Journal of African Economies, Vol. 22, number 5, pp. 693–731
doi:10.1093/jae/ejt007 online date 22 May 2013

Determinants of Schooling: Empirical Evidence from Rural Ethiopia

Subha Mani[a,*], John Hoddinott[b] and John Strauss[c]

[a]Fordham University, Bronx, NY, USA
[b]International Food Policy Research Institute, Washington, D.C., USA
[c]University of Southern California, Los Angeles, CA, USA

*Corresponding author: Subha Mani, Department of Economics and Center for International Policy Studies, Fordham University, 441 East Fordham Road, Bronx, New York, NY 10458, USA. E-mail: smani@fordham.edu

Abstract

We examine the determinants of current enrolment status and relative grade attainment among primary school children in rural Ethiopia. We use repeated cross-sectional data from 15 rural villages in Ethiopia to capture the impact of changing household and child characteristics on enrolment status and relative grade attainment between 1994 and 2004. Using instrument variable (IV) estimation, we find, first, a positive income effect for schooling enrolments and an even stronger effect for relative grade attainment. Second, the effect of income is larger for girls compared to boys. Third, OLS estimates of the impact of household income are biased downwards relative to IV results. Finally, observable community characteristics have little role in explaining schooling. These findings suggest that policies that address the demand-side constraints with a special focus on girls will have the potential to improve schooling attainments as well as to reduce gender differences in schooling attainments found in Ethiopia and elsewhere in sub-Saharan Africa.

JEL classification: I20, I25

1. Introduction

Sub-Saharan African countries have some of the lowest primary school enrolment and school completion rates in the world. For example, the World

Bank's 2009 World Development Report indicates that only 60% of Sub-Saharan African children completed primary school. Economic theory predicts and many empirical studies confirm that investments in schooling lead to higher economic and non-economic gains in the future for the individual, her household and the aggregate economy. Variants of the Mincer (1974) equation have been widely used to estimate the private returns from schooling. Micro-level findings indicate that in developing countries, the private returns to schooling lie between 5 and 15% (Orazem and King, 2008).[1] While difficult to estimate, there is a widespread view that schooling conveys additional social returns, for example through improvements in maternal and child health and lower fertility rate (Strauss and Thomas, 1995; Schultz, 2002).

The conjunction of high private and social returns to schooling combined with low enrolments and completion rates points to the value of understanding the determinants of schooling enrolments and grade attainments in Africa. Although this appears to be a well-trodden territory with studies identifying labour market returns, parental education, household income, child health and school characteristics as key determinants of schooling,[2] closer inspection reveals important limitations in the extant literature. Consider Behrman and Knowles's (1999) review of studies that examine the association between income and schooling. Out of the forty-two studies they examine, twenty-one analysed determinants of final schooling attainments (grades or years of schooling) using current socio-economic characteristics. Such an empirical specification is potentially misspecified since the right-hand-side characteristics do not map to the year in which the schooling investment decision is made.[3] Most of the studies they consider rely on single-year cross-sectional data which cannot capture the impact of changing socio-economic environment on schooling investments across cohorts. Nearly 90% of these studies treat household characteristics such as income as predetermined, thereby ignoring potential correlation between income and unobserved characteristics that also affect schooling. Finally, only a quarter of the papers reviewed by Behrman and Knowles used both short-run (enrolment, drop-out) and long-run (completed grades, test scores, years of schooling) measures of schooling in their analysis.

1 See Psacharopoulos and Patrinos (2004) for the estimated private returns to schooling for over forty countries.

2 See Schultz (1988), Rosenzweig (1990), Lillard and Willis (1994), Parish and Willis (1993), Glewwe and Jacoby (1995), Behrman and Knowles (1999) and Alderman *et al.* (2006).

3 See Bommier and Lambert (2000) for more recent evidence based on this approach.

This paper contributes to the literature on the determinants of schooling in sub-Saharan Africa by taking these methodological issues seriously. Using repeated cross-sectional data from rural Ethiopia, we estimate the determinants of schooling including both current enrolment and a longer term measure of schooling progression, relative grade attainment. Second, our repeated cross-sectional data allow us to capture the impact of the changing socio-economic environment on schooling among primary school children between 1994 and 2004. Third, we restrict our analysis sample to primary school-age children for whom the socio-economic characteristics map to the year in which the schooling investment decision is made. Restricting the sample to include primary school-age children only also addresses individual specific out-migration-related selection concerns. Our sample has very little attrition at the household level, addressing potential biases associated with household migration. Fourth, we construct a measure of schooling progression—relative grade attainment—that allows for delays in grade accumulation. This measure is computed as actual completed grades of schooling divided by potential grades. Where potential grade is calculated as total number of grades accumulated had the individual completed one grade of schooling by age 7 and continued to accumulate an additional grade of schooling in each subsequent year. Fifth, we use an instrument variable (IV) estimation strategy to address the endogeneity problem in our measure of household income.

Ethiopia is an appropriate setting for a study of schooling. Historically, the level of schooling attainments has been abysmal with more than 50% of the population having never attended school (Central Statistical Agency [Ethiopia] and ORC Macro, 2006). However, the last 15 years has witnessed a significant increase in primary school enrolment, particularly in rural areas. In 1995, only 15% of the primary school children from rural areas were currently enrolled in school (Schaffner, 2004). By 2005, this has more than doubled to 38.8%. We find that in rural Ethiopia: (a) the impact of household income on schooling is not time invariant, (b) the demand for schooling progression is relatively more elastic compared with the demand for schooling enrolment in all three years, (c) these income effects are substantially larger for girls compared with boys, and (d) in these data, community characteristics explain little of the changes in enrolment and relative grade attainment, possibly reflecting limited observed variability in these variables. These findings suggest that policies that address the demand-side constraints with a special focus on girls will have the potential to improve schooling attainments as well as to reduce gender differences in schooling attainments found in Ethiopia and elsewhere in sub-Saharan Africa.

2. Data

We use data from the Ethiopian Rural Household Survey (ERHS), a longitudinal socio-economic survey administered in selected rural peasant associations of Ethiopia between 1989 and 2004.[4] The first wave of the ERHS was fielded in 1989 during which households from seven farming villages in central and southern Ethiopia were surveyed. In 1994, the sample was expanded, retaining six of the original villages and adding nine additional villages. These fifteen rural villages are representative of the diverse farming systems practiced in rural Ethiopia (Dercon *et al.*, 2009). The ERHS provides extensive information on household composition, income, consumption expenditure, farm and non-farm assets, ownership and value of land and livestock units, anthropometrics, crop production and schooling. In 1997 and 2004, the survey also collected detailed village-level information on infrastructure, wages and prices of consumption goods.

We use observations on primary school-age children from the 1994, 1999 and 2004 waves of the ERHS. There are a number of reasons why we focus on primary school-age children. First, fewer than 10% of school-age children complete primary schooling. Hence, it is the socio-economic environment during primary school age that matters most in determining a child's complete future trajectory of schooling. Second, including observations on high school-age children would create considerable selection bias, since there occurs huge out-migration among high school-age females due to early marriage. Ezra and Kiros (2001) document that 79% of female migrants from Ethiopia migrated at the time of marriage with the average age at the time of marriage being 16 years. Fafchamps and Quisumbing (2005) find that in the ERHS sample, the average age of first marriage for women is 17 years. While household attrition could be problematic for our analysis, the low levels of household attrition in the ERHS allay this concern. Only 13% of the household sample was lost between 1994 and 2004. This partly reflects the relative immobility of the sample (it is difficult to obtain land if households migrate) and partly a high degree of institutional continuity in the development of these surveys (see Dercon *et al.*, 2012).

We use data collected during the 1994, 1999, and 2004 waves to avoid the confounding effects of the irregular spacing of the survey rounds. Mirroring the levels and patterns observed in national statistics, in 1994 only 12.7% of primary school children were enrolled in school. By 2004, this percentage

[4] The smallest administrative unit in Ethiopia is called a 'peasant association', which is sometimes equivalent to one village or a cluster of villages. We use the term 'villages' and 'peasant association' interchangeably throughout this paper.

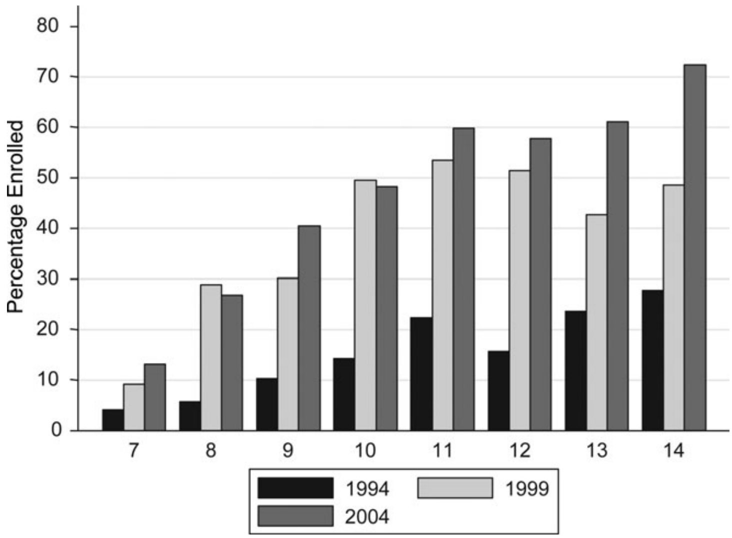

Figure 1: Male Enrolment Rate

increased to 45.5% (see Figures 1 and 2). Four additional observations emerge from these figures. First, the large improvements in enrolments during 1994–99 reflect mostly new enrolments, whereas improvements in enrolments during 1999–2004 reflect both new and continued enrolments.[5] Second, few children are enrolled at young ages. Third, there exists a strong positive association between male enrolments and age. However, for females, the relationship between age and enrolment is non-linear, with enrolments peaking around age 12 and declining thereafter. Lastly, gender gaps in enrolment narrow over time. Figures 3 and 4 depict trends in relative grade attainment among primary school-age children. More timely enrolments in 1999 led to a steep increase in the relative grades accumulated. The pattern of improvement in relative grades is similar to the pattern observed for enrolment rates. Male–female differences in relative grades increase between 1994 and 1999 and then narrow during 1999–2004.

[5] A table of age distributions (available on request) illustrates the entry of new children into the sample used here in 1999 and 2004. The age distribution in our sample is fairly constant over time, reflecting that new, younger students are replacing the prior young students, as they age.

