MATERNAL EDUCATION AND CHILD HEALTH: IS THERE A STRONG CAUSAL RELATIONSHIP?*

SONALDE DESAI AND SOUMYA ALVA

Using data from the first round of Demographic and Health Surveys for 22 developing countries, we examine the effect of maternal education on three markers of child health: infant mortality, children's height-for-age, and immunization status. In contrast to other studies, we argue that although there is a strong correlation between maternal education and markers of child health, a causal relationship is far from established. Education acts as a proxy for the socioeconomic status of the family and geographic area of residence. Introducing controls for husband's education and access to piped water and toilet attenuate the impact of maternal education on infant mortality and children's height-for-age. This effect is further reduced by controlling for area of residence through the use of fixed-effects models. In the final model, maternal education has a statistically significant impact on infant mortality and height-for-age in only a handful of countries. In contrast, maternal education remains statistically significant for chidren's immunization status in about one-half of the countries even after individual-level and community-level controls are introduced.

Following a pioneering paper by Caldwell (1979), it has often been argued that, *ceteris paribus*, children of educated mothers experience lower mortality than do children of uneducated mothers. Relying on the many demographic studies that demonstrate a strong correlation between maternal education and child health, public policy discourse has increasingly assumed that investments in women's education are important for lowering infant and child mortality and improving child health (Schultz 1993; World Bank 1993).

Nonetheless, a few scholars are uncomfortable about concluding from this correlation that there is a strong causal relationship between maternal education and child health (Basu 1994; Hobcraft 1993; Kunstadter 1995). Unfortunately, their discomfort has been largely ignored in most of the policy discussions because the link between maternal education and child health has rarely been analyzed using appropriate statistical models. Even scholars who point out the lack of conclusive evidence supporting a strong maternal education/child health link (Hobcraft 1993) do so rather hesitantly (Hobcraft 1994). This hesitation stems from a lack of appropriate control variables in large data sets such as the World Fertility Surveys and Demographic and Health Surveys.

In contrast to prior research, using community-level fixed-effects models, we argue that maternal education may be a proxy for the socioeconomic status of the household as well as for characteristics of the community of residence. Hence, we consider the possibility that the observed correlation between maternal education and various markers of child health may be spurious. Using data from the first round of Demographic and Health Surveys, we demonstrate that controls for a few socioeconomic variables and for community of residence substantially attenuate the maternal education/child health link. Based on these results, we argue that the relationship between maternal education and child health is considerably weaker than is commonly believed. Consequently, public policies emphasizing investments in women's education as an important avenue for improving child health are inefficient, although such investments may have other societal benefits.

We also distinguish between parental health-seeking behavior—measured by children's immunization status—and health outcomes measured by infant mortality and children's height-for-age. Our results show that whereas introduction of individual-level and community-level controls significantly attenuates the relationship between maternal education and child health, the link between maternal education and child immunization status remains fairly strong. Thus, educated mothers are more likely to engage in health-seeking behavior, but its impact on actual health outcomes seems to be rather weak, possibly because the impact of environmental conditions supersedes the impact of parental behavior in shaping child health.

PAST RESEARCH

Much recent demographic research on child health focuses on maternal education as a precursor to improved child health. Support for this hypothesis has come from large cross-national surveys conducted under the rubric of the World Fertility Surveys and the Demographic and Health Surveys (Bicego and Boerma 1991; Boerma, Sommerfelt, and Rutstein 1991; Cleland and Kaufmann 1993; Hobcraft, McDonald, and Rutstein 1984; Mensch, Lentzner, and Preston 1985).

Virtually all studies based on these two large data sets have shown strong correlations between maternal education and child health and survival. However, given the close link between education and other favorable socioeconomic con-

*Sonalde Desai and Soumya Alva, Center on Population, Gender and Social Inequality, University of Maryland at College Park, 3114 Art-Sociology Building, College Park, MD 20742-1315; e-mail:sdesai@bss1.umd.edu. This research was supported in part by funds from The William and Flora Hewlett Foundation to the Center on Population, Gender and Social Inequality, including a traineeship for Soumya Alva. We thank Karen Mason, Susheela Singh, and Joan Kahn for comments on an early draft.

ditions, researchers vary in their propensity to move beyond this correlation to argue that maternal education *causes* low child mortality. Although a majority of scholars working in this area unequivocally state that maternal education has a strong independent effect on child health and survival (Caldwell 1994; Cleland and Kaufmann 1993; World Bank 1993), a minority are less sure about the causal nature of this relationship and advocate caution (Hobcraft 1993).

Why do we expect maternal education to improve child health? Education is linked to family socioeconomic situation, which in itself is a determinant of child health. But above and beyond this, maternal education is hypothesized to bring about certain changes in individual behavior that result in better child health (Cleland 1990). A recent review of the literature by Caldwell and Caldwell (1993) suggests two potential paths: (1) education improves child health solely by enhancing the use of modern health services; and (2) education results in a wide range of favorable behaviors—mostly connected with child care—that play a role in improving child health.

