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Sibling rivalry and the gender gap: Evidence from child health outcomes in Ghana

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Abstract. When capital and labor markets are imperfect, choice sets narrow, and parents must choose how to ration available funds and time between their children. One consequence is that children become rivals for household resources. In economies with pro-male bias, such rivalries can yield gains to having relatively more sisters than brothers. Using a rich household survey from Ghana, we find that on average if children had all sisters (and no brothers) they would do roughly 25-40% better on measured health indicators than if they had all brothers (and no sisters). The effects are as large as typical quantity-quality trade-offs, and they do not differ significantly by gender.

JEL classification: I12 , J16, O12

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1. Introduction

Most children of poor parents are doubly disadvantaged. Their parents often lack the personal resources to invest optimally in the human capital of their children, and the problem is compounded by constraints faced by parents in obtaining credit and in flexibly allocating time. As a result, even when there are high returns to investing in the human capital of children, education and health levels often remain low.

A further consequence of constraints in capital and labor use is that parents must ration available funds and time to each of their children. Children thus become rivals, even when there is no explicit strategic behavior on the part of any family member. As with psychological notions of sibling rivalry, the relative genders and ages of siblings can be central in determining the outcomes of these rivalries, creating human capital differences between siblings from the same family that can be as great as those between siblings from different families (Dunn and Plomin 1990).

While only a handful of papers consider sibling rivalry in economic contexts, many common variants of the economic theory of the household imply that the genders and ages of a child's siblings help determine human capital allocations.[1] This paper considers these issues using a rich household survey from Ghana, focusing largely on the place of gender. We pay particular attention to whether children with relatively more sisters than brothers are healthier than their peers.

An explanation follows naturally from the pure investment model of Becker (1991), generalized to allow imperfections in labor and capital markets. Where market constraints bind, children will fare better when pitted against siblings with fewer intrinsic advantages. In societies with pro-male biases, children with relatively more sisters than brothers will then benefit most. This explanation, for example, has been used to explain why Taiwanese children with relatively more older sisters have higher education rates (Parish and Willis 1994).

The advantages to having sisters will be mitigated (or perhaps reversed) when there are positive spillovers in the "production" of human capital within families. In addition, advantages might also be reversed if parents act with concern for the fairness of investments across children (e.g. Behrman et al. 1982). When parents desire to equalize outcomes across their children, *ceteris paribus* all children gain when a "high return child" (son) replaces a "low return child" (daughter). Brothers can also be helpful if they bring relatively more resources into the household, either directly by working when young or by raising the household's permanent income.

The relative strength of these competing forces is ultimately an empirical issue. Similarly, there is no general theoretical presumption that, for example, boys will benefit more from having sisters than will girls. But the preponderance of the evidence points to fairly clear patterns: In looking across a number of anthropometric measures, we find that if children had all sisters (and no brothers) they would do roughly 25-40% better than if they had all brothers (and no sisters). The findings can be seen both in simple bivariate relationships and in econometric specifications that allow controls for birth order; cultural factors; both observed and unobserved family heterogeneity; and systematic biases due to the use of U.S. anthropometric standardizations. While the data show strong advantages to having sisters, we find very little systematic difference between impacts on boys and girls. Nor do we find that birth order matters very much in our sample. Contrary to our expectations, the gains to having sisters are similar and sometimes weaker in poorer households relative to richer households.

2. Human capital investment and the gender gap

Male-female differentials in returns to investing in human capital are often invoked to explain gender gaps in health and education. In Ghana, the gender gap in returns is due to both labor market forces and, to a large degree, cultural practices. In many households, women move out of the family when they marry, while men stay within the household with their wives. Thus, the full return to investing in sons is more likely to be reaped by parents than the return to investments in daughters.[2]

This economic logic can be extended to explain why rising income often leads to increased accumulation of human capital and the narrowing of gender gaps (even when the pattern of returns remains unchanged). As long as the human capital of children is valued intrinsically, rising income will lead to rising human capital (assuming that human capital is treated like a normal good). And gender gaps will close under the common assumption that parents' aversion to the unequal treatment of their children also increases with income.

However, we show below that preference-based arguments are not necessary to explain why gender gaps close. Nor is it always so that gender gaps will narrow as constraints are eased or income rises. Narrowing or widening depends largely on the relative curvature of health returns functions for males and females, an attribute of human capital returns functions about which very little has been established empirically.

The implications of sibling sex composition

We begin by assuming that parents perceive that the returns to investing in the health of sons is higher than investing in daughters. The central elements of the discussion, however, extend to any situation in which children differ in their endowments (see, e.g., Behrman et al. 1982). The total financial return to investing H units in each male is given by the concave health returns function $R_m = R_m(H)$ where $R' > 0$ $R'' < 0$. Similarly, the total return to investments in females is given by the concave function $R_f = R_f(H)$. Pro-male advantage is captured by the assumption that $R_m(H_i) > R_f(H_i)$ for all investment levels i.[3]

We begin by considering the pure investment case, in which parents make investments in their children based solely on expected economic returns. When there are neither constraints on time nor credit, investments in children will reflect solely their returns relative to the cost of the funds. When unit child costs are equal for both sons and daughters, parents will invest in each child until the marginal value product equals the gross cost of borrowing, $1 + r$. Thus, $R'_f(H_f) = R'_m(H_m) = (1 + r)$. Since the invest-

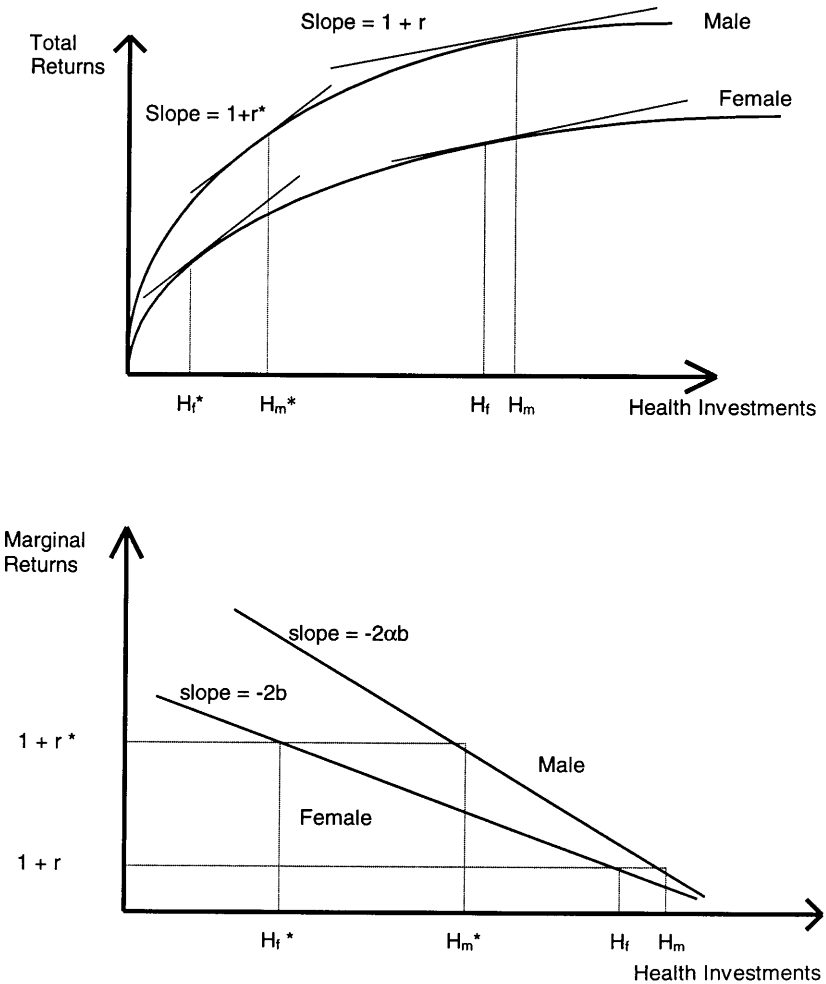

Fig. 1. Optimal health investments and a narrowing gender gap

ments depend solely on the cost of borrowing and expected returns, the gender composition of siblings will make no difference to investments here – for boys or for girls.