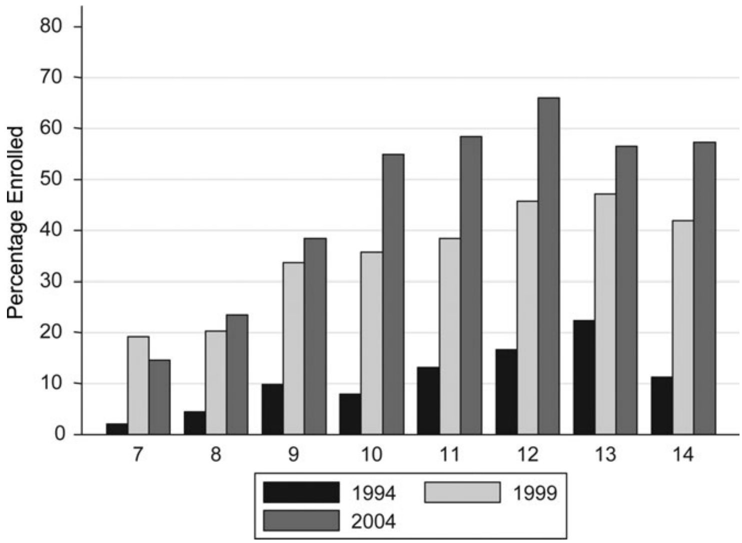

Figure 2: Female Enrolment Rate

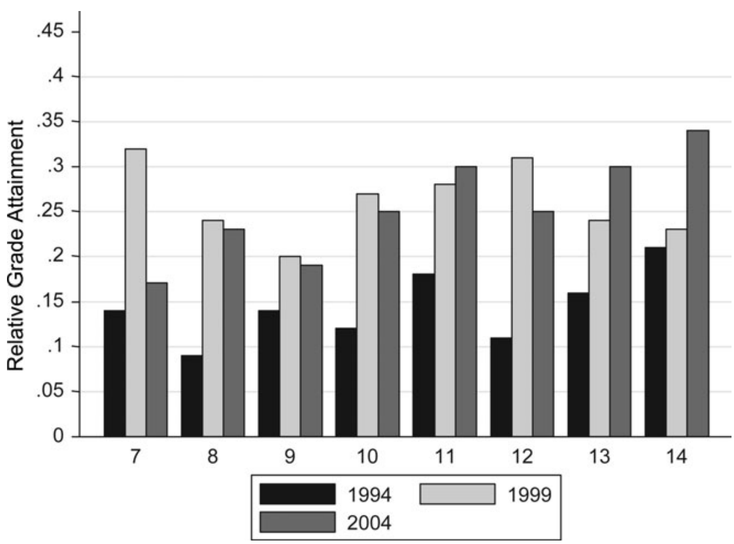

Figure 3: Male Relative Grade Attainment

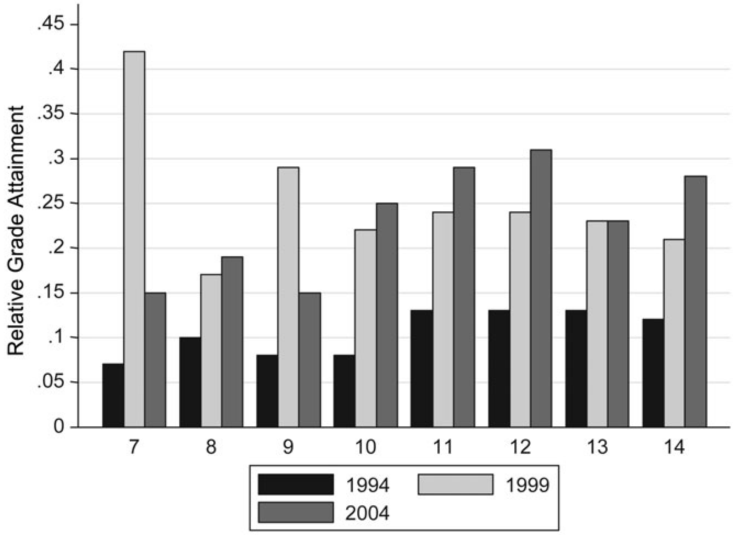

Figure 4: Female Relative Grade Attainment

3. Conceptual framework

We model the determinants of enrolment and relative grade attainment using a conditional schooling demand function. These are appropriate when there exists pre-allocated or non-market goods.[6] Specifically, following Pollak (1969), Thomas *et al.* (1990) and Glewwe and Miguel (2008), we assume that the household maximises utility subject to an income constraint and a schooling production function, where the schooling production function transforms available home and schooling inputs into schooling outcomes. The solution to this optimisation problem is a set of conditional demand functions for all the marketed goods (food, books) where the demand for each marketed good depends upon prices of all market goods, pre-allocated consumption of the non-marketed goods (mother's time spent on child care, availability of free primary schools), total expenditure on the marketed goods and demographic factors that affect household preferences.

The conditional schooling demand function is expressed in terms of price for schooling inputs (P_t^s), price of leisure (w_t) and prices of other consumption goods (P_t^c) as well as total household expenditure (X_t), demographic characteristics that affect household preferences and life-cycle position

[6] In our sample, every village has a primary school and these do not charge fees for attendance. Their availability is likely to affect the way in which household allocates resources towards the purchase of school books and other consumption goods (Pollak, 1969).

(D_t) and the availability of schools and other village infrastructure (I_t). It is also a function of children's observable characteristics such as age and gender and unobserved cognitive abilities (θ_{ct}, θ_c) that affect parents' allocation of schooling inputs. Time-varying and time-invariant rearing and caring practices (μ_{ht}, μ_h) also affect schooling:

$$S_t^* = f(P_t^c, P_t^s, w_t, D_t, X_t, \theta_{ct}, \theta_c, \mu_{ht}, \mu_h; I_t) \quad (1)$$

The empirical counterpart to the conditional schooling demand function is:

$$S_i = \beta_0 + \sum_{j=1}^{R} \beta_j C_{ij} + \sum_{j=1}^{R} \gamma_j H_{ij} + \varepsilon_i + \varepsilon_h + \varepsilon_v \quad (2)$$

where S_i is enrolment and relative grade attainment of child i. Enrolment takes a value of 1 if the child is enrolled in school at the time of the survey, 0 otherwise.[7] Relative grade attainment is defined as actual grades divided by potential grades, where potential grades is calculated as total number of grades accumulated had the individual completed one grade of schooling by age 7 and continued to accumulate an additional grade of schooling in each subsequent year. Means and standard deviations are given in Table 1.

C includes individual-level variables and H captures household-level variables. We control for the age of the child, male dummy, mother's age and measure of parental schooling. Age is specified using dummy variables, with a separate dummy variable assigned for each year between 7 and 14 years; the omitted category is 7 years. The age dummies capture for age-specific differences in schooling. The male dummy equals 1 if male, 0 if female to capture gender-specific differences in schooling. We interact the age dummies with the male dummy to capture age–gender-specific differences in schooling.

We include parental schooling as a demand-side determinant. Parents affect children's schooling through time spent teaching at home and through the choice of schooling inputs. Studies find that there is positive partial relationship between parental schooling and child's schooling (Strauss and Thomas, 1995; Tansel, 1997; Schultz, 1988; Brown and Park, 2002; Schaffner, 2004; Orazem and King, 2008). Some studies in the literature specifically find that mother's schooling has a greater impact on children's

[7] Some children in our sample are enrolled in religious schools. Our interest is limited to measuring human capital accumulated through learning subjects like mathematics, science and social science; none of which is taught in religious schools. For this reason, we treat children enrolled in religious schools as not enrolled.

Table 1: Descriptive Statistics

Variable name: definition	Mean (standard deviation) 1994	1999	2004
Household and individual characteristics			
Enrolment = 1 if currently enrolled in school, 0 otherwise	0.13	0.38	0.45
	(0.34)	(0.48)	(0.49)
Schooling: actual grades of schooling completed	0.60	1.13	1.17
	(1.43)	(1.70)	(1.67)
Relative grade attained: actual grade completed divided by potential grade given age	0.13	0.27	0.24
	(0.38)	(0.47)	(0.41)
Household size: total number of individuals in the household	8.29	7.30	7.15
	(3.17)	(2.90)	(2.33)
PCE: log of real per capita household consumption expenditure	3.79	4.05	4.05
	(0.73)	(0.76)	(0.75)
Mother's schooling = 1 if the mother has at least one completed grade of schooling, 0 otherwise	0.17	0.22	0.31
	(0.37)	(0.42)	(0.46)
Father's schooling = 1 if the father has at least one completed grade of schooling, 0 otherwise	0.39	0.44	0.51
	(0.49)	(0.49)	(0.50)
Male = 1 if male, 0 otherwise	0.50	0.50	0.51
	(0.50)	(0.50)	(0.50)
Age: in years	10.80	10.81	10.70
	(2.29)	(2.21)	(2.41)
Land: in hectare per adult member	0.57	0.47	0.64
	(0.56)	(0.43)	(0.56)
Livestock units: standardised livestock units	3.16	3.32	3.42
	(4.12)	(3.10)	(3.61)
Number of adult males: sum of all males 18 years and older	1.65	1.58	1.49
	(1.18)	(1.15)	(0.95)
Number of adult females: sum of all females 18 years and older	1.71	1.70	1.52
	(1.05)	(1.03)	(0.83)
Mother's age: age of the mother in years	38.64	39.76	40.05
	(9.60)	(9.91)	(8.65)
Age of the household head: in years	48.78	49.27	49.31
	(13.49)	(12.64)	(12.36)
Observations	2,047	1,877	1,629
Village characteristics			
Distance: distance to primary school in kilometres			2.10
			(1.31)
Piped water = 1 if piped water is available in the village, 0 otherwise			0.18
			(0.38)
Electricity = 1 if electricity available in the village, 0 otherwise			0.31
			(0.46)

education compared with father's schooling (Singh, 1992; Alderman *et al.*, 2001; Dostie and Jayaraman, 2006). We are cautious in interpreting these variables since the presence of common intergenerational unobservables (such as ability) affect both children's and parents' schooling.[8] In the ERHS, parental grade attainment is low, averaging only one grade in 1994 and two grades by 2004. The majority of parents have no formal schooling, thus there is limited variation in parental grade attainments, making it suitable to characterise schooling using dummy variables, equalling 1 if the child's mother (father) has at least one grade of formal schooling, 0 otherwise.[9] Mother's age is included in the regressions to capture mother's experience and knowledge, which affects her ability to make household decisions and schooling-related decisions.

