**Sibling Rivalry in Africa**

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# Sibling Rivalry in Africa

*By* JONATHAN MORDUCH*

The most basic neoclassical theory of human capital holds that, when markets are complete, parents will educate their children up to the point at which the expected marginal return equals the marginal cost. The theory thus relates schooling levels, expected wages (and nonpecuniary returns), and the costs of schooling. In practice, schooling levels are apt to be determined by a much broader range of factors, including the demographic composition of the students' households (Kristen Butcher and Anne Case, 1994). For example, as William Parish and Robert Willis (1993) show using data from Taiwan, having sisters can be an important predictor of schooling outcomes. They find that older sisters in particular confer benefits by helping to care for younger children; older sisters are also more likely to take wage employment that helps to pay school fees and allows younger children to postpone employment. Robert Kaestner (1996) similarly finds that completed schooling would hypothetically rise by 0.77 years in a sample of African-American men if they had all sisters rather than all brothers (although the causal mechanism is not clear). Turning to child health, Ashish Garg and Morduch (1998a) find predicted improvements of as much as 40 percent in anthropometric measures of children in Ghana when shifting from a scenario where all siblings are brothers to one in which all are sisters. Garg and Morduch (1998b) also find that predicted enrollments in secondary school are increased by more than 50 percent when shifting from the all-brothers to all-sisters scenario but find negligible impacts on primary-school enrollments.

This paper uses data on young teenagers to investigate how sibling composition affects schooling outcomes in South Africa and Tanzania. Completed years of schooling are explained as a function of sibling composition and a variety of household and individual characteristics. In the data from Tanzania, moving from an all-brothers to all-sisters scenario raises completed years of school by 0.44 years, while the data from South Africa show very limited effects of sibling composition. The Tanzania sample shows no strong difference between advantages for boys or girls, and having older sisters confers no particular advantages.

While the present data do not allow a precise test of alternative theories of human-capital investment, the results are consistent with a situation in which sons are favored by parents and households face constraints in the time or financial resources available to expend on their children's education. This sets in motion rivalry for scarce resources in which parents favor sons, and children do better with sisters. This resource-based view of sibling rivalry is similar to personality-based theories in explaining heterogeneity in outcomes within families, and it suggests a strong role for demography in explaining human-capital investment.

It is not the only way to explain the results, however. Even where resource constraints are not binding, gender-specific nonconvexities in returns functions could lead to asymmetries of the sort found here. The results might also emerge from demonstration effects or social pressures, where the presence of sisters promotes attention to education. Further, the results could arise as statistical artifacts arising from the endogeneity of household structures. By this argument, the presence of a relatively greater number of sisters would need to be correlated with omitted determinants of education.

The objective here is modest but fundamental: to describe basic patterns in the data that remain poorly appreciated. Additional steps (and additional data) will be needed to disentangle causal effects.

## I. Methods and Data

The association of sibling composition and school outcomes is quantified through a series

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of ordinary least-squares (OLS) regressions, where the dependent variable is the number of years of completed schooling, and indicator variables for age (by year) and sex are included on the right-hand side. A similar specification, augmented with school quality measures, is used by Anne Case and Angus Deaton (1999) to analyze South African schooling data.

The main explanatory variables are the number of sisters and the number squared, with alternative specifications as shown in the tables. All specifications also include the total number of siblings (and the square), birth order (i.e., the number of older siblings), household head's education, household size, and indicator variables for sex, age (by year), region, urban residence, and whether the household head is female. Similar results were found when estimating tobit models instead.

The focus is on the conditional expectation of the role of having sisters, taking family structure as given. The number of siblings is tallied as the number of biological brothers and sisters of each child living in the household. All other residents are included in the household-size measure.

The samples are restricted to children aged 13–16. The lower limit of the samples (age 13) was chosen to include children at a stage in their schooling where variations in outcomes emerge, while the upper bound (age 16) was chosen to limit the extent to which older siblings have left the household. An advantage of considering these issues in the sub-Saharan African context is that there is little evidence of either gender-biased mortality or fertility patterns, so the fraction of sisters is not apt to be correlated strongly with the number of siblings. The correlation is 0.02 in the Tanzania sample and 0.04 in the South Africa sample. There still may be concern with gender-selective child-fostering, migration, and marriage patterns, and these issues are discussed further in Section III.

