# In Sickness and in Health...
# Risk-sharing within households in rural Ethiopia


## Stefan Dercon and Pramila Krishnan



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Centre for the Study of African Economies
Institute of Economics and Statistics
University of Oxford
St Cross Building
Manor Road
Oxford OX1 3UL.

Stefan Dercon is at the Centre for the Study of African Economies and Katholieke Universiteit Leuven. Pramila Krishnan is at the Centre for the Study of African Economies (e-mail: Pramila.Krishnan@economics.ox.ac.uk).



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**Abstract:** To investigate risk-sharing within the household, we model nutritional status as a durable good and we look at the consequences of individual health shocks. For household allocation to be pareto-efficient, households should pool shocks to income. We also investigate whether households can smooth nutritional levels over time. Using data from rural Ethiopia on adult nutritional status, we find that poor households are affected by idiosyncratic agricultural shocks, while richer households are more successful in smoothing nutritional levels. All individuals adjust to predictable changes in earnings and the nutritional status of poor individuals is responsive to seasonal food price fluctuations. Poor southern households are not sharing risk; women in these households bear the brunt of adverse shocks. Finally, we look at the role of inside and outside options in determining the intrahousehold allocation of nutrition of married couples. We find that wives' relative position improves with a smaller age gap between partners, in younger marriages, as well as by favourable customary laws on settlements upon divorce – but the most important variable affecting allocation is household wealth.

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## 1. 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 due to factors beyond their control. Keeping consumption smooth in the face of such fluctuations is difficult and the imperfect markets in credit and insurance markets 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 data from rural Ethiopia, on individual nutritional status, to test whether individuals can smooth nutritional levels over time. Furthermore, we investigate whether households act as risk-sharing institutions 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 examine whether, in the face of fluctuating incomes, the burden of adjustment is borne by women, particularly in poorer households. We use panel data from Ethiopia to ask whether women in poor households are likely to experience more fluctuations in their nutritional status and consumption than men. There is little to suggest that there are substantial differences on average between the nutritional well-being of men and women in Sub-Saharan Africa, in contrast with Asia[1] – but does this hide the fact that women's consumption fluctuates more than that of men?

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: for example, Paxson (1993) and Chaudhuri and Paxson (1994) investigate consumption-smoothing in Thailand and in India over the seasons and establish that the consumption of different groups moves together, suggesting that variation in consumption is determined by seasonal variations in prices or tastes. The presence of liquidity constraints for some households, however, qualifies these findings, for it does limit the ability of such households to smooth their consumption effectively over time (e.g. Morduch (1990), 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 thereby to share the risk of volatile income with others. This is the cross-sectional counterpart to consumption-smoothing across time. If complete markets or some other mechanism, including non-market institutions, implement a full-information pareto-optimal allocation, household consumption should not respond to idiosyncratic shocks in income. The evidence from the United States (Cochrane (1991), Attanasio and Weber (1992), Altonji et al. (1992), Hayashi et al. (1996)) suggests that there is little evidence that such risk-sharing takes place either at the national level or within extended families – but the evidence for developing countries favours at least partial risk-sharing amongst households. Empirical evidence comes from Townsend (1993, 1994) for Thailand and India (accompanied by a dissenting note from Ravallion and Chaudhuri, 1997), Morduch (1991) for India, Deaton (1995) for the Côte d'Ivoire and Udry (1990, 1994) who examines the use of credit-cum-insurance arrangements to pool risks across households in Northern Nigeria. All these tests use data on the household to examine its behaviour in the

[1] There is considerable evidence on this point, for instance, Deolalikar (1995) for Kenya; Vosti and Witcover (1993) for Ethiopia, Pitt, Rosenzweig and Hassan (1990) for Bangladesh and Harriss (1995) for India.

aggregate, but clearly, can be extended to examine risk-sharing within the household. Kotlikoff and Spivak (1981) used simulations to discuss how within households insurance could substitute to a large extent for the purchase of insurance in the form of annuities, without providing a test. Data limitations are a likely cause for the lack of tests within households, in part due to the difficulty of obtaining information on individual incomes and individual levels of consumption.

Nevertheless, there is substantial indirect evidence that in some parts of the world, boys and men are favoured 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 in many cases substantially and significantly lower 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. Rose (1994) finds that in rural India negative rainfall shocks are associated with higher boy and girl mortality rates 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[2].

The extension of risk-sharing models to the examination of risk-sharing within the household is a natural approach to the examination of intrahousehold allocation. In maximising 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 (for instance, the amount of wealth brought in at marriage by each of the partners) and thus maximise the appropriately weighted sum of individual utilities. As Hayashi (1988) and Townsend (1994) point out, a natural test of altruism when household members are able to pool risk amongst 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, controlling for the household's resources, individual shocks do not affect individual outcomes: 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 (Browning et al. (1994), Bourguignon et al. (1993), Chiappori (1992), Deaton (1995)). Chiappori develops models consistent both with the familiar dictatorial model of household allocation and with cooperative bargaining and suggests ways of discriminating between them. The weights determining a pareto-efficient allocation ought to be functions of total resources, prices, and factors that might affect

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[2] Foster employs a model of intertemporal resource allocation under liquidity constraints and examines the ability of households to smooth consumption (or more indirectly, the anthropometric well-being of the children) over time. He finds that changes in health outcomes were influenced by liquidity constraints linked to credit market imperfections.

distribution such as the share of income earned by individuals or extra-environmental effects (as in McElroy (1990)). In essence, testing the dictatorial model involves testing whether incomes are pooled amongst members. However, this is essentially 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, must allow for the pooling of shocks to income.

The use of anthropometric indicators in examining the problem of whether individuals can smooth consumption has substantial advantages. They are relatively easy to collect, 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, using nutritional status to do so, 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[3]. 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, who are liquidity-constrained, might, in the face of a shock to their incomes, be forced to allocate limited resources towards those members who are more productive or more likely to survive.

The paper is organised as follows. In the next section, the basic findings on adult nutritional status in rural Ethiopia are presented. Section 3 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 4, we present the empirical specification and in section 5 a discussion of the data used here, followed by a discussion of the results in section 6. Section 7 discusses the intrahousehold allocation. Section 8 concludes.

[3] Pitt, Rosenzweig and Hassan (1990) examine the impact of higher endowments on nutrient intakes. They find that higher endowments increase nutrient intakes only for men and not women. That is, households seem to allocate more resources to members with better health endowments, who are more likely to work at energy-intensive activities which pay relatively high wages.

## 2. Adult nutrition using the Quetelet index in Ethiopia

The Quetelet or body mass index (QI) is now recognised as a reliable measure of the current nutritional status of adults and was proposed by the Working Party of the International Dietary Energy Consultative Group (James et al. (1988)). It is measured as weight in kilograms divided by squared height in metres[4]. 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 cut-off points for identifying the degree of malnutrition[5], extremely low values of the QI (below 17 or 18) have been associated with higher adult mortality (Waaler (1984)). However, the link between illness and the QI is less clear. For instance, Adams (1995) finds no significant correlation between the QI and the duration and incidence of adult morbidity across seasons. The QI is also a recognised measure of the amount of energy stored in the body and is liable to some fluctuation over a short horizon[6]. It is likely therefore that the index is affected by changes in prices and income (Strauss and Thomas (1995)).

There is increasing evidence of a relationship between QI and productivity, especially at lower levels of the QI[7]. 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 1477 households in 15 areas of the country, with extensive health and socioeconomic data on over 9000 adults and children, interviewed thrice between 1994 and 1995. Further details are given in annex 1.

Table 1 presents mean levels of QI in rural Ethiopia for individuals aged 20 and over. Data are based on 2343 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 per cent of men and 9 per cent of women display a QI below 17[8]. Mean levels of QI are between 19 and 20, which is lower than the usual national averages in

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4 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 social*. He conceived of it as an aggregate measure of human faculties but we do not take this position.

5 See Dasgupta (1993) and James and François (1994) for a discussion.

6 It ought to be stressed that there is little information on morbidity risk of low levels of the QI in developing countries (Satyanarayana (1991)) but there appears to be general agreement that large fluctuations may pose unacceptable risks to health and have functional consequences.

7 The basis for this relationship is its correlation with the maximum oxygen intake during physical work which directly determines the total amount of (physical) work a person can perform. For an excellent discussion, see Dasgupta (1993), pp. 432–436. The QI has been found to be related to physical functioning or the ability to perform certain tasks, in Indonesia (Strauss and Thomas (1995)), and to perform strenuous work (Pitt et al. (1990) for Bangladesh and Bhargava (1997) for Rwanda), while it was found to be positively related to wages and earnings in India and Bangladesh (Behrman and Deolalikar (1989) and Pitt et al. (1990)) and to the productivity of women in Brazil (Thomas and Strauss (1995)).

8 In comparison, data from Indonesia suggest that 7 per cent of men and 8 per cent of women are below 18.

poor countries of between 21 and 23, although not particularly low in comparison to other rural areas in developing countries[9].

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 QI** | 19.9 | 19.5 |

Source: pooled data from round 1 to 3 of ERHS 1994–1995. Adult Population (above 20 years of age). 2343 observations in each round. Grade I Chronic Energy Deficiency is 17<QI<18.4; Grade II: 16-17; Grade III: <16; Normal: >18.5 (see James et al. (1988)). Sample excludes pregnant and lactating women.

We are interested in the fluctuations of the QI over a relatively short time horizon. Table 2 presents measures of fluctuations in the QI for individuals in the sample and comparisons with other studies. The lowest level of the QI as a percentage of the highest is level 91.6 per cent for men and at 90.7 for women[10]. These measures are somewhat lower than for other countries for which comparable data could be found, suggesting very large fluctuations over time in QI, with women being adversely affected. Since these results imply a weight loss of on average 8 to 9 per cent, this does not look like consumption-smoothing.

Table 2 Fluctuations in QI in Ethiopia and in other studies

| **Mean of the lowest QI as percentage of highest over time** | **Men** | **Women** |
|---|---|---|
| ERHS 1994–95 | 91.6 | 90.7 |
| Ethiopia 1990 | | 96.5 |
| South India | 97.2 | 98.8 |
| Burma | 93.4 | |
| Bangladesh | | 91.4 |

Source: data collected from ERHS, Ferro-Luzzi et al. (1990) for a relatively rich area in Ethiopia, Gillespie and McNeill (1992) for South India, Payne and Dugdale (1987) for Burma, and Chowdhury et al. (1981) on Bangladeshi women in landless households. See also Brun (1984) and Lawrence et al. (1985). Data exclude pregnant and lactating women.

The observed scale of adult malnutrition is likely to have functional consequences and affect labour productivity. Data collected on physical functioning do show clear correlations between

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9 See Strauss and Thomas (1995) on Brazil, Indonesia and Cote d'Ivoire for national levels. Dugdale and Payne report means of 20.6 for male farmers in The Gambia and 19.7 for Burmese farmers while Adams (1995) obtains an average of 19.8 for men and 19.3 for women in Mali. Vosti and Witcover (1993) report means for a particular area in South-Central Ethiopia which 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 Southern India of 19.1 for both men and women. In the UK, mean values are close to 24, while in the US the mean value is about 25.

