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Intra-Household Resource Allocation
An Inferential Approach

Duncan Thomas

ABSTRACT

If household income is pooled and then allocated to maximize welfare then income under the control of mothers and fathers should have the same impact on demand. With survey data on family health and nutrition in Brazil, the equality of parental income effects is rejected. Unearned income in the hands of a mother has a bigger effect on her family's health than income under the control of a father; for child survival probabilities the effect is almost twenty times bigger. The common preference (or neoclassical) model of the household is rejected. If unearned income is measured with error and income is pooled then the ratio of maternal to paternal income effects should be the same; equality of the ratios cannot be rejected. There is also evidence for gender preference: mothers prefer to devote resources to improving the nutritional status of their daughters, fathers to sons.

I. Introduction

There are remarkably few empirical examinations of how households allocate resources among their members. In most (but not all) household surveys, consumption and expenditure data are collected at

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the household rather than individual level and so individual consumption is not directly observed. This poses no problem for goods, such as leisure, which are consumed by only one member in the household: but there are few such goods and their identification is not trivial. Most empirical tests of household allocation models have, therefore, focused on leisure demand (or labor supply).

The theoretical literature on economic models of household behavior dates back at least to Becker's (1964) extension of the neoclassical model of (individual) consumer demand to families. All members of the household are assumed to jointly maximize some household level welfare function and income is allocated so that the marginal rate of substitution between any two goods is the same as for any other pair. Essentially, as long as the household remains intact, it may be treated as if it acts as a single individual; put another way, all resources are pooled and then reallocated according to some common rule.

Recently, research has focused on explicitly modelling intrahousehold allocation within a bargaining framework,[1] as a Pareto efficient outcome,[2] by assuming a particular structure for parental preferences,[3] or within a class-based model of conflict.[4] All these models permit heterogeneity in preferences among household members but differ in assumptions about the allocation mechanism. Unlike the Beckerian model, there is an incentive for household members not to pool income but rather to allocate resources over which they have discretion toward goods they especially care about.

The aim of this paper is to determine whether the Beckerian model of common preferences is consistent with data; rejection of this simple model should be a precursor to a search for an alternative structural model of household decisionmaking. Instead of examining individual leisure demands, we shall attempt to infer how resources are allocated by focusing on a series of outcomes of household resource allocations: in particular, nutrient intake, child health, survival, and fertility. According to the Beckerian model, the effect on these outcomes of unearned income should be the same, independent of who controls it. However, if the number and healthiness of children enter parents' utility functions differently and if household resources are not pooled, then each parent will want to allocate a different quantity of resources to these outcomes; the

1. See Manser and Brown (1980), McElroy and Horney (1981, 1990), McElroy (1990), Ulph (1988) in the economic literature and Blumberg (1988) for a review of the sociological literature. Behrman (1990) provides an excellent survey of this literature.
2. Chiappori (1988a) and Kooreman (1988).
3. Behrman, Pollak, and Taubman (1982, 1986).
4. Folbre (1986).

effect of unearned income of the father may be different from that accruing to the mother. It is this joint hypothesis which will be tested using a large scale household survey from Brazil.

Second, we shall present evidence on whether parents display gender preferences. If the common preference model of household resource allocation is incorrect then these gender preferences should also be reflected in differential resource allocations depending on who controls income.

II. Models of Household Behavior

Models of household behavior considered in this paper can all be cast in the same framework as a Bergson–Samuelson social welfare function. For a particular household, welfare in any period, W, is a function of M individual utility functions, $U_1, U_2, \ldots, U_M$, where M is the number of household members:

$$(1) \quad W = W[U_1(X, Z), \ldots, U_M(X, Z)]$$

X is a vector of commodity demands including leisure, and Z is a vector of home-produced (or nonmarketed) goods such as child health and quality. If there are N goods then X will be of dimension N^*M with typical element, X_{im}, the consumption of the ith good by the mth member. The model is quite general; it imposes weak separability between the felicity functions and separability between the inputs into the Z-good production functions and the elements of X in each felicity function.

If the utility of member m does not enter the aggregator function, then it will be assigned a weight of zero. In the special case that the utility of a member, m, depends only on his own consumption then X_{il} will carry weight zero for all $l \neq m$ in the utility function U_m. This represents *egotistical preferences* (Chiappori 1988a).

In most survey data, it is only household consumption of good i, $X_i = \Sigma_m X_{im}$, rather than individual consumption, which is observed, although leisure is the exception to this rule. Household welfare is maximized subject to the budget constraint:

$$(2) \quad pX = \sum_m w_m T + y_m$$

and production function for each Z:

$$(3) \quad Z = Z(X)$$

where all prices are included in vector p and y_m is the unearned income of member m; w_m, which is an element of p, is the price of time of household member m. Then there exists a household demand function for each element of X and Z, denoted X_i^*, which depends on all prices, including wages, and unearned income of each member:

$$(4)\quad X_i^* = \sum_m^M X_{im}^* = g(p, y_1, \ldots, y_M).$$

In the empirical work below, we shall maintain the strong assumption that unearned income is exogenous and therefore ignore the fact that current unearned income probably reflects past labor supply decisions. It would be preferable to model household resource allocation within a dynamic framework or, perhaps, using the level of resources brought to the marriage (Schultz 1989); neither option is feasible with the data used in this paper.

The simplest static model of the household assumes either that all household members have exactly the *same preferences* or that a *dictator* makes all allocation decisions in which case the aggregator function $W(\cdot)$ assigns a zero weight to all but one member of the household's utility function. For the purposes of this paper, the two assumptions are observationally equivalent and we refer to them as the *common preference* model; it is often called the neoclassical model and underlies Becker's (1964, 1974, 1981) discussion of household formation. For this model to be correct, it is necessary that the household act as if it pools all unearned income and thus the demand functions depend only on *total* household unearned income and *not* its components.

$$(5)\quad X_i^* = \sum_m^M X_{im}^* = g\left(p_m, \sum_m^M y_m\right).$$

These are testable restrictions, not observed in demands based on the general model, (4).