The data from South Africa come from the 1993 South Africa Integrated Household Survey, completed just prior to the elections of 1994 which brought Nelson Mandela to power. The survey covers about 9,000 households drawn from across the country and was managed by the South Africa Labour Development Research Unit (SALDRU) of the Department of Economics at the University of Cape Town, with assistance from the World Bank and funding from the governments of Denmark, the Netherlands, and Norway. The data from Tanzania come from a nationally representative survey of 5,000 households, also completed in 1993. The survey was a joint effort undertaken by the Department of Economics of the University of Dar es Salaam, the Government of Tanzania, and the World Bank and was funded by the World Bank, the government of Japan, and the British Overseas Development Agency. Like the Ghana data used by Garg and Morduch (1998a, b) the surveys are part of the World Bank's Living Standards Measurement Survey (LSMS) program, and information on obtaining the data can be obtained from the LSMS program at the World Bank.

## II. Results

The first column of Table 1 shows a positive association of having sisters and schooling outcomes in Tanzania. The coefficient on the number of sisters is statistically significant with 90-percent confidence, as are both coefficients jointly. The negative coefficient on the quadratic term shows that the average advantage falls slowly with the number of sisters, but it remains substantial over the relevant ranges.

A simple calculation illustrates the estimated magnitudes. Imagine that all children had four siblings (the median for the sample) and hold all else the same. Completed years of education would be raised by 0.44 years by shifting from a scenario in which all four siblings were brothers to one in which all four were instead sisters. To put the magnitude in perspective, compare it to the coefficient on the female dummy, which picks up the first-order effect that girls tend to get more schooling in Tanzania (after controlling for other variables): being a girl raises completed years of schooling by 0.37 years. The magnitude of the sisters variables are thus relatively large, given that they reflect second-order effects.

The second column sheds light on the proposition that it is the presence of older sisters that drives the result in the first column. The coefficient on the number of older sisters is negative but relatively small, and it has a low $t$ statistic. The coefficient on the number of sisters (both older and younger) is little affected. The third

TABLE 1—EFFECTS OF SIBLING COMPOSITION ON COMPLETED EDUCATION, TANZANIA, 1993

| | Regression | | |
|---|---|---|---|
| Independent variable | (i) | (ii) | (iii) |
| Number of sisters | 0.15 | 0.19 | 0.12 |
| | (1.67) | (1.86) | (1.01) |
| Number of sisters squared | −0.01 | −0.02 | −0.01 |
| | (0.71) | (1.07) | (0.45) |
| Number of older sisters | | −0.09 | −0.05 |
| | | (0.71) | (0.36) |
| Number of older sisters squared | | 0.05 | 0.04 |
| | | (1.06) | (0.82) |
| Female × number of sisters | | | 0.16 |
| | | | (0.85) |
| Female × number of sisters squared | | | −0.03 |
| | | | (0.70) |
| Female × older sisters | | | −0.04 |
| | | | (0.06) |
| Female × older sisters squared | | | −0.02 |
| | | | (0.05) |
| Number of observations: | 2,104 | 1,806 | 1,806 |

*Notes:* The dependent variable is completed years of education of children aged 13–16. OLS regression coefficients are shown, with robust absolute values of $t$ statistics (in parentheses). Specifications include total number of siblings (and the square), birth order, household head's education, household size, and indicator variables for sex, age (by year), region, urban residence, and whether the head is female.

TABLE 2—EFFECTS OF SIBLING COMPOSITION ON COMPLETED EDUCATION, SOUTH AFRICA, 1993

| | Regression | | |
|---|---|---|---|
| Independent variable | (i) | (ii) | (iii) |
| Number of sisters | −0.003 | −0.04 | −0.002 |
| | (0.03) | (0.34) | (0.01) |
| Number of sisters squared | 0.01 | 0.007 | 0.002 |
| | (0.43) | (0.28) | (0.06) |
| Number of older sisters | | 0.06 | 0.15 |
| | | (0.49) | (1.06) |
| Number of older sisters squared | | 0.02 | −0.02 |
| | | (0.57) | (0.52) |
| Female × number of sisters | | | −0.05 |
| | | | (0.27) |
| Female × number of sisters squared | | | 0.005 |
| | | | (0.15) |
| Female × older sisters | | | −0.23 |
| | | | (1.27) |
| Female × older sisters squared | | | 0.09 |
| | | | (2.14) |
| Number of observations: | 1,625 | 1,625 | 1,625 |

*Notes:* The dependent variable is completed years of education of children aged 13–16. The sample is restricted to the African population. OLS regression coefficients are reported, with robust absolute values of $t$ statistics (in parentheses). Specifications include control variables in Table 1 and log of average monthly household expenditures.

column is ambitious in attempting to tease out independent impacts by gender as well. Adding the interacted variables reduces the measured association with having sisters (both older and younger), and the precision of the point estimates falls sharply. None of the variables added in rows 3–8 is large, and none has a $t$ statistics greater than 1.0.

