Dads, Disease and Death:
Decomposing Daughter Discrimination

Joyce J. Chen
Harvard University

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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. I utilize a household fixed effects model which controls for unobservable characteristics which may be correlated with household type and permits identification of precisely the education gap between sons and daughters within a household. Preliminary results indicate that, controlling for household fixed effects, reductions in the amount of time available for household production reduce the probability that daughters will be enrolled in school, relative to their brothers, whereas reductions in household income 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 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.

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 and careful measurement of their relative magnitudes can have a significant impact on policy and program design. In addition to asking if growth is good for the poor, we might ask if growth is good for women or how policies might make growth more pro-women. Many governments have already undertaken policies with this goal in mind - Progresa 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 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." (Sen, 2000) 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 interventions and economic development more generally.

Allocations to girls and boys may differ for a variety of reasons. The most difficult to address is pure 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 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 have a disproportionate effect on girls when the nature of the shock requires some reallocation of time spent in household activities.

Critical investments in human capital are generally made early on by parents; thus it is appropriate to focus our attention on the household as the unit of observation. 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 actual sources of discrimination, *i.e.* whether differential treatment is the result of standard maximizing behavior on the part of households or a more deeply entrenched notion of gender bias. The relative magnitude and independent effect of each of these factors remains an open question.

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; (2) fathers with a temporary debilitating illness; and (3) widowed mothers. 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 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 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 point to the conclusion that differences in rates of return 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. Yet 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. (2004b) examine directly the impact of parental death on child health and education. Death of the father increases the probability of leaving school, and death of the mother decreases the probability of entering school for both sons and daughters, controlling for community fixed effects. 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. These findings suggest that parental loss indeed has adverse effects, but changes in income play a relatively small role in the reductions in child health and education. 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 is a shock to the aggregate household time endowment as well as income, and the remaining parent may have different preferences for children's schooling. The analysis also does not explore the impact of parental death on gender gaps.

It is important to note that the nature of the shock experienced by the household matters. Health shocks affect the production function for health and thus the time inputs required of other household members. In the face of missing markets, the household that is both producer and consumer of the same good(s) finds its profit and utility maximization problems inextricably linked. 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. 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.

Alternatively, children may contribute income directly to the household through wage labor or work in self-employment activities, *e.g.* the household farm or "cottage industry". Pitt and Khandker (1998) find that women's participation in microcredit programs 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. Conversely, higher income may allow the household to purchase market substitutes for certain home-produced goods (*e.g.* prepared foods, nannies), and higher profitability or earnings potential may reduce the demand for children's time in income generating activities. The relationship between household income and the opportunity cost of children's schooling is ambiguous, and the non-separable nature of household production precisely implies that the roles of income and time allocation cannot be disentangled in these studies.

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. 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.

These results suggest that men's and women's preferences for allocations to children differ, but it is less clear that there is a gender component to these preferences. If women prefer to invest more in children generally, positive income shocks may disproportionately favor girls simply because equality is a normal good, and not as a result of women's direct preferences for daughters relative to sons. Unearned income is also likely to reflect past consumption decisions which, in turn, are indicative of heterogeneous preferences; these preferences may also extend to investments in children and equality within the household.

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 have a smaller impact on children's time allocation.

In summary, the existing literature provides some evidence that income, time allocation and preferences all affect gender discrimination 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.[1,2]

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 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

$$U = \lambda_m[U_m(\boldsymbol{x}_m,\boldsymbol{z},h_m,h_f,s_b,s_g,e_b,e_g;\theta_m)] + \lambda_f[U_f(\boldsymbol{x}_f,\boldsymbol{z},h_m,h_f,s_b,s_g,e_b,e_g;\theta_f)] \qquad \lambda_m + \lambda_f = 1$$

where $\boldsymbol{x}_k$ denotes a vector of market goods specific to the individual; $\boldsymbol{z}$ denotes a vector of shared household goods; h_m and h_f denote health of the respective adult decision-makers; h_b and h_g denote health of sons and daughters; s_b and s_g denote human capital of sons and daughters[3]; e_b, e_g denote future earnings of children, and θ_m, θ_f denote preferences. 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.[4]

The parameters λ_m and λ_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 $\boldsymbol{\omega}_k$.

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

If preferences are identical, the utility function reduces to a unitary model of the household. Note that one component of the individual's outside option is potential or expected future

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[1] 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.

[2] 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 must be made jointly as they can be easily monitored.

[3] This model could easily be extended to include a vector of allocations to children, including health.

[4] 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.

income.[5] Thus a transitory reduction in income, as in the case of temporary debilitating illness, will have a smaller effect on the λ-weights than a permanent change in earnings potential, as in the case of death. For a purely transitory shock to income, the λ-weights should be unaffected. The goods $\mathbf{z}$, 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.

$$\mathbf{z} = \mathbf{z}(t_{qm},t_{qf},t_{qb},t_{qg},\mathbf{x}_z)$$
$$h_k = h(t_{hk,m},t_{hk,f},t_{hk,b},t_{hk,g},x_{hk};\gamma_k,\upsilon_k) \quad \text{for } t_{hk,k}>0 \text{ and } k = m, f$$
$$s_i = s(t_{si,m},t_{si,f},t_{si,i},x_{si};\gamma_i,\upsilon_i) \quad \text{for } t_{si,,i}>0 \text{ and } i = b, g$$

The production functions for health and human capital depend on the individual's characteristics (γ_k or γ_i) and good-specific endowments (υ_k or υ_i), where each individual endowment is a composite of the household-level endowment (υ) and person-specific heterogeneity[6]. With regard to health, the elasticity of substitution between time inputs of different types of household members is assumed to be less than one but greater than the elasticity of substitution between own and others' time. With regard to children's human capital, the elasticity of substitution between fathers' and mothers' time is assumed to be less than the elasticity of substitution between the child's time and parents' time, but I do not rule out the possibility that mothers' and fathers' time are perfect substitutes for each other. For both health and children's human capital, it is also assumed that market substitutes are imperfect, with an elasticity of substitution less than that between household members' time inputs. Finally, I assume that daughters are better substitutes for mothers than are sons and that sons are better substitutes for fathers than are daughters.