We consider two classes of models which differ from the neoclassical or common preference model and refer to them under the general rubric of *individual preference* models since they emphasize the role of individuals within the household. The first relies on notions of *bargaining* within the household. Manser and Brown (1980) discuss both a Nash and Kalay–Smorodinsky definition of bargaining equilibrium; McElroy and Horney (1981, 1990) and Horney and McElroy (1988) focus on Nash equilibrium; Bjorn and Vuong (1984, 1985) examine both Nash and Stackelberg bargaining models. Ulph (1988) considers a noncooperative Nash model.

Assume, for the moment, that the Nash equilibrium is the appropriate concept and there are M players in the household allocation game. They will choose X and Z to maximize the product of the differences between the utility level each achieves, U^m, and the threat point or reservation utility level, V_o^m, each would achieve outside the household.

$$\text{MAX} \prod_{m=1}^{M} [U^m(X, Z) - V_o^m(p, y_m, A_m)],$$

subject to the budget constraint, (2), and production function for Z, (3). A_m represents nonprice characteristics of the environment an individual would face if he withdrew from the household and y_m is the unearned income he would take away with him. It is assumed that nonwage income is perfectly transportable so that the amount he would carry with him is equal to the amount he controls while a member of the household.[5] There exist a set of household demands for each good, i:

$$(6) \quad X_i^* = \sum_{m}^{M} X_{im}^* = g(p, y_1, \ldots, A_1, \ldots, A_M),$$

which depends on prices, unearned income of each member, and characteristics A_m. In contrast with the common preference model, each element of household unearned income, $y_1, \ldots, y_M$, enters the bargaining model demand functions separately (see Manser and Brown 1980, McElroy and Horney 1981).

Chiappori (1988a, 1988b) has made the appealing argument that imposing a bargaining concept on the household allocation problem is quite restrictive and these restrictions are hard to test. He proposes assuming only that household allocations are *Pareto-efficient* (see, also, Kooreman 1990). As long as members are not purely egotistical then Chiappori argues the implications of his model cannot be distinguished from the neoclassical model except within a revealed preference or nonparametric framework (Afriat 1967, Varian 1982).

This seems overly strong. In the absence of complete pooling of income, a change in nonwage income under the control of different members will have the same effect on demands only if the dictatorial model is correct. If household allocations are assumed to be Pareto-efficient, but parents have different preferences, then demands should depend on prices and individual components of unearned income:

5. Ulph (1988), Folbre (1984), and Blumberg (1988) argue that bargaining power depends on the proportion of total income each member controls. This would be true if labor supplies are exogenously given and no household member has a claim on the earnings of another member if the household splits up. The empirical results reviewed in Blumberg make this assumption and are, therefore, hard to interpret.

$$(7)\quad X_i^* = \sum_{m}^{M} X_{im}^* = g(p, y_1, \ldots, y_M).$$

The implication that unearned income should enter the demand functions (4) in the same way, *independent of the source* of income is a key feature of the dictatorial model, which is not shared by the general model (4) or either of the class of individual preference models discussed above. This suggests a simple and appealing test of the common preference model against a broad class of alternatives.[6] Rejection of the equality of income effects does not imply acceptance of any one of these alternatives; the fact there are plausible alternatives does, however, suggest the test has some power.

McElroy and Horney (1981) report equality of income effects as one of the tests of the bargaining model. Using National Longitudinal Survey data, they find they cannot reject the hypothesis that nonwage income accruing to the husband, wife and other members of the household have the same effect on male and female labor supply.

Using household expenditure data from Thailand, Schultz (1988, 1989, 1990) demonstrates that a woman's unearned income has a significantly larger negative effect on the probability that she enters the wage labor force than does her husband's unearned income. The reverse is true for men. He also examines the impact of nonwage income on fertility rates: more unearned income in the hands of women tends to (significantly) raise fertility; it is little affected by husband's nonwage income.

III. Nutrient Intake and Indicators of Child Health

This paper will study three levels of data from a large-scale household survey. First, at the household level, nutrient intakes—in particular calories and protein—are considered. Second, for each woman who has ever borne a child, we examine the determinants of fertility and child survival rates. Third, two anthropometric indicators—height (conditional on age) and weight (conditional on height)—for children less than 8 years old are analyzed. The two nutrient intakes are X-goods and the four health indicators are Z-goods in the demand functions, (4).

Analysis of nutrient intakes is a straightforward extension of demand analysis; see Behrman and Deolalikar 1988 for a review and discussion

6. Some of the Marxian-feminist models assert that women's bargaining power is independent of the resources she controls but depends on societal norms (Hartmann 1981). Within the context of the test on unearned income, this model is observationally equivalent to the dictator model.

of the current debate about the size of income effects. A small number of studies have used individual nutrient intake data to study intrahousehold inequality in nutrient intake (Behrman 1988, Behrman and Deolalikar 1989, Pitt, Rosenzweig, and Hassan 1990). No studies, however, have estimated elasticities separately for income in the hands of males and females in spite of the implications for policies aimed at raising nutrition levels in households.

There is a large literature on levels and determinants of fertility and mortality; for recent surveys see, *inter alia,* Schultz (1984), and Mensch, Lentzner, and Preston (1986). Mother's education, and possibly father's education, have a negative impact on fertility and positive effect on child survival probabilities. A smaller number of studies have examined the effect of income on these outcomes (Casterline, Cooksey, and Ismail 1989).