10 Alternatively, one could calculate a coefficient of variation although the number of observations over time is only three. The mean coefficient of variation gives the same picture: 4.8 per cent for men and 5.2 per cent for women.

malnutrition and the ability to perform standard tasks. Table 3 shows the relationship between measured energy deficiency and answers to the questions on functioning asked of each adult.

Table 3 Frequency distribution of functioning by level of malnutrition

| | **Level of Malnutrition** | | | |
|---|---|---|---|---|
| | **Normal** | **Grade I** | **Grade II** | **Grade III** |
| **Can person carry 20 lt water for 20 metres?** | | | | |
| Easily | 89.5 | 79.0 | 67.0 | 64.6 |
| With a little difficulty | 3.3 | 7.3 | 11.0 | 10.3 |
| With a lot of difficulty | 3.7 | 6.8 | 10.7 | 7.0 |
| Not at all | 3.5 | 6.9 | 11.3 | 18.1 |
| **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: ERHS, rounds 2 and 3. Questions were phrased with relevance to local units and if necessary rephrased to suit division of labour by sex.

The correlations are striking. A third of those measured as grade II or grade III malnourished have some difficulty hoeing a field or carrying water, compared with 10 to 15 per cent of normally nourished adults. Almost a fifth of the adults who are grade III malnourished are not able to carry water and cannot hoe a field.

These productivity effects suggest that there are substantial costs to large fluctuations in nutritional status, beyond the direct effects on well-being. However, if labour requirements and returns fluctuate over time, then these productivity effects may also create incentives to reduce food intakes in some seasons to be able to boost nutrition in peak periods. In annex 1, Table A.2 shows the exact timing of peak labour periods (mainly land preparation and harvesting). As in other rural areas, labour requirements are skewed towards certain months. The high fluctuations observed in Table 2 may therefore be compatible with optimising behaviour. Limited means may also result in a skewed allocation of food towards more productive members. If men have higher returns in farming, then higher fluctuations of the QI in women may be a consequence. In analysing the observed nutritional fluctuations, we will need to distinguish productivity differences from differences in preferences.

Table 4 presents the average QI and the average of the minimum QI as a percentage of its maximum level, for men and women, by consumption per capita in quartiles, land holdings per capita in terciles and by geographical location.

Table 4 QI and consumption, land and geographical location

| | Average QI | | Mean minimum as percentage of maximum | |
|---|---|---|---|---|
| | **Men** | **Women** | **Men** | **Women** |
| **by consumption p.c. 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 land holdings p.c. 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 first round of ERHS. Mean consumption per capita is 72 birr per month (about 12 US dollars). Land holdings are land suitable for cultivation and its mean is about 0.3 ha per capita. North includes all villages in Tigray (2), Wollo (1), Gojjam (1) and Northern Shoa (2); Central includes villages in East Shoa (1), South Shoa (1), Arssi (1) and Hararghe (1); South are all villages in the current Southern Ethiopia Peoples' Association (5). People in the North are mainly Amhara and Tigrayan, in Central mainly Oromo and in the South different groups such as Gurage, Kembata, Gedeo and Wolaita.

Average QI increases while its variability falls with consumption levels per capita. Land holdings per capita show a similar pattern: more variability and lower mean QI for households with lower holdings. Finally, mean levels of QI 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 QI 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 level. We explore whether the higher fluctuation in the QI 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 or individual health endowments, or a consequence of a bias against women? We test whether a bargaining view of intrahousehold allocation is consistent with the observed differences. In the next section, we describe our formal approach to these questions.

## 3. The model

We begin by assuming that the household maximises an expected utility function based on 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 non-food items denoted by *O*, which is assumed to be a public good within the household and weakly separable from allocations to *N*. Let β be the discount rate (β<1) and let the household consist of *J* members. The household maximises the following intertemporally additive utility function at each point in time τ[11]:

$$\sum_{t=\tau}^{T} \beta^{t-\tau}.U(N_{1t}(C_{1t}),N_{2t}(C_{2t}))...,N_{Jt}(C_{Jt}),O_t) \quad (1)$$

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

Households maximise 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*[12]. Let $L_{jt}$ be the total available labour of household member *j* at time *t*. Using non-food prices as the numeraire, *r* as the return on asset holdings between periods and nutrient prices as *P*, we can write this equation as:

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

For convenience we will 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 labour supply *L* in the previous period as is likely in agricultural production. Income earned by each member is a function of individual labour supply and is pooled by the household. Each member's income depends on own productivity, on the division of labour across members and on general labour market conditions. In the model we assume that leisure is exogenous; however, labour supply becomes endogenous through its link with nutrition and illness, hence affecting productivity. In particular, we assume that labour supply is the sum of total work time available *T* minus the time lost due to illness *Z*. Time available for work is made dependent on nutrition, as in the usual literature on nutrition-productivity links: better nourished people in poor societies can endure longer hours and perform tasks in less time (Dasgupta (1993), Strauss and Thomas (1995)):

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

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11 Uncertainty will be introduced below.

12 We assume that consumption is always non-negative and the household is not in debt in the final period T (the transversality condition). There is no bequest motive.

with $T_N$>0, $T_{NN}$<0. In principle, it is likely that $T_N$ decreases after some value of $N$; obesity would affect the ability to work, but this is unlikely to affect our sample of households in rural Ethiopia. The links between income, labour supply and nutrition results in some efficient productivity level of nutrition intake.

Nutritional status must be considered as a stock or a 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 without performing activity, the body will use energy, although it only depreciates gradually over time (Foster (1995), Dasgupta (1993)). The function $f$ in (4) 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 $L$ (Higgins and Alderman (1995), Dasgupta (1993)). It is assumed that $f_N$ is between zero and one – some but not total depreciation takes place, while consumption always implies some improvement in nutritional levels. Equation (4) gives the relationship in full:

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

with $0<f_N<1$, $m_C>0$, $m_Z<0$ and $m_{CZ}<0$.

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, an alternative asset to financial assets. While the latter has 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 will have to define the opportunity cost of boosting nutritional status in one period relative to another period (Grossman (1972)). 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 health status in one period only, by boosting consumption (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)[13]:

$$\Pi_{jt} = P_t.\frac{(1+r_{t+1})}{m^{j}_{c_t}} - \frac{f^{j}_{Nt}}{m^{j}_{c_{t+1}}}.P_{t+1} - \frac{\partial Y_{t+1}}{\partial L_{jt}}.T_{N_{jt}} + a^{j}_{L_{jt}}.T_{N_{jt}}.P_t.\frac{(1+r_{t+1})}{m^{j}_{c_t}} \qquad (5)$$

The first term of this user cost 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 will persist in the future, the actual user cost in period t is lower

[13] This user cost is the cost of increasing nutritional status in period t, leaving nutrition in period t+1 unaffected. In annex 2 we presents its derivation.

by the second term, in which the value of what will be carried over to the future is given, controlling for depreciation. Furthermore, increased nutritional status now yields increased returns, since more nutrition increases endurance and therefore income in the future. This marginal income effect has to be subtracted from the cost of boosting nutritional status. Note that this return could be high if the individual is currently far below the optimal level of labour productivity. Working longer or more intensively will result in some additional use of nutrition (since more energy is expended), which needs to be taken into account when considering the net gain in productivity and earnings.

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. In accord with Behrman (1993) and Dasgupta (1993), we will argue that these differences in prices for nutritional status are fundamental in establishing the meaning of observed differences in nutritional status and intakes, observed in many studies on developing countries. To see this, consider the optimal allocation of nutritional status between two individuals *i* and *j* of the same household from optimising (1) subject to the constraints.

$$\frac{U_{N_{it}}}{U_{N_{jt}}} = \frac{\Pi_{it}}{\Pi_{jt}} \qquad \textbf{((6))}$$

It states that in each period households will so allocate consumption 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 both on preferences and the price of nutrition. Even if the household attaches the same weight to the utility of each member, and even if the processes by which the body stores, uses and transforms nutrition are the same between the members, the allocation within the household will not be equal if there are differences in the productivity of different members[14]. Obviously, equal user costs will not necessarily imply equal nutrition either, since the actual shape and the position relative to the 45 degree line will affect the intrahousehold allocation. 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

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[14] Formally, if indifference curves are linear, so that the sum of utilities are maximised 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 (1993) refer to this as the degree of inequality aversion within the household, with L-shaped indifference curves implying maximum concern.

household prefers smooth nutrition to fluctuations for its members[15], the household may choose to switch resources to particular members in particular periods[16].

This weighing of preferences and prices can be further illustrated by writing the instantaneous utility function in (1) as in Samuelson's (1956) consensus model, mirroring a social welfare function with pareto-weights. In particular, assume that this utility function can be written as:

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

in which θ 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 (equation (6)) can be written as:

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

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 if a social welfare function is maximised subject to a pooled resource. In general, the weights could depend on prices, incomes, preferences and other variables, including variables which do not affect the total budget constraint or individual preferences, but still affect the decision process[17] or arise from a cooperative bargaining within the household.

Typically, following 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

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15 This is implied by the additive separable nature of the intertemporal utility function, combined with concave instantaneous utility.

16 For example, it might be optimal to boost the nutritional status of men during the wet season when land needs to be prepared for the next harvest (see also Pitt, Rosenzweig and Hassan (1990)).

17 Examples are 'extra-environmental parameters', variables which do not directly affect marital utility as in McElroy (1990). Other examples are individual (non-earned) incomes, which in the common preference model will be pooled (as in (2)), so the source of this income should not affect the allocation. If extra-environmental variables or individual incomes affect the allocation within the family then this has been taken to be evidence against the common preference model, since, if the decision process could be presented as if one standard utility function was maximised subject to a pooled resource constraint, these variables should not enter the demand or labour supply equations.

marriage' in which spouses may revert to a traditional sexual division of labour or, as Bergstrom (1996) puts it, 'burnt toast and harsh words'. 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[18].

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[19]. 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 labour 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 labour, but also more positive effects 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 modelled. Equation (8) also shows that such productivity and 'collective weights' effects may not be easily identified in a cross-section.

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 or transient changes in variables like income. Nevertheless, whether someone's current income increase 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: by imposing pareto-efficiency 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. Idiosyncratic shocks in the income (or in other variables) of one of the individuals will need to be insured for pareto-optimality to hold. *Collective models are consistent with the absence of income pooling but require income shock pooling: they*

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18 For example, the often-cited example of increased own earnings for the women in households and the positive effects on children's health and food (Kennedy (1989)) is not straightforwardly interpreted as referring to increased 'outside options' if the marriage were to be dissolved. Given norms on patterns of labour allocation and child care, it is unlikely that if the women were to leave the husband, she would retain the same possibility to earn income and care for her children as if she stayed within the household. Rather, it is plausible that her higher earnings give her better inside options within the household and that therefore she is able to exert her influence on expenditure patterns. If this is a shift in her threat point, then the Nash bargaining solution would be affected (Bergstrom (1996)). However, if it only changes her best possible outcome within the marriage, then the Nash bargaining solution is not affected, suggesting that a Kalai-Smorodinsky bargaining solution may be the more appropriate way to characterise observed outcomes (Dasgupta (1993)).

