Dads, disease, and death: determinants of daughter discrimination

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MIGRATION
Immigrants assimilate as communities, not just as individuals
T.J. Hatton · A. Leigh 389
The savings behavior of temporary and permanent migrants in Germany
T.K. Bauer · M.G. Sinning 421
Immigrant selection and short-term labor market outcomes by visa category
A. Aydemir 451
Dual-earner migration. Earnings gains, employment and self-selection
B. Rabe 477
Immigrant over-education: evidence from Denmark
C.P. Nielsen 499
USA immigration policy, source-country social programs, and the skill composition of legal USA immigration
M.J. Greenwood · J.M. McDowell 521
Change and continuity among minority communities in Britain
A. Georgiadis · A. Manning 541
When nature rebels: international migration, climate change, and inequality
L. Marchiori · I. Schumacher 569
HEALTH
Child mortality in rural India
B. van der Klaauw · L. Wang 601
Uncovering the impact of the HIV epidemic on fertility in Sub-Saharan Africa: the case of Malawi
D. Durevall · A. Lindskog 629
Household allocation decisions and child health: can behavioral responses to vitamin A supplementation programs explain heterogeneous effects?
B. Yarnoff 657
EDUCATION
The dynamics of school attainment of England's ethnic minorities
D. Wilson · S. Burgess · A. Briggs 681
Does teenage childbearing reduce investment in human capital?
D. Webbink · N.G. Martin · P.M. Visscher 701
The effects of parental leave extension on training for young women
P.A. Puhani · K. Sonderhof 731
The effect of schooling on teenage childbearing: evidence using changes in compulsory education laws
M.A. Silles 761

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DOI 10.1007/s00148-011-0357-9

ORIGINAL PAPER

Dads, disease, and death: determinants of daughter discrimination

Joyce J. Chen

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Abstract Existing evidence suggests that girls are differentially affected by income shocks and changes in bargaining power. Most studies, however, ignore household production and confound differential opportunity costs with changes in income or bargaining power. I disentangle these determinants of gender discrimination—preferences, income and time allocation—by comparing households with varying degrees of parental involvement. Results indicate that, controlling for household fixed effects, reducing the time available for household production has a disproportionately negative effect on daughters. But, for a transitory income shock, daughters' education is less income-elastic. Increasing mothers' bargaining power is most effective in narrowing the gender gap.

Keywords Intra-household allocation · Gender · Household production

JEL Classification D13 · J16 · O15

1 Introduction

Gender gaps in schooling, nutrition, health and even survival remain persistent in many countries around the world. Girls constitute nearly two-thirds of the children excluded from a basic education (UNESCO 2000), and more than

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100 million women are "missing" from global population figures (UNICEF 2000). The experience of women in developed countries suggests that innate differences in abilities and pre-dispositions cannot be the sole reason for such discriminatory behavior. Gaps between men and women seem to decline almost naturally with the process of economic development, which leads us to inquire into the determinants of discrimination and how they may be ameliorated. The focus of this paper is to identify and quantify the potential causes of discriminatory behavior, with particular attention to differences in investments in children's human capital.

A better understanding of the root causes of discrimination and careful measurement of their relative magnitudes can have a significant impact on policy and program design. Many governments have already undertaken policies aimed at reducing gender disparities—Progresa in Mexico and the Female Secondary Schooling Assistance Program in Bangladesh are well-studied examples which provide larger subsidies for girls' schooling than boys'. Microcredit programs also began with this goal in mind, targeting gender discrimination on two levels: bringing opportunity to women who have been excluded from traditional economic activities and strengthening women's position in the family. As Sen (2000) describes, increasing women's agency through economic empowerment is "both a reward on its own (with associated reduction of gender bias in the treatment of women in family decisions), and a major influence for social change in general." While these programs have had some success in closing the gender gap, it is important to consider whether other forms of intervention or targeting can be more effective. To assess this, we must acknowledge the varied sources of discrimination and consider how each will be affected by policy interventions and broader economic development.

Allocations to girls and boys may differ for a variety of reasons. The most difficult to address is preference-based gender bias, rooted in socio-cultural norms which favor males. However, some degree of discrimination may be efficient (leaving aside the question of whether it is *utility* maximizing), given the prevailing returns and opportunity costs. The theory of comparative advantage suggests that boys will receive more human capital investment for two reasons: (1) expected returns to human capital are typically lower for females, perhaps as a result of lower labor force participation, wage discrimination, or provision of old-age support through sons; and (2) when girls are trained from a very young age in household tasks, daughters will be more productive than sons in the household. The former suggests that, with diminishing marginal returns to human capital investment, gender equality is a normal good for the household; discrimination may be due to low income and poverty (see also Garg and Morduch 1998). The latter suggests that the demand for household labor is a key determinant of daughters' ability to invest in human capital; household activities and schooling are competing activities for girls. Thus, even in the absence of socio-cultural bias against women, gender discrimination may still be prevalent.

Income, time allocation and parental preferences all affect gender differences in human capital investments. But much of the existing literature neglects the role of household production and children's time allocation, thereby confounding the estimated effects of income or bargaining power with simultaneous changes in time allocation. This paper disentangles the potential causes of discrimination—preferences, income, and time allocation—in a context that allows direct comparison of the magnitudes of these effects. Four household types are utilized, highlighting different levels of parental involvement: (1) both parents present in the household and in good health (fathers contribute time to household production, income, and influence over allocation decisions) (2) fathers as long-term migrants (contribute only income and influence over allocation decisions); (3) fathers with a temporary debilitating illness (contribute only influence over allocation decisions); and (4) widowed mothers (fathers contribute no time, income or influence). Comparison of boys' and girls' schooling outcomes across these groups allows us to separately identify the effects of time allocation, income and bargaining power.

Controlling for observable characteristics and unobservable household fixed effects, I find that increasing mothers' bargaining power has the largest impact on closing the gender gap. A reduction in the amount of time available for household production reduces the probability that daughters will be enrolled in school, relative to their brothers, whereas a temporary reduction in household income, holding constant the time available for household production, has a positive effect on girls' school enrollment relative to their male siblings. These findings suggest that policies aimed at empowering women and increasing gender equity will be most effective when constraints on women and children's time allocation are taken into account. On the other hand, girls' schooling is found to be less income-elastic, at least with regard to transitory shocks, which suggests that short-term cash transfer programs may be ineffective at reducing gender disparities unless specific conditions are imposed.

The following section discusses the roles of household income, household production and parental preferences in determining gender disparities in human capital, as explored in the existing literature on intra-household allocation. Section 3 describes the data, empirical methodology, and challenges to identification. Results are presented in Section 4, with a discussion of robustness in Section 5.

2 Income, time allocation, parental preferences and children's schooling

Schooling is typically believed to be a normal good and, with diminishing marginal returns to schooling investments per child, higher income will also lead to more equality within the household. Equality of allocations among children may also, itself, be a normal good. Behrman and Deolalikar (1990) estimate more negative food price elasticities for females in India. This suggests that, while women benefit disproportionately from falling food prices, the nutritional burden of a rise in food prices is also borne disproportionately by women. The authors note that, to the extent that food intake is more critical in lean seasons when agricultural prices are high, women are thus put at greater risk for malnutrition or starvation. Cameron and Worswick (2001) utilize data on crop loss to directly estimate transitory income for households in Indonesia. The authors find that expenditure on girls' education is more responsive to fluctuations in transitory income, suggesting that it is a luxury good for the household.

These findings suggest that equality within the household is a normal good, increasing in income. But it is difficult to rule out the possibility that the estimated effects are driven by concurrent changes in time allocation induced by these shocks, not the changes in income per se. Rising food prices and crop loss likely increase time spent in income-generating activities. Kochar (1999) finds that, when faced with negative crop income shocks, households in rural India smooth consumption via changes in market labor supply, with changes in wage income compensating for as much as one-third of the crop income shock. To the extent that wage labor consists predominantly of labor-intensive activities with high returns to nutrition and higher returns to males than females, the nutritional requirements for males will also rise relative to female household members. Increased time in income-generating activities also reduces the time available for household production, which will have a disproportionate negative effect on girls' schooling if girls are more productive in the home than are boys.

Rose (1999) finds that, in rural India, rainfall shocks early in life affect the probability of survival for girls relative to boys, with girls' survival being more income-elastic, i.e., girls suffer more from negative shocks but also benefit more from positive shocks. Rainfall shocks are likely to affect time allocation and the demand for nutrients among other household members, but shocks occurring in infancy should not *differentially* affect the returns to nutrition/health for boys and girls. However, if negative rainfall shocks reduce the proportion of household income earned by women, e.g., due to decreased opportunities for wage labor or reduced time available for non-crop activities (see Kochar 1995), the disproportionately negative effect on girls may be the result of a decline in mothers' bargaining power rather than an income effect per se.

