Marriage Market, Parents' Bargaining Powers, and Children's Education

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Abstract

This paper suggests that the marriage market conditions, summarized by the sex ratio, can be used to test the unitary and the non-unitary household models and that the test may be combined with the income pooling test to produce more reliable inferences. Using data from Indonesia, this paper conducts the combined test by estimating the effects of the provincial sex ratio and the parents' nonlabor incomes/premarital assets on Indonesian household's investment in children's education. I find that in urban areas the sex ratio has a strong positive effect on education expenditures, but not in rural areas. Being consistent to the sex ratio effect estimates, the income pooling hypothesis is mostly rejected in urban areas, but not in rural areas. In addition, premarital assets are found to have significant effects on investment in children's education in urban areas, but not in rural areas. I find that the estimation results are robust against alternative definitions of the sex ratio and additional controls for women's fertility choices and community-level income/wealth differences.

1 Introduction

Traditionally a family is regarded as a single decision-making unit in the economy. The traditional unitary household model abstracts away the differences among the family members and treats a household as if it is a collection of identical individuals or it is governed by the dictatorial householder who optimizes, subject to the 'pooled' budget constraint. More recently, however, models that preserve intra-family differences are advanced and appear to be gaining popularity. Such models treat a household as a collection of individuals with divergent preferences and individual resources and view intrahousehold resource allocations as bargaining outcomes (McElroy and Horney 1981), or as Pareto-efficient outcomes reached through some (unspecified) collective decision-making processes (Chiappori 1988).

The unitary and the non-unitary household models have vastly different implications on social policies, in particular on effects of transfer programs. The unitary model suggests that, controlling for the total amount of transfers, the identity of the beneficiary should have no effect on household consumptions, because all the household members' resources are pooled. The non-unitary models suggest otherwise, because the change of the recipient's individual resources will shift the household members' bargaining powers, and subsequently the household consumption decisions. For example, the non-unitary models suggest that welfare benefits have different effects on children's welfare depending on which parent receives them, while the unitary model suggest that as long as the amount itself does not change, who receives them should not make any difference[1]. Furthermore, according to the non-unitary models, policies that change the social environment where the value of the household members' outside options is determined—for example, divorce laws—may change the intrahousehold resource allocations (Chiappori et al. 2002).

Considering the theoretical as well as the practical importance of the issue, one may find only fitting that there are many studies that test validity of the competing household models. Although their data sets and variables of interest differ, most of the early empirical studies focus on testing whether nonlabor incomes of household members are pooled for household consumptions. For example, Thomas (1990, 1994) tests whether in Brazil the mother's and the father's nonlabor incomes have different impacts on children's health and nutritional intake. Schultz (1990) tests whether in Thailand unearned income and transfers received by the husband and the wife have different effects on female labor supply and fertility. Duflo (2000) tests whether in South Africa pensions received by women and men have different effects on children's height and weight.

As pointed out elsewhere (Thomas et al. 2002), however, in many cases it is difficult to defend exogeneity of the nonlabor income variables, the critical assumption which validity of the income pooling tests rests on. Pensions, for example, are likely to be the outcome of the intertemporal household resource allocation decisions. Despite that the problem is well known, it is difficult to solve it because the incidence of purely exogenous nonlabor incomes—for example, lottery winnings—is extremely rare and the proper instruments—some variables which determine the household members' nonlabor incomes but not the household resource allocation decisions—seem to be nonexistent in most data sets. Facing the problem in using the nonlabor income variables, several studies suggest that the value of premarital assets may be used instead of the amount of current nonlabor incomes, assuming that it is exogenous to postmarital household decisions. For example, Thomas et al. (2002) test whether in Indonesia resources brought to marriage by the husband and the wife have any effect on the children's health, controlling for the current economic status. Quisumbing and Maluccio (2003) estimate the effects of the husband's and the wife's assets at marriage on the shares of the food, health, education, child clothing, and tobacco expenditures in the total household expenditures using data from Bangladesh, Ethiopia, Indonesia, and South Africa. It should be noted that inferences based on this alternative are valid only if the size of the assets brought to the marriage is not determined through a premarital bargaining between the bride and the groom or if the postmarital bargainings are uncorrelated with the premarital bargainings.

Though most widely used, the income pooling hypothesis is not the only testable implication that can be used to test the unitary household model against the non-unitary models. The non-unitary household models suggest that any factor outside the household that shifts individual household members' household bargaining powers or Pareto weights in the household—called "extrahousehold environmental parameters (EEPs)" by McElroy and Horney (1981) and McElroy (1990) and "distribution factors" by Chiappori et al. (2002)—such as the sex ratio in the marriage market, should change the intrahousehold resource allocation in a predictable way. The unitary model predicts that they should have no effect at all (McElroy 1990, Chiappori et al. 2002).

There are several empirical studies on intrahousehold allocations that exploit the implication and test whether the sex ratio has any impact on household consumptions. As far as I know, all the studies use data from the US and focus on women's labor supply[2]. Grossbard-Shechtman (1993), using data from the US Census in 1930 and in 1980, finds that the city sex ratio is negatively correlated with the labor force participation rate of married women. Grossbard-Shechtman and Neideffer (1997), using a microsample from the 1990 Census, find that the sex ratio in metropolitan areas is negatively correlated with the likelihood of local women's being in the labor force or with local women's working hours. Chiappori et al. (2002) find that, using data from the Panel Study of Income Dynamics (PSID) of 1988 and the Census of 1990, the state sex ratio is negatively correlated with married women's labor supply but positively correlated with married men's labor supply. Angrist (2002) studies the impact of the immigration flow into America on the marriage market of the second generation in early twentieth century and finds that the increased sex ratio reduces female labor force participation and tilts the balance of household bargaining power toward women. These studies provide evidence that the local sex ratio, whose increase should contribute to strengthening wives' bargaining power, has a significant effect on the intrahousehold resource allocation to the expected direction. The linkage between the sex ratio and household decisions, however, has not been studied so far in other areas than labor supply.

