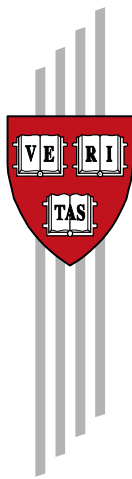


# **Dads, Disease and Death: Decomposing Daughter Discrimination**

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# Dads, Disease and Death: Decomposing Daughter Discrimination

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## *Abstract*

Existing literature suggests that girls are differentially affected by (1) income shocks and (2) changes in bargaining power. However, these analyses do not shed light on the actual sources of discrimination, *i.e.* whether differential treatment is the result of maximization given differential returns and opportunity costs or a more deeply entrenched notion of gender bias. In particular, the majority of studies neglect the role of household production and children's time allocation in gender discrimination. Using data from the 1990 Indonesian Population Census and 1993 Indonesian Socioeconomic Survey, this paper seeks to identify and quantify the potential sources of discrimination, namely preferences, income and time allocation. To identify the distinct sources of discrimination, I utilize three household types, each highlighting different levels of parental involvement and exogenous shocks to income, as well as their interaction. A household fixed effects model is employed to control for unobservable characteristics which may be correlated with household type. Results indicate that a reduction in the amount of time available for household production, in most cases, reduces the probability that daughters will be enrolled in school, relative to their brothers, but a reduction in household income alone may have a positive effect on girls' school enrollment relative to their male siblings. Increasing mothers' bargaining power has the largest impact on closing the gender gap.

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## **I. Introduction**

Gender gaps in schooling, nutrition, health and even survival remain persistent. Girls constitute nearly two-thirds of the children excluded from a basic education (UNESCO, 2000), and more than 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 sources 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 - Progreso in Mexico and FSSAP (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 Amartya 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 economic development more generally.

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' time in the household will be much more productive and a better substitute for mothers' time than will sons' time (see Pitt and Rosenzweig, 1990). 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. The latter suggests that the demand for time in the household 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. Furthermore, even if parity in schooling is achieved, shocks to the household may still have a disproportionate effect on girls.

The existing literature on intrahousehold allocation and gender neglects the role of household production and children's time allocation and thus fails to answer the question of how gender discrimination is independently affected by factors such as income and bargaining power. Theory suggests that income, time allocation and preferences all affect the gender gap in human capital investments, but the current literature cannot shed light on the relative magnitude and independent contribution of each of these factors. Utilizing cross-sectional surveys from

Indonesia, this paper seeks to quantify these potential sources of discrimination, namely preferences, income and time allocation, by examining sons and daughters' schooling outcomes under varied conditions. To identify the distinct sources of discrimination, I utilize three household types, each highlighting different levels of parental involvement and exogenous shocks to income, as well as their interaction: (1) fathers as long-term migrants (reduction in the time available for household production); (2) fathers with a temporary debilitating illness (reduction in time and income); and (3) widowed mothers (reduction in time, income and a shift in bargaining power). Comparison of boys' and girls' schooling outcomes across these groups will allow us to separate and identify the effects of time allocation, income and control over income. The analysis will also provide definitive evidence of the role of preferences by utilizing a forced change in bargaining, rather than one which is inferred from changes in the sources or amounts of income accruing to men versus women.

Controlling for observable characteristics and unobservable household fixed effects, I find that 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 reduction in household income alone may have a positive effect on girls' school enrollment relative to their male siblings. Increasing mothers' bargaining power has the largest impact on closing the gender gap, but, as in the case of widowed mothers, this effect can be entirely offset by concurrent changes in income and time allocation. These findings suggest that policies aimed at empowering women will be most effective when constraints on women and children's time allocation are taken into account.

## **II. Children's Schooling and Parental Inputs**

Investments in schooling are unique in that they may provide some consumption value beyond the market return. For younger children, these investment decisions are largely made by

parents, or at least made possible by parents. Parental inputs to children's schooling are multi-dimensional and interlinked. Higher income lessens the burden of schooling costs in the sense that education expenditures are less likely to compete directly with goods necessary for subsistence, and schooling is typically believed to be a normal good. With diminishing marginal returns to schooling investments per child, higher income will also lead to more equality within the household.

A related body of literature suggests that shocks to household income differentially affect daughters. 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. Extending upon this, Rose (1999) finds that favorable rainfall shocks increase girls' probability of survival relative to boys, while girls suffer disproportionately from negative shocks.

These findings suggest that equality within the household is a normal good, increasing in income. But differences in rates of return also affect gender-specific investments in health and nutrition; when resource constraints are binding, parents can invest only in the children with the highest expected returns. It is difficult to rule out the possibility that it is changes in time allocation induced by these shocks, not the changes in income per se, which lead to discriminatory behavior. Rising food prices and adverse rainfall shocks likely increase time spent in income-generating activities. In rural areas, these tend to be labor-intensive activities with high returns to nutrition and higher returns to males than females; when men allocate more time to such activities, their nutritional requirements also rise relative to female household members. Increased time in income-generating activities also reduces the time available for

household production which may differentially affect girls' survival probabilities, particularly if girls are more likely to be malnourished even prior to the shock.

Gertler et. al. (2004a, 2004b) investigate the effect of parental death on children's school enrollment and find that it has a large and significant adverse effect. Children in high asset households do not experience a smaller reduction in school enrollment following parental death (2004a), 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 (2004b). 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 which 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. This paper aims to disentangle these three factors and determine their signs and relative magnitudes.

The possible consumption value associated with investments in children's schooling suggests that preferences 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 for women should lead to higher investments in children. Girls in particular will benefit if mothers also have stronger preferences for gender equality, especially if there is some gender discrimination ex ante. Duflo (1999) finds that the extension of the old age pension in South Africa led to an improvement in the health and nutrition of girls and no discernible effect on boys, with the effect being entirely due to pensions received by women (grandmothers). Men's pension income had no effect on children's health and nutrition for either girls or boys. Similarly, Thomas (1990) finds that both parents' unearned income is positively correlated with

household per capita calorie consumption and protein intake, but the effect of maternal income is four to seven times larger.

With regard to schooling, Pitt and Khandker (1998) find that credit provided to women increases the probability of school enrollment for both boys and girls, while participation by men has a positive effect on boys only. The common notion is that empowerment of women via the provision of credit has positive effects on children and daughters in particular. But, participation in credit programs by men may induce different changes in time allocation than participation by women. The investments and micro-enterprises chosen by men may affect the returns to schooling for boys or increase the demand for girls' time in household activities, whereas women's investments may be more complementary to household production and thus free up more of children's time for schooling. Women's participation in microcredit programs also has a larger effect on sons' schooling than 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. In this case, the (positive) pure income effect is offset by the change in time allocation induced by the change in this particular source of income. The relationship between household income and the opportunity cost of children's schooling is ambiguous, and the non-separable nature of household production implies that the roles of income and time allocation cannot be disentangled in these studies.

Clearly, the nature of the shock experienced by the household also matters. All households produce some goods in the home, and these goods invariably require time inputs from household members. Children may be valuable in the production of household goods, *e.g.* caring for younger siblings or assisting in household chores, and this, in turn, may free up parents' time for more productive opportunities. Health shocks affect the production function for health and thus the time inputs required of other household members. 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 relative to both wage labor and school attendance, 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 can be differentially affected by a change in the demand for household time.

In summary, the existing literature provides some 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. To examine and disentangle these effects, this paper will examine boys' and girls' respective schooling outcomes under varied household structures: (1) father is present and contributing income; (2) father is not present but is actively earning and contributing income; (3) father is present but temporarily debilitated; and (4) father is deceased. The first is the baseline case; the second captures solely the effect of the change in time allocation induced by the father's absence, with no effect on income or bargaining power; the third represents a shock which affects time allocation and income but does not affect bargaining power; while the fourth represents a shock that affects time allocation, income and bargaining power.<sup>1,2</sup>

### **III. Theory and Methodology**

To organize the discussion of methodology, I will outline a simple model that incorporates various aspects of the existing literature on schooling. Define a representative

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<sup>1</sup> Households in which the father is deceased but the family continues to collect some of his asset or insurance income are included in the fourth group, because expenditure decisions are unlikely to be influenced by the father's preferences after his death.