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

$$w_m\Sigma t_{qm} + w_f\Sigma t_{qf} + w_b\Sigma t_{qb} + w_g\Sigma t_{qg} - p\cdot x_m - p\cdot x_f - \Sigma p_q\cdot q = w_mT_m + w_fT_f + w_bT_b + w_gT_g + \sigma_be_b + \sigma_ge_g + v$$
$$\text{for } q\in\{\mathbf{z},h_m,h_f, s_b,s_g\}$$

where T_k, T_i are the time endowments[7], and v denotes wealth. w_k denotes the (shadow) price of time for adults, which depends on endowments and past production of human capital.

$$w_k = r_ky_k \qquad \text{where } y_k = y(s_k;\gamma_k,\upsilon_k) \qquad \text{for } k = m, f$$

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

[5] 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 *m* in household *j* have preferences identical to individual *m*, at least with regard to allocations within household *j*.

[6] For adults, υ represents the degree of assortative mating; for children, υ represents heritability.

[7] $T = \Sigma t_q + t_w$ where t_w is time spent in income-generating activities.

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

where r_k, r_i are gender-specific market returns to skill units. σ_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.

Baseline: Father Present and Contributing Income

This group will be represented by married parents who are both currently residing in the same household without any significant periods of absence. In these households, decisions regarding children's schooling may be made jointly by both parents, according to their respective bargaining weights ($\lambda_m \geq 0$ and $\lambda_f \geq 0$). 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 $q \in \{\mathbf{z}, h_m, h_f, s_b, s_g\}$).

Case 2: Father Not Present but Contributing Income

This group will be comprised of children with fathers who are long-term migrants[8]. These fathers are still actively contributing income ($T_m - \Sigma t_q > 0$) 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. The difference in children's schooling outcomes between this group and the baseline group can be attributed to the change time allocation induced by the withdrawal of fathers' time available for 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 impose the further limitation that this shock does not affect bargaining power within the household, the illness must be temporary.[9] In that case, as in the baseline case, both parents may have some influence on schooling decisions. This is analogous to the case presented above, with two exceptions: (1) fathers' earned income is negatively affected, and (2) instead of reducing

[8] Away from the household for more than six months of the preceding year.

[9] Illness may also lead to a decline in ω_m; identification requires only that the change in λ_m relative to λ_f be less than in the case of death.

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 interhousehold 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. Mothers may allocate more time to income-generating activities, thereby requiring daughters to increase their time in household activities, or daughters may be pulled into income-generating activities directly. Existing literature (Rose, 1999; Pitt and Rosenzweig, 1990) suggests that, in households experiencing a shock to income, allocations to daughters will fall disproportionately compared to their brothers. However, to the extent that sons are also required to substitute for fathers in household and/or work activities, the boy-girl differential will be attenuated. In this respect, this sample will only permit the identification of an uncompensated income effect, *i.e.* it will not net out the substitution effect induced by changes in children's time allocation between schooling and the household. The difference in children's schooling outcomes between this group and the previous group will provide a lower bound estimate of the compensated income effect.[10]

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, but there is an additional effect on money income which is equivalent to that in Case 2.[11,12] 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 ω_m and shifts decision-making power to the mother ($\lambda_f = 1$ and $\lambda_m = 0$).[13] Existing literature suggests that this will lead to an increase in allocations to children, especially daughters. This will offset the negative effect due to the reduction in household income and time available for household production. The extent to which this occurs, *i.e.* the difference in children's schooling outcomes between this group and the previous group, will provide a lower bound estimate of the extent to which control over income matters in girls' and boys' schooling outcomes.

The implications of relaxing the above assumptions will be discussed further in Sections III, IV and V.

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[10] Assuming that the reduction in fathers' time spent in household production is sufficiently large, relative to the increase in other household members' time spent in the production of his health.

[11] 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.

[12] 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.

[13] Identification requires only that $\lambda'_m < \lambda_m$ and $\lambda'_f > \lambda_f$, where λ'_f represents the mother's bargaining power after the death of the husband, and λ'_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.

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.[14] 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.[15] 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 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.

Percent of Children Living with Mother

Age	Overall	Boys	Girls
10	0.908	0.910	0.905
11	0.899	0.901	0.898
12	0.884	0.887	0.881
13	0.862	0.869	0.855
14	0.844	0.854	0.833
15	0.809	0.826	0.790
16	**0.772**	**0.802**	**0.741**
17	0.722	0.766	0.676
18	0.670	0.739	0.603
19	0.628	0.717	0.544
20	0.530	0.648	0.428

Percent of Children Attending School

	With Mom		Not With Mom	
Age	Boys	Girls	Boys	Girls
10	0.953	0.954	0.907	0.913
11	0.940	0.941	0.880	0.889
12	0.887	0.876	0.815	0.802
13	0.791	0.761	0.745	0.693
14	0.687	0.651	0.651	0.586
15	0.586	0.566	0.580	0.483
16	**0.518**	**0.497**	**0.564**	**0.448**
17	0.434	0.433	0.498	0.346
18	0.375	0.367	0.449	0.248
19	0.286	0.255	0.343	0.161
20	0.163	0.147	0.189	0.073

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. 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

[14] 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.

[15] More cumulative measures of schooling, such as years of school attended, are not relevant because identification relies on transitory shocks.

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 1 and 2.

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. 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 higher endowments likely have higher earnings potential and thus higher 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. 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 appear to be censored at 30 days. The outcomes of interest involve long-term investments, and thus 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 gain bargaining power, and this may attenuate the son-daughter differential.

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. In this manner, 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.