Household-level regressors include number of adult ($>$18 years) males and number of adult ($>$18 years) females, capturing household demographic composition. Age of the head of the household is included as an additional regressor to capture household experience and life-cycle position.[10]

Household resources are often measured in terms of household income or wealth or expenditures. Increased household income improves enrolment probabilities (Dostie and Jayaraman, 2006), lowers delays in schooling enrolment (Glewwe and Jacoby, 1995), increases levels of schooling completion (King and Lillard, 1987), lowers school withdrawal rates (Glick and Sahn, 2000) and improves test scores (Brown and Park, 2002). In our specification, we include household consumption expenditure as a representation of permanent income (Thomas *et al.*, 1990; Behrman and Knowles, 1999). Total household consumption expenditure is computed as the sum of value of food items consumed, including purchased and non-purchased consumption goods (consumption out of own stock), and value of non-investment type non-food items purchased. Non-food items include consumables such as matches, batteries, kerosene, but exclude expenditure on durables

[8] We do not have data that address this problem. For further discussion, see Lillard and Willis (1994) and Behrman and Rosenzweig (2002).

[9] In 1994, 92% of mothers and 80% of fathers had zero completed grades of schooling in our sample. By 2004, this improved a little to 80% of mothers and 60% of fathers having zero grades of schooling completed.

[10] We exclude for birth order, since birth order is strongly correlated with the number of children ever born creating fertility-related selection concerns. For robustness purposes, we re-estimated our specification including birth order, finding that our coefficient estimates remain unchanged. We also estimated alternate specifications that indirectly control for birth order by including the number of older and younger children, again finding that our coefficient estimates remain unchanged.

such as housing (Dercon *et al*., 2009. Consumption is valued using prices obtained from the market surveys fielded at the same time as the household surveys. Total household consumption expenditure is divided by household size to capture the per-person resource availability in the household. Nominal per capita consumption values are converted to real per capita consumption expenditure using a food price index. We use the logarithm of real per capita household consumption expenditure (PCE) to capture non-linearities in the relationship between this characteristic and the outcome variables of interest. While prices, demographic characteristics and local village infrastructure can be treated as exogenous to the household, total expenditure cannot be treated as exogenous. Total household expenditure on all market goods is jointly determined with the demand for schooling. Hence, in the empirical work to follow, we will treat PCE as endogenous.

The disturbance term ε_v captures village-specific characteristics that affect schooling. A number of studies find that improvements in school quality result in higher enrolments (Glewwe and Jacoby, 1994) and early/timely enrolments (Bommier and Lambert, 2000). Glewwe and Jacoby (1994, 1995) show that school characteristics like building materials, writing materials and teaching materials are important determinants of test scores and other measures of cognitive development. However, there are econometric concerns associated with using measures of school quality. These are endogenous when there is non-random programme placement (Rosenzweig and Wolpin, 1986). Given this, we adopt two approaches. First, following Dostie and Jayaraman (2006) and others, we use village fixed effects to sweep out the impact of all time-invariant village-level characteristics.[11] Second, mindful of these econometric issues, for the 2004 survey round (the only round where we have observed village-level characteristics), we include three observed village characteristics in an alternative specification. These are distance to the primary school (a measure of the price of schooling) and the availability of electricity and piped water.

The remaining unobservables in equation (2) are ε_i and ε_h. ε_h captures household specific unobservables such as parental preferences and their discount rate.[12] ε_i is assumed to be a random i.i.d. error term.

[11] See Glewwe (2002) for detailed review on the supply-side determinants of schooling.

[12] We cannot use household fixed effects to remove ε_h, since we do not have enough cases of multiple children from the same household.

4. Results

4.1 Enrolment

To estimate the determinants of enrolment in each survey round, we estimate a linear probability model (LPM) with village fixed effects. The village fixed effects allow us to (a) remove all sources of common village-level unobservables, and (b) address cluster-related issues in the standard errors since common village-level unobservables are also cluster effects (Wooldridge, 2003).[13] We use a LPM, as logit and probit fixed effect estimators suffer from the incidental parameter problem (Greene, 2004). Our estimates account for arbitrary forms of heteroskedasticity using the White (1980) formulation (see Wooldridge, 2002).

Before we move to our preferred estimates, we report results for enrolment and grade progression using land (hectares per adult member) as a measure of household resources (Table 2). In Ethiopia, land allocations are based on local administrative decisions; it is illegal to buy or sell land. These allocations are determined independent of household's schooling investment decision and hence can be treated as exogenous to the household. We interact land with rainfall to capture non-linear effects of land on schooling. Following a formulation set out in Maccini and Yang (2009), rainfall is calculated as the natural log of mean rainfall observed over the kirmet (June–September) months in year t minus the natural log of mean annual rainfall computed over 1994–2004. Since more than 90% of the agricultural output is produced during this season, we construct our rainfall measure to include the kirmet rains only. We allow these effects to vary by gender. Results are reported in Table 2. The impact of land while positively associated with schooling is statistically insignificant and has limited role in explaining variations in enrolment. The interaction term between land and rainfall suggests that the causal effect of land is related to the amount of rainfall a plot receives where these effects vary by gender. The coefficient estimate on the interaction term indicates that the more rainfall a plot receives the lower is the probability of being enrolled among male relative to female children. This is not completely surprising since rural households often pull out male children from school to get some help on the farm. Similar effects are observed for our measure of grade progression as well for all three waves of our sample. Rainfall generates both substitution and income effects on schooling; as a result, the net effect of rainfall on schooling depends in part on other assets that the household can use for meeting demand for labour.

[13] We thank Jeffrey Wooldridge for his assistance on this point.

Table 2: LPM of Enrolment and Relative Grade Attainment

Variables	Enrolment			Relative grade attainment		
	1994 (1)	1999 (2)	2004 (3)	1994 (4)	1999 (5)	2004 (6)
Mother's schooling	0.0961**	0.0696*	0.0898**	0.0846*	0.014	0.0880**
	(0.03)	(0.039)	(0.03)	(0.047)	(0.040)	(0.036)
Father's schooling	0.108***	0.0952***	0.0771**	0.114***	0.111***	0.047
	(0.022)	(0.028)	(0.030)	(0.025)	(0.031)	(0.034)
Land	−0.019	0.070	0.040	−0.030	−0.128**	0.042
	(0.049)	(0.072)	(0.037)	(0.049)	(0.058)	(0.029)
Land × rainfall	0.036	−0.122	0.0777*	0.076	0.130	0.003
	(0.066)	(0.093)	(0.042)	(0.064)	(0.079)	(0.039)
Land × rainfall × male	−0.128*	−0.113	−0.170***	−0.238**	−0.228***	−0.135***
	(0.069)	(0.097)	(0.043)	(0.094)	(0.079)	(0.032)
Land × male	0.045	0.038	0.026	0.143**	0.142**	0.038
	(0.055)	(0.087)	(0.046)	(0.070)	(0.065)	(0.035)
Male	0.037	−0.054	0.033	0.064	−0.093	0.040
	(0.026)	(0.053)	(0.049)	(0.079)	(0.125)	(0.078)
Child is 8 years	0.032	0.057	0.125**	0.037	−0.231***	0.058
	(0.022)	(0.052)	(0.051)	(0.048)	(0.081)	(0.073)
Child is 9 years	0.0819***	0.187***	0.249***	0.009	−0.113	0.003
	(0.028)	(0.054)	(0.057)	(0.040)	(0.087)	(0.066)
Child is 10 years	0.0617**	0.209***	0.415***	0.008	−0.194**	0.103
	(0.025)	(0.054)	(0.058)	(0.038)	(0.081)	(0.066)
Child is 11 years	0.0962***	0.266***	0.426***	0.024	−0.156*	0.108
	(0.035)	(0.056)	(0.066)	(0.045)	(0.082)	(0.073)

(*continued on next page*)

Table 2: *Continued*

Variables	Enrolment			Relative grade attainment		
	1994 **(1)**	**1999** **(2)**	**2004** **(3)**	**1994** **(4)**	**1999** **(5)**	**2004** **(6)**
Child is 12 years	0.139***	0.313***	0.512***	0.043	−0.179**	0.152**
	(0.034)	(0.055)	(0.056)	(0.041)	(0.081)	(0.065)
Child is 13 years	0.205***	0.344***	0.440***	0.060	−0.164**	0.079
	(0.039)	(0.057)	(0.060)	(0.042)	(0.079)	(0.066)
Child is 14 years	0.0941***	0.305***	0.442***	0.043	−0.193**	0.131**
	(0.034)	(0.059)	(0.059)	(0.042)	(0.081)	(0.063)
Number of adult males	0.014	−0.009	0.008	0.004	0.011	−0.003
	(0.009)	(0.010)	(0.013)	(0.008)	(0.012)	(0.011)
Number of adult females	−0.0192**	0.005	0.0300*	−0.0163*	−0.0319**	0.008
	(0.008)	(0.012)	(0.016)	(0.009)	(0.013)	(0.013)
Mother's age	0.001	−0.00320***	−0.002	0.001	0.0001	−0.001
	(0.001)	(0.001)	(0.002)	(0.001)	(0.001)	(0.002)
Age of the household head	0.0001	0.0001	0.0001	0.0001	0.002	0.001
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Sample size	2,047	1,877	1,629	2,047	1,877	1,629

Notes: Robust standard errors in parentheses; ***significant at 1%, **significant at 5%, *significant at 10%. Reference category is female aged 7 years; village fixed effects and age interacted with male dummies are included but not reported.

The regression results from the preferred conditional demand function (1) are reported in Tables 3–5 for 1994, 1999 and 2004, respectively.[14] Our preferred IV estimates for enrolment are reported in columns 2 and 4 of Tables 3–5. We control for age, gender and the interaction of these to account for age, gender and age–gender-specific differences in enrolments.[15] There is a positive relationship between age and enrolment. The coefficient estimates on age dummies from column 2 of Table 3 imply that the probability of a 13-year-old being enrolled is 11 percentage points higher than the probability of an 11-year-old being enrolled. The parameter estimates from the 1999 regressions depict increase in enrolment probabilities among all age groups capturing both timely enrolments and continued enrolments. By 2004, the probability of an 8-year-old being enrolled is statistically significant and is 9 percentage points higher than the probability of an 8-year-old being enrolled in 1994, reflecting improvements in timely enrolments.