However, when resource constraints are binding, the story changes since children must compete now for the resources currently available to the household. In this sibling rivalry, the children with lower returns lose out. Boys then gain the advantage when there is pro-male bias in returns. But both boys and girls then do better the more sisters they have, holding constant the total number of siblings.

It is ambiguous whether boys or girls stand to gain more from having sisters – so that the effect on the gender gap is ambiguous in theory. We illustrate the possibilities with two simple competing cases, continuing under the assumption that parents' choices are driven by pure investment motives.

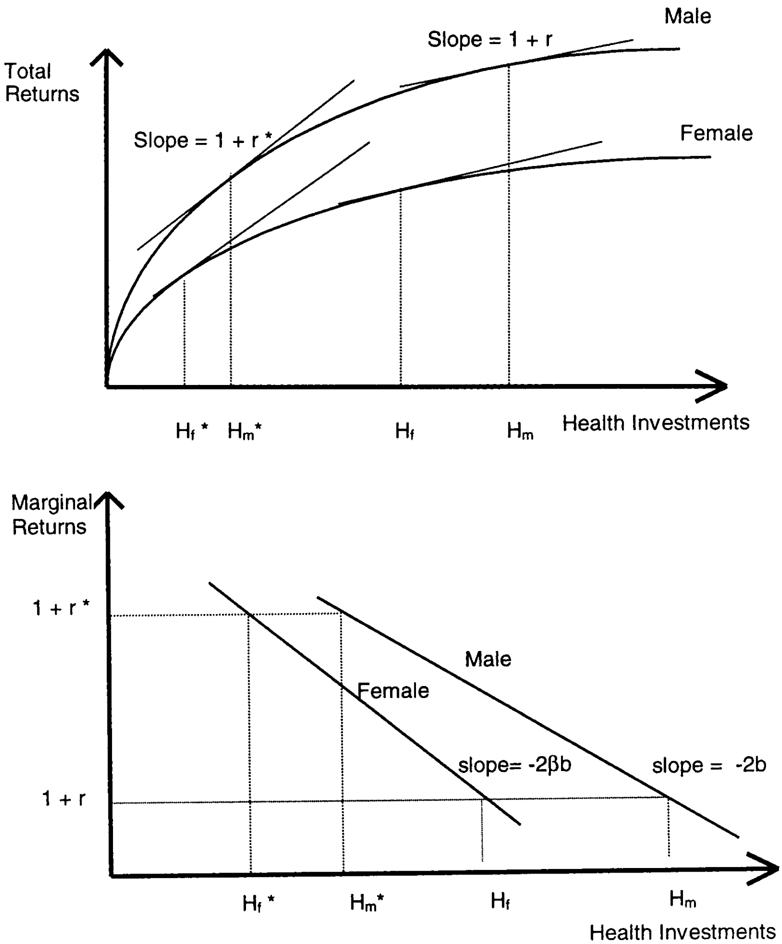

Fig. 2. Optimal health investments and a widening gender gap

(These cases are admittedly artificial; see Garg and Morduch 1997, for a general theoretical model with comparative static results.)

In the first case, assume that net returns to the parents for investments in their sons take the quadratic form: $R_m(H) = aH - bH^2$, where a and b are positive numbers which satisfy concavity of the function. The returns to investing in daughters are assumed to be $R_f(H) = \alpha R_m(H) = \alpha(aH - bH^2)$, so that with $\alpha < 1$ the returns to females are below those of males for every level of investment, but the returns decline at a slower rate. This is depicted in the top panel of Fig. 1. The lower panel gives the corresponding marginal returns to health investments. Both relationships are negative and linear, given the quadratic functions. Because $\alpha < 1$, the slope of the marginal returns for females is less than that of males. When resource constraints do not bind, marginal returns are set equal to the cost of borrowing, $(1 + r)$, and this leads to investment levels of H_f and H_m. There is a gender gap because $\alpha < 1$, but it is relatively small. However, when resource constraints bind, marginal returns are set at levels higher than the cost of borrowing, depicted in Fig. 1 at level $1 + r^*$. With tightening resource constraints, the gender gap widens from $(H_m - H_f)$ to $(H_m^* - H_f^*)$. Thus, resource constraints exacerbate initial differences in male-female returns. The positive link between the relaxing of constraints and the narrowing of the gender gap is reinforced under the plausible assumption that discrimination falls at higher health levels.

A counter example is given under the alternative assumption that, while total returns to males remain higher everywhere, marginal returns to investments in daughters fall more quickly than for investments in sons. This case is depicted in Fig. 2. It can be characterized by the assumption that the returns to females are instead given, for example, by the relationship:

$$R_f(H) = (aH - \beta bH^2) \ , \quad \text{where } \beta > 1 \ .$$

Figure 2 shows that the gender gap gets larger as resource constraints are lifted. Thus, prediction of impacts on the gender gap requires more than knowledge of the concavity of returns functions for males and females: it is necessary to know the relative degrees of concavity as well.[4]

Alternative explanations and issues

The scenario above captures only one part of the way that human capital is formed. As described in the introduction, there are counter-balancing advantages to having brothers rather than sisters even when biases favor males. This is particularly so when parents are concerned with the fairness of allocations to children, if there are positive within-family spillovers, and if boys bring relatively more resources into the family.

The relationship between sibling composition and human capital may also be influenced directly by psychological and anthropological phenomena. Two mechanisms stand out in this regard. First, there may be spillovers in the way that children are socialized, such that having at least one brother may lead parents to instill more “masculine” traits in their daughters. Some researchers include among those traits greater self-confidence and enhanced physical activity. This may then affect the way that girls with brothers are treated – and expect to be treated – relative to girls with only sisters. We capture this in the empirical work with a dummy variable that indicates whether a child has at least one brother. For girls, this captures possible spillover effects, while for boys it captures possible “reference group effects” (see below). The dummy variable should be non-negative for both boys and girls.

