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## Like Father, Like Son: Gender Differences in Household Resource Allocations

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CENTER DISCUSSION PAPER NO. 619

LIKE FATHER, LIKE SON:

GENDER DIFFERENCES IN HOUSEHOLD RESOURCE ALLOCATIONS

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LIKE FATHER, LIKE SON:  
GENDER DIFFERENCES IN HOUSEHOLD RESOURCE ALLOCATIONS

Abstract

Using household survey data from the United States, Brazil and Ghana, we examine the relationship between parental education and child height, an indicator of health and nutritional status. In all three countries, the education of the mother has a bigger effect on her daughter's height; paternal education, in contrast, has a bigger impact on his son's height. There are, apparently, differences in the allocations of household resources depending on the gender of the child and these differences vary with the gender of the parent. In Ghana, relative to other women, the education of a woman who is better educated than her husband has a bigger impact on the height of her daughter than her son. In Brazil, women's non-labor income has a positive impact on the health of her daughter but not on her son's health. If relative education of parents and non-labor income are indicators of power in a household bargaining game, then these results suggest that gender differences in resource allocations reflect both technological differences in child rearing and differences in the preferences of parents.

KEY WORDS: Gender, Child Health, Household Economics

## 1. INTRODUCTION

How are resources allocated within the household? Whereas sociologists and anthropologists have attempted to peer into this black box, economists have paid less attention to the question. In almost every study of household behavior, economic models, implicitly or explicitly, treat resource allocations as the outcome of maximizing a well-behaved household welfare function (Becker, 1964, 1981) without specifying the mechanisms underlying the allocation process.<sup>1</sup>

Economists have paid more attention to a second, related issue: is there evidence for gender bias in the allocation of resources? A mini-industry revolves around discrimination in the labor market and, within the household economics literature, several studies have looked at gender differences in human capital investments. In South and, possibly, Southeast Asia, girls tend to fare worse than boys,<sup>2</sup> but the evidence elsewhere seems quite weak.<sup>3</sup>

Our aim is to determine whether there is evidence for differences by gender in the allocation of household resources. We will focus on child health, as measured by height for age. Tests for gender differences

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<sup>1</sup>Ashenfelter and Heckman (1974) report one of the first tests of the household model. McElroy and Horney (1981) and Manser and Brown (1980) propose alternative models of household decision-making and recently there has been a resurgence of interest in these models; see Chiappori (1988, 1989), McElroy and Horney, (1988), McElroy (1990), Ulph (1988), Folbre (1986) and empirical work by Bjorn and Vuong (1984, 1985), Kooreman (1990), Schultz (1990) and Thomas (1990). Behrman (1990) provides an excellent review of the literature. Blumberg (1988) reviews several empirical studies in the sociology literature.

<sup>2</sup>D'Souza and Chen (1980) and Rosenzweig and Schultz, (1982) find infant and child mortality in India is lower among boys; Sen, (1984), Sen and Sengupta (1983) and Behrman (1988) argue, on the basis of anthropometric indicators, that boys receive preferential treatment in India. Several studies indicate that boys tend to be favored in the intrahousehold distribution of nutrients (Rosenzweig and Schultz, 1982, and Behrman and Deolalikar, 1989, for India; Evenson *et al.*, 1980, and Senauer *et al.*, 1988, for the Philippines; Chen, Huq and D'Souza, 1981, for Bangladesh; Chernikovskiy *et al.*, 1983, for India) although part of these differences can be ascribed to different activity levels (Pitt, Rosenzweig and Hassan, 1989, using data from Bangladesh). Alderman and Gertler (1989) report the income and price elasticities of the demand for health care are larger for girls than boys in Pakistan. Finally, Subramanian and Deaton (1990) argue there is evidence in Indian NSS data that parents make more room in their household expenditures for boys rather than girls.

<sup>3</sup>There is little evidence other than in Asia for gender differences in infant and child mortality outcomes; gender differences in levels of anthropometric outcomes are small and often not significant; see, for example, Strauss, (1988), and Svedberg, (1990), on Africa and Schofield, (1979), on Latin America. In many countries, school enrolment ratios are higher for boys: Schultz (1987) argues that gender bias in schooling enrollments and attainments tends to decline with income. Psacharopoulos and Arriagada (1989) present evidence for discrimination against boys in school attendance and performance in Brazil; Chernichovsky (1985) argues there is discrimination against girls in school attendance in Botswana. 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). See Behrman (1990) for a comprehensive review.

based on anthropometric data typically compare the mean standardized *levels* of boys with girls, where the standards take account of differential growth patterns over age. Since these growth patterns also distinguish boys from girls, gender differences can only be identified *relative* to the standards.<sup>4</sup> Rather than follow this approach, we will examine the *determinants* of child anthropometric outcomes and test for differential impacts of mother's and father's education on the heights of sons and daughters. Our results will be suggestive of how resources are allocated within households.

The data are drawn from four household surveys carried out in the United States, Africa and Latin America. It turns out that in a variety of cultural and economic settings, the effect of a mother's education is larger on the health of her daughters, than her sons, and that the education of the father has a bigger impact on the health of his sons than his daughters. There are, then, gender differences in the allocation of household resources to child health and these differences vary with the gender of the parent.

This may simply reflect differences in the technology of child rearing; for example, it may be efficient for mothers to spend more time with daughters and fathers with sons. It turns out, however, that relative to other women, the education of a woman who is better educated than her husband has a bigger impact on the health of her daughters. Unearned income in the hands of women has a positive effect on daughter's height but no effect on sons. If unearned income and education differences between husbands and wives are indicative of power in the household bargaining game, then this additional evidence suggests that the gender differences reflect *both* differences in the technology of child rearing and differences in the preferences of parents.

Public policy can change the balance of power within the household. Government transfer programs may be directed at particular individuals within the household. Development programs and public investments may favor one gender over another (see, for example, Boserup 1970). Our results suggest that the impact on child health of a change in household resource allocation mechanisms is unlikely to be gender neutral.

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<sup>4</sup>These are usually the National Center for Health Statistics standards for a well-nourished child in the United States (NCHS, 1976). Sample based standardizations obviously preclude even these simple sorts of tests for gender bias.

## 2. THEORETICAL MODEL

The model motivating this study is in the tradition of Becker's (1981) model of household utility maximization but extended to permit household members to have different preferences (Manser and Brown, 1980; McElroy and Horney, 1981). Essentially, assume each household member,  $m=1...M$ , wishes to maximize his own utility given by:

$$v_m(x, \ell, \theta; \mu, \varepsilon) \quad [1]$$

where  $x$  is a vector of commodity demands of each individual in the household,  $\ell$  is a vector of leisure of each individual and  $\theta$  is a vector of home produced goods. This paper will focus on one element of  $\theta$ , namely an indicator of child health status.

$\mu$  represents the background characteristics, such as education, of all household members and  $\varepsilon$  is a vector representing the unobserved tastes of household members. If all elements of consumption, leisure, health and tastes, other than those relating to individual  $m$ , are given zero weight in the utility function, then that individual has egotistical preferences. Evidence suggests, however, that household members do value the consumption of others and, furthermore, household members make many joint decisions.<sup>5</sup> Let household welfare be given by the maximum of a Bergson-Samuelson type social welfare function, the arguments of which are each member's preference functions:

$$U(v_1, v_2, \dots, v_M) \quad [2]$$

If the household chooses an optimal bundle through some bargaining mechanism, then resource allocations are affected both by members' preferences and their power to assert their preferences. The latter is related to the individual's threat point utility: the utility enjoyed upon quitting the household.

If all household members have the same preferences then the household welfare maximization problem is observationally equivalent to [1]; alternatively, if one member,  $m$ , dictates all resource allocations then a zero weight is attached to all other felicity functions in [2] and the household optimization problem is, once again, [1]. In the empirical model estimated below, the two assumptions are observationally equivalent and we will refer to this restricted model as the *common preference* (or *dictatorial*) model. These quite severe restrictions have

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<sup>5</sup>In fact, the usual survey definition of a household (that is, eating from the same kitchen) may not be appropriate. See, for example, Altonji *et al.*, (1989), for a study of the extended family in the United States and Udry (1990) which suggests there is significant risk sharing among villagers in northern Nigeria.

been imposed in virtually every economic model of the household optimization problem; see Folbre (1986) and Fapohunda and Todaro (1988) for a critical discussion.

In only a small number of cases, the restrictions of the common preference model have been tested. Ashenfelter and Heckman (1974) reported one of the first tests of Slutsky symmetry for labor force participation of men and women using aggregate data from the United States; McElroy and Horney (1980) performed similar tests with individual survey data. In both cases, the common preference assumption is not rejected. In contrast, recent work with household survey data from two poor countries, indicates that in those cases the common preferences assumption is violated by the data (Schultz, 1990; Thomas, 1990).

Whatever the mechanism, household welfare [2] is maximized subject to a budget constraint<sup>6</sup> and a production function for each element of the home production vector,  $\theta$ . In the case of child health, it depends on a set of inputs,  $M$ , (such as nutrient intake and the quantity and quality of child care), individual characteristics,  $\mu_i$ , (including age and sex), family characteristics,  $\mu_f$ , (such as parental human capital) and community characteristics,  $\mu_c$ , (such as the healthiness of the environment):

$$\theta_i = \theta_i(M, \mu_i, \mu_f, \mu_c, \eta_i) \quad [3]$$

where  $\eta_i$  represents individual-specific unobserved heterogeneity in health. The technology underlying the health production function may vary with the gender of the child and, also, the impact of maternal and paternal characteristics (including time inputs) may differ. For example, if mothers spend more time with daughters than sons, and fathers spend more time with sons than daughters, then the impact of parental characteristics on child health may differ depending on the gender of the child and the parent. If it is more efficient for fathers to spend time with sons, say, then these differences may simply reflect the technology of health production and have nothing to do with parental preferences.