Shocks that affect the demand for non-farm household labor may also affect girls and boys differently. Children often engage in household production, e.g., caring for younger siblings or assisting in household chores which, in turn, can free up parents' time for more productive tasks. If productivity differs among sons and daughters, a shock affecting the demand for household labor will have a larger effect on the opportunity cost of education for the more productive of the two sexes. Dammert (2010) and Edmonds (2006) find that girls' household labor is more sensitive to sibling size and composition, compared to their male siblings. And, using data from Indonesia, Pitt and Rosenzweig (1990) find that infant morbidity in the household causes teenaged daughters to increase their time in household care and reduce time in school attendance, wage labor and leisure, with the effect being more pronounced than for their teenaged male siblings. Even without a change in the household time endowment, income, or decision-making power, girls are differentially affected by a change in the demand for household labor, with infant morbidity leading to a 15% decline in the likelihood of school attendance for girls in Indonesia.

The possible consumption value associated with investments in children's schooling suggests that preferences also matter, and thus control over income matters. If women prefer to allocate more goods to the household and to children than do men, increased control over income (empowerment) for women should lead to higher investments in children. Girls will also benefit disproportionately from an increase in women's bargaining power if women have stronger preferences for daughters or for gender equality. Both hypotheses have been well-documented in a number of papers (see, for example, Thomas 1990a, b; Duflo 1999; Qian 2008). However, Pitt and Khandker (1998) find that women's participation in microcredit programs has a larger positive effect on schooling for sons than for daughters. The authors suggest that this is due to the fact that boys are poor substitutes for women's time and thus are less likely to be drawn into household and/or self-employment activities. Here, the (positive) effect of an increase in women's bargaining power is offset by concurrent changes in time allocation induced by the change in income. The relationship between children's schooling and household income depends critically on both who controls the income as well as the source from which the income is derived and its effect on the demand for child labor.

Gertler et al. (2004a, b) find that parental death has large and significant adverse effects on children's health and school enrollment, and Gertler et al. (2004a, b) find an adverse effect on child schooling as well. Children in households with high permanent income do not experience a smaller reduction in school enrollment following parental death (Gertler et al. 2004a, b), and the inclusion of changes in household consumption before and after parental death does not significantly reduce the magnitude or significance of the coefficient on parental death (Gertler et al. 2004a, b). These findings suggest that parental loss indeed has adverse effects, but the associated changes in income play a relatively small role in the reductions in children's human capital. Instead, the authors suggest that it is the *presence* of parents that plays a key role in children's human capital accumulation. However, it is not clear what specifically this "presence" entails. Parental death affects the aggregate household time endowment and the distribution of decision-making power among household members, as well as the level of income.

In summary, the existing literature provides evidence that income, time allocation, and preferences all affect gender disparities in human capital investments. What is less clear, however, is the relative magnitude and independent effect of each of these factors. This paper aims to disentangle these three factors and to determine their relative magnitudes.

3 Data and methodology

Data are drawn from the 1990 Indonesian Population Census and the 1993 Indonesian Socio-Economic Survey (SUSENAS), and current school enrollment is the outcome of interest. The analysis does not focus on more cumulative measures of schooling because identification relies, in part, on transitory shocks. The population of interest is households in which parents are the primary decision-makers, and estimation is limited to children between the ages of 10 and 16, inclusive. In 1984, the government of Indonesia instituted compulsory schooling of 6 years, equivalent to completion of the primary level. Thus, variation in school enrollment at younger ages largely reflects preferences regarding the timing of school attendance rather than the level of schooling attainment. As children age, preferences of the parents are likely to play a diminishing role in schooling decisions, especially as these children approach the ages at which they will leave the household. Practical data limitations also motivate the truncation at age sixteen. Children not in residence cannot be matched with mothers and thus cannot be included in the selected samples; this attrition may be selective with regard to the outcomes of interest.

Tabulations from the full 5% census sample indicate that, of 16-year olds, slightly more than 77% live with mothers, compared to approximately 72% and 67% for 17- and 18-year olds, respectively (see Table 1). Sons are more likely to live with their mothers than are daughters, and the age gradient is much flatter for sons than for daughters. A relationship between school attendance and co-residence with one's mother is also evident in the population. At younger ages, children co-residing with mothers are more likely to be enrolled in school, while at older ages, children not living with their mothers are more likely to be enrolled in school. The change in the sign of this correlation at age sixteen for males suggests that the inclusion of children age seventeen

Table 1 Residence and schooling status

Proportion live with mom				Proportion attend school			
Age	Overall	Boys	Girls	Live with mom		Not with mom	
				Boys	Girls	Boys	Girls
10	0.908	0.910	0.905	0.953	0.954	0.907	0.913
11	0.899	0.901	0.898	0.940	0.941	0.880	0.889
12	0.884	0.887	0.881	0.887	0.876	0.815	0.802
13	0.862	0.869	0.855	0.791	0.761	0.745	0.693
14	0.844	0.854	0.833	0.687	0.651	0.651	0.586
15	0.809	0.826	0.790	0.586	0.566	0.580	0.483
16	0.772	0.802	0.741	0.518	0.497	0.564	0.448
17	0.722	0.766	0.676	0.434	0.433	0.498	0.346
18	0.670	0.739	0.603	0.375	0.367	0.449	0.248
19	0.628	0.717	0.544	0.286	0.255	0.343	0.161
20	0.530	0.648	0.428	0.163	0.147	0.189	0.073

Data drawn from 1990 Indonesian Population Census and older, living with their mothers, will overstate enrollment of girls relative to boys. Children ages 17 and older will be omitted from the sample, but observations for younger children in the same household will be retained.

The following sub-section describes the estimation strategy and the conditions under which unbiased parameter estimates may be obtained. The second sub-section discusses the validity of comparing across samples to isolate the effects of time allocation, income and bargaining power and sets some a priori bounds on the estimates, given the limitations of the data and methodology.

3.1 Estimation strategy

Four types of households are selected to examine variation in the time available for household production, the amount of income available to the household, and the distribution of bargaining power among mothers and fathers. The baseline group is represented by children with married parents both currently residing in the same household. In these households, decisions regarding children's schooling are made jointly, and both parents are contributing to the household by earning income and/or engaging in production of household public goods. The second group is comprised of children with fathers who are long-term migrants, living away from the household for the majority of the preceding year. These fathers are still actively contributing income and thus actively participating in decisions regarding children's schooling. However, the length of their absences implies that they cannot contribute to the production of household public goods, and both sons and daughters may be required to allocate more time to household activities. Comparison of this group with the baseline provides an estimate of the effect of a reduction in time available for household production on gender differences in schooling.[1]

The third group is comprised of married, dual-parent households in which the father experiences an illness that prevents him from carrying out his daily activities. This sample is analogous to the case presented above, except that fathers' earned income is also negatively affected. Comparison of the two will therefore provide an estimate of the effect of a reduction in household income on gender disparities in education, holding constant the demand for child labor in the household. Households in which the mother is currently widowed comprise the fourth group. Death of the father reduces household income and the time available for household production, analogous to the case of a debilitating illness. But death has an additional effect on household bargaining, shifting decision-making power in favor of the mother.[2] The difference in

[1] It should be noted that these distinctions are specific to the analysis of children's schooling, which cannot be easily hidden from others. Mothers may adjust some allocations to children when fathers are not present on a daily basis, but schooling decisions are more constrained as they can be easily monitored (see Chen 2009).

[2] The mother's preferences may be mediated by members of the extended family, e.g. parents and/or in-laws, after the father's death. Identification requires only that the mother's bargaining power increases after her spouse dies.