Empirical evidence for these paths, however, is rather weak and mixed. For example, if maternal education has a positive effect on child health by increasing the use of medical services, a greater effect of maternal education should be observed in urban areas, where better quality services are located. However, empirical research shows that the effect of maternal education is greater in rural areas than in urban areas (Caldwell and Caldwell 1993; Schultz 1993). Similarly, arguments regarding the reduction in diarrheal deaths with better implementation of oral rehydration therapy by educated mothers fail to focus on the importance of this behavior to actual mortality. A study sponsored by the World Health Organization (Victora et al. 1993) suggests that although oral rehydration therapy reduces deaths due to dehydration for children with acute diarrhea, it fails to reduce deaths due to chronic diarrhea or dysentery.

Historical research on mortality transition adds a further note of caution. Two arguments deserve special attention. First, compositional analysis conducted by Cleland, Bicego, and Fegan (1991) shows that between the 1970s (when World Fertility Surveys were conducted) and the 1980s (when Demographic and Health Surveys were conducted), female education increased and infant mortality decreased in 12 developing countries with comparative data. However, this decline in mortality occurred in all education groups with the result that differences in child mortality by maternal education remain relatively constant. Moreover, only a modest percentage of the decline in child mortality can be attributed to increasing education as reflected in the compositional change. Ten percent or less of the mortality decline in 7 of 12 countries studied can be attributed to increasing education (Cleland et al. 1991).

Second, historical research documents situations in which high maternal education does not lead to differential improvements in child health; that is, improvements in mortality seem to occur independent of improvements in female education. Preston and Haines (1991) found few differences by maternal education in infant mortality patterns in the United States at the turn of the century.[1] Conversely, Kunstadter (1995), examining changes in mortality among Hmong in Thailand, shows that mortality can decline without concurrent improvements in education.

In light of these caveats, it is important to reevaluate the link between maternal education and child health and to control for some key explanatory variables. In particular, it is important to recognize that educated mothers come from high socioeconomic strata and that controlling for socioeconomic status of the family is likely to attenuate the impact of maternal education on child health. Educated mothers also tend to live in more economically developed areas. Areas that are rich and powerful enough to have schools are probably also rich and powerful enough to have medical facilities (Palloni 1981). Given the disproportionate concentration of educated women in cities and other advantaged regions, maternal education may be a proxy for better health systems as well as water and sanitation systems in intracountry comparisons.

DATA AND METHODOLOGY

The empirical results presented here are based on data from 22 Demographic and Health Surveys (DHS) in different parts of the developing world. Following a stratified cluster-sampling design, these surveys randomly selected women aged 15–49 within each cluster.[2] In some countries, the sample was restricted to ever-married women; in others, it included all women of reproductive age. Information was also collected regarding their fertility history, maternity care, and some aspects of their socioeconomic background. In addition, a health history was collected for children aged 5 and under and anthropometric data were collected for children aged 36 months and under.

Although the socioeconomic information collected by DHS is rather limited, these surveys are among the few large cross-national surveys to collect health histories for young children. This makes DHS ideal for comparing determinants of child health across a diverse range of countries. In addition, DHS relies on cluster sampling. Although sampling procedures varied across countries, on average about 10–50 children aged 12–120 months from each cluster are included in our sample. We take advantage of this cluster-sampling procedure to control for unobserved community-level factors by estimating cluster-level fixed-effects models.

1. Preston and Haines (1991) argue that because germ theory was not well known at the turn of the century, educated mothers could do little to reduce mortality. Similar arguments could be advanced for developing countries. In countries with water shortages and difficulties in finding fuel to boil water, theremay be little that educated mothers can do to protect thir children by following better hygienic practices. Preston and Haines (1991) also suggest that modern medicine had little to contribute toward the decline of mortality, particularly when compared to the contributions of public health programs.

2. Cluster sampling introduces within-cluster correlations among variables of interest, thereby reducing the efficiency of the sample and increasing the standard errors of the coefficients. Few studies using these data correct for this and, hence, tend to overestimate statistical significance.

Because the community information in the DHS is rather limited, direct controls for characteristics of the communities, except for urban/rural residence, is not possible for most countries. Hence, we rely on an alternative strategy: We compare women of different educational levels in each cluster and thereby examine education effects net of community effects.[3]

Our strategy underestimates the effects of education because it ignores migration selectivity. Individuals with high levels of education are more likely to migrate to communities with better amenities. Moreover, a control for a community effect in this analysis also controls for some individual characteristics, particularly race/ethnicity and income. However, these shortcomings do not affect our central purpose. Although these indirect effects are important for a study of individual-level determinants of child health, our primary purpose is to examine the validity of investments in maternal education as a strategy for improving child health. Because public policy cannot rely on migration—particularly urbanization—to improve child health, we argue that our approach of focusing on the effect of education net of indirect effects via residential mobility is reasonable.