Estimates for the parental schooling variables reported in column 2 of Tables 3–5 indicate a strong positive relationship between parental schooling and enrolment status. A child whose mother has had any schooling is 7 percentage points more likely to be enrolled in 1994, 3 percentage points more likely to be enrolled in 1999 and 7 percentage points more likely to be enrolled in 2004 compared with a child whose mother has no schooling. A child whose father has any schooling is 10 percentage points more likely to be enrolled in 1994, 6 percentage points more likely to be enrolled in 1999 and 7 percentage points more likely to be enrolled in 2004. As children of educated parents may be more likely to be enrolled on time (at an earlier age) compared with a child of a less educated parent, we added interaction terms between the age dummies and the parental schooling variables as additional regressors in our preferred specification. These interaction terms were all jointly insignifi-

[14] We include dummy variables to capture missing observations in our socio-economic characteristics that were replaced by the sample mean in our regression results. These imputations were done for parental schooling, mother's age and age of the head of the household. Fewer than 2% of observations are missing for age of the household head, about 10% for mother's age and 15% for parental schooling.

[15] We test whether the impact of socio-economic characteristics varied by gender. A joint test on the interaction between the gender dummy and the socio-economic characteristics from 1994 yields an F-statistic of 1.11 (P-value $= 0.35$). A similar test on the pooled sample from 1999 yields an F-statistic of 1.39 (P-value $= 0.20$) and that from 2004 yields an F-statistic of 0.62 (P-value $= 0.80$). As the impact of these socio-economic characteristics does not vary by gender, we pool males and females. The age and gender interaction terms in the pooled specifications allow for age–gender-specific differences in schooling attainments.

Table 3: Determinants of Enrolment and Relative Grade Attainment, 1994

Covariates	Enrolment				Relative grade attainment			
	OLS	IV	OLS	IV	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
PCE	0.031***	0.055	0.05***	0.123*	0.024**	0.188***	0.03**	0.270***
	(0.01)	(0.06)	(0.01)	(0.07)	(0.011)	(0.059)	(0.015)	(0.07)
PCE × male			−0.04**	−0.110*			−0.018	−0.164**
			(0.019)	(0.06)			(0.019)	(0.066)
Mother's schooling	0.085**	0.07*	0.11**	0.09*	0.074	0.029	0.14**	0.076
	(0.04)	(0.043)	(0.046)	(0.05)	(0.04)	(0.05)	(0.05)	(0.06)
Mother's schooling × male			−0.055	−0.039			−0.13***	−0.098**
			(0.04)	(0.04)			(0.04)	(0.05)
Father's schooling	0.107***	0.10***	0.08***	0.08***	0.113***	0.116***	0.095***	0.097***
	(0.02)	(0.02)	(0.025)	(0.026)	(0.02)	(0.02)	(0.02)	(0.02)
Father's schooling × male			0.05	0.04			0.03	0.027
			(0.03)	(0.03)			(0.03)	(0.03)
Mother's age	0.0009	0.0008	0.0009	0.0008	0.0001	0.0004	0.0001	0.0004
	(0.0008)	(0.0008)	(0.0007)	(0.0008)	(0.0008)	(0.0008)	(0.0008)	(0.0009)
Age of the household head	−0.0001	−0.00004	0.00009	0.00002	0.0001	0.0004	0.00002	0.0004
	(0.0005)	(0.00005)	(0.00007)	(0.00005)	(0.0004)	(0.0005)	(0.0002)	(0.0005)
Number of adult males	0.014	0.013	0.014	0.013	0.002	−0.003	0.002	−0.002
	(0.009)	(0.0086)	(0.009)	(0.0086)	(0.007)	(0.008)	(0.007)	(0.008)
Number of adult females	−0.017**	−0.016*	−0.017**	−0.016*	−0.017*	−0.008	−0.016*	−0.009
	(0.008)	(0.008)	(0.008)	(0.008)	(0.009)	(0.009)	(0.009)	(0.009)
Male	0.017	0.018	0.18**	0.429*	0.056	0.066	0.140	0.694**
	(0.02)	(0.02)	(0.076)	(0.23)	(0.07)	(0.07)	(0.11)	(0.27)

Dummy variables for child age								
Child is 8 years	0.029	0.029	0.034	0.034	0.035	0.035	0.040	0.043
	(0.02)	(0.02)	(0.02)	(0.02)	(0.04)	(0.04)	(0.04)	(0.05)
Child is 9 years	0.083***	0.084***	0.087***	0.094***	0.009	0.02	0.011	0.033
	(0.02)	(0.02)	(0.02)	(0.02)	(0.04)	(0.04)	(0.04)	(0.03)
Child is 10 years	0.058**	0.058**	0.062**	0.062**	0.004	0.003	0.006	0.008
	(0.025)	(0.025)	(0.025)	(0.025)	(0.03)	(0.03)	(0.03)	(0.04)
Child is 11 years	0.093***	0.094***	0.096***	0.099***	0.019	0.028	0.024	0.037
	(0.03)	(0.03)	(0.03)	(0.03)	(0.04)	(0.04)	(0.04)	(0.04)
Child is 12 years	0.137***	0.136***	0.138***	0.135***	0.04	0.034	0.043	0.032
	(0.03)	(0.03)	(0.03)	(0.03)	(0.04)	(0.04)	(0.04)	(0.04)
Child is 13 years	0.204***	0.206***	0.207***	0.211***	0.05	0.06	0.058	0.072
	(0.03)	(0.03)	(0.03)	(0.03)	(0.04)	(0.044)	(0.04)	(0.047)
Child is 14 years	0.092***	0.092***	0.093***	0.096***	0.04	0.044	0.041	0.05
	(0.03)	(0.03)	(0.03)	(0.03)	(0.04)	(0.04)	(0.04)	(0.04)
Kleibergen–Paap Wald *F*-statistic		41.97		20.58		41.97		20.58
Hansen *J*-statistic		2.53		4.03		0.40		0.79
		[0.11]		[0.13]		[0.52]		[0.67]
Chi-squared test for equality of maternal and paternal schooling parameters	0.23	0.33			0.67	2.62		
	[0.63]	[0.56]			[0.41]	[0.10]		
Chi-squared test for equality of maternal and paternal schooling effects on male children			1.55	1.37			4.81	6.15
			[0.21]	[0.24]			[0.02]	[0.01]
Chi-squared test for equality of maternal and paternal schooling effects on female children			0.36	0.06			0.67	0.09
			[0.54]	[0.81]			[0.41]	[0.76]

(*continued on next page*)

Table 3: *Continued*

Covariates	Enrolment				Relative grade attainment			
	OLS	IV	OLS	IV	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Hausman's test: PCE		−0.19		−0.18		−0.09		−0.11
		(0.11)		(0.12)		(0.12)		(0.11)
Hausman's test: PCE × male				−0.05				0.02
				(0.04)				(0.04)

Notes: Robust standard errors in round brackets; ***significant at 1%, **significant at 5%, *significant at 10%. Reference category is female aged 7 years. Village fixed effects and age interacted with male dummies are included but not reported; *P*-values reported in square brackets for the Hansen *J*-statistic, and the chi-squared tests. Sample size is 2,047. IVs used in columns 2 and 6 are land and lagged livestock units. IVs used in columns 4 and 8 are land, lagged livestock units, land × male dummy and lagged livestock units × male dummy. The first-stage regression results are reported in Tables A1 and A2.

Table 4: Determinants of Enrolment and Relative Grade Attainment, 1999

Covariates	Enrolment				Relative grade attainment			
	OLS	IV	OLS	IV	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
PCE	0.041**	0.379*	0.05**	0.495**	0.045***	0.498**	0.06***	0.561***
	(0.018)	(0.19)	(0.02)	(0.20)	(0.01)	(0.21)	(0.02)	(0.20)
PCE × male			−0.018	−0.14*			−0.045*	−0.084
			(0.02)	(0.08)			(0.026)	(0.07)
Mother's schooling	0.061	0.031	0.06	0.027	0.006	−0.034	−0.01	−0.05
	(0.04)	(0.04)	(0.04)	(0.05)	(0.03)	(0.04)	(0.05)	(0.06)
Mother's schooling × male			0.003	0.010			0.045	0.041
			(0.05)	(0.06)			(0.07)	(0.07)
Father's schooling	0.093***	0.061*	0.07**	0.04	0.106***	0.064*	0.088**	0.047
	(0.027)	(0.035)	(0.03)	(0.44)	(0.03)	(0.037)	(0.037)	(0.04)
Father's schooling × male			0.03	0.026			0.03	0.03
			(0.04)	(0.05)			(0.04)	(0.05)
Mother's age	−0.003***	−0.004***	−0.003***	−0.004***	−0.00002	−0.001	−0.0003	−0.001
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Age of the household head	0.0004	−0.0004	0.0002	−0.0006	0.001	0.001	0.001	0.001
	(0.0009)	(0.001)	(0.0009)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Number of adult males	−0.003	0.021	−0.0028	0.025	0.017	0.051**	0.018	0.053**
	(0.01)	(0.01)	(0.01)	(0.01)	(0.011)	(0.022)	(0.11)	(0.02)
Number of adult females	0.008	0.033*	0.008	0.036*	−0.025**	0.008	−0.024**	0.009
	(0.01)	(0.01)	(0.01)	(0.019)	(0.012)	(0.018)	(0.012)	(0.02)

(*continued on next page*)