19 Tests along these lines are in Thomas (1990, 1993), Rose (1994), Bourguignon et al. (1993), Browning et al. (1994), Schultz (1990), McElroy and Horney (1981). For reviews, see Lundberg and Pollak (1996), Alderman et al. (1995). Despite the mounting evidence against a simple neoclassical common preference model, some authors have questioned whether empirically these tests indeed suggest that the proposed alternatives, bargaining models or more general pareto-efficient collective household models, are to be favoured over appropriately (i.e. less restrictively) specified common preference household models (Kapteyn and Kooreman (1992), Behrman (1992)).

*view the household as a risk-sharing institution*[20]. This also suggests that a more appropriate test of the collective model may be a test of endowment pooling rather than income pooling[21].

To see this formally, we derive the perfect risk-sharing solution (Townsend (1994), Udry (1994), Cochrane (1991)). Assume that uncertainty is summarised 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:

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

The household can then be viewed as if 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[22]. 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 realisations being state-dependent. In each state and in each period a household level multiplier can be defined related to equation (2) which is the same for each member of the household[23]. 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:

$$\frac{\dfrac{U_{N_{it}}^{s}}{\Pi_{it}^{s}}}{\dfrac{U_{N_{jt}}^{s}}{\Pi_{jt}^{s}}} = \frac{\theta_j}{\theta_i} = \frac{\dfrac{U_{N_{it}}^{s'}}{\Pi_{it}^{s'}}}{\dfrac{U_{N_{jt}}^{s'}}{\Pi_{jt}^{s'}}} \qquad \textbf{((10))}$$

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20 Cross-section evidence on the effects of changes in wage rates or incomes on the allocation within the household may then not only suggest a rejection of the neoclassical common preference model but also of the very assumption of pareto-optimality of collective models.

21 Of course, by using unearned income as the term on which to test income pooling (e.g. Thomas (1990)), some of these problems may not occur. Nevertheless, unearned income flows may also be risky.

22 This assumption could be relaxed to allowing 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 realisations and savings (see Fafchamps (1996)).

23 In annex 3, we describe the intrahousehold and intertemporal first order conditions for the problem under uncertainty.

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 will be obtained[24]. 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 amongst members of the same family in the form of an informal insurance arrangement[25]. In fact, risk-sharing opportunities give powerful incentives for marriage (Kotlikoff and Spivak (1981)). The problems caused by asymmetric information are likely to be limited. Also, it is likely that members care about each other which will reduce incentives to cease participation in agreements and renege on commitments. (This does not necessarily mean that enforcement problems of the risk-sharing arrangement after the realisation of each state of the world would not exist.)

Viewing the household as a risk-sharing institution is a straightforward extension of imposing pareto-efficiency onto all allocations within the household as in the collective approach to modelling the household. Enforcement problems after each realisation of the state of the world imply, however, that the derivation from a one-period game-theoretic model of the pareto-weights determining each individual's outcome is not possible. Nevertheless, such arrangements can come about from repeated interaction and several contributions have determined or characterised constrained pareto-efficient equilibrium informal insurance contracts using a repeated game structure[26]. 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 modelling the household[27].

Incomes within households are likely to covariate, offering substantial scope for risk-trading with other households. Rural villages may be suitable settings for further risk-sharing to take place: information flows relatively freely, enforcement may be relatively 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

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[24] From (5) it can be seen that under uncertainty user costs may vary across states and time as well. It is assumed that once that 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.

[25] See Kotlikoff and Spivak (1981), Hayashi et al. (1996), Foster and Rosenzweig (1995), Fafchamps (1992, 1996).

[26] See Ravallion and Coate (1993), Kimball (1988), Thomas and Worrall (1994), Fafchamps (1996).

[27] In empirical tests, full risk-sharing within (extended) families has been rejected in the US 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. Bourgignon 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 the production allocations within households in Burkina Faso.

family members and evidence has been found to support this[28]. 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 much the same problems.

For much risk self-insurance – through the accumulation and depletion of assets – may be the only alternative for the household, even if it operates as a risk-sharing institution. 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 has to be added:

$$A_t \geq 0 \qquad \textbf{((11))}$$

The Euler equation describing the optimal path over time for the nutritional status of each individual j in the household can then be written as:

$$\frac{U'(N_{jt})}{\Pi_{jt}} = Max\left[\frac{U'(N_{jt}\ (C^*_{jt}))}{\Pi_{jt}},\ E_t\ [\beta.(1+r_{t+1}).\frac{U'(N_{jt+1})}{\Pi_{jt+1}}\ ]\right] \qquad (12)$$

with

$$C^*_{jt} = \frac{A_t + P_t.\sum_{i=1,\ j\neq i}^{J} C_{it} - O_t}{P_t}$$

If households are unconstrained, the standard result of intertemporal allocation is obtained, in which marginal utilities are equated across time, appropriately discounted and controlling for possible price changes over time[29]. 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.

Despite the existence of a risk-sharing arrangement across individuals, liquidity constraints may cause substantial fluctuations in the nutrition of households and its members. Under plausible

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28 Lucas and Stark (1985) focus on migrant remittances; Rosenzweig and Stark (1988) emphasise the role of marriage patterns and transfers between in-laws; Ravallion and Dearden (1988) and Rosenzweig (1988) investigate household transfers.

29 See annex 3.

assumptions, households will hold precautionary savings to deal with possible liquidity constraints, but despite these savings the possibility exists that households have to draw down all their liquid asset holdings[30]. Furthermore, if holding liquid assets is costly or impeded by risk, lumpiness or other factors, their use as a buffer for consumption will be less than perfect, resulting in further fluctuations in consumption (Dercon (1996), Fafchamps (1996)). There is extensive evidence on the use of savings to smooth consumption in developing countries,even if resulting in less than perfect smoothing[31]. Evidence also suggests that poorer households are less able to smooth consumption, suggesting differential liquidity constraints between poorer and richer households[32].

However, the tendency to link the absence of smooth consumption over time with the existence of liquidity constraints may be misleading[33]. Equation (11) shows how this might happen: even if interest rates are constant over time, relative user costs of nutritional status could affect intertemporal allocation. Food prices are one term in relative user costs and if food prices are systematically lower in certain seasons then fluctuations in consumption and therefore in nutritional status may be optimal.

The issue of seasonal user costs is rather more complicated. Abstracting from seasonal differences in the transformation of food into nutrition (for example due to climatic variability over the seasons), different returns to labour 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 labour seasons usually coincide with high food prices, since substantial labour 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 labour seasons, it may make consumption intakes even higher in slack seasons, resulting in even more, but optimal, variability in household consumption over the seasons[34].

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30 In particular, households will respond to increased risk by holding precautionary savings if marginal utility is convex, which is satisfied by, for example, constant relative risk aversion utility functions (Deaton (1991), Zeldes (1989)).

31 See Alderman and Paxson (1992), Paxson (1992), Rosenzweig and Wolpin (1993), Morduch (1990), Fafchamps et al. (1996))

32 Townsend (1994) finds more fluctuations in consumption for landless households in the ICRISAT sample. Findings by Morduch (1991) suggests a link between landlessness and liquidity constraints. See also Dercon (1996) on the link between lumpy or risky liquid asset holdings and limited endowments in Tanzania, as well as the discussion in Dasgupta (1993).

33 Paxson (1993) and Chaudhuri and Paxson (1994) look at fluctuations in household consumption over the seasons and find that liquidity constraints cannot explain the observed patterns. They suggest that price or taste differences over the seasons are likely explanations.

34 Dugdale and Payne (1987) calibrate a model of optimal seasonal fluctuations in nutrition intake and outcomes using data from The Gambia and Burma. They allow for different seasonal workloads, for depreciation but not for price fluctuations – but they include storage losses over the seasons, which results in much the same results as referred to above. They find that their data are quite consistent with such a model. Consumption intakes and nutritional outcomes fluctuations over the seasons are optimal.

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. Empirically, substantial variations in weight-related nutritional status measures have been established in many developing countries[35]. In many cases, members of poor households show more fluctuations than the richer households. Unless there is evidence of differential seasonal productivity and food prices between richer and poorer households, it is hard to see why the observed patterns would reflect optimal consumption smoothing on the part of the poor, rather than suggesting some form of liquidity constraints faced by the latter (Dasgupta (1993)). This is, of course, an empirical question.

## 4. Empirical model and econometric specification

Our empirical approach focuses on the behaviour of nutritional status for individuals across households and over time. Risk-sharing and consumption-smoothing are investigated using the responsiveness to shocks of the path of nutritional status over time and between household members. In particular, we use a specification that allows us to test whether risk-sharing takes places within the household: is the path of the QI within the household smooth across members up to movements in user cost, as required by pareto-efficiency? If we find evidence against it, we have also found evidence against a 'collective' specification of the household allocation process. As in other studies on risk-sharing, we identify idiosyncratic shocks to individual members of a household. Risk-sharing requires that allocations are unaffected by such shocks, except for their effect on the household budget constraint.

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 distinguish idiosyncratic household-level shocks from community-wide shocks. The first are likely to be readily insured within the community through some form of informal insurance arrangement, such as state-contingent credit or other mechanisms, while with poorly functioning markets, for the latter, only self-insurance may be effective. We make a distinction between poorer and richer households, to proxy for liquidity constraints. In particular, if altruism is a luxury good (rather than a moral stance), then individuals in richer households may be less sensitive to household-level shocks.

We also 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'. As we show below, to the extent that idiosyncratic shocks do not affect individual permanent income, the pareto-weights can be retrieved from the tests of risk-sharing. Focusing specifically on married or cohabiting couples, we will proceed in the next section to investigate which factors determine their respective weight in the allocations: assets, altruism or user-cost related efficiencies?

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[35] See for example Chambers, Longhurst and Pacey (1991), Payne (1985a, 1985b), Ferro-Luzzi et al. (1990), Norgan et al. (1992), Adams (1995), Abdullah and Wheeler (1985), Gillespie and McNeill (1992).

Suppose that the utility function of each member of a household is additive-separable in non-food 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 ρ the coefficient of relative risk aversion. An alternative (structural) way of writing the first order conditions of our model is that the weighted marginal utilities for each individual will be equated 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 using (10), the optimal solution for each household member in each period can be written as:

$$\frac{\theta_j . U'(N_{jt})}{\Pi_{jt}} = \lambda_{Jt} = \frac{\theta_j . N_{jt}^{-\rho}}{\Pi_{jt}} \tag{13}$$

First-order conditions linking period *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:

$$Ln\left(\frac{N_{jt}}{N_{jt-1}}\right) = \frac{1}{\rho} Ln\left(\frac{\Pi_{jt-1}}{\Pi_{jt}}\right) + \frac{1}{\rho} Ln\left(\frac{\lambda_{Jt-1}}{\lambda_{Jt}}\right) + \epsilon_{jt} \tag{14}$$

If households do not face credit constraints, then, as in (12), the shadow price of the household budget constraint will be equalised 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 only affect the Lagrangean multipliers to the extent that they shift the life-cycle asset constraint[36]. However, if credit constraints bind, then the intertemporal relationship between the multipliers breaks down as in the first term in (12), and the relevant household-level budget constraint is determined by current asset levels[37]. Shocks to household income would have a larger effect on the evolution of

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[36] One would need to distinguish persistent from non-persistent changes in income, since the implications for the interpretation of the results will be quite different. See Deaton (1992), Muellbauer (1994).