We focus on three dependent variables: probability of death within the first year, height-for-age, and number of vaccinations received by the index child aged 12 months and over. Sample means and standard deviations of the variables for each country are presented in Table 1. For all analyses the sample is restricted to births occurring 12 to 120 months before the survey. However, the sample for each outcome variable is different: Infant mortality analyses focus on children born within 10 years of the survey date, analyses of height-for-age rely on children between the ages of 12 and 36 months, and immunization analyses are conducted for all children between 12 and 60 months for whom the health card is available.[4] Given our focus on within-cluster variation, we excluded clusters in which only one child was available for any given analysis.[5] We also excluded children whose mothers had not lived in the present cluster since the birth of the child. Because parental migration is related to child mortality (Brockerhoff 1990), this selection criterion biases our results somewhat. Additional selection bias is attributable to the omission of dead mothers—and their children—from our sample. Thus, our focus is to examine changes in the coefficient for maternal education within this particular sample as successive levels of controls are introduced.

For each dependent variable, we present three models. To take account of within-cluster correlation, the standard errors for Models 1 and 2 are adjusted using the Huber (1967) procedure.

Model	Estimation Technique	Independent Variables
Model 1	OLS or logit, standard errors adjusted for intracluster correlations	Mother's education, child's age, urban residence
Model 2	OLS or logit, standard errors adjusted for intracluster correlations	Mother's education, child's age, urban residence, father's education, mother does not have a partner, water, toilet
Model 3	Fixed effects	Mother's education, child's age, father's education, mother does not have a partner, water, toilet (water and toilet variables dropped for the countries in which there is no variation within clusters on these variables)

Fixed-effects linear regressions are easily estimated for continuous dependent variables using ordinary-least-squares estimation by subtracting the value of each variable of interest from the cluster-specific mean and regressing deviations from the mean for the dependent variable on deviations from the mean in the independent variables. This is mathematically equivalent to adding one dummy variable for each cluster (Greene 1993). Standard errors are adjusted for the loss of freedom associated with addition of one variable for each cluster.[6]

For logit regressions, however, given the nonlinear nature of the relationship, this solution is not applicable. For infant mortality, a dichotomous variable, we rely on conditional logits, also known as the case-control technique (Chamberlain 1980). We randomly selected pairs from each sample consisting of one dead and one live infant and exam-

3. We use the terms communities and clusters interchangeably. Although in rural areas a village was often selected as a cluster and thus community and cluster are synonymous, a city may contain more than one cluster. However, because our primary interest is in determining whether a control for residential location attenuates the effect of maternal education, we believe that this is a reasonable approximation. For studies with a deeper interest in the nature of communities and their impact on child health, a more nuanced measure of community may be required.

4. Because immunization cards are not available for all children, considerable sample attenuation occurs for this outcome. On average, only about 54% of the eligible children are included in this analysis.

5. This criterion has little impact. With the exception of Botswana, less than 1% of the sample children were excluded from the analysis of infant mortality. A slightly higher percentage of cases were omitted from the height-for-age and immunization analyses.

6. Standard errors are inflated using an adjustment factor, I, calculated using the following formula:

$$I = \frac{\sqrt{NC-K}}{\sqrt{NC-N-C-K+1}},$$

where N is the number of cases, C is the number of clusters, and K is the number of independent variables. Because the number of cases is large relative to the number of clusters, the increase in the standard errors associated with the addition of cluster-specific dummy variables is extremely small. Coefficients for maternal education lost statistical significance as a result of the adjustment in standard errors in only one country each for Tables 4 and 5.