Table 2 Descriptive statistics, children co-residing with mother, ages 10 to 16

	Baseline	Migrant	Sick dad	Widow[a]
Boys				
Age	12.74	12.76	12.77	12.66
	(1.97)	(2.01)	(1.99)	(2.00)
Schooling attainment	5.80	5.58**	5.75*	5.25***
(highest grade attended/attending)	(2.26)	(2.37)	(2.28)	(2.33)
Schooling attainment	0.654	0.630***	0.647***	0.598***
(% of potential)	(0.181)	(0.197)	(0.186)	(0.207)
School enrollment	0.821	0.754***	0.795***	0.694***
	(0.384)	(0.431)	(0.403)	(0.462)
School as primary activity	0.809	0.739***	0.781***	0.680***
	(0.393)	(0.440)	(0.414)	(0.467)
Work (any income-generating activity)	0.177	0.228***	0.215***	0.279***
	(0.382)	(0.420)	(0.411)	(0.449)
Work hours (if work)	25.58	27.29	25.77	29.88***
	(15.02)	(16.39)	(15.86)	(15.12)
Number of observations	51,053	574	6,479	359
Girls				
Age	12.73	12.87*	12.79**	12.58
	(1.95)	(1.92)	(1.97)	(1.89)
Schooling attainment	5.90	5.88	5.89	5.18***
(highest grade attended/attending)	(2.26)	(2.32)	(2.25)	(2.33)
Schooling attainment	0.668	0.656	0.663**	0.601***
(% of potential)	(0.179)	(0.195)	(0.181)	(0.219)
School enrollment	0.820	0.761***	0.793***	0.655***
	(0.384)	(0.427)	(0.405)	(0.476)
School as primary activity	0.810	0.752***	0.781***	0.652***
	(0.392)	(0.432)	(0.414)	(0.477)
Work (any income-generating activity)	0.121	0.161***	0.139***	0.243***
	(0.326)	(0.368)	(0.346)	(0.429)
Work hours (if work)	24.15	27.72**	25.64**	26.60
	(15.21)	(16.83)	(15.76)	(13.44)
Number of observations	47,711	577	6,037	342

For comparability, includes only data from the 1993 SUSENAS. Standard deviations reported in parentheses

*$p = 0.10$ significant difference from column [1]; **$p = 0.05$ significant difference from column [1], ***$p = 0.01$ significant difference from column [1]

[a] Includes only households in which the maximum length of widowhood does not exceed 5 years

children's schooling outcomes between this group and the previous group will reveal the extent to which control over income matters in girls' and boys' schooling outcomes, net of changes in time allocation and income. All samples are drawn from both the 1990 Population Census and the 1993 SUSENAS, with the exception of the third group because data on temporary illness is available only in the SUSENAS.[3] Descriptive statistics are presented in Tables 2, 3, and 4.

[3] Limiting estimation to the SUSENAS, which covers all four groups, is not feasible given the rarity of widowhood and temporary migration among survey households.

Table 3 Descriptive statistics, household-level characteristics

	Baseline	Migrant	Sick dad	Widow[a]
Household members	5.82	4.36***	5.86*	5.71
	(1.73)	(1.59)	(1.78)	(1.58)
Children at home	3.47	3.02***	3.49	4.39***
	(1.54)	(1.44)	(1.57)	(1.44)
Sons/all children at home	0.525	0.509	0.526	0.532
	(0.299)	(0.325)	(0.301)	(0.247)
Males, 17–35	0.599	0.401***	0.628***	0.390***
	(0.762)	(0.682)	(0.792)	(0.678)
Females, 17–35	0.743	0.633***	0.737	0.682*
	(0.678)	(0.644)	(0.700)	(0.678)
Males, 36–54	0.720	0.014***	0.673***	0.005***
	(0.459)	(0.116)	(0.478)	(0.072)
Females, 36–54	0.587	0.660***	0.604***	0.667***
	(0.508)	(0.485)	(0.504)	(0.488)
Males, 55 and above	0.146	0.020***	0.212***	0.010***
	(0.359)	(0.141)	(0.415)	(0.101)
Females, 55 and above	0.104	0.140***	0.123***	0.074**
	(0.316)	(0.362)	(0.340)	(0.263)
Floor area of owned home	65.34	56.93***	62.65***	51.16***
	(48.01)	(41.14)	(47.03)	(30.76)
Urban residence	0.300	0.312	0.282***	0.187***
	(0.458)	(0.463)	(0.450)	(0.391)
Mothers' age	38.55	40.41***	39.55***	38.88
	(7.34)	(7.90)	(7.88)	(6.56)
Mothers' schooling attainment	5.51	5.52	4.90***	4.29***
(highest grade attended/attending)	(3.78)	(3.98)	(3.69)	(3.60)
Mother works (any income-generating activity)	0.528	0.753***	0.562***	0.831***
	(0.499)	(0.431)	(0.496)	(0.375)
Mothers' work hours (if work)	32.75	37.18***	32.83	36.59***
	(15.91)	(16.92)	(16.88)	(14.24)
Fathers' Age	44.03		45.78***	
	(8.59)		(9.68)	
Fathers' schooling attainment	6.57		5.98***	
(highest grade attended/attending)	(3.88)		(3.852)	
Father works (any income-generating activity)	0.956		0.823***	
	(0.205)		(0.382)	
Fathers' work hours (if work)	42.96		40.14***	
	(14.14)		(15.45)	
Maximum length of widowhood				3.59
				(1.37)
Fathers' length of illness			7.80	
			(7.800)	
Number of observations	59,199	738	7,600	390

For comparability, includes only data from the 1993 SUSENAS. Includes only households with at least one child between the ages of 10 and 16 in residence. Standard deviations reported in parentheses

$^{*}p = 0.10$ significant difference from column [1]; $^{**}p = 0.05$ significant difference from column [1], $^{***}p = 0.01$ significant difference from column [1]

[a] Includes only households in which the maximum length of widowhood does not exceed 5 years

Table 4 Per capita expenditure/transfers

	Baseline	Migrant	Sick dad	Widow[a]	Married, spouse away	Separated
Last month[b]						
Education expenditures (per child)	2,870 (9003)	2,488** (4540)	2,501*** (6181)	982*** (1571)		
Health expenditures	559 (5557)	558 (1733)	2,538*** (12797)	264*** (871)		
Total expenditure	37,196 (34352)	36,486 (26739)	37,746 (32571)	25,811*** (15971)		
Transfers received (if received, excl. charity and gifts)	6,738 (17390)	35,632*** (53083)	8,829 (40914)	13,585 (22668)		
Last year[b]						
Education expenditures (per child)	33,554 (97894)	29,991* (55518)	28,843*** (74027)	11,102*** (17459)		
Health expenditures	4,534 (20287)	4,575 (10639)	10,286*** (60925)	2,999*** (8387)		
Total expenditure	438,025 (338928)	439,754 (317718)	424,202*** (292071)	296,120*** (147905)		
Transfers received (if received, excl. charity and gifts)	80,489 (177365)	380,124*** (495161)	74,475 (223664)	148,777 (257204)		
Fraction w/positive transfers in last year	0.023	0.161	0.034	0.036		
Number of observations	59,198	738	7,600	390		
IFLS-1 sample						
Education expenditures (per child, monthly)	20,782 (65138)				18,571 (45400)	15,062 (16543)
Health expenditures (monthly)	1,194 (6283)				1,102 (2411)	458*** (683)
Total expenditure (monthly)	64,337 (71206)				68,165 (60039)	60,097 (47099)
Transfers received from family (annual; if received, excl. parents, sibs, kids)	42.30 (275)				113.07 (253)	4.53 (4.47)
Fraction w/positive transfers in last year	0.191				0.190	0.286
Number of observations	2,622				79	21

Includes only households with at least one child between the ages of 10 and 16 in residence. Standard deviations reported in parentheses

*$p = 0.10$ significant difference from column [1]; **$p = 0.05$ significant difference from column [1], ***$p = 0.01$ significant difference from column [1]

[a]Includes only households in which the maximum length of widowhood does not exceed 5 years

[b]For comparability, includes only data from the 1993 SUSENAS

Clearly, the selected household types are not exogenous with respect to child schooling. A household fixed effects model is employed to account for this heterogeneity. Because this strategy exploits within-family rather than inter-temporal variation, it does not assume that household type is independent of time-varying shocks to the household. Rather, the key identifying assumption is simply that, conditional on observed characteristics and the household fixed effect, status of the father—co-resident, migrant, debilitated by illness or deceased—is exogenous with respect to the boy–girl differential within the household. That is, unobserved characteristics of the household that are correlated with household type should not differentially affect sons and daughters.

Each household is assumed to maximize a utility function in which schooling of children provides a market return and/or some direct utility, subject to a time-income constraint and a [set of] household production function[s] that utilizes child labor. The reduced-form linearized demand equation for schooling of child i in household j of type t (migrant, sick, or widow) can then be expressed as

$$S_{ij} = \alpha + \alpha^t + \left(\delta + \delta^t\right) D_{ij} + \left(\beta + \beta^t\right) \mathbf{P}_{ij} + \left(\beta_g + \beta_g^t\right) \left(D_{ij} \cdot \mathbf{P}_{ij}\right) + \left(\gamma + \gamma^t\right) \mathbf{H}_j + \left(\gamma_g + \gamma_g^t\right) \left(D_{ij} \cdot \mathbf{H}_j\right) + \upsilon_{ij} \tag{1}$$

where D_{ij} is an indicator for gender which takes on a value of one for girls; $\mathbf{P}_{ij}$ denotes the vector of individual characteristics (six age-specific dummies); $\mathbf{H}_j$ denotes the vector of observed household characteristics (parents' ages and ages squared, parents' schooling, land holdings, urban residence, number and sex ratio of siblings in co-residence, household size, the number of adult household members in six age–sex specific groups, and year of survey). The error term (υ_{ij}) consists of two additive components—a household-level unobservable (μ_j) that is fixed across individuals within the household, and an i.i.d. disturbance (ε_{ij}). Parameters on household-level variables are allowed to vary by gender, denoted with subscript g, and all parameters are allowed to vary by household type. All households are pooled in the estimation to allow for tests of parameter constancy across types, and standard errors are clustered at the household level. Because the focus of this analysis is on a binary outcome variable, school enrollment, a conditional logit model (Chamberlain 1980) is estimated.