Taken together, the data from Tanzania suggest a substantial positive association of educational outcomes and the number of sisters that a child has, holding constant the total number of siblings. The association does not vary significantly with the sex of the child or with whether sisters are older or younger. The precision of the estimates is less inspiring, though: none of the main coefficients is statistically significant with 95-percent confidence.

The data from South Africa allow an additional view into the role of sibling composition on educational outcomes in sub-Saharan Africa. The data set reflects the realities of South Africa under apartheid, and the observations are divided into four categories by race: "African," "Colored," "Indian," and "White." The distinction is followed here, with capital letters used to denote that these are official categorizations, important in signaling structural asymmetries that affect livelihoods, aspirations, access, and other elements that influence educational outcomes. The focus will be on Africans, comprising 73 percent of children aged 13–16 in the survey.

Table 2 gives the results in a form that mirrors Table 1. The dependent variable is again the years of completed education, but the independent variables now also include the logarithm of per capita monthly expenditures, for which data were not available from Tanzania. The argument for inclusion is that total

household expenditures are often a strong predictor of education outcomes, and exclusion can yield omitted-variable bias. An argument for exclusion is that part of the beneficial roles of siblings (via working to add to household resources, for example) will be absorbed by the expenditure variable, leaving a possible misimpression of the net role that sisters play (relative to brothers). By this argument, expenditures should be left out of the reduced-form equation. Quantitatively, though, including the expenditure variable makes a negligible difference to the point estimates, just as in the Garg-Morduch results from Ghana.

Unlike the results from Tanzania, the point estimates in the first column of Table 2 show no strong association between the number of sisters and schooling outcomes, and the second column shows that the story is unchanged when breaking out the role of older sisters. The result is a function of restricting attention to young teenagers. When the same regression is run on a sample of children aged 13–20, the coefficient on the number of sisters changes little; the coefficient on older sisters rises, however, to 0.24 and is statistically significant with 95-percent confidence (not reported in the tables). This is a large and striking result. But it may be "too large" to be fully believable as stemming just from causal relationships. The coefficient may be partly picking up omitted variables: in particular, evidence that a child in the late teens still has older sisters at home may signal that the household is one in which daughters tend to marry later and where education is especially privileged. The results on the younger teenagers, presented in Table 2, are thus likely to be more informative.

The third column, which breaks out the independent associations with females, also shows little in this sample. The final two rows yield a surprisingly large negative coefficient on the interaction of being a girl and having older sisters (although the standard error is also large) and a large, significant positive coefficient on the quadratic term. However, the sibling composition variables, taken as a whole, are not jointly significant.

## III. Discussion

The results from Tanzania, while not estimated very precisely, establish additional evidence of positive associations between school completion and the number of sisters a child has (controlling for the total number of siblings), but the evidence from South Africa shows that they are not general findings.

The estimates are conditional on the given family structure, and of course, family structure may not be fully exogenous to schooling choices. There is no evidence that the number of living sisters (controlling for total siblings) is endogenous in these samples, but parents may make fertility choices with trade-offs between the quality and quantity of children in mind. Migration or child-fostering arrangements may also partly hinge on concerns with schooling. Based upon detailed fertility histories not available here, Garg and Morduch (1998a, b) use as instruments the number and gender of children ever born (rather than children currently in the household) and find little qualitative change to their results, although the point estimates are less precisely measured under the two-stage procedure. This addresses migration and fostering, but not quality–quantity trade-offs. Satisfactorily addressing the latter requires finding variables that determine fertility choices but not education choices, and none are obvious here; convincingly identifying causal effects will thus require other data.

Investigating the timely completion of primary education in sub-Saharan Africa places attention on determinants of delay and repetition, as well as on dropouts. A more complete analysis would be aided by information on adults and their siblings (as in the data used by Butcher and Case [1994]), but data on siblings are only available for children. Still, data are available on current enrollment and on the last year of enrollment, and these can be used in future work to partially distinguish delay and repetition from dropouts.

Data on the siblings of adults would also allow sharper tests of propositions regarding older sisters, since the number of resident older siblings falls with age and some of the effects here will be absorbed by the age dummies. With the available data though, little evidence was found to suggest that it is the presence of older sisters that drives the results in Ghana or Tanzania.

The next step is to better understand the specific mechanisms through which these house-

hold-level demographic factors affect patterns of human-capital accumulation. Increasing education levels have been accorded great importance in sub-Saharan Africa, and effective policy will depend on more fully understanding the nature of schooling decisions. To extent that the model of sibling rivalry holds, equitably raising education levels will require relaxing resource constraints, and not just raising the perceived returns to education.

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