The two approaches to the research on intrahousehold resource allocations have been taken separately so far. The theory, however, does not suggest that the two tests—the income pooling and the EEP effect tests—should be done separately. It does suggest, to the contrary, that the two tests be combined to be a more rigorous test of the household models. An important advantage of the combined test over the separate ones is that cross-checks of the two test results for whether they lead to consistent inferences regarding the household behavior will make the conclusions drawn from the combined test more reliable than those from the separate tests.

Adopting the strategy of combining the two tests in this paper, I estimate the effects of the provincial sex ratio and of the mother's and the father's nonlabor incomes on household resources allocated to children's education in Indonesia. In another specification I replace the nonlabor incomes with the parents' premarital assets. Using three different samples, I estimate the effects of the sex ratio, the nonlabor incomes, and the premarital assets on children's school enrollment status and education expenditures on the individual and on the household level, controlling for the child's and the parents' characteristics and the communities' education infrastructure. The results are used to test validity of the household models separately for the urban and the rural area.

The estimation results show that in urban areas the sex ratio has strong positive effects on the households' education expenditures, but not in rural areas. The income pooling is rejected mostly in urban areas, but never rejected in rural areas. The estimation results are found to be robust even if the effects of the sex ratio on mothers' fertility choices and the average household income and asset holdings of communities are taken into account in estimations. It is also found that controlling for the current household income and assets, the parents' premarital assets, when they replace the nonlabor income variables, have significant effects on urban households' investment in children's education, but not on rural households' investment. Note the consistency of the estimation results—in urban areas, they all agree in rejecting the unitary household model, while in rural areas they agree in *not* rejecting the unitary model. I also find evidence that in Indonesia both the mother and the father prefer educating sons to educating daughters.

The balance of the paper is organized as follows. Section 2 describes the unitary and the non-unitary household models and discusses their empirical implications this study is based on. Section 3 describes the data and how some key variables are constructed. Then Section 4 presents and discusses the estimation results. The conclusion of this paper is provided in Section 5.

2 The Models of Intrahousehold Resource Allocation and Children's Education

Think of a household that consists of the father, the mother, and K children. For the time being, let us think of the household as a unit whose decision is made by only one of the parents—the householder. The householder's utility function is given by

$$u^d = U^d(c, l_f, l_m, q), \tag{1}$$

where c is the pooled household consumption, l_f the father's leisure, l_m the mother's leisure, and q the average quality of the children[3]. The children's average quality is determined by the following 'children's quality production function:'

$$q = Q(e_1, e_2, \ldots, e_K, t_f, t_m, h_1, h_2, \ldots, h_K), \tag{2}$$

where $e_1, \ldots, e_K$ are the goods input for the child's education, t_f and t_m are the father's and the mother's time inputs, and $h_1, \ldots, h_K$ are the children's quality endowment. The budget constraint is

$$I_f + I_m + w_f(T - l_f - t_f) + w_m(T - l_m - t_m) = \sum_{i=1}^{K} p_{e_i} e_i + p_c c, \tag{3}$$

where I_f and I_m are the parents' nonlabor incomes, w_f and w_m the wages, T is the total available time, and $p_{e_1}, \ldots, p_{e_K}, p_c$ are the prices. The householder's problem is to choose c, l_f, l_m, t_f, t_m, and $e_1, \ldots, e_K$ to maximize the utility (1) subject to the children's quality production function and the pooled budget constraint. Our interest lies in the optimal choice of e_i for $i = 1, 2, \ldots, K$, e_i^*, which can be written in the following general function form:

$$e_i^* = f(I_f + I_m, w_f, w_m, \mathbf{p}', \mathbf{h}', K), \tag{4}$$

where $\mathbf{p}$ is the $(K+1) \times 1$ price vector and $\mathbf{h}$ is the $K \times 1$ children's quality endowment vector. This function shows that according to the unitary household model, the parents' nonlabor incomes are pooled.

Now let us turn our attention to the non-unitary models in which one parent's utility function may be different from the other's. The father's and the mother's utility is a function of his or her own consumption and leisure, the spouse's consumption and leisure, and the household's public consumption. One element of the public household goods is the children's average quality. We can write the father's ($j = f$) and the mother's ($j = m$) utility function as follows:

$$u^j = U^j(q, c_0, c_f, c_m, l_f, l_m), \tag{5}$$

for $j = f, m$ where c_0 is the public household goods other than the children's quality, c_f and l_f are the father's consumption and leisure, and c_m and l_m are the mother's consumption and leisure.

There are two well-known models that explain how the final household decisions are made between the two individuals of different preferences. One is the Nash household bargaining model (McElroy and Horney 1981, McElroy 1990) and the other is the Pareto-efficient household model (Chiappori 1988, Chiappori et al. 2002). While the assumptions of the two models differ, the empirical implications derived from the two models to be tested in this paper are identical. Since the empirical implications can be derived more directly from the Nash household bargaining model than from the other, the bargaining model is explained in more detail here than the Pareto-efficient model is.

