16) implies that (18) also contains the fixed health effects,  $\pi_j - \pi_i$ ; to eliminate them requires further differencing. We can derive these fixed health effects as the average error in the estimated regression (18). This is, however, still not sufficient to derive the fixed Pareto weights from (17). Effectively, the derived fixed effects from (18) allow us to determine a predictor for

$$\frac{1}{\rho} \cdot \left( \ln \frac{\Pi_{it}}{\Pi_{jt}} - \ln \frac{\Pi_{it-1}}{\Pi_{jt-1}} \right),$$

not for  $(1/\rho) \cdot \ln(\Pi_{it}/\Pi_{jt})$  itself. In other words, we know the difference for each period  $t$  and  $t-1$ , but not the underlying levels. Note that this difference already contains a health fixed effect, but the entire term is not necessarily constant since it also contains time-varying effects such as the patterns of relative changes in the returns to labor over the seasons. Nevertheless, since we know the pattern of change over time, we can derive the levels for each period as well, up to a constant. To see this, let us simplify the notation as follows, by using  $x_{t+1}$ ,  $x_t$ , and  $s_{t+1}$  defined as

$$x_{t+1} - x_t = \frac{1}{\rho} \left( \ln \frac{\Pi_{it+1}}{\Pi_{jt+1}} - \ln \frac{\Pi_{it}}{\Pi_{jt}} \right) = s_{t+1}. \quad (\text{A2})$$

We are interested in an expression for all  $x_t$  to solve (A1). From (A2) it follows that

$$x_\tau = x_0 + \sum_{t=1}^{\tau} s_t, \quad \tau = 1, \dots, T, \quad (\text{A3})$$

and that

$$\begin{aligned} \frac{\sum_{t=1}^T x_t}{T} &= \frac{T \cdot x_0 + T \cdot s_1 + (T-1) \cdot s_2 + (T-2) \cdot s_3 + \dots + s_T}{T} \\ &= x_0 + \sum_{t=1}^T \frac{T-t+1}{T} \cdot s_t. \end{aligned} \quad (\text{A4})$$

Using (A4) and (A1), we can derive the relative Pareto weights up to a constant. The constant term  $x_0$  reflects initial relative health endowments of husband and wife, which are not reflected in the path of nutrition over time. We cannot directly identify this and therefore cannot purge this effect from the Pareto weights, but we can control for some kinds of individual heterogeneity in the regression on the Pareto weights by use of variables such as the relative age gap, health history (whether the husband or the wife had serious illnesses in the five years preceding the survey), the education of their parents, and the availability of health care facilities. While improving the fit, they did not alter the findings and are not reported in table 7. Note further that all models were estimated with controls for heterogeneity in the path of nutrition, providing extensive controls for differences in health endowments between husbands and wives.

## References

- Adams, A. M. "Seasonal Variations in Energy Balance among Agriculturalists in Central Mali: Compromise or Adaptation?" *European J. Clinical Nutrition* 49 (November 1995): 809–23.
- Altonji, Joseph G.; Hayashi, Fumio; and Kotlikoff, Laurence J. "Is the Extended Family Altruistically Linked? Direct Tests Using Micro Data." *A.E.R.* 82 (December 1992): 1177–98.
- Arellano, Manuel, and Bond, Stephen. "Some Tests of Specification for Panel Data: Monte Carlo Evidence and an Application to Employment Equations." *Rev. Econ. Studies* 58 (April 1991): 277–97.
- Baltagi, Badi H. *Econometric Analysis of Panel Data*. Chichester: Wiley, 1995.
- Behrman, Jere R. "Intrahousehold Distribution and the Family." In *Handbook of Population and Family Economics*, edited by Mark R. Rosenzweig and Oded Stark. Amsterdam: North-Holland, 1993.
- Behrman, Jere R., and Deolalikar, Anil B. "Health and Nutrition." In *Handbook of Development Economics*, vol. 1, edited by Hollis Chenery and T. N. Srinivasan. Amsterdam: North-Holland, 1988.
- . "The Intrahousehold Demand for Nutrients in Rural South India: Individual Estimates, Fixed Effects, and Permanent Income." *J. Human Resources* 25 (Fall 1990): 665–96.
- . "The Intrahousehold Distribution of Market Labour Supply in Rural South India." *Oxford Bull. Econ. and Statis.* 55 (November 1993): 409–20.
- Behrman, Jere R.; Deolalikar, Anil B.; and Lavy, Victor. "Dynamic Decision Rules for Child Growth in Rural India and the Philippines: Catching Up or Staying Behind?" Manuscript. Philadelphia: Univ. Pennsylvania, 1995.
- Bergstrom, Theodore C. "Economics in a Family Way." *J. Econ. Literature* 36 (December 1996): 1903–34.
- Bernanke, Ben S. "Permanent Income, Liquidity, and Expenditure on Automobiles: Evidence from Panel Data." *Q.J.E.* 99 (August 1984): 587–614.
- Bertola, Giuseppe, and Caballero, Ricardo. "Kinked Adjustment Costs and Aggregate Dynamics." In *NBER Macroeconomics Annual 1990*, edited by Olivier J. Blanchard and Stanley Fischer. Cambridge, Mass.: MIT Press, 1990.
- Bhargava, Alok. "Nutritional Status and the Allocation of Time in Rwandese Households." *J. Econometrics* 77 (March 1997): 277–95.
- Bourguignon, François; Browning, Martin; Chiappori, Pierre-André; and Lechene, Valérie. "Intra Household Allocation of Consumption: A Model and Some Evidence from French Data." *Annales d'Economie et de Statistique*, no. 29 (January–March 1993), pp. 137–56.
- Browning, Martin; Bourguignon, François; Chiappori, Pierre-André; and Lechene, Valérie. "Income and Outcomes: A Structural Model of Intrahousehold Allocation." *J.P.E.* 102 (December 1994): 1067–96.
- Chiappori, Pierre-André. "Rational Household Labor Supply." *Econometrica* 56 (January 1988): 63–90.
- . "Collective Labor Supply and Welfare." *J.P.E.* 100 (June 1992): 437–67.
- Coate, Stephen, and Ravallion, Martin. "Reciprocity without Commitment: Characterization and Performance of Informal Insurance Arrangements." *J. Development Econ.* 40 (February 1993): 1–24.
- Cochrane, John H. "A Simple Test of Consumption Insurance." *J.P.E.* 99 (October 1991): 957–76.
- Dasgupta, Partha. *An Inquiry into Well-Being and Destitution*. Oxford: Oxford Univ. Press, 1993.