As an outcome of this household optimization program, there is a household demand for each element of consumption, leisure and home production. The focus here is on one element of  $\theta$ , namely child height,  $h$ . Among nutritionists, it is considered a long run measure of nutritional status (Waterlow *et al.*, 1977); both nutritionists and economists have suggested that height is also a useful indicator of child health and welfare. In

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<sup>6</sup>Each individual may face his own budget constraint with credit transactions taking place between household members. Some elements of the vector  $x$  are public goods and need to be appropriately evaluated in the budget constraint. Neither of these complications are important for this study and so they are ignored.

this model, child height depends on exogenous child, family and community characteristics,  $\mu_i$ ,  $\mu_f$  and  $\mu_c$  respectively:

$$h_i = h_i(\mu_i, \mu_f, \mu_c, \xi_i) \quad [4]$$

where  $\xi_i$  represents individual specific heterogeneity, part of which may be common across individuals within a household because of, for example, family specific health variation.<sup>7</sup>

The demand for height functions, [4], estimated below will depend on the child's gender and age,  $\mu_i$ , parental education, height and non-labor income,  $\mu_f$ . To control, albeit crudely, for community heterogeneity,  $\mu_c$ , the functions will be estimated separately for rural and urban sectors or, in one case, the race of the child.

We are interested in the effect of parental education on child height and, in particular, whether maternal (and paternal) education has the same impact on the height of a daughter and a son. Different effects can arise for two reasons. Firstly, mothers (and fathers) may have different preferences with respect to the health of their daughters and sons as embodied in [1]. Secondly, as discussed above, the technology underlying the health production function, [3], may differ for boys and girls: father's time may have a bigger impact on a son and mother's time on a daughter. It is not possible from estimates of the impact of parental education on child height in the reduced form [4] to distinguish these hypotheses. If a mother's education has a bigger impact on the height of her daughter, this simply indicates that she allocates more resources to the health of girls than boys.

We will, however, also attempt to disentangle the role of technology from preferences in two ways. Firstly, education is highly correlated with potential earnings and so, *ceteris paribus*, a better educated spouse is likely to have more power in the determination of household resource allocations. In game-theoretic terms, a better educated spouse will have a higher threat point utility. For one survey, we will compare the estimated education effects of women who are better educated than their husbands with the estimated effects for all other women.

A small number of studies have considered the role of a husband's and wife's earnings on household resource allocations. In the economic model of the household, however, the allocations of time to market work,

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<sup>7</sup>It has been demonstrated that parental education and, to a less extent, household income, often have a significant positive impact on child height for age even after controlling for genetic endowment; see Horton (1986), Behrman and Deolalikar (1990), Barrera (1989), Behrman (1988), Thomas, Strauss and Henriques (1990). For reviews of the determinants of child anthropometric indicators, see Cochrane, Leslie and O'Hara (1982), Martorell and Habicht (1986), Behrman and Deolalikar (1988) and Behrman (1990).

home production and child care are jointly determined and so labor earnings are properly treated as endogenous. We will assume non-labor income is exogenous<sup>8</sup> and since it is likely to affect threat point utility, it may be an indicator of power in the household game. We will test for differential effects of unearned income in the hands of mothers and fathers on sons and daughters. We will also compare the evidence for gender preference in Brazil in 1974 with 1986, a period during which women's opportunities outside the home have increased substantially.

These are more subtle tests for gender bias in resource allocations than simply comparing mean (standardized) heights of boys and girls or, analogously in the multivariate regression context, testing for the significance of a gender dummy in [4]. Whereas the conventional tests can only identify gender bias relative to some (arbitrarily chosen) standards, the tests proposed here are independent of standards (the effects of which are captured in gender and age dummies). Instead, we look at interactions between family characteristics and child gender. The results will have implications for how household allocations are made and also how changes in the economic environment are likely to affect the balance between investments in sons and daughters. A campaign to raise the education level of women, for example, may have significantly different implications for the healthiness of the next generation of girls relative to boys.

Using data on another indicator of child health, weight for age, Bhuiya *et al.* (1986) find that in Matlab, Bangladesh, mother's education has a significant positive effect on son's weight but not on daughter's. King and Lillard (1987) report that among Malays in Malaysia, mother's and father's education have a significant effect on daughter's schooling attainment, but not on sons. Among the Chinese in Malaysia, mother's education has a positive effect on both but father's education affects only son's schooling attainment. King and Bellew (1989) report that, in Peru, both parents' education significantly raises the probability a child (aged 8 through 19) attends school. Maternal education has a bigger effect on the probability a daughter is at school, relative to a son; paternal education has a bigger impact on the son. Desai *et al.* (1989) find that a mother's education significantly affects her daughter's intellectual ability (measured by the Peabody Picture Vocabulary Test) but not her son's ability. However, none of Bhuiya *et al.*, King and Lillard, King and Bellew or Desai *et al.* reports whether the effects of maternal (or paternal) education are significantly different on girls relative to boys. In urban Brazil,

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<sup>8</sup>This is a strong assumption since current unearned income reflects, in part, past labor supply. Since our analysis will be restricted to families with young children, most of the parents are early in the life cycle and so their unearned income is likely to be more than just saved past earnings.

Thomas (1990) finds that non-labor income attributed to mothers has a (significantly) bigger effect on the weight for height of daughters, relative to sons, and that unearned income in the hands of fathers has a bigger impact on his sons' weight for height. A similar pattern is reported for the impact of parental education on height for age. These results, together with those reported for Malaysia, Peru and the United States, suggest there may be gender preference within the household with mothers devoting more resources towards their daughters and fathers towards their sons.

### 3. EMPIRICAL RESULTS

There are a limited number of surveys which include information on both child anthropometry and household socio-economic characteristics. Four household surveys will be analyzed in this paper: one each from the United States and Ghana and two from Brazil. All of the analyses draw on the theoretical model outlined above and estimate the demand for height function, [4], following a similar empirical strategy. Heights are converted to age and gender specific z-scores using the NCHS (1976) standards for a population of well nourished children in the United States: these z-scores are used in all the analyses below. Since each survey involves its own peculiar empirical issues, the data and results are described separately in the next four subsections.

#### *(a) National Longitudinal Survey : Mother and Child Sample, 1986.*

In the United States, information has been collected on the labor market, education, training and marital experiences of young men and women since 1979, on an annual basis, as part of the National Longitudinal Survey (NLS) Youth Cohort. In 1986, a series of Child Assessment instruments were completed for almost 5,000 children of women in the NLS Youth cohort. These instruments included a set of cognitive, socio-emotional and physiological tests, information from the mother on the child's health, as well as child anthropometric measures. Analysis is restricted to information on 4,704 children less than twelve years of age.<sup>9</sup> These children are not drawn from a random sample of the United States population: 38% of the children in the sample lived in families with income below the poverty line in 1986.<sup>10</sup> Relative to the average United States child, those in the sample tend to have younger mothers (with an average age of 25 years) and more minority mothers (slightly less than 50% are white, 32% are black and the rest are hispanic). Their mothers are also less well educated: 60% completed high school and only 18% proceeded to further education.

Mean standardized z-scores of child height are reported in Table 1; a z-score of zero implies that the child's height is equal to the median height of a well-nourished child. Girls tend to be slightly taller than boys, relative to the standards, but none of these differences is significant. The average white and hispanic child is the

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<sup>9</sup>This is the same data source used in Desai *et al.* (1989) who restricted attention to 4 year old children.

<sup>10</sup>For details on the sampling frame, attrition and interviews, see Baker and Mott (1989).

**Table 1**

**Heights of male and female children:  
Tests for differences in z-scores**

	# of children	per cent stunted	Mean child height female	Mean child height male	z-scores difference
<b>United States</b>					
<i>National Longitudinal Survey: Mother &amp; Child sample, 1986</i>					
(All children < 12 years)					
All children	4704	3.4	0.127 [0.04]	0.033 [0.04]	0.094 [0.05]
Blacks	1513	4.2	0.254 [0.07]	0.190 [0.08]	0.064 [0.11]
Hispanics	856	13.1	0.018 [0.09]	-0.102 [0.08]	0.120 [0.12]
Whites	2335	3.3	0.083 [0.05]	-0.015 [0.06]	0.098 [0.07]
<b>Ghana</b>					
<i>Living Standards Survey, 1988-9</i>					
(Children < 12 years)					
Urban	990	15.7	-0.975 [0.09]	-1.209 [0.07]	0.234 [0.11]
<b>Brazil</b>					
<i>Household expenditure survey (ENDEF), 1974-5</i>					
(Children < 8 years)					
Northeast urban	9266	24.8	-1.321 [0.02]	-1.473 [0.02]	0.152 [0.03]
(Children 3 - 60 months)	6176	22.8	-1.386 [0.03]	-1.538 [0.03]	0.152 [0.04]
Northeast urban					
<i>Demographic and Health Survey, 1986</i>					
(Children 3 - 60 months)					
Northeast urban	647	11.6	-0.959 [0.08]	-1.203 [0.07]	0.244 [0.10]
Northeast rural	659	22.0	-1.321 [0.02]	-1.473 [0.02]	0.152 [0.03]

**Notes**

% stunted is proportion below 90% of median height of well nourished child of same age and sex (based on NCHS standards). z-scores computed with same standards. [Standard errors in parentheses.]

same height as the median child in the NCHS standards; blacks, however, are significantly taller, reflecting genetic differences. About 3% of the sample children are stunted (that is, their heights, given age and sex, are less than 90% of the median height of a US child; Waterlow *et al.*, 1977). The distribution of heights of hispanics is more fat tailed: 13% of hispanic children are stunted.

Many poor, young women have children out of wedlock; in the NLS Mother Child sample, mothers of one third of all children had never been married by the first survey date after the child's birth. Furthermore, a substantial number of women have changed partners since the birth of the child. Although the child assessment instruments have been employed only once, the NLS is a panel survey and so it is possible to determine the education of the mother's partner, or husband, in each year of the survey. Among those women who ever report a partner, his level of education changes for 34% of all women and 55% of black women. This information will be exploited in order to separately study the impact on child height of the education of the *father*,<sup>11</sup> the *current partner* (in 1986) and the average education of all partners or spouses present in the household since the child's birth (weighted by the number of years the person is reported to be present); this variable will be referred to as *partners' mean education*.