<sup>2</sup> 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.

household composed of four types of individuals, with one member of each type: two adult decision makers, male ( $m$ ) and female ( $f$ ), and two children, a boy ( $b$ ) and a girl ( $g$ ). To abstract from issues of intertemporal substitution, consider the lifetime utility function for this household, which can be specified as a weighted sum of each adult decision maker's individual utility function

$$U = \lambda_m [U_m(\mathbf{x}_m, \mathbf{z}, h_m, h_f, s_b, s_g, e_b, e_g)] + \lambda_f [U_f(\mathbf{x}_f, \mathbf{z}, h_m, h_f, s_b, s_g, e_b, e_g)] \quad \lambda_m + \lambda_f = 1$$

where  $\mathbf{x}_k$  denotes a vector of market goods specific to the individual;  $\mathbf{z}$  denotes a vector of shared household goods;  $h_m$  and  $h_f$  denote health of the respective adult decision-makers;  $s_b$  and  $s_g$  denote human capital of sons and daughters<sup>3</sup>; and  $e_b, e_g$  denote future earnings of children. Allocations to children are permitted to affect the overall household utility function differentially by gender. In other words, the marginal utility of investment in daughters' human capital may be different from that of investment in sons' human capital, and this too may vary according to the identity (gender) of the decision maker.<sup>4</sup>

The parameters  $\lambda_m$  and  $\lambda_f$  indicate the bargaining power of each respective parent and are determined through negotiation in which relative threat points and power within the household depend on the individual's outside option, represented by the vector  $\boldsymbol{\omega}_k$ .

$$\lambda_k = \lambda(\boldsymbol{\omega}_k) \quad \text{for } k = m, f$$

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<sup>3</sup> This model could easily be extended to include a vector of allocations to children, including health.

<sup>4</sup> It is also possible that individual utility functions are interdependent, whereby other household members' utility functions would enter into the decision maker's utility function, however this specification is not essential to the following analysis.

If preferences are identical, the utility function reduces to a unitary model of the household. One component of the individual's outside option is potential or expected future income.<sup>5</sup> Thus a transitory reduction in income, as in the case of temporary debilitating illness, will have a smaller effect on the  $\lambda$ -weights than a permanent change in earnings potential, as in the case of death. For a purely transitory shock to income, the  $\lambda$ -weights should be unaffected.

The goods  $z$ ,  $h_m$ ,  $h_f$ ,  $s_b$  and  $s_g$  are produced by the household and require time inputs from household members as well as market goods.

$$z = z(t_{qm}, t_{qf}, t_{qb}, t_{qg}, x_z)$$

$$h_k = h(t_{hk,m}, t_{hk,f}, t_{hk,b}, t_{hk,g}, x_{hk}, \gamma_k, \upsilon_k) \quad \text{for } t_{hk,k} > 0 \text{ and } k = m, f$$

$$s_i = s(t_{si,m}, t_{si,f}, t_{si,b}, t_{si,g}, x_{si}, \gamma_i, \upsilon_i) \quad \text{for } t_{si,i} > 0 \text{ and } i = b, g$$

The production functions for health and human capital depend on the individual's characteristics ( $\gamma_k$  or  $\gamma_i$ ) and good-specific endowments ( $\upsilon_k$  or  $\upsilon_i$ ), where each individual endowment is a composite of the household-level endowment ( $\upsilon$ ) and person-specific heterogeneity.<sup>6</sup> With regard to both health and schooling, the elasticity of substitution between time inputs of different types of household members is assumed to be greater than the elasticity of substitution between own and others' time. I do not rule out the possibility that other household members' time contributions are perfect substitutes for each other in the production of either health or human capital.

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<sup>5</sup> Another component of the individual's outside option is the number of relatives within close proximity. These relatives can provide transfers and/or insurance to lessen the cost of departing the household, should the individual choose to exercise his/her outside option. The provision of transfers may also allow relatives some leverage in directly influencing allocations. Without loss of generality, I assume that relatives in household  $l$  providing transfers to member  $k$  in household  $j$  have preferences identical to individual  $k$ , at least with regard to allocations within household  $j$ .

<sup>6</sup> For adults,  $\upsilon$  represents the degree of assortative mating; for children,  $\upsilon$  represents heritability.

The household seeks to maximize the above utility function subject to these production technologies and a time-budget constraint,

$$w_m \sum t_{qm} + w_f \sum t_{qf} + w_b \sum t_{qb} + w_g \sum t_{qg} - p x_m - p x_f - \sum p_q x_q = w_m T_m + w_f T_f + w_b T_b + w_g T_g + \sigma_b e_b + \sigma_g e_g + v$$

for  $q \in \{z, h_m, h_f, s_b, s_g\}$

where  $v$  denotes wealth, and  $T_k, T_i$  are the time endowments with  $T = \sum t_q + t_w$ ,  $t_w$  equal to the time spent in income-generating activities.  $w_k$  denotes the (shadow) price of time for adults, which depends on endowments and past production of human capital.

$$w_k = r_k y_k \quad \text{where } y_k = y(s_k; \gamma_k, \nu_k) \quad \text{for } k = m, f$$

Similarly, children's future earnings depend on endowments and the current production of human capital

$$e_i = r_i y_i \quad \text{where } y_i = y(s_i; \gamma_i, \nu_i) \quad \text{for } i = b, g$$

where  $r_k, r_i$  are gender-specific market returns to skill units.  $\sigma_i$  denotes the pre-determined share (e.g. through social norms) of children's future earnings which are used to support parents in old age. The price of children's unskilled time  $w_i$  is the (exogenous) market wage for child labor.

Four types of households are selected to examine variation in the time available for household production, the amount of income household members are able to earn, and the distribution of bargaining power among mothers and fathers. Here, I present a brief description of each household type within the context of the simple model. Section III provides a more detailed description with respect to the limitations of the selected data sources. The implications of relaxing these assumptions will be discussed with the empirical results in Section V.

#### *Baseline: Father Present and Contributing Income*

This group will be represented by children with married parents both currently residing in the same household without any significant periods of absence. In these households, decisions regarding children's schooling are made jointly by both parents, according to their respective

bargaining weights ( $\lambda_m > 0$  and  $\lambda_f > 0$ ). Both parents are contributing to the household by earning income and/or engaging in production of household goods, and both parents are able to provide inputs to the production of children's human capital ( $T_m = T_f = T$  and  $t_q \geq 0$  for  $q \in \{z, h_m, h_f, s_b, s_g\}$ ).

*Case 2: Father Not Present but Contributing Income*

This group will be comprised of children with fathers who are long-term migrants.<sup>7</sup> These fathers are still actively contributing income ( $T_m - \sum t_q > 0$ ) and thus actively participating in decisions regarding children's schooling outcomes ( $\lambda_m > 0$ ). However, the length of their absence(s) implies that they cannot contribute to the production of household goods other than their own health ( $t_{zm} = t_{hf,m} = t_{sb,m} = t_{sg,m} = 0$ ), and both sons and daughters may be required to allocate more time to household activities. Changes in the allocation of household members' time will depend on the relative elasticities of substitution. If mothers' time is the closest substitute for fathers' time and daughters' time is a better substitute for mothers' time than sons' time, girls' schooling will suffer relative to their male siblings. The father's absence, however, also implies that other household members need not contribute time to the production of the his health ( $t_{hm,f} = t_{hm,b} = t_{hm,g} = 0$ ); this will partially offset the increased demand for time in household production.

*Case 3: Father Present and Not Contributing Income*

This group will be comprised of married, dual-parent households in which the father experiences an illness. Illness is a shock to the health production function, via a change in the individual's health endowment. This shock reduces the productivity of time inputs to production, at a minimum forcing the sick individual to increase time spent in the production of health ( $t'_{hk,k} > t_{hk,k}$ ). An illness that prevents the father from carrying out his daily activities will also affect household income and the time available for household production as  $t_{hm,m} \rightarrow T$ . To

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<sup>7</sup> Defined as living away from the household for more than six months of the preceding year.

impose the further limitation that this shock does not affect bargaining power within the household, the illness must be temporary.<sup>8</sup> This sample is analogous to the case presented above, with two exceptions: (1) fathers' earned income is negatively affected, and (2) instead of reducing other household members' time in the production of fathers' health, illness is likely to increase the demand for that time.

In the absence of complete income insurance (via markets, informal risk pooling, or inter-household transfers) or a sufficiently large buffer stock of assets to facilitate consumption smoothing, households may cope with this shock in a variety of ways. The household may decide to temporarily forgo expenditures on children's schooling in order to cover consumption of necessities. However, the change in income may itself induce an additional change in time allocation. Mothers may allocate more time to income-generating activities, thereby requiring children to increase their time in household activities, or children may be pulled into income-generating activities directly. In this respect, this sample will only permit the 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 time available for household production. When there are fixed costs associated with increasing time allocation to income-generating activities, *e.g.* search costs in the labor market, such a change will be less likely to occur for smaller income shocks. The analysis will focus on illnesses no longer than one month in duration, which is likely to be a relatively small income shock.

Relative to the previous case, the change in time allocation is larger in households where the father experiences a debilitating illness. There are two reasons for this: illness increases the demand for time in the production of father's health in addition to reducing the total amount of time available for household production, and the loss of income may induce an additional change

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<sup>8</sup> Illness may also lead to a decline in  $\omega_m$ ; identification requires only that the change in  $\lambda_m$  relative to  $\lambda_f$  be less than in the case of death.

in time allocation. If increasing the demand for children's time in household production has a negative effect on schooling, the difference in children's outcomes between this group and the previous group will provide a downward-biased estimate of the compensated income effect.