The household fixed effect captures the common endowment across household members and is, therefore, positively correlated with parents' wages.[4] Controls for full income are included, but individual-level data on wages are not available. If gender equality within the household is a normal good, the estimated differential effect on daughters will be biased downward when the probability of falling into a specific household type is negatively correlated with endowments, and vice versa (please see Table 6 for a summary of

[4]The market rates of return to children's skill units are not directly observable, thus they will also be subsumed in the household fixed effect. Provided that the gender gap in rates of return is exogenous to household type, conditional on the household fixed effect, this is not problematic. The estimated coefficient for the baseline girl dummy variable will also capture differences in the returns to human capital across males and females.

the following discussion). The type of illness is recorded in the SUSENAS, but the categories are not sufficiently detailed to permit the identification of exogenously occurring afflictions. This suggests that households with low endowments may be over-represented. Summary statistics are consistent with this; both mothers and fathers in these households are slightly older and have less schooling, relative to the baseline sample (see Table 3). Average total expenditure over the last 12 months is also significantly lower in these households, but the mean is 97% of that in baseline households (see Table 4). Even omitting health expenditures, average total expenditure in these households is still 95% of that in baseline households. Thus, the degree of selection appears to be modest.

However, ill fathers report being debilitated for just 8 days, on average, which may cast some doubt on the severity of the income shock. Self-reported measures of debilitation may also overstate the true impact of an illness, particularly if the individual can continue to work even though his daily activities have been "disrupted". Of fathers reporting a health-related disruption to their daily activities, approximately 18% report no work at all in the previous week, compared to 4% of healthy fathers. Roughly 65% of the working ill fathers report hours of work in the previous week below the mean for healthy fathers. These statistics suggest that a large majority of these households experience a reduction in the quantity of time the father can allocate to productive activities, either household or market, with many more likely experiencing some decline in the quality of work time. Health expenditures in the previous month are also roughly three times higher (see Table 4), suggesting that these illnesses are indeed quite severe.

With regard to parental death, evidence from the 1993 and 1997 Indonesian Family Life Survey (IFLS) suggest that it is largely an unanticipated event (Gertler et al. 2004a, b). Controlling for age, area of residence, household composition and mean village consumption, parents who die between 1993 and 1997 do not have significantly worse health in 1993, as measured by activities of daily living (ADLs), than those who do not. However, certain occupations and behaviors are associated with higher risk of death and may be more prevalent among lower endowment individuals or may directly affect the health endowment over time, even if they do not affect current physical functionality. Lower average schooling attainment in the sample of widows from the 1993 SUSENAS supports this hypothesis, assuming a sufficient degree of assortative mating.

The choice to remain unmarried following widowhood may also be endogenous. Data from the first wave of the IFLS indicate that widowed mothers who remarry have considerably less schooling, on average (see Table 5). Current widows have somewhat less schooling than married, never-widowed mothers, but those who have been recently widowed (5 years or less) have schooling attainment very similar to the baseline group. This suggests that, among widows, those who remarry and those remain unmarried for several years are negatively selected. Although the duration of widow status is unknown in the census and SUSENAS, the maximum length of widowhood can

Table 5 Schooling attainment of mothers, by marital status

	Married women[a]	Current widows	Recent widows[b]	Remarried widows[c]
Primary or less (%)	78.67	83.21	78.48	93.83
Junior high school (%)	11.16	8.76	12.66	2.47
High school (%)	8.25	7.3	7.59	3.7
College and graduate (%)	1.91	0.73	1.27	–
Highest grade completed	4.52	3.77	3.92	2.90
Number of observations	2,508	137	79	81

Data drawn from Indonesian Family Life Survey, 1993. Includes only women with at least one child between the ages of 10 and 16 in residence

[a] Excludes women previously widowed

[b] Includes only households in which the maximum length of widowhood does not exceed 5 years

[c] Includes only currently married women who were widowed in the previous marriage

be approximated by the age of the mother's youngest child. Based on the descriptive statistics from the IFLS, this sample is restricted to recent widows, i.e. those who cannot have been widowed for more than 5 years.

Conversely, girls will appear to be better off, relative to their male siblings, in households in which the father is a migrant, since migration only occurs when wages are higher at the destination than at the origin (Borjas 1987). Moreover, data is not collected on absent fathers in the census and SUSENAS, so key characteristics such as age and education are also unobserved for this sample. Fortunately, the IFLS collects retrospective migration histories, which allows us to examine the schooling attainment of men who have migrated for work purposes. To make this sample more comparable with the samples from the census and SUSENAS, I further limit it to the most recent migration episodes that (1) occurred while the individual had at least one child in the 10–16 age range and (2) did not include any other household members. These migrants are found to have nearly three more years of schooling (mean = 7.3, s.d. = 3.59, $n = 30$) than fathers with children age 10–16 who have never migrated (mean = 4.52, s.d. = 4.18, $n = 12450$). And, although the sample size is quite small, a test for equality of means easily rejects the hypothesis (t statistic = 4.24, p value = 0.002). This suggests that migrants are positively selected on unobserved endowments. The selected sample of migrant households will, therefore, understate (overstate) the negative (positive) effect of a reduction in the household time endowment on gender disparities in schooling.

A second challenge arises from the fact that migration is not explicitly enumerated in the data, so migrant status of absent individuals can only be inferred. For the purposes of this study, the father is identified as a temporary migrant when the mother reports that she is both married and the head of the household but her spouse is not currently in residence. Data is not collected for individuals who have been away from the household for 6 months or more at the time of enumeration, even if they are still considered members of the household; therefore, I cannot directly confirm that these fathers are, in fact, economic migrants. Remittances are crucial in this sample, as the intent is to identify households in which income and bargaining power are unaffected while fathers' non-monetary inputs are withdrawn. A much larger proportion of households in this sample receive transfers, and the average value of transfers, conditional upon receipt, is approximately five times larger than for baseline households (see Table 4). Still, only 16.1% of these households report receiving transfers in the last 12 months. In part, this may reflect the ambiguous status of temporary migrants with respect to the sending household, i.e., there may be some inconsistency in how income earned by migrants is reported when the he/she continues to be a member of the household and retains influence over allocation decisions.

Alternatively, some of these women may have been abandoned by their husbands. This would also imply a reduction in the time available for household production, but should lead to a reduction in income as well. Consistent with this, a larger proportion of mothers and children in the migrant sample engage in income-generating activities (see Tables 2 and 3). However, mean per capita household expenditure in these households is not significantly different from that in baseline households (see Table 4). In contrast, in households in which the father is deceased, per capita expenditure is much lower and an even larger proportion of mothers and children work. Widows, who are effectively abandoned, are unable to fully compensate for the income loss, even with larger increases in labor supply. Thus, despite the low level of reported remittances, households classified as having a migrant closely resemble baseline households in terms of expenditure, yet the observed labor force participation among these mothers and children does not seem sufficient to compensate for a complete loss of the father's income.

The IFLS also does not record migrant status for absent and/or departed household members, but it does allow individuals to report their marital status as "separated", whereas the census and SUSENAS do not. A comparison of these two groups can provide some additional insight into whether women who report being currently married but living separate from their spouses—the so-called "migrant" households—are still receiving support from their spouses and, if so, to what extent. As in the SUSENAS, women who report being married and head of the household but living separately from their spouses exhibit expenditure patterns very similar to those observed among households in which the parents are married and both in residence (see Table 4, bottom panel). The per capita value of transfers received from family members other than parents, siblings or children is also roughly three times greater for households in which the husband is currently absent. In contrast, female-headed households in which the woman reports her marital status as "separated" have considerably lower per capita monthly expenditure and, although a larger proportion report receiving transfers from family members other than parents, siblings or children, the value of those transfers is significantly lower. However, *t* tests for equality of means across groups are largely inconclusive due to the very small sample size.

The panel feature of the IFLS further allows us to examine whether women who report being married eventually resume co-residence with their spouses.