Schooling Level	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
N	4716	107	146

Certain occupations are associated with higher risk of death, particularly those with low wages or hazardous working conditions. These types of occupations may be more appealing to lower endowment individuals, or may directly affect the health endowment over time by constraining the ability to purchase health inputs. Lower average schooling attainment among widows, 3.5 years versus 5.4 years for currently married women, 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 overestimated, 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

From the model outlined in Section II, the reduced-form linearized demand equation for human capital of sons and daughters in household *j*, conditional on the father's state (co-resident, long-term migrant, temporarily ill or deceased), can be expressed as

$$S_{bj} = \alpha_b + \beta_b \mathbf{P}_{bj} + \delta_b \mathbf{H}_j + \pi\mu_j + \varepsilon_{bj}$$
$$S_{gj} = \alpha_g + \beta_g \mathbf{P}_{gj} + \delta_g \mathbf{H}_j + \pi\mu_j + \varepsilon_{gj}$$

where $\mathbf{P}_{ij}$ 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. μ_j is a household-level unobservable, and ε_{ij} is an i.i.d. error term. Parameters on household-level variables are allowed to vary by gender, and all parameters are allowed to vary by household type.

To minimize the heterogeneity bias caused by μ_j, a household fixed effects model is employed. Taking account of the household fixed effect, the exogeneity condition is simply that household type, as represented by the father's state - co-resident, migrant, temporarily ill or deceased - be 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 μ_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.[16] 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. However, the common household-level endowment affects the quantity of parents' skill units and thus affects wage rates. With the inclusion of fixed effects, parents' earnings potential need only be exogenous to the boy-girl differential within a household. This restriction also implies that changes in income induced by illness or death must be exogenous with respect to the boy-girl differential, but not necessarily with respect to the level of schooling. Changes in income are not explicitly measured, and thus the estimated effect of fathers' illness or death on the educational gender gap will be attenuated. It should be noted that, when gender equality within the household is a normal good, this restriction may be problematic. The differential effect on daughters will be biased downward for households that experience a negative shock to income, *i.e.* households in which the father is deceased or temporarily ill, if the negative shock is more likely to occur in households with low endowments. Given that death is the most severe outcome of illness, it is expected that the magnitude of this bias will be larger for 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. However, individuals usually choose to migrate when better work opportunities are available elsewhere; long-term migrants are likely to have been successful in securing more lucrative employment than that available locally, which may be indicative of higher endowments, on average. This may generate an upward bias of the differential effect on daughters in these households and may lead to some ambiguity in the interpretation of regression coefficients - the estimated differential effect on daughters could be the result of girls being worse (better) off or boys being better (worse) off within the household. Whether sons and daughters are differentially affected by the withdrawal of fathers' time available for household production will depend on the production functions for these goods. Results presented below will provide information about changes in time allocation induced by this shock; from this, the relative elasticities of substitution can be inferred. It should be noted, however, that because migrant fathers cannot contribute time to household activities and the production of household goods, the reduction in household *full* income may not be fully offset by any gains in monetary income due to long-term migration.

[16] The market rates of return to skill units are not directly observable, thus they will also be subsumed in the household fixed effect. This should not be problematic, as these rates of return are exogenous to the household.

Parental preferences for children's education will also be captured in the household fixed effect, provided that these preferences 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 structure. The estimation strategy, however, imposes the restriction that these unobservable gender preferences do not vary within a given household type. Estimated coefficients for the girl dummy variables will also capture the differences in the returns to human capital across males and females. Simple regressions of household type on the sex composition of children living at home, controlling for mother's age and education, suggest that there is no relationship between the sex ratio of children and the probability of migration, temporary illness, or death of the father, at least with respect to child health and fostering (see Table A1).

The net effect of parental preferences can be inferred from a comparison of these parameter estimates 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 ill will experience a negative shock to both income and time. There is no discernible[17] 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:

$$\alpha_m - \alpha_b \leq \text{effect of changes in time allocation}$$
$$\alpha_s - \alpha_m \geq \text{effect of changes in income}$$
$$\alpha_w - \alpha_s \leq \text{effect of changes in bargaining power}$$

where b denotes the baseline parameters; m denotes estimated parameters for children with migrant fathers; s denotes estimated parameters for children with ill fathers, and w denotes estimated parameters for children of widowed mothers [18].

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. Estimated parameters for children with migrant fathers reflect the fact these fathers are constrained in the amount of time they can devote to household activities. If migrants have higher income than the baseline and gender equality is a normal good, the effect of changes in time allocation will be underestimated (*i.e.*, algebraically less than the true effect). If wives of migrants gain decision-making power while their husbands are away and women have stronger preferences for gender equality, the effect of time allocation will again be underestimated.

Temporary illness imposes a similar constraint on the amount of time fathers can devote to household production, but also requires an increase in other household members' time spent in the production of fathers' health. Thus, the difference between α_s and α_m will overestimate of

[17] It is assumed that any change in bargaining power that might be induced by one parent's temporary illness is not sufficient to affect children's schooling, given that schooling requires long-term investments.
[18] $\alpha_k < 0$ for $k = b,m,s,w$ indicates that girls are worse off relative to their male siblings.

the effect of changes in income on children's schooling. However, if there is some complementarity between fathers' time spent in production of own health and household goods **z**, the change in time available for household production may be less than that for migrant households, which would lead to downward bias in the estimated effect of income on the gender gap. Alternatively, if ill fathers are still able to engage in labor market activities, the loss of income will be less than if the father were totally debilitated, and the effect of income will be biased upward, given that the decomposition seeks to identify the effect of a *complete* loss of father's income. Results from the Tobit regression of mothers' hours of work indicate that wives in these households work considerably *fewer* hours than their counterparts with migrant husbands[19]. Thus, the possible complementarities between fathers' production of own health and household goods do not appear sufficient to offset the increased demand for adult women's time in the production of men's health, and the difference between α_s and α_m will be an overestimate (*i.e.* algebraically more than the true effect) of the effect of income.