*Case 4: Father Not Present and Not Contributing Income*

Households in which the mother is currently widowed will comprise this group. Death of the father implies that  $T_m = 0$  and thus  $t_{zm} = t_{hf,m} = t_{sb,m} = t_{sg,m} = 0$  and  $t_{hm,f} = t_{hm,b} = t_{hm,g} = 0$ . The latter constraints are equivalent to those in Case 1, and there is an additional effect on money income analogous to that in Case 2.<sup>9,10</sup> When the father is either ill or deceased, the household experiences the withdrawal of both time and monetary contributions previously made by the father but, in the case of death, this withdrawal is permanent. This permanence entails an additional effect on household bargaining; the death of the father reduces  $\omega_m$  and shifts decision-making power to the mother ( $\lambda_f = 1$  and  $\lambda_m = 0$ ).<sup>11</sup> The difference in 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. However, death of the father affects the household's current and permanent income, and the reduction in work hours due to death is likely to be larger than in the case of illness. If the effect of a reduction in income is negative, the estimated effect of women's preferences, *i.e.* the difference in schooling outcomes between

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<sup>9</sup> Note that, even if the death was perceived to be imminent by the household (*e.g.* the result of an extended illness), the household still experiences an unanticipated shock to income, although it may not occur precisely at the time of death. In that case, this sample would permit identification of the consequences in the medium-run, rather than the short-run adjustments made by the household.

<sup>10</sup> For both Case 2 and Case 3, households may also have access to various consumption smoothing mechanisms, but they are unlikely to be comprehensive enough to attenuate the effect of the shock to zero.

<sup>11</sup> Identification requires only that  $\lambda'_m < \lambda_m$  and  $\lambda'_f > \lambda_f$ , where  $\lambda'_f$  represents the mother's bargaining power after the death of the husband, and  $\lambda'_m$  captures the extent to which the mother's preferences are consequently mediated by members of the extended family, *e.g.* parents and/or in-laws.

this case and the previous case, will be biased downward. Illness is also likely to increase the time spent in the production of  $h_m$ , whereas death of the father implies  $t_{hm,f} = t_{hm,b} = t_{hm,g} = 0$ . If reducing the time available for household production has a negative effect on children's schooling, then the estimated effect of women's preferences will be biased upward.

### III. Data

Data are drawn from the 1990 Indonesian Population Census and the 1993 Indonesian Socio-Economic Survey (SUSENAS). Both datasets include individual-level information on demographic characteristics, schooling, time allocation, and household characteristics. The SUSENAS also includes information on morbidity and health care in the month prior to enumeration. Children can be matched to co-resident mothers by a unique identifier based on their biological relationship, and fathers' characteristics may be inferred from the mother's spouse identification number and household relationship variables.<sup>12</sup> Data is not collected for individuals who have been away from the household for six months or more at the time of enumeration, although they are still considered members of the household.

Current school enrollment is the outcome of interest.<sup>13</sup> The sample of interest is children between the ages of ten and sixteen, inclusive. In 1984, the government of Indonesia instituted compulsory schooling of six years, equivalent to completion of the primary level. Thus variation in school attendance is minimal between the ages of five and nine, and much of the variation at

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<sup>12</sup> Matching children to fathers requires that children first be matched to their mothers and then to the spouse of the mother. The relationship is further verified by ensuring that the child's reported relationship to the household head is "child". Nonetheless, because fathers and children cannot be matched by a unique identifier based on biological relationship, it is possible that some stepfathers are included.

<sup>13</sup> The analysis does not focus on more cumulative measures of schooling, such as highest level of schooling completed, because identification relies on transitory shocks.

such young ages likely reflects unobservable characteristics of the child, rather than preferences of the parents. As children grow older, preferences of the parents are also 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 sixteen year olds, slightly more than 77% live with mothers, compared to approximately 72% and 67% for seventeen and eighteen year olds, respectively (see Table 1). Sons tend to remain at home with mothers longer than do 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 and older will overstate school attendance and completion of girls relative to boys. Setting the truncation point lower will result in a substantial loss of variation in schooling outcomes, and a higher truncation point induces clear selection bias. Children ages seventeen and older will be omitted from the sample; these observations will not be included in estimation, but observations for younger children in the same household will be retained.

As discussed in the preceding section, samples of four household types will be selected. The baseline group will correspond to children with married parents both in co-residence for at least half of the previous year, selected from both datasets. A sample of families in which the father is away for six months or more in the year preceding the enumeration date is selected to control for the effect of fathers' time contribution, holding income and bargaining power constant. From the SUSENAS, I can identify fathers who have been ill in the last month and

confirm that the illness has disrupted their daily activities; this will distinguish the sample experiencing a shock to both income and time which does not affect bargaining power within the household. Children of currently widowed mothers will comprise the group subjected to changes in income and time as well as changes in parental control over income. The possible selectivity of each sample will be discussed in turn. Descriptive statistics are presented in Tables 2 and 3.

The SUSENAS data provides information on illness in the last month, as well as whether the illness disrupted the person's daily activities and the length of time disrupted. Of fathers reporting a health-related disruption to their daily activities, approximately 40% report no work in the previous week, compared to 5% of all fathers. Roughly 60% of the working ill fathers report fewer hours of work than the mean for all fathers. This suggests that a large majority of these households experience some reduction in the quantity of time the father can allocate to productive activities, and all presumably experience some decline in the quality of work time. Health expenditures in the previous month are roughly three times higher in households in which the father experiences an illness which disrupts his daily activities, suggesting that these illnesses are indeed quite severe.

The type of illness is also recorded, but the categories are not sufficiently detailed to permit the identification of (1) exogenously occurring afflictions or (2) temporary versus permanent illnesses. With regard to the former, this suggests that households with low levels of health are likely to be over-represented in this sample. Low levels of health may also signal that these households have low endowments which, in turn, may lead to lower levels of schooling for all children in these households. Summary statistics suggest that this may be true; both mothers and fathers in these households are slightly older and have fewer years of schooling, relative to the baseline sample. Households with lower endowments likely have lower earnings potential and thus lower income; to the extent that gender equality within the household is a normal good,

the gap between sons and daughters will also be larger in these households even prior to the illness. However, approximately 70% of households experiencing such an illness had lower health expenditures in the previous year than the sample average, whereas average total expenditure in these households is 95% of that in baseline households. This suggests that these households do not have substantially lower endowments in health or otherwise, although it is possible that maintaining a certain level of overall consumption is coming at the cost of lower health.

In the sample of interest, married men living with their wives and at least one school-age child, approximately 85% report being disrupted for two weeks or less, and over 92% report being disrupted for three weeks or less; roughly 6% of observations may be censored at 30 days. It is unlikely that illnesses lasting less than one 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 these disruptions are caused by chronic illnesses which are only periodically debilitating. In those cases, husbands contribute less income and may not be able to play as large a role in the daily management of the household. Consequently, wives may have higher bargaining power, and the son-daughter differential may be smaller prior to the illness.

Children with currently widowed mothers are selected from both data sources. Unfortunately, the data do not provide cause of death or length of widowhood, nor do they permit identification of remarried widows. This is problematic for several reasons: (1) widowhood may not be random; (2) children's schooling may exhibit some persistence; and (3) current widow status may be endogenous. The choice to remain unmarried following widowhood may be a reflection of the mother's inability to find a suitable match in the marriage market or unwillingness to have her preferences mediated by a new husband. It is possible that differences in allocations between sons and daughters are the cause of current widowhood, as

observed in the data, and not vice versa. The direction of bias in this case is not clear; women who choose to remain unmarried may have strong preferences for daughters' education, whereas women with poor marriage prospects may have experienced more gender discrimination from their own parents and thus have less inequality aversion. Data from the first wave of the Indonesian Family Life Survey indicate that roughly 45% of widowed women do not remarry, and those who do remarry have, on average, less schooling whereas those who remain unmarried have a distribution of schooling similar to the baseline (see Table 4). The lower average schooling of widows in the estimation sample can be partially explained by the difference in average age and the fact that the baseline sample includes observations from both the 1990 census and the 1993 SUSENAS, whereas widows are drawn only from the 1990 data.

Certain occupations are associated with higher risk of death. These types of occupations may be more appealing to lower endowment individuals or may directly affect the health endowment over time. Lower average schooling attainment in the sample of widows supports this hypothesis, assuming a sufficient degree of assortative mating. Conversely, hazardous occupations may entail a compensating differential. Widows may also be representative of women who did not fare well in the marriage market and thus married older husbands, although the more advanced age of the husband may have been compensated by greater wealth. However, the ex ante *expected* present discounted value of income for these households is still likely to be less than that for the baseline sample. To the extent that the lower level of income for these households cannot be attributed to an exogenous shock, the son-daughter differential is likely to be biased upward, again assuming that equality is a normal good.