Examinations of child anthropometric indicators are reviewed in Cochrane, Leslie, and O'Hara (1982), and Martorell and Habicht (1986). Among nutritionists, child height for age is considered to be a long-run measure of nutritional status and weight for height a shorter-run indicator (Waterlow et al. 1977). Parental education and, to a less extent, household income, typically have a significant positive impact on both anthropometric outcomes even after controlling for genetic endowment (Behrman and Deolalikar 1988, Thomas, Strauss, and Henriques 1989, 1990, Horton 1986). In order to account for the fact that children grow, one might include child's age and transformations of it in a regression; a more parsimonious approach is used here. Child height is standardized by the median height of a well-nourished child of the same age and sex in a reference population; we use the United States as the reference (National Center for Health Statistics 1976). Weight, conditional on height, is similarly standardized.

It is maintained that parents are concerned about all six household and child health outcomes; the question is do their preferences differ? If so, and if income is not pooled in the household, then these differences should be transmitted into differential parental income effects.

The empirical model[7] follows directly from (4):

7. The number of children ever born takes on only a discrete number of values and so ordinary least squares may not be appropriate. Since there are over 20,000 observations and considerably more heterogeneity in fertility in these data than is observed in, say, the United States, taking account of the integer problem in estimation is probably of second order importance. The truncation of survival rates at 0 and 1 has been handled by a Tobit (Trussell and Preston 1982, Thomas, Strauss, and Henriques 1989) and modelled as a binomial process (Thomas, Strauss, and Henriques 1990); in both cases, the results are virtually identical to those based on least squares. Least squares is, therefore, adopted.

(12) $X = M\beta + \varepsilon,$

where M is a vector of prices, including wages, and unearned income of household members. In the absence of price data, location dummies are included; instead of mother's and father's wages—which are observed only for those participating in the wage market—parental education is included. Unearned income accruing to the mother, father, and all others are entered separately. Estimated variance-covariance matrices are corrected for arbitrary heteroskedasticity using the infinitesimal jackknife (Jaeckel 1972, Efron 1982) also proposed by White (1980).

IV. Data and Empirical Model

The common preference model will be tested with Brazilian survey data collected in 1974/75. Family structure in Brazil has changed substantially over the last three decades; see Oliveira and Berquo (1988) for a review and Goldani (1989) for a discussion. Whereas in 1960, 46 percent of the population was urbanized, this proportion had risen to almost 80 percent in 1984 and much of this growth may be attributed to migration. Furthermore, in Brazil, as in much of Latin America, women are more likely to migrate so that the ratio of women to men is higher in cities (1.05) than the countryside (0.94) and is very high in some cities (1.16 in the Northeast city of Fortaleza, for example). At the same time, there has been a tripling of female labor force participation rates to about 36 percent in 1984 and an even more dramatic increase in the proportion of women earning an income (from 7 percent in 1960 to 33 percent in 1984). Over 40 percent of women earn less than the minimum wage and, on average, they earn less than men; more surprisingly, perhaps, the ratio of female to male wages is smallest for illiterates (93 percent) and greatest for those with college education (only 36 percent) (Neuhouser 1989).

While fertility rates have declined from 6.3 in 1960 to 3.6 in 1984 and the infant mortality rate has been cut in half (IBGE 1988), the age at first union and probability of being married have remained remarkably stable. Over this period, 88 percent of men aged 40–49 were married as were 81 percent of women in the 30–39 year age group. The influence of the Catholic church has, however, been eroded: the proportion of marriages performed solely in church has fallen by 50 percent; civil marriages have risen in popularity and there has been a doubling of consensual unions—especially among young adults (Goldani 1989, Henriques 1988).

In spite of these changes, Brazil remains a *machista* society (Neuhouser 1989), suggesting the dictatorial model of household decision-making should perform well. Sociologists, however, argue that even within this society, resource allocation decisions are quite complex. They claim that women have a good deal of control over expenditures, in particular food expenditures, as well as over the distribution of food (Goldani 1989, Neuhouser 1989). If this is true and if nonwage income is a good indicator of power within the household, then differences in the preferences of men and women should be reflected in differential effects of parental income on indicators of household and child health.

The *Estudo Nacional da Despesa Familiar* (ENDEF) is a random national sample of nearly 55,000 households and is very comprehensive by expenditure survey standards.[8] In addition to the usual income, expenditure, and demographic information, the amount of food consumed by the household during seven consecutive 24-hour periods was recorded. This information has been converted into nutrient intakes taking account of wastage. The height and weight of all members of the household were measured. For each woman, the number of children ever born and the number alive at the survey date are also recorded. The data are presented in tabulations by the Brazilian statistical agency (IBGE 1982); fertility and mortality tables are discussed in National Research Council (1983); in addition, nutrient intakes and anthropometric outcomes are discussed in Knight, Mahar, and Moran (1979). Using the household level data, Thomas, Strauss, and Henriques (1989, 1990) examine the determinants of the child health outcomes; Strauss and Thomas (1990) consider the relationship between nutrient intakes and measures of income.

In the survey, each member of the household was asked about their own income and nonwage income was broken down into income from pensions, social security and workers compensation, rents and income from physical assets, financial assets, gifts, and other irregular income. Among urban households in the survey, 44 percent of all men and 36 percent of all women report some income other than earnings. In rural households, however, there is very little reported unearned income; 29 percent of all men and only 11 percent of women report positive unearned income. The very low proportion of women with unearned income is probably due, in part, to the survey design. All income from rural enterprises (including farms) was attributed to the head of the household. Women's unearned income is, therefore, probably underreported and men's overreported. This will seriously contaminate tests of the allocation model and so, in this paper, only urban households are included: a sample of over 25,000 households.[9]

8. The data used in this study cover the Northeast, Southeast (Rio Grande do Sul, Santa Catarina, Parana), Rio de Janeiro, Sao Paulo, and Brasilia.

9. A subset of the estimates reported below have been replicated with the rural data. In most models, we cannot reject the common preference model because income effects are very imprecisely measured.