Secondly, “reference group effects” may exist such that girls with only brothers are treated differently from girls with at least one sister. Without sisters, a single daughter may be treated similarly to the boys in the family, but differences may widen once another girl is added to the family, altering the yardstick for comparison of treatment (the dummy variable for having at least one brother captures these reference group effects for boys). This effect has been found by Butcher and Case (1994) to explain completed education in a sample of U.S. women. They find that having at least one sister is associated with a decline in average female educational levels of half a year (and, correspondingly, 9% lower incomes.) We do not, however find these sorts of effects in the data. Kaestner (1996) also cannot find such effects when investigating the education of later born cohorts in the United States. Interestingly, however, he does find that in a sample of black adults, those with relatively more sisters had greater educational outcomes, similar to the findings here.

Another mechanism through which sibling composition can matter is that sisters may help other siblings directly – perhaps by working and bringing extra resources into the household or by taking care of younger siblings. Parish and Willis (1994), find strong evidence that in Taiwan having older sisters is associated with higher educational investments in younger children (see Parish and Willis 1994, for a survey of the related literature.) In contrast, Das Gupta (1987) finds that in rural Punjab it is girls with older sisters that suffer most in the face of pro-son bias.

In considering these explanations, we find that spillover/reference effects matter consistently in explaining patterns of child health in Ghana. However, we find little evidence in support of reference group effects for girls, nor of the particular advantage of either having older sisters or of having younger sisters – although they might emerge in a larger data set. Below, we refer to "spillover effects" as shorthand for "spillover effects for girls/reference effects for boys".

3. Data description

The data we analyze come from the 1988-1989 Ghana Living Standards Survey (GLSS) which includes an extensive household questionnaire. The survey was completed as part of the World Bank's Living Standard Measurement Survey program, and a variety of quality assurance features were built into the survey design. Other studies which use the GLSS data on health include Thomas (1994); Glewwe and Jacoby (1995); Pitt and Lavy (1996), and Behrman and Lavy (1994).

The survey consists of nearly 3,200 households drawn to form a sample that is representative of the ten regional and four ecological zones in the country. The data permit us to analyze the gender and age composition of children living at home at the time of the survey. This may understate the total number of the siblings, especially older siblings. Our full sibling data set contains basic anthropometric outcomes for 5,203 children. Out of these children, 3,354 are under age eleven, and 2,458 have consistent and available anthropometric and parental data. Appendix Table 1 provides summary statistics of the data.

One adult woman in each household was selected randomly and questioned about her fertility history (providing information to use in forming instrumental variables). Of the 7,230 children born to the women that responded, 83% were currently alive and 60% of those were currently resident in the household. Many of the children in question were already adults by the time of the survey and had left home upon marriage or to find employment. Nearly half (48.9%) of children born were female and a very similar fraction (49.1%) remained alive at the time of the survey.

Thus, neither the fertility nor mortality data indicate signs of gross differences in the treatment of sons vs. daughters.

Height and weight measurements are available for most young children, and we have converted them to U.S. standards using the National Center for Health Statistics (NCHS) standards. The general health indicators are height-for-age, a measure of longer-term health status; and weight-for-height and weight-for-age, two measures of medium-term health status. The indicators of extreme health outcomes are stunting, wasting, and being underweight. These are defined by World Health Organization (1986) as being at least two (U.S.) standard deviations below the reference U.S. population.[5] Following best-practices advocated by the World Health Organization, we focus where feasible just on the health of children under age seven. This minimizes the risk that the data reflect mostly genetic variation, as is more likely to be the case for health outcomes of older children. The data set is unusual in also providing anthropometric data for parents. A parent's physical status is an important determinant of the physical status of their children, and this provides an important control for family-specific heterogeneity.

The raw anthropometric data show strong evidence of malnutrition according to U.S. standards. The standardized measures for height-for-age, weight-for-age and weight-for-height were well below comparable standards in the United States. The median child, for example, was 93% of the U.S. median in terms of height-for-age, which is generally regarded as a robust indicator of long-term nutrition; nearly 31% of the sample under age seven is stunted and 5% wasted (Appendix Table 1 A). These levels are comparable to those found in earlier studies of Ghana (e.g., Alderman 1991).

The role of siblings: bivariate analyses

Before turning to the econometric results, we show the broad patterns in bivariate analyses. Because the gender composition of children is not chosen by the household – and is unlikely to be correlated with most determinants of health – the bivariate analyses tell most of the story.[6] Since there is little evidence of either excess female mortality or son-preference in fertility in Ghana, sibling sex composition should be orthogonal to the total number of siblings in the sample – unless there are sex-based patterns of child departure from the household. The Spearman rank correlation coefficient for total siblings and the percent that are female is –0.013, essentially zero. Non-parametric regression of the two variables yields a flat line at just under 50% female.

In the sections below, we corroborate the bivariate analyses with regression analyses. The regressions control for possible biases due to the inappropriateness of U.S. standardizations and to omitted factors beyond family structure. The regression results provide much the same qualitative and quantitative results seen here.

Table 1 presents mean values of two anthropometric outcomes, height-for-age and weight-for-height. Samples have been expanded beyond the preferred group of children under age seven used in the regression analyses in order to maintain reasonable sizes in each cell. The sample for height-for-age includes all children 15 and below and the sample for weight-for-

Table 1. How the number of sisters affects nutritional status: bivariate analyses

Sisters	Number of siblings							
	One	Two	Three	Four	Five	Six	Seven	Eight
Height-for-age								
None	–1.28	–1.44	–1.32	–1.74	–1.60	–1.85	–2.40	–2.12
One	–1.22	–1.49	–1.52	–1.62	–1.58	–1.65	–1.91	–1.75
Two		–1.31	–1.44	–1.28	–1.50	–1.68	–1.90	–1.50
Three			–1.39	–1.34	–1.27	–1.59	–1.94	–1.73
Four					–1.22	–1.25	–1.97	–0.66
Five						–1.09	–1.86	–1.54
Six								
Mean	–1.24	–1.43	–1.44	–1.53	–1.49	–1.52	–1.74	–1.72
Observations	401	595	681	535	330	237	161	112
Weight-for-height								
None	–0.681	–0.632	–0.669	–0.669	.	–0.523	–0.546	
One	–0.691	–0.636	–0.721	–0.641	–0.773	–0.684	–1.02	
Two		–0.690	–0.661	–0.629	–0.746	–0.654	–0.782	
Three			–0.570	–0.594	–0.750	–0.640	–0.717	
Four				–0.593	–0.713	–0.632	–0.677	
Five					–0.590	–0.601	.	
Six							–0.350	
Mean	–0.686	–0.640	–0.662	–0.644	–0.727	–0.645	–0.789	
Observations	239	354	397	269	159	116	87	

Height-for-age is calculated for all children age 15 years and below.
Weight-for-height is calculated for children age 10 years and below.

height includes only children age 10 and below. The latter sample is cut more sharply since weight-for-height is particularly sensitive to genetic factors at higher ages.

The average values are displayed by total sibling size and, within each sibling group, by the number of sisters. The negative numbers in the table reflect that levels are below U.S. median standards. Thus, improvements in the measures occur as they become less negative. The numbers in the tables are simple sample averages (not predictions) of the values for children in particular cells. The tables can be read across as showing quality-quantity trade-offs and downward showing the impact of substituting sisters for brothers.

The results show that, for the most part, children in larger families are less healthy. Consistent with the idea of a quality-quantity trade-off (Becker and Tomes 1976), when sibling size increases from two to seven, average child health status worsens by 17% as measured by height-for-age and by 19% as measured by weight-for-height.