Table 4: *Continued*

Covariates	Enrolment				Relative grade attainment			
	OLS	IV	OLS	IV	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Male	−0.068	−0.077	−0.12	0.50	−0.091	−0.10	0.061	0.21
	(0.05)	(0.05)	(0.12)	(0.34)	(0.12)	(0.12)	(0.15)	(0.32)
Dummy variables for child age								
Child is 8 years	0.056	0.062	0.057	0.063	−0.22***	−0.218**	−0.224***	−0.216**
	(0.05)	(0.05)	(0.05)	(0.06)	(0.08)	(0.08)	(0.08)	(0.08)
Child is 9 years	0.182***	0.16***	0.182***	0.15**	−0.11	−0.145	−0.116	−0.146
	(0.05)	(0.05)	(0.05)	(0.06)	(0.08)	(0.08)	(0.08)	(0.09)
Child is 10 years	0.202***	0.16***	0.201***	0.154**	−0.20**	−0.25***	−0.20**	−0.25***
	(0.05)	(0.06)	(0.05)	(0.06)	(0.08)	(0.08)	(0.08)	(0.09)
Child is 11 years	0.258***	0.20***	0.254***	0.178**	−0.16**	−0.238***	−0.16**	−0.25***
	(0.05)	(0.07)	(0.05)	(0.07)	(0.08)	(0.09)	(0.08)	(0.09)
Child is 12 years	0.312***	0.311***	0.315***	0.311***	−0.18**	−0.181**	−0.18**	−0.18**
	(0.05)	(0.05)	(0.05)	(0.06)	(0.08)	(0.08)	(0.08)	(0.08)
Child is 13 years	0.338***	0.301***	0.336***	0.294***	−0.168**	−0.217**	−0.17**	−0.22**
	(0.05)	(0.06)	(0.05)	(0.06)	(0.07)	(0.08)	(0.07)	(0.08)
Child is 14 years	0.30***	0.271***	0.30***	0.26***	−0.195**	−0.23***	−0.19**	−0.23**
	(0.05)	(0.06)	(0.05)	(0.07)	(0.07)	(0.09)	(0.08)	(0.09)
Kleibergen–Paap Wald *F*-statistic		8.56		4.35		8.56		4.35
Hansen *J*-statistic		3.29		3.66		7.13		6.98
		[0.07]		[0.15]		[0.007]		[0.03]

Chi-squared test for equality of maternal and paternal schooling parameters	0.33 [0.55]	0.29 [0.58]			2.96 [0.08]	2.30 [0.12]		
Chi-squared test for equality of maternal and paternal schooling effects on male children			0.46 [0.49]	0.20 [0.65]			1.57 [0.20]	1.12 [0.29]
Chi-squared test for equality of maternal and paternal schooling effects on female children			0.04 [0.84]	0.06 [0.81]			1.67 [0.19]	1.35 [0.24]
Hausman test: PCE		−0.71 (0.41)		−0.79 (0.42)		−1.06 (0.44)		−1.10 (0.45)
Hausman test: PCE × male				0.08 (0.06)				0.05 (0.06)

Notes: See Table 3. Sample size is 1,877.

Table 5: Determinants of Enrolment and Relative Grade Attainment, 2004

Covariates	Enrolment					Relative grade attainment				
	OLS (1)	IV (2)	OLS (3)	IV (4)	IV (5)	OLS (6)	IV (7)	OLS (8)	IV (9)	IV (10)
PCE	0.06***	0.14	0.08***	0.25**	0.25***	0.041**	0.19*	0.039**	0.21**	0.15**
	(0.01)	(0.10)	(0.02)	(0.10)	(0.07)	(0.019)	(0.09)	(0.019)	(0.09)	(0.05)
PCE × male			−0.03	−0.19**	−0.19**			0.005	−0.024	−0.03
			(0.02)	(0.08)	(0.08)			(0.03)	(0.07)	(0.07)
Mother's schooling	0.077**	0.0682**	0.13***	0.11**	0.14**	0.081**	0.064	0.13***	0.115***	0.13***
	(0.03)	(0.03)	(0.04)	(0.047)	(0.048)	(0.03)	(0.04)	(0.03)	(0.04)	(0.03)
Mother's schooling × male			−0.11**	−0.08	−0.08			−0.094*	−0.097*	−0.09*
			(0.05)	(0.05)	(0.05)			(0.04)	(0.05)	(0.05)
Father's schooling	0.077**	0.068**	0.03	0.028	0.04	0.04	0.03	−0.002	−0.019	−0.008
	(0.03)	(0.03)	(0.03)	(0.04)	(0.03)	(0.03)	(0.03)	(0.03)	(0.03)	(0.03)
Father's schooling × male			0.08*	0.082	0.087*			0.09**	0.094**	0.097**
			(0.05)	(0.05)	(0.05)			(0.04)	(0.04)	(0.04)
Mother's age	−0.002	−0.002	−0.002	−0.001	−0.001	−0.0012	−0.001	−0.001	−0.001	−0.0009
	(0.01)	(0.01)	(0.01)	(0.001)	(0.001)	(0.002)	(0.002)	(0.002)	(0.002)	(0.001)
Age of the household head	−0.0003	−0.0008	−0.0003	−0.0009	−0.0006	0.0009	0.0001	0.0009	0.0001	0.0006
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.0009)
Number of adult males	0.001	0.0006	0.001	0.004	0.004	−0.006	−0.006	−0.006	−0.006	−0.009
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Number of adult females	0.029*	0.038**	0.029*	0.040**	0.059***	0.008	0.023**	0.008	0.023**	0.025**
	(0.015)	(0.018)	(0.015)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Male	0.009	0.198	0.14	0.762**	0.764**	0.031**	0.049	−0.010	0.126	0.161
	(0.04)	(0.04)	(0.11)	(0.32)	(0.33)	(0.07)	(0.07)	(0.12)	(0.29)	(0.29)

Child age										
Child is 8 years	0.12**	0.118**	0.11**	0.114**	0.112**	0.054	0.049	0.052	0.051	0.060
	(0.05)	(0.05)	(0.05)	(0.05)	(0.05)	(0.07)	(0.07)	(0.07)	(0.07)	(0.06)
Child is 9 years	0.247***	0.253***	0.24***	0.259***	0.265***	0.001	0.051	−0.003	0.009	0.009
	(0.05)	(0.05)	(0.05)	(0.06)	(0.06)	(0.066)	(0.07)	(0.06)	(0.06)	(0.06)
Child is 10 years	0.415***	0.421***	0.41***	0.414***	0.417***	0.105	0.11*	0.102	0.11*	0.11*
	(0.05)	(0.05)	(0.05)	(0.05)	(0.05)	(0.06)	(0.06)	(0.06)	(0.06)	(0.06)
Child is 11 years	0.431***	0.418***	0.43***	0.439***	0.445***	0.114	0.13*	0.11	0.13*	0.13*
	(0.06)	(0.05)	(0.06)	(0.06)	(0.06)	(0.07)	(0.06)	(0.07)	(0.07)	(0.07)
Child is 12 years	0.528***	0.539***	0.52***	0.544***	0.55***	0.160**	0.179***	0.156**	0.17***	0.17***
	(0.08)	(0.05)	(0.05)	(0.05)	(0.05)	(0.06)	(0.07)	(0.06)	(0.06)	(0.06)
Child is 13 years	0.458***	0.463***	0.45***	0.461***	0.45***	0.089	0.09	0.08	0.09	0.09
	(0.06)	(0.06)	(0.06)	(0.06)	(0.06)	(0.06)	(0.06)	(0.06)	(0.06)	(0.06)
Child is 14 years	0.440***	0.445***	0.43***	0.442***	0.45***	0.129**	0.137**	0.127**	0.13**	0.13**
	(0.05)	(0.05)	(0.05)	(0.05)	(0.06)	(0.06)	(0.06)	(0.06)	(0.06)	(0.06)
Village fixed effects	Yes	Yes	Yes	Yes	No	Yes	Yes	Yes	Yes	No
Distance					−0.020					−0.02
					(0.01)					(0.01)
Distance × male					0.002					−0.01
					(0.01)					(0.01)
Piped water					0.053					0.014
					(0.06)					(0.02)
Electricity					0.0003					0.10
					(0.02)					(0.07)
Kleibergen–Paap Wald *F*-statistic		26.24		13.34	38.37		26.24		13.34	38.37
Hansen *J*-statistic		3.42		3.07	0.66		0.182		0.42	0.88
		[0.06]		[0.21]	[0.72]		[0.66]		[0.81]	[0.64]

(*continued on next page*)

Table 5: *Continued*

Covariates	**Enrolment**					**Relative grade attainment**				
	OLS (1)	**IV (2)**	**OLS (3)**	**IV (4)**	**IV (5)**	**OLS (6)**	**IV (7)**	**OLS (8)**	**IV (9)**	**IV (10)**
Chi-squared test for equality of maternal and paternal schooling parameter	0.00 [0.99]	0.00 [0.99]				0.33 [0.56]	0.32 [0.57]			
Chi-squared test for equality of maternal and paternal schooling effects on male children			1.91 [0.16]	1.47 [0.22]	1.04 [0.30]			0.46 [0.49]	0.48 [0.49]	0.36 [0.54]
Chi-squared test for equality of maternal and paternal schooling effects on female children			2.29 [0.13]	1.30 [0.25]	2.15 [0.14]			4.96 [0.02]	4.57 [0.03]	5.84 [0.01]
Hausman test: PCE		0.39 (0.25)		0.33 (0.21)			0.09 (0.23)		0.09 (0.21)	
Hausman test: PCE $\times$ male				0.05 (0.16)					−0.04 (0.11)	

Notes: See Table 3. IVs used in columns 2 and 7 are land and lagged livestock units. IVs used in columns 4, 5, 9 and 10 are land, lagged livestock units, land $\times$ male dummy and lagged livestock units $\times$ male dummy. Sample size is 1,629.

cant at 1% significance level for all the three enrolment regressions.[16] To assess whether the estimated impact of parental schooling differs by gender (see Glick, 2008), we interacted parental schooling with the male dummy. The chi-squared tests reported in Tables 3–5 indicate that in most specifications we cannot reject the null that the estimated impact of mothers' and fathers' schooling is equal across children of the same gender. That said, the magnitude of these coefficients suggests that mother's schooling has marginally (though not statistically significant) higher impacts on girl's enrolment, while father's schooling has higher estimated effects on boy's enrolment.

An OLS estimate of the coefficient on PCE is likely to be biased and inconsistent due to the following: (a) simultaneity bias arising from the joint determination of household expenditures and the demand for schooling inputs—a feature of conditional demand functions, (b) omitted variables bias resulting from the potential correlation between household-specific time-invariant unobservables (parent's preferences and time discount rate) and PCE, and (c) the presence of random measurement error in data which biases the estimated coefficient on PCE towards zero.[17] For these reasons, we use an IV estimator where we use land and lagged livestock units as instruments.[18] As described above, land allocations are exogenous. Livestock, however, can be bought and sold to meet current period consumption needs and thus jointly determined with PCE. For this reason, we use lagged livestock holdings.