Among urban households, almost a quarter of men's income comes from nonearned sources. A little less than half of that is from pensions and social security; a quarter is returns on financial or physical assets. About 40 percent of income reported by women is not earned: relative to men, a higher proportion is from pensions (over 50 percent) and social security (almost 20 percent); rather less is from financial or physical assets (13 percent).

Table 1a presents mean total and unearned income by deciles of *per capita* expenditure (PCE).[10] Total annual household income is, on average, Cr$31,401.[11] Almost a quarter is from nonearned sources and this proportion tends to increase with PCE, although households in the bottom decile report relatively high proportions of unearned income.[12] Fathers account for 76 percent of total household income, mothers receive 13 percent, and the rest accrues to other household members. These ratios are remarkably stable across the expenditure distribution. In contrast, the proportion of unearned income attributed to the father rises from about 55 percent in the bottom decile to 75 percent at the top; the mother's share *falls* from 37 percent at the bottom to about half that at the top. If the individual preference model is correct then mother's unearned income should have a bigger influence on consumption patterns at the bottom of the PCE distribution.

On average, men report Cr$5,500 in unearned income; women report about a quarter of that. A third of all men and a fifth of women report some income from nonwage sources so that conditional on receiving nonwage income, the mother–father differential is much smaller: on average, the mother receives just over Cr$7,500 and the father about twice as much.

Table 1b presents mean *per capita* calorie and protein intakes measured at the household level. Both tend to rise with expenditure although the relationship between calories and expenditure is apparently quite nonlinear. When daily *per capita* calories reaches about 2,400, it remains constant even as *per capita* expenditure increases.

Information on children ever born and the proportion who survived to the survey date are recorded for each woman aged 14 to 55. The income characteristics of this dataset are remarkably similar to the household level data in Table 1a. Mean fertility levels and survival rates by deciles

10. Since measured income is likely to reflect a larger transitory component than expenditure, *per capita* expenditure can be thought of as a first approximation to a longer-run welfare indicator.

11. Approximately U.S.$2,000.

12. It is likely that some of these households suffered a temporary earnings shock, part of which was transmitted into expenditures; hence, the relatively low share of earned income.

Table 1a
Mean Household, Father, and Mother Income by Deciles of Household Per Capita *Expenditure*

Decile of PCE	Total Income (in Cr $) HH	Father	Mother	Unearned Income (in Cr $) HH	Father	Mother
1	5,412	4,073	853	1,028	566	385
2	8,921	6,576	1,126	1,360	856	411
3	11,385	8,118	1,394	1,835	1,158	534
4	14,496	10,190	1,696	2,426	1,605	648
5	17,482	12,201	2,158	3,099	2,025	751
6	21,708	15,441	2,725	4,022	2,815	960
7	27,488	20,182	3,345	5,694	4,150	1,127
8	36,971	27,667	4,759	8,652	6,281	1,824
9	52,507	39,407	7,610	12,281	8,633	3,043
10	117,667	95,142	16,464	36,528	27,445	7,077
Average	31,401	23,897	4,213	7,692	5,553	1,676

of PCE are reported in Table 1b. The average urban woman has 4.7 children and fertility declines with PCE; although 60 percent of women have four children or fewer, more than 10 percent have ten or more children. The average survival probability is 0.86; about 63 percent of all women do not report the death of a child. In the bottom decile of PCE, over 75 percent of women have lost at least one child, in the top decile about 15 percent and, on average, survival rates tend to rise with PCE.

The analysis of anthropometric data is restricted to children less than 8 years old and so younger (and poorer) households are included in this level of data. The average urban child in Brazil has the same weight, conditional on height, as the median child in the United States; the longer run indicator of nutritional status, height for age, suggests, however, that Brazilian children are not as well nourished. There is considerably more heterogeneity in the shorter than longer-run nutritional indicators; the standard deviations of weight for height and height for age are 13.7 and 6.6, respectively. Weight for height and, particularly, height for age, increase with PCE; on average, a child in a household at the top decile is the same height as the U.S. median child.

Relative to the household level data, in the child level sample, average

Table 1b
Mean Nutrient Intakes and Child Health Outcomes by Deciles of Household PCE

	Household Level		Mother Level		Child Level	
Decile of HH PCE	Per capita intake of Calories	Protein	# Children Ever Born	Survival Rate	Height for Age	Weight for Height
1	1,460	425	7.36	0.771	91.10	99.21
2	1,744	518	6.56	0.809	92.61	98.72
3	1,931	574	5.80	0.825	93.66	99.70
4	2,067	630	5.19	0.850	94.61	100.10
5	2,185	683	4.65	0.860	95.58	99.82
6	2,276	723	4.28	0.880	96.34	100.10
7	2,329	762	3.75	0.887	97.38	100.69
8	2,407	806	3.40	0.907	98.33	102.19
9	2,427	843	3.11	0.915	99.45	103.37
10	2,417	889	2.77	0.938	100.69	105.35
Average	2,116	681	4.69	0.864	95.97	100.92
# observations	27,547		24,240		26,538	

Notes: Per capita expenditure (PCE) defined to exclude all infrequent purchases such as durables. Protein intake measured in grams. Height for age and weight for height are percentages of United States NCHS median.

income is about 20 percent lower. Unearned income accounts for a smaller proportion of total income (17 percent), the father controls relatively more (80 percent), and the proportions of mothers and fathers reporting any income from nonwage sources are slightly smaller (14 percent and 34 percent, respectively). The relationship between all these variables and *per capita* expenditure is, however, very similar in all three levels of data.

V. Testing the Common Preference Model

A. Tests of Equality of Income Effects

The reduced form coefficient estimates for unearned income are reported in Table 2 for each of the six resource allocation outcomes. Household nutrient intakes are in logarithms and the anthropometric indicators are logarithms of U.S. medians. All the regressions include unearned income, dummy variables for the education of both parents, whether or not a mother or father exists and a dummy variable for each of 15 states. Also included are the age (and age squared) of the household head (in the nutrient regressions), dummies for the age of the mother (in the fertility and survival regressions), and dummies for age and sex of the child (in the anthropometric regressions).