If the parents' marriage dissolved, the father and the mother would draw utility from their private consumption goods, own leisure, and the portion of the household public goods allocated to them by law or the social custom. If they were not married, the father and the mother would maximize his or her own utility:

$$u_0^j = U_0^j(q_j, c_{0_j}, c_j, l_j), \tag{6}$$

where q_j and c_{0_j} for $j = f, m$ is the amount of the public goods j would keep if their marriage dissolved, subject to the individual budget constraint:

$$I_j + w_j(T - l_j - t_j) = \frac{q_j}{q} \sum_{i=1}^{K} p_{e_i} e_i + p_c(c_{0_j} + c_j). \tag{7}$$

The father's and the mother's maximum utility level that would be obtained if their marriage dissolved is, therefore, written in a value function:

$$V_0^j = V_0^j(w_j, I_j, \mathbf{p}', \mathbf{h}', K|\boldsymbol{\alpha}_j), \tag{8}$$

for $j = f, m$ where $\boldsymbol{\alpha}_j$ is the vector of variables that affect the maximum value of utility attainable by j outside of the marriage, so-called extrahousehold environmental parameters (EEPs). The EEPs certainly include the marriage market situations. One important parameter is the sex ratio, the number of males divided by that of females. Once we view the marriage as the bargaining outcome between a man and a woman and apply the sequential bargaining model of Rubinstein and Wolinsky (1985), we can show that as the number of males relative to that of the females increases, the husband is willing to concede a larger share of the gains from the marriage with the wife. It implies that, as the sex ratio increases, the wife's expected maximum utility outside the marriage increases, while the husband's decreases. That is,

$$\frac{\partial V_0^f}{\partial r} < 0 \quad \text{and} \quad \frac{\partial V_0^m}{\partial r} > 0, \tag{9}$$

where r denotes the the sex ratio in the marriage market.

The couple's objective is to maximize the product of their gains from the marriage:

$$u_1 = \left[U^f(q, c_0, c_f, c_m, l_f, l_m) - V_0^f(w_f, I_f, \mathbf{p}', \mathbf{h}', K|\boldsymbol{\alpha}_f)\right] \times \\ \left[U^m(q, c_0, c_f, c_m, l_f, l_m) - V_0^m(w_m, I_m, \mathbf{p}', \mathbf{h}', K|\boldsymbol{\alpha}_m)\right], \tag{10}$$

subject to the combined budget constraint

$$I_f + I_m + (w_f + w_m)T = \sum_{i=1}^{K} p_{e_i} e_i + p_c(c_0 + c_f + c_m) + w_f(l_f + t_f) + w_m(l_m + t_m) \quad (11)$$

and the children's quality production function (2). By solving the utility maximization problem we can derive the demand function of e_i for $i = 1, 2, \ldots, K$:

$$e_i^* = g(w_f, w_m, I_f, I_m, \mathbf{p}', \mathbf{h}', K|\boldsymbol{\alpha}_f, \boldsymbol{\alpha}_m). \quad (12)$$

The equation (12) derived from the Nash household bargaining model is different from the equation (4) derived from the unitary household utility model in two aspects. First, the equation (12) implies an *incomplete* income pooling while the equation (4) a complete income pooling. Since the mother's and the father's nonlabor income directly affects their respective threat points in the household bargaining, they have different effects on the household demand for goods. As the mother's (father's) nonlabor income increases, the household consumption becomes more aligned to the mother's (father's) preference, everything else equal. Second, the equation (12) implies that the EEPs have effects on the demand while the equation 4 does not. McElroy and Horney (1981, equation 13) show that if EEPs change to increase the mother's threat point and lower the father's threat point, the household consumption shifts to the goods valued relatively more by the mother than by the father. Therefore, if the sex ratio in the marriage market increases and the child quality is valued more (less) relative to other goods by the mother than by the father, the model predicts that the demand for e_i will increase (decrease). This implication, when it is combined with the incomplete income pooling, leads to the 'consistency' condition. That is, an increase of the mother's nonlabor income and an increase of the sex ratio should have the same qualitative effect, though they may be different quantitatively, on the child quality demand, because both changes increase the mother's utility outside the marriage. On the other hand, the consistency condition implies that the sex ratio and the father's nonlabor income should have the opposite qualitative effects.

The Pareto-efficient household model, the other non-unitary household model, does not assume any specific intrahousehold decision making mechanism as the household bargaining model does, but assumes that the intrahousehold resource allocation is Pareto-efficient. Therefore, for any given $(w_f, w_m, I_f, I_m, \mathbf{p}', \mathbf{h}', K, \boldsymbol{\alpha}_f, \boldsymbol{\alpha}_m)$, there exists a weighting factor $\mu(w_f, w_m, I_f, I_m, \mathbf{p}', \mathbf{h}', K, \boldsymbol{\alpha}_f, \boldsymbol{\alpha}_m)$ belonging to $[0, 1]$ such that the chosen levels of $(l_f, t_f, l_m, t_m, c_0, c_f, c_m, e_1, \ldots, e_K)$ maximize the following 'collective' utility:

$$\mu u^f + (1 - \mu)u^m \tag{13}$$

subject to the equations (2) and (11). This leads to the demand function for e_i for $i = 1, \ldots, K$ which can be written in the following general form:

$$e_i^* = h[w_f, w_m, I_f, I_m, \mathbf{p}', \mathbf{h}', K, \mu(w_f, w_m, I_f, I_m, \mathbf{p}', \mathbf{h}', K, \boldsymbol{\alpha}_f, \boldsymbol{\alpha}_m)]. \tag{14}$$

Note that this equation, like the equation (4), implies an incomplete income pooling and the effect of the EEPs (distribution factors) on the demand for e_i. The EEPs affect the demand by shifting the Pareto weight. If the individual preference is egotistic, the consistency condition is implied (Chiappori et al. 2002).