By netting out the effects of the reduction in income *and* the withdrawal of fathers' time in household production, we can infer the true effect of mothers' preferences, given that intrahousehold allocations more closely reflect mothers' preferences after the death of the father. Because mothers and children do not need to allocate time to the production of fathers' health after death, the difference between α_w and α_s will underestimate of the effect of changes in bargaining power on children's schooling. If temporary debilitating illness is indicative of chronic morbidity which in turn reduces fathers' bargaining power, the bargaining power effect will again be underestimated. Conversely, if ill fathers are still able to contribute productive time to household activities and wage labor, the difference between α_w and α_s will again understate the effect of women's preferences as long as equality within the household is a normal good.

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 3[20]; coefficients in columns 2 through 4 should be interpreted relative to the baseline (un-interacted) coefficients. As expected, widows work considerably more hours, and the age profile is flatter and less concave. Descriptive statistics indicate that these households tend to have more male household members in the 17-35 age range; older widows may also have older sons who remain in the household to contribute wage income. The negative coefficients on wealth (floor area of owned home) and males aged 17 to 35 is consistent with this. The number of children living at home has a large positive effect, but male and female children between the ages of five and sixteen have significant negative effects on widowed mothers' hours of paid work. This pattern suggests that resource constraints are tight in these households; more "mouths to feed" leads to more time in the labor market, while young children require more time spent in household activities. Conversely, females aged 17 and older may be available to provide time in household care, which is consistent with the negative coefficient on each. Years of schooling and the (imputed)

[19] The 95% confidence intervals for predicted hours of work indicate that women with ill husbands work 3 to 10 hours less than women with migrant husbands and 2.5 to 11 hours less than the baseline sample.
[20] Results from a logit regression of mothers' labor market participation are reported in Table A2.

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. Predicted values at the sample-specific mean indicate that mothers in this sample work 20 more hours, relative to the baseline. Widowed mothers cannot spend as much time in household, which implies that sons and/or daughters must be pulled in to household activities that compete with schooling.

Women with husbands who are temporarily ill work fewer hours, consistent with the theoretical implication that they are required to spend more time caring for the household and their sick husbands. Children younger than five have a negative effect on work hours, which again suggests that these young children require a significant amount of mothers' time in the home. Males aged 36 to 54 and the husband's years of schooling are negatively related to work hours; both of these factors may mitigate the loss of income due to illness. The length of the husband's illness, however, has a significant positive effect - the more severe the shock to income, the more mothers are pulled into wage labor. Urban residence has a positive effect, also found for widowed women, 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. Predicted values at the sample-specific mean indicate that mothers in this sample work 7-8 fewer hours, relative to the baseline. When these women do not or cannot enter the labor market in response to husband's illness, the shock to household income is even 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.

Wives of long-term migrants work 1-2 fewer hours relative to the baseline but work 6-7 more hours than women with ill husbands. This is consistent with (1) a higher demand for adult women's time in the production of household goods and (2) a lower demand for adult women's time in the production of men's health. Taken together, these results suggest that a reduction in the amount of time available for household production increases demand for women's time in household production, but health shocks lead to a substantially larger increase in the demand for women's time in the home.

Current School Enrollment

The probability of being enrolled in school at the time of enumeration[21] is estimated with a conditional logit to account for household fixed effects. Results are summarized in Table 4. Specification I includes a quadratic in the child's age, age and its square interacted with sex, and household level covariates (wealth, mother's characteristics, and father's characteristics if

[21] It is not clear whether school enrollment refers to current attendance and, because the average length of illness is roughly 8 days, this measure may not reflect true changes in school-going. An alternative measure can be derived from the child's "primary activity" in the last week, but some children who do not report school as their primary activity may still be attending nonetheless. Estimates using "school as primary activity" as the dependent variable are presented in Table A4. Parameter estimates are not substantially different from those presented in Table 4 and are more difficult to interpret since they conflate changes in school attendance with changes in labor supply.

present) interacted with the child's sex, and specification II adds the length of widowhood or illness. Specification III allows all household-level covariates to vary by household type and sex of the child and is the preferred specification. A joint test of significance for the covariates added in Specification IV cannot reject the null that the parameters are jointly equal to zero for all household types. 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 activities in the previous one week.

The full set of parameter estimates for Specification III is reported in Table A3. Baseline parameters indicate that observable household characteristics do not differentially affect daughters, with the exception of mothers' years of schooling, which 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 is innocuous. The coefficient on the dummy variable for girls is positive and significant, but the probability of being enrolled in school declines rapidly for girls as they age.

There are no statistically significant differential effects for daughters with widowed mothers. The age profile of school enrollment is convex for children in households with either deceased or temporarily ill fathers, which indicates that the probability of attending school is lower for all children in the relevant age range, with the effect being more pronounced for younger children. If interruptions to school attendance are more detrimental at higher levels of schooling, households facing an adverse shock may be more reluctant to pull older children out school. The large and significant negative effect on daughters in households that experience a temporary illness is the most striking result in this regression and is consistent with the existing literature. For a shock which increases the demand for time in household production and/or decreases household income, daughters suffer disproportionately. 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. Note that the concavity in the effect of father's age implies that, even at the maximum, this does not offset the negative level effect for daughters.

Table 5 displays the odds of school enrollment for girls relative to boys, by age and household type. Relative to the baseline, daughters in widowed households are more likely to be enrolled in school than their male siblings at all ages except sixteen, which suggest that the increasing women's bargaining power may have a large impact on gender discrimination. Girls are relatively worse off in migrant households, which is consistent with increased demand for these daughters' time in household production. Gender inequality appears most pronounced for the youngest and oldest children in households experiencing a temporary debilitating illness. In the 12-14 age range, daughters in these households are actually more likely to be in school than brothers, and the gender differential is smaller than in the baseline case.

Comparing estimates across these samples will permit a decomposition of the sources of discrimination, and the independent effects of changes in time allocation, income and bargaining can identified. Results in Table 6 confirm that the elimination of fathers' time available for household production increases gender discrimination in school enrollment for younger children, and perhaps a slight decrease in discrimination among older children. This suggests that, at younger ages, $\xi_{fm} > \xi_{bm}$ and $\xi_{gf} > \xi_{bf}$ whereas, at older ages, $\xi_{bm} > \xi_{fm}$ where ξ_{ik} is the elasticity of substitution of person type *i*'s time for person type *k*'s time, *m* represents adult males, *f* represents adult females, *b* represents boy children and *g* represents girl children. That is, mothers are better substitutes for fathers in household production when children are young, and daughters in turn are better substitutes for mothers than are sons, but older sons are better substitutes for fathers than mothers and daughters.