The effects of mother's and her partner's education on z scores of child height are reported for all children in Table 2. Levels of parental education are represented by three dummies (with the excluded category being six years of education or less).<sup>12</sup> Asymptotic *t*-statistics are reported below the regression coefficients and  $\chi^2$  test statistics for joint significance of covariates are reported at the foot of each panel. All the test statistics are based on jackknifed estimates of the variance-covariance matrix (Efron, 1982).

Maternal education has a bigger effect on child height than her partner's education presumably because mothers spend more time in child rearing activities. More important, for this paper, is the fact that independent of the definition of partner's education, mother's education has a bigger effect on the height of her daughter than her son. In all regressions, the education effect on daughter's height is positive and significant, but the effect on

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<sup>11</sup>Defined as the man reported as the mother's husband or partner in the first year of the survey after the child's birth. For children born prior to the first interview in 1979, the husband or partner in that interview is assigned the status of father.

<sup>12</sup>In addition to parental education dummies, the regressions include the mother's height and age at menarche (which affect the health of the child at birth; see Martorell *et al.*), whether a partner or spouse was present during the year of the survey and dummies for the race, gender and age of the child.

Table 2

Effect of education on child height  
United States (NLS Mother-Child, 1986)

	Females	Males	Difference	Females	Males	Difference
<b>Mother's education</b>						
(1) if 7 -12 yrs	0.739 [2.05]	-0.084 [0.26]	-0.823 [1.70]	0.773 [2.04]	-0.096 [0.28]	-0.870 [1.71]
13-20 yrs	0.937 [2.52]	-0.026 [0.07]	-0.963 [1.88]	0.954 [2.44]	-0.088 [0.24]	-1.041 [1.93]
<b>Partners' mean education</b>						
(1) if 7 -12 yrs	-0.086 [0.63]	0.137 [0.87]	0.223 [1.07]	.	.	.
13-20 yrs	-0.183 [1.11]	0.235 [1.17]	0.418 [1.61]	.	.	.
<b>Father's education</b>						
(1) if 7 -12 yrs	.	.	.	-0.242 [0.74]	0.213 [0.50]	0.455 [0.85]
13-20 yrs	.	.	.	-0.257 [0.74]	0.467 [1.05]	0.724 [1.28]
<b><math>\chi^2</math> for joint significance of</b>						
mother's education	8.386 [0.01]	0.322 [0.43]	3.621 [0.08]	7.660 [0.01]	0.088 [0.48]	3.966 [0.07]
partners' mean education	1.349 [0.26]	1.370 [0.25]	2.623 [0.14]	.	.	.
father's education	.	.	.	0.556 [0.38]	2.993 [0.11]	2.660 [0.13]
all covariates	209.447 [0.00]	120.217 [0.00]	329.594 [0.00]	202.196 [0.00]	126.460 [0.00]	328.586 [0.00]
<hr/>						
<b>Mother's education</b>						
(1) if 7 -12 yrs	0.682 [1.75]	-0.025 [0.07]	-0.707 [1.31]	0.728 [1.86]	-0.012 [0.03]	-0.740 [1.37]
13-20 yrs	0.846 [2.12]	0.019 [0.05]	-0.827 [1.47]	0.899 [2.24]	-0.001 [0.00]	-0.899 [1.58]
<b>Father's education</b>						
(1) if 7 -12 years	.	.	.	-0.608 [1.34]	0.504 [1.52]	1.112 [1.98]
13-20 years	.	.	.	-0.675 [1.41]	0.875 [2.42]	1.550 [2.58]
<b>Current partner's education</b>						
(1) if 7 -12 yrs	0.085 [0.23]	-0.143 [0.31]	-0.227 [0.39]	0.509 [1.05]	-0.472 [1.23]	-0.981 [1.59]
13-20 yrs	0.100 [0.26]	-0.038 [0.08]	-0.139 [0.23]	0.569 [1.14]	-0.629 [1.52]	-1.198 [1.85]
<b><math>\chi^2</math> for joint significance of</b>						
mother's education	5.971 [0.03]	0.140 [0.47]	2.275 [0.16]	6.521 [0.02]	0.009 [0.50]	2.793 [0.12]
father's education	.	.	.	1.324 [0.26]	2.417 [0.15]	7.639 [0.01]
current partner's education	0.074 [0.48]	0.628 [0.37]	0.377 [0.41]	1.986 [0.19]	6.938 [0.02]	3.584 [0.08]
all covariates	207.986 [0.00]	116.616 [0.00]	324.532 [0.00]	209.453 [0.00]	136.309 [0.00]	345.688 [0.08]

**Notes**

[t statistics] in parentheses below estimated effects; [p-values] below  $\chi^2$  statistics. All statistics based on jackknifed estimates of variance-covariance matrix. Father's education is education of male spouse or partner reported in first survey after child's birth. Current partner is spouse or partner reported in 1986. Partners' mean education is mean of education of all partners/spouses reported by mother since first survey after birth of child.

son's height is essentially zero. These gender differences are significant (at a 10% size of test) in two cases: when partners' mean or father's education are included in the regression. Mothers apparently devote more resources to improving the health of their daughters than to their sons.

In contrast, both the partner's mean education and father's education have a positive effect on son's height but a negative effect on daughter's height. The effect of the education of the current partner is much like maternal education, and has a bigger impact on girls than boys. Current partners, who differ from the father, may have preferences regarding child health closer to those of the mother or at least behave as if they do. This may be because of better matching of preferences. Alternatively it may reflect that the mother has more power in asserting her preferences over child health in the absence of the biological father.

The effects of the father's and the current partner's education are significant only in the regression which includes both sets of variables; in this case father's education has a significant positive effect on his son's height and this effect is significantly larger than the effect on his daughter's height. Fathers devote more resources to improving the health of their sons than their daughters. Current partners, on the other hand, have a significantly larger effect on girls relative to boys (with a 10% size of test).

Stratification by race turns out to be informative (Table 3). The larger impact of maternal education on her daughter's height is significant only for blacks and hispanics<sup>13</sup> and the education of black mothers has an absolutely large and significantly negative impact on her son's height. Partners' mean education has a significantly bigger impact on sons, relative to daughters, only for whites.<sup>14</sup> Black and hispanic women prefer to allocate resources towards their daughters and white men towards their sons.

Regressions for black children only are reported in the lower panel of Table 3 using alternative indicators of education of the man in the home. Maternal education effects are very robust to the specification of partner's education: in all cases, she allocates more resources to her daughters. Whether or not current partner's education is included in the regressions, father's education has a large positive and significant impact on son's height although this effect is not significantly larger than it is on daughters.

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<sup>13</sup>In the case of hispanics, this is true only for all women relative to those with 0 - 6 years of education. Individually, the maternal education dummies are insignificant, but taken together they are jointly significant with a  $\chi^2$  of 5.0 and a p-value of 0.04.

<sup>14</sup>This is also true if partners' mean education is replaced by father's or current partner's education.

Table 3

**Effect of education on child height  
United States (NLS Mother-Child, 1986)**

## (a) Stratified by race of child

Effect of mother's education and mean of all partners' education

	Blacks			Hispanics			Whites		
	Females	Males	Difference	Females	Males	Difference	Females	Males	Difference
<b>Mother's education</b>									
(1) if 7 -12 yrs	2.28 [1.3]	-2.06 [4.3]	-4.35 [2.5]	0.66 [1.3]	-0.31 [1.6]	-0.98 [1.8]	0.11 [0.3]	1.56 [1.1]	1.46 [1.0]
13-20 yrs	2.43 [1.4]	-2.19 [4.3]	-4.61 [2.6]	0.79 [1.4]	0.32 [1.1]	-0.47 [0.8]	0.37 [1.0]	1.61 [1.1]	1.23 [0.8]
<b>Partners' mean education</b>									
(1) if 7 -12 yrs	0.06 [0.2]	0.06 [0.1]	0.00 [0.0]	-0.01 [0.0]	-0.07 [0.3]	-0.06 [0.2]	-0.29 [1.6]	0.27 [1.7]	0.56 [2.3]
13-20 yrs	-0.16 [0.6]	-0.05 [0.1]	0.12 [0.2]	-0.11 [0.3]	0.20 [0.7]	0.30 [0.7]	-0.40 [1.8]	0.33 [1.4]	0.73 [2.2]
<b><math>\chi^2</math> for joint significance of</b>									
mother's education	2.29 [0.2]	19.27 [0.00]	6.91 [0.02]	2.09 [0.2]	7.21 [0.01]	5.00 [0.04]	3.83 [0.07]	1.11 [0.3]	2.09 [0.2]
partners' mean education	1.17 [0.3]	0.27 [0.4]	0.15 [0.5]	0.14 [0.5]	1.43 [0.3]	1.15 [0.3]	3.38 [0.09]	3.23 [0.10]	6.43 [0.02]
all covariates	138.36 [0.00]	108.33 [0.06]	246.89 [0.00]	32.96 [0.00]	20.24 [0.05]	53.26 [0.01]	52.32 [0.00]	38.39 [0.00]	220.08 [0.00]