The IFLS sample is again limited to women with children between the ages of 10 and 16 living at home and identifying themselves as the head of the household to ensure comparability. Of the women who fall into this group in 1993 and are re-surveyed, slightly more than one-third (24 out of 73) report co-residing with their spouses in 1997, and slightly more than half (41 out of 80) report co-residing with their spouses in 2000. Of those who are still living separately from their spouses, 75% report still being married in 1997, 50% report being still married in 2000, and 35% of those who did not resume co-residence in 1997 have resumed co-residence with their spouse by 2000. These figures, combined with the data on household expenditures and transfers, suggest that the majority of women who report being married but living separately from their spouses are not experiencing de facto marital separations, nor have they been abandoned. However, to the extent that the sample of migrant households inadvertently includes such women, this will partially offset the bias due to high unobserved endowments among true migrants, assuming marital separation and abandonment are indicative of low endowments.

3.2 Isolating the effects of time allocation, income, and bargaining power

The second major challenge to the methodology of this paper lies in the fact that, even where the inclusion of household fixed effects is sufficient to generate unbiased regression coefficients, estimates of the effects of time allocation, income and bargaining power may not be unbiased. These are obtained by comparing outcomes across household types, analogous to a differences-in-differences strategy. Thus, an analogous assumption of "parallel trends" must also be satisfied. That is, the change in time allocation in households with migrant and temporarily ill fathers must be identical in order for the comparison across these two types to generate an accurate estimate of the income effect net of changes in time allocation. Similarly, the changes in both time allocation and income in households with ill and deceased fathers must be identical in order for the comparison across these two types to generate an accurate estimate of the bargaining power effect net of changes in time allocation and income. The effect of changes in time allocation is derived from a comparison of migrant households to baseline households who, by definition, have experienced no changes in the demand for household labor. However, in order to obtain an unbiased estimate of the effect of changes in time allocation on gender gaps in schooling, migrant households must not have experienced any other changes in addition to the changes in time allocation.

These assumptions are quite strong and not likely to be satisfied with natural experiments such as these. Still, relaxing these assumptions allows us to place some a priori bounds on the estimates. The sources of potential bias are described in detail below and summarized in Table 6 (for ease of reference, sub-section headings correspond to row numbers), where the stated direction of bias is based on the condition given in the far right column.

Table 6 Potential biases in estimated net effects

Source	Time allocation	Income	Bargaining power	Condition
Demand for household labor		− bias Other household members increase time caring for the father when he is sick but decrease that time when he is living away from home	+ bias Other household members increase time caring for the father when he is sick but decrease that time when he is living away from home	Increased demand for household labor has a disproportionate negative effect on girls' schooling
Labor supply		− bias Mothers are pulled into the labor market, and daughters must spend more time in household production	− bias Mothers are pulled into the labor market, and the change in income is permanent in the event of death	Reducing household income induces an additional change in time allocation that has a disproportionate negative effect on girls' schooling
Bargaining power	+ bias Women may have more influence over intra-household allocation when their spouses are away	+ bias Women may have more influence over intra-household allocation when their spouses are debilitated		Increasing women's bargaining power within the household has a positive effect on daughters relative to sons
Unobserved endowments	+ bias Migrants are self-selected and have high unobservable endowments	− bias Illness may be indicative of low unobservable endowments	− bias Death may be indicative of low unobservable endowments	High endowment households have higher income and therefore smaller gender disparities in education
Overall	+ bias	− bias Limited duration of illnesses suggests that any shifts in bargaining power are dominated by changes in time allocation due to debilitation of the father	− bias Large increase in mothers' labor supply suggests that increase in demand for household labor exceeds any reduction associated with not having to care for fathers	

Effect of time allocation is calculated as the ratio of the odds of school enrollment for girls relative to boys between households in which the father is a migrant and baseline households. Effect of income is calculated as the ratio of the odds between households in which the father is temporarily ill and those in which he is a migrant. Effect of bargaining power is calculated as the ratio of the odds between households in which the father is deceased and those in which he is temporarily ill

3.2.1 Demand for household labor

Illness, migration, and death impose similar constraints on the amount of time fathers can devote to household production. However, illness requires others to increase the amount of time spent in the production of fathers' health, whereas absence of the father, via migration or death, reduces the demand for household goods. Whether there is a net increase (decrease) in the demand for children's household labor depends on whether fathers are net producers (consumers) of household goods and on the elasticities of substitution across worker types in the household production function. Time allocation data from the IFLS indicate that approximately 32% of fathers living with children age 10–16 engaged in housework during the previous week, and those who did housework allocated roughly 11.6 hours (s.d. $= 9.66$, $n = 841$) to these tasks. While this seems to be a rather modest contribution to household production, the anthropological evidence suggests that men do engage in tasks around the home (e.g., collecting firewood and water), although they are often not perceived as "housework" (Dawson 2008, p. 50). Moreover, men often participate in childcare, which is typically viewed as distinct from "housework" (Williams 1991, p. 198) but is, nonetheless, a task that others must take over in his absence. Children with absent fathers are also more likely to report engaging in housework, particularly boys.[5] Forty-six percent of boys living with both parents report doing some housework in the last week compared to 67% of boys whose fathers are absent, and a test for equality of means rejects the hypothesis at the 10% level (t statistic $= 1.93$, p value $= 0.0625$). Thus, it appears that absence of the father does, on average, induce an increase in the demand for child household labor. No significant difference is observed among girls, but this may again be related to the implicit distinction between housework and childcare.

If increasing the demand for household labor has a disproportionately negative effect on girls, they will appear to be worse off where additional time must be devoted to the production of fathers' health and better off where fathers consume fewer (or no) home-produced goods. Consequently, the difference in girls' schooling relative to boys' between households with ill fathers and those with migrant fathers will provide a downward-biased estimate of the effect of a change in household income, and the difference in girls' schooling relative to boys' between households with deceased fathers and those with ill fathers will provide an upward-biased estimate of the effect of an increase in mothers' bargaining power.

3.2.2 Labor supply

A reduction in household income may induce additional changes in time allocation. Mothers may allocate more time to income-generating activities,

[5]Time allocation data was only recorded in the IFLS for individuals aged 15 and older. Therefore, this sample is limited to children between the ages of 15 and 20.

thereby requiring children to reduce time in school and increase time in household activities, or children may be pulled out of school and into income-generating activities directly. This is in addition to the direct effect of the reduction in monetary income, which operates as a constraint on the household's ability to purchase schooling inputs. Thus, comparing children with ill fathers to those with migrant fathers will only permit identification of an uncompensated income effect, i.e. it will not net out the changes in children's time allocation induced by the fall in income, holding constant the total time available for household production.

If children are less productive than adult women in the labor market and daughters are better substitutes for mothers in the home, girls will appear worse off in households in which there is a reduction in fathers' labor market hours and a corresponding increase (reduction) in mothers' market (household) labor hours. In this case, the estimated effect of a reduction in income on girls' schooling relative to boys' will be biased downward, relative to the true compensated income effect. Additionally, death of the father is a permanent shock to household income, whereas illness is a transitory shock. If there are fixed costs associated with increasing time allocation to income-generating activities, e.g., search costs in the labor market, a permanent income shock would be more likely to induce a secondary effect whereby mothers are pulled into the labor market and daughters must substitute for mothers in household production.

3.2.3 Bargaining power

Lastly, migration of the father may affect the mother's bargaining power. Wives may gain de facto control over intra-household allocations simply because husbands are not in close proximity. Alternatively, husbands may grant wives more control over their children's activities to compensate for their reduced contribution to household production. These factors, however, are unlikely to affect school enrollment because it is an easily monitored and thus perfectly contractible good (see Chen 2009). For similar reasons, women may also have more influence over household decisions when husbands are debilitated by illness, although this effect should be relatively small for temporary conditions. In the sample of interest, married men living with their wives and at least one school-age child, approximately 85% report being disrupted for 2 weeks or less, and 92% report being disrupted for 3 weeks or less; only about 7% of observations may be censored at 30 days. It is unlikely that illnesses lasting less than 1 month will induce sufficient changes in bargaining power to affect children's schooling. However, because respondents are asked about "health complaints" rather than illness, it is possible that disruptions are caused by chronic illnesses which are only periodically debilitating. In those cases, husbands may have lower bargaining power, and the estimated effect of income on gender disparities will be biased upward if mothers have stronger preferences for girls' schooling.