- Deaton, Angus. "Saving and Liquidity Constraints." *Econometrica* 59 (September 1991): 1221-48.
- Deaton, Angus, and Muellbauer, John. *Economics and Consumer Behavior*. Cambridge: Cambridge Univ. Press, 1980.
- Dercon, Stefan, and Krishnan, Pramila. "In Sickness and in Health: Risk-Sharing within Households in Rural Ethiopia." Working Paper no. 97/12. Oxford: Centre Study African Economies, 1997.
- Dugdale, A. E., and Payne, P. R. "A Model of Seasonal Changes in Energy Balance." *Ecology Food and Nutrition* 19, no. 3 (1987): 231-45.
- Fafchamps, Marcel; Udry, Christopher; and Czukas, Katherine. "Drought and Saving in West Africa: Are Livestock a Buffer Stock?" *J. Development Econ.* 55 (April 1998): 273-305.
- Ferro-Luzzi, Anna, et al. "Seasonal Energy Deficiency in Ethiopian Rural Women." *European J. Clinical Nutrition* 44 (suppl. 1; 1990): 7-18.
- Foster, Andrew D. "Prices, Credit Markets and Child Growth in Low-Income Rural Areas." *Econ. J.* 105 (May 1995): 551-70.
- Gertler, Paul, and Gruber, Jonathan. "Insuring Consumption against Illness." Working Paper no. 6035. Cambridge, Mass.: NBER, May 1997.
- Gillespie, Stuart, and McNeill, Geraldine. *Food, Health, and Survival in India and Developing Countries*. Delhi: Oxford Univ. Press, 1992.
- Grossman, Michael. *The Demand for Health: A Theoretical and Empirical Investigation*. Occasional Paper no. 119. New York: Columbia Univ. Press (for NBER), 1972.
- Hayashi, Fumio. "Tests for Liquidity Constraints: A Critical Survey and Some New Observations." In *Advances in Econometrics: Fifth World Congress*, vol. 2, edited by Truman F. Bewley. Cambridge: Cambridge Univ. Press, 1987.
- Hayashi, Fumio; Altonji, Joseph G.; and Kotlikoff, Laurence J. "Risk-Sharing between and within Families." *Econometrica* 64 (March 1996): 261-94.
- James, W. P. T.; Ferro-Luzzi, Anna; and Waterlow, J. C. "Definition of Chronic Energy Deficiency in Adults: Report of a Working Party of the International Dietary Energy Consultative Group." *European J. Clinical Nutrition* 42 (December 1988): 969-81.
- Kaplan, Irving, et al. *Area Handbook for Ethiopia*. Washington: Government Printing Office (for Foreign Area Studies, American Univ.), 1971.
- Levine, Donald N. *Wax and Gold: Tradition and Innovation in Ethiopian Culture*. Chicago: Univ. Chicago Press, 1965.
- Lucas, Robert E. B., and Stark, Oded. "Motivations to Remit: Evidence from Botswana." *J.P.E.* 93 (October 1985): 901-18.
- Lundberg, Shelly, and Pollak, Robert A. "Separate Spheres Bargaining and the Marriage Market." *J.P.E.* 101 (December 1993): 988-1010.
- . "Bargaining and Distribution in Marriage." *J. Econ. Perspectives* 10 (Fall 1996): 139-58.
- Lusky, Ayala, et al. "Relationship between Morbidity and Extreme Values of Body Mass Index in Adolescents." *Internat. J. Epidemiology* 25 (August 1996): 829-34.
- McElroy, Marjorie B., and Horney, Mary Jean. "Nash-Bargained Household Decisions: Toward a Generalization of the Theory of Demand." *Internat. Econ. Rev.* 22 (June 1981): 333-49.
- Mankiw, N. Gregory. "Hall's Consumption Hypothesis and Durable Goods." *J. Monetary Econ.* 10 (November 1982): 417-25.
- Morduch, Jonathan. "Consumption Smoothing across Space: Tests for Village-Level Responses to Risk." Manuscript. Cambridge, Mass.: Harvard Univ., 1991.
- Pankhurst, Helen. *Gender, Development, and Identity: An Ethiopian Study*. London: Zed Books, 1992.