## (b) Black children

Effect of education of mother, father and current partner

	Education of current partner			Education of child's father			Education of current partner		
	Females	Males	Difference	Females	Males	Difference	Females	Males	Difference
<b>Mother's education</b>									
(1) if 7 -12 years	2.32 [1.4]	-2.03 [4.1]	-4.36 [2.5]	2.30 [1.4]	-2.03 [4.4]	-4.33 [2.5]	2.29 [1.4]	-1.98 [4.3]	-4.27 [2.4]
13-20 years	2.44 [1.4]	-2.14 [4.0]	-4.58 [2.6]	2.42 [1.4]	-2.20 [4.4]	-4.62 [2.6]	2.42 [1.4]	-2.13 [4.3]	-4.55 [2.6]
<b>Current partner's education</b>									
(1) if 7 -12 years	0.12 [0.2]	-0.45 [0.7]	-0.57 [0.6]	.	.	.	0.08 [0.2]	-0.40 [0.8]	-0.48 [0.6]
13-20 years	0.02 [0.0]	-0.68 [1.0]	-0.70 [0.8]	.	.	.	0.02 [0.0]	-0.77 [1.3]	-0.79 [0.9]
<b>Father's education</b>									
(1) if 7 -12 years	.	.	.	0.19 [0.0]	1.59 [4.4]	1.39 [0.4]	0.20 [0.0]	1.57 [4.1]	1.37 [0.0]
13-20 years	.	.	.	0.03 [0.0]	1.56 [3.7]	1.53 [0.5]	0.08 [0.0]	1.81 [3.5]	1.73 [0.0]
<b><math>\chi^2</math> for joint significance of</b>									
mother's education	2.22 [0.2]	17.10 [0.00]	6.68 [0.02]	2.26 [0.2]	20.36 [0.00]	7.17 [0.01]	2.24 [0.2]	19.30 [0.00]	6.84 [0.02]
current partner's education	0.22 [0.5]	1.34 [0.3]	0.59 [0.4]	.	.	.	0.07 [0.5]	1.72 [0.2]	0.91 [0.3]
father's education	.	.	.	0.43 [0.00]	19.91 [0.00]	0.31 [0.4]	0.20 [0.5]	19.00 [0.00]	0.52 [0.4]
all covariates	133.09 [0.00]	107.69 [0.00]	240.94 [0.00]	136.88 [0.00]	118.86 [0.00]	255.99 [0.08]	136.33 [0.00]	132.57 [0.00]	269.07 [0.00]

## Notes

See Table 2.

### *Robustness of results*

Interactions between mother's and her partner's education are small and insignificant. The inclusion of these interactions has no impact on the results described above. Maternal height is included in each regression to capture both phenotype and genotype influences on child height; the effect of mother's height is positive and slightly (but not significantly) larger for girls.

The NLS data also report weight and length at birth for each child. If the results for child height reflect behavioral differences, then the effect of parental education on the birthweight and length of a girl should be the same as that on a boy. In all cases, this turns out to be true. In fact, in a regression of *growth* in height (current height less birth length), the estimate effects of parental education are almost identical to those described above although the coefficients of interest are estimated slightly more precisely.

### *Statistical issues*

The sample includes 2,446 women, of whom just over half have more than one child under twelve; regression errors,  $\xi_i$ , which incorporate unobserved family specific health heterogeneity, may not be independent across observations. The estimation of family fixed effects make little sense in this study. Firstly, it will induce a selected sample removing all households with one child less than twelve and there are no valid instruments available to control for this selectivity. Secondly, it will not be possible to determine how maternal education affects child height. Random effects estimates, however, exploit the efficiency gains associated with the structure of the errors. Corresponding with the results in panel a of Table 3, random effects estimates are reported in panel a of Table A.1. A Lagrange Multiplier test for the presence of error components (Breusch and Pagan, 1980) is significantly different from zero only for whites in the fully interacted model; a Hausman test for correlation between error components and regressors is rejected.<sup>15</sup> In all regressions, the differences between the ordinary least squares and random effects estimates are very small. Since this is true in all the models reported in this paper, we report only least squares estimates. The assumption that the regression errors are

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<sup>15</sup>For white children, the LM test is 7.9 (with a critical value of 3.84); the Hausman test statistic is 38.4 which is less than the critical value of 53.4 for 38 covariates.

homoskedastic is, however, unambiguously rejected.<sup>16</sup> A comparison of the jackknife, White and ordinary least squares estimators of the variance-covariance matrix is discussed in the appendix; in view of that discussion, all test statistics in this paper are based on the jackknife.

*(b) Ghana Living Standards Survey, 1988-9*

The Ghana Living Standards Survey (GLSS), which was carried out by the Government of Ghana, in collaboration with The World Bank, collected very detailed information on a nationwide random sample of about 3,200 households. The Ghana LSS is part of a program being implemented in several developing countries which, using common instruments, collects extensive socio-economic information from households on expenditure, income and labor supply, and assets. Demographic data on each individual in the household includes age, sex, education and anthropometric measures.<sup>17</sup>

The survey is an ongoing effort with a panel dimension since half the households are enumerated for two years; we will focus on urban households enumerated during the first year of the survey, 1987-88. For an African society, women are unusually active in the Ghanaian economy, especially in trading and extra-domestic activities: they thus wield a good deal of power in the household (Peil, 1977; Hill, 1975; Bleek, 1987). Husbands and wives seldom pool their resources but "prefer to keep their money separate in order to avoid conflicts" (Oppong, 1982; see also Guyer, 1988). Women are typically responsible for child rearing as well as food production and distribution: they are likely, therefore, to have a large influence on the nutritional status of their children. Fathers, on the other hand, tend to be peripheral figures in the lives of their children (Oppong and Bleek, 1922; Nsamenang, 1987); this reflects, in part, the fact that Ghana is a matrilineal society, at least in the south.

Fathers are, however, usually responsible for their children's schooling (Bleek, 1982; Ghana Teaching Service, 1975): investments in the education of women are low (Schultz, 1989). Almost half the women in our

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<sup>16</sup>Lagrange Multiplier test statistics for heteroskedasticity (Breusch and Pagan, 1980) are about 1,000 for each regression in Table 2. The  $\chi^2$  are all over 400 for the regressions in panel b of Table 3 and for panel a the test statistics are 182, 458 and 1,245 for blacks, hispanics and whites respectively. The  $\chi^2$ s have between 30 and 40 degrees of freedom: clearly homoskedasticity is rejected in all models.

<sup>17</sup>Data from the survey have been tabulated in Ghana Statistical Service (1989). The anthropometric and health data are described in Alderman, (1989).

sample have received no formal education, about 12% received some primary education and 40% went beyond primary school. Historically, men have received much more education; in this sample, 65% have more than primary education and only 30% have no education at all.

As in the United States sample discussed above, the Ghanaian sample is restricted to children aged less than twelve years, of whom there are 990 living in 412 households. In Ghana, both boys and girls are significantly shorter, on average, than a well nourished United States child of the same age; relative to the US standards, girls are taller than boys, and this difference is also significant (Table 1). About 16.2% of children in urban Ghana are stunted (have heights below 90% of the US median).

Regression results are reported in Table 4.<sup>18</sup> Maternal education has a significant positive effect on the height of daughters, which declines with education and this effect is significantly greater than the effect on sons. Paternal education has a significant negative effect on daughters' heights and, for fathers with primary education, also on sons' heights although the latter effect is absolutely smaller. Fathers with more than primary education have a significantly greater effect on their sons' height relative to daughters' and, taking both the father education covariates together, the impact is significantly larger on sons (with a p-value of 0.08).

In a matrilineal society, children inherit not from their fathers but their mother's brothers. Fathers may, therefore, divert resources away from their own children to their nephews. Unfortunately no data are available on members of the extended family to test this hypothesis.

If education is indicative of power in the household bargaining game, then women who are better educated than their husbands should be able to assert their preferences and direct more resources towards commodities they care about. In the lower panel of Table 4, a dummy variable is included to identify these women (who account for 8% of the sample). Relative to other mothers, the education of a woman who is better educated than her husband has a large and significant (at 10%) effect on her daughter's height, no effect on her sons height and this difference is significant. This does not simply reflect non-linearities in maternal education

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<sup>18</sup>No education is the excluded category. The restriction that primary and secondary education have the same impact on child height is not rejected; they are therefore grouped together. In addition to parental education, the regressions include maternal and paternal heights as well as dummies for the child's age and sex. Since data are available on both parents' height and they are significant determinants of child height, the sample has been restricted to children with both parents present in the household. This restriction also permits an examination of the effect on child height of the differential between the educational attainment of mothers and fathers. If paternal height is excluded (so that female headed households are included) then the estimated education effects are smaller in magnitude. All of this difference can be explained by the change in the sample composition. If the larger sample is used, the substance of the results reported below is unchanged.

Table 4

## Effect of parental education on child height

*Ghana Living Standards Survey, 1988-9: Urban sector*

	Females	Males	Difference
<b>Mother's education</b>			
(1) if primary	1.024 [1.86]	-0.181 [0.99]	-1.204 [2.07]
(1) if more than primary	0.595 [3.00]	0.079 [0.44]	-0.516 [1.93]
<b>Father's education</b>			
(1) if primary	-0.917 [2.49]	-0.798 [2.54]	0.119 [0.25]
(1) if more than primary	-0.549 [2.09]	0.034 [0.17]	0.583 [1.75]
$\chi^2$ for joint significance of			
mother's education	9.687 [0.00]	1.656 [0.22]	6.059 [0.02]
father's education	6.864 [0.02]	8.706 [0.01]	3.675 [0.08]
all covariates	368.722 [0.00]	511.518 [0.00]	879.355 [0.00]
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<b>Mother's education</b>			
(1) if primary	0.779 [1.63]	-0.074 [0.30]	-0.853 [1.64]
(1) if more than primary	0.450 [2.08]	0.258 [1.61]	-0.292 [1.08]
(1) if mother better educated	0.910 [1.66]	-0.234 [0.78]	-1.184 [1.87]
<b>Father's education</b>			
(1) if primary	-0.943 [2.53]	-0.790 [1.94]	0.153 [0.31]
(1) if more than primary	-0.368 [1.67]	0.044 [0.22]	0.324 [1.11]
$\chi^2$ for joint significance of			
mother's education	14.267 [0.00]	4.560 [0.10]	7.942 [0.02]
father's education	7.364 [0.03]	5.856 [0.06]	5.710 [0.06]
all covariates	340.594 [0.00]	461.549 [0.00]	852.965 [0.00]

**Notes**

[t statistics] in parentheses below estimated effects; [p-values] below  $\chi^2$  statistics. All statistics based on jackknifed estimates of variance-covariance matrix.

effects. The result is robust to the inclusion of years of education, of more dummies and both years of education and the dummies.