3 Data

The samples used for estimations in this study are drawn from the Second Indonesian Family Life Survey data of 1997 (IFLS–2). The IFLS is a longitudinal socioeconomic and health survey of Indonesian households representing about 83 percent of the Indonesian population living in 13 of the nation's 26 provinces. Two waves of the IFLS data, collected in 1993 and 1997, are now available. The IFLS provides us with rich socioeconomic information on individuals and households, including adults' nonlabor income, children's school enrollment status, household-level education expenditures, and individual-level education expenditure for selected 7 to 15 year-old children. In addition to the individual and household information, the IFLS collects community (village or township) information, for example, information on transportation, industries, schools, and hospitals[4].

The sex ratio variable used in the study is computed using the provincial population counts from Indonesia's Population Census in 2000[5]. In the regression analyses whose results are reported in this article, the single-age sex ratio at the father's age is used. That is, for a household k in province π with the father who is τ years old at the time of the IFLS–2 survey, its associated sex ratio is

$$r_k = \frac{M_{\pi,\tau+3}}{F_{\pi,\tau+3}}, \tag{15}$$

where $M_{\pi,\tau}$ is the number of τ-year old males in province π as counted in the Population Census of 2000 and $F_{\pi,\tau}$ is the number of τ-year old females in province π in the Census. Some alternative definitions of the sex ratio, for example, $\sum_{\iota=1}^{5} M_{\pi,\tau+\iota} / \sum_{\iota=1}^{5} F_{\pi,\tau+\iota}$, are tried, but the estimation results remain almost the same so that they are not reported in this article. I also find similar results from estimations done using sex ratios computed with the community population size reported in the IFLS–2 data instead of those from the Census[6].

The nonlabor income variable used in the study is the sum of pension, scholarship, insurance claims, lottery winnings, and other nonlabor income received in the previous year by the respondents to the survey. Gifts or transfers received are not counted, because they are likely to depend on the economic status and needs of the household, and therefore clearly endogenous.

Three different samples are extracted from the IFLS–2 data. The first sample comprises of the household heads' children 7 to 24 years old. The children's school enrollment status and other personal characteristics are drawn from the household rosters. This sample is used to estimate the effects of the provincial sex ratio and parental nonlabor incomes on children's school enrollment status. The second sample is a subset of the first sample that comprises of the household heads' children 7 to 16 years old—7 to 14 years old except for 6 children—on whom much detailed survey was carried out in the IFLS–2. The detailed survey provides us with information on the children's health status and the expenditures spent for their education in the 1997–1998 school year. The expenditure variable is the sum of the school fees, expenditure on school supplies, transportation and pocket money, and other school related expenses. It should be noted that the IFLS–2 collected the detailed information on children for up to only three children in each household who are (supposed to be) 7 to 15 years old. Since the education expenditure information is available only for a subset of children who are relatively young—mostly of primary school age—I complement the analysis of individual-level education expenditures with an analysis of household-level education expenditures. The household expenditure information is obtained from the questionnaire on household consumption. For the household-level analysis, a sample of households that have at least one member who is 7 to 24 years old is used. The households included in the three samples share the following characteristics: the household head is male and married, and both he and his spouse are 25 years old or older[7]. Observations that have missing information on any necessary variable are excluded from the samples.

[Table 1 here.]

Table 1 shows the summary statistics of the dependent variables—children's school enrollment status and the individual and the household education expenditures—and some key explanatory variables—sex ratio, mother's and father's nonlabor incomes, child's age and sex—used in the estimations whose results are reported in the panel (A) of Table 2. The upper panel (A) shows the statistics for observations from urban areas and the lower panel (B) for observations from rural areas. Comparing the statistics across the two panels, we can spot urban-rural differences, somewhat expected, in children's education. Urban children are more likely to be enrolled in school than rural children—74 percent of the urban children are enrolled in a school, while 69 percent of the rural children are. Urban households also spend much more for their children's education than rural households. The individual education expenditure data indicates that the urban households spend, on average, twice as much as the rural households per child—mostly of elementary school age—on education. On the household level, the rural households spend about one third of what the urban households spend for children's education. While the urban households and the rural households appear to spend greatly different amount of money on children's education in absolute terms, in relative terms to the household earnings, the rural-urban difference in education expenditures is a lot smaller—the median ratio of the household education expenditure to the household earnings is 9.6 percent for the urban households and 7.0 percent for the rural households (not shown).

The average provincial sex ratio is slightly slightly higher than 1.0. Although not shown in the table, the distribution of the sex ratio ranges wide from .7 to 1.6. Urban parents are likely to have higher nonlabor income than their rural counterparts. Urban mothers' average nonlabor income ranges from 18,000 to 48,000 rupiah a year depending on the sample, while rural mother's from 3,000 to 5,000 rupiah only. Fathers tend to receive higher nonlabor incomes than mothers. Urban fathers' average nonlabor income ranges from 95,000 to 216,000 rupiah a year, while rural father's from 56,000 to 66,000 rupiah. It should be noted here that the vast majority of the parents do not have any nonlabor income. 97 percent of the urban mothers, 93 percent of the urban fathers, 99 percent of the rural mothers, and 97 percent of the rural fathers report zero nonlabor income.