TABLE 1. DESCRIPTIVE STATISTICS FOR SELECTED VARIABLES USED IN THE ANALYSIS: 22 DEVELOPING COUNTRIES

	Number of Immunizations		Height-for-Age		Mother's Education		
Country	Mean	Standard Deviation	Mean	Standard Deviation	Percent Not Educated	Percent With Primary Education	Percent With Secondary Education
Africa							
Botswana	7.8	.9	—	—	29.7	50.6	19.7
Burundi	6.8	1.9	–204.8	138.9	80.1	16.7	3.2
Ghana	6.2	2.2	–166.9	130.3	50.1	46.2	3.7
Kenya	7.4	1.3	—	—	35.1	50.8	14.1
Liberia	4.3	2.3	—	—	69.7	15.9	14.4
Malil	4.9	2.0	–132.0	144.1	84.9	13.8	1.3
Ondo State	7.2	1.5	–166.4	124.7	52.2	32.4	15.4
Senegal	4.9	2.4	–136.5	121.4	86.6	8.9	4.5
Zimbabwe	7.7	1.1	–159.3	121.8	19.1	69.6	11.3
$\bar{X}$	6.4	1.7	–161.0	130.2	56.4	33.9	9.7
Asia/North Africa							
Egypt	6.1	2.9	–154.5	138.0	54.8	32.5	12.7
Indonesia	—	—	—	—	20.7	61.9	17.4
Morocco	6.1	3.0	–144.5	136.4	86.8	8.7	4.5
Sri Lanka	7.4	1.5	–165.0	119.8	12.6	33.5	53.9
Thailand	7.1	1.5	–120.6	112.1	9.4	76.7	13.9
Tunisia	7.3	1.9	—	—	61.1	30.4	8.5
$\bar{X}$	6.8	2.1	–146.2	126.6	40.9	40.6	18.5
Latin America/Caribbean							
Bolivia	5.2	2.7	–176.7	142.7	22.3	48.4	29.4
Brazil	6.2	2.6	–141.7	134.4	14.2	71.4	14.4
Columbia	7.3	1.5	–141.6	128.6	9.0	59.7	31.2
Dominican Republic	4.7	1.8	–113.6	156.9	8.1	73.1	18.8
Ecuador	—	—	—	—	12.6	59.8	27.6
Guatemala	5.4	2.8	–256.7	137.5	51.8	41.7	6.5
Peru	6.4	2.0	—	—	20.5	47.8	31.7
$\bar{X}$	5.9	2.2	–138.4	116.7	19.8	57.4	22.8
Number of Countries	20	20	15	15	22	22	22
$\bar{X}$	6.3	2.0	–148.8	124.2	40.5	43.2	16.3

ined the impact of maternal education on infant mortality, given that only one child in each pair experienced death.[7] In this model, the probability of death for child i and survival for child j, given that only one member of the pair experienced mortality has the following form:

$$\text{Prob}\ (\text{death}_i = 1, \text{death}_j = 0 | \text{death}_i + \text{death}_j = 1)$$
$$= \exp(\beta X_i) \,/\, \exp(\beta X_i) + \exp(\beta X_j),$$

where X_i refers to a vector of independent variables for child i.

7. To maximize sample size, we selected as many pairs from each cluster as we could. In general, the number of pairs was determined by the number of infant deaths in a cluster.

TABLE 2. CHILDREN'S HEIGHT-FOR-AGE BY MATERNAL EDUCATION AND AREA OF RESIDENCE

	High-Education Cluster			Low-Education Cluster		
		Children With:			Children With:	
Country	Total (1)	Mother Not Educated (2)	Mother With Secondary Education (3)	Total (4)	Mother Not Educated (5)	Mother With Secondary Education (6)
Africa						
Burundi	−108	−166	−48	−187*	−191	−139*
Ghana	−105	−119	−79	−139*	−143	−123
Mali	−90	−114	−100	−107	−108	−75
Ondo State (Nigeria)	−133	−149	−119	−141	−141	−142
Senegal	−89	−95	−61	−125*	−134*	−11
Zimbabwe	−108	−111	−75	−157*	−171*	−131*
Asia/North Africa						
Egypt	−117	−150	−84	−143*	−152	−114*
Morocco	−80	−97	−48	−134*	−138*	−73*
Sri Lanka	−137	−207	−121	−150*	−203	−110
Thailand	−68	−135	−37	−122*	−148	−75*
Latin America/Caribbean						
Bolivia	−112	−171	−92	−168*	−200	−111
Brazil	−62	−66	−45	−150*	−180	−76*
Colombia	−108	−167	−86	−138*	−167	−101
Dominican Republic	−64	−118	−38	−118*	−155	−62
Guatemala	−164	−222	−103	−238*	−264*	−147*

*$p < .05$ (Kolmogrov-Smirnov test of whether the distribution of values in column 4 is lower than that in column 1, that for column 6 is lower than that for column 3, and that for column 5 is lower than that for column 2).

Table 2 shows the importance of introducing community-level controls. Table 2 shows children's height-for-age by area of residence. We divided the clusters in each country into two categories: clusters with a high concentration of secondary education and clusters with a low concentration of secondary education. High-education clusters were determined by ranking the clusters according to the number of mothers with a secondary education living in each cluster. The top clusters were selected until 75% of the sample women with a secondary education were included. These are defined as high-education clusters; the rest are called low-education clusters.

A comparison of columns 1 and 4 clearly indicates that children in low-education areas are considerably shorter than children in high-education areas. Moreover, the effect of residence is not simply a function of different educational distributions. Within the same educational category, children who live in low-education clusters are shorter than children who live in high-education clusters. In almost all countries, mothers with a secondary education fail to realize the full advantage of education if they live in low-education clusters. In contrast, in many countries children of uneducated mothers benefit if they live in high-education clusters.

RESULTS

Infant Mortality

Consistent with other studies, Model 1 in Table 3 shows a consistent negative relationship between maternal education and the probability of infant death. This model contains only variables for maternal education, urban residence, and child's age, which serves as a marker for historical period of birth. Children of mothers who attended primary school are less likely to die than are children of mothers with no education. Children of mothers with a secondary-school education are the least likely to experience infant deaths. Among the 22 countries, this effect is statistically significant in 11 countries for primary education and in 15 countries for secondary education. The education variables are jointly significant in 14 countries.