The variation due to shifting sibling composition is at least as striking. This is seen by comparing measures down any given column. For example, for height-for-age, shifting form having four siblings, all of which are brothers, to having one brother and three sisters leads to a 23% improvement in height-for-age. The corresponding improvement for children with five siblings is 24%, and with six siblings average height-for-age improves by 41% when moving to having all but one sister. Similar qualitative patterns can be seen in the weight-for-height table with smaller magnitudes. For example, moving from having three brothers to three sisters improves weight-for-height by 15%. Moving from four brothers to four sisters improves the measure by 11%.

The bivariate tables show some anomalies as well, especially for larger households and very small households. Partly this arises because cell sizes diminish sharply for both larger and smaller sibling groups, allowing outliers to have a greater influence. With these exceptions, the bivariate tables show clear trends in the data.

4. Empirical specification

The regression specifications below yield similar conclusions to those in the bivariate analyses. The regression specification is useful in providing an explicit test of the hypotheses, and it allows comparison of the relative magnitudes of the effects of sibling composition, income, birth order, and socio-cultural variables. These elements are captured by a series of health outcome equations that take the general form:

$$\begin{aligned} H_{ij} = {} & \alpha_0 + \alpha_1 \boldsymbol{X}_{ij} + \alpha_2 \boldsymbol{Z}_j + \alpha_3 \boldsymbol{Y}_j + \alpha_4 N_j + \alpha_5 S_{ij} + \alpha_6 S_{ij}^2 \\ & + \alpha_7 R_{ij} + \sum_{\text{sex}} \sum_{\text{age}} \alpha_{as} \text{Sex}_{ij} \text{Age}_{ij} + \eta_j + \mu_{ij} \ , \end{aligned} \tag{1}$$

where H_{ij} is a measure of the health of child i in household j; $\boldsymbol{X}_{ij}$ is a vector of child-specific variables like birth order; $\boldsymbol{Z}_j$ is a vector of household variables like the height and education of parents; and $\boldsymbol{Y}_j$ is total household expenditure per capita. We use expenditure rather than income since we lack reliable income data and expenditure provides a more accurate gauge of current resources given the possibility of consumption smoothing. The total number of siblings is N_j, and the number of sisters is S_{ij}. Spillover effects are captured by R_{ij}, a dummy variable that equals one if the child has at least one brother.

Equation (1) is a linear approximation to a fundamentally nonlinear relationship, and we also estimate it with a quadratic term in the number of siblings, in logarithmic form, and without the "one brother" dummy variable. The central results are robust to these permutations (Appendix Tables 2 A and 3 A and Garg 1997).

The η_j accounts for unobserved family fixed effects common to all siblings. This includes parental preferences for health, knowledge about health, and access to health providers. We deal with these unobservables explicitly by estimating a random effects generalized least squares (GLS) regression and testing whether the η_j term is different from zero and uncorrelated with the μ_{ij}, the individual-specific errors. In no specification can we reject that the coefficients for the random effects model differ systematically from coefficients estimated using fixed effects (Hausman 1978).

Where we cannot use random effects (e.g., in the probits on extreme health outcomes), we use Huber's (1964) heteroscedasticity correction to control for the family effect and cluster-based sampling.

The analysis centers on health outcomes conditional on a given household structure. That structure may be partly endogenous, though. Most importantly, the quality-quantity tradeoff implies that parents that care more about the health of their children will also have fewer children. As a result, there may be a negative relationship between sibling size, N_j and the unobserved household effects, η_j. Similarly, the existence of at least one son may be endogenous. Since there are no convincing instruments for sibling size, an alternative approach would be to estimate a reduced form equation that excludes these potentially endogenous variables (and consider sisters as a fraction of siblings instead of the number of sisters). However, our interest is in how parents act given their current constraints, and we follow standard approaches in empirical studies of consumer behavior by conditioning on the present structure of the household.

A related issue arises in that we only consider children of the household head (and head spouse) currently residing in the household. We do not consider other children temporarily residing in the household, nor children that have left the household. These exclusions best approximate the short term concerns of most households, but the data set allows us to test robustness to excluding people no longer resident. As mentioned above, we have full fertility histories from one woman in the household (not necessarily the head woman) and can match those histories to half of the children here. In alternative specifications we use the number (and gender) of children born alive as instruments for the number of children and the number of sisters. The two-stage least squares estimates are less precise (partly since the samples are cut) but their size is slightly larger than the OLS estimates. These instrumented equations do not, however, address the more fundamental endogeneity reflected in the quality-quantity trade-off.

To test differences in outcomes for males and females, we also interact a sex dummy (1 = female) with per capita expenditure, total siblings, the "one brother" dummy variable, and the quadratic in the number of sisters. As described above, whether the impacts on girls exceed those on boys depends on the relative concavity of male versus female returns functions.

The health variables are standardized according to U.S. nutritional standards. However, because U.S. standards may not be appropriate for nutrition in Ghana, we include a full set of dummy variables that allow age-sex intercepts specific to Ghana. This is a conservative approach since it controls for the possibility that the results are influenced by systematic biases in the U.S. standardizations. The cost is that the dummy variables absorb substantial variation in the dependent variables.

By including total expenditures per capita as an explanatory variable, we control implicitly for the impact of sibling composition on total household resources. Thus, the coefficient on the sibling variable S_{ij} reflects the "allocation effect" – i.e., the impact given available resources. It captures the effect on how the "pie is split", not on the size of the pie. In general, households with a higher fraction of daughters have slightly higher expenditures (Garg and Morduch 1997). This may be because girls work more than their brothers or because parents reduce labor supply more when they have sons rather than daughters. Thus, the results below slightly under-estimate the total positive impact of sisters on their siblings.

5. Empirical results

Results on the base anthropometric measures

Table 2 gives the results for the base anthropometric measures. Following World Health Organization recommendations, the sample is restricted to children under age seven (since the indicators are most sensitive to environmental versus genetic factors for younger children). Sibling composition affects child health significantly in all specifications without interactions. Coefficients on the number of sisters are positive and large, although quadratic terms are seldom statistically significant. Discussion of magnitudes is deferred until the section on predicted effects (Table 5). Both parents' heights matter strongly, and increasing per capita expenditures is consistently helpful for health status, while increasing family size is deleterious.

None of the interacted terms on the "sister variables" are significant, and we cannot reject the hypothesis that effects for girls and boys are the same. The (non-interacted) number of sisters remains significant once the interactions are added, with the exception of the weight-for-height equation. The "spillover/reference effect" indicated by the dummy variable of having at least one brother affects both boys and girls (positively, as expected). Unlike the Butcher and Case (1994) study, we find that the effect is sometimes slightly larger for boys than for girls; the result indicates that there may be an important reference group effect for boys also.[7]

Extreme health outcomes

Table 3 provides results on the determinants of stunting, wasting, and being underweight. These measures reflect children who fall two standard deviations below the U.S. medians for height-for-age, weight-for-height, and weight-for-age, respectively. While these are "extreme" health outcomes, they are not uncommon in the Ghanaian context. Of children under age 7, over 30% are stunted, 5% are wasted and 26% are underweight (Appendix Table 1). Unlike the previous measures, progress here is indicated by declining values in the dependent variables. Thus, the expected coefficients uniformly take the opposite sign to those above.