Observed outcomes for children will require some time to adjust and respond to the change in resource allocation due to widowhood and the mother's subsequent increase in control over income. Although the duration of widow status is unknown, the maximum length of widowhood can be approximated by the age of the mother's youngest child, assuming no out-of-

wedlock pregnancies. Schooling is also unique in that it requires continuous investment; children that have been out of school for some time may not return, even if mothers' prefer that they have more education. Consequently, if children and particularly daughters of widows were disadvantaged even prior to the father's death, this may not be overcome with a later increase in the mother's decision-making power. That is, if widowhood occurs some time after children have left school, the effect of the change in household bargaining will be attenuated, given the nature of schooling and human capital accumulation.

Households in which the father is away for six months or more in the year preceding the enumeration date comprise the fourth sample. When building the household roster, all members are listed next to individual identification numbers, with the head of the household listed first (person identifier=1). Household members are defined as "persons who usually live and eat in this household". After the list has been completed, enumerators are instructed to then add individuals "who usually live here but have been away for less than six months" and delete those "who have been away for more than six months". Data are then collected for all individuals that have not been crossed off the roster. Observationally, these households are identified by the fact that there is no record for the household head, *i.e.* the individual with identifier equal to one, even though the children's mother reports that she is currently married and the spouse of the household head. The characteristics and activities of these missing household heads are not reported, therefore I cannot confirm that they are in fact economic migrants who remit money to the household. Remittances are crucial in this sample, as the intent is to identify households in which income is unaffected while fathers' non-monetary inputs are withdrawn. Descriptive statistics, however, indicate that these households have observable characteristics very similar to the baseline sample. Fewer children and mothers in this sample work outside the home, suggesting that the absent fathers are still contributing considerable amounts of income.

#### IV. Estimation Strategy

The reduced-form linearized demand equation for human capital of child  $i$  in household  $j$  of type  $t$  (migrant, sick, or widow) can be expressed as

$$S_{ij} = \alpha + (\delta + \delta')D_{ij} + (\beta + \beta')\mathbf{P}_{ij} + (\beta_g + \beta_g')(D_{ij} \cdot \mathbf{P}_{ij}) + (\gamma + \gamma')\mathbf{H}_j + (\gamma_g + \gamma_g')(D_{ij} \cdot \mathbf{H}_j) + \mu_j + \varepsilon_{ij}.$$

where  $D_{ij}$  takes on a value of one for girls and zero for boys;  $\mathbf{P}_{ij}$  denotes the vector of individual characteristics including age and age squared;  $\mathbf{H}_j$  denotes the vector of observable household characteristics including parents' ages and ages squared, parents' schooling, wealth, urban residence, mother's age at first marriage, number and sex ratio of siblings in co-residence, and the number of household members in each of twelve age-sex specific groups.  $\mu_j$  is a household-level unobservable, and  $\varepsilon_{ij}$  is an i.i.d. error term. Parameters on individual- and household-level variables are allowed to vary by gender, and all parameters are allowed to vary by household type. All household types are pooled in the estimation to obtain estimates of the entire parameter covariance matrix and tests of parameter constancy across household types.

To minimize the heterogeneity bias caused by  $\mu_j$ , a household fixed effects model is utilized. Taking the case of a household with one boy and one girl child, the differenced equation can be written as

$$S_{gj} - S_{bj} = (\delta + \delta') + (\beta + \beta')(\mathbf{P}_{gj} - \mathbf{P}_{bj}) + (\beta_g + \beta_g')\mathbf{P}_{gj} + (\gamma_g + \gamma_g')\mathbf{H}_j + (\varepsilon_{bj} - \varepsilon_{gj}).$$

Because the focus of this analysis is on a binary outcome variable, school enrollment, a conditional logit model (Chamberlain, 1980) is estimated instead of a linear model, such that school enrollment  $S_{ij}$  equals one if  $S_{ij}^* > 0$  and  $S_{ij}$  equals zero otherwise. The key identifying assumption is simply that household type, as represented by the father's state - co-resident, migrant, temporarily ill or deceased - is exogenous with respect to the boy-girl differential within a household. It must also be the case that unobservable characteristics which are common across household members do not differentially affect sons and daughters. To clarify the implications

of this assumption, let us consider what is represented by the household specific term  $\mu_j$ . The household fixed effect captures the role of (1) the common endowment across household members and (2) preferences in the demand for children's schooling that do not vary between sons and daughters. Any gender-based preference will be subsumed in the indicator variable for girls, and this parameter is allowed to vary by household type.<sup>14</sup> Simple regressions of household type on the sex composition of children living at home reveal no significant relationship between the two (see Table A1). These results suggest that there is no relationship between gender preferences and the probability of migration, temporary illness, or death of the father, at least with respect to child survival and fostering.

For a reduced-form demand equation, household income does not appear as a right-hand side variable due to the endogeneity of time spent in income-generating activities. Controls for the household's full income are included, but individual-level data on wages are not available. The common household-level endowment affects the quantity of parents' skill units and thus wage rates. With the inclusion of fixed effects, parents' earnings potential, *i.e.* their unobservable skill endowments, need only be exogenous to the boy-girl differential within a household. When gender equality within the household is a normal good, this restriction may be problematic. If high endowment households have smaller gender disparities in education, the estimated differential effect on daughters in households experiencing either paternal illness or death will be biased downward, *i.e.* girls appear to be worse off relative to their male siblings, if these events are more likely to occur in households with low endowments. Given that death is a more severe outcome than illness, it is expected that the magnitude of this bias will be larger for

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<sup>14</sup> The market rates of return to 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 the household, this is not problematic. The estimated coefficient for the baseline girl dummy variable will also capture the differences in the returns to human capital across males and females.

children with deceased fathers than children with ill fathers. With regard to households in which the father is a long-term migrant, the inclusion of household fixed effects now requires only that the migration decision be exogenous with respect to the gender gap in schooling. If individuals with high endowments are more likely to migrate, the estimated differential effect on daughters will be biased upwards in these households.

The net effect of changes in time available for household production, household income, and the distribution of bargaining power can be inferred from a comparison of outcomes across samples. As stated above, married co-resident parents represent the baseline case. Children whose fathers are long-term migrants suffer only the withdrawal of fathers' time, specifically time spent in household production, while children whose fathers are temporarily ill will experience a negative shock to both income and time. These fathers can still contribute income to the household, so there is no discernible effect on bargaining power between parents in these households. Finally, children with widowed mothers experience a shock to income and time that simultaneously shifts bargaining power to the mother. These samples are designed to permit the following decompositions:

$$\hat{S}_g^M / \hat{S}_b^M - \hat{S}_g / \hat{S}_b = \text{effect of changes in time allocation}$$

$$\hat{S}_g^I / \hat{S}_b^I - \hat{S}_g^M / \hat{S}_b^M = \text{effect of changes in income}$$

$$\hat{S}_g^W / \hat{S}_b^W - \hat{S}_g^I / \hat{S}_b^I = \text{effect of changes in bargaining power}$$

where  $\hat{S}$  denotes predicted school enrollment for the baseline sample,  $M$  denotes children with migrant fathers,  $I$  denotes children with ill fathers, and  $W$  denotes children of widowed mothers.

## V. Results

### *Mothers' Hours of Paid Work*

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' hours of

paid work in the previous week. Parameter estimates are reported in Table 5; coefficients in columns 2 through 4 should be interpreted relative to the baseline (un-interacted) coefficients. Because all parameters are allowed to vary by household type, the results are best summarized with predicted values calculated at the overall sample mean. Wives of long-term migrants worked approximately the same number of hours in the previous week as women in the baseline sample. The reduction in time available for household production due to the husband's absence does not appear to induce much substitution out of paid work. However, this effect is partially offset by the simultaneous reduction in demand for time spent in the production of husbands' health.

Women with husbands who were temporarily ill worked two hours more in the previous week than women in the baseline or migrant samples. Husband's years of schooling has a slightly larger negative effect on wife's work hours when he is ill; schooling may mitigate the loss of income due to illness via occupational choice or improved access to insurance. The number of children living at home has a larger positive effect on mothers' work hours, but very young children (age 5 and younger) have a larger negative effect. This suggests that resource constraints are tight in these households; more "mouths to feed" lead to more time in the labor market, while young children require more time spent in household activities. The larger negative effect of young children further suggests that sick husbands do not take over the household chores typically performed by their wives, *i.e.* the complementarity in production of own health and other household goods is relatively low. The length of the husband's illness has a significant positive effect - the more severe the shock to income, the more mothers are pulled into wage labor. When women in these households do not or cannot enter the labor market in response to husband's illness, the shock to household income is more pronounced. It is possible, however, that the limited labor supply response on the part of wives indicates that the loss of husband's income is either negligible or partially offset by transfers and/or insurance. Mitigation

of the income shock would imply that allocations to children should be no different than in the baseline case, once the change in time allocation is taken into account.