Both parents' unearned incomes are significantly and positively associated with household *per capita* calorie and protein intakes although the relationship is quite nonlinear.[13] In fact, the income pooling hypothesis cannot be rejected in the calorie regression with linear income terms; relaxing the linearity restriction results in a rejection of the hypothesis. In the protein regression, the equality of income effects is rejected in both cases. Furthermore, the effect of maternal income on nutrient demand is between four and seven times larger than income in the hands of fathers (evaluated at mean household unearned income). Relative to men, women apparently direct more resources under their control toward improving household nutrition.

Both men and women use unearned income to reduce fertility and increase the probability that their children survive although the male income effects on survival are not significant. The absolute magnitudes of women's income effects are much larger than men's: almost 20 times bigger for survival rates in both the linear and quadratic models. The income pooling hypothesis is again rejected in all cases.

Mortality risk is a rapidly declining function of a child's age and so it would be desirable to control for exposure in these regressions (Trussell and Preston 1982). In the absence of information on which to compute these controls (such as age at first marriage or even age at first birth), the estimates for younger women may be misleading to the extent that more educated, higher-income women tend to delay child-bearing, thus imparting a positive bias in the estimated income coefficients. Among older women, failure to standardize is unlikely to have much impact on the estimates and so the regressions have been repeated for women who have completed their fertility (45–55 year olds). The results are virtually identical: the impact of parental income is not the same on either fertility or child survival.

Mother's unearned income also positively affects both anthropometric outcomes; father's income is associated with taller children. In both cases, linearity of the income effects cannot be rejected. Maternal income effects are four to eight times bigger than paternal effects. In the anthropometric regressions, the equality of income effects can only be rejected in the weight for height regression.

13. Interactions between maternal and paternal unearned incomes are insignificant (and very close to zero); they are, therefore, excluded.

Table 2
Effect of Mother's and Father's Unearned Income on Household and Child Health

	Log (caloric intake)	Log (protein intake)	# Children Ever Born	Survival Rate	Log (weight for height)	Log (height for age)
Linear Model						
Unearned income of						
Mother	0.456	1.218	−5.911	0.437	0.317	0.110
	[2.4]	[3.7]	[3.7]	[3.3]	[3.6]	[2.0]
Father	0.063	0.170	−0.734	0.024	0.039	0.026
	[1.8]	[2.0]	[2.3]	[1.9]	[1.7]	[2.8]
Other	−0.100	0.074	−1.169	0.066	−0.027	0.022
	[0.7]	[0.7]	[2.6]	[1.4]	[1.3]	[2.7]
Tests for equality of income effects						
Mother = father	2.75	14.17	10.26	9.69	9.18	2.17
[p-value]	[9.7]	[0.0]	[0.1]	[0.2]	[0.2]	[14.0]
Tests for joint significance						
All coefficients	102.80	132.87	339.85	168.82	38.84	187.12
Income	8.96	26.34	26.50	17.30	12.49	7.68
Quadratic Model						
Unearned Income of						
Mother	1.519	3.146	−12.774	0.998		
	[6.4]	[10.0]	[7.2]	[6.5]		

squared	−1.892	−3.487	10.395	−0.850
	[4.9]	[5.6]	[5.0]	[4.2]
Father	0.243	0.651	−2.714	0.045
	[3.7]	[7.2]	[4.1]	[0.9]
squared	−0.054	−0.145	0.559	−0.000
	[3.6]	[6.0]	[3.8]	[0.6]
Other	−0.961	−0.238	−4.489	0.570
	[2.4]	[0.4]	[0.7]	[1.4]
squared	0.524	0.178	1.890	−0.283
	[2.5]	[0.6]	[0.6]	[1.3]
Slopes of income effects at mean				
Mother	0.89	1.99	−9.24	0.71
Father	0.18	0.49	−2.07	0.04
Tests for equality of income effects				
Mother = father	4.14	12.00	94.20	11.71
[p-value]	[1.6]	[0.0]	[0.0]	[0.0]
Tests for joint significance				
All coefficients	96.45	126.73	311.53	150.49
Mother's income	20.80	64.30	59.11	44.02
Father's income	6.49	32.65	17.10	8.92
Other income	2.99	0.26	0.40	1.00

Notes: Income in Cr$ mills. Includes dummies for state of residence, parental education and existence included. Household head's age (and age^2) in nutrient intake regressions; dummies for mother's age in fertility and survival regressions; dummies for child's age and sex in anthropometric regressions. Weight for height and height for age are percentages of U.S. NCHS medians. [*t*-statistics] below estimates; [p values*100] below test statistics; ts and χ^2s based on heteroskedasticity consistent estimates of covariance matrix.

Unearned income accruing to other household members positively affects height for age and is associated with lower fertility; the estimates are quite close to those of father's unearned income. The hypothesis that income effects of mothers, fathers, and others is the same is rejected in all cases except height for age.

If the impact of income on the six health outcomes is not linear then differential estimated income effects may be due to differences, across the income distribution, in the proportion of men and women who report any unearned income. The income effects have been reestimated conditional on the mother or father reporting positive unearned income.[14] Mother's income still has an (absolutely) bigger effect on health outcomes than father's income (Table 3) and these differences are significant in the protein intake, fertility, child survival, and weight for height regressions. Permitting quadratics in income, then the estimated income effects are significantly different in both the nutrient intake regressions (with $F_{2,27502}$ statistics of 10.21 and 33.13 for calories and protein, respectively). Even conditional on reporting any unearned income, that under the control of women has a bigger impact on health than income in the hands of men.