Reducing fathers' contribution to household income increases discrimination in school enrollment for younger and older daughters, but may decrease discrimination in the middle age range. Because this is an overestimate of the true effect of eliminating fathers' income-generating ability, it is possible that the income effect (weakly) increases discrimination at all ages, as suggested by the existing literature. The potential decrease in discrimination at ages 11-14 should be interpreted as a disproportionate negative effect on sons, rather than a differential positive effect on daughters. This counterintuitive result suggests that sons in the middle age group spend more time caring for ill fathers than do their female siblings of the same age, *i.e.* these sons are more productive than daughters with regard to the production of adult male health ($\xi_{bf} > \xi_{gf}$). Finally, the decompositions indicate that increasing women's bargaining power decreases the gender gap in school enrollment, and the effect is more pronounced for the youngest and oldest girls. Mothers appear to have a strong preference for gender equality, and perhaps even a slight preference for daughters.

Years of Schooling

To verify the assumptions and results presented above, I examine the gender gap in years of schooling. 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, however, induces permanent changes in income, time allocation and bargaining power. Table A4 reveals that the age profile is steeper and more concave for children in widowed households, with years of schooling being higher for all children of widowed mothers in the 10 to 16 age range, all else equal. Interestingly, the baseline differential effect for daughters is not statistically significant, which suggests that daughters may experience more school interruption and/or grade repetition, given that daughters are more likely to be enrolled in school (see Tables 4 and A3). There are no significant differential effects for daughters' years of schooling in households where 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. This is consistent with the earlier assertion that the difference between α_m and α_b underestimates the true effect of changes in time allocation on the educational gender gap. However, this finding may also suggest that migrant households have higher endowments, higher earnings potential, and thus less 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. By symmetry, these results suggest that there are no differential effects on sons and thus the absence of fathers is not affecting the long-term accumulation of sons' human capital relative to daughters.

VI. Conclusion

This paper seeks to disentangle the various factors contributing to gender discrimination in investment in children's human capital. The results provide a careful accounting of the role of preferences, income and time allocation in the educational gender gap. Increasing women's bargaining power will have the largest effect on reducing discrimination, but policymakers must be aware that the induced changes in time allocation and/or income may work in the opposite direction. Future work will utilize panel data to extend this methodology to children's nutrition and non-cooperative bargaining among spouses.

Table 1. Descriptive Statistics for Children[a]

Sons

	Married	Widow	Sick Dad	Migrant
Age	12.76	13.31 ***	12.77	12.78
	(1.97)	(1.97)	(1.99)	(1.97)
Years of Schooling	66.39%	64.91% ***	64.65% ***	67.19% ***
(% of Potential)	(0.19)	(0.21)	(0.19)	(0.19)
Enrolled in School	80.55%	65.55% ***	79.53% **	79.33% ***
	(0.40)	(0.48)	(0.40)	(0.40)
School as Primary Activity	79.50%	64.75% ***	78.03% ***	78.29% ***
	(0.40)	(0.48)	(0.41)	(0.41)
Work for Pay	16.90%	26.30% ***	21.52% ***	15.18% ***
	(0.37)	(0.44)	(0.41)	(0.36)
Work in Addition to School	5.43%	4.45% ***	8.73% ***	3.64% ***
	(0.23)	(0.21)	(0.28)	(0.19)
Number of Observations	104,189	34,527	6,483	10,741

Daughters

	Married	Widow	Sick Dad	Migrant
Age	12.73	13.27 ***	12.79 **	12.70
	(1.96)	(1.96)	(1.97)	(1.97)
Years of Schooling	67.76%	66.22% ***	66.31% ***	68.31% ***
(% of Potential)	(0.19)	(0.21)	(0.18)	(0.20)
Enrolled in School	80.11%	65.71% ***	79.30%	78.99% ***
	(0.40)	(0.47)	(0.41)	(0.41)
School as Primary Activity	79.16%	64.89% ***	78.04% ***	77.74% ***
	(0.41)	(0.48)	(0.41)	(0.42)
Work for Pay	11.40%	18.05% ***	13.87% ***	10.35% ***
	(0.32)	(0.38)	(0.35)	(0.30)
Work in Addition to School	3.75%	3.13% ***	4.88% ***	2.75% ***
	(0.19)	(0.17)	(0.22)	(0.16)
Number of Observations	97,512	31,158	6,043	9,943

Significantly different from column 1 at the 1% (***), 5% (**), or 10% (*) level.

[a] Limited to unmarried children between the ages of five and sixteen, co-residing with mothers.

Table 2. Descriptive Statistics for Households

	Married	Widow	Sick Dad	Migrant
Children at Home	3.58	2.98 ***	3.15 ***	3.21 ***
	(1.66)	(1.51)	(1.53)	(1.58)
Sons at Home/	0.52	0.54 ***	0.52	0.52
Children at Home	(0.30)	(0.33)	(0.32)	(0.32)
Mom's Years of Schooling	5.40	3.51 ***	5.20 ***	5.39
	(3.79)	(3.71)	(3.76)	(3.80)
Floor Area of Owned Home	58.64	54.01 ***	59.80 **	54.21 ***
	(49.80)	(48.10)	(45.32)	(52.22)
Mom's Age	36.19	44.50 ***	37.11 ***	35.69 ***
	(7.79)	(9.23)	(8.53)	(8.56)
Mom's Age at First Marriage	18.37	18.50 ***	18.62 ***	18.30 *
	(4.74)	(5.41)	(3.75)	(5.48)
Urban Residence	28.28%	24.93% ***	28.68%	28.86% *
	(0.45)	(0.43)	(0.45)	(0.45)
Mother Works for Pay	46.21%	71.37% ***	54.00% ***	43.51% ***
	(0.50)	(0.45)	(0.50)	(0.50)
Mom's Hours of Paid Work	33.05	36.62 ***	32.55 ***	33.35 **
	(15.92)	(15.93)	(16.76)	(15.99)
Max Length of Widowhood		10.13		
		(4.26)		
Dad's Years of Schooling	6.38		6.16 ***	
	(3.89)		(3.88)	
Dad's Age	41.79		43.17 ***	
	(9.13)		(10.34)	
Dad's Length of Illness			7.57	
			(8.00)	
Number of Observations	260,867	50,058	10,325	18,364

Significantly different from column 1 at the 1% (***), 5% (**), or 10% (*) level.