The IV estimates of log(PCE) reported in column 2 of Tables 3–5 are generally larger than the OLS estimates found in the first column of these tables. This is consistent with any upward bias in the OLS estimates due to omitted variables being outweighed by downward bias due to measurement error. We

[16] The chi-squared statistic values on the interaction between the age dummies and mother's schooling for the enrolment regressions in 1994, 1999 and 2004 are (*P*-values in parentheses) 5.06 (0.65), 10.01 (0.18) and 9.18 (0.24) respectively. The chi-squared statistic values on the interaction between the age dummies and father's schooling in the enrolment regressions from 1994, 1999 and 2004 are (*P*-values in parentheses) 9.74 (0.20), 6.83 (0.44) and 5.44 (0.60), respectively.

[17] Under the strong assumption that there is no random measurement error in the other regressors, random measurement error in PCE causes an attenuation bias (see Wooldridge, 2002, pp. 73–76).

[18] The first-stage regression results are provided in Tables A1 and A2. Table A1 corresponds to the IV estimates reported in columns 2 and 6 of Tables 3 and 4 and columns 2 and 7 of Table 5. Table A2 corresponds to the IV estimates reported in columns 4 and 8 of Tables 3 and 4 and columns 4 and 9 of Table 5.

evaluate these estimates at average enrolment rates to obtain our measure on income elasticity of demand for schooling. Using the IV estimates reported in column 2 of Tables 3–5 and the mean enrolment rates reported in Table 1, the enrolment-income elasticity values are 0.38 (1994), 0.97 (1999) and 0.31 (2004). These estimates suggest that the demand for schooling is significantly responsive to changes in household income. We also allow these income effects to vary by gender. In column 4 of Tables 3–5, PCE and its interaction with the male dummy are both treated as endogenous. We find that the estimated impact of income is smaller for males when compared with females. This is consistent with Mani *et al.* (2012), who find that the history of schooling inputs and resources matters more for females than for males in rural Ethiopia.

Our IV estimates are robust to concerns regarding instrument validity. Instruments are considered valid only if they are strongly correlated with the endogenous regressor and uncorrelated with the error term in the second-stage regressions. The instruments used here are both strongly correlated with the endogenous regressor and uncorrelated with the error term in the second-stage regression. We first verify that the endogenous regressor is strongly correlated with the IV using the Kleibergen–Paap Wald *F*-statistic, which is robust to the presence of heteroskedasticity, autocorrelation and clustering (Kleibergen and Paap, 2006). In the presence of a single endogenous regressor, the Kleibergen–Paap test statistic reduces to the first-stage *F*-statistic on the excluded instruments. Staiger and Stock (1997) suggest that in the presence of a single endogenous regressor, instruments are deemed to be weak if the first-stage *F*-statistic on the excluded instruments is less than 10. Here, the Kleibergen–Paap Wald *F*-statistic on the excluded instruments reported in our regressions exceeds 10 in nearly all cases, rejecting the null of weak correlation between the instruments and the endogenous regressor.

Second, we verify that the IVs are uncorrelated with the error term in the second-stage regressions. Land can be treated as exogenous given the institutional policies which do not allow households to influence land allocation decisions. However, lagged livestock units can still be correlated with time-invariant household-specific unobservables. To ensure that the instruments are uncorrelated with the error term in the second-stage regression, we report both the Hansen *J*-statistic and the Hausman (1978) specification test results. We cannot reject the Hansen *J*-statistic for majority of the specifications except for the 1999 results. To check further, we use the Hausman test, where, under the null of exogeneity in lagged livestock units, PCE estimates obtained using both land and lagged livestock units as IVs are both consistent and efficient. However, if this assumption fails, the alternate estimator, that is, where PCE estimates are obtained only using land as an IV, must be chosen. We report the Hausman test statistic on PCE in Tables 3–5. In all cases, except for grade progression in 1999, we do not reject the null of exogeneity of lagged livestock units.[19]

The 2004 survey round collected information on village-level supply-side characteristics. In column 5 of Table 5, we drop the village dummy variables and replace them with three covariates: distance to the nearest primary school (measured in kilometres), a dummy variable to capture availability of electricity and a dummy variable to capture availability of piped water. However, while these supply-side characteristics are correctly signed, they are not statistically significant.

We have assumed that the coefficient estimates on the right-hand-side variables do not remain constant over time. As a check on this assumption, we report the results from pooling all 3 years' data together in Table A3. The chi-squared test rejects the null of equality of estimates for all categories of right-hand-side covariates (household consumption, parental characteristics and village effects).

4.2 Relative grade attainment

Schooling enrolments depict short-term investments in education. They do not capture regular attendance and grade advancement. To capture long-term investments in schooling and grade advancement, authors have previously used completed grades of schooling as an outcome variable of interest. However, where delayed enrolment is widespread—as is the case here—many children will have not yet started schooling, thus a large number of observations are censored at zero. Further, observations on completed grades of schooling will be right-censored for children currently enrolled in school. Both sources of censoring make standard OLS estimates on the right-hand-side variables biased.

One way to address this problem is to restrict the sample to include observations on children with completed schooling spells alone and estimate an ordered probit model. However, selective out-migration of such individuals

[19] As a further consistency check for the 1999 results, we estimated an alternate specification where PCE is instrumented with land $\times$ rainfall and lagged livestock units. The Hansen J using these alternate IVs for the specification reported in column 8 of Table 4 is 1.32 (P-value $= 0.51$), which does not reject the null of exogeneity in lagged livestock units. We also find that PCE reported in this alternate specification is not statistically significantly different from the PCE estimates reported in column 8 of Table 4.

is a concern as is the fact that the right-hand-side variables used to characterise the determinants of completed grades may not be representative of the actual socio-economic environment that affected these attainments. King and Lillard (1983, 1987) use a censored ordered probit specification where children who have completed their entire schooling spells (uncensored observations) and children who have not completed their entire schooling spell (censored observations) enter the likelihood function separately. While this addresses both sources of censoring bias, it relies on the strong assumption that children who belong to the uncensored category do not re-enter schools. To use a censored ordered probit model, we would have to include high school-age children and this would create sample selection bias of unknown magnitude and direction. Finally, even though our estimates are likely to suffer from downward censoring bias for children whose final completed grades of schooling are unknown, our estimates represent lower bounds on the estimated effect of household income and parental schooling on child's relative grade attainment.

Given all this, relative grade attainment (actual grade attained divided by potential grade given age) is an attractive measure to use as it accounts for delays in enrolments as well as grade attained conditional on age. Individuals with the same completed grades of schooling are treated differently depending upon their age, except if the actual completed grade is zero. As pooling tests reject the null of pooling all three waves together (see Table A3), we focus our attention on the results reported in columns 6 of Tables 3 and 4 and column 7 of Table 5. We use the same control variables in the relative grade attainment regressions as we used in the enrolment regressions.[20]

Parental schooling is positively associated with relative grade attainment, though it is not consistently statistically significant. To capture possible nonlinear effects, we added interaction terms between the age dummies and the parental schooling variables to our specification. These were jointly insignificant at the 5% significance level in all years.[21] We interacted parental

[20] A joint test on the interaction between the gender dummy and the socio-economic characteristics from 1994 yields an F-statistic of 1.11 (P-value $= 0.35$). A similar test on the pooled sample from 1999 yields an F-statistic of 0.65 (P-value $= 0.71$) and that from 2004 yields an F-statistic of 1.38 (P-value $= 0.18$). We conclude that the impact of the socio-economic characteristics included in these regressions does not vary by gender.

[21] The chi-squared statistic values on the interaction between the age dummies and mother's schooling for the relative grade attainment regressions in 1994, 1999 and 2004 are (P-values in parentheses) 1.30 (0.98), 7.54 (0.37) and 7.15 (0.41). The chi-squared statistic values on the interaction between the age dummies and father's schooling in the relative grade

schooling with the male dummy as well (see column 8, Tables 3 and 4 and column 9, Table 5). There are no statistically significant gender differences in the estimated impact of parental schooling.

As income increases, the gap between actual grades and potential grades decreases, improving grade progression. We evaluate these estimates at average relative grade attainment to obtain our measure on income elasticity of demand for schooling. The IV estimates reported in column 6 of Tables 3 and 4 and column 7 of Table 5 are evaluated at the average relative grade attainment reported in Table 1, to obtain the relative grade attainment—income elasticity for each survey round: 1.38 (1994), 1.81 (1999) and 0.79 (2004). Children from higher income households are always more likely to progress on time compared with children from low-income households. We also find that the demand for grade progression is more elastic compared with the demand for enrolment.

We can use the results in Tables 3 and 5 to explore why enrolment and relative grade attainment increased over the period considered here. We begin with three observations. First, between 1994 and 2004, enrolment increased by 32.7 percentage points and relative grade attainment increased by 0.11 (Table 1). Second, when we examine the mean values of the regressors (Table 1), we see while some increase substantially in value (such as log of real per capita consumption expenditure increases by 26%), others remain largely unchanged (age of household head) and some fall (number of adult females in the household). Third, comparing the enrolment regressions results from Tables 3 and 5, we see that for any given specification, there are marked differences in the coefficients on the dummy variables for children's age in 1994 and 2004, with the latter almost always much larger.

These observations suggest that increases in enrolment and relative grade attainment are being driven by both changes in parameter estimates and average value of the regressors. We examine this idea more formally by implementing an Oaxaca–Blinder type decomposition. This method decomposes differences in mean values across groups (here, survey years) into those attributable to changes in mean values of the explanatory variables, those attributable to changes in the parameter estimates associated with those explanatory variables and the fact that differences in explanatory variables and coefficients exist simultaneously between the two groups (the interaction effect). The original formulation of this approach suffered from four

attainment regressions from 1994, 1999 and 2004 are (*P*-values in the parentheses) 6.54 (0.47), 12.34 (0.08) and 8.03 (0.33).

weaknesses: (a) the dependence of the decomposition on the reference group (Oaxaca, 1973); (b) the dependence of categorical variables on the omitted category (Oaxaca and Ransom, 1999); (c) the need to construct standard errors to assess the statistical significance of the contributors to the differences in groups; and (d) the need to extend these methods beyond ordinary least squares. There is a large literature on all these issues. Jann (2008) summarises this body of work and outlines an integrated solution to all of them. Our results are based on Jann's (2008) decomposition method.