Unearned income, as measured above, includes income from social security and pensions, thus probably incorporating past labor supply behavior. For each respondent in the survey, it is possible to identify income from physical and financial assets; asset income is a more appealing measure of nonwage resources although not even it is purged of previous labor supply and savings decisions. More important, perhaps, very few people report any asset income—about 12 percent of men and 5 percent of women. The lower half of Table 3 reports asset income effects on the six health outcomes. Both parents' asset income has a positive effect on calorie and protein intake; mothers' income effects are significantly bigger than fathers'. Anthropometric measures are positively affected by asset income, although not significantly, and there is no difference between maternal and paternal income effects. Fertility and child survival are unaffected by asset income. The lack of precision in these estimates is hardly surprising given the proportion of nonzero observations. Nevertheless, the nutrient intake results, at least, support those based on the broad definition of unearned income.

The common preference (or neoclassical) model of household resource allocation does not seem to perform well in these health outcome regressions. Relative to fathers (and other household members), mothers appear to be more effective at using the income over which they have control to improve the health of their families.

14. Two dummy variables are included in the regression, one each to identify mothers and fathers who report any unearned income.

Table 3
Robustness of Estimated Parental Income Effects and Tests for Equality of Effects

	Maternal Income	Paternal Income	χ^2 for Equal Income Effects
Impact of Unearned Income Conditional on Reporting any Unearned Income			
Nutrient demand			
Log (per capita calories)	0.356	0.044	3.29
	[2.1]	[1.6]	[19.3]
Log (per capita protein)	1.133	0.141	9.78
	[3.7]	[1.9]	[0.8]
Fertility and child survival			
Children ever born	−6.137	−0.595	6.92
	[0.2]	[0.3]	[98.9]
Survival rate	0.462	0.022	7.95
	[0.0]	[0.2]	[100.0]
Anthropometrics			
Log (height for age)	0.135	0.032	2.70
	[2.0]	[3.1]	[25.9]
Log (weight for height)	0.309	0.039	9.52
	[3.6]	[1.7]	[0.9]
Impact of asset income			
Nutrient demand			
Log (per capita calories)	1.64	0.25	10.59
	[4.1]	[2.0]	[0.0]
Log (per capita protein)	2.80	0.79	5.99
	[3.6]	[2.8]	[5.0]
Fertility and child survival			
Children ever born	−0.68	−0.24	0.02
	[0.2]	[0.3]	[98.9]
Survival rate	−0.01	−0.01	0.00
	[0.0]	[0.2]	[100.0]
Anthropometrics			
Log (height for age)	0.85	0.12	2.03
	[1.7]	[1.8]	[36.3]
Log (weight for height)	2.19	0.14	2.29
	[1.6]	[1.4]	[31.8]

Notes: See Table 2. Asset income is income from financial and physical assets; income measured in Cr$1,000,000.

B. Tests When Income is Measured with Error

In addition to the assumption of exogeneity, there are at least two problems with relying on unearned income to test the common preference model. First, according to theory, in the absence of credit constraints, the present discounted value of lifetime nonlabor income is the appropriate income measure to be included in these regressions; current unearned income is a noisy indicator of this value.[15] Second, even current unearned income is difficult to measure in household surveys. Rejection of the equality of income effects may be due solely to differential measurement errors across individuals in the survey.

Consider a linear version of (4) and, for ease of exposition, let there be two components to household unearned income: maternal and paternal, y_m and y_p, respectively. Letting prices (and all other covariates) be $\bar{M}$ and dropping the i subscripts, then (4) becomes:

$$(8)\quad X^* = \alpha + \beta_m y_m^* + \beta_p y_p^* + \gamma' \bar{M} + \varepsilon$$

where an asterisk on income denotes the correct measure of wealth. Assume that observed nonlabor income is equal to the correct measure contaminated by a linear error, η:

$$(9)\quad y_j = y_j^* + \eta_j \quad j = m, p$$

where η_j has a zero first moment and finite second moment $\sigma^2_{\eta j}$. If η_m, η_p, y_m^* and y_p^* are uncorrelated with each other, then the bias in least squares estimates is:

$$(10)\quad \hat{\beta}_j = \beta_j \cdot \frac{\sigma^2_{yj}}{\sigma^2_{yj} + \sigma^2_{\eta j}} \quad j = m, p.$$

Since β_j is a vector, one element for each dependent variable, $X^* = \{XZ\}$, and the contamination term in (10) is common across equations, the ratio of maternal (or paternal) income effects across two equations, r and q, will be unbiased:

$$(11)\quad \hat{\theta}^j_{rq} = \frac{\hat{\beta}_{jr}}{\hat{\beta}_{jq}} = \frac{\beta_{jr}}{\beta_{jq}} = \theta^j_{rq} \quad j = m, p.$$

If maternal and paternal income effects are equivalent, then it must be that their ratios will also be equivalent. This is an alternative test of the common preference model which permits measurement error in unearned income. The cross-equation restrictions, $h = \theta^m - \theta^p$, can be tested with

15. If one of the bargaining models is correct, then resources carried away from the household in the event it splits up should enter the calculation.

a nonlinear Wald test, $\chi^2 = h'[HVH']^{-1}h$, where H is the matrix of derivatives of the restriction vector, h, and V is the variance covariance matrix of the estimated income effects.

Rejection of the equality of the θs implies rejection of the common preference model. The test embodies strong assumptions: the regressions must be linear in unearned income and the measurement error must be uncorrelated with the correct measure of wealth, all other covariates and the regression error term. Unfortunately, failure to reject the equality of ratios does not have an unambiguous interpretation. It may be that the income effects are equal; alternatively, if mothers value all health outputs included in the test proportionately more than fathers, then the ratio of the income effects will be equal.