In the survey, 60% of mothers have more than one child under twelve; mother-specific random effects estimates are almost identical to those in Table 4.<sup>19</sup> Since marriages in Ghana are sometimes polygenous, 12% of women in the sample live in the same household; household-specific random effects are also essentially identical to the OLS estimates.<sup>20</sup>

In rural Ghana, children are significantly shorter than their urban counterparts: 28% are stunted. Less than half the mothers have any education and her education has no impact on child height. Paternal education has a significant negative effect and this effect is (absolutely) larger for daughters. For both parents the interaction between education and the gender of the child is not significant.

### (c) *Brazilian household surveys 1974-5 and 1986*

Two household surveys from Brazil will be examined. The first survey, the *Estudo Nacional da Despesa Familiar* (ENDEF), is a random national sample of nearly 55,000 households carried out by the Brazilian statistical agency, IBGE, in 1974-5. It is very comprehensive by large scale expenditure survey standards with data collected on household expenditures and food consumption; in addition to the usual socio-demographic information, each individual reports labor supply and earnings as well as nonlabor income. The height and weight of all members of the household were measured by an anthropometrist.<sup>21</sup> The second survey was collected in 1986 as part of the *Demographic and Health Survey* (DHS) program which has been implemented in about fifty countries worldwide and focusses on fertility, contraception, maternal and child health.

Brazil, like much of Latin America, is a *machista* society (Neuhouser, 1989) although sociologists argue that household allocation decisions are complex and women have a good deal of control over food expenditures and the distribution of food (Goldani, 1989; Neuhouser, 1989). As in Ghana, women are likely to have a big influence on their children's nutrition. During the twelve years between these two surveys, Brazilian society has

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<sup>19</sup>The hypothesis of no group variance is rejected with an LM test statistic of 4.11 but the Hausman specification test statistic (28.67) is less than the critical value.

<sup>20</sup>The LM test statistic for no group variance is 5.596 and the Hausman test is 28.596.

<sup>21</sup>The data are presented in tabulations by the Brazilian statistical agency, IBGE (1982).

seen dramatic changes. Female labor force participation rates have risen by about 20% to 37% in 1986 (IBGE, 1987), while total fertility rates have declined from above 6 in 1974 to about 3.5 in 1986 (National Research Council, 1983). The influence of the Catholic church has been eroded with a rise in the popularity of civil marriages, divorce rates and of consensual unions, especially among young adults. (Goldani, 1989; Henriques, 1988). All these changes suggest women have more opportunities outside the home and so their threat points will have risen during this period: a bargaining model of household resource allocations would suggest women have a bigger influence on outcomes in the later survey.

Anthropometric data in the DHS were recorded for children aged 3 to 60 months living in the Northeast, the poorest region of Brazil.<sup>22</sup> There are 1316 children in this sample. To facilitate comparisons, the ENDEF sample will also be restricted to the Northeast; results for both urban and rural children will be discussed although we will focus on urban households. Since the heights of all household members were recorded in ENDEF, the sample will be restricted to 9,266 children less than eight years old, both of whose parents are present in the household.

Although Brazil has invested relatively little in education given its level of income (Schultz, 1987) there is no apparent difference in enrolment rates by sex. In the 1974-5 ENDEF sample, almost 30% of mothers are illiterate, 16% have completed (four years of) elementary school and 9% have more education; fathers are very slightly better educated.<sup>23</sup> According to the 1980 Brazilian census, women have recently overtaken men in levels of education and this is reflected in the 1986 DHS: urban women have 4.8 years of education on average, their partners have 4.6 years; in the rural sector, the average woman has spent only 1.9 years at school, her partner 2.1 years.

The average child in the Northeast of Brazil is substantially shorter than a well-nourished child in the United States but slightly taller than a Ghanaian child. There have been dramatic improvements in the nutritional status of Brazilian children over the last fifteen years. The proportion of 3 to 60 month old urban children who were stunted halved between 1974 and 1986 (to 11%). Rural children are significantly shorter; 38% were stunted in 1974 and 22% remained stunted in 1986.

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<sup>22</sup>Tabulations of these data are presented in BEMFAM-DHS, (1989).

<sup>23</sup>Education in this survey is reported in discrete intervals only; respondents who claimed to be able to read and write but had less than four years of education were deemed 'literate'.

**Table 5**  
**Effect of parental education on child height**  
**Brazil, ENDEF 1974-5 : Urban Northeast**

	Females	Males	Difference
<b>Mother's education</b>			
(1) if literate	0.344 [6.60]	0.217 [3.93]	-0.127 [1.67]
(1) if completed elementary school	0.445 [6.17]	0.334 [4.09]	-0.111 [1.01]
(1) if completed secondary school or higher	0.835 [8.10]	0.572 [5.38]	-0.264 [1.78]
<b>Father's education</b>			
(1) if literate	0.189 [3.54]	0.268 [4.80]	0.079 [1.02]
(1) if completed elementary school	0.399 [5.39]	0.595 [7.24]	0.196 [1.78]
(1) if completed secondary school or higher	0.666 [6.96]	1.132 [11.12]	0.466 [3.34]
<b><math>\chi^2</math> for joint significance of</b>			
mother's education	80.117 [0.00]	33.139 [0.00]	4.357 [0.11]
father's education	52.984 [0.00]	128.040 [0.00]	11.475 [0.01]
all covariates	6242.888 [0.00]	6791.399 [0.00]	13032.880 [0.00]
<b>Mother's nonlabor income (Cr\$ 000000)</b>			
	0.232 [2.72]	0.002 [0.02]	-0.231 [1.80]
<b>Father's nonlabor income (Cr\$ 000000)</b>			
	0.039 [1.75]	0.029 [1.02]	-0.010 [0.26]
<b>Mother's education</b>			
(1) if literate	0.341 [6.54]	0.216 [3.90]	-0.125 [1.64]
(1) if completed elementary school	0.434 [5.99]	0.330 [4.03]	-0.104 [0.95]
(1) if completed secondary school or higher	0.803 [7.81]	0.556 [5.25]	-0.247 [1.67]
<b>Father's education</b>			
(1) if literate	0.189 [3.53]	0.267 [4.80]	0.079 [1.02]
(1) if completed elementary school	0.396 [5.36]	0.591 [7.20]	0.195 [1.76]
(1) if completed secondary school or higher	0.649 [6.77]	1.122 [10.85]	0.473 [3.36]
<b><math>\chi^2</math> for joint significance of</b>			
mother's education	75.744 [0.00]	31.863 [0.00]	4.019 [0.13]
father's education	50.579 [0.00]	122.487 [0.00]	11.589 [0.01]
all covariates	6275.592 [0.00]	6800.600 [0.00]	13074.781 [0.00]

**Notes**

[t statistics] in parentheses below estimated effects; [p-values] below  $\chi^2$  statistics. All statistics based on jackknifed estimates of variance-covariance matrix.

In both Brazilian surveys and in the Ghanaian survey, girls are significantly taller than boys, relative to a well nourished child in the United States. While it may be that there is gender bias in Ghana and Brazil, it seems likely that growth paths there are systematically different from those in the United States. Comparisons of levels of standardized anthropometric measures seems to be a weak basis for statements about the presence or absence of gender bias in a society.

*(i) Estudo Nacional da Despesa Familiar, 1974-5*

In the urban ENDEF sample (Table 5), both parents' education have a significant positive effect on child height. The impact of mother's education is bigger on girls although the significance of this difference is marginal. Father's education has a bigger effect on boys and this difference is clearly significant (taken jointly); for boys, the effect of paternal education is significantly larger than the impact of maternal education.

In this sample, 35% of fathers and 14% of mothers report any unearned income.<sup>24</sup> Each parent's unearned income is added to the regression in the lower panel of Table 5; the parental education coefficients hardly change. Father's nonlabor income has no impact on child height. Mother's unearned income, however, has a positive and significant impact on her daughter's height and a negative (but insignificant) effect on her son's height. This difference is also significant (at a 10% size of test). Women apparently prefer to allocate resources to improving the health of their daughters.

In the rural sector, parental education is a significant determinant of child height. None of the interactions between education and gender of the child is significant.

*(ii) Demographic and Health Survey, 1986*

Table 6 reports the regression results for the 1986 DHS sample. In the urban sector, maternal education has a significant positive effect on her daughter's height and this is significantly larger than the positive effect on her son's height. Father's education also has a positive effect on his children's heights but the effect is significant only for sons; it is (marginally) significantly bigger for sons than daughters. Removing the restriction that the impact of parental education be linear does not affect any of the results except that the differential effect of father's education on sons and daughters is no longer significant.

In the rural sector, child height is unresponsive to father's education. Maternal education has a significantly positive effect on both sons and daughters, which is larger on sons, but this difference is not

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<sup>24</sup>Unearned income includes income from financial and physical assets, gifts and income from pensions, social security and workers compensation. For a fuller description of the data, see Thomas, (1990).