The results on stunting and being underweight in Table 3 show that these outcomes are more likely in households with more children, and less likely in richer households. Again, parents' height matters, and younger children tend to fare better. With one exception (the interaction with "one brother" in the "wasting" regression), none of the interactions take significant coefficients and most are small, yielding little impact on gender gaps.

Strikingly, the impact of having more sisters is strong, reducing the incidence of these extreme health outcomes. It matters significantly in explaining stunting and being underweight, but not in explaining wasting (which may not be surprising given that wasting affects just 5% of the population.)

Table 2. GLS – Random effects estimates: base health outcomes

Coefficient	Specifications					
	Height-for-age		Weight-for-age		Weight-for-height	
	Pooled	Interacted	Pooled	Interacted	Pooled	Interacted
At least one brother	0.4191** (0.109)	0.3839** (0.152)	0.3327** (0.089)	0.4106** (0.113)	0.1055 (0.074)	0.1564 (0.102)
Birth order	0.06714*** (0.037)	0.0668*** (0.037)	0.0353 (0.031)	0.0329 (0.031)	–0.1036 (0.026)	–0.0300 (0.033)
Per capita expend (million cedis)	4.30** (1.07)	4.19** (1.35)	4.00** (0.866)	4.23** (1.10)	1.86** (0.710)	1.10** (0.0910)
Mother's education	0.0037 (0.009)	0.0045 (0.009)	–0.0035 (0.0078)	–0.0036 (0.008)	–0.0044 (0.006)	0.0014 (0.009)
Father's education	0.0060 (0.008)	0.0061 (0.008)	0.0144 (0.007)	0.0147 (0.007)	0.0137** (0.005)	0.0086 (0.008)
Total siblings	–0.1404** (0.043)	–0.1170** (0.049)	–0.1122** (0.035)	–0.1098** (0.035)	–0.0209 (0.029)	0.0005 (0.033)
Sisters	0.1772* (0.082)	0.2302* (0.104)	0.1882** (0.067)	0.2514** (0.084)	0.1113* (0.055)	0.0654 (0.073)
Sisters squared	–0.01336 (0.015)	–0.0323*** (0.019)	–0.0115 (0.013)	–0.0224 (0.015)	–0.0061 (0.010)	–0.0022 (0.012)
Mother's height (m)	3.329** (0.604)	3.34** (0.776)	2.12** (0.491)	2.16** (0.498)	2.67** (0.461)	2.77** (0.498)
Father's height (m)	2.854** (0.679)	2.79** (0.083)	2.19** (0.550)	2.17** (0.552)	3.56** (0.780)	3.27** (0.422)
Female* one brother	–	0.0380 (0.222)	–	–0.1760 (0.151)	–	–0.0128 (0.150)
Female* total siblings	–	–0.0491 (0.051)	–	–0.0331 (0.081)	–	–0.0422 (0.035)
Female* sisters	–	–0.2107 (0.158)	–	–0.1508 (0.119)	–	0.1060 (0.108)
Female* sisters squared	–	–0.0545 (0.031)	–	0.0279 (0.026)	–	–0.1033 (0.021)
Female* per capita expenditure	–	–0.121 (1.85)	–	–0.588 (1.52)	–	–1.09 (1.27)
Observations	1427	1427	1427	1427	1427	1427
Pseudo R^2	0.2288	0.2306	0.2184	0.2203	0.0857	0.0915
Chi-square	431.36	436.36	376.23	378.89	127.87	131.13

Standard errors in parentheses. Additional variables include all sex and age interactions, Akan, Christian, traditional religion, rural residence, and dummy variables for missing parental height and education variables.
* Significant at the 5% level, ** significant at the 1% level, *** significant at the 10% level

Instrumental variables estimates

Table 4 gives results in which the number and sex of children ever born are used as instruments for the number and sex of children currently resident. The instruments control for endogeneity due to choices about who leaves and enters the household. To further show robustness of the results above, we include a quadratic term for the total number of siblings (and in-

Table 3. Probability of extreme health outcomes, probit estimates

Coefficient	Specifications					
	Stunting		Wasting		Underweight	
	Pooled	Interacted	Pooled	Interacted	Pooled	Interacted
At least one brother	–0.3588 ** (0.111)	–0.2471 (0.156)	–0.1152 (0.170)	–0.5824 * (0.261)	–0.4492 ** (0.112)	–0.5649 ** (0.164)
Birth order	–0.0806 ** (0.040)	–0.0840 * (0.040)	–0.0336 (0.054)	–0.042 (0.059)	–0.0833 * (0.041)	–0.0818 * (0.041)
Per capita expend (million cedis)	–3.45 ** (1.18)	–3.64 * (1.53)	–4.45 * (2.07)	– 4.77 * (2.33)	–5.10 ** (1.63)	–4.15 ** (1.76)
Mother's education	0.0055 (0.009)	–0.280 ** (0.068)	0.0048 (0.015)	0.0085 (0.015)	–0.0020 (0.010)	–0.0015 (0.011)
Father's education	–0.005 (0.008)	–0.3039 ** (0.066)	–0.0200 (0.014)	–0.0211 (0.014)	–0.0121 (0.009)	–0.0115 (0.009)
Total siblings	0.1700 ** (0.043)	0.1638 ** (0.048)	0.0573 (0.060)	0.0594 (0.068)	0.1685 ** (0.044)	0.1715 ** (0.053)
Sisters	–0.2291 ** (0.082)	–0.1810 *** (0.107)	–0.1080 (0.120)	–0.1282 (0.198)	–0.2104 ** (0.0811)	–0.2458 * (0.110)
Sisters squared	0.0226 (0.015)	0.01773 (0.018)	0.0087 (0.024)	0.0266 (0.028)	0.01116 (0.014)	0.0167 (0.017)
Mother's height (m)	–0.282 ** (0.66)	–0.280 ** (0.068)	0.0162 (0.09)	0.0447 (0.092)	–0.215 ** (0.06)	0.2200 ** (0.063)
Father's height (m)	0.3053 ** (0.66)	–0.3039 ** (0.066)	0.0116 (0.096)	0.0100 (0.100)	–0.328 ** (0.076)	0.3295 ** (0.063)
Female* one brother	–	–0.2206 (0.224)	–	0.8897 * (0.364)	–	0.2183 (0.236)
Female* total siblings	–	0.0188 (0.059)	–	0.0260 (0.090)	–	–0.0071 (0.062)
Female* sister	–	–0.0954 (0.167)	–	0.1980 (0.262)	–	0.0731 (0.168)
Female* sister squared	–	0.0084 (0.033)	–	–0.0946 (0.049)	–	–0.0121 (0.0311)
Female* per capita expenditure	–	0.399 (2.35)	–	0.470 (4.04)	–	–2.14 (2.38)
Observations	1427	1427	1427	1427	1427	1427
Pseudo R^2	0.1239	0.1247	0.0917	0.1132	0.1031	0.1045
log likelihood	–0.781.96	–781.25	–256.36	–250.30	–737.17	–0.736.01

Huber-corrected standard errors in parentheses. Additional variables include all sex and age interactions, Akan, Christian, traditional religion, rural residence, and dummy variables for missing parental height and education variables.
* Significant at the 5% level, ** significant at the 1% level, *** significant at the 10% level

strument it). All specifications are estimated with two-stage least squares, so we assume a linear probability model rather than the probit for the extreme health outcomes.