[37] Equation (13) is invalid since an additional multiplier $\mu_{Jt}$ needs to be added, related to the non-negativity constraint on assets, so the first-order condition can be written as:

$$\frac{\theta_j . U'(N_{jt})}{\pi_{jt}} = \lambda_{Jt} + \mu_{Jt} = \frac{\theta_j . N_{jt}^{-\rho}}{\pi_{jt}} \tag{15}$$

Equation (14) can then 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 current asset position changes can produce relatively large effects on $\mu_{Jt}$ compared to $\lambda_{Jt}$.

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 from household-level shocks, for the former cannot be insured within the village. As in Udry (1994), to measure shocks to income we will focus on the causes of income shocks rather than 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. Controlling for the effect on total household life-cycle wealth or, for constrained households, on current wealth levels, under full risk-sharing, 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[38]. 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 the 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 $w_{jt+1}$ 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 that the individual member can commit to earn this return. Idiosyncratic income shocks will be defined as being related to time lost due to illness shocks, to the extent that they cannot be predicted from past history. Formally, we define:

$$Y_{jt} = \varphi_{Jt}.w_{jt}.L_{jt-1} \tag{16}$$

with $w_j$ the return to labour time for individual *j* and $\varphi_{Jt}$ multiplicative risk to returns to labour, which is the same for all members *j* of household *J* and with $E_{t-1}(\varphi_{Jt})$ equal to one. Using equation (3) and defining $Z_{jt-1}$ as the individual specific time lost due to illness, we can define idiosyncratic income shocks as:

$$\frac{\partial Y_{jt}}{\partial Z_{jt-1}} = -\varphi_{Jt}.w_{jt} \tag{17}$$

Working time lost due to individual illness shocks will reduce household earnings by a multiple of (17). 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, while controlling for the effect on household-level earnings in equation (16), ought not to be significant. If these terms matter, then full risk-sharing does not take place, although this test cannot distinguish partial risk-sharing from the absence of it.

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[38] If household members each own separate plots, as in parts of West Africa (see for example Udry (1995)), then one may have more potential to derive direct measures of idiosyncratic income shocks.

However, this test lacks power to distinguish risk-sharing. Illness may not only affect the ability to earn income, but it could also have a direct effect on nutritional status. For example, illnesses such as gastro-intestinal ailments could result in direct weight loss from an inability to absorb nutrients[39]. Within the model this is captured via equation (4), in which it was assumed that $m_{cz}$, the cross-derivative of the nutrient transformation function between the effects of food consumption and illness, is negative: illness may reduce the nutritional status effects of more food consumption. This effect will cause the user cost of nutritional status to change. Formally, even if risk is shared within the household, so that equation (14) holds, and using equation (16) illness shocks in *t-1* will have two effects on the path of nutritional status:

$$\frac{\partial \ln \frac{N_{jt}}{N_{jt-1}}}{\partial Z_{jt-1}} = \frac{1}{\rho}.\frac{\partial \ln \Pi_{jt-1}}{\partial Z_{jt-1}} - \frac{1}{\rho}.\frac{\partial \ln \lambda_{Jt}}{\partial Y_{jt}}.\frac{\partial Y_{jt}}{\partial Z_{jt-1}} \tag{18}$$
$$= -\frac{1}{\rho}.\frac{P_t.(1 + a^j_{L_{jt}}.T_{N_{jt}})(1+ r_{t+1}).m^j_{zc}}{(m^j_{c_{t+1}})^2} + \frac{1}{\rho}.\frac{\partial \ln \lambda_{Jt}}{\partial Y_{jt}}.\varphi_{Jt}.w_{jt}$$

The first term is positive: illness in period *t-1* will result in a lower nutritional status in *t-1* relative to *t*. Without credit constraints, the latter term is likely to be close to zero; with credit constraints it is likely to be positive for constrained households. This means that even if one controls for the effect on household-level income, using an overidentifying restriction for individual illness shocks may not readily be interpreted as testing full risk-sharing. In the empirical application we will therefore have to control for illness shocks which affect the user cost of nutrition.

Equation (14) shows that the path of nutritional status will be affected by the intertemporal evolution of the user cost of nutrition, as defined by (5). To estimate the regression, we introduce variables that can account for these user costs. Since an exact parametrisation of the logarithm of (5) cannot be derived, we content ourselves with writing the path of user costs of nutrition as the sum of time-varying but not individual-specific variables (prices), variables which are similar for individuals with particular characteristics (wages) and of time-varying individual-specific variables to capture individual differences in physiology. In particular, we write this term as the sum of a function of prices, which are assumed to be constant for individuals in each community, a function of wages or returns *w*, which are assumed to be the same for all individuals with characteristics *a* in the same community, and of individual specific factors thus:

$$\ln\frac{\Pi_{jt-1}}{\Pi_{jt}} = \ln\frac{\upsilon_{t-1}(P_{t-1},P_t,r_t)}{\upsilon_t(P_t,P_{t+1},r_{t+1})} + \ln\frac{\upsilon^{a_j}_{t-1}(w_t)}{\upsilon^{a_j}_t(w_{t+1})} + \pi_j\ (m^j_c,\ f^j_N,\ a^j_L,\ T_{N_j}) \tag{19}$$

Higher prices at *t-1* or at *t+1* relative to *t* would increase user costs at *t-1* compared to *t*, as would higher returns to savings today compared to tomorrow, while higher wages at *t* compared to *t+1*

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[39] Lusky et al. (1996) use data on 17-year-old Israeli males to examine the relationship between morbidity and extreme values of the QI. They find that functional disorders associated with underweight are bronchial and lung conditions, intestinal conditions and emotional neuroses. They point out that the links between underweight and intestinal disorders may be a consequence rather than an aetiological effect.

would reduce user costs at *t-1* compared to *t*. The other term will depend on individual physiological characteristics[40].

*The econometric specification of the intertemporal model*

From equations (14), (15), (18) and (19), the equation to be estimated can be written as:

$$\ln(\frac{N_{jt}}{N_{jt-1}}) = \frac{1}{\rho}.\pi_j + \frac{1}{\rho}.\ln\frac{\upsilon_{t-1}(P_{t-1},P_t,r_t)}{\upsilon_t(P_t,P_{t+1},r_{t+1})} + \frac{1}{\rho}.\ln\frac{\upsilon^{a_j}_{t-1}(w_t)}{\upsilon^{a_j}_{t}(w_{t+1})} + \beta_l.S_{Ct} + \gamma_l.S_{Jt} + \delta_j.Z_{jt-1} + \epsilon_{jt} \quad (20)$$

in which $S_{ct}$ and $S_{jt}$ are insurable shocks affecting income occurring at the community and household level respectively in the period between *t-1* and *t*. The coefficients $\beta_l$ and $\gamma_l$ are expected to be non-negative and are allowed to differ (or be larger) for liquidity-constrained households. A test for consumption-smoothing involves testing whether $\beta_l$ and $\gamma_l$ are equal to zero. Risk-sharing within the household would imply that $\delta$ in (20) is zero once the household-level effect of the shock (see (18)) is controlled for. The test will allow for differences between poor and rich households and different effects on different members in the households. Note that the presence of $\pi_j$ requires fixed-effect estimation methods to control for individual heterogeneity[41].

*The econometric specification of the intrahousehold model*

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 (13) 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 (21) in $u_{ijt}$ is an independently distributed random error:

$$\ln(\frac{N_{jt}}{N_{it}}) = \frac{1}{\rho}\ln(\frac{\Pi_{it}}{\Pi_{jt}}) + \frac{1}{\rho}\ln(\frac{\theta_j}{\theta_i}) + u_{ijt} \quad (21)$$

By first-differencing this equation we eliminate the relative pareto-weights to get:

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[40] Note that in deriving (18) from (5), we implicitly assume that at different levels of nutrition and consumption, the slopes of these individual specific functions are different. If they were constant across individuals (so that these functions are linear and equal for different people), then they could be solved out of (5) and no individual-specific effects are needed. Individual heterogeneity is however a major finding in most health analysis and a potential source of serious misspecification (Behrman and Deolalikar (1990), Strauss and Thomas (1995)).

[41] Strauss and Thomas (1995) comment on the problems related to ignoring heterogeneity in the Euler equations estimated by Foster (1995). Behrman et al. (1995) argue strongly for the need to control for heterogeneity in dynamic models of nutritional status.

$$\ln(\frac{N_{jt}}{N_{jt-1}}) - \ln(\frac{N_{it}}{N_{it-1}}) = \frac{1}{\rho}.\ln(\frac{\Pi_{jt-1}}{\Pi_{jt}}) - \frac{1}{\rho}.\ln(\frac{\Pi_{it-1}}{\Pi_{it}}) + u_{ijt} - u_{ijt-1} \quad (22)$$

Using (19) and adding an overidentifying variable related to illness shocks at *t-1*, we get:

$$\ln(\frac{N_{jt}}{N_{jt-1}}) - \ln(\frac{N_{it}}{N_{it-1}}) = \frac{1}{\rho}.\pi_j - \frac{1}{\rho}.\pi_i + \frac{1}{\rho}.\ln\frac{\upsilon^{a_j}_{t-1}}{\upsilon^{a_j}_{t}} - \frac{1}{\rho}.\ln\frac{\upsilon^{a_i}_{t}}{\upsilon^{a}} + \delta_j.Z_{jt-1} - \delta_i.Z_{it-1} + u_{ijt} - u_{ijt-1} \quad (23)$$

Significance of $\delta_i$ or $\delta_j$ would signify the absence of risk-sharing. Since by definition this coefficient ought to be equal to the same variables in (20), this constitutes a test of robustness for the risk-sharing results. Note that in obtaining the estimates, the user-cost-related fixed effects in (23) must be eliminated.