TABLE 3. COEFFICIENT FOR MATERNAL EDUCATION FROM LOGISTIC REGRESSION OF INFANT MORTALITY ON SELECTED INDEPENDENT VARIABLES

	Model 1				Model 2				Model 3			
Country	Mother Has Primary Education		Mother Has Secondary Education	*N*	Mother Has Primary Education		Mother Has Secondary Education	*N*	Mother Has Primary Education		Mother Has Secondary Education	*N*[a]
Africa												
Botswana	–.082		–.261	5,282	–.112		–.459*	5,282	.382		.609	400
Burundi	–.011		–.761*	4,714	.042		–.589	4,714	.095		.317	956
Ghana	–.205*		.013	5,276	–.157		.171	5,276	–.272		–.051	1,070
Kenya	–.266**	†	–.705**	7,811	–.209*	†	–.557**	7,811	–.036		–.158	1,098
Liberia	–.058	†	–.361**	7,435	–.010	†	–.267**	7,435	–.027		–.311*	2,752
Mali	–.455**	†	–.952	4,421	–.436**	†	–.849	4,421	–.403*		–.885	1,456
Ondo State	.061		–.220	2,221	.040		–.308	2,221	.385		.224	282
Senegal	–.144		–.372	5,343	–.022		–.033	5,343	–.124		.217	1,204
Zimbabwe	–.357*		–.486	3,728	–.382*		–.546*	3,728	–.391		–.030	450
Mean–Africa	–.169		–.456		–.138		–.382		–.043		–.008	
Asia/North Africa												
Egypt	–.274**	†	–.856**	14,029	–.178**	†	–.407**	14,029	–.198*		–.124	2,946
Indonesia	–.167*	†	–.762**	13,832	–.082	†	–.487**	13,832	.097		–.160	2,056
Morocco	–.281		–.533	4,506	–.238		–.513	4,506	–.340		.368	754
Sri Lanka	–.598**	†	–.766**	6,610	–.435		–.313	6,610	–.813**	†	–.428	464
Thailand	–.126	†	–1.169**	5,490	.141		–.770*	5,490	–.209		–.943	408
Tunisia	–.026	†	–.936**	6,416	.066	†	–.730*	6,416	.143		–.387	770
Mean–Asia/ North Africa	–.245		–.837		–.121		–.537		–.220		–.279	
Latin America/ Caribbean												
Bolivia	–.087	†	–.816**	8,278	–.029	†	–.669**	8,278	–.086	†	–.523**	1,710
Brazil	–.263*	†	–1.408**	4,849	–.159	†	–1.003**	4,849	–.084		–.835*	836
Colombia	–.203		–.607*	3,639	–.037		–.265	3,639	–.125		–.508	292
Dominican Republic	.488**	†	–.868**	6,104	–.406**	†	–.751**	6,104	–.549*		–.841*	870
Ecuador	–.478**	†	–.905**	4,720	–.444**	†	–.719**	4,720	–.291		–.769*	674
Guatemala	–.015	†	–.742**	6,345	.047		–.348	6,345	–.147		–.995*	1,124
Peru	–.331**	†	–1.001**	5,045	–.205*	†	–.738**	5,045	–.095	†	–.649*	900
Mean–Latin America/ Caribbean	–.127		–.907		–.176		–.642		–.197		–.731	
Mean–Total	–.181		–.703		–.145		–.520		–.153		–.339	

*$p < .05$; **$p < .01$ (one-tailed test)

†Education variables are jointly significant.

[a]Number of pairs of one dead infant and one live infant.

Model 2 shows results of ordinary logit regressions that include an expanded set of predictors: access to piped water, access to any type of toilet facilities, whether mother has ever had a partner, whether father/stepfather attended primary school, whether father/stepfather ever attended secondary school. Because mother's education serves as a proxy for family's socioeconomic situation, introducing direct controls for some of the socioeconomic variables reduces the effect of maternal education. Averaging across the sample countries, the coefficient for primary education changes from –.181 to –.145 and for secondary education from –.703 to –.520. In Model 2, the effect of maternal education is statistically significant for 7 and 13 countries for primary and secondary education respectively. The two education variables are jointly significant in 11 of the 22 countries.

Fixed-effects models with the same set of variables show further declines in educational differentials. In Model 3, averaging across the sample countries, the coefficient for primary education is –.153 and for secondary education is –.339. A coefficient of –.153 reflects a 14% reduction in child mortality with only 9 out of the 22 countries showing this level of reduction. Moreover, this effect is statistically significant in only 4 of the 22 countries, even after using a generous one-tailed test. Obtaining secondary education—as opposed to remaining uneducated—has a larger effect, with a mean reduction in the odds of infant mortality of about 27%, but the effect is statistically significant in only 7 of the 22 countries. The decline in the impact of maternal education is particularly marked in Sub-Saharan Africa.