As expected, widows work considerably more, approximately 12 hours more in the previous week, and the age profile is flatter and less concave. Urban residence has a positive effect, also found for the sample of women with ill husbands, which suggests that, where labor market opportunities are better and more readily available, women take advantage of these opportunities when faced with a negative shock to income. Years of schooling and the (imputed) length of widowhood exhibit an inverse relationship with hours of work; because the dependent variable is hours rather than wages, it may be that women with more schooling and more labor market experience are able to earn more per hour and thus work fewer hours, given the demands on their time in the home. Widowed mothers cannot spend as much time in the household, which is likely to pull sons and/or daughters into household activities that compete with schooling.

#### *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 6. Specification I includes the child's age, age interacted with sex, and household level covariates (wealth, demographic composition, urban residence, mother's characteristics, and father's characteristics if present) interacted with the child's sex. Specification II adds the length of widowhood or illness, and specification III allows all household-level covariates to vary by household type. A joint test of significance for a subset of the covariates in Specification III (number and sex ratio of siblings at home, household members in 12 age-sex categories, mother's age at first marriage and urban residence interacted with the child's sex and household type) cannot reject the null that the parameters are jointly equal to zero for all household types. The preferred estimates, specification IV, do not control for these additional covariates.

Specification V considers an alternative measure of schooling which better captures actual attendance in the previous week. However, because this measure is based on the child's reported *primary* activity, 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 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. Furthermore, illness is reported for the previous month and thus may not coincide with the child's reported activities in the previous one week.

The full set of parameter estimates for Specification IV is reported in Table A2. Baseline parameters indicate that observable household characteristics do not differentially affect daughters, with the exception of mothers' years of schooling and floor area of owned homes, which both have a positive effect. Fathers' characteristics have a small and insignificant differential effect for girls with married co-resident parents, which suggests that the omission of these covariates for migrant fathers does not generate significant bias. The coefficient on the dummy variable for girls is positive and insignificant, but the probability of being enrolled in school declines more rapidly for girls as they age. There are no statistically significant differential effects for daughters with migrant fathers. Age effects are not significantly different from the baseline, with the exception of a positive effect for children age 16. This may reflect higher income in migrant households which is used to finance more education for all children.

The large and significant direct negative effect on daughters in households that experience a temporary illness is the most striking result in this regression. The effect is exacerbated by the length of the father's illness, as evidenced by the negative coefficient on this covariate. However, the positive effect of father's schooling and the positive but diminishing effect of father's age on daughters' school enrollment suggest that older and more educated fathers may be better able to mitigate the loss of income due to illness. The probability of being

enrolled in school declines more rapidly for boys and less rapidly for girls in this group, although most of these effects are not significant. There is some evidence that schooling investments are protected for children nearing completion of primary school; children age 12 or 14 are significantly less likely to be enrolled in school, relative to the baseline, but the point estimate is considerably smaller and not statistically significant for age 13.

For girls with deceased fathers, there are again no statistically significant differential effects, with the exception of a positive effect for the youngest girls. Other age effects for girls in this sample are also positive but not significant. This pattern suggests that widows may be attempting to increase investments in girls' schooling, but only the youngest daughters are able to benefit from this when there is some cost to either returning to school after a prolonged absence or beginning school at much older ages. The pattern of age effects for all children again suggests that schooling investments are protected for children nearing completion of a schooling level. The relative age coefficients for this sample are significant and negative for all ages except ages 13 and 16, the ages at which most children complete primary and secondary school, respectively. Length of widowhood, as proxied by the age of the mother's youngest child, has a positive but insignificant effect on girls' school enrollment relative to boys, suggesting that outcomes require some time to adjust. The measurement error in this variable biases the estimated parameter towards zero.

Estimating partial effects and imputing predicted probabilities requires the selecting an appropriate value of  $\mu_j$  to utilize in the calculation. These parameters are not directly estimated by the conditional logit estimator, nor are any restrictions place on the distribution of  $\mu_j$ . To circumvent this issue, I calculate the odds ratio of school enrollment for girls relative to boys; taking the ratio of the odds ratios eliminates  $\mu_j$  from the calculation. Comparing these estimates across samples will then permit a decomposition of the sources of discrimination, and the independent (net) effects of changes in time allocation, income and bargaining can identified.

Table 6 presents this decomposition for the oldest children in the sample. The net effect of time allocation is calculated as the difference in the odds of school enrollment for girls relative to boys between household in which the father is a migrant and baseline households. The net effect of time allocation is calculated as the difference between households in which the father is temporarily ill and those in which he is a migrant, and the effect of bargaining power is calculated as the difference between households in which the father is deceased and those in which he is temporarily ill. Under the preferred estimates, specification IV, reducing the time available for household production has a small and insignificant negative effect on daughters relative to sons. Reducing household income has a larger and significant negative effect, whereas increasing mothers' bargaining power has a significant positive effect which is the largest in magnitude of the three.

These effects, however, are heterogeneous with respect to the child's age, as shown in Table 7. The effect of a reduction in time available for household production has a large negative effect on girls at ages 11 and 12, and a small positive effect at ages 13, 14 and 15 before again turning negative at age 16. This age pattern suggests that, for younger children, mothers are better substitutes for fathers in household production, and daughters, in turn, are better substitutes for mothers than are sons. In the 13-15 age range, boys and girls are similarly productive in the household but, as children near completion of secondary school at age 16, boys are less likely than girls to be pulled out of school and into household production. The point estimates, however, are quite small in magnitude, and none are significantly different from zero. Reducing household income, while holding the time available for household production constant, has a significant differential effect on daughters age 12-16, with the effect being positive at ages 12 and 14. Surprisingly, younger girls' schooling investments appear to be less income-elastic than boys', with the exception of age 13, which is typically the age at which primary schooling is completed. After children have started in secondary school, however, boys' schooling appears to

be better insulated from income fluctuations. Increasing mothers' bargaining power, holding time allocation and household income constant, increases the probability that girls are enrolled in school, relative to their brothers. This is true at all ages except age 12, which again suggests that boys' schooling is more protected as they near completion of primary school.

As a comparison, Table 7 also presents the odds of school enrollment for girls relative to boys in each household type, relative to baseline households. 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. It is clear that the estimates in column (2) are not equivalent to those in column (4), and the estimates in column (3) are not equivalent to those in column (5). In other words, the sample of children with ill fathers is not, alone, an appropriate sample from which to identify an income effect, and the sample of children with deceased fathers is not, alone, an appropriate sample from which to identify a bargaining power effect, even controlling for unobservable characteristics of the household that are constant across children. The distinct effects of changes in time allocation, income and bargaining power must be carefully accounted for. Focusing on the critical ages of 13 and 16, *i.e.* the ages at which, respectively, primary and secondary schooling are typically completed, we see that girls with deceased fathers fare worse than girls in the baseline, relative to their male siblings. However, once the reduction in household income and the reduction in time available for household production are netted out, I find that increasing mothers' bargaining power in fact has a strong positive effect on girls' school enrollment relative to boys'. Similarly, we see that girls at ages 13 and 16 fare worse in households with ill fathers but, once the change in time allocation is accounted for, the magnitude of this effect changes substantially; the effect of a change income net of changes in time allocation is larger at age 13 and smaller at age 16 when compared with the effect of simply having an ill father.

Relaxing the assumptions on changes in time allocation, income and household bargaining for each sample implies that these decompositions can yield only bounded estimates of the effects described above. However, the direction of bias can be signed in most cases. The main categories of potential bias are summarized in Table 10, in which the stated directions of bias are based on the empirical conditions given in the rightmost column. As discussed in the previous section, if the unobservable household-level skill endowment  $\mu_j$  is correlated with gender disparities in education, the estimates will be biased upwards, *i.e.* girls appear to be better off relative to their male siblings, for households with high endowments and downwards for households with low endowments, provided that endowments are negatively correlated with gender disparities. Thus, if migrants have high endowments, the estimated effect of a reduction in time available for household production on girls' school enrollment relative to boys' will be biased upward. If ill fathers have low endowments, the differential effect on girls will be biased downwards and, subtracting the upward biased effect found for migrant households will yield a net downward bias in the estimated effect of a reduction in household income on girls' school enrollment relative to boys'. Similarly, the estimated effect of an increase in mothers' bargaining power will be biased downwards, provided that paternal death signals even lower skill endowments than does illness.