First-differencing equation (21) obtains consistent estimates of the parameters, but loses an important variable of interest: the ratio between the pareto-weights which determine the intrahousehold allocation of nutrition between *i* and *j*. Since one of the objectives of this paper is to interpret the nature of the arrangements within the family, we need to retrieve this information. We do this by combining our regression results on the intertemporal and intrahousehold allocation (equations (20) and (23)). From the estimates of the parameters in (20) using fixed-effect estimation techniques, we can derive the individual fixed effects $\pi_j$ and $\pi_i$ (see annex 4 for the derivation). By definition (see (19)), we can write each fixed effect as the difference between period-specific individual effects:

$$\pi_j = \ln\upsilon_{jt-1} - \ln\upsilon_{jt} = (L - 1).\ln\upsilon_j \quad (24)$$

By lagging (24) and by adding (24) to it, we obtain:

$$(L^2 - 1).\ln\upsilon_j = 2.\pi_j \quad (25)$$

Dividing (25) by (24) and subtracting (24) from it, we then obtain:

$$\ln\upsilon_{jt} = 1 - \frac{\pi_j}{2}$$
$$\ln\upsilon_{jt-\varrho} = 1 - \frac{\pi_j}{2} + \varrho.\pi_j \quad (26)$$

Estimates of these effects can be obtained as the average of the residuals in an estimation of equation (20). Further, using (19), we can write (21) as:

$$\ln(\frac{N_{jt}}{N_{it}}) = \frac{1}{\rho}\ln(\frac{\upsilon^{a_i}{}_t.\upsilon_{it}}{\upsilon^{a_j}{}_t.\upsilon_{jt}}) + \frac{1}{\rho}\ln(\frac{\theta_j}{\theta_i}) + u_{ijt} \qquad (27)$$

The estimates of the pareto-weights $1/\rho.ln(\theta_j/\theta_i)$ are obtained as the average residuals in (21), after subtracting the estimated health-fixed effects (as in annex 4). The factors determining the relative weights of household members are investigated by running a regression of individual endowments on this fixed effect.

The assumption that individual effects $\pi_j$ in the relative user costs are fixed is overly restrictive. It means that fixed endowments matter but the history of such accumulation does not. We investigate an alternative specification in which $\pi_j$ in (19) also depends on $N_{jt-1}$. Specifically, we assume that $\pi_j$ is the sum of a fixed effect $\pi^*_j$ and a function of lagged nutritional status[42]:

$$\pi_j = \pi^*{}_j + \gamma.\ln N_{jt-1} \qquad (28)$$

While equations (20) and (23) can be consistently estimated using GLS on the equation in first differences, this is no longer valid in the presence of a lagged endogenous variable. To estimate (20) and (23) we use the optimal GMM estimator derived by Arrellano and Bond (1991). Using $X_{jt}$ to denote all right-hand-side variables besides the fixed effect and the lagged nutritional status variable, the empirical model of (20) can be written as:

$$\ln\frac{N_{jt}}{N_{jt-1}} = \pi_j^* + \gamma.\ln N_{jt-1} + \kappa.X_{jt} + \epsilon_{jt} \qquad (29)$$

where $\epsilon_{jt}$ ~ ID $(0, \sigma_\epsilon^{\,2})$ and $\pi_j^*$ ~ IID$(0, \sigma_\pi^{\,2})$ and where $E[\epsilon_{jt}\,\pi_j] = E[\epsilon_{js}\,\epsilon_{jt}] = 0$ for $t \neq s$ (the residuals are not serially correlated) and $\epsilon$ and $\pi$ are assumed to be independently distributed.

When T=3, $\gamma$ is just identified. Using the equation above in first differences, we obtain:

$$\ln\frac{N_{jt}}{N_{jt-1}} - \ln\frac{N_{jt-1}}{N_{jt-2}} = \gamma.\ln(\frac{N_{jt-1}}{N_{jt-2}}) + \kappa.(X_{jt} - X_{jt-1}) + \epsilon_{jt} - \epsilon_{jt-1} \qquad (30)$$

[42] 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 optimisation model, further simplifying assumptions will need to be made regarding (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.

We use $ln\ N_{jt-2}$ and a set of exogenous and pre-determined 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 Arrellano and Bond estimator, which corrects for possible heteroscedasticity in $\epsilon_{jt}$, is used to obtain consistent estimates of $\gamma$ and $\kappa$. We report the Sargan test for overidentifying restrictions to test the null hypothesis of the validity of the instruments.

## 5. The empirical specification

To estimate equations (20) and (23), we construct a series of variables measuring individual-, household- and village-level shocks in incomes and prices. The variables used are summarised in table 5. First, we need to define variables reflecting price and wage variables relevant for the evolution of user costs of nutrition, as in (20). We assume that returns on assets *r* are relatively constant over the seasons. Since complete data on wages were not available, we construct proxies to describe prices and wages[43]. In (20), with the time-lag in earnings from labour, higher future wages relative to given current levels would increase current nutritional levels compared to past levels. In highly seasonal agriculture (see annex 1), peak labour 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 labour period relative to the previous observation of QI, it would have positive effects on the difference in QI between periods. We introduce sex-specific interaction terms since returns to labour may be different between 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*; similarly, lower food prices in *t* relative to *t-1* reduce the current user cost of nutrition at t. Part of seasonal changes in food prices is predictable, with prices lower in post-harvest periods. Consequently, we introduce dummies indicating whether the measurement of QI was taken in the post-harvest period, when prices were low. The post-harvest period was defined as within four months of the start of the harvest (see annex 1, Table A.2).

Rainfall data for each village, measured as differences in 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 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 community-wide positive income shocks. We also interact rainfall with an indicator of whether the area has an all-weather road. In a country with poor road infrastructure, the presence of a good road may have different effects on prices and incomes.

[43] An alternative specification, using site-specific dummies as in Deaton (1992a) on Côte d'Ivoire, 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 5 Variables used in the regression of changes in the log of QI

| | |
|---|---|
| *Community-level variables* | |
| Rainfall | Change in rainfall as a percentage of mean rainfall in periods relevant for harvest before t and t-1. |
| Rainfall * Road | Change in rainfall multiplied by whether the village has an all-weather road. Rainfall may have different effects on prices and incomes depending on its links with other markets. |
| Post-harvest | Difference in dummies each defined as 1 if QI measurement was taken in post-harvest period, defined as within 4 months of start of harvest. |
| Peak labour | Difference in dummies each defined as 1 if QI measurement was taken during peak period for male labour, such as ploughing 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 ploughing 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 ploughing. 1 if constrained, 0 if not. Used as difference of dummies. |
| No labour | Whether labour could not be hired when needed on the farm. 1 if constrained. Used as difference of dummies. |
| No off-farm | Whether no off-farm wage labour could be found when wanted in the period between observations. 1 if constrained, 0 if not. Used as difference of dummies. |
| Livestock shock | Farm-specific index of problems with livestock related to access to grazing land and to livestock diseases. The higher the worse. 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 Labour Supply* | |
| Male adult illness days | Number of days in the last 28 days preceding the survey that male adults in the households were too ill to work, as a percentage of male adults in the family. The higher the worse. |
| Female adult illness days | Idem, but for female adults. |

| | |
|---|---|
| Number of dead male adults | Number of male adults who died between round t and t-1. |
| Number of dead female adults | Idem, but for female adults. |
| *Individual-Level Shocks* | |
| Idiosyncratic illness shock | Unpredicted shock in working days lost due to illness in last 28 days preceding round t. Residuals obtained from a fixed-effect regression using previous illness experience, lagged QI and controls for pregnancy and breastfeeding. |
| *Time-varying control variables* | |
| Breastfeeding | 1 if women is breastfeeding. Used as change in dummies. |
| Pregnancy | 1 if women is pregnant. Used as change in dummies. |
| Lagged log QI | Lagged dependent variable. Instrumented using Arrellano and Bond estimator in fixed-effects model. 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 log QI. |
| *Interaction Terms* | |
| Male or female | Sex-specific effects |
| South | 1 if village in SEPA, i.e. Southern Ethiopia. |
| Low land | 1 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 1). |

Next, household-level income shock variables were constructed as an index of self-reported adverse occurrences affecting crops and livestock production, reported in Table 5. 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 ploughing. A fourth index measures problems in hiring labour[44]. The last two indices are proxies for the degree of imperfection in oxen and agricultural labour markets and the effect on the farmer. We also constructed an index of problems associated with livestock diseases and sufficient access to

[44] Kochar (1995) finds that households in semi-arid India are well protected from idiosyncratic crop shocks – but this is achieved through well-functioning labour markets rather than well-functioning credit markets.

grazing land. To measure shocks, we use the first difference of each index in each period[45]. 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 the household labour supply. We used the total number of days of illness reported by adults in the household divided by the number of adults as a measure of the reduction in household income from illness[46]. The measures were separately constructed for males and females. We also include the number of male and female adults who died between rounds.

To measure individual idiosyncratic illness shocks in the period preceding the QI measurement, we base ourselves on working days lost due to illness in the previous month, as discussed before. There are two important problems with using this variable[47]. First, as (18) showed, illness not only affects income but also the user cost of nutrition; the same food intake would result in direct weight loss (via a reduced efficiency in transforming food into nutrition). To handle this problem, we investigated the nature of the illness symptoms reported. Just under 30 per cent of adults in the sample reported at least one illness episode in any of the rounds used in this paper. The most common symptoms reported by those ill were severe headaches (20 per cent of cases) and fever (19 per cent). Symptoms such as diarrhoea, vomiting, abdominal pains and other symptoms related to gastro-intestinal illnesses were reported in 15 per cent of the cases. General weakness and fainting (9 per cent) and respiratory problems (8 per cent) are the next most important. Most of these implied illnesses do not appear to result in a direct weight loss with the clear exception of gastro-intestinal diseases. Less than 7 per cent 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 gastro-intestinal disease symptoms. A second problem relates to the endogeneity of illness. Besides the econometric problems of including illness in such regressions, an important issue is that a correct measure of an idiosyncratic shock in a risk-sharing test ought to exclude predictable illness episodes. A measure of unpredictable illness was obtained by using a individual fixed-effect regression of illness days on lagged QI, lagged illness and variables related to breastfeeding and pregnancy.

All the household- and community-level variables describing income shocks were interacted with whether the household was in the lower half of the land per capita distribution in each village.

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[45] 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.

[46] In regressions on health variables, household illness variables may be endogenous. Endogeneity is less likely to be a problem for the household-level variables than for the individual illness variables included as well. Nevertheless, we investigated the consequences of instrumenting the household illness variables. The regression results were not noticeably affected.

[47] Another problem is the self-reporting bias often observed with illness in poor, with more educated or richer people reporting more illness episodes (Behrman and Deolalikar (1988)). In Dercon (1996b), it was found that this problem does not occur in questions related to working days lost due to illness as opposed to questions asking whether the person was ill and for how long. Regressions on duration of illness show predictable correlations with educated and richer people being ill for a shorter time than uneducated and poor people.

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 land-holdings may be a good proxy of liquidity constraints. The land distribution per village is used to allow for the fact that the returns to land are quite dependent on the local agro-climatic circumstances (see annex 1, Table A.1). For instance, in some villages with low land-holdings, highly remunerative permanent crops such as coffee or chat (q'at) can be grown.

To test whether there is risk-sharing within families, the individual illness shocks were interacted with sex. Table 4 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 low land-holdings. Other control variables include changes in the breastfeeding or pregnancy status of women.

Equation (20) was estimated for all adults aged 20 and over for whom we have complete information in all rounds, excluding those suffering gastro-intestinal related illnesses in any round of the survey, leaving 1787 individuals[48]. Equation (20) was also estimated on a smaller sample of heads and spouses of non-polygamous households, a sample of 816 individuals[49]. This sample was also used to estimate (23), providing a consistency test, as well as a means to retrieve 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.