To summarize, the estimated effect of time allocation is unambiguously biased upward, given migrants' high endowments and the possibility that mothers have higher bargaining power when fathers are away. The estimated effect of income on gender differences in school enrollment is biased downward, provided that the effect of any change in bargaining power is dominated by the effect of changes in time allocation induced by illness, above and beyond the changes in time allocation induced by migration. Given that the large majority of illnesses experienced by fathers in this sample appear to be temporary, shifts in bargaining power are expected to be minimal. The estimated effect of bargaining power is biased downward provided that the reduction in demand

Table 7 Tobit regression of mothers' hours of paid work

		Migrant	Sick dad	Widow
Constant	1.703	17.34***	1.484**	26.13***
	(1.841)	(0.454)	(0.592)	(1.299)
Age[a]	1.276***			
	(0.140)			
Age squared[a]	−0.015***			
	(0.002)			
Schooling attainment	0.102***			
	(0.029)			
# of children at home	−2.186***			
	(0.144)			
Ratio of sons to all children at home	1.611***			
	(0.602)			
Floor area of home	0.000			
	(0.002)			
Urban residence	−4.655***			
	(0.295)			
Household size	3.179***			
	(0.312)			
Max. length of widowhood[b]				0.487
				(0.424)
Length of illness			0.136***	
			(0.052)	
$E(Y \mid Y > 0)$[c]	35.84***	46.56***	37.24***	54.11***
	(1.912)	(1.907)	(1.947)	(1.948)
$Pr(Y > 0)$[c]	0.756	0.891	0.780	0.939
Number of observations				

Includes kabupaten-level fixed effects, controls for year of survey, and number of household members in 11 age–sex categories with females age 36–54 as the omitted category. Standard errors in parentheses

$^{*}p = 0.15$; $^{**}p = 0.10$; $^{***}p = 0.05$

[a] Standardized to (age – 20)

[b] Widow sample includes only households in which the maximum length of widowhood does not exceed 5 years

[c] Calculated at sample mean for baseline households in the school attendance conditional logit estimation sample

for children's time in the production of father's health is dominated by the increase in demand for children's time due to the mother's increased labor market hours. Estimates in Table 7 indicate that, on average, widowed mothers work 16–17 hours more per week than women with ill husbands (see following section for additional discussion), which is likely larger than the amount of time children would devote to the care of temporarily ill fathers.

4 Results

To confirm that the selected samples indeed experience shocks to the time–budget constraint as described above, I first report results from a Tobit regression of mothers' time in income-generating activities, followed by the main results on children's school enrollment.

4.1 Mothers' hours of paid work

Controlling for observable characteristics of the household and individual, migration of the husband is associated with approximately ten more hours of work per week, as compared to mothers in the baseline sample (see Table 7; coefficients in columns 2 through 4 are relative to the baseline (un-interacted) coefficients in column 1). Note, however, that these estimates should be interpreted only as adjusted means and not as causal relationships because the endogeneity of household status has not been accounted for. In fact, data from the IFLS indicate that mothers are more likely to be engaged in income-generating activities even prior to the husband's migration; that is, among married women living with their spouses in 1993, those who will be living apart from their spouses in 1997 also report more hours of work for pay, work on the family farm and/or work for the family business in 1993 (49 versus 38 h, t statistic for test of equality of means = 2.0748, p value = 0.0481).

Temporary debilitation of the husband is also associated with significantly greater work hours for women, but the magnitude of this difference is quite small. Length of the husband's illness has a larger marginal effect, suggesting that the more severe the shock to income, the more mothers are pulled into wage labor. A limited market labor supply response is consistent with the counterbalancing effects of illness on time allocation of other household members. There is both an increase in the demand for household labor, as the father increases time spent in the production of own health and reduces time spent in the production of other household goods, and an increase in the demand for market labor, as the father reduces time spent in wage labor and household income falls. Finally, death of the husband is associated with significantly greater labor supply, with widowed mothers working roughly 18 hours more per week. Although the demand for household labor also increases when the father dies, a permanent income shock appears to induce a much larger market labor supply response from mothers.

4.2 Current school enrollment

The probability of being enrolled in school at the time of enumeration is estimated with a conditional logit to account for household fixed effects. Results are summarized in Table 8. All specifications include the child's age and household-level covariates (wealth, household size, demographic composition, urban residence, mother's characteristics, and father's characteristics if present), all interacted with the child's sex. Specification I is the most parsimonious and does not allow parameters on household characteristics to vary with household type. Specification II relaxes this assumption, and a test of joint significance rejects the null that the effect of household characteristics is constant across gender and household type ($\chi^2(45) = 82.20$, p value =

Table 8 Children's school enrollment, conditional logit estimates

	I	II	III	IV
δ	−0.167	0.298	−0.160	0.250
	(0.431)	(0.458)	(0.161)	(0.440)
δ^M	−0.062	−2.768**	0.241	−2.632**
	(0.365)	(1.396)	(0.450)	(1.306)
δ^S	0.262	−6.388***	−0.757	−7.792***
	(0.439)	(2.133)	(0.658)	(2.321)
δ^W	−0.169	−2.769	−0.001	−2.649
	(0.499)	(2.093)	(0.529)	(2.062)
Net effect on odds of enrollment for girls relative to boys				
Time allocation[a]	0.820	0.693	1.106	0.714
	(1.08)[b]	(0.60)	(0.35)	(0.57)
Income	1.076	1.753	0.920	1.987
	(0.31)	(0.85)	(0.27)	(1.10)
Bargaining power	0.983	6.591**	0.913	5.658**
	−(0.07)	(2.37)	(0.18)	(2.19)
Odds of enrollment for girls relative to boys, relative to baseline				
Sick dad	0.883	1.216	1.018	1.419
	(0.759)[b]	(0.767)	(0.178)	(1.548)
Widow	0.868	8.015***	0.929	8.028***
	(0.690)	(2.751)	(0.145)	(2.737)
Household fixed effects	Yes	Yes	No	Yes
Other parameters vary by household type	No	Yes	Yes	Yes
Alternative measure of attendance	No	No	No	Yes

Includes controls for age, number and sex ratio of siblings at home, floor area of owned home, parents' ages and years of schooling, household size and composition, and year of survey. Net effect on odds of school enrollment for girls relative to boys calculated at the sample mean. Standard errors clustered at the household level and reported in parentheses

$^{*}p = 0.10$; $^{**}p = 0.05$; $^{***}p = 0.01$

[a] Net effect of time allocation is equivalent to the odds of enrollment for girls relative to boys in migrant households, relative to baseline households

[b] Values in parentheses are robust t statistics for test that value is equal to one

0.001). For comparison, specification III excludes household fixed effects, but a Hausman test indicates that the estimates without fixed effects are inconsistent ($\chi^2(92) = 1167.27$, p value $= 0.000$). Finally, specification IV considers an alternative measure of schooling which better captures actual attendance in the previous week. However, this measure is based on the child's reported *primary* activity, so children who work, in or outside the home, in addition to attending to school are assigned a value of zero if they allocated more time to work than to school in the previous week. Estimated parameters are therefore difficult to interpret, as decisions regarding school attendance and time spent in productive activities are confounded in the outcome measure. Furthermore, illness is reported for the previous month and thus may not coincide with the child's reported activities in the previous 1 week.

The full set of parameter estimates for Specification II is reported in the Appendix, Table 10. The probability of being enrolled in school falls rapidly and steadily as children age. However, age effects are more heterogeneous when the household is faced with additional time and/or income constraints. Children are less likely to be enrolled in school when the father migrates, but the magnitude of this effect is relatively small and only significant for ages 11 and 12. This is consistent with an increased demand for household labor, with both boys and girls bearing some of the burden of the father's absence. The age gradient for school enrollment is also steeper when the father is temporarily debilitated, and a similar but much more pronounced pattern is observed when the father is deceased. These estimates are consistent with a reduction in the time available for household production, as well as a reduction in household income. The change in income is much larger when the father is deceased, leading to much larger and more precisely estimated negative coefficients. Furthermore, the age effects suggest that the increase in mothers' bargaining power associated with the death of the father is not sufficient to offset the income effect, at least with regard to levels.

Baseline parameters indicate that observable household characteristics do not differentially affect daughters, with the exception of mothers' years of schooling, which has a positive effect. The coefficient on the dummy variable for girls is positive and insignificant, but daughters with migrant fathers are significantly less likely to be enrolled in school, which suggests that reducing the time available for household production has a differentially negative effect on girls. This effect is mediated, but not entirely offset, by the age of the mother, with prime-aged mothers ($\sim$35 to 50) having the largest positive effect on daughters' school enrollment. Mother's age effects are evident for all three comparison groups but statistically significant only for the sample of children with migrant fathers, which perhaps suggests that more experienced mothers are better able to protect daughters from shocks to household production than shocks to household income.

Temporary debilitation of the father has a very large and statistically significant negative effect on daughters' school enrollment, which is exacerbated by the length of the father's illness and the number of siblings in the household, both of which increase demand for girls' time in household production.