Table 3. Tobit Regression of Mothers' Hours of Paid Work[b]

		Widow	Sick Dad	Migrant
Intercept	-40.954 ***	55.225 ***	-26.385 **	-17.720 ***
	(2.437)	(4.104)	(10.562)	(5.907)
Age	2.421 ***	-1.561 ***	1.176 **	0.066
	(0.097)	(0.192)	(0.518)	(0.297)
Age^2	-0.026 ***	0.011 ***	-0.017 ***	0.000
	(0.001)	(0.002)	(0.006)	(0.003)
Years of Schooling	0.590 ***	-0.488 ***	-0.129	-0.103
	(0.028)	(0.053)	(0.133)	(0.084)
# of Children at Home	-2.535 ***	2.959 ***	0.956	-0.306
	(0.114)	(0.218)	(0.588)	(0.409)
Sons at Home/	0.518 *	-0.907	-0.047	0.798
Children at Home	(0.311)	(0.615)	(1.395)	(1.106)
Floor Area of Home	0.010 ***	-0.025 ***	-0.021 ***	-0.004
	(0.002)	(0.004)	(0.008)	(0.006)
Dummy for Urban	-6.078 ***	9.412 ***	1.579 *	-0.774
	(0.218)	(0.431)	(0.894)	(0.714)
Husband's Age	-0.401 ***		0.129	
	(0.088)		(0.382)	
Husband's Age^2	0.003 ***		-0.001	
	(0.001)		(0.004)	
Husband's Schooling	-0.544 ***		-0.592 ***	
	(0.027)		(0.127)	
Max. Length of Widowhood		-0.142 **		
		(0.056)		
Length of Illness			0.111 **	
			(0.045)	

Number of obs =	339317	LR chi2(384) =	56111.22
left-censored	169748	Log-likelihood =	-943591.2
uncensored	169569	Prob>chi2 =	0.00
		Psuedo R2 =	0.0289

Significant at the 1%(***), 5%(**), or 10%(*) level.

[b] Includes controls for kabupaten-level fixed effects and demographic characteristics of other household members.

Table 4. Conditional Logit Regressions of Children's School Enrollment[c,d]

	I	II	III	IV	V
$\alpha_{baseline}$	2.901	2.916	**5.972** *	5.648	5.332
	(2.250)	(2.253)	**(3.182)**	(3.724)	(3.422)
$\alpha_{migrant}$	-0.094	-0.094	-1.470	0.166	-2.080
	(0.182)	(0.182)	(6.443)	(7.164)	(6.663)
α_{ssick}	-0.095	0.109	**-19.954** **	**-32.261** ***	**-38.633** ***
	(0.158)	(0.218)	**(10.177)**	**(12.376)**	**(10.794)**
α_{widow}	-0.038	-0.123	-3.363	-2.360	-1.844
	(0.199)	(0.250)	(4.833)	(5.392)	(5.094)
$\alpha_{migrant}$ - $\alpha_{baseline}$	-0.094	-0.094	-0.028	-0.453	-1.394
α_{sick} - $\alpha_{migrant}$	-0.001	0.023	0.107	-9.742	-7.156
α_{widow} - α_{sick}	0.057	0.052	0.055	3.579	1.804
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
Additional household characteristics	no	no	no	yes	yes
Alternative measure of attendance	no	no	no	no	yes

Significant at the 1%(***), 5%(**), or 10%(*) level.

[c] Additional household-level characteristics include sibling size, sibling sex ratio, number of household members in twelve age-sex specific categories, mother's age at first marriage, and urban residence.

[d] Alternative measure of school attendance is equal to one when the child reports schooling as his/her primary activity in the previous week.

Table 5. Odds of School Enrollment for Girls Relative to Boys[e]

	Baseline	Widows	Sick Dad	Migrant
10	1.540	1.642	0.966	1.253
11	1.241	1.443	1.150	1.061
12	1.054	1.277	1.220	0.942
13	0.943	1.139	1.151	0.879
14	0.889	1.023	0.967	0.861
15	0.882	0.925	0.724	0.885
16	0.923	0.843	0.482	0.955

	Baseline		Widows		Sick Dad		Migrant	
	Lower	Upper	Lower	Upper	Lower	Upper	Lower	Upper
10	1.300	1.825	1.290	2.090	0.548	1.704	0.892	1.762
11	1.190	1.295	1.322	1.574	0.970	1.365	0.965	1.166
12	1.092	1.018	1.291	1.263	1.320	1.127	0.998	0.889
13	1.009	0.882	1.205	1.076	1.393	0.951	0.992	0.778
14	0.938	0.842	1.077	0.972	1.143	0.819	0.951	0.779
15	0.878	0.886	0.921	0.930	0.728	0.719	0.877	0.893
16	0.826	1.030	0.752	0.946	0.358	0.650	0.775	1.178

Table 6. Decomposition of Sources of Discrimination by Age[f]

	$\alpha_{migrant} - \alpha_{baseline}$	$\alpha_{sick\ dad} - \alpha_{migrant}$	$\alpha_{widow} - \alpha_{sick\ dad}$
10	-0.286	-0.287	0.676
11	-0.181	0.090	0.292
12	-0.112	0.277	0.058
13	-0.064	0.272	-0.012
14	-0.028	0.107	0.055
15	0.003	-0.161	0.202
16	0.033	-0.473	0.361

[e] Based on parameter estimates in Specification III.
[f] Based on parameter estimates in Specification III. Calculated as the difference between samples in the odds of school enrollment for girls relative to boys, as described in Section IV.