## 6. Estimation results

Table 6 gives the results of the estimation of the intertemporal path of nutritional status. The first set of results are for the entire sample, the second set only for married or cohabiting couples. The dependent variable is the difference in the logarithm of QI at *t* and *t-1*. The model was estimated using OLS and proxies for the lagged value of QI (in first differences, to remove the fixed effects) and also using the generalised method-of-moments estimator developed by Arrellano and Bond (1991). The GMM estimator produces more efficient estimates but the point estimates do not differ significantly between the naive OLS estimator (using proxies) 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, while breastfeeding is not significant. The lagged dependent variable is highly significant, rejecting a pure random walk

[48] We also restricted the sample to those households consisting of at least one male and one female adult to make sure we are not drawing conclusions on intrahousehold allocations between men and women on the basis of households for which this interpretation is invalid.

[49] Polygamy is not very common, so that few observations are lost through this restriction. No detailed information on the nature of the polygamous relationship was collected so little would be gained from including them, except for possible complications in the interpretation of the findings.

specification[50]. Its coefficient suggests a positive correlation over time in the path of nutritional status. Interpreting it as reflecting user costs in period *t-1* relative to *t*, the negative coefficient in the regression suggests user costs are higher at *t* for a higher level of QI obtained at *t-1*. This is consistent with expectations: boosting nutritional status will be more costly at higher levels of QI[51].

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50 Measurement error in the QI could pose a problem in obtaining consistent estimates. However, it is likely that measurement errors are uncorrelated over time: weighing machines 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 weighing machine. If so, this allows a test of the robustness of coefficient estimates by comparing the estimates of Equation (23) (on differences across couples, where identical measurement errors are netted out), with those of Equation (30), for individuals. The coefficients on lagged QI in (30) are -.87 compared with Equation (23) where they are -.80. In short, if measurement error were a problem, we would expect the coefficient on lagged QI in Equation (30) to be biased towards zero compared to the estimates obtained by differencing between husbands and wives. Since there is little difference in the coefficients of both lagged QI and the individual-specific shocks, it might be concluded that household-specific measurement error is unlikely to be an issue.

51 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 0.89 in Bernanke's model, suggesting low adjustment costs.

Table 6. GMM estimation of intertemporal path. Dependent variable is change in logarithm of QI at t and t-1

| | Full sample (1787 adults) | | Couples only (816 adults) | |
|---|---|---|---|---|
| **Variables** | **Coefficient** | **t-value** | **Coefficient** | **t-value** |
| constant | -0.0066 | -1.81+ | -0.0046 | -0.95 |
| log QI t-1 | -0.8925 | -14.51** | -0.8720 | -9.53** |
| breastfeeding t,t-1 | 0.0004 | 0.08 | 0.0080 | 1.09 |
| pregnancy t, t-1 | 0.0187 | 3.11** | 0.0233 | 3.05* |
| *community-level effects* | | | | |
| peak male t, t-1 | 0.0039 | 1.70+ | 0.0057 | 1.79+ |
| peak female t, t-1 | 0.0050 | 2.10* | 0.0060 | 1.75+ |
| high land post harvest t, t-1 | 0.0015 | 0.69 | 0.0049 | 1.34 |
| low land post harvest t,t-1 | 0.0049 | 2.45** | 0.0081 | 2.48* |
| high land rainfall t, t-1 | -0.0010 | -0.09 | -0.0186 | -1.01 |
| low land rainfall t, t-1 | -0.0075 | -0.65 | -0.0164 | -1.02 |
| high land rainfall * road t, t-1 | 0.0235 | 1.28 | 0.0307 | 1.20 |
| low land rainfall * road t,t-1 | 0.0227 | 1.19 | 0.0263 | 0.86 |
| *household level-effects* | | | | |
| high land rain shock t, t-1 | -0.0012 | -0.14 | 0.0020 | 0.17 |
| low land rain shock t, t-1 | 0.0089 | 2.01* | 0.0082 | 1.26 |
| high land crop shock t, t-1 | 0.0239 | 2.23* | 0.0100 | 0.61 |
| low land crop shock t, t-1 | -0.0235 | -2.04* | -0.0237 | -1.26 |
| high land no oxen t,t-1 | -0.0015 | -0.28 | 0.0091 | 1.07 |
| low land no oxen t, t-1 | -0.0104 | -1.76+ | -0.0110 | -1.14 |
| high land no labour t,t-1 | 0.0119 | 1.69+ | 0.0090 | 0.64 |
| low land no labour t, t-1 | 0.0020 | 0.49 | 0.0011 | 0.22 |
| high land livestock shock t, t-1 | -0.0095 | -1.55 | -0.0052 | -0.55 |
| low land livestock shock t, t-1 | -0.0057 | -1.00 | -0.0026 | -0.25 |
| high land livestock loss t, t-1 | -0.0039 | -0.71 | -0.0209 | -0.91 |
| low land livestock loss t, t-1 | -0.0123 | -0.23 | -0.0143 | -0.74 |
| high land no off-farm t,t-1 | -0.0027 | -0.25 | -0.0043 | -0.77 |
| low land no off-farm t, t-1 | -0.0008 | -0.14 | 0.0026 | 0.51 |
| *labour supply shocks* | | | | |
| high land male adults died t , t-1 | -0.0162 | -0.84 | | |
| low land male adults died t, t-1 | -0.0964 | -1.19 | -0.0313 | -0.26 |
| high land female adults died t, t-1 | 0.0155 | 0.70 | | |
| low land female adults died t, t-1 | 0.0034 | 0.22 | | |
| high land male adults ill days t , t-1 | -0.0008 | -1.00 | -0.0013 | -1.47 |
| low land male adults ill days t, t-1 | -0.0005 | -0.58 | -0.0005 | -0.43 |
| high land female adults ill days t,t-1 | -0.0010 | -0.99 | -0.0003 | -0.27 |
| low land female adults ill days t, t-1 | 0.0012 | 1.56 | 0.0011 | 0.77 |
| *idiosyncratic individual shocks* | | | | |
| high land, south, male ill t,t-1 | -0.0010 | -0.95 | -0.0003 | -0.23 |
| low land, south, male ill t,t-1 | 0.0001 | 0.04 | -0.0018 | -0.56 |
| high land, south, female ill t,t-1 | -0.0022 | -1.17 | -0.0003 | -0.23 |
| low land, south, female ill, t,t-1 | -0.0042 | -5.90** | -0.0030 | -3.34** |
| high land, not south, male ill t,t-1 | 0.0013 | 1.17 | -0.0002 | -0.17 |
| low land, not south, male ill t,t-1 | -0.0016 | -1.35 | -0.0013 | -0.69 |
| high land, not south, female ill t,t-1 | 0.0004 | 0.32 | 0.0013 | 0.77 |
| low land, not south, female ill t,t-1 | -0.0007 | -0.81 | -0.0001 | -0.51 |
| Wald Joint Significance | 748.23 | p=0.000 | 404.60 | p=0.000 |
| Sargan Test | 2.05 | p=0.842 | 6.40 | p=0.269 |

**=significant at 1%    *=significant at 5%    +=significant at 10%

The change in the dummy describing whether the current period is a peak period for male labour needs has its expected positive and significant effect on nutritional status. The effect is valid for both for males and for females, suggesting that peak periods are not just relevant for men. Households appear therefore to be adjusting their nutrition to cash in on higher returns to labour in particular periods. Seasonal productivity differences are partly responsible for observed fluctuations. The change in dummies related to the post-harvest period also has the expected sign, suggesting that lower food prices in the post-harvest period will induce increases in nutritional status, as part of arbitrage in user costs of the seasons, as suggested by Payne and Dugdale (1987). The effect is however only significant for households with low land holdings: 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[52]. Rainfall shocks appear to have little effect, although this may be a reflection of the fact that rainfall was good between 1993 to 1995, in most areas. Rainfall in areas with an all-weather road appears to be more favourable to nutritional status than rainfall in areas with poor infrastructure.

Although not all shocks at the household level appear to 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 low land-holdings. Since the farming system is entirely rain-fed and in most cases dependent on ploughing, this indicates that farmers with little land cannot insure themselves or 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 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.

Turning to 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, while 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 for women in southern households with little land. The coefficients suggests that an extra day of unexpected illness for these women would reduce QI in the next period by 0.3-0.4 per cent. 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[53].

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[52] The combined effects of post-harvest and peak period variables could explain a more than one per cent decline in nutritional status over about six months (i.e. between rounds of the survey).

[53] Since the result is asymmetric (men appear to be insured, but not women), it would reflect a set of sustainability constraints resulting in women fully insuring men, but not vice versa.

Equation (23) 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=0.877). The significant coefficients on the control variables, on the lagged dependent variable and on the individual shock variables were again significant with similar values. The coefficient on lagged QI 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[54].

## 7. Nutritional allocation within the marriage

The regressions allowed us to calculate the pareto-weights in the allocation of nutrition using the approach outlined in the section 5. The estimates are an increasing function of the ratio of husband's weight in the allocation of nutrition relative to the wife: 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 (divorce as in McElroy and Horney (1981)) or uncooperative behaviour inside the marriage ('burnt toast and harsh words' as in Lundberg and Pollak (1993)) influence allocations. We also look for evidence that systematic productivity differences rather than bargaining between partners determine allocations within a marriage.

Before turning to testing these propositions, a brief digression on the institution of marriage and its variation by ethnic group and region is warranted, for they are quite different from 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. Amongst the Amhara-Tigre (and to a great extent, the Oromo), the basic unit of organisation 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 are usually between families of equal standing. A formal marriage contract is drawn up and property division upon divorce is equitable[55]. Most studies of the Amhara-Tigre stress that

[54] The coefficient for men in poor southern households was also significantly negative in the robustness test, suggesting that they, contrary to the intertemporal results, 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.

[55] 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.'

the wife is very much her husband's equal within the household[56]. The customs amongst the Oromo are not dissimilar, particularly where there has been close association with the Amhara. However, amongst the Muslim Oromo, customs do deviate from the mainly Orthodox Christian Amhara. Divorce is less common and husbands have the advantage in initiating divorce. Divorce settlements are often less generous towards the wife and the custom of bride price is more prevalent at marriage.

The customs of the southern groups in our sample both 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 upon divorce. The strong patrilineal links seem 'to make the marital relationship a discordant one' (Kaplan et al., 1971). Women cannot sue for divorce but men may do so at whim[57]. Divorce is rare and wives get no share of assets upon divorce.

This anthropological evidence suggests strong differences between 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 per cent of the Northern households in the sample were headed by divorced or separated women, compared to less than 1 per cent 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 QI 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 QI was only 19.42[58].

All this provides some support for a bargaining view with divorce as the threat point. To explore this, a regression was estimated with the relative weights at the left-hand side. Data definitions of the variables used in the determinants of the intrahousehold allocation are in table 7. Descriptive statistics of the variables are given in table A.4 in annex 1. Variables included are characteristics of husband and wife: the age of the husband and the age gap between husband and wife, the difference in their length of residence at the village, and their difference in schooling. 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 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

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56 See Pankhurst (1992), p. 109: 'Transactions appear to be constructed to facilitate divorce and the redivision of wealth between husband and wife. Theoretically, they are in an equal position.' Ibid., p. 112: 'Overall, the position of endowments in the marital scene is determined in a way that allows mobility. Feasts are, in general, of an almost equal nature, although the parents of the bride can spend more on the first serg ceremony; wedding gifts are on a small scale and are made by the groom; land is not exchanged, livestock rights are clearly defined for ease of separation, and labour expended on a homestead can theoretically be translated into grain or money, and thus be repaid.'