The increase in the predicted probability of enrolment, 0.327, is attributable to changes in the mean values of the regressors (0.026), the interaction effect (−0.012 and not statistically significant) and changes in the coefficients (0.313). Changes in the coefficients account for nearly all, 95.7%, of the observed change in predicted enrolment probabilities. The change in the coefficients for log real per capita consumption expenditure and the coefficients for the dummy variables for children's age accounts for 33.3 and 70.6% of the change in predicted probability enrolment, respectively. With the caveat that the change in coefficients for log real per capita consumption expenditure is statistically significant only at the 15% level, the changes in predicted probabilities of enrolment are driven entirely by changes in these two characteristics. The increase in predicted relative grade attainment, 0.11, is also attributable to changes in the mean values of the regressors (0.057), the interaction effect (−0.007 and not statistically significant) and changes in the coefficients (0.06). Changes in the coefficients account for nearly half, 54.5%, of the observed change in predicted relative grade attainment. Notice that changes in the mean value of the regressors account for the remaining half of the changes, of which, log of per capita consumption expenditure (0.049) alone accounts for 85% of the observed change. Hence, both changes in income and changes in the responsiveness of schooling to income explain for observed improvements in relative grade attainment during this period.

Explaining why we observe these changes is tricky; effectively we have one observation for the change in the coefficient on log real per capita consumption expenditure and one observation for changes in the coefficients on the age dummy variables. So our ability to provide an explanation grounded in further statistical analysis is greatly constrained. Instead, we offer several observations about these changes.

Recent work by King *et al.* (2012) shows a positive correlation between the returns to human capital and what they term economic freedom (openness to trade; absence of restrictions on market functioning and so on). As discussed by Dercon (2006), in Ethiopia the fall of the Marxist-oriented

Dergue government meant that the period after 1994 coincided with the elimination of many restrictions on market activities in both the tradable and non-tradable sectors including those such as coffee production and wage employment which are skill intensive. Taken together, these studies are suggestive that returns to schooling either may have been rising or at the very least, parents perceived that they were likely to rise. If this was the case, parents would have, at the margin, allocated more resources to schooling. If parents care about schooling directly (in their utility function) or if there are credit market constraints, then income will affect schooling choices. With credit market constraints, a higher demand for schooling would result in higher income elasticities because of increasing market returns due to economic liberalisation. These perceived higher returns to schooling might also encourage parents to let their children who previously would not have been enrolled to 'try out' schooling (hence the increased elasticity observed for enrolment between 1994 and 1999) or to persist in school (hence the increased elasticity for relative grade attainment in the same period). Other factors may also be at play. The period 1994–99 is one in which income of households in this sample grows (Dercon *et al*., 2012). Again, with credit market constraints, the income elasticity of schooling will be positively related to the rate of income growth provided that income growth itself raises the demand for schooling. This too would account for the higher elasticity that we observe in 1999 relative to 1994. However, income growth stagnates between 1999 and 2004 and there is also a gradual easing of access to credit. Following the logic of this argument, we would expect the elasticity to fall back towards its earlier level, which is what we observe for both enrolment and relative grade attainment. In addition, Mani *et al*. (2012) document that school enrolment in Ethiopia is path dependent. The positive income shock observed between 1994 and 1999—the increased income elasticity—together with the path dependency in enrolment creates a hysteresis-type effect whereby the likelihood that a child at a given age is enrolled rises, which is also what we observe here. These hysteresis effects will also raise relative grade attainment. That said, while we regard these observations as plausible explanations, we stress that this argument is speculative.

5. Conclusion

UNESCO's 2008 global monitoring report indicates that primary school net enrolment rate in sub-Saharan Africa have increased from 59% in 1999 to 70% in 2005. Despite this progress, the report notes that nearly 33 million children of primary school age are still not enrolled in school. The region as a whole was home to 45% of the world's out-of-school children in 2005, of whom 54% were girls. The low levels of schooling combined with significant gender differences in schooling points to the value of identifying the determinants of schooling for both males and females in the region.

Our objective in this paper is to contribute to the literature on schooling in sub-Saharan Africa by using cross-sectional data rich in schooling information, available in 5-year intervals between 1994 and 2004 for rural Ethiopia. We analyse the determinants of both current enrolment status and relative grade attainment for primary school-age children, addressing the potential endogeneity of household income in our regression analysis.

Using IV estimation, we find, first, a positive income effect for schooling enrolments and an even stronger effect for relative grade attainment. The income effect is masked by OLS estimation. Second, these income effects are larger for girls than for boys. Third, there exist gender-specific differences in the estimated impact of parental schooling on children's schooling, though these effects may not always be statistically significant. Finally, where we are able to control for observable community characteristics, we do not find these supply-side factors to have an important role in explaining the variation in schooling. These results suggest, therefore, that policies that address the demand-side constraints such as income with special focus on girls are likely to improve schooling attainments as well as reduce the gender differences in schooling attainments found in Ethiopia and elsewhere in sub-Saharan Africa.

We end on two speculative notes. First, much of the literature examining the determinants of schooling implicitly assumes that these determinants are time invariant. Our results call this implicit assumption into question, and suggest the need for repeated cross-sectional data analysis in order to capture episodic effects. Second, the rise in schooling induced by the income growth between 1994 and 1999, together with changes in perceived returns to schooling, is likely to have contributed to increasing children's initial contact with the school system. Such an outcome may have induced hysteresis-type effects, resulting in the impressive increases in school enrolments and relative grade attainment observed in rural Ethiopia over this period.

Acknowledgements

We thank the following organisations for funding the survey work on which this paper is based: SIDA, USAID and the World Food Programme. We thank John Ham, Jeffrey B Nugent and Jeffrey Wooldridge for comments and suggestions. We are grateful to the editor and two anonymous referees for providing excellent suggestions. We are thankful to Yisehac Yohannes for clarifying several data related questions. Financial support was provided by the College of Letters, Arts and Sciences, University of Southern California and International Food Policy Research Institute.

References

Alderman, H., J. R. Behrman and R. Menon (2001) 'Child Health and School Enrollment: A Longitudinal Analysis', *The Journal of Human Resources*, 56: 185–205.

Alderman, H., J. Hoddinott and B. Kinsey (2006) 'Long-term Consequences of Early Childhood Malnutrition', *Oxford Economic Papers*, 58 (3): 450–74.

Behrman, J. R. and J. C. Knowles (1999) 'Household Income and Child Schooling in Vietnam', *The World Bank Economic Review*, 13 (2): 211–56.

Behrman, J. R. and M. R. Rosenzweig (2002) 'Does Increasing Women's Schooling Raise the Schooling of the Next Generation?', *American Economic Review*, 92: 323–34.

Bommier, A. and S. Lambert (2000) 'Education Demand and Age at School Enrollment in Tanzania', *The Journal of Human Resources*, 35 (1): 177–203.

Brown, P. H. and A. Park (2002) 'Education and Poverty in Rural China', *Economics of Education Review*, 21: 523–41.

Central Statistical Agency [Ethiopia] and ORC Macro (2006) 'Ethiopia Demographic and Health Survey 2005.' Addis Ababa, Ethiopia and Calverton, Maryland, USA: Central Statistical Agency and ORC Macro.

Dercon, S. (2006) 'Economic Reform, Growth and the Poor: Evidence from Rural Ethiopia', *Journal of Development Economics*, 81 (1): 1–24.

Dercon, S., D. O. Gilligan, J. Hoddinott and T. Woldehanna (2009) 'The Impact of Agricultural Extension and Roads on Poverty and Consumption Growth in Fifteen Ethiopian Villages', *American Journal of Agricultural Economics*, 91 (4): 1007–21.

Dercon, S., J. Hoddinott and T. Woldehanna (2012) 'Growth and Chronic Poverty: Evidence from Rural Communities in Ethiopia', *Journal of Development Studies*, 48 (2): 238–53.

Dostie, B. and R. Jayaraman (2006) 'Determinants of School Enrollment in Indian Villages', *Economic Development and Cultural Change*, 54: 405–21.

Ezra, M. and G. E. Kiros (2001) 'Rural Out-migration in the Drought Prone Areas of Ethiopia: A Multilevel Analysis', *International Migration Review*, 35 (3): 749–71.

Fafchamps, M. and A. Quisumbing (2005) 'Marriage, Bequest and Assertive Matching in Rural Ethiopia', *Economic Development and Cultural Change*, 53: 347–80.

Glewwe, P. (2002) 'Schools and Skills in Developing Countries: Education Policies and Socioeconomic Outcomes', *Journal of Economic Literature*, XI: 463–82.

Glewwe, P. and H. Jacoby (1994) 'Student Achievement and Schooling Choice in Low-income Countries: Evidence from Ghana', *Journal of Human Resources*, 29 (3): 843–64.

Glewwe, P. and H. Jacoby (1995) 'An Economic Analysis of Delayed Primary School Enrollment in a Low Income Country: The Role of Early Childhood Nutrition', *The Review of Economics and Statistics*, 77: 156–69.

Glewwe, P. and E. Miguel (2008) 'The impact of child health and nutrition on education in less developed countries', in T. P. Schultz and John Strauss (eds), *Handbook of Development Economics*, vol. 4. Amsterdam: North Holland Press.

Glick, P. (2008) 'What Policies Will Reduce Gender Schooling Gaps in Developing Countries: Evidence and Interpretation', *World Development*, 36 (9): 1623–46.

Glick, P. and D. Sahn (2000) 'Schooling of Girls and Boys in a West African Country: The Effects of Parental Education, Income, and Household Structure', *Economics of Education Review*, 19 (1): 63–87.

Greene, W. (2004) 'The Behaviour of the Maximum Likelihood Estimator of Limited Dependent Variable Models in the Presence of Fixed Effects', *The Econometrics Journal*, 7: 98–119.

Hausman, J. A. (1978) 'Specification Tests in Econometrics', *Econometrica*, 46 (6): 1251–71.

Jann, B. (2008) 'The Blinder-Oaxaca Decomposition for Linear Regression Models', *The Stata Journal*, 8 (4): 453–79.

King, E. M. and L. A. Lillard (1983) *Determinants of Schooling Attainment and Enrollment Rates in the Philippines*. A Rand Note. Santa Monica, CA: Rand Corporation.

King, E. M. and L. A. Lillard (1987) 'Education Policy and Schooling Attainment in Malaysia and the Philippines', *Economics of Education Review*, 6 (2): 167–81.