- Paxson, Christina H. "Using Weather Variability to Estimate the Response of Savings to Transitory Income in Thailand." *A.E.R.* 82 (March 1992): 15–33.
- . "Consumption and Income Seasonality in Thailand." *J.P.E.* 101 (February 1993): 39–72.
- Pitt, Mark M.; Rosenzweig, Mark R.; and Hassan, Md. Nazmul. "Productivity, Health, and Inequality in the Intrahousehold Distribution of Food in Low-Income Countries." *A.E.R.* 80 (December 1990): 1139–56.
- Quetelet, Lambert Adolphe Jacques. *Sur l'homme et le développement de ses facultés, ou Essai de physique social*. Paris, 1835. Translated as *A Treatise on Man and the Development of His Faculties*, by Robert Knox. Edinburgh: Chambers, 1842.
- Ravallion, Martin, and Dearden, Lorraine. "Social Security in a 'Moral Economy': An Empirical Analysis for Java." *Rev. Econ. and Statis.* 70 (February 1988): 36–44.
- Rose, Elaina. "Consumption Smoothing and Excess Female Mortality in Rural India." Manuscript. Seattle: Univ. Washington, 1994.
- Rosenzweig, Mark R., and Stark, Oded. "Consumption Smoothing, Migration, and Marriage: Evidence from Rural India." *J.P.E.* 97 (August 1989): 905–26.
- Rosenzweig, Mark R., and Wolpin, Kenneth I. "Credit Market Constraints, Consumption Smoothing, and the Accumulation of Durable Production Assets in Low-Income Countries: Investments in Bullocks in India." *J.P.E.* 101 (April 1993): 223–44.
- Samuelson, Paul A. "Social Indifference Curves." *Q.J.E.* 70 (February 1956): 1–22.
- Satyanarayana, K.; Rao, S. S.; Radhial, G.; and Reddy, V. "Body Mass Index and Mortality Rates." *Nutrition News*. Hyderabad, India: Nat. Inst. Nutrition, 1991.
- Schultz, T. Paul, and Tansel, Aysit. "Wage and Labor Supply Effects of Illness in Côte d'Ivoire and Ghana: Instrumental Variable Estimates for Days Disabled." *J. Development Econ.* 53 (August 1997): 251–86.
- Strauss, John, and Beegle, Kathleen G. "Intrahousehold Allocations: A Review of Theories, Empirical Evidence and Policy Issues." Working paper. East Lansing: Michigan State Univ., 1995.
- Strauss, John, and Thomas, Duncan. "Health, Nutrition and Economic Development." Working Paper no. 95-23. Santa Monica, Calif.: Rand Corp., Labor and Population Program, 1995. (a)
- . "Human Resources: Empirical Modeling of Household and Family Decisions." In *Handbook of Development Economics*, vol. 3A, edited by Jere R. Behrman and T. N. Srinivasan. Amsterdam: North-Holland, 1995. (b)
- Thomas, Duncan, and Chen, C.-L. "Income Shares and Shares of Income: Empirical Tests of Models of Household Resource Allocations." Working Paper no. 94-08. Santa Monica, Calif.: Rand Corp., Labor and Population Program, 1994.
- Thomas, Jonathan, and Worrall, Tim. "Informal Insurance Arrangements in Village Economies." Manuscript. Coventry: Univ. Warwick, 1994.
- Townsend, Robert M. "Intra-regional Risk-Sharing in Thailand: An Uneven but Growing Economy." Manuscript. Washington: Inst. Policy Reform, 1993.
- . "Risk and Insurance in Village India." *Econometrica* 62 (May 1994): 539–91.
- Udry, Christopher. "Risk and Insurance in a Rural Credit Market: An Empirical Investigation in Northern Nigeria." *Rev. Econ. Studies* 61 (July 1994): 495–526.
- . "Gender, Agricultural Production, and the Theory of the Household." *J.P.E.* 104 (October 1996): 1010–46.
- Vosti, Stephen A., and Witcover, J. "Gender Differences in Levels, Fluctuations

and Determinants of Nutritional Status: Evidence from South-Central Ethiopia." Manuscript. Washington: Internat. Food Policy Res. Inst., 1993.

Waler, Hans T. *Height, Weight and Mortality: The Norwegian Experience*. Stockholm: Almqvist & Wiksell (for Oslo Nat. Inst. Public Health), 1984.