Temporary illness, migration and death impose similar constraints on the amount of time fathers can devote to household production. However, illness requires an increase in other household members' time spent in the production of fathers' health ( $\Delta t_{hm} > 0$ ) whereas migration and death both imply a decrease ( $t_{hm} = 0$ ). When reducing the time available for household production has a disproportionately negative effect on girls, girls will appear to be worse off in households in which time must be diverted away from household production and towards the production of fathers' health. 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. Illness and death also reduce the amount of time fathers can spend in income-generating activities ( $\Delta t_{w,m} < 0$ ), and this may, in turn, pull other household members into the labor market and induce an additional change in children's time allocation. This effect 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. If children are less productive than mothers 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. In this case, the estimated effect of a reduction in income on girls' schooling relative to boys' will be biased downward. The estimated effect of an increase in mothers' bargaining power will also be biased downward, given that fathers' labor market hours necessarily fall to zero when he is deceased and are weakly greater than zero when he is debilitated. Related to the issue of changes in household income is the fact that migration, illness and death have differing implications for changes in fathers' labor supply versus changes in fathers' wages. Temporary illness affects only labor supply, but fathers' wages are effectively zero after death. And, conversely, fathers likely choose to migrate precisely because it provides the opportunity to earn higher wages. Under the assumption that reducing household income has a negative effect on girls' schooling relative to boys', this will generate an upward bias in the estimated effect of time allocation and a downward bias in the estimated effect of bargaining power.

Lastly, migration and illness of the father may affect mothers' bargaining power. Wives may gain de facto control over intra-household allocations simply because husbands are not in close proximity; this, however, is unlikely to affect school enrollment because it is an easily monitored and thus perfectly contractible good. Alternatively, husbands living away from the

home may grant wives more control over their children's activities to compensate for their reduced contribution to household production. 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. If women have stronger preferences for gender equality, this will generate an upward bias in the estimated effect of both changes in time allocation and changes in household income.

### *Schooling Attainment*

To verify the assumptions and results presented above, I 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 unaffected by short-term changes in income and 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.

Tables 8 and 9 reveal no significant effects by household type on daughters' schooling attainment relative to sons'. Empirical specifications are analogous those used in the analysis of school enrollment, and a joint test of significance for the additional household characteristics included in specifications I through III cannot reject the null that the parameters are jointly equal to zero for all household types. The full set of parameter estimates for the preferred specification (IV) is reported in Table A3, and these estimates are used for the calculations presented in Table 9. For households in which the father is deceased, there are no significant differential effects on daughters' schooling attainment, although all older children have less schooling relative to those in baseline households. There are again no significant differential effects for daughters' years of schooling in households in which the father experiences a temporary illness. The same is true for migrant households with the exception of mothers' schooling, which has a positive

differential effect on daughters. More educated mothers appear to be better able to influence children's schooling decisions when husbands are living outside the household, which is consistent with the earlier assertion that the difference between migrant and baseline households is an upper bound on the true effect of changes in time allocation on the educational gender gap. This finding may simply be reflecting the correlations between migration, unobservable skill endowments, wives' schooling (via assortative mating), and gender discrimination.

A final hypothesis to consider is that the absence of the father from the household may differentially affect the production function for sons' and daughters' human capital. The selected sample of migrants are away from the home for six months or more out of the year, and thus comparison of parameters between widow and migrant households can, at a minimum, reveal the direction of this bias. As mentioned above, there are no statistically significant differential effects for daughters in migrant households, with the exception of mothers' schooling, and the point estimates are generally quite small in magnitude. By symmetry, these results suggest that there are no differential effects on sons; absence of fathers does not appear to affect the long-term accumulation of sons' human capital relative to daughters.

## **VI. Conclusion**

This paper seeks to disentangle the various factors contributing to gender disparities in investment 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. Reducing the time available for household production has an insignificant and relatively small impact on gender disparities in school enrollment. However, failing to account for the change in time allocation associated with an income shock can lead to a sizable bias in the 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 will also result in a large bias, even after controlling for unobservable characteristics of the household that affect all children identically. In this particular case, the sign of the effect is reversed – daughters of deceased fathers are worse off, relative to their male siblings, but, in fact, the increase in mothers' bargaining power has a large and significant positive effect on daughters relative to sons.

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. Increased demand for household labor generally has a gender-neutral effect on children's school enrollment, but girls are more vulnerable to being pulled out of school to compensate for a reduction in the total time available for household production. Low income has a heterogeneous effect on gender disparities in education that varies with the age of the child. Overall, girls' schooling is less income-elastic than boys' schooling, but boys' schooling is much more likely to be protected from income fluctuations when they are nearing the completion of either primary or secondary schooling. Increasing women's bargaining power has the largest effect on reducing gender disparities. This finding indicates that women and men have differing preferences with regard to gender equality, although not necessarily with regard to the level of children's human capital. Programs aimed at bolstering women's position in the household (*e.g.* microfinance) should, thus, have the largest effect on reducing gender disparities in human capital investment. However, policymakers must also be aware that the induced changes in time allocation and/or income may work in the opposite direction.

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Table 1. Children's Residence and Schooling Status

Age	Children Living with Mom (%)			Children Attending School (%)			
	Overall	Boys	Girls	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

Table 2. Descriptive Statistics, Children Ages 10 to 16

	Boys	Baseline	Migrant	Sick Dad	Widow
Age		12.76 (1.97)	12.78 (1.97)	12.77 (1.99)	13.31 *** (1.97)
Schooling Attainment		5.88 (2.28)	5.96 *** (2.30)	5.75 *** (2.28)	6.08 *** (2.44)
Schooling Attainment (% of Potential)		0.664 (0.186)	0.672 *** (0.192)	0.647 *** (0.186)	0.649 *** (0.211)
School Enrollment		0.806 (0.396)	0.793 *** (0.405)	0.795 ** (0.404)	0.656 *** (0.475)
School as Primary Activity		0.795 (0.404)	0.783 *** (0.412)	0.780 *** (0.414)	0.647 *** (0.478)
Work for Pay		0.169 (0.375)	0.152 *** (0.359)	0.215 *** (0.411)	0.263 *** (0.440)
Work in Addition to School		0.054 (0.227)	0.036 *** (0.187)	0.087 *** (0.282)	0.045 *** (0.206)
Number of Observations		104,189	10,741	6,483	34,527
	Girls	Baseline	Migrant	Sick Dad	Widow
Age		12.73 (1.96)	12.70 (1.97)	12.79 ** (1.97)	13.27 *** (1.96)
Schooling Attainment		5.98 (2.28)	6.00 (2.32)	5.89 *** (2.25)	6.17 *** (2.44)
Schooling Attainment (% of Potential)		0.678 (0.187)	0.683 *** (0.196)	0.663 *** (0.181)	0.662 *** (0.212)
School Enrollment		0.801 (0.399)	0.790 *** (0.407)	0.793 (0.405)	0.657 *** (0.475)
School as Primary Activity		0.792 (0.406)	0.777 *** (0.416)	0.780 ** (0.414)	0.649 *** (0.477)
Work for Pay		0.114 (0.318)	0.103 *** (0.305)	0.139 *** (0.346)	0.181 *** (0.385)
Work in Addition to School		0.037 (0.190)	0.027 *** (0.163)	0.049 *** (0.216)	0.031 *** (0.174)
Number of Observations		97,512	9,943	6,043	31,158

Includes only children currently co-residing with their mothers. Standard deviations reported in parentheses. (\*) indicates significantly different from column [1] at the 15%, (\*\*) 10%, or (\*\*\*) 5% level.

Table 3. Descriptive Statistics, Household-Level Characteristics

	Baseline	Migrant	Sick Dad	Widow
Number of Children at Home	3.81 (1.67)	3.63 *** (1.63)	3.48 *** (1.57)	3.10 *** (1.53)
Ratio of Sons at Home to All Children at Home	0.524 (0.286)	0.523 (0.295)	0.526 (0.301)	0.540 *** (0.324)
Floor Area of Owned Home	62.70 (51.60)	59.34 *** (55.49)	62.66 (47.05)	54.72 *** (48.20)
Urban Residence	0.289 (0.453)	0.292 (0.455)	0.282 (0.450)	0.255 *** (0.436)
Mothers' Age	38.74 (7.35)	38.64 (7.91)	39.56 *** (7.88)	45.58 *** (8.62)
Mothers' Schooling Attainment	5.19 (3.78)	5.09 *** (3.78)	4.90 *** (3.69)	3.42 *** (3.68)
Mothers' Age at First Marriage	18.25 (4.56)	18.12 *** (5.42)	18.44 *** (3.77)	18.42 *** (5.37)
Mother Works for Pay	0.480 (0.500)	0.453 *** (0.498)	0.562 *** (0.496)	0.710 *** (0.454)
Fathers' Age	44.47 (8.69)		45.79 *** (9.70)	
Fathers' Schooling Attainment	6.27 (3.91)		5.98 *** (3.85)	
Fathers' Length of Illness			7.90 (8.34)	
Maximum Length of Widowhood				10.76 (4.07)
Number of Observations	157012	12345	7608	43872

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

Standard deviations reported in parentheses. (\*) indicates significantly different from column [1] at the 15%, (\*\*) 10%, or (\*\*\*) 5% level.