Ratios of maternal to paternal income effects are presented in Table 4 together with Wald tests for equality of these ratios.[16] The broader definition of unearned income is used in the top half of the table. In the nutrient and anthropometric regressions, the ratios are quite similar and the χ^2 statistics are small. In the fertility and survival rate regressions, the ratios differ by a factor of 2 to 3 but the equality test cannot be rejected.[17] When asset income is used, all the ratios differ by a factor of about 2 but since the estimates are imprecise, the χ^2 are very small.

Maintaining that nonlabor income is measured correctly, then the common preference model is not consistent with these data. The ratios of income effects are not significantly different from each other: it is not possible to reject the joint hypothesis that unearned income is measured with error, that there is no correlation among the errors and appropriate measures of wealth, that the health outcomes are linear in income *and* that the common preference model is correct.

VI. Testing for Gender Effects in Anthropometric Outcomes

We turn to a second, related issue: gender bias in household resource allocations and focus on child anthropometric indicators.[18] Differential allocations to boys and girls can arise in both the common prefer-

16. Since the tests are based on heteroskedasticity–consistent variance–covariance matrices, the matrix is not block diagonal. The level of observations varies across the three datasets and so the test statistics can only be calculated for each pair.
17. This is because the father's income effect is imprecisely estimated and the covariance between father's and mother's unearned income is high.
18. Nutrient intakes are only observed at the household level; it is not possible to test for gender preference in their allocation. Similarly, the fertility and child survival data preclude tests for sex discrimination.

Table 4
Non-linear Wald Tests for Equality of Ratio of Mother's to Father's Income Effects

	Ratio of Income Effects	Wald Test	
		χ^2	p-value
Unearned Income			
Nutrient demand			
Log (per capita calories)	7.23	0.32	(0.85)
Log (per capita protein)	7.16		
Fertility and child survival			
Children ever born	8.05	1.10	(0.58)
Child survival	18.21		
Anthropometrics			
Log (height for age)	8.12	0.53	(0.77)
Log (weight for height)	4.23		
Asset Income			
Nutrient demand			
Log (per capita calories)	6.44	1.24	(0.87)
Log (per capita protein)	3.53		
Fertility and child survival			
Children ever born	2.80	0.00	(1.00)
Child survival	0.93		
Anthropometrics			
Log (height for age)	15.57	0.29	(0.99)
Log (weight for height)	7.04		

Notes: See Tables 2 and 3.

ence and individual preference models. Parents may have differential preferences with respect to investments in boys relative to girls, there may be differential returns to these investments in the labor market or parents may have differential claims on the returns their children are expected to receive.

Many recent studies have compared levels of child health outcomes by sex. In some cases differences have been observed, most notably in the Indian subcontinent (D'Souza and Chen 1980, Rosenzweig and Schultz 1982 for mortality; Sen 1984, Sen and Sengupta 1983, and Behrman 1988 for anthropometric indicators) although the significance of some of these differences has been questioned (Kakwani 1986). Outside of Asia, however, it appears that differences in levels of outcomes are small and often not significant; see, for example, Strauss (1990) and Svedberg (1987) on Africa, and Schofield (1979) on Latin America. Schultz (1987) argues that there is evidence for gender bias in schooling enrollments and attainments and this bias tends to decline with income. Psacharopolous and Arriagada (1989) present evidence for discrimination against boys in school attendance and performance in Brazil.

Attempts to measure gender bias in the intrahousehold distribution of nutrients suggest boys tend to be favored, at least in South Asia (Rosenzweig and Schultz 1982, and Behrman and Deolalikar 1990, for India; Evenson, Popkin, and Quizon 1980, and Senauer, Garcia, and Jacinto 1988, for the Philippines; Chen, Huq, and D'Souza 1981, for Bangladesh) although part of these differences can be ascribed to different activity levels (Pitt, Rosenzweig, and Hassan 1990, using data from Bangladesh). In contrast, in the equivalence scale literature, there is little evidence for gender bias in the allocation of expenditures in the Côte d'Ivoire and Thailand (Deaton 1989), or in the United States (Gronau 1985).

In the case of child anthropometric outcomes, gender bias can only be identified relative to another population: standardizations based on the sample preclude tests for discrimination.[19] We shall, therefore, examine the *determinants* of child anthropometric outcomes and test for differential impacts of mother's and father's unearned income on the outcomes of sons and daughters. Assume mothers prefer their daughters to be healthy and fathers are more concerned about the health of their sons. The effect of household unearned income on these outcomes will be a weighted average of the impact of mother's and father's income where the weights are given by the bargaining strength of each member. If the common preference model is correct, then there is no reason to distinguish mother's from father's unearned income.

Mean height for age and weight for height of male and female children are presented in Table 5. Relative to the U.S. standards, on average, girls tend to be taller (given age) and heavier (given height) than boys and these differences are significant. When four age groups are distinguished, girls are significantly taller (given age) than boys only among infants (0–5 months) and older children (5–8 years). Older girls are heavier, given height, than boys. One interpretation of these results would be that households care more about the health status of girls than boys. It is rather hard, however, to also explain the fact that infant boys (0–5 month

19. If there is systematic gender bias in the reference population but not in the observations, then the standardizations could impart spurious gender bias.