57 Kaplan et al. 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.'

58 Since very few females were heads of households in the South equivalent means could not be calculated.

to divorce: whether the assets are shared equally or given to the husband if the couple split up (see Table A.5 in annex 1). Finally, we include a measure of the wealth of the household (land per capita) and the distance of the village to the nearest town.

Table 7 Variables used for determining intrahousehold allocation

| | |
|---|---|
| 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 '000 birr of goods given to the couple as wedding gifts to help them start up household |
| Bride price | Value in '000 birr of goods given by the husband to the bride's family at marriage |
| Divorce settlements | 1 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 (1 if yes; no if husband takes everything) |
| Land | The value of land owned by the household in ha per capita |
| South * Land | Land interacted with a dummy for South |
| Distance town | Distance to nearest town in km |

Some of these variables can be interpreted as reflecting outside options at 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 duration of the marriage, the age gap between spouses and the difference in the years of residence in the village, are not as clearly interpreted as related to outside options. For example, controlling for the age of the husband, the duration of the marriage and the age gap is unlikely to affect opportunities outside the marriage. A spouse who is considerably older may be able to rely on a higher allocation sustained by social norms, while 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 land-holdings owned by the household are not easily interpreted as reflecting relative bargaining power. Land claims may traditionally have been more biased towards males, but at present land-holdings 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 land-holdings with the dummy variable indicating a southern household. Data were only available on the first marriage, so we had to restrict ourselves to husbands and wives each in their first marriage, limiting the sample to 191 couples. Table 8 shows the results.

Table 8: OLS estimation. Dependent variable is the household fixed effect in (23), i.e. the predicted log of the ratio of the pareto-weights of husband over wife, times 1/ρ, using regression on couples (191 couples)

| Variable | Coefficient | t-value |
|---|---|---|
| Age gap | 0.010 | 2.15* |
| Age husband | -0.005 | -1.72+ |
| Difference in years of residence | 0.001 | 0.87 |
| Schooling difference | -0.003 | -0.95 |
| Duration of marriage | 0.004 | 1.26 |
| Husband richer | 0.060 | 1.52 |
| Husband same | 0.101 | 2.77* |
| Value of joint gifts on marriage | -0.034 | -1.54 |
| Bride price | 0.037 | 0.41 |
| Divorce settlements expected | -0.019 | -0.62 |
| Customary divorce rules | -0.055 | -1.39 |
| Land | -0.009 | -0.23 |
| South * Land | -0.353 | -1.82+ |
| Distance to town | 0.002 | 0.69 |
| Constant | 0.060 | 0.73 |
| +=significant at 10% *=significant at 5% | Joint F-test = 1.68+ | |

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 women's position is improved, although the effects are only significant at 20 per cent. Marrying a husband from the same or from a wealthier background is disadvantageous for the women. However, there is also some evidence that variables more likely to affect the norms for 'uncooperative marriage' matter as well. The duration of the marriage and a large age gap between husband and wife worsen the wife's allocation. We also find that while land-holdings are entirely insignificant in the north, they are strongly significant in the south. There is a surprisingly large positive effect on allocations to southern wives of increased landholdings. If land-holdings were to matter as a proxy for individual wealth, then one would have expected that it should improve the men's position, since the division of assets is strongly in favour of the husband in the south. An alternative interpretation is that it is related to productivity effects: if land-holdings are small, the returns to additional labour 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 (see table A.5 in annex 1), he may not find it in his (or the family's) interest to increase the family's labour supply by improving his wife's nutrition. With increasing land-holdings, the marginal returns to labour 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, would also appear 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 land-holdings, and not just preferences and bargaining power, particularly in the south.

## 8. Conclusions

In poor countries, where consumption expenditures are almost entirely on food, testing for consumption-smoothing over the seasons needs to take into account seasonal differences in prices, workloads and returns to labour. 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 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 labour. 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 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 duration of the marriage and ages of husband and wife mattered for allocations, as did the relative wealth of the husband and the customary divorce settlement rules. Some of the evidence appears consistent with a Nash-bargaining solution in which the options available outside marriage matter for allocations, but equally, there is evidence consistent with descriptions of 'uncooperative marriages'. However, the wealth of the household – measured by its land-holding – had a very large positive effect on the wife's allocation in part of the country, suggesting that productivity-related effects rather than bargaining may be at the root of the relative bias against women in some parts of the country.

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## Annex 1: The Ethiopian Rural Household Survey

Since 1994, the Economics Department of Addis Ababa University and the Centre for the Study of African Economies, Oxford University, have been collecting an extensive rural household data set in 15 villages around Ethiopia to study rural welfare dynamics. Six of the villages included in the survey were also surveyed in 1989 by the International Food Policy Research Institute, Washington, D.C. As a consequence, a panel data set for 400 households, going back to 1989, has been constructed. Since the start of the extended survey three rounds have been completed, in 1994 and 1995, covering a total of 1477 households. The attrition rate between 1989 and 1994 was 7 per cent; between the first round in 1994 and the third in 1995, the attrition rate was less than 3 per cent. A further round of data collection is taking place in 1997 with future rounds envisaged.

No attempt has been made to have a 'representative sample' of rural Ethiopia and its more than 50 million people. Instead, a sample of clusters, representative of main agro-ecological zones in the country was drawn (see Bereket Kebede, 1994). The importance of agriculture, and the problem of comparing different farming systems, encouraged cluster-based sampling. Nevertheless, given the geographical spread over the main rural areas of Ethiopia, the sample is likely to provide a very relevant, if not a representative, picture of the state of rural Ethiopia. Households were randomly sampled within each site, and the number of households interviewed in each site was proportional to the population of the region relative to the national population.

The data were obtained from three rounds of interviews spread over an 18-month period, with on average of six months between rounds in each site. Each of the rounds of surveying involved up to 5 hours of interviews, conducted in one-hour blocks during a period of several weeks. Enumerators resided in survey sites during the entire survey period and were usually from the surrounding area. The questionnaire consists of a series of core modules, such as consumption expenditure, wealth changes, incomes, health outcomes including anthropometrics on all individuals and sources of income shocks. Furthermore, in each round data were collected on specific issues, such as agricultural innovation, community sharing mechanisms, long recall on health, parental background information, migration, off-farm activities and skill aquisition, education, fertility, marriage history, etc. By focusing on a series of villages, some of the logistical and financial problems of running a long-term panel survey were avoided, at the cost of representativeness. However, the advantage is that it is much easier to collect and control for community-level characteristics. By having fifteen different areas with well-defined differences within each areas and fairly large samples, the problems of small samples in some existing village-level panels are avoided.

Table A.1 gives the characteristics of the sample sites. Details on the sites surveyed in 1989 can be found in Webb, von Braun and Yohannes (1992). Details on the new sites from the 1994 survey can be found in Bereket Kebede (1994) and in table A.1. Details on the timing of the survey relative to peak labour times for males and harvest periods are in table A.2. Tables A.3 and A.4 give descriptive statistics for the variables used in the regressions. Table A.5 gives details on the position of the women. All data are from the ERHS and the community surveys by the survey team and complementary sociological research (for details see Bevan and Pankhurst (1996)).

Table A.1: characteristics of the sample sites

| **Survey site** | **Location** | **Background** | **Main crops** | **Perennial crops?** | **Mean rainfall in mm** |
|---|---|---|---|---|---|
| Haresaw | Tigray | Poor and vulnerable area. | cereals | no | 558 |
| Geblen | Tigray | Poor and vulnerable area; used to be quite wealthy. | cereals | no | 504 |
| Dinki | N. Shoa | Badly affected in famine in 84/85; not easily accessible even though near Debre Berhan. | millet, teff | no | 1664 |
| Debre Berhan | N. Shoa | Highland site. Near town. | teff, barley, beans | no | 919 |
| Yetmen | Gojjam | Near Bichena. Ox-plough cereal farming system of highlands. | teff, wheat and beans | no | 1241 |
| Shumsha | S. Wollo | Poor area in neighbourhood of airport near Lalibela. | cereals | no | 654 |
| Sirbana Godeti | Shoa | Near Debre Zeit. Rich area. Much targeted by agricultural policy. Cereal, ox-plough system. | teff | no | 672 |
| Adele Keke | Hararghe | Highland site. Drought in 85/86 | millet, maize, coffee, chat | yes, no food | 748 |
| Korodegaga | Arssi | Poor cropping area in neighbourhood of rich valley. | cereals | no | 874 |
| Turfe Kechemane | S.Shoa | Near Shashemene. Ox-plough, rich cereal area. Highlands. | wheat, barley, teff, potatoes | yes, some | 812 |
| Imdibir | Shoa (Gurage) | Densely populated enset area. | enset, chat, coffee, maize | yes, including food | 2205 |
| Aze Deboa | Shoa (Kembata) | Densely populated. Long tradition of substantial seasonal and temporary migration. | enset, coffee, maize, teff, sorghum | yes, including food | 1509 |
| Addado | Sidamo (Dilla) | Rich coffee producing area; densely populated. | coffee, enset | yes, including food | 1417 |
| Gara Godo | Sidamo (Wolayta) | Densely packed enset-farming area. Famine in 83/84. Malaria in mid-88. | barley, enset | yes, including food | 1245 |
| Doma | Gama Gofa | Resettlement area (1985); semi-arid; droughts in 85, 88, 89, 90; remote. | enset, maize | yes, some | 1150 |

Source: Community survey ERHS, Webb and von Braun (1994), Bevan and Pankhurst (1996).

Table A.2: timing of activities and of the survey in 1994–1995

| **Survey site** | **Location** | **Main harvest period** | **Peak labourtime for men** | **Time of interview** | | |
|---|---|---|---|---|---|---|
| | | | | **R1** | **R2** | **R3** |
| Haresaw | Tigray | October–November | March–May, October–November | June–July | January | March |
| Geblen | Tigray | October–November | March–May, October–November | June–July | January | March |
| Dinki | N. Shoa | December | June–September, December | March–April | November | January |
| Debre Berhan | N. Shoa | November–December | June–July, November–December | March–April | October | March |
| Yetmen | Gojjam | November–December | June–July, November–December, January–February | March–April | October | March |
| Shumsha | S.Wollo | October–December | April–June, October–December, January–February | June–July | December–January | May |
| Sirbana Godeti | Shoa | November–December | April–July, November–January, February | March–April | November | March |
| Adele Keke | Hararghe | November–December | March–April, November–December | May–June | October | April |
| Koro-degaga | Arssi | October–November | July–November, October–November | May–June | November–December | May– June |
| Turfe Kechemane | S.Shoa | December | December, June–July | March–April | September–October | March–April |
| Imdibir | Shoa (Gurage) | October–December | October–December, January–February | March–April | October | March |
| Aze Deboa | Shoa (Kembata) | October–November | May–June, October–November | March–April | September–October | March |
| Addado | Sidamo (Dilla) | December–January | December–January, February | March–April | January | March |
| Gara Godo | Sidamo (Wolayta) | August–December | September–October, November–December | March–May | October | March |
| Doma | Gama Gofa | September–December | June–July, October–December | April–May | December–January | May–June |

Source: Community survey ERHS and Bevan and Pankhurst (1996).