Table A1. Logit Regression of Household Type on Child Sex Ratio[g]

	Sick Dad		Migrant		Widow	
Age	**-0.032**	***	**-0.118**	***	**0.089**	***
	(0.007)		**(0.005)**		**(0.011)**	
Age^2	**0.001**	***	**0.001**	***	**-0.001**	***
	(0.000)		**(0.000)**		**(0.000)**	
Mom's Years of Schooling	**-0.008**	***	-0.001		**-0.084**	***
	(0.003)		(0.002)		**(0.003)**	
Sons at Home/Kids at Home	0.026		0.003		0.008	
	(0.034)		(0.025)		(0.037)	
Constant	**-2.823**	***	**-0.300**	***	**-4.685**	***
	(0.143)		**(0.104)**		**(0.215)**	
Number of obs.	271192		279231		269434	
LR chi2(384) =	182.43		480.5		870.24	
Prob>chi2 =	0		0		0	
Log-likelihood	-43779.47		-67486.52		-37563.71	
Psuedo R2	0.0021		0.0035		0.0115	

Significant at the 1%(***), 5%(**), or 10%(*) level.

[g] Includes only widows who have been widowed for five years or less, in order to capture the sex ratio at the time of the change in household type.

Table A2. Mothers' Probability of Working for Pay[h]

		Widow	Sick Dad	Migrant
Intercept	**-2.205 *****	**3.301 *****	**-1.372 ****	**-0.797 ****
	(0.141)	**(0.270)**	**(0.624)**	**(0.348)**
Age	**0.130 *****	**-0.070 *****	**0.066 ****	-0.005
	(0.006)	**(0.013)**	**(0.031)**	(0.018)
Age^2	**-0.001 *****	**0.000 *****	**-0.001 *****	0.000
	(0.000)	**(0.000)**	**(0.000)**	(0.000)
Years of Schooling	**0.029 *****	**-0.033 *****	-0.011	**-0.011 ****
	(0.002)	**(0.004)**	(0.008)	**(0.005)**
# of Children at Home	**-0.132 *****	**0.156 *****	0.020	-0.015
	(0.007)	**(0.014)**	(0.036)	(0.024)
Sons at Home/	0.028	**-0.097 ****	0.027	0.013
Children at Home	(0.018)	**(0.042)**	(0.085)	(0.064)
Floor Area of Home	**0.001 *****	**-0.002 *****	**-0.001 *****	0.000
	(0.000)	**(0.000)**	**(0.001)**	(0.000)
Dummy for Urban	**-0.483 *****	**0.173 *****	0.023	-0.040
	(0.012)	**(0.028)**	(0.052)	(0.041)
Husband's Schooling	**-0.040 *****		**-0.048 *****	
	(0.002)		**(0.008)**	
Husband's Age	**-0.022 *****		0.007	
	(0.005)		(0.023)	
Husband's Age^2	**0.000 *****		0.000	
	(0.000)		(0.000)	
Max. Length of Widowhood		**-0.015 *****		
		(0.004)		
Length of Illness			0.003	
			(0.003)	
	Number of obs =	339361	LR chi2(384) =	63576.66
			Log-likelihood =	0
	Psuedo R2 =	0.1351	Prob>chi2 =	-203438.8

Significantly at the 1% (***), 5% (**), or 10% (*) level.

[h] Includes controls for kabupaten-level fixed effects and demographic characteristics of other household members.

Table A3. Conditional Logit Regression of Children's School Enrollment

		Widow	Sick Dad	Migrant
Dummy for Girl	**5.972** *	-3.363	**-19.954** **	-1.470
	(3.182)	(4.833)	**(10.177)**	(6.443)
Age	0.030	**-1.116** ***	**-1.586** *	-0.269
	(0.268)	**(0.413)**	**(0.878)**	(0.591)
Age^2	**-0.037** ***	**0.043** ***	**0.058** *	0.014
	(0.010)	**(0.016)**	**(0.033)**	(0.022)
Age*Girl	**-0.762** *	0.557	2.157	0.083
	(0.426)	(0.659)	(1.353)	(0.923)
(Age^2)*Girl	0.026	-0.022	-0.084	-0.002
	(0.016)	(0.025)	(0.052)	(0.035)
Floor Area of Home	0.002	-0.002	0.002	0.001
	(0.001)	(0.002)	(0.005)	(0.002)
Mom's Age	-0.058	0.024	-0.203	0.044
	(0.084)	(0.108)	(0.293)	(0.112)
Mom's Age^2	0.001	0.000	0.003	0.000
	(0.001)	(0.001)	(0.003)	(0.001)
Mom's Years of Schooling	**0.045** **	0.001	-0.040	-0.010
	(0.019)	(0.027)	(0.061)	(0.036)
Dad's Age	0.010		**0.402** **	
	(0.066)		**(0.190)**	
Dad's Age^2	0.000		**-0.004** **	
	(0.001)		**(0.002)**	
Dad's Years of Schooling	-0.001		**0.117** **	
	(0.019)		**(0.053)**	
Max. Length of Widowhood		0.025		
		(0.018)		
Length of Illness			**-0.033** *	
			(0.019)	

Number of obs =	46195	LR chi2(45) =	21098.47
Number of groups =	19162	Log-likelihood =	-5970.804
		Prob>chi2 =	0

Significant at the 1%(***), 5%(**), or 10%(*) level.