Table 6

## Effect of parental education on child height

## Brazilian Demographic Health Survey, 1986 : Northeast Region

	Females	Males	Difference	Females	Males	Difference
<i>Urban</i>						
Mother's education	1.165	0.250	-0.915	2.183	0.776	-1.407
(in years)	[4.08]	[0.98]	[2.39]	[3.22]	[1.15]	[1.47]
-"- squared				-0.079	-0.044	0.034
				[1.78]	[0.95]	[0.53]
Partner's education	0.150	0.797	0.647	0.193	1.222	0.103
(in years)	[0.57]	[2.99]	[1.73]	[0.30]	[1.88]	[1.12]
-"- squared				0.000	-0.033	0.033
				[0.01]	[0.61]	[0.49]
$\chi^2$ for joint significance of						
Mother's education				18.169	1.41	6.611
				[0.00]	[0.25]	[0.02]
Partner's education				0.561	11.60	2.837
				[0.38]	[0.00]	[0.12]
All covariates						
<hr/>						
<i>Rural</i>						
Mother's education	0.987	1.312	0.325	1.901	-0.169	-2.070
(in years)	[2.67]	[3.28]	[0.60]	[2.50]	[0.22]	[1.91]
-"- squared				-0.128	0.159	0.287
				[1.48]	[1.90]	[2.38]
Partner's education	0.095	0.565	0.470	-1.125	0.497	1.623
(in years)	[0.23]	[1.45]	[0.82]	[1.26]	[0.51]	[1.23]
-"- squared				0.169	-0.038	-0.208
				[1.58]	[0.38]	[1.41]
$\chi^2$ for joint significance of						
Mother's education				7.485	11.396	5.724
				[0.01]	[0.00]	[0.03]
Partner's education				2.541	0.313	2.003
				[0.14]	[0.43]	[0.18]
All covariates						
<hr/>						
Notes	See Table 4.					

significant in the linear model. Permitting quadratics changes the results dramatically. Maternal education effects on both sons and daughters remain significant; for daughters, there are decreasing marginal returns to additional maternal education (with the effect turning negative when education exceeds 7.4 years); for sons, there are increasing marginal returns to education. The difference, however, is significant overall and positive for the vast majority of mothers (namely the 83% who have less than 3.6 years of education).

The evidence from both Brazilian surveys tell a similar story. Mothers devote more resources to the health of their daughters and fathers prefer to allocate resources to their sons. The differential impact of maternal education on sons and daughters turned from being insignificant in 1974 to significant in 1986. This is consistent with women having more power in 1986 and thus being able to assert their preferences as their opportunities outside of the home improved.

#### 4. DISCUSSION

Our evidence indicates that mothers and fathers invest different amounts of resources in the human capital of the children. Mothers allocate more resources to their daughters and fathers channel resources towards their sons. This is true in the United States, Brazil and Ghana.

This surely reflects, at least in part, the technology of child rearing which Durkheim (1893) called the sexual division of labor. Over the last two decades there has been a good deal of research by psychologists indicating that fathers play a bigger role in the development of their sons than their daughters. Fathers spend more time with their sons and sons show preference for their fathers at an early age (at least from the second year of life) (Lamb, 1976; Morgan, Lye and Condran, 1988). Mothers, on the other hand, tend to spend more time with their daughters and have a closer relationship with a daughter than a son. For a synthesis of this research, see Lamb (1976, 1987). Longitudinal data on child development indicates that the absence of a father (because of divorce) has a more severe and enduring impact on boys than girls (Hetherington, Hetherington and Cox, 1978). Similar differences by gender are reported for the impact of divorce on child health (Mauldon, 1990).

Do gender specific investments by mothers and fathers reflect only technological differences? In Brazil, women devote unearned income towards improving the health of their daughters but not their sons. The difference between the maternal education effects on girls and boys turned from being insignificant in the first survey (in 1974) to being significant in the second (in 1986); during the same period, women's opportunities outside of the home probably improved substantially. In Ghana, if a woman is better educated than her husband, then her daughter benefits more and her son benefits less from her education than if the husband is better educated than his wife. If unearned income, opportunities outside the home and relative educational status are indicative of power in the household bargaining game, then these results suggest that more powerful women are able to assert their own preferences in the allocation of household resources.

There is other evidence in the literature which suggests we may be picking up differences in preferences of mothers and fathers. Sociologists and demographers have pointed out that the probability of marital dissolution is lower if a couple has a son (Spanier and Glick, 1981). It is argued that this reflects a higher price of marital dissolution to the father with sons (with whom he spends more time) (Morgan, Lye and Condran, 1988). Fertility preferences are also often gender specific. In the United States, the birth interval between first and second born is independent of gender of the first child among white women but, among black women, it is smaller if the first child is a son. Black women apparently prefer daughters (Teachman and Schollaert, 1989).

In several developing countries, the data suggest that fathers with sons are more willing to limit their family size than fathers without sons, (Caldwell and Caldwell, 1978; Mason and Taj, 1987).

Different investments in sons and daughters by mothers and fathers might also reflect different returns to these investments. Children participate in work within the house and family business from an early age in Ghana and girls tend to help mothers, boys work with fathers (Kaye, 1962). Children in Africa are also frequently a source of old age security for parents (Caldwell, 1979); this is likely to be more important for women who live longer than men. There is, as far as we can tell, no evidence on whether sons or daughters provide more assistance to mothers or fathers although in a matrilineal society, like Ghana, where women and daughters retain close ties, it seems plausible that mothers will realize higher returns from investments in daughters than sons. In the United States, Spitze and Logan (1990) report that daughters are more likely to be in contact with and assist their parents in old age and that a single mother is significantly more likely to be in contact with her children if at least one is a daughter. After the father's death, daughters tend to give more attention to their widowed mothers and sons give less (Hess and Waring, 1978).

Taken together, all this evidence suggests that in a variety of economic, social and cultural settings, mothers prefer to allocate resources towards daughters. Fathers treat their sons preferentially. As economists, we need to find richer models of household behavior.

## References

- Alderman, H. (1989). Nutritional status in Ghana and its determinants, mimeo, Cornell University Food and Nutrition Policy Program.
- Alderman, H. and P. Gertler. (1989). Family resources and gender differences in human capital, mimeo.
- Altonji, J., F. Hayashi and L. Kotlikoff. (1989). Is the extended family altruistically linked? Direct tests using micro data, mimeo.
- Ashenfelter, O. and J. Heckman. (1974). The estimation of income and substitution effects in a model of family labor supply. *Econometrica* 42.1:73-85.
- Baker, P. C. and F. L. Mott. (1989). *NLSY Child Handbook, 1989: A guide and resource document for the National Longitudinal Survey of Youth 1986 Child Data*, Center for Human Resource Research, Ohio State University.
- Barrera, A. (1990). The role of maternal schooling and its interactions with public health programs in child health production. *Journal of Development Economics*, 32:69-91.
- Becker, G. S. (1964). *Human Capital*, Columbia University Press, New York
- Becker, G. S. (1981). *A Treatise on the Family*, Harvard University Press, Cambridge, MA.
- Behrman, J. R. (1988). Intrahousehold allocation of nutrients in rural India: Are boys favored? Do parents exhibit inequality aversion?, *Oxford Economic Papers*, 40(1):32-54.
- Behrman, J. R. (1990). Intrahousehold allocation of nutrients and gender effects: A survey of structural and reduced form estimates, in *Nutrition and Poverty*, ed. S. Osmani, Oxford: Oxford University Press.
- Behrman, J. R. (1990). Interactions among human resources and poverty: What we know and what we do not know. Mimeo, University of Pennsylvania.
- Behrman, J. R. and A. Deolalikar (1988). Health and nutrition, in H. Chenery and T. N. Srinivasan (eds.). *Handbook of Development Economics*, North Holland, Amsterdam.
- Behrman, J. R. and A. Deolalikar (1989). Seasonal demands for nutrient intakes and health status in rural South India. In D. Sahn (ed.) *Causes and Implications of Seasonal Variability in Household Food Security*, Washington, D. C.:Johns Hopkins.
- Behrman, J. R. and A. Deolalikar (1990). The intrahousehold demand for nutrients in rural South India: individual estimates, fixed effects and permanent income, forthcoming, *Journal of Human Resources*.
- Belsley, D. A., E. Kuh, and R. E. Welsch. (1980). *Regressions Diagnostics : Identifying influential data and sources of collinearity*. New York:Wiley.
- BEMFAM-IRD. (1988). *Brasil. Pesquisa sobre saude materno-infantil e planejamento familiar*.
- Bhuyia, A., B. Yojtniak, S. D'Souza and S. Zimicki. (1986). Socio-economic determinants of child nutritional status: boys versus girls, *Food and Nutrition Bulletin*, 8.
- Bjorn, P. A. and Q. H. Vuong (1984). Simultaneous equations models for dummy endogenous variables: a game theoretic formulation with an application to labor force participation, working paper, California Institute of Technology.
- Bjorn, P. A. and Q. H. Vuong (1985). Econometric modelling of a Stackelberg game with an application to labor force participation, working paper, California Institute of Technology.
- Bleck, W. (1987). Family and family planning in Southern Ghana. In C. Oppong (ed.), *Sex Roles, Population and Development in West Africa*, 138-53. London:James Currey.