The qualitative results carry through to this specification, with coefficients for the most part being about the same size and taking the same signs as before. The standard errors, however, are larger. While the "sisters" variables are jointly significant in the specifications for height-for-age,

Table 4. Instrumental variables estimates

Coefficient	Specifications					
	Height-for-age	Weight-for-age	Weight-for-height	Stunting	Under-weight	Wasting
At least one brother	0.3039**	0.2480**	0.0803	–0.0772***	0.1323**	–0.0126
	(0.107)	(0.09)	(0.075)	(0.042)	(0.042)	(0.015)
Birth order	–0.0350	–0.0416***	–0.0216	0.0089	0.0075	–0.0036
	(0.029)	(0.024)	(0.021)	(0.012)	(0.01)	(0.003)
Per capita expend (million cedis)	4.14**	3.34**	1.01	–1.16**	–1.18**	–3.02**
	(1.15)	(0.095)	(8.05)	(0.45)	(0.44)	(1.22)
Mother's education	–0.0018	–0.0041	–0.0013	0.0025	–0.0002	–0.0002
	(0.010)	(0.008)	(0.007)	(0.004)	(0.003)	(0.001)
Father's education	0.0014	0.0103	0.0111***	0.0013	–0.0004	–0.0002
	(0.009)	(0.007)	(0.006)	(0.003)	(0.003)	(0.001)
Total siblings	–0.210	–0.198	–0.870	0.0360	0.1343*	0.0387*
	(0.169)	(0.140)	(0.118)	(0.062)	(0.058)	(0.018)
Total siblings squared	0.0345	0.0205	–0.0015	–0.0028	–0.003	–0.01343
	(0.024)	(0.02)	(0.017)	(0.009)	(0.002)	(0.019)
Sisters	0.1852	0.1163	0.1146	–0.0326	0.011	–0.019
	(0.194)	(0.161)	(0.136)	(0.071)	(0.071)	(0.025)
Sisters squared	0.0156	0.0213	0.0240	–0.0026	–0.0342***	–0.0026
	(0.052)	(0.043)	(0.037)	(0.019)	(0.019)	(0.008)
Mother's height (m)	2.82**	1.67**	0.086	–0.804**	–0.512**	0.0364
	(0.63)	(0.053)	(0.44)	(0.025)	(0.199)	(0.070)
Father's height (m)	3.64	2.60**	0.62	–1.170	–1.34**	0.0190
	(0.70)	(0.58)	(0.49)	(0.265)	(0.247)	(0.074)
Observations	1136	1136	1136	1136	1136	1136
R^2	0.255	0.232	0.096	0.127	0.1221	0.110
F(sisters, sisters squared) jointly significant	2.12***	3.40*	1.89	0.65	11.26**	0.73

Standard errors in parentheses. Additional variables include all sex and age interactions, Akan, Christian, traditional religion, rural residence, and dummy variables for missing parental height and education variables.
* significant at the 5% level, ** Significant at the 1% level, *** significant at the 10% level

weight-for-age, and being underweight, none is individually significant in any of the specifications (except for the quadratic in the specification explaining being underweight). The next section describes the size of the impacts.

Predicted impacts

Table 5 gives predicted effects for all-sister versus all-brother sibling groups. The top panel corresponds to the point estimates in Tables 2 and 3, while the bottom half corresponds to the estimates in Table 4. The experiment involves predicting each child's health status if all of their siblings were brothers versus if all were sisters. This provides a gauge of the magnitude of the estimated coefficients. Formally, the predictions follow equation (1) and fix all variables at the mean. But in the "all sisters" case we set

Table 5. Predicted effects of sibling composition on health indicators

Indicator	Sibling composition	Boys	% change	Girls	% change	Ratio M/F
Height-for-age	All brothers	–1.75	26.29	–1.39	25.18	1.259
	All sisters	–1.29		–1.04		1.240
Weight-for-age	All brothers	–1.60	29.38	–1.40	40.0	1.143
	All sisters	–1.13		–0.100		1.130
Weight-for-height[a]	All brothers	–0.713	37.31	–0.734	33.65	0.972
	All sisters	–0.447		–0.487		0.918
Stunting	All brothers	0.422	27.49	0.377	35.55	1.119
	All sisters	0.306		0.243		1.259
Wasting[a]	All brothers	0.039	–7.69	0.083	53.01	0.470
	All sisters	0.042		0.039		1.077
Underweight	All brothers	0.359	43.18	0.338	40.23	1.062
	All sisters	0.204		0.202		1.010
Height-for-age	All brothers	–1.42	30.9	–1.25	33.9	1.136
	All sisters	–0.983		–0.826		1.190
Weight-for-age	All brothers	–1.39	36.7	–1.26	39.0	1.103
	All sisters	–0.880		–0.768		1.146
Weight-for-height[a]	All brothers	–0.699	50.9	–0.634	54.2	1.103
	All sisters	–0.343		–0.290		1.183
Stunting[a]	All brothers	0.346	20.5	0.292	21.2	1.185
	All sisters	0.275		0.230		1.196
Wasting[a]	All brothers	0.043	62.8	0.063	60.3	0.683
	All sisters	0.016		0.025		0.640
Underweight	All brothers	0.274	68.2	0.258	66.7	1.062
	All sisters	0.087		0.086		1.012

[a] Coefficients on sisters and sisters squared not jointly significant in the base regressions or probits.

$S_i = N_i$ and $R_{ij} = 0$, and in the "all brothers" case we set $S_i = 0$ and $R_{ij} = 1$. The results presented are averages of predictions across the sample. Coefficients are taken from the pooled linear models and from the probits on extreme health outcomes.

The first and third columns of numbers give predicted outcomes, and the second and fourth give the percentage change in moving from the "all brothers" to "all sisters" scenarios. The final column gives the impact on the gender gap, with movement toward one indicating a narrowing of the gap.

The table shows that varying sibling composition can explain differences of roughly 25% for height-for-age, and 30% to 40% for weight-for-age and weight-for-height. The weight-for-height predictions have a wider range and are based on much more noisily measured coefficients. In general, the base impacts are large (and the impacts on extreme health outcomes are even larger). The final column shows a narrowing of the gender gap in the top panel (in moving to the "all sisters" case), but the size is modest with the exception of the noisy predictions of wasting. This is not surprising given the similarity of estimated impacts for boys and girls in Tables 2 and 3.

The predicted effects for the instrumented results range more widely than for the base results, but they are generally of similar magnitudes. The wider range for the base anthropometric measures (31% to 54%) is partly due to the increase in standard errors due to using a smaller sample, and it is partly due to the loss of precision inherent in two-stage estimation. Again, little impact is seen on the gender gap. The predicted impacts show a slight widening, but the magnitudes are small. The gender gap finding is mainly due to the lack of gender-specific interactions in Table 4, so that by construction there is apt to be little difference in gender-specific estimates in this set of predictions.