4 Estimation and Results

4.1 Empirical Model

Using the three samples described in the previous section, I estimate three equations: the children's school enrollment equation, the individual school expenditure equation, and the household school expenditure equation. Let e^*_{ki} denote the parents' 'desired' education expenditure for child i in household k. Based on the equation (12), we write the equation for e^*_{ki} as follows:

$$e^*_{ki} = \beta_0 + \beta_1 r_k + \beta_2 I_{k_m} + \beta_3 I_{k_f} + \boldsymbol{\beta}_4 \mathbf{H}_{ki} + \boldsymbol{\beta}_5 \mathbf{X}_{k_m} + \boldsymbol{\beta}_6 \mathbf{X}_{k_f} + \boldsymbol{\beta}_7 \mathbf{Z}_k + \boldsymbol{\beta}_8 \mathbf{S}_k + \varepsilon_{ki}, \quad (16)$$

where r_k is the provincial sex ratio at the father's age computed by equation (15), I_{k_m} and I_{k_f} are the mother's and the father's nonlabor income, $\mathbf{H}_{ki}$ is the vector of the child's characteristics (age, age squared, and sex), $\mathbf{X}_{k_m}$ and $\mathbf{X}_{k_f}$ are the vectors of the mother's and the father's characteristics (age, age squared, and education), $\mathbf{Z}_k$ is the vector of household characteristics (number of household members by age groups—0 to 6, 7 to 12, 13 to 18, 19 to 55, and 56 or older), $\mathbf{S}_k$ is the vector of community characteristics indicating the community's educational infrastructure (the number of primary, junior secondary and senior secondary schools and the total population size), and ε_{ki} is the random factor that is uncorrelated across households but may be correlated within the household. Note that here I assume that the choice of education expenditure for child i does not depend on other children's quality endowment and that the child i's quality endowment is controlled for by the child's and the parents' observed characteristics. Wages and price differences of education goods are controlled for by the parents' characteristics, the urban dummy, and the community characteristics.

According to the unitary household model, $\beta_1 = 0$ and $\beta_2 = \beta_3$ should hold. The household bargaining model and the Pareto-efficient household model, on the other hand, reject the implications of the unitary household model. As discussed in section 2, the household bargaining model demands consistency between the sex ratio and the parental nonlabor income parameters, that is, $\beta_1\beta_2 > 0$ and $\beta_1\beta_3 < 0$.

The enrollment status of a child is determined by whether the parents' desired education expenditure for her is greater than zero or not. That is,

$$a_{ki} = \begin{cases} 1 & \text{if } e^*_{ki} > 0, \\ 0 & \text{otherwise,} \end{cases} \tag{17}$$

where $a_{ki} = 1$ if the child i in household k is enrolled in a school, and $a_{ki} = 0$ otherwise. Assuming that ε_{ki} is normally distributed, probit model is used to estimate the parameters of the equation (16) using the sample of 7 to 24 year children.

The parameters of the equation (16) can be also estimated by using education expenditure data, assuming the desired education expenditure is equal to the observed education expenditure. Two equations of education expenditures, one for individual-level expenditures using the sample of 7 to 16 year old children and the other for household-level expenditures using the sample of households that have one or more 7 to 24 year old members, are estimated. Since education expenditures are observed only for children enrolled in a school, sample-selection corrected education expenditure equation is estimated by Heckman's method for the individual-level expenditures. For identification purposes, the sample selection equation includes nineteen dummy variables indicating whether a child suffers from various disease symptoms—fever, nausea, diarrhea, and infections in body parts—in addition to the variables in the expenditure amount equation. As the disease symptom variables are available only for those in the second sample, they are not included in the general enrollment equation (17) estimated using the household roster information. The equation of household-level education expenditures, equivalent to the aggregation of equation (16) across children within households, is estimated by OLS.

In order to investigate urban-rural differences in parents' investment in children's education, the three equations described above are estimated separately for the urban sample and the rural sample. Statistical tests reject equality of the parameters estimated using the two different samples in most specifications.

4.2 Estimation Results

[Table 2 here.]

Table 2 shows the estimated coefficients of the sex ratio, the mother's nonlabor income, and the father's nonlabor income in the three equations for the urban and the rural households. The left panel (A) of the table shows the coefficient estimates using the setup of equation (16) and the right panel (B) shows the coefficient estimates using an augmented setup that has the current household earnings and the asset holdings as additional explanatory variables. Panels (A) and (B) are also divided into the upper and the lower subpanels showing the estimation results for the urban households and the rural households respectively. For all the coefficient estimates, standard errors are adjusted for heteroscedasticity and for possible correlations among children within the same household[8].

The first columns of the panels (A) and (B) show the coefficient estimates of the school enrollment equation for the household heads' 7–to–24–year old children. In the both panels (A) and (B), the estimated sex ratio coefficient is positive and statistically insignificant at any conventional level for the urban households, while the coefficient for the rural households is negative and statistically significant at the 5 percent level for the rural households. The estimated coefficients imply that, at the sample mean, an increase of the local sex ratio by 0.1—which is approximately its standard deviation—raises the urban children's school enrollment probability by 0.3 percentage points, but decreases the rural children's by 2 percentage points.