Table 5
Tests for Differences in Levels of Height for Age and Weight for Height by Gender of Child

	Mean Height for Age					
Age in Months	Female	Male	Difference	Standard Error	*t*-statistic	p-value
0–5	100.39	99.08	1.308	0.34	3.9	0.0
6–23	96.13	96.20	−0.068	0.21	0.3	37.3
24–59	95.97	95.91	0.058	0.14	0.4	33.9
60–107	95.70	95.39	0.306	0.11	2.8	0.3
All	96.08	95.87	0.209	0.07	2.6	0.1

	Mean Weight for Height					
Age in Months	Female	Male	Difference	Standard Error	*t*-statistic	p-value
0–5	104.21	108.48	−4.266	1.11	3.9	0.0
6–23	104.85	104.42	0.423	0.49	0.9	19.4
24–59	100.70	100.26	0.443	0.26	1.7	4.4
60–107	99.94	98.91	1.034	0.22	4.8	0.0
All	101.17	100.67	0.498	0.17	3.0	0.2

Notes: See Table 2. Height for age and weight for height are relative to U.S. median. Difference is female-male; SE is standard error and *t* the associated *t*-statistic. p-values are multiplied by 100.

olds) are significantly heavier, given height, than infant girls. An alternative, and entirely plausible, interpretation is that the standards chosen (in this case the United States NCHS standards) are inappropriate for urban Brazilian children and the sex differentials are spurious. In spite of this difficulty in interpretation, almost all the tests for gender bias in the anthropometric literature are based on these sorts of comparisons.

The anthropometric regressions have been repeated for girls and boys separately (Table 6). Since differences in mean standardized heights and weights are accounted for by differences in intercepts, the impact of covariates should not be affected by the standards chosen.

Both parents' unearned income are positively associated with each outcome. Mother's unearned income has a significant effect on her daughter's weight for height which is about five times larger than the (significant) effect on her son's weight: this difference is also significant. Although the effect of the mother's income on height is larger for daughters, it is not significant for either child. Relative to the effect on daughters, father's unearned income has a significantly bigger effect on his son's weight for height and a slightly larger effect on height (but this difference is not significant). In spite of this evidence for gender preference, in the case of *both* sons and daughters, mother's income has a bigger impact on heights and weights than father's income. The hypothesis that income from each source has the same impact on either boys' or girls' weight for height is rejected: the common preference model of household resource allocation fails to be supported.

If there is no gender preference, then the impact of either parent's education should be the same for both sons and daughters. Mother's education has a bigger impact on daughter's than son's height; father's education affects son's height more. For mothers, the differences are individually significant at low levels of education but disappear as education increases; in the case of fathers, the differences are significant only at higher levels of education. The weight for height regressions suggest similar patterns although parental education tends to have a small and seldom significant effect.

Mothers appear to devote resources to daughters and fathers to sons and, in many cases, the differences are significant. There is, then, evidence for gender bias in the allocation of resources and this evidence is more subtle than that considered in most other studies which only compare levels of outcomes.

VII. Conclusions

The common preference model of household resource allocation postulates that all income is pooled and a dictator determines the allocation (or all household members have the same preferences): the effect of income under the control of different household members should be the same assuming it is measured without error. An equality restriction on unearned income effects is rejected for five of the six outcomes examined: nutrient intakes, fertility, child survival, and child weight for height. Equality of asset income effects is rejected for the nutrient intake regressions.

Ratios of income effects are not significantly different from each other. This is consistent with the common preference model as long as income is measured with error. It is also consistent with differential intrahousehold

Table 6

Testing for Gender Bias: Determinants of Anthropometric Outcomes by Sex of Child

	Weight for Height			Height for Age		
	Females	Males	Difference	Females	Males	Difference
Unearned income						
Mother	1.097	0.198	0.899	0.243	0.094	0.149
	[2.7]	[3.5]	[2.2]	[1.4]	[1.9]	[0.8]
Father	0.006	0.070	−0.064	0.023	0.031	−0.008
	[0.3]	[2.8]	[2.0]	[3.1]	[1.6]	[0.4]
Other	−0.045	2.923	−2.968	0.020	0.793	−0.045
	[6.7]	[3.0]	[3.1]	[4.0]	[1.6]	[1.6]
Education (1) if Mother						
Literate	−0.338	0.126	−0.464	1.756	1.261	0.495
	[1.1]	[0.4]	[1.1]	[11.5]	[8.2]	[2.3]
Completed elementary	0.116	0.641	−0.525	2.681	2.322	0.359
	[0.3]	[1.6]	[0.9]	[13.8]	[11.7]	[1.3]
Completed secondary	2.499	1.979	0.520	4.101	3.705	0.396
	[4.3]	[3.8]	[0.7]	[15.8]	[14.6]	[1.1]

Father						
Literate	−0.705	−0.282	−0.423	1.094	1.517	−0.423
	[2.0]	[0.8]	[0.8]	[6.2]	[8.6]	[1.7]
Completed elementary	−0.406	0.411	0.821	2.297	2.671	−0.374
	[0.9]	[1.0]	[1.3]	[10.7]	[12.3]	[1.2]
Completed secondary	0.845	1.949	−1.104	3.558	4.319	−0.761
	[1.5]	[3.6]	[1.4]	[13.5]	[16.6]	[2.1]
Tests of joint significance						
Income	6.71	10.89		2.53	3.54	
	[0.1]	[0.1]		[11.1]	[6.0]	
Education						
Mother	4.59	7.33		275.30	203.20	
Father	0.05	3.32		146.90	201.30	
Tests of equality of effects						
Income						
Mother = father	7.14	4.14		1.51	1.34	
	[0.3]	[4.2]		[21.8]	[24.7]	
Mother = father = other	13.40	11.95		1.67	3.74	
	[0.1]	[0.3]		[43.4]	[15.4]	
Education						
Mother = father	2.78	0.79		7.03	1.93	

Notes: See Table 2.

preferences but homogeneous in relative weights mothers and fathers attach to the health outcomes.

It is the case, however, that unearned income in the hands of the mother is estimated to have a bigger impact on her family's health than income attributed to the father. For child survival probabilities, the effect is almost 20 times bigger.

There is some evidence for gender preference: mothers prefer to devote resources to improving the heights and weights of their daughters, fathers to sons. The maternal income effects for *both* sons and daughters are much bigger than the effect of paternal income. It may be wise for programs aimed at improving the healthiness of urban households in Brazil—and especially children—to take into account the suggestion that resources in the hands of mothers appear to have a bigger impact on household and child health than resources controlled by fathers.

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