- Blumberg, R. (1988). Income under female versus male control: Hypotheses from a theory of gender stratification and data from the Third World, *Journal of Family Issues*, 9.1:51-84.
- Boserup, E. (1970). *Women's role in economic development*, New York:St Martin's Press.
- Breusch T. S. and A. R. Pagan. (1980). The lagrange multiplier test and its applications to model specification in econometrics, *Review of Economic Studies*, 47:225-38.
- Caldwell, J. C. (1977). The economic rationality of high fertility: An investigation illustrated with Nigerian data, *Population Studies*, 31.1:5-27.
- Caldwell, J. C. and P. Caldwell. (1978). The achieved small family: Early fertility transition in an African society, *Studies in Family Planning*, 9.1:2-18.
- Chen, L., E. Huq and S. D'Souza (1981). Sex bias in the family allocation of food and health care in rural Bangladesh, *Population and Development Review*, 7.1:55-70.
- Chernichovsky, D. (1985). Socioeconomic and demographic aspects of school enrollment and attendance in rural Botswana, *Economic Development and Cultural Change*, 33.2:319-332.
- Chiappori, D. A. (1988a). Rational household labor supply, *Econometrica* 56.1:63-89.
- Chiappori, D. A. (1988b). Nash bargained household decisions, *International Economic Review*, 29(4):791-6.
- Cochrane, S., J. Leslie and D. O'Hara (1982). Parental education and child health: intercountry evidence, *Health Policy and Education*, 2:213-50.
- D'Souza, S. and L Chen (1980). Sex differentials in mortality in rural Bangladesh, *Population and Development Review*, 6.2:257-70.
- Deaton, A. (1989). Looking for boy-girl discrimination in household expenditure data, *World Bank Economic Review*, 3.1:1-16.
- Desai, S., P. L. Chase-Lansdale and R. T. Michael. (1989). Mother or market? Effects of maternal employment on the intellectual ability of 4-year-old children, *Demography*, 26.4:545-63.
- Durkheim, E. (1933). *The Division of Labor in Society* translated by G. Simpson from *De la division du travail social: Etudes sur l'Organisation des Societes Superieures* (1893), London:MacMillan.
- Efron, B. (1982). *The jackknife, the bootstrap and other resampling plans*, SIAM CBMS-NSF Monograph, 38.
- Evenson, R. E., B. M. Popkin and E. K. Quizon (1980). Nutrition, work and demographic behavior in rural Philippine households, in H. Binswanger et al. (eds). *Rural Household Studies in Asia*, Singapore University Press, Singapore.
- Fapohunda, E. and M. Todaro. (1988). Family structure, implicit contracts and the demand for children in Southern Nigeria, *Population and Development Review*, 14.4:571-94.
- Folbre, N. (1986). Cleaning house, *Journal of Development Economics*, 22, 5-40.
- Gertler, P. and H. Alderman. (1989). Family resources and gender differences in human capital investments: The demand for children's medicalcare in Pakistan, mimeo, Rand Corporation.
- Ghana Statistical Service and World Bank Social Dimensions of Adjustment Project Unit. (1989). *Ghana Living Standards Survey: First Year Report September 1987 - August 1988*.
- Ghana Teaching Service. (1975). *Report on population and family life education survey in schools and colleges*. Accra:Ministry of Education.
- Goldani, A. M. (1989). Women's transitions: the intersection of female life course, family and demographic transition in Twentieth Century Brazil, unpublished Ph.D. dissertation, University of Texas at Austin.

- Gronau, R. (1985). The allocation of goods within the household and estimation of adult equivalence scales: how to separate the men from the boys, manuscript, Hebrew University.
- Guyer, J. (1988). Dynamic approaches to domestic budgeting: Cases and methods from Africa, in J. Bruce and D. Dwyer, (eds.), *A Home Divided: Women and Income in the Third World*, Stanford:Stanford University Press.
- Handschin, E., F. C. Schweppe, J. Kohlas, and A. Fiechter. (1975). Bad data analysis for power system state estimators. *IEEE Trans. Power Appar. Sys., PAS-94*, 329-337.
- Henriques, M. H. (1988). Marriage systems changes in Brazil, mimeo, Fordham University.
- Hess, B. and J. Waring. (1978). Parent and child in later life: Rethinking the relationships. In R. Lerner and G. Spanier (eds.) *Child Influences on Marital and Family Interactions: A Life-Span Perspective*, 241-73. New York:Academic Press.
- Hetherington, E. M. Cox and R. Cox. (1978). The aftermath of divorce. In J. Stevens and M. Mathews, *Mother/Child-Father/Child Relationships*, 149-76. Washington D.C.:National Association for the Education of Young Children.
- Hill, P. (1975). The West African farming household. In J. Goody, (ed.), *Changing Social Structure in Ghana*, London:International Africa Institute.
- Horton, S. (1986). Child nutrition and family size in the Philippines, *Journal of Development Economics*, 23:161-176.
- IBGE (1978). *Estudo Nacional da Despesa Familiar: Dados Preliminares*, Rio de Janeiro.
- IBGE. (1988). *Anuario Estatístico Brasil*, Rio de Janeiro.
- Jaeckel, L. A. (1972)., Estimating regression coefficients by minimizing the dispersion of residuals, *Annals of Mathematical Statistics*, 42, 1020-1034.
- Kakwani, N. (1986). Is sex bias significant?, Wider Working Paper 9, Helsinki.
- Kaye, B. (1962). *Bringing up children in Ghana*. London:Allen and Unwin.
- King, E. and R. Bellwew. (1989). Gains in the education of Peruvian women in 1940's-1980's -- Patterns and explanations. Mimeo. PHR, World Bank.
- King, E. and L. Lillard. (1987). Education policy and schooling attainment in Malaysia and the Philippines, *Economics of Education Review*, 6.2:167-81.
- Kooreman, P. (1990). Household labor force participation as a cooperative game: an empirical model, *Journal of Human Resources*, forthcoming.
- Lamb, M. E. (1976). The role of the father: an overview. In M. Lamb, (ed.), *The Role of the Father in Child Development*, New York:Wiley.
- Lamb, M. E. (1987). *The Father's Role: Cross Cultural Perspectives*, London:Lawrence Erlbaum Associates.
- MacKinnon, J., and H. White. (1982). Some heteroskedasticity consistent covariance matrix estimators with improved finite sample properties, *Journal of Econometrics*, 29, 305-325.
- Manser, M. and M. Brown (1980). Marriage and household decision making: a bargaining analysis, *International Economic Review*, 21.1:31-44.
- Martorell, R and J.-P. Habicht (1986). Growth in Early Childhood in Developing Countries in F. Falkner and J. Tanner (eds.). *Human Growth: A Comprehensive Treatise*, Vol 3, 2nd edition, Plenum Press, New York.
- Mason, K. O. and A. M. Taj. (1987). Difference between women's and men's reproductive goals in developing countries, *Population and Development Review*, 13.4:611-38.

- Mauldon, J. (1990). The effect of marital disruption on children's health, *Demography*, 27.3:431-47.
- McElroy, M. B. (1990). The empirical content of Nash-bargained household behavior, *Journal of Human Resources*, 25.4:559-83.
- McElroy, M. B. and M. J. Horney (1981). Nash bargained household decisions, *International Economic Review*, 22(2):333-50.
- McElroy, M. B. and M. J. Horney (1988). Nash bargained household decisions: Reply, *International Economic Review*, 31(1):237-42.
- Morgan, S. P., D. Lye and G. Condran. (1988). Sons, daughters and the risk of marital disruption, *American Journal of Sociology*, 94.1:110-29.
- Mensch, B., H. Lentzner and S. Preston (1986). Socio-Economic differentials in Child Mortality in Developing Countries, United Nations, New York.
- National Center for Health Statistics (1976). *Growth Charts*, US Department of Health and Human Services, Public Health Service, Health Resources Administration.
- National Research Council (1983). *Levels and Recent Trends in Fertility and Mortality in Brazil*, Report # 21. Washington D.C.: National Academy Press.
- National Research Council. (1983). *The Determinants of Brazil's Recent Rapid Decline in Fertility*, Washington D. C.:National Academy Press.
- Neuhouser, K (1989). Sources of women's power and status among the urban poor in contemporary Brazil, *Signs*, 14.3 685-702.
- Nsamenang, A. B. (1987). A West African perspective. In M. Lamb, (ed.), *The Father's Role: Cross cultural perspectives*, London:Lawrence Erlbaum.
- Oppong, C. (1982). *Middle class African marriage*, London:George Allen and Unwin.
- Oppong, C. and W. Bleek. (1982). Economic models and having children: Some evidence from Kwahu, Ghana, *Africa*, 52.3
- Peil, M. (1977). *Consensus and conflict in African societies*. London:Longman.
- Pitt, M., M. Rosenzweig and N. Hassan (1990). Productivity, health and inequality in the intrahousehold distribution of food in low-income countries, *International Economic Review*, forthcoming.
- Psacharopoulos G. and A. M. Arriagada (1989). The determinants of early age human capital formation: evidence from Brazil, *Economic Development and Cultural Change*, 37:683-708.
- Rosenzweig, M. and T. P. Schultz (1982). Market opportunities, genetic endowments and the intrafamily distribution of resources, *American Economic Review*, 72(4):803-15.
- Schofield (1979). *Development and the Problems of Village Nutrition*, Croom Helm, London.
- Schultz, T. P. (1987). School expenditures and enrollments, 1960-1980: The effects of income, prices and population. In *Population Growth and Economic Development*, eds., D. G. Johnson and R. Lee. Madison: University of Wisconsin Press, 413-76.
- Schultz, T. P. (1989). Women and Development: Objectives, frameworks and policy interventions. WPS 200, PHR, The World Bank, Washington D.C.
- Schultz, T. P. (1990). Testing the Neoclassical model of family labor supply and fertility, *Journal of Human Resources*, 25.4:599-634.
- Sen, A. (1984). Family and food: sex bias in poverty, in A. Sen, *Resources, Value and Development*, Blackwell, London.