In urban areas both parents' nonlabor incomes are estimated to have positive effects on children's school enrollment probability. The magnitudes of the effects are, however, quite different. An increase of the mother's nonlabor income by one million rupiah a year is estimated to increase the school enrollment probability, at the sample mean, by 4 percent, while the same increase of the father's nonlabor income increases the probability by only 0.4 percent. The statistical test rejects equality of the father's and the mother's nonlabor income coefficients, or the income pooling hypothesis, at the 1 percent level in the panel (A) and at the 5 percent level in the panel (B) for the urban households. In rural areas, on the other hand, the mother's nonlabor income coefficient is estimated to be negative and the father's positive. However, neither of them is statistically significant even at the 10 percent level. For rural households, equality of the parental nonlabor income coefficients is not rejected at any conventional level.

The joint hypothesis that the sex ratio has no effect and the parents' incomes are pooled is rejected at the 10 percent or the smaller levels both in the urban and the rural areas. In urban areas it is rejected at the 2 percent and at the 8 percent levels in the panels (A) and (B) respectively; in rural areas at the 7 percent and at the 10 percent levels. The signs of the sex ratio coefficient and the maternal nonlabor income coefficient seem to suggest that the mother's household bargaining power is correlated with the children's school enrollment probability positively in urban areas, but negatively in rural areas.

The second column of each panel shows the coefficient estimates of the individual education expenditure equation for children 7 to 16 years old. The sex ratio coefficient estimates are all positive. According to the estimates in the panel (A), however, the sex ratio has much greater effect on the education expenditure in urban areas than in rural areas—an increase of the provincial sex ratio by 0.1 increases the annual education expenditure for an urban child by 29,000 rupiah, or 15 percent from the sample mean, but for a rural child only by 3,400 rupiah, or 4 percent from the mean. Note also that the sex ratio coefficient is statistically significant at the 1 percent level in urban areas in both panels, while, in rural areas, statistically insignificant at any popular level in the panel (A) and significant only at the 10 percent level in the other panel.

As in the enrollment equation, the income pooling hypothesis is rejected soundly in urban areas. The mother's nonlabor income coefficient is positive and statistically significant at the 1 percent level, while the father's nonlabor income coefficient is slightly negative but statistically insignificant. In rural areas neither parent's nonlabor income coefficient is statistically significant and the income pooling hypothesis is not rejected. The joint test of the sex ratio coefficient and the income pooling hypothesis rejects the null only in urban areas. The unitary household model is, therefore, rejected in urban areas, but not in rural areas.

The last columns show the coefficient estimates of the household-level education expenditure equation. The sex ratio coefficient estimation results are similar to those of the individual expenditure equation. That is, in urban areas, the sex ratio coefficient is positive and statistically significant, while in rural areas it is small, positive, but statistically insignificant. The correlation of the sex ratio and the household education expenditure is strong in urban areas—an increase of the sex ratio by 0.1 is associated with an increase of the household education expenditure by 130,000 to 156,000 rupiah a year or about 20 percent of the average expenditure. Unlike other equations, however, the income pooling hypothesis is rejected in neither area. As with the individual expenditure equation, note that the joint test rejects the unitary household model in urban areas, but not in rural areas.

Collating the estimation results in Table 2 described above, we find that the unitary household model is rejected in urban areas in every specification. The estimated sex ratio coefficient is positive in urban areas without an exception and the mother's nonlabor income coefficient estimate is always positive and larger than the father's nonlabor income coefficient estimate. The estimation results, therefore, strongly suggest that in urban areas the mother's household bargaining power has positive effects on the amount of household resources devoted to the children's education, while the father's has little or negative effects on the investment in children's education. On the other hand, in rural areas the unitary household model is rejected only in the enrollment equation. The estimated sign of the sex ratio coefficient is negative in the enrollment equation, but positive in the other two equations. The mother's nonlabor income seems to be negatively correlated with the investment in children's education, but neither parent's nonlabor income coefficient is statistically significant in any specification.

As discussed before, the household bargaining model demands consistency between the sex ratio and the parental nonlabor income parameters. There are two consistency conditions. One is that the sex ratio coefficient and the mother's nonlabor income coefficient should have the same sign. The other is that the sex ratio coefficient and the father's nonlabor income coefficient should have the opposite sign. In urban areas where the unitary household model is rejected and the sex ratio and the mother's nonlabor income coefficients are estimated with some precision, the first condition seems to be satisfied in all equations. The second condition is, however, satisfied only in the individual expenditure equation. In rural areas, on the other hand, the first and the second conditions are met in the enrollment equation in which the unitary household model is rejected. In other equations where the unitary model is not rejected, the conditions are not met in general. It is notable that the fist consistency condition, which can be assessed more reliably than the second condition since the father's nonlabor income coefficient is never estimated with enough precision, is satisfied whenever the unitary household model is rejected.

So far the sex ratio coefficient has been interpreted as the effect of the mother's household bargaining power on the parents' household resource allocation decisions for children's education. However, one may object to such interpretation of the coefficient on the following two grounds. One is that the households' fertility choices may be inadequately controlled for in estimations. If there is a negative relationship between the mother's bargaining power and the mother's 'desired' fertility, we may find a positive relationship between the sex ratio and the investment in 'existing' children's education, not because of the direct effect of the mother's bargaining power on education expenditures, but because of the fertility choices. The other is that the sex ratio may be correlated with income differences across communities which may be insufficiently controlled for. Considering that men, who are main earners in households, are likely to be attracted to areas where they can receive high wages, the sex ratio is likely to be positively correlated with the level of local income and asset holdings. Therefore, the positive relationship between the sex ratio and expenditures on children's education may be nothing but the income effect.