However, additional non-sibling household members reduce the burden on daughters, either by taking on some of the father's household production tasks or by providing an additional channel for consumption smoothing. Older and more educated fathers are also better able to smooth consumption, and perhaps income as well, again mitigating the negative effect of illness on daughters. Death of the father does not have a statistically significant direct effect on daughters' school enrollment. Although the point estimate is relatively large, it is considerably smaller than that for girls' with ill fathers and more similar in magnitude to the estimate for households with migrant fathers. Thus, while death of the father has a large negative effect on the probability of school enrollment for all children, the increase in mothers' bargaining power allows them to offset the negative income and time allocation effects that otherwise would be borne disproportionately by daughters. Consistent with this, siblings have a significant and positive effect on daughters' school enrollment, rather than the negative effect observed in households in which the father is debilitated. It appears that mothers, when granted more influence over household decision-making, tend to allocate relatively more resources to daughters.

4.3 Calculation of net effects

Comparing across the four household types generates estimates of the net effects of time allocation, income, and bargaining power on gender differences in human capital investment. Since there are both direct effects and effects operating through other household characteristics, this comparison is best done by calculating predicted probabilities for school enrollment. However, calculating predicted probabilities requires selecting an appropriate value of μ_j, and these parameters are not directly estimated with conditional logit, nor are any restrictions placed on the distribution of μ_j. Instead, I calculate the odds ratio of school enrollment for girls relative to boys, which eliminates the need to assign a value for μ_j. Taking the ratio of these odds ratios will then permit a decomposition of the sources of discrimination, as follows, and the independent (net) effects of changes in time allocation, income and bargaining can be identified.

$$\mathrm{OR}^{\mathrm{M}} \big/ \mathrm{OR} = \text{net effect of changes in time allocation}$$

$$\mathrm{OR}^{\mathrm{S}} \big/ \mathrm{OR}^{\mathrm{M}} = \text{net effect of changes in income}$$

$$\mathrm{OR}^{\mathrm{W}} \big/ \mathrm{OR}^{\mathrm{S}} = \text{net effect of changes in bargaining power,}$$

where OR denotes the odds ratio of school enrollment for girls relative to boys, M refers to the sample of children with migrant fathers, S refers to the sample of children with sick fathers, and W refers to the sample of children with widowed mothers.

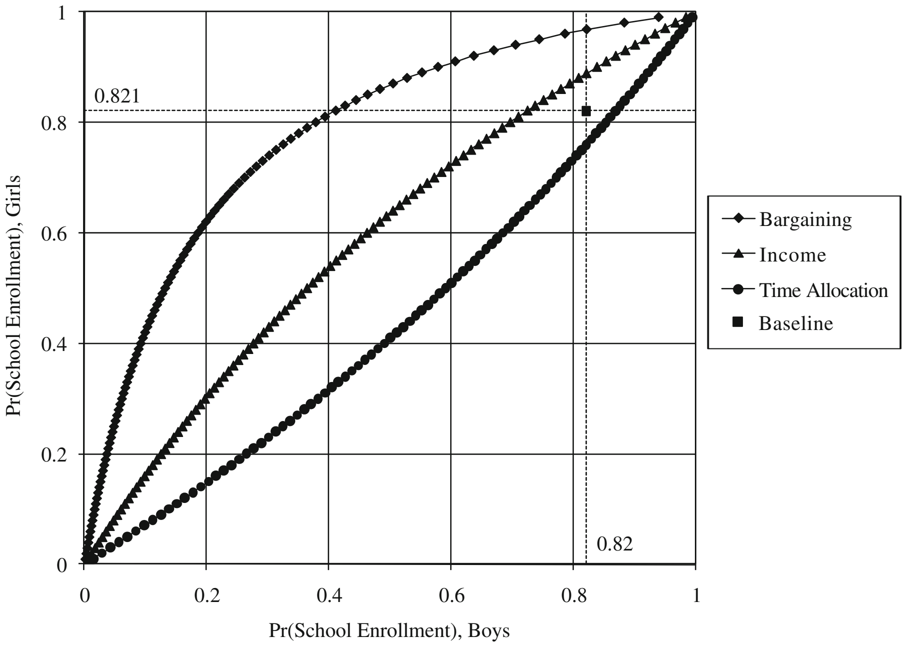

Fig. 1 Net effect on odds of school enrollment

These values are calculated at the sample mean for baseline households and presented in the center portion of Table 8. Values less than one indicate that daughters are less likely to be enrolled in school than sons, and vice versa for values greater than one. Probabilities consistent with the predicted odds ratios are also presented in Fig. 1. Under the preferred estimates, specification II, reducing the time available for household production has a disproportionately negative effect on daughters. If we assume that the probability of enrollment remains constant for boys, this estimate implies that the probability of enrollment for girls would fall from 82% to 76%. Reducing household income has a modest positive effect on daughters relative to sons, whereas increasing mothers' bargaining power has a very large positive effect, the largest of the three, in terms of magnitude. Again assuming that the probability of enrollment remains constant for boys, a fall in household income, net of changes in time allocation, increases girls' probability of enrollment to 89%. In comparison, an increase in bargaining power, net of changes in time allocation and household income, increases girls' enrollment to 97%.

Unfortunately, point estimates are imprecise with this type of decomposition because the standard errors are based on two sets of parameter estimates, one for each household type utilized in the comparison. However, the discussion of potential sources of bias in Section 3.3 (see also Table 6), allows us to bracket the point estimates and, in each case, provide an accurate characterization of the sign of the effect if not the magnitude. The biases implicit in the estimated time allocation effect are unambiguously positive, and the point estimate is negative. Thus, reducing the time available for household production clearly has a disproportionate negative effect on daughters; girls are more likely to be withdrawn from school in order to compensate for the increased need for household labor. And, in fact, the estimate obtained here understates the magnitude of the true effect. The biases implicit in the estimated income effect are negative on net, provided that any shifts in bargaining power are dominated by changes in time allocation, whereas the point estimate is positive. Thus, the point estimate, although imprecise, again provides an accurate characterization of the sign of the true effect, although understating the magnitude. A reduction in income unambiguously increases the probability that girls are enrolled in school, relative to their male siblings, which suggests that girls' schooling is less income-elastic, at least with regard to transitory income. Finally, the biases implicit in the estimated bargaining power effect are negative on net, as long as death of the father leads to a net increase in the demand for household labor. The point estimate is positive, statistically significant and quite large in magnitude, which suggests that shifting bargaining power to mothers will, in fact, have an even larger effect on reducing gender disparities in education than what is indicated by the point estimate presented here.

5 Robustness and discussion

The above decomposition also allows us to evaluate the bias introduced by neglecting the role of concurrent changes in time allocation and/or income. The bottom portion of Table 8 presents the odds of school enrollment for girls relative to their male siblings, relative to the baseline. The net effect of time allocation is, by construction, identical to the odds of school enrollment for girls relative to boys in migrant households, relative to baseline households. Based on the preferred specification, column II, temporary debilitation of the father has, at most, a slight positive effect on the probability of school enrollment for girls relative to boys. However, a simple comparison with baseline households confounds changes in income and time allocation and, in this case, leads to substantial downward bias in the estimated effect of household income on gender differences in human capital investment. Death of the father has a large and significant positive differential effect on daughters' school enrollment. But, once concurrent changes in time allocation and income have been taken into account, it is clear that the net effect of shifting bargaining power to the mother has a much smaller, though still positive and significant, effect on daughters. From a policy perspective, this exercise reiterates how important it is to consider the ways in which a program may affect all aspects of household decision-making, particularly changes in time allocation which are often overlooked.

5.1 Schooling attainment

To help validate the assumptions and results presented above, I also examine the gender gap in schooling attainment, defined as the highest grade completed. This is a longer term measure of human capital investment and thus should be largely unaffected by short-term changes in income and the time available for household production, as long as illness and migration are in fact exogenous to gender preferences, conditional on the household fixed effect. Widowhood, in contrast, does induce permanent changes in income, time allocation and bargaining power which may, in turn, affect schooling attainment.