- Sen, A. and S. Sengupta (1983). Malnutrition of rural children and the sex bias, *Economic and Political Weekly*, 18:855-864.
- Senauer, B., M. Garcia and E. Jacinto (1988). Determinants of the intrahousehold allocation of food in the rural Philippines, *American Journal of Agricultural Economics*, 70.1:170-80.
- Spanier, G. B. and P. C. Glick. (1981). Marital instability in the United States: Some correlates and recent changes, *Family Relations*, 31:319-38.
- Spitze, G. and J. Logan. (1990). Sons, daughters and intergenerational social support. *Journal of Marriage and the Family*, 52:420-30.
- Strauss, J. (1990). Households, communities and preschool children's nutrition outcomes: evidence from rural Côte d'Ivoire, *Economic Development and Cultural Change*, 38(2):231-62.
- Subramanian, S. and A. Deaton. (1990). Gender effects in Indian consumption patterns. Research Program in Development Studies Discussion Paper 147, Princeton University.
- Svedberg, P. (1990). Undernutrition in sub-Saharan Africa: is there a gender bias?, *Journal of Development Studies*, 26.3:469-86.
- Teachman, J. D. and P. T. Schollaert. (1989). Gender of children and birth timing, *Demography*, 26.3:411-24.
- Thomas, D. (1990). Intra-household resource allocation: an inferential approach, *Journal of Human Resources*, 25.4:635-664.
- Thomas, D., J. Strauss and M. H. Henriques (1990). Survival rates, height for age and household characteristics in Brazil, *Journal of Development Economics*, forthcoming.
- Townsend, R. (1989). Risk and insurance in Village India. Manuscript, University of Chicago.
- Udry, C. (1990). Rural credit markets in northern Nigeria: Credit as insurance in a rural economy, *World Bank Economic Review*, forthcoming.
- Ulph, D. (1988). A general non-cooperative nash model of household consumption behaviour, mimeo, Bristol University.
- Waterlow, J., R. Buzina, W. Keller, J. Lane, M. Nichman and J. Tanner (1977). The presentation and use of height and weight data for comparing the nutritional status of groups of children under the age of ten years, *Bulletin of the World Health Organization*, 55:489-98.
- White, H. (1980). A heteroskedasticity-consistent covariance matrix and a direct test for heteroskedasticity, *Econometrica*, 48:817-38.

## Appendix

This appendix compares estimates of quasi- $t$  statistics based on the jackknife, infinitesimal jackknife (Jaeckel, 1974, also suggested by White, 1980) and the usual formula for the variance-covariance matrix. Estimates are reported both for least squares and for random effects; for the latter, the variance covariance matrix for the final weighted least squares estimates is computed by the jackknife, infinitesimal jackknife and usual methods. As an example, the estimates using the NLS data on black children (including the education of the father and the mother's current partner) (third column of panel b in Table 3) are repeated in Table A.1b.

Whereas the difference between least squares and random effects is small, differences between the standard  $t$  statistics and jackknifed  $t$ s are in many cases quite large and inference changes in several instances. The jackknife and White estimates are often quite similar but there are three instances in which they differ dramatically: the estimate of the effect of father's education interacted with the sex of the child is significant if White  $t$ -statistics are used (in both the OLS and random effects estimates) but not significant if the jackknife or standard estimates are used.

This result is caused by a small number of observations which are leverage points and, therefore, have a very large influence on the estimates of the variance covariance matrix. The primary difference between the jackknife and infinitesimal jackknife lies in the weighting of residuals by a transform of the diagonal element of the hat matrix. This weight has been proposed as a useful indicator of the influence of observations in the regression model (Belsley, Kuh and Welsch, 1980; Handschin *et al.*, 1975). Observations with large values of this weight tend to be outliers in the covariate space and are called leverage points.

Clearly the infinitesimal jackknife or White estimate of standard errors can be *seriously misleading*. This fact has been pointed out by MacKinnon and White, (1982), on the basis of a Monte Carlo study and, from our own experience, the problem arises in a variety of empirical models: adoption of the jackknife estimator of the variance covariance matrix is not just an empirical nicety. Since the computational burden of estimating White and jackknife estimates is virtually identical, all  $t$  and  $\chi^2$  test statistics in the tables are based on the jackknife.

**Table A.1**  
Random effects estimates and estimates of quasi-t-statistics

*(a) Random effects estimates : NLS (1986) stratified by race of child*

Effect of mother's education and mean of all partners' education

	Blacks			Hispanics			Whites		
	Females	Males	Difference	Females	Males	Difference	Females	Males	Difference
<b>Mother's education</b>									
(1) if 7 -12 yrs	2.317 [1.38]	-2.063 [4.33]	-4.227 [2.72]	0.648 [1.27]	-0.314 [1.56]	-1.020 [2.00]	0.147 [0.39]	1.560 [1.05]	1.288 [0.86]
13-20 yrs	2.446 [1.45]	-2.188 [4.27]	-4.502 [2.86]	0.776 [1.43]	0.321 [1.10]	-0.499 [0.87]	0.372 [0.94]	1.599 [1.05]	1.092 [0.71]
<b>Partners' mean education</b>									
(1) if 7 -12 yrs	0.037 [0.14]	0.060 [0.12]	0.066 [0.11]	-0.061 [0.22]	-0.068 [0.27]	-0.016 [0.04]	-0.254 [1.40]	0.267 [1.69]	0.494 [2.13]
13-20 yrs	-0.159 [0.51]	-0.046 [0.08]	0.146 [0.23]	-0.182 [0.52]	0.200 [0.68]	0.392 [0.86]	-0.408 [1.82]	0.333 [1.38]	0.702 [2.21]
<b><math>\chi^2</math> for joint significance of</b>									
mother's education	2.294 [0.16]	19.272 [0.00]	8.464 [0.01]	2.092 [0.18]	7.209 [0.01]	5.792 [0.03]	3.833 [0.07]	1.113 [0.29]	1.711 [0.21]
partners' mean education	1.167 [0.28]	0.272 [0.44]	0.093 [0.48]	0.137 [0.47]	1.425 [0.25]	1.530 [0.23]	3.382 [0.09]	3.225 [0.10]	5.860 [0.03]
all covariates	138.358 [0.00]	108.334 [0.06]	237.283 [0.00]	32.961 [0.00]	20.235 [0.05]	53.288 [0.01]	52.317 [0.00]	38.391 [0.00]	199.493 [0.00]
LM test		1.812			3.023			7.897	
Hausman test		-14.764			-29.338			38.434	

*(b) Quasi t-stastics : OLS and random effects estimates : NLS (1986) : Black children*

Effect of education of mother, father and current partner

	Coefficient estimates		Quasi t statistics						
	OLS	Random effects	OLS estimates			Random effects estimates			
			Stand-ard	White	Jack-knife	Stand-ard	White	Jack-knife	
<b>Education</b>									
(1) if mother's education 7-12 years	2.293	2.377	2.31	1.84	1.36	2.37	1.94	1.43	
13-20 years	2.422	2.484	2.41	1.92	1.42	2.45	2.00	1.48	
(1) if father's education 7-12 years	0.202	0.156	0.10	1.11	0.00	0.08	0.77	0.00	
13-20 years	0.078	0.033	0.04	0.25	0.00	0.02	0.10	0.00	
(1) if partner's education 7-12 years	0.083	0.069	0.07	0.21	0.15	0.06	0.17	0.12	
13-20 years	0.023	0.051	0.02	0.05	0.04	0.04	0.11	0.08	
<b>Education * male interaction</b>									
(1) if mother's educ 7-12 years * male	-4.274	-4.195	3.04	3.30	2.44	3.14	3.68	2.72	
13-20 years * male	-4.549	-4.471	3.20	3.44	2.56	3.32	3.83	2.86	
(1) if father's educ 7-12 years * male	1.367	1.441	0.56	4.21	0.01	0.63	4.35	0.01	
13-20 years * male	1.734	1.819	0.70	3.31	0.02	0.78	3.40	0.01	
(1) if partner's educ 7-12 years * male	-0.482	-0.387	0.30	0.91	0.64	0.25	0.78	0.54	
13-20 years * male	-0.792	-0.769	0.47	1.21	0.92	0.48	1.22	0.93	
<b>Presence of partner/father</b>									
(1) if current partner exists	-0.053	-0.026	0.05	0.14	0.10	0.02	0.07	0.04	
"-" * male	0.347	0.250	0.21	0.68	0.47	0.16	0.53	0.36	
(1) if father present after birth	-0.013	0.009	0.01	0.07	0.00	0.01	0.04	0.00	
"-" * male	-1.245	-1.303	0.51	3.83	0.01	0.57	3.98	0.01	
<b>Mother's characteristics</b>									
Mother's height (cm)	0.109	0.107	4.16	3.71	3.60	3.94	3.51	3.41	
"-" * male	-0.044	-0.041	1.19	1.01	0.98	1.12	0.96	0.93	
Age of mother at menarche	-0.095	-0.099	2.16	2.42	2.36	2.17	2.43	2.35	
"-" * male	0.017	0.019	0.27	0.30	0.29	0.31	0.34	0.33	
<b>Child characteristics</b>									
(1) if child is male	7.363	7.173	2.55	2.22	2.06	2.52	2.21	2.06	
aged 6 - 11 months	-0.639	-0.593	1.32	1.41	1.36	1.23	1.33	1.28	
12 - 17 months	-0.837	-0.855	1.78	1.96	1.86	1.82	2.01	1.91	
18 - 23 months	-0.816	-0.789	1.74	1.69	1.63	1.68	1.66	1.60	
24 - 35 months	-0.054	0.048	0.13	0.14	0.13	0.12	0.13	0.12	
36 - 47 months	0.612	0.612	1.49	1.81	1.74	1.49	1.83	1.77	
48 - 59 months	0.692	0.737	1.64	2.11	2.03	1.75	2.28	2.20	
60 -107 months	0.737	0.778	1.93	2.47	2.37	2.03	2.66	2.55	
108 -143 months	0.656	0.728	1.56	2.14	2.05	1.73	2.44	2.34	
aged 6 - 11 months * male	-0.789	-0.891	1.19	0.89	0.86	1.35	1.01	0.98	
12 - 17 months * male	-0.284	-0.291	0.44	0.31	0.30	0.45	0.32	0.31	
18 - 23 months * male	-0.283	-0.387	0.43	0.30	0.29	0.58	0.41	0.39	
24 - 35 months * male	0.080	-0.131	0.14	0.10	0.09	0.23	0.16	0.15	
36 - 47 months * male	-0.913	-1.010	1.60	1.14	1.10	1.76	1.26	1.22	
48 - 59 months * male	-0.949	-1.030	1.63	1.19	1.15	1.77	1.29	1.24	
60 -107 months * male	-0.689	-0.795	1.32	0.91	0.88	1.52	1.05	1.02	
108 -143 months * male	-0.273	-0.429	0.48	0.36	0.34	0.75	0.56	0.54	
Constant	-8.235	-8.122	4.02	3.51	3.14	3.87	3.39	3.04	