The two points raised above are actually backed up by the data. Controlling for the parents' education and age, the number of existing male and female children, and the urban area dummy, a probit analysis suggests that an increase of the provincial sex ratio by 0.1 decreases the probability that the mother wants another child by 1.1 percentage points. The relationship of the provincial sex ratio with the local income and wealth level is very strong. Controlling for the father's age, age squared, and education level and the urban area dummy, an increase of the provincial sex ratio by 0.1 is estimated to be associated with an increase of the average household earnings of a community by 160,000 rupiah and of the average household assets by 6 million rupiah.

[Table 3 here.]

To examine whether the coefficients estimated in Table 2 substantially change if the two points are taken into account in estimations, the three equations are re-estimated, including in the right hand side the average household earnings and asset holdings of the community and the dummy variable indicating the mother's desire for another child. Since the mother's fertility choice is an endogenous variable, the dummy variable is instrumented by nine dummies indicating whether the mother can handle activities of daily living—carrying water, walking, kneeling, etc.—with ease. The estimation results are in Table 3. Due to collinearity of the individual household earnings and assets with the average values, the individual household earnings and assets variables are not included in the equation.

It is found that inclusion of the three new variables reduces the magnitude of the sex ratio coefficient in general, especially in urban areas. Restricting the samples to those included in the estimations in Table 3 and doing OLS estimations without the three variables, the sex ratio coefficient (robust standard error) of the three equations is estimated to be 0.043 (0.075), 0.271 (0.064), and 1.847 (0.464) in urban areas and -0.169

(0.080), 0.057 (0.026), and -0.0001 (0.132)[9]. In the IV estimation results of Table 3, those estimates become 0.001, 0.189, 1.475, -0.176, 0.037, and -0.011 respectively. It implies that the fertility decisions and the income and wealth differences across communities are partly responsible for the positive effect of the sex ratio on the investment in children's education, particularly in urban areas. It is notable that, although rarely statistically significant, all the estimated coefficients of the 'mother wants another child' dummy variable are negative in both urban and rural areas. Furthermore, the coefficients of the average household earnings and asset variables are estimated to be positive in urban areas.

However, although the estimates differ in the magnitude, note that the joint statistical test results regarding the household models we draw from the estimation results in Table 3 are identical to those from the results in Table 2—the unitary household model is rejected in every specification in urban areas, but it is rejected, somewhat marginally, only in the enrollment equation in rural areas. The income pooling hypothesis test results does not change, either. All in all, the estimation results suggest that although the fertility choices and the income differences are partly responsible for the positive sex ratio coefficient, they are not the main reason for it and the interpretation that the sex ratio is an indicator of the mother's household bargaining power withstand the attempts to falsify it.

One may now raise a question on exogeneity of the nonlabor income variable. As discussed previously, it is very difficult to solve the possible endogeneity problem of the nonlabor income variable. In other studies (Thomas et al. 2002; Quisumbing and Maluccio 2003), the size of individual assets brought to the marriage by the husband and the wife is suggested as an alternative to the postmarital nonlabor income variable. Following their lead, in Table 4 I estimate again the three equations with the value of individual assets at the time of marriage instead of the nonlabor income. The current household earnings and asset holdings are included in the right hand side to control for their effects on education investment. If the unitary household model is correct, the sex ratio and the individual premarital asset variables should have no effect on the household resource allocation, controlling for the household's current earnings and the wealth (Thomas et al. 2002). The estimation results are shown in Table 4.

[Table 4 here.]

Note first of all that the sex ratio coefficient estimates in Table 4 are little different from those in the panel (B) in Table 2. In urban areas the asset coefficients are jointly significant in all equations. On the other hand, in rural areas the parents' assets at the marriage appear to be largely uncorrelated with the size of household resources devoted to the children's education. Like the finding from the previous estimations using the nonlabor income variables, the unitary household model is rejected in urban areas in every equation, but in rural areas it is rejected, at the 9 percent level, only in the enrollment equation. The estimation results suggest that in urban areas the education expenditures increase as the mother's household bargaining power increases, while the relationship between the parents' bargaining powers and the children's school enrollment is not so clear. It shows that the overall implication of the equations do not change even if the parents' nonlabor income variables are replaced with their assets at the marriage variables.

It should be noted, however, that the consistency conditions are violated more frequently in Table 4 than in Table 2. In urban areas, none of the consistency conditions is satisfied in the enrollment equation, while in other equations they seem to hold in general. In the enrollment equation in the rural areas, on the other hand, none of the conditions is met. In other equations where the unitary household model is not rejected, the results are mixed.

The equations estimated so far assume that the parents regard equally the son's and the daughter's education. Previous studies have found, however, evidence to the contrary, though not in education. For example, Thomas (1994) finds that the daughter's health and nutritional status are correlated closer to the mother's education and nonlabor income than to the father's and that the son's are to the father's than to the mother's. Duflo (2000) also finds that the pension received by elderly women in South Africa has a large impact on health status of girls, but little effect on that of boys. Whether the same pattern holds for the children's education is examined here. I re-estimate the enrollment status equation and the individual education expenditure equation using the augmented setup of the panel (B) in Table 2, having in the right hand side the interaction terms between the child's sex dummy (1 if female, 0 otherwise) and each of the sex ratio, the maternal nonlabor income, and the paternal nonlabor income variables.

[Table 5 here.]