Table A.3 Descriptive statistics

| | Mean | Std. Dev. |
|---|---|---|
| Log QI r1 | 2.98 | 0.12 |
| Log QI r2 | 3.00 | 0.11 |
| Log QI r3 | 2.99 | 0.11 |
| Road? | 0.56 | 0.50 |
| Breastfeeding r1 | 0.17 | 0.37 |
| Breastfeeding r2 | 0.15 | 0.36 |
| Breastfeeding r3 | 0.14 | 0.35 |
| Pregnancy r1 | 0.04 | 0.20 |
| Pregnancy r2 | 0.04 | 0.18 |
| Pregnancy r3 | 0.03 | 0.17 |
| Peak female r1 | 0.11 | 0.31 |
| Peak male r1 | 0.07 | 0.25 |
| Peak female r2 | 0.72 | 0.31 |
| Peak male r2 | 0.74 | 0.44 |
| Peak female r3 | 0.11 | 0.31 |
| Peak male r3 | 0.21 | 0.41 |
| Post-harvest r1 | 0.36 | 0.48 |
| Post-harvest r2 | 0.59 | 0.49 |
| Post-harvest r3 | 0.25 | 0.43 |
| No labour r1 | 0.18 | 0.39 |
| No labour r2 | 0.18 | 0.39 |
| No labour r3 | 0.14 | 0.43 |
| No oxen r1 | 0.42 | 0.49 |
| No oxen r3 | 0.28 | 0.45 |
| Crop shock r1 | 0.66 | 0.22 |
| Crop shock r3 | 0.82 | 0.18 |
| No off-farm? r1 | 0.43 | 0.50 |
| No off-farm? r2 | 0.42 | 0.49 |
| No off-farm? r3 | 0.43 | 0.49 |
| Livestock shock | 0.68 | 0.26 |
| Livestock shock | 0.74 | 0.23 |
| Livestock shock | 0.78 | 0.22 |
| Livestock loss r3 | 0.02 | 0.11 |
| Dead females r2? | 0.01 | 0.10 |
| Dead females r3? | 0.01 | 0.11 |
| Dead males r2? | 0.01 | 0.08 |
| Dead males r3? | 0.01 | 0.09 |
| Ind illness female r1 | 0.66 | 3.23 |
| Ind illness male r1 | 0.67 | 3.47 |
| Male adults illness r1 | 1.21 | 3.51 |
| Female adults ill r1 | 1.21 | 3.58 |
| Ind illness female r2 | 0.39 | 2.52 |
| Ind illness male r2 | 0.57 | 3.31 |
| Male adults illness r2 | 1.05 | 3.45 |
| Female adults ill r2 | 1.01 | 3.57 |
| Ind illness female r3 | 0.49 | 3.05 |
| Ind illness male r3 | 0.31 | 2.47 |
| Male adults illness r3 | 0.53 | 2.36 |
| Female adults ill r3 | 1.01 | 3.57 |
| South | 0.34 | 0.47 |
| Lowland | 0.50 | |
| Land per capita | 0.30 | 0.34 |

Table A.4 Descriptive statistics of variables used for determining intrahousehold allocation

| | **mean** | **std. dev** |
|---|---|---|
| Age gap years | 8.09 | 5.53 |
| Age husband years | 44.59 | 13.58 |
| Years residence gap years | 12.01 | 14.30 |
| Schooling difference years | 0.13 | 0.43 |
| Duration marriage years | 18.75 | 12.01 |
| Husband richer | 0.29 | 0.46 |
| Husband same | 0.49 | 0.50 |
| Joint goods ’000 | 0.17 | 0.64 |
| Bride price ’000 | 0.58 | 0.17 |
| Divorce settlements | 0.39 | 0.49 |
| Divorce rules | 0.54 | 0.50 |
| Land ha p.c. | 0.28 | 0.40 |
| Distance town | 7.42 | 4.66 |

Table A.5: Summary of indicators of women's bargaining position within a household by region

| Sites | Primary decision maker | Bride wealth | Dowry | Gifts to couple | Arrangements upon divorce |
|---|---|---|---|---|---|
| *Northern sites* | | | | | |
| Yetmen *(Amhara)* | Joint | Jewellery, clothing | None | Gifts of livestock and household goods from both families | Land and property shared equally |
| Dinki *(Amhara)* | Joint | Jewellery, cash, clothing, livestock | None | Livestock, cash for bride | Christian: gifts, goods divided equally<br>Muslim: clothes + 30 Birr |
| Shumsha *(Amhara)* | Father/ Male head | Clothes | Clothing | Livestock, grain | Property and land divided equally |
| Debra Berhan *(Amhara)* | Joint | None | None | Livestock, clothes | Property shared equally |
| Geblen *(Tigray)* | Joint but age determines hierarchy | None | Livestock and cash | Grain, household goods | Property equally divided |
| Haresaw *(Tigray)* | Joint | Livestock, jewellery, clothing | Cattle, cash | Livestock | House remains husband's property – all else divided equally |
| *Orominya sites* | | | | | |
| Adele Keke | Husband/male head | Flour, firewood, salt, cash | None | None | Wife retains bridal gifts and parental gifts |
| Sirbana Godeti | Joint | Livestock and clothing | None | Household goods | Land and house to the husband; remainder shared equally |
| Turufe Kecheme (Shashemene) | Husband/male head | Cattle, clothes | None | Clothes, furniture (for bride) | Property equally shared if Amhara/Tigrinya household; wife gets no assets if Orominya |
| *Southern Sites* | | | | | |
| Gara Godo | Male head | Clothing, watch | None | House, household goods | Wife receives no assets; land shared if partner in fault |
| Adado, Dilla | Male head, men | None | None | Land, tools for groom; household goods for bride | Wife receives no property/ assets upon divorce |
| Imdibir | Male head | Clothes, cash, special mat | None | None | Wife claims her own property; land divided only if spouse at fault |
| Doma | Male head | Cattle, clothes, jewellery | None | Clothes for bride, household goods | Wife receives no property or assets; only her own property |
| Aze Deboa | Husband, Male head | Clothes, honey, sheep, butter, etc | None | Bride's parents: clothing, household goods, cattle | Wife receives no property or assets |

Source: Community survey ERHS and Bevan and Pankhurst (1996).

## Annex 2: The user cost of nutrition

The user cost is defined as the cost of boosting nutritional status in one period, without affecting future nutritional status. To boost nutrition by one unit at t ($dN_t$), one would need to consume more food (as in (4)), but one also becomes able to work more time during which some energy is expended, as in (3) and (4). We can write this as:

$$dN_{jt} = m^j_{C_{jt}}.dC_{jt} - a^j_{L_{jt}}.L_{N_{jt}}.dN_{jt} \qquad \text{(A1)}$$

Nutrition in period t+1 should remain unchanged. Consequently, we can write this condition, using (4):

$$dN_{jt+1} = f^j_{N_t}.dN_{jt} + m^j_{C_{jt+1}}.dC_{jt+1} = 0 \qquad \text{(A2)}$$

Rewriting these equations, it can be seen that:

$$dC_{jt} = \frac{(1 + a^j_{L_{jt}}.L_{N_{jt}})}{m^j_{C_{jt}}}.dN_{jt} \qquad \text{(A3)}$$

The reduction in assets in period t+1 from this temporary boost in nutritional status can be written as:

$$dA_{jt+1} = -P_t.(1+r_{t+1}).dC_t - P_{t+1}.dC_{t+1} + \frac{\partial Y_{t+1}}{\partial L_{jt}}.T_{N_t}.dN_{jt} \qquad \text{(A4)}$$

Assets are affected by consumption changes in t and t+1, but also by the earnings from the nutrition–income link. This equation can be rewritten as:

$$dA_{jt+1} = -[P_t.\frac{(1+r_{t+1})}{m^j_{c_t}} - \frac{f^j_{N_t}}{m^j_{c_{t+1}}}.P_{t+1} - \frac{\partial Y_{t+1}}{\partial L_{jt}}.T_{N_{jt}} + a^j_{L_{jt}}.L_{N_{jt}}.P_t.\frac{(1+r_{t+1})}{m^j_{c_t}}].dN_{jt} \qquad \text{(A5)}$$

The term in rectangular brackets is the cost of boosting nutrition in one period only.

## Annex 3: Optimal solutions under uncertainty

The optimal solution at $\tau$ for the problem of allocating nutrition over time and across individuals from maximising (9), subject to (2), (3) and (4), and using (5), can be written as:

$$E_{\tau}\left(\frac{\theta_j . U_{N_{jt}}}{\Pi_{jt}} - \frac{\theta_i . U_{N_{it}}}{\Pi_{it}}\right) = 0 \qquad for\ t = \tau, \ldots, T\ ;\ i, j = 1, \ldots, J \tag{A6}$$

$$E_{\tau}\left(\frac{U_{N_{jt}}}{\Pi_{jt}} - \beta.(1+r_{t+1})\frac{U_{N_{j,t+1}}}{\Pi_{j,t+1}}\right) = 0 \qquad for\ t = \tau, \ldots, T\ ;\ j = 1, \ldots, J \tag{A7}$$

## Annex 4: The derivation of the fixed effects

Let the residuals be denoted by:

$$u_{it} = \mu_i + v_{it} \tag{A8}$$

with i = 1,2, .... N and t = 1,....T. Equivalently, we can write this using matrices as:

$$u = Z_{\mu}.\mu + v \tag{A9}$$

with dimensions u and v (NT x 1), $Z_{\mu}$ (NT x N) and $\mu$ (N x 1). $Z_{\mu}$ is a stacked matrix of T identity matrices $I_N$ with dimension (N x N). Defining $J_T$ as a matrix of values 1, we can write:

$$\begin{aligned} Z_{\mu}.Z_{\mu}' &= (I_N \otimes J_T) \\ Z_{\mu}'.Z_{\mu} &= T.I_N \\ (Z_{\mu}'.Z_{\mu})^{-1} &= \frac{1}{T}.I_N \end{aligned} \tag{A10}$$

The fixed effects can then be calculated as:

$$\begin{aligned} Z_{\mu}.\mu &= Z_{\mu}[(Z_{\mu}'Z_{\mu})^{-1}.Z_{\mu}'.u] \\ &= Z_{\mu}[(\frac{1}{T}.I_N).Z_{\mu}'.u] \\ &= \frac{1}{T}.(I_N \otimes J_T).u \end{aligned} \tag{A11}$$

The typical element of this vector is:

$$[\;.\;] = \frac{\sum_{t=1}^{T} u_{it}}{T} \qquad \text{(A12)}$$

In short, the fixed effect is the average error for each individual i from equation (A8).