These results show substantial narrowing of maternal educational differentials in infant mortality[8] based on: (1) incorporation of a relatively small set of variables used as an index of family socioeconomic conditions and (2) control for cluster-level effects by focusing on within-cluster educational differentials. Note that even without including cluster-level effects, after controlling for socioeconomic conditions, primary education per se has a statistically significant negative effect in only 7 of the 22 countries studied (Model 2).

Height-for-Age

DHS data are unique in providing comparable anthropometric measurements for a wide range of developing countries. Our analyses are restricted to children aged 12 to 36 months for whom data on height and age are available. Sex and an age-specific standardized score on height have often been used to indicate children's nutritional status. This variable reflects both food intake and freedom from chronic gastrointestinal diseases.[9] Socioeconomic differences in height-for-age have been shown to be quite striking in a variety of national contexts (Martorell and Habicht 1986).

Results presented in Table 4 show that children's height-for-age is closely linked to mother's education in Model 1. In 6 of the 15 countries, the coefficient for primary education is significant and positive, and in 13 countries the coefficient for secondary education is significant and positive. The introduction of variables reflecting family socioeconomic background attenuates these coefficients (Model 2). In the final model, the effects of primary and secondary education are small and are statistically significant in only a few countries. On average, the coefficients in Model 3 are only about one-third as large as those in Model 1.

In interpreting educational effects on children's height-for-age, it is important to remember that anthropometric measurements are available only for living children. Because maternal education affects children's survival probability, the sample of living children overrepresents children of educated mothers. Consequently, the effect of maternal education is likely to be slightly underestimated.

Immunization Status

DHS collected data on the number of immunizations received by children born in the five years before the survey. A detailed immunization record is available only for children for whom immunization cards are available. Thus, a substantial number of eligible children do not have data on the exact number of immunizations received. Moreover, the availability of the immunization card is likely to be related to maternal education (Boerma et al. 1990). Thus, the analyses presented here are subject to substantial sample selectivity and the results should be treated with caution.

We created a count of immunizations received by the index child as of the interview date. The immunizations considered here are three doses of polio, three doses of DPT, a measles vaccine, and BCG (Bacillus Calmette-Guerin) vaccine. Because our sample is restricted to children aged 12 months and over, in spite of some cross-national variation in recommended ages for some of the immunizations, it is generally expected that all children should have received these eight doses.

The results presented in Table 5 differ from results for the other dependent variables. Maternal education continues to have a statistically significant effect on children's immunization status regardless of the inclusion of cluster-level fixed effects or other socioeconomic factors. Thus, in our final model in Table 5, the effect of primary education is statistically significant in 9 of the 20 countries studied, whereas the effect of secondary education is statistically significant in 11 of the 20 countries studied. Although the magnitudes of the education coefficients for specific countries decline slightly across the models, this decline is relatively small.

Although there are considerable educational differences in immunization status, these differences do not seem to translate into mortality or anthropometric differences of similar magnitudes. Educational differences in immunization status may be result from an increase in one or two types of immunization that do not have a significant effect on mortality. For example, research on vaccine efficacy shows that the efficacy of BCG in preventing tuberculosis ranges from 0 to 80% (Fine 1988). So if the educational difference in immunizational status is a result of higher BCG vaccination rates for educated mothers, this may not translate into reduced mortality. Alternatively, if the educational differences are a result of measles immunization, this would have a greater impact at ages 1–5 rather than in infancy.

8. Maternal behavior may be more important to child mortality than to infant mortality. Hence, the effect of maternal education may increase with the age of the child (Caldwell and Caldwell 1993).

9. The height-for-age variable refers to the index child's standardized score on a sex-specific height-for-age scale developed by WHO/NCHS. In the reference population of North American children, the mean is 0 and the standard deviation is 100. The mean score for most children in developing countries is negative and the range is between –597 and +598.