King, E., C. Montenegro and P. Orazem (2012) 'Economic Freedom, Human Rights, and the Returns to Human Capital: An Evaluation of the Schultz Hypothesis', *Economic Development and Cultural Change*, 61 (1): 39–72.

Kleibergen, F. and R. Paap (2006) 'Generalized Reduced Rank Tests Using the Singular Value Decomposition', *Journal of Econometrics*, 133 (1): 97–126.

Lillard, L. and R. J. Willis (1994) 'Intergenerational Educational Mobility: Effects of Family and State in Malaysia', *The Journal of Human Resources*, 29 (4): 1126–66.

Maccini, S. and D. Yang (2009) 'Under the Weather: Health, Schooling, and Economic Consequences of Early-Life Rainfall', *American Economic Review*, 99 (3): 1006–26.

Mani, S., J. Hoddinott and J. Strauss (2012) 'Long-term Impact of Investments in Early Schooling – Empirical Evidence from Rural Ethiopia', *Journal of Development Economics*, 99 (2): 292–9.

Mincer, J. (1974) *Schooling, Experience, and Earnings*. New York, NY: Columbia University Press.

Oaxaca, R. (1973) 'Male-Female Wage Differentials in Urban Labor Markets', *International Economic Review*, 14: 693–709.

Oaxaca, R. L. and M. R. Ransom (1999) 'Identification in Detailed Wage Decompositions', *The Review of Economics and Statistics*, 81: 154–7.

Orazem, P. F. and E. M. King (2008) 'Schooling in developing countries: the roles of supply, demand and government policy', in T. P. Schultz and J. Strauss (eds), *Handbook of Development Economics*, vol. 4. Amsterdam: Elsevier.

Parish, W. L. and R. J. Willis (1993) 'Daughters, Education, and Family Budgets: Taiwan Experiences', *Journal of Human Resources*, 28 (4): 863–98.

Pollak, R. A. (1969) 'Conditional Demand Functions and Consumption Theory', *The Quarterly Journal of Economics*, 83 (1): 60–78.

Psacharopoulos, G. and H. A. Patrinos (2004) 'Returns to Investment in Education: A Further Update', *Education Economics*, 12 (2): 111–34.

Rosenzweig, M. R. (1990) 'Population Growth and Human Capital Investments: Theory and Evidence', *Journal of Political Economy*, 98: S38–70.

Rosenzweig, M. R. and K. J. Wolpin (1986) 'Evaluating the Effects of Optimally Distributed Public Programs', *American Economic Review*, 76 (3): 470–87.

Schaffner, J. (2004) 'Determinants of Schooling Investments among Primary School Aged Children in Ethiopia', Africa Region Human Development, Working Paper Series.

Schultz, T. P. (1988) 'Educational investments and returns', in H. Chenery and T. N. Srinivasan (eds), *Handbook of Development Economics*, vol. 1. Amsterdam: North Holland Press.

Schultz, T. P. (2002) 'Why Governments Should Invest More to Educate Girls', *World Development*, 30 (2): 207–25.

Singh, R. D. (1992) 'Underinvestment, Low Economic Returns to Education, and the Schooling of Rural Children: Some Evidence from Brazil', *Economic Development and Cultural Change*, 40 (3): 645–64.

Staiger, D. and J. H. Stock (1997) 'Instrumental Variables Regression with Weak Instruments', *Econometrica*, 65 (May): 557–86.

Strauss, J. and D. Thomas (1995) 'Human resources: empirical modeling of household and family decisions', in H. Chenery and T. N. Srinivasan (eds), *Handbook of Development Economics*, edn. 1, vol. 3. Amsterdam: Elsevier, pp. 1883–2023.

Tansel, A. (1997) 'Schooling Attainment, Parental Education, and Gender in Cote d'Ivoire and Ghana', *Economic Development and Cultural Change*, 45 (4): 825–56.

Thomas, D., J. Strauss and M. Henriques (1990) 'Child Survival, Height for Age and Household Characteristics in Brazil', *Journal of Development Economics*, 33 (2): 197–234.

White, H. (1980) 'A Heteroskedasticity-Consistent Covariance Matrix and a Direct Test for Heteroskedasticity', *Econometrica*, 48: 817–38.

Wooldridge, J. (2002) *Econometric Analysis of Cross Section and Panel Data*. Cambridge, MA: MIT Press.

Wooldridge, J. M. (2003) 'Cluster-sample Methods in Applied Econometrics', *American Economic Review*, 93: 133–8.

Appendix

Table A1: Selected First-stage Estimates for Log per Capita Consumption in Columns (2) and (6) of Tables 3 and 4 and Columns (2) and (7) of Table 5

	1994	1999	2004
	(1)	(2)	(3)
Excluded instruments			
Land	0.07*	0.08	0.05**
	(0.04)	(0.05)	(0.02)
Lagged livestock units	0.04**	0.01**	0.04**
	(0.004)	(0.004)	(0.007)
Second-stage controls			
Male	−0.02	0.03	−0.12
	(0.07)	(0.07)	(0.08)
Child is 8 years	0.01	0.002	0.02
	(0.07)	(0.07)	(0.08)
Child is 9 years	−0.04	0.06	−0.06
	(0.07)	(0.07)	(0.09)
Child is 10 years	0.02	0.12	−0.04
	(0.07)	(0.07)	(0.08)
Child is 11 years	−0.01	0.17**	−0.09
	(0.08)	(0.07)	(0.09)
Child is 12 years	0.04	0.009	−0.13
	(0.07)	(0.07)	(0.08)
Child is 13 years	−0.06	0.10	−0.05
	(0.08)	(0.07)	(0.08)
Child is 14 years	−0.0001	0.09	−0.04
	(0.08)	(0.08)	(0.08)
Number of adult males	0.015	−0.07**	−0.01
	(0.01)	(0.01)	(0.01)
Number of adult females	−0.05**	−0.07**	−0.09**
	(0.01)	(0.01)	(0.02)
Age of the household head	−0.001	0.001	0.0048**
	(0.001)	(0.001)	(0.001)
Mother's age	0.003**	0.002	0.0001
	(0.001)	(0.001)	(0.002)
Mother's schooling	0.27**	0.09**	0.10**
	(0.06)	(0.04)	(0.05)
Father's schooling	0.006	0.09**	0.09*
	(0.03)	(0.03)	(0.04)
R-square	0.29	0.49	0.30

Notes: Robust standard errors in parentheses; **significant at 5%; *significant at 10%. Village fixed effects included but not reported.

Table A2: Selected First-stage Estimates for Log per Capita Consumption and Its Interaction with Male in Columns (4) and (8) of Tables 3 and 4 and Columns (4) and (9) of Table 5

	1994	1999	2004
	(1)	**(2)**	**(3)**
PCE			
Excluded instruments			
Land	0.27**	0.09**	0.10**
	(0.06)	(0.04)	(0.05)
Land × male	0.006	0.09**	0.09*
	(0.03)	(0.03)	(0.04)
Lagged livestock units	0.07*	0.08	0.05**
	(0.04)	(0.05)	(0.02)
Lagged livestock units × male	0.04**	0.01**	0.04**
	(0.004)	(0.004)	(0.007)
PCE × male dummy			
Excluded instruments			
Land	0.27**	0.09**	0.10**
	(0.06)	(0.04)	(0.05)
Land × male	0.006	0.09**	0.09*
	(0.03)	(0.03)	(0.04)
Lagged livestock units	0.07*	0.08	0.05**
	(0.04)	(0.05)	(0.02)
Lagged livestock units × male	0.04**	0.01**	0.04**
	(0.004)	(0.004)	(0.007)

Notes: See Table A1 for notes. Second-stage controls included but not reported.

Table A3: Pooled IV Results for Enrolment and Relative Grade Attainment

	Enrolment	**Relative grade attainment**
PCE	0.46**	0.65**
	(0.21)	(0.25)
PCE × 1994	−0.40*	−0.48*
	(0.22)	(0.27)
PCE × 2004	−0.36	−0.54*
	(0.23)	(0.28)
Mother's schooling	−0.08*	−0.08*
	(0.03)	(0.04)
Mother's schooling × 1994	0.16**	0.12
	(0.05)	(0.07)
Mother's schooling × 2004	0.16**	0.16**
	(0.05)	(0.04)
Father's schooling	0.003	0.03
	(0.03)	(0.03)
Father's schooling × 1994	0.11***	0.07*
	(0.03)	(0.05)
Father's schooling × 2004	0.07*	0.003
	(0.04)	(0.04)
Chi-squared tests		
PCE and year interaction terms	7.76	19.13
	[0.05]	[0.00]
Maternal schooling and year interaction terms	12.13	8.24
	[0.00]	[0.04]
Paternal schooling and year interaction terms	29.54	23.03
	[0.00]	[0.00]
Village characteristics	328.06	236.31
	[0.00]	[0.00]
Household characteristics and year interaction terms	22.59	12.54
	[0.00]	[0.00]
Village characteristics do not change with time	72.68	76.86
	[0.00]	[0.00]
PCE has equal effect in 1994 and 1999	3.16	3.12
	[0.07]	[0.07]
PCE has equal effect in 1999 and 2004	2.38	4.02
	[0.12]	[0.07]
PCE has equal effect in 1994 and 2004	0.18	0.46
	[0.66]	[0.49]
Village × year dummies	Yes	Yes
Kleibergen–Paap rank Wald *F*-statistic	4.83	4.83

(*continued on next page*)

Table A3: *Continued*

	Enrolment	Relative grade attainment
Hansen *J*-statistic	1.75 [0.18]	3.10 [0.08]

Notes: Robust standard errors adjusted for clustering at the household level reported in round brackets; **significant at 5%; *significant at 10%; *P*-values reported in square brackets for the chi-squared tests and the Hansen *J*-statistic. In all the columns, PCE is instrumented with land, lagged livestock units, lagged livestock units interacted with the 1994 and 2004 year dummies. All other right-hand-side variables are as specified in regression Tables 2–4 along with a full set of interaction terms with the year dummies; year dummy = 1 if year is 1994, 0 otherwise; 2004 dummy = 1 if year is 2004, 0 otherwise; omitted year is category is 1999; the *F*-statistic values on the excluded IVs for the PCE, PCE $\times$ 1994 dummy and PCE $\times$ 2004 dummy are 45.6, 20.9 and 20.0, respectively.