Disaggregation by income group

The impact of sisters is driven mainly by capital constraints, impacts should be stronger for households that face greater constraints – likely, poorer households. To investigate consistency with this proposition, we divide the sample at a per capita expenditure of 60,000 cedis, roughly the average annual per capita expenditure in the sample. Tables 6 and 7 give the results for the three main indicators (there are too few children with extremely poor health among the richer subsample to disaggregate the regressions for extreme health outcomes).

Table 6 shows a similar pattern of signs to those reported for the full sample in Table 2. However, precision is lower in the disaggregated regression due to smaller sample size. The general result remains, however: sisters matter and having more of them (while controlling for total sibling size) leads to improved health status.

The all-sister/all-brother comparisons in Table 7 show similar results for boys and girls. However, only for weight-for-height do we find that effects are noticeably greater for poorer households than richer: 45% improvement for poor boys versus 9% improvement for rich boys, and 38% improvement for poor girls versus 21% improvement for rich girls. However, the coefficients in the base regressions are noisily estimated, and, otherwise, the impact of sisters is notably larger for richer households than poorer (double the size for height-for-age). This suggests that richer households may not in fact be so unconstrained relative to the size of their desired investments – or that the previous results are driven as much by time constraints as capital constraints. The effects for per capita expenditure, on the other hand, take the predicted pattern, falling by over half for the richer sample.

The tables also show that spillover/reference effects are consistently smaller for richer groups than for poorer groups: they are measured with much more noise and are roughly half the size of effects in poorer groups for height-for-age and weight-for-age. They are larger, though, for weight-for-height. With the data at hand, we are not able to determine the reasons behind the differences, but they are consistent with multiple explanations. We would have predicted that the fact of diminishing marginal returns alone would yield smaller impacts on health at higher levels of investment than at lower levels, suggesting bigger effects for poorer children than richer. The results by total expenditure are thus equivocal and do not allow us to better pinpoint the underlying causes of the sibling rivalry.

Table 6. GLS – random effects estimates: health outcomes in richer and poorer households

Coefficient	Level specification					
	Height-for-age		Weight-for-age		Weight-for-height	
	Poorer households	Richest households	Poorer households	Richest households	Poorer households	Richest households
At least one	0.5318**	0.2994***	0.4299**	0.2081	0.0175***	0.0197
brother	(0.154)	(0.164)	(0.123)	(0.138)	(0.100)	(0.118)
Birth order	0.01600	0.1643*	0.0094	0.0613	–0.0001	–0.0345
	(0.046)	(0.70)	(0.037)	(0.059)	(0.029)	(0.051)
Per capita	9.54*	2.31	8.63**	2.34***	3.66	0.630
expenditure	(4.07)	(1.58)	(3.19)	(1.33)	(2.61)	(1.10)
Mother's	–0.1135**	0.0113	–0.0080	0.0022	–0.0081	–0.0025
education	(0.012)	(0.015)	(0.010)	(0.013)	(0.008)	(0.011)
Father's	0.0091	0.0070	0.0190	0.0110	0.0017	0.0083
education	(0.011)	(0.012)	(0.009)	(0.010)	(0.007)	(0.0086)
Total siblings	–0.1028*	–0.1920**	–0.1001*	–0.0781	–0.0517	0.0504
	(0.053)	(0.082)	(0.041)	(0.069)	(0.034)	(0.059)
Sisters	0.1891***	0.1489	0.1508 ***	0.2069 ***	0.0540	0.1567***
	(0.106)	(0.131)	(0.084)	(0.111)	(0.069)	(0.093)
Sisters squared	–0.0199	0.0033	–0.0067	–0.0163	0.0078	–0.0310
	(0.020)	(0.025)	(0.016)	(0.021)	(0.0129)	(0.018)
Mother's height	2.50**	4.75**	1.41*	3.24**	–0.2212	0.524
(m)	(0.75)	(1.2)	(0.59)	(0.88)	(0.487)	(0.72)
Father's height	3.16**	2.76**	2.86**	1.77*	1.23*	0.250
(m)	(0.92)	(1.0)	(0.71)	(0.86)	(0.590)	(0.704)
Observations	857	570	857	570	857	570
Adjusted R^2	0.2482	0.2231	0.2482	0.1978	0.1217	0.1012
Chi-square	275.75	173.68	260.93	130.62	105.50	52.55
Number of households	528	380	528	380	528	380

Standard errors in parentheses. Additional variables include all sex and age interactions, Akan, Christian, traditional religion, rural residence, and dummy variables for missing parental height and education variables.
* Significant at the 5% level, ** significant at the 1% level, *** significant at the 10% level.

6. Conclusions

The economic models that lie behind nearly all studies of human capital investment suggest that sibling composition can affect investment patterns. The evidence from Ghana shows how sibling composition can matter quantitatively. We find that having more sisters (holding sibling size constant) substantially raises the health status of children. In the benchmark case, moving from a hypothetical case in which children have all brothers to one in which they have all sisters increases predicted health outcomes by roughly 25% to 40%.

While impacts are large overall, there is little evidence of strong differences between impacts by gender. In economies in which boys are favored by parents, both boys and girls benefit from having more sisters: boys because they can draw away resources from their sisters and girls because sis-

Table 7. Predicted effects of sibling composition on health indicators in richer and poorer households

Indicator	Sibling composition	Poorer		% Change		Richer		% Change	
		Boys	Girls	Boys	Girls	Boys	Girls	Boys	Girls
Height-for-age[b]	All brothers	–1.76	–1.62	19.3	21.0	–1.55	–1.31	36.8	38.2
	All sisters	–1.42	–1.28			–0.98	–0.81		
Weight-for-age	All brothers	–1.60	–1.55	26.2	25.2	–1.46	–1.27	30.8	33.8
	All sisters	–1.18	–1.16			–1.01	–0.084		
Weight-for-height[a]	All brothers	–0.74	–0.76	45.6	38.2	–0.65	–0.57	9.2	21.1
	All sisters	–0.40	–0.47			–0.59	–0.45		

[a] Coefficients on sisters and sisters squared not jointly significant at the 10% level in regressions for the poorer sample.
[b] Coefficients on sisters and sisters squared not jointly significant at the 10% level in regressions for the richer sample.

ters help share burdens and offer less competition. Theory yields ambiguous predictions on whether boys should benefit more than girls, and the lack of difference suggests that (in the relevant range) the shape of returns functions is fairly similar by gender.

Interpretation of the results has started with the implicit assumption that parents favor sons over daughters. The sort of inequalities described here are typically hidden in simple tabulations of health status by gender, and the results thus reveal some of the non-obvious ways in which gender biases are manifested. While sibling rivalry does not explain differences in average health outcomes by gender in Ghana, it helps to explain inequality in health outcomes within gender groups. Mitigating the initial biases should go far in reducing the resulting rivalries. In turn, reducing the rivalries should help narrow these inequalities.