Table 4. Schooling Level by Marital Status (%)

	Married Women	Current Widows	Remarried Widows
Primary or Less	72.48	72.90	93.83
Junior High School	12.23	14.95	3.42
High School	12.32	11.21	2.74
College and Graduate	2.97	0.93	0.00
Number of Observations	4716	107	146

Table 5. Tobit Regression of Mothers' Hours of Paid Work

	Baseline	Relative Effects		
		Migrant	Sick Dad	Widow
Constant	-37.929 *** (2.438)	-16.898 *** (5.902)	-29.794 *** (10.555)	56.067 *** (4.102)
Age	2.390 *** (0.097)	0.094 (0.297)	1.229 *** (0.518)	-1.530 *** (0.192)
Age Squared	-0.026 *** (0.001)	-0.001 (0.003)	-0.018 *** (0.006)	0.011 *** (0.002)
# of Children at Home	-2.536 *** (0.114)	-0.315 (0.408)	0.941 * (0.587)	2.955 *** (0.217)
Ratio of Sons to All Children at Home	0.501 * (0.311)	0.813 (1.105)	-0.008 (1.394)	-0.898 * (0.615)
Schooling Attainment	0.569 *** (0.028)	-0.086 (0.084)	-0.131 (0.133)	-0.470 *** (0.053)
Floor Area of Home	0.007 *** (0.002)	-0.003 (0.006)	-0.020 *** (0.008)	-0.024 *** (0.004)
Dummy for Urban	-6.061 *** (0.218)	-0.858 (0.713)	1.931 *** (0.894)	9.306 *** (0.430)
Husband's Schooling Attainment	-0.562 *** (0.027)		-0.597 *** (0.127)	
Husband's Age	-0.413 *** (0.088)		0.133 (0.381)	
Husband's Age Squared	0.003 *** (0.001)		-0.001 (0.004)	
Max. Length of Widowhood				-0.145 *** (0.056)
Length of Illness			0.113 *** (0.045)	
Predicted Hours at Sample Mean	13.12 (1.13)	13.02 (1.22)	15.23 (1.30)	25.28 (1.30)
F-Test that Household Type-Specific Parameters Jointly Equal Zero		3.63	2.76	86.8

Includes kabupaten-level fixed effects, control for year of survey, and number of household members in 12 age-sex categories. Parameters on husbands' characteristics are estimated only for households in which the husband is present, and those reported in column three are direct, not relative, effects. Standard errors in parentheses. (\*) indicates significant at 15%, (\*\*) 10% or (\*\*\*) 5% level.

Table 6. Children's School Enrollment, Conditional Logit Estimates

	I	II	III	IV	V
$\delta$	0.074 (1.038)	0.081 (1.040)	0.969 (2.303)	1.125 (1.473)	1.185 (2.178)
$\delta^M$	-0.097 (0.201)	-0.097 (0.201)	-0.202 (3.491)	-0.561 (2.253)	-1.387 (3.404)
$\delta^S$	-0.079 (0.148)	0.151 (0.205)	-18.386 *** (7.718)	-7.000 * (4.565)	-24.685 *** (8.085)
$\delta^W$	-0.041 (0.215)	-0.107 (0.256)	0.130 (2.912)	-0.459 (1.988)	-0.914 (2.770)
Net Effect on Odds of School Enrollment for Girls Relative to Boys					
Time Allocation	-0.086 * (0.054)	-0.086 * (0.055)	-0.032 (0.694)	-0.117 (0.136)	-0.061 (1.267)
Income	0.016 (0.045)	0.037 (0.044)	0.083 (0.603)	-0.292 (0.170) **	0.629 (1.107)
Bargaining Power	0.033 (0.029)	0.025 (0.029)	0.149 (0.189)	0.370 (0.114) ***	-0.245 (0.190)
Household Fixed Effects	yes	yes	yes	yes	yes
Length of Widowhood/Illness	no	yes	yes	yes	yes
Parameters Vary by Household Type	no	no	yes	yes	yes
Alternative Measure of Attendance	no	no	no	no	yes
Additional Household Characteristics	yes	yes	yes	no	yes

Includes controls for age, floor area of owned home, parents' age and years of schooling and year of survey. Additional household characteristics include number and sex ratio of siblings at home, household members in 12 age-sex categories, mother's age at first marriage, and urban residence. Net effect on odds of school enrollment for girls relative to boys calculated at the overall sample mean for age 16. Standard errors clustered at the household level and reported in parentheses. (\*) indicates significant at 15%, (\*\*) 10% or (\*\*\*) 5% level.

Table 7. Odds of School Enrollment for Girls Relative to Boys by Age

Age	Net Effects			Relative to Baseline	
	Time Allocation	Income	Bargaining	Sick Dad	Widow
10	0.161 (0.420)	-0.606 (0.677)	0.307 (0.447)	-0.446 (0.336)	-0.139 (0.194)
11	-0.390 (0.821)	0.376 (0.583)	0.821 *** (0.153)	-0.014 (0.534)	0.807 *** (0.378)
12	-0.389 (0.598)	1.052 ** (0.552)	-0.518 *** (0.052)	0.663 (0.531)	0.145 * (0.095)
13	0.113 (0.145)	-0.312 * (0.198)	0.159 (0.113)	-0.199 ** (0.106)	-0.040 (0.061)
14	0.022 (0.218)	0.324 * (0.199)	0.022 (0.061)	0.346 *** (0.152)	0.368 *** (0.079)
15	0.098 (0.094)	-0.355 *** (0.150)	0.176 ** (0.098)	-0.258 *** (0.074)	-0.082 ** (0.044)
16	-0.117 (0.136)	-0.292 ** (0.170)	0.370 *** (0.114)	-0.409 *** (0.102)	-0.040 (0.080)

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. Standard errors reported in parentheses. (\*) indicates significant at 15%, (\*\*) 10% or (\*\*\*) 5% level.

Table 8. Children's Schooling Attainment, Household Fixed Effects Estimates

	I	II	III	IV
$\delta$	0.116 (0.634)	0.118 (0.636)	0.392 (1.335)	0.803 (0.833)
$\delta^M$	-0.045 (0.092)	-0.044 (0.092)	-1.210 (2.071)	-0.718 (1.436)
$\delta^S$	0.026 (0.074)	0.041 (0.102)	0.171 (3.434)	-1.590 (2.391)
$\delta^W$	-0.008 (0.104)	-0.048 (0.148)	-0.267 (1.743)	-0.631 (1.224)
<b>Net Effect on Schooling Attainment for Girls Relative to Boys</b>				
Time Allocation	-0.045 (0.106)	-0.044 (0.106)	0.111 (0.265)	-0.049 (0.237)
Income	0.071 (0.116)	0.071 (0.116)	-0.057 (0.381)	0.151 (0.318)
Bargaining Power	-0.034 (0.137)	-0.027 (0.138)	-0.080 (0.388)	-0.112 (0.280)
Household Fixed Effects	yes	yes	yes	yes
Length of Widowhood/Illness	no	yes	yes	yes
Parameters Vary by Household Type	no	no	yes	yes
Additional Household Characteristics	yes	yes	yes	no

Includes controls for age, floor area of owned home, parents' age and years of schooling and year of survey. Additional household characteristics include number and sex ratio of siblings at home, household members in 12 age-sex categories, mother's age at first marriage, and urban residence. Net effect on schooling attainment for girls relative to boys calculated at the overall sample mean for age 16. Standard errors clustered at the household level and reported in parentheses. (\*) indicates significant at 15%, (\*\*) 10% or (\*\*\*) 5% level.

Table 9. Schooling Attainment for Girls Relative to Boys by Age

Age	Net Effect			Relative to Baseline	
	Time Allocation	Income	Bargaining	Sick Dad	Widow
10	-0.091 (0.161)	0.018 (0.214)	0.028 (0.209)	-0.073 (0.157)	-0.045 (0.134)
11	0.029 (0.178)	0.047 (0.243)	-0.015 (0.234)	0.076 (0.184)	0.061 (0.148)
12	0.047 (0.167)	-0.089 (0.221)	0.121 (0.213)	-0.042 (0.163)	0.078 (0.140)
13	-0.019 (0.174)	0.103 (0.240)	-0.111 (0.229)	0.084 (0.182)	-0.028 (0.142)
14	-0.056 (0.187)	0.231 (0.259)	-0.063 (0.242)	0.175 (0.197)	0.112 (0.144)
15	-0.045 (0.203)	0.023 (0.292)	0.029 (0.269)	-0.022 (0.228)	0.008 (0.146)
16	-0.049 (0.237)	0.151 (0.318)	-0.112 (0.280)	0.101 (0.234)	-0.011 (0.157)

The net effect of time allocation is, by construction, identical to the schooling attainment for girls relative to boys in migrant households relative to baseline households. Standard errors reported in parentheses. (\*) indicates significant at 15%, (\*\*) 10% or (\*\*\*) 5% level.