TABLE 4. COEFFICIENT FOR MATERNAL EDUCATION FROM REGRESSIONS OF HEIGHT-FOR-AGE ON SELECTED INDEPENDENT VARIABLES

	Model 1		Model 2		Model 3			
Country	Mother Has Primary Education	Mother Has Secondary Education	Mother Has Primary Education	Mother Has Secondary Education	Mother Has Primary Education	Mother Has Secondary Education	Number of Cases	Number of Clusters
Africa								
Botswana	—	—	—	—	—	—	—	—
Burundi	7.15 †	72.26**	5.30	55.14*	7.16	34.09	995	124
Ghana	18.48 †	37.40	9.57	13.02	9.31	12.04	1,044	142
Kenya	—	—	—	—	—	—	—	—
Liberia	—	—	—	—	—	—	—	—
Mali	14.20	17.00	11.41	–33.15	4.23	–30.58	417	84
Ondo State	–2.54	26.56*	–1.79	19.89	–.78	12.71	584	73
Senegal	43.03 †	58.27*	52.64** †	77.51*	30.42	55.34	330	87
Zimbabwe	16.91 †	39.34**	11.70	28.51*	8.26	17.18	812	140
Mean–Africa	16.21	41.81	14.81	26.82	9.77	16.80		
Asia/North Africa								
Egypt	10.16 †	47.57**	5.27	26.15*	–10.67	–8.60	1,128	285
Indonesia	—	—	—	—	—	—	—	—
Morocco	25.26* †	56.85**	8.48	29.41*	2.74	1.38	942	163
Sri Lanka	44.98** †	91.03**	30.63** †	55.98**	30.19** †	45.46**	1,245	265
Thailand	45.68** †	98.73**	29.53* †	69.32**	21.48 †	64.48**	1,120	255
Tunisia	—	—	—	—	—	—	—	—
Mean–Asia/North Africa	31.52	73.55	18.48	45.22	10.94	25.68		
Latin America/Caribbean								
Bolivia	25.75** †	87.03**	24.27* †	77.70**	–1.57 †	27.53*	1,365	395
Brazil (T)	61.88** †	104.72**	64.56** †	77.98*	43.29** †	18.93	373	85
Colombia	21.60 †	65.85**	10.04 †	37.76*	27.84* †	56.23**	775	139
Dominican Republic	29.12 †	79.79**	17.54 †	55.74**	–20.18 †	17.52	1,082	334
Ecuador	—	—	—	—	—	—	—	—
Guatemala	53.13** †	156.95**	40.88** †	104.95**	14.65* †	65.85**	1,277	203
Peru	—	—	—	—	—	—	—	—
Mean–Latin America/Caribbean	39.78	90.57	33.80	65.76	8.57	27.91		
Mean–Total	28.88	68.03	22.85	46.02	9.61	23.19		

*$p < .05$; **$p < .01$ (one-tailed test)

†Education variables jointly significant.

Note: T = Access to toilet facility was omitted from Model 3.

DISCUSSION

We have shown that although maternal education is strongly correlated with three different markers of child health—infant mortality, height-for-age, and number of immunizations—for two of the three outcome variables, this relationship is considerably attenuated once controls for individual socioeconomic characteristics and community effects are introduced. The magnitudes of the coefficients decline substan-

TABLE 5. COEFFICIENT FOR MATERNAL EDUCATION FROM REGRESSIONS OF NUMBER OF IMMUNIZATIONS ON SELECTED INDEPENDENT VARIABLES

Country	Model 1: Mother Has Primary Education		Model 1: Mother Has Secondary Education	Model 2: Mother Has Primary Education		Model 2: Mother Has Secondary Education	Model 3: Mother Has Primary Education		Model 3: Mother Has Secondary Education	Number of Cases	Number of Clusters
Africa											
Botswana	.37**		.27	.26*		.18	.02		.08	195	90
Burundi	–.01		.29	.09		.12	.00		–.35	594	103
Ghana	.27	†	1.53**	.15	†	1.05**	.43*		.82*	492	113
Kenya	.53**	†	.54**	.44**	†	.33*	.48**	†	.41**	874	231
Liberia	–.26	†	1.05**	–.33	†	1.05**	–.13	†	1.04**	455	99
Mali	.14		1.10	–.03		.31	–.09		.27	125	35
Ondo State	.18	†	.90**	.04		.57	–.05		.23	176	41
Senegal	1.01**	†	1.41**	.82*	†	1.24**	.78*	†	1.78**	283	66
Zimbabwe	.03		.06	.05		.05	.02		.15	743	142
Mean–Africa	.25		.79	.17		.54	.16		.49		
Asia/North Africa											
Egypt	.69**	†	.87**	.41**	†	.47*	.05		–.04	1,687	332
Indonesia	—		—	—		—	—		—	—	—
Morocco	.79**	†	.65**	.65**	†	.54*	.15		.21	976	167
Sri Lanka	.39*	†	.90**	.35*	†	.76**	.49**	†	.81**	1,346	279
Thailand	.28		.41	.20		.15	.22		.48	422	138
Tunisia	.67**	†	.65**	.42**	†	.37**	.24**		.24	1,478	145
Mean–Asia/North Africa	.56		.70	.41		.46	.23		.34		
Latin America/Caribbean											
Bolivia	1.27**	†	2.04**	1.17**	†	1.67**	.89**	†	1.30**	1,645	447
Brazil (T)	.78**	†	1.48**	.48	†	.94**	.34		.67*	950	246
Colombia	.73	†	1.12**	.58	†	.86*	.52*		.68*	426	107
Dominican Republic	2.24**	†	2.48**	1.27		1.82*	4.52**	†	5.17**	88	35
Ecuador	—		—	—		—	—		—	—	—
Guatemala	.97**	†	1.27**	.79**	†	.76**	.50**	†	.76*	1,405	206
Peru	.11	†	.91**	.23	†	1.04**	.17	†	1.30**	371	48
Mean–Latin America/Caribbean	.90		1.41	.66		1.09	.99		1.47		
Mean–Total	.54		.98	.39		.71	.45		.78		

*$p < .05$; **$p < .01$ (one-tailed test)

†Education variables jointly significant.

tially, and the differences between children of educated and uneducated mothers remain statistically significant in only a handful of countries. We have been unable to control for several factors known to be associated with both education and health outcomes, primarily race/ethnicity and income. Addition of these control variables would probably further attenuate the remaining relationship between maternal education and child health. Thus, our results suggest that there is little empirical justification for such bold assertions as "a large number of studies have shown, almost as convincingly as anything can in the social sciences, that a mother's education has an independent, strong and positive impact on the survival of her children" (Caldwell 1994).