Table A4. Children’s Years of Schooling[i]

		Widow	Sick Dad	Migrant
Dummy for Girl	0.353	-1.773	-1.375	-2.289
	(0.844)	(1.319)	(2.679)	(1.854)
Age	**1.603** ***	**0.382** ***	0.159	0.152
	(0.060)	**(0.105)**	(0.190)	(0.152)
Age^2	**-0.030** ***	**-0.019** ***	-0.007	-0.007
	(0.002)	**(0.004)**	(0.007)	(0.006)
Age*Girl	0.007	0.241	0.214	0.192
	(0.092)	(0.163)	(0.292)	(0.234)
(Age^2)*Girl	0.000	-0.009	-0.008	-0.008
	(0.004)	(0.006)	(0.011)	(0.009)
Floor Area of Home	0.000	0.000	0.000	-0.001
	(0.000)	(0.000)	(0.001)	(0.001)
Mom's Age	-0.002	-0.009	0.006	0.028
	(0.026)	(0.035)	(0.093)	(0.050)
Mom's Age^2	0.000	0.000	0.000	0.000
	(0.000)	(0.000)	(0.001)	(0.001)
Mom's Years of Schooling	0.007	0.004	0.021	**0.029** ***
	(0.005)	(0.007)	(0.015)	**(0.010)**
Dad's Age	-0.011		-0.018	
	(0.019)		(0.049)	
Dad's Age^2	0.000		0.000	
	(0.000)		(0.001)	
Dad's Years of Schooling	0.000		-0.009	
	(0.005)		(0.014)	
Max. Length of Widowhood		0.000		
		(0.009)		
Length of Illness			-0.001	
			(0.005)	

				R-squared:	
Number of obs =	300394	F(109,79611)=	1301.47	within	0.6405
Number of groups =	220674	Prob>F = 0	0	between	0.335
		corr(u_i, Xb) =	-0.2266	overall	0.3644

Significant at the 1%(***), 5%(**), or 10%(*) level.

[i] Includes controls for sibling size, sibling sex ratio, number of household members in 12 age-sex specific categories, mother’s age at first marriage, and urban residence.

Table A5. Probability of Working for Pay[j]

	Married		Widow		Sick Dad		Migrant
Dummy for Girl	**-9.077**	**	1.762		3.641		5.133
	(4.282)		(6.285)		(11.378)		(10.659)
Age	**1.058**	***	0.183		-0.364		-0.422
	(0.298)		(0.505)		(0.861)		(0.868)
Age^2	-0.003		-0.003		0.013		0.020
	(0.011)		(0.019)		(0.033)		(0.033)
Age*Girl	**1.178**	**	0.058		0.388		-0.281
	(0.499)		(0.809)		(1.367)		(1.408)
(Age^2)*Girl	**-0.045**	**	-0.007		-0.017		0.012
	(0.019)		(0.030)		(0.052)		(0.053)
Floor Area of Home	**-0.004**	***	**0.005**	**	-0.005		0.003
	(0.001)		**(0.002)**		(0.005)		(0.003)
Mom's Age	0.052		-0.063		-0.505		-0.223
	(0.129)		(0.153)		(0.359)		(0.256)
Mom's Age^2	-0.001		0.001		0.005		0.002
	(0.001)		(0.002)		(0.004)		(0.003)
Mom's Years of Schooling	0.003		-0.025		-0.044		0.035
	(0.021)		(0.031)		(0.060)		(0.050)
Dad's Age	-0.026				0.239		
	(0.069)				(0.212)		
Dad's Age^2	0.000				-0.002		
	(0.001)				(0.002)		
Dad's Years of Schooling	**0.038**	*			**0.156**	***	
	(0.021)				**(0.055)**		
Max. Length of Widowhood			-0.014				
			(0.035)				
Length of Illness					-0.005		
					(0.019)		

Number of obs =	35477	LR chi2(109) =	16541.98
Number of groups =	14671	Log-likelihood =	-4404.12
		Prob>chi2 =	0.000

Significant at the 1%(***), 5%(**), or 10%(*) level.

[j] Includes controls for sibling size, sibling sex ratio, number of household members in 12 age-sex specific categories, mother's age at first marriage, and urban residence.

Table A6. Probability of School as Primary Activity[k]

	Married	Widow	Sick Dad	Migrant
Dummy for Girl	5.332	-1.844	**-38.633 *****	-2.080
	(3.422)	(5.094)	**(10.794)**	(6.663)
Age	0.221	**-1.071 *****	**-1.368 ***	-0.308
	(0.246)	**(0.392)**	**(0.799)**	(0.563)
Age^2	**-0.042 *****	**0.040 *****	0.050	0.014
	(0.009)	**(0.015)**	(0.030)	(0.021)
Age*Girl	**-0.670 ***	0.231	**2.448 ****	0.098
	(0.389)	(0.625)	**(1.233)**	(0.868)
(Age^2)*Girl	0.023	-0.010	**-0.094 ****	-0.003
	(0.015)	(0.024)	**(0.047)**	(0.033)
Floor Area of Home	0.001	0.000	0.000	0.001
	(0.001)	(0.002)	(0.005)	(0.002)
Mom's Age	-0.029	0.020	0.391	0.064
	(0.108)	(0.141)	(0.316)	(0.167)
Mom's Age^2	0.000	0.000	-0.004	-0.001
	(0.001)	(0.002)	(0.004)	(0.002)
Mom's Years of Schooling	**0.041 ****	0.017	-0.049	0.005
	(0.018)	(0.026)	(0.059)	(0.034)
Dad's Age	-0.019		**0.561 *****	
	(0.075)		**(0.218)**	
Dad's Age^2	0.000		**-0.006 *****	
	(0.001)		**(0.002)**	
Dad's Years of Schooling	-0.009		**0.096 ***	
	(0.018)		**(0.053)**	
Max. Length of Widowhood		**0.073 *****		
		(0.028)		
Length of Illness			**-0.034 ***	
			(0.017)	

Number of obs =	47262	LR chi2(109) =	20665.96
Number of groups =	19594	Log-likelihood =	-6568.654
		Prob>chi2 =	0.000

Significant at the 1%(***), 5%(**), or 10%(*) level.

[k] Includes controls for sibling size, sibling sex ratio, number of household members in 12 age-sex specific categories, mother's age at first marriage, and urban residence.

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