Table 10. Potential Biases in Estimated Net Effects

	Time Allocation	Income	Bargaining Power	Condition
$\mu_j$	+ 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 smaller gender disparities in education.
$\Delta t_{hm}$		- bias Other household members increase time in the production of father's health when he is sick but decrease that time when he is living away from home.	+ bias Other household members' time spent in the production of father's health falls to zero when he is deceased but increases when he is ill.	Reducing time available for household production has a negative effect on daughters relative to sons.
$\Delta t_{w,m}$		- 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 daughters must spend more time in household production.	Reducing household income induces an additional change in time allocation that has a negative effect on daughters relative to sons.
$\Delta w_m$	+ bias Migrants earn higher wages than they would have if they had remained at home.		- bias Loss of income is larger when father dies than for a temporarily debilitating illness.	Reducing household income has a negative effect on daughters relative to sons.
$\Delta \lambda_f$	+ bias Women may have more influence over intrahousehold allocation when their spouses are away.	+ bias Women may have more influence over intrahousehold allocation when their spouses are debilitated.		Increasing women's bargaining power within the household has a positive effect on daughters relative to sons.

Effect of time allocation is calculated as the difference in 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 difference 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 difference between households in which the father is deceased and those in which he is temporarily ill.

Table A1. Logit Regression of Household Type on Observable Characteristics

	Migrant	Sick Dad	Widow
Mother's Age	-0.072 *** (0.006)	0.012 * (0.008)	0.093 *** (0.016)
Mother's Age Squared	0.001 *** (0.000)	0.000 (0.000)	-0.001 *** (0.000)
# of Children at Home	-0.151 *** (0.006)	-0.178 *** (0.009)	0.010 (0.009)
Ratio of Sons to All Children at Home	-0.001 (0.025)	0.032 (0.033)	0.013 (0.040)
Mother's Schooling Attainment	-0.012 *** (0.003)	-0.008 *** (0.003)	-0.074 *** (0.004)
Floor Area of Home	-0.001 *** (0.000)	0.001 *** (0.000)	-0.005 *** (0.001)
Dummy for Urban	0.042 (0.033)	0.072 ** (0.041)	-0.174 *** (0.041)

Includes kabupaten-level fixed effects. Column 3 omits women who may have been widowed for more than five years, as imputed from the age of the youngest child, in order to reflect the child sex ratio at the time of widowhood. Standard errors clustered at the kabupaten level and reported in parentheses. (\*) indicates significant at 15%, (\*\*) 10% or (\*\*\*) 5% level.

Table A2. Conditional Logit Regression of Children's School Enrollment

		Relative Effects		
		Migrant	Sick Dad	Widow
Age 11	-0.166 (0.164)	0.021 (0.357)	-0.637 (0.520)	-0.864 *** (0.260)
Age 12	-1.003 *** (0.131)	0.259 (0.304)	-0.798 ** (0.448)	-0.512 *** (0.217)
Age 13	-2.346 *** (0.136)	0.183 (0.295)	-0.443 (0.467)	-0.240 (0.223)
Age 14	-3.187 *** (0.137)	0.350 (0.300)	-1.049 *** (0.521)	-0.679 *** (0.230)
Age 15	-4.089 *** (0.148)	0.258 (0.336)	-0.601 (0.501)	-0.429 ** (0.240)
Age 16	-5.192 *** (0.165)	0.743 *** (0.345)	-0.508 (0.530)	-0.098 (0.260)
Relative Effects for Daughters				
Constant ( $\delta$ )	1.125 (1.473)	-0.561 (2.253)	-7.000 * (4.565)	-0.459 (1.988)
Age 11	-0.104 (0.251)	-0.471 (0.533)	0.366 (0.796)	0.592 * (0.395)
Age 12	-0.315 * (0.214)	-0.579 (0.487)	0.872 (0.664)	0.234 (0.342)
Age 13	-0.292 * (0.202)	-0.006 (0.452)	0.169 (0.640)	0.065 (0.330)
Age 14	-0.562 *** (0.207)	-0.081 (0.452)	0.733 (0.674)	0.478 (0.337)
Age 15	-0.524 *** (0.219)	0.003 (0.481)	0.011 (0.717)	0.000 (0.352)
Age 16	-0.354 * (0.230)	-0.233 (0.510)	-0.152 (0.731)	0.062 (0.371)
Floor Area of Owned Home	0.002 * (0.001)	0.000 (0.002)	0.003 (0.005)	-0.002 (0.002)
Mother's Age	-0.067 (0.082)	0.045 (0.108)	-0.232 (0.258)	0.033 (0.100)
Mother's Age Squared	0.001 (0.001)	0.000 (0.001)	0.003 (0.003)	-0.001 (0.001)
Mother's Schooling Attainment	0.046 *** (0.018)	-0.010 (0.038)	-0.049 (0.057)	0.003 (0.026)
Father's Age	0.007 (0.061)		0.437 *** (0.216)	
Father's Age Squared	0.000 (0.001)		-0.005 *** (0.002)	
Father's Schooling Attainment	0.001 (0.019)		0.124 *** (0.052)	
Maximum Length of Widowhood				0.023 (0.017)
Length of Illness			-0.037 * (0.023)	
Survey Year=1990	-0.116 (0.176)			

Parameters on fathers' characteristics are estimated only for households in which the husband is present, and those reported in column three are direct, not relative, effects. Standard errors clustered at the household level and reported in parentheses. (\*) indicates significant at 15%, (\*\*) 10% or (\*\*\*) 5% level.

Table A3. Household Fixed Effects Regression of Children's Schooling Attainment

		Relative Effects		
		Migrant	Sick Dad	Widow
Age 11	0.884 *** (0.053)	-0.047 (0.141)	-0.088 (0.170)	-0.002 (0.109)
Age 12	1.833 *** (0.044)	-0.019 (0.124)	-0.045 (0.146)	-0.094 (0.098)
Age 13	2.672 *** (0.051)	-0.073 (0.136)	-0.001 (0.170)	-0.117 (0.104)
Age 14	3.552 *** (0.054)	0.003 (0.147)	-0.048 (0.180)	-0.301 *** (0.107)
Age 15	4.207 *** (0.058)	-0.117 (0.152)	-0.110 (0.189)	-0.465 *** (0.111)
Age 16	4.857 *** (0.064)	-0.185 (0.176)	-0.168 (0.199)	-0.628 *** (0.121)
Relative Effects for Daughters				
Constant ( $\delta$ )	0.803 (0.833)	-0.718 (1.436)	-1.590 (2.391)	-0.631 (1.224)
Age 11	-0.072 (0.076)	0.121 (0.202)	0.150 (0.247)	0.106 (0.162)
Age 12	-0.063 (0.070)	0.138 (0.189)	0.031 (0.227)	0.123 (0.152)
Age 13	0.009 (0.076)	0.072 (0.203)	0.157 (0.241)	0.018 (0.157)
Age 14	-0.018 (0.079)	0.036 (0.210)	0.249 (0.258)	0.157 (0.160)
Age 15	-0.003 (0.087)	0.047 (0.227)	0.052 (0.282)	0.053 (0.167)
Age 16	-0.061 (0.095)	0.042 (0.262)	0.175 (0.287)	0.034 (0.181)
Floor Area of Owned Home	0.000 (0.000)	-0.001 (0.001)	-0.001 (0.001)	0.000 (0.001)
Mother's Age	-0.010 (0.042)	0.000 (0.070)	0.037 (0.126)	-0.001 (0.057)
Mother's Age Squared	0.000 (0.000)	0.000 (0.001)	0.000 (0.002)	0.000 (0.001)
Mother's Schooling Attainment	0.008 (0.009)	0.026 * (0.018)	0.017 (0.027)	0.005 (0.014)
Father's Age	-0.024 (0.033)		0.008 (0.078)	
Father's Age Squared	0.000 (0.000)		0.000 (0.001)	
Father's Schooling Attainment	0.001 (0.009)		-0.010 (0.026)	
Maximum Length of Widowhood				0.005 (0.012)
Length of Illness			-0.001 (0.010)	
Survey Year=1990	0.010 (0.077)			
Intercept	3.520 *** (0.029)			

Parameters on fathers' characteristics are estimated only for households in which the husband is present, and those reported in column three are direct, not relative, effects. Standard errors clustered at the household level and reported in parentheses. (\*) indicates significant at 15%, (\*\*) 10% or (\*\*\*) 5% level.