Several important caveats must be noted, however. The education coefficients, although small, are in the expected direction and continue to be statistically significant in a small number of countries—they are particularly large for secondary education in Latin America. Moreover, this analysis masks potentially positive effects of maternal education through two avenues: (1) education may affect access to health facilities at the community level, thereby improving the health of children of educated as well as uneducated mothers in communities with high levels of education, and (2) higher immunization levels for children of educated mothers may reduce the likelihood of diseases like measles for all children in the community, thereby reducing mortality for children of educated and uneducated mothers in a given community through spillover effects.

Our results also suggest that although educational differentials in infant mortality and height-for-age narrow substantially after controlling for household-level socioeconomic variables and the community of residence, educational differentials in number of immunizations continue to be large. This suggests that educated mothers are more likely to engage in health-promoting behavior.[10] However, in many instances external factors supersede the abilities of families to enhance the health of their members. Thus, individual-level improvement in health-enhancing behavior may fail to translate into actual improvements in health. For example, Bicego and Boerma (1991) find few differences in the incidence of diarrhea by maternal education levels. Onset of diarrhea is probably related to the prevalence of disease in the broader environment and educated mothers may be able to do little to limit the exposure of their children—particularly toddlers—to diarrhea.

These observations suggest a need for further research and reevaluation of the available evidence before engaging in policy advocacy. Because investment in women's education is desirable on grounds other than health improvements, we have tended to evaluate the evidence for an education/health link less critically. However, this lack of critical attention may result in public policies based on unrealistic expectations from mothers and may result in lower levels of child health and an increased burden on mothers. This issue is particularly crucial in the current climate of privatization in which public investments in health-delivery systems, food subsidies, water supply, and sanitation systems are declining and a focus on primary health care is replaced by a focus on selective health care (Pebley 1993; World Bank 1993).

Our results also suggest a need for explicitly measuring the community context and incorporating it in studies of child health. Our findings indicate considerable variation in child health outcomes across clusters, but given the paucity of community-level information in the core DHS questionnaire, we have not attempted to identify the sources of this variation. However, our findings underscore the importance of directing explicit attention to community characteristics such as access to primary health-care facilities, number of physicians and nurses in the community, immunization campaigns, and the nature of water and sanitation systems in order to develop effective health policies that incorporate public and private dimensions of investments in child health.

10. In some instances, maternal education may also be related to harmful health-related behavior. Educated mothers tend to be less likely to breast-feed or to breast-feed for short durations (Benefo and Parnell 1991; Stewart et al. 1991), and breast-feeding is associated with low infant mortality rates (Palloni and Tienda 1986).

APPENDIX TABLE A1. DEFINITIONS OF VARIABLES USED IN THE ANALYSES

Variable	Definition
Infant Mortality	Death during the first year of life. (Base = children born alive one to ten years before the survey)[a]
Number of Immunizations	Number of immunizations received. Immunizations considered include 3 doses of DPT, 3 doses of polio, measles and BCG. (Base = children born 1 to 5 years before the survey and have a health record)[a]
Height-for-Age	Standardized score on sex-specific height-for-age scale developed by WHO/NCHS. In the reference population consisting of North American children, the mean is 0 and the standard deviation is 100. The mean score for most children in developing countries is below 0. (Base = children born 1 to 5 years before the survey and have exact data on date of birth as well as height)[a]
Mother's Primary Education	1 if the highest enrollment level for mother is primary school. The number of years included in primary school varies by country, but is usually 4 years. 0 if uneducated or ever enrolled in secondary school.
Mother's Secondary Education	1 if the highest enrollment level for mother is secondary school or higher. 0 if never enrolled in secondary school. No schooling is the omitted category.
Father's/ Stepfather's Education	Similar to mother's education described above: There is one variable for Father's/Stepfather's Primary Education and one for Father's/Stepfather's Secondary Education (both as described above for mother's education). Set to 0 for women who do not have a partner.
Ever Had a Partner	1 if mother is not married formally or consensually, 0 otherwise.
Child's Age	Child's age at the time of the interview, in months.
Urban Residence	1 if the current residence is urban, 0 otherwise.
Water	1 if the household has access to piped water, 0 otherwise.
Toilet	1 if there is a toilet (any kind) in the household, 0 otherwise.

[a]Dependent variables.

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