We predict that the rivalries will also be reduced by lifting the time and credit constraints faced by parents. Improving market performance could have substantial effects on health levels even if everything else about households were to remain unchanged, including lifetime resources. The available data, however, offer no way to determine the specific source of the constraints. Results from a poorer sub-sample compared to results from a richer sub-sample do not support a simple credit constraint story. Additional data are required to better pin down the economic structures that explain why having relatively more sisters appears to be such an advantage in Ghana. Additional data can also help address potential biases introduced by endogenous household formation.

Appendix

Table 1 A. Means and standard deviations of variables in regressions

Health variable	Full sample		Girls		Boys	
	Mean	Std dev.	Mean	Std dev.	Mean	Std dev.
Sex (male=1)	0.539	0.498	–	–	–	–
Total siblings	3.48	2.32	3.33	2.28	3.60	2.35
Number of brothers	1.877	1.66	1.807	1.66	1.93	1.66
Number of sisters	1.605	1.32	1.525	1.25	1.67	1.37
Percentage sisters	0.481	0.332	0.484	0.338	0.479	0.327
Per capita expenditure	59,717.9	37,621	60,387	36,536	59,146	38,528
Mother’s education	3.557	4.555	3.737	4.4492	3.132	4.478
Father’s education	6.798	5.789	6.7887	5.2367	6.8702	5.675
Mother’s education missing	0.2563	0.4367	0.2453	0.4305	0.2670	0.4427
Father’s education missing	0.2237	0.2237	0.2241	0.4172	0.2234	0.4167
Age (years)	4.75	2.98	4.36	2.78	5.089	3.112
Have at least one brother	0.815	0.387	0.8044	0.396	0.825	0.379
Urban residence	0.444	0.430	0.443	0.429	0.446	0.430
Christian religion	0.5442	0.498	0.5728	0.494	0.5169	0.5000
Traditional religion	0.2566	0.2566	0.2377	0.426	0.2748	0.4466
Akan tribe	0.490	0.500	0.492	0.500	0.488	0.500
Height-for-age	–1.34	1.39	–1.23	1.42	–1.43	1.36
Weight-for-age	–1.28	1.11	–1.21	1.13	–1.34	1.08
Weight-for-height	–0.657	0.896	–0.6108	0.909	–0.623	0.884
Stunted	0.308	0.461	0.272	0.445	0.3388	0.473
Wasted	0.0523	0.222	0.0625	0.242	0.0436	0.204
Underweight	0.2647	0.441	0.243	0.429	0.283	0.450
Mother’s height	1503.29	349.17	1518.59	311.42	1488.17	381.66
Father’s height	1276.26	727.75	1286.96	720.27	1265.93	735.14
Mother’s height missing	0.0495	0.217	0.0387	0.1928	0.0601	0.2378
Father’s height missing	0.2441	0.4297	0.2371	0.4255	0.2507	0.434
Health sample	2458	2458	1133	1133	1325	1325

Exchange rate in 1989 \$1 = C250

Table 2 A. Predicted effects of sibling compositions on health outcomes "reference effect" excluded

Indicator	Sibling composition	Boys	% Change	Girls	% Change
Height-for-age	All brothers	–1.405	29.25	–1.572	9.41
	All sisters	–0.994		–1.424	
Weight-for-age	All brothers	–1.419	8.81	–1.439	37.1
	All sisters	–1.294		–0.905	
Weight-for-height	All brothers	–0.6202	12.5	–0.7706	47.1
	All sisters	–0.5426		–0.4071	
Stunting	All brothers	0.3806	33.6	0.3375	24.1
	All sisters	0.2526		0.2562	
Wasting	All brothers	0.0364	–17.9	0.0776	48.6
	All sisters	0.0429		0.0399	
Underweight	All brothers	0.2868	19.5	0.3058	28.2
	All sisters	0.2310		0.2196	

Table 3 A. Predicted effects of sibling composition on health indicators by richer and poorer, "reference effect" excluded

Indicator	Sibling composition	Poorer		% Change		Richer		% Change	
		Boys	Girls	Boys	Girls	Boys	Girls	Boys	Girls
Height-for-age	All brothers	–1.622	–1.762	20.8	19.5	–1.545	–1.307	36.5	37.9
	All sisters	–1.284	–1.419			–0.9805	–0.8117		
Weight-for-age	All brothers	–1.554	–1.598	25.2	26.3	–1.458	–1.274	30.5	34.0
	All sisters	–1.163	–1.178			–1.013	–0.8406		
Weight-for-height	All brothers	–0.7363	–0.7592	45.6	38.7	–0.6510	–0.5724	9.7	20.5
	All sisters	–0.4006	–0.4657			–0.5880	–0.4548		

Endnotes

1 Papers in economics and sociology that consider the role of siblings in human capital allocation are Parish and Willis (1994) for Taiwan; Lillard and Willis (1994) for Malaysia; and, for the United States, Powell and Steelman (1989), Butcher and Case (1994), Kaestner (1996), and Hauser and Kuo (1996). These papers focus on educational attainment. Muhuri and Preston (1991) and Das Gupta (1987) consider sibling composition and excess female mortality in South Asia. Rose (1997) considers links between demographics, markets, and investments that parallel the discussion here.

2 If marriage markets functioned perfectly, parents should be able to recoup the full returns to investments in the human capital of daughters, but in practice, bride prices and dowries value human capital only imperfectly. In cultural groups with matrilineal structures, daughters may retain close connections with their families after marriage and, especially, after divorce. The effects we find here give average effects across cultural groups.

3 We assume that the returns to investing in any girl or any boy is the same. In the empirical section, we consider the roles of age and birth order. Alternatively, returns functions might feature a convex and then concave portion. This can explain sharp differences between investments in boys and girls, but it is not consistent with our findings that girls and boys benefit in similar ways from having more sisters. Note that here we abstract from time inputs in the returns functions.

[4] The presence of non-convexities in returns reinforces the ambiguity of theoretical predictions. Imagine that there are two health technologies. One has a modest return (e.g., using local, traditional healers), while the other may have higher returns but a sizable fixed cost (e.g., visiting a medical doctor). In poor households, the fixed investments may be so great that all children are treated with the traditional methods. When households get richer – and boys have even a slight advantage over girls – gender gaps will emerge as boys get the "non-convex" treatment while sisters must suffice with the traditional one. Eventually with increasing wealth, all children will be treated with the non-convex treatment and the gender gap will narrow. While this scenario is plausible, our data do not indicate that it is a likely case.

[5] Alternative measures of acute and chronic malnutrition have been proposed. The alternative criteria for low nutrition are to use 90 percent of the median as a cutoff point for height-for-age and 80% for weight-for-age. The standard deviation measure we employ is age invariant and thus is preferable to the other criteria. Alderman (1991) provides a discussion in the Ghanaian context.

[6] Correlations of sibling composition and total income (and total expenditure) may occur if there are sex-based differences in child labor or parental labor supply response. We control for total expenditures in the regressions; thus, the "sibling composition" variables reflect impacts on the allocation of given resources, rather than effects on the size of overall resources. The bivariate analyses combine both effects and yield similar results.

[7] The dummy variable may also pick up non-linear effects of sisters (beyond the quadratic term). The qualitative results remain when we drop the "one brother" dummy variables from the regressions, with the predicted effects showing greater variability but falling in the same general range (Appendix Tables 2 and 3).

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