Estimates of the effect of household type on schooling attainment are presented in Table 9. Again, all specifications include the child's age and household-level covariates interacted with the child's sex. Specification I is the most parsimonious and does not allow parameters on household characteristics to vary with household type. Specification II relaxes this assumption, but a test of joint significance does not reject the null that the effect of household characteristics is constant across gender and household type ($F(45{,}145581) = 0.41$, p value $= 0.9999$). For comparison, specification III excludes household fixed effects, but a Hausman test indicates that the estimates without fixed effects are inconsistent ($\chi^2(52) = 69.31$, p value $= 0.0545$). The full set of parameter estimates for the preferred specification (I) is reported in the Appendix,

Table 9 Children's schooling attainment, household fixed effects estimates

	I	II	III
δ	0.233	0.194	0.108
	(0.198)	(0.208)	(0.078)
δ^M	0.071	0.321	−0.016
	(0.178)	(0.566)	(0.074)
δ^S	0.003	0.162	−0.039
	(0.200)	(0.759)	(0.084)
δ^W	0.043	0.141	0.009
	(0.246)	(0.819)	(0.112)
Net effect on attainment for girls relative to boys			
Time allocation	0.071	−0.223	−0.038
	(0.178)	(0.320)	(0.160)
Income	−0.078	0.270	−0.027
	(0.098)	(0.336)	(0.167)
Bargaining power	−0.038	0.517	−0.079
	(0.109)	(0.696)	(0.054)
Household fixed effects	Yes	Yes	No
Parameters vary by household type	No	Yes	No

Includes controls for age, number and sex ratio of siblings at home, floor area of owned home, parents' ages and years of schooling, household size and composition, and year of survey. Net effect on attainment for girls relative to boys calculated at the sample mean. Standard errors clustered at the household level and reported in parentheses

$^{*}p = 0.10$; $^{**}p = 0.05$; $^{***}p = 0.01$

Table 11. Coefficients on the indicator variables for girls by household type are not statistically significant and, with the exception of baseline parameters, the point estimates are quite small in magnitude. These estimates reveal no significant gender differences in schooling attainment across household types, lending empirical support to the main identifying assumption—household type is exogenous to gender differences in human capital investment once household fixed effects have been taken into account. Net effects, calculated as for school enrollment, are also small in magnitude and not statistically significant.

The absence of the father from the household may also differentially affect the production function for sons' and daughters' human capital. Fathers' time may be a more productive input in the production of boys' human capital, or there may be stronger complementarities between sons' and fathers' time in this production process. Migrants in the sample are away from the home for six months or more out of the year. Thus, if boys' human capital is particularly sensitive to fathers' time inputs, we would expect to see a slight deterioration, on average, in schooling attainment among boys in this sample, even given changes in school enrollment that slightly favor boys. Again, the estimates reveal no significant gender differences, although there is some indication of differences in the level of schooling attainment for all children in migrant and widow households.

6 Conclusion

This paper seeks to disentangle the various factors contributing to gender disparities in children's human capital. Measurement of the independent contribution of each of these factors is essential if gender disparities are to be effectively targeted. The methodology provides a careful accounting of the role of preferences, income and time allocation. Increasing mothers' influence over household decision-making has the largest impact on closing gender gaps in school enrollment. Conversely, reducing the time available for household production increases the enrollment gap between daughters and sons. But, surprisingly, a transitory reduction in household income increases the probability that girls will be enrolled in school, relative to their male siblings; girls' schooling appears to be better insulated from temporary income shocks. Moreover, this finding highlights the fact that failing to account for changes in time allocation associated with an income shock can lead to a sizable bias in estimated income effects. Similarly, failing to account for changes in both income and time allocation that are associated with a change in women's bargaining power can also result in a significant bias, even after controlling for unobservable characteristics of the household that affect all children identically.

The decomposition presented in this paper provides insight into how gender disparities arise and whether they can be best ameliorated by addressing poverty, the demand for household labor, or the preferences of parents. Based on the net effects of time allocation, income and bargaining power calculated above, programs aimed at bolstering women's position in the household would have the largest effect on reducing gender disparities in human capital investment. However, policymakers must also be aware that concurrent changes in time allocation and/or income induced by such a program may work in the opposite direction unless some conditionality is imposed.

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Appendix

Table 10 Conditional logit regression of children's school enrollment

		Relative effects		
		Migrant	Sick dad	Widow[c]
Age 11	−0.220*	−0.730**	−0.468	−0.845***
	(0.114)	(0.318)	(0.360)	(0.327)
Age 12	−1.185***	−0.468*	−0.210	−0.702**
	(0.093)	(0.262)	(0.305)	(0.277)
Age 13	−2.474***	−0.348	−0.275	−0.585*
	(0.097)	(0.281)	(0.322)	(0.303)
Age 14	−3.455***	−0.415	−0.678*	−0.873***
	(0.098)	(0.300)	(0.349)	(0.334)
Age 15	−4.225***	−0.141	−0.549	−0.700**
	(0.104)	(0.302)	(0.356)	(0.341)
Age 16	−5.198***	−0.307	−0.741*	−0.600
	(0.114)	(0.346)	(0.401)	(0.394)
Relative effects for daughters				
Constant (δ)	0.298	−2.768**	−6.388***	−2.769
	(0.458)	(1.396)	(2.133)	(2.093)
Number of siblings	0.055	0.164	−0.562**	0.416*
	(0.063)	(0.177)	(0.246)	(0.246)
Male siblings/all siblings	0.169	0.819	−0.213	0.408
	(0.235)	(0.741)	(1.046)	(0.924)
Floor area of owned home	0.001	0.003	0.001	−0.003
	(0.001)	(0.003)	(0.004)	(0.004)
Urban residence	0.066	0.332	0.617	− 0.019
	(0.105)	(0.385)	(0.392)	(0.487)
Household size	−0.074	−0.192	0.563**	−0.270
	(0.062)	(0.181)	(0.235)	(0.278)
Mother's age[a]	−0.032	0.228**	0.055	0.282
	(0.036)	(0.106)	(0.173)	(0.178)
Mother's age squared[a]	0.001	−0.005**	0.000	−0.006*
	(0.001)	(0.002)	(0.003)	(0.004)

Table 10 (continued)

		Relative effects		
		Migrant	Sick dad	Widow[c]
Mother's schooling attainment	0.042***	0.051	−0.064	0.075
	(0.013)	(0.040)	(0.056)	(0.050)
Father's age[a, b]	0.029		0.364**	
	(0.027)		(0.149)	
Father's age squared[a, b]	0.000		−0.007**	
	(0.000)		(.003)	
Father's schooling attainment[b]	−0.014		0.123**	
	(0.014)		(0.055)	
Length of illness			−0.039	
			(0.021)	
Maximum length of widowhood				0.049
				(0.110)
Number of observations	44,227	3,491	2,695	3,062
Wald test that parameters jointly = 0 (p value)		31.45*	39.10**	44.12***
		(0.067)	(0.036)	(0.003)

Includes interactions between age and gender and controls for year of survey and number of adult household members in six age–sex categories. Standard errors clustered at the household level and reported in parentheses

$*p = 0.10$; $**p = 0.05$; $***p = 0.01$

[a]Standardized to age – 20

[b]Parameters estimated only for households in which the father is present, and those reported in column three are direct, not relative, effects

[c]Includes only households in which the maximum length of widowhood does not exceed 5 years

Table 11 Fixed effects regression of children's schooling attainment

		Relative effects		
		Migrant	Sick dad	Widow[c]
Age 11	0.860***	−0.129	−0.039	−0.003
	(0.032)	(0.095)	(0.094)	(0.118)
Age 12	1.822***	0.001	−0.052	−0.047
	(0.028)	(0.080)	(0.076)	(0.099)
Age 13	2.669***	−0.072	0.057	−0.026
	(0.031)	(0.089)	(0.090)	(0.123)
Age 14	3.508***	−0.077	0.084	−0.262**
	(0.032)	(0.095)	(0.097)	(0.123)
Age 15	4.157***	−0.190*	−0.037	−0.442***
	(0.034)		(0.108)	(0.130)
Age 16	4.772***	−0.179	−0.021	−0.640***
	(0.038)	(0.111)	(0.112)	(0.148)
Relative effects for daughters				
Constant (δ)	0.233	0.071	0.003	0.043
	(0.198)	(0.178)	(0.200)	(0.246)
Number of siblings	0.007			
	(0.021)			
Male siblings/all siblings	0.089			
	(0.090)			
Floor area of owned home	0.000			
	(0.000)			

Table 11 (continued)

		Relative effects		
		Migrant	Sick dad	Widow[c]
Urban residence	0.057**			
	(0.029)			
Household size	−0.009			
	(0.022)			
Mother's age[a]	−0.024*			
	(0.014)			
Mother's age squared[a]	0.000			
	(0.000)			
Mother's schooling attainment	0.009*			
	(0.005)			
Father's age[a, b]	−0.002			
	(0.014)			
Father's age squared[a, b]	0.000			
	(0.000)			
Father's schooling attainment[b]	−0.001			
	(0.005)			
Length of illness			−0.001	
			(0.009)	
Maximum length of widowhood				−0.026
				(0.049)
Number of observations	201,999	17,591	12,516	9,247

Includes interactions between age and gender and controls for year of survey and number of adult household members in six age–sex categories. Standard errors clustered at the household level and reported in parentheses

$*p = 0.10$; $**p = 0.05$; $***p = 0.01$

[a]Standardized to age – 20

[b]Parameters estimated only for households in which the father is present, and those reported in column three are direct, not relative, effects

[c]Includes only households in which the maximum length of widowhood does not exceed 5 years

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