Table 5 shows the estimation results. In urban areas, the sex ratio and the both parents' nonlabor income are estimated to be all positively correlated with the son's school enrollment probability. The daughter's school enrollment is, however, negatively correlated with the parents' nonlabor incomes. In the individual expenditure equation, while both the sex ratio and the mother's nonlabor income have positive relationships with the the son's and the daughter's education expenditures, their relationship with the son's education expenditures is estimated to be stronger than that with the daughter's education expenditures. The father's nonlabor income is estimated to be weakly correlated with any child's education expenditures. It implies that in urban areas an increase of the mother's household bargaining power is likely to increase an investment in the son's education more than an investment in the daughter's. In rural areas, it appears that the son's and the daughter's education are treated more or less equally. The interaction terms are jointly statistically insignificant even at the 10 percent level in each equation.

We can find an explanation for it in the parents', especially the mother's, needs for old age security in Indonesia. Lacking social security, most old Indonesians depend on their own savings, the family's support, and especially the children's transfers for economic security. The need for old age security is greater for mothers than for fathers, because mothers are likely to live longer and more likely to be disadvantaged economically in old ages than fathers. That the mother's strong bargaining power is more positively correlated with sons' education than with daughters' education should be attributed to that the primary economic support for mothers comes from sons, either by cohabitation or by remittances, rather than from daughters (Park 2003). In a similar vein, Schultz (1990) finds that the mother's nonlabor income is positively correlated with fertility in Thailand and attributes it to the mothers' needs for old age security. The finding that the son's education is preferred to the daughter's in Indonesia coincides with that of Quisumbing and Maluccio (2003).

5 Conclusion

In this paper I estimate how each parent's household bargaining power affects children's school enrollment status and household education expenditures, using data from the Second Indonesian Family Life Survey and the Indonesia's Population Census in 2000. The estimation results are used to test validity of the unitary and the non-unitary household models. Instead of measuring an individual's household bargaining power only by his or her nonlabor income which may be endogenous, in this paper I exploit the implication of the non-unitary models that the marriage market conditions should affect intrahousehold resource allocations. Thus I estimate the effect of the provincial sex ratio (male-female ratio) in addition to that of each parent's nonlabor income on the household resource allocation.

I find that in urban areas the local sex ratio has a strong positive effect on the amount of expenditures on children's education. I also find that in urban areas the income pooling hypothesis is rejected in most estimated equations and that the mother's nonlabor income is more strongly correlated than the father's with the household's investment in children's education. In rural areas, on the other hand, I find that the parents' nonlabor incomes are generally uncorrelated with the children's school enrollment status or the education expenditures and that the income pooling hypothesis cannot be rejected. Furthermore, while the provincial sex ratio is negatively correlated with the children's school enrollment, its effects on the education expenditures are found to be statistically insignificant. In sum, the unitary household model is rejected in all specifications in urban areas, but in rural areas it is rejected only in the school enrollment equation. The finding holds even if the nonlabor income variables are replaced with the values of the parents' individual assets brought to the marriage. The estimation results change little even when the potential effects of the sex ratio on women's fertility choices and the average income and assets of the communities are controlled for. I also find that in urban areas both parents prefer educating sons to daughters.

The finding of this paper complements other research findings that the parents' nonlabor incomes have different effects on investment in children's welfare. The finding of this paper and of the previous studies suggest that implementing policies to strengthen the mother's household bargaining power will improve the children's welfare in many areas. For example, social welfare benefits can become more effective in improving welfare of children by assigning mothers rather than fathers to be the recipients. Expansion of women's opportunities in the labor market is also likely to improve the children's education and health. The findings of this paper, however, indicate that the 'mother-empowering' policies may not work universally. In Indonesia, those policies are likely to be effective only in urban areas. Another evidence found in this study indicates that while the mother-empowering policies will increase investment in education both of the son and of the daughter, they are likely to have bigger impact on the son's education than on the daughter's.

Notes

[1]For an example from an actual policy, look at Lundberg et al. (1997).

[2]Rao (1993) studies the effect of the sex ratio on the value of dowries in India, which is not directly related to the household resource allocation we are interested in. Look at Edlund (2000) and Rao (2000) for a critique on the work.

[3]For simplicity, following Becker and Lewis (1973), the quality is assumed to be the

same for all of the children.

[4]The 'community' of IFLS is the enumeration area (EA) defined by the sampling framework designed by the Indonesian Central Bureau of Statistics (BPS) and used for Indonesia's SUSENAS of 1993, a socio-economic survey of about 60,000 households.

[5]The population counts are obtained from *Results of the 2000 Population Census* published by the BPS.

[6]The unreported estimation results are available upon request from the author.

[7]A few polygynous households are excluded from the sample.

[8]Statistical tests strongly reject the homoscedasticity assumption in any setup.

[9]The full estimation results are available upon request from the author.

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Table 1: Summary Statistics of Some Key Variables

Variable	Enrollment data	Ind. expenditure data	HH expenditure data
		(A) Urban areas	
Enrolled in school (1 if enrolled)	.740	–	–
Individual education expenditure† $\times 10^{-6}$	–	.187 (.191)	–
Household education expenditure $\times 10^{-6}$	–	–	.760 (1.466)
District sex ratio in 2000 at the father's age	1.046 (.104)	1.038 (.102)	1.030 (.111)
Mother's nonlabor income $\times 10^{-6}$	.030 (.361)	.018 (.235)	.048 (.454)
Father's nonlabor income $\times 10^{-6}$	.169 (.845)	.095 (.589)	.216 (.971)
Child's age	14.5 (4.6)	10.6 (2.3)	–
Child's sex (1 if female)	.479	.490	–
Number of children	2909	1511	–
Number of households	1291	977	1381
Number of communities	165	162	164
		(B) Rural areas	
Enrolled in school (1 if enrolled)	.687	–	–
Individual education expenditure† $\times 10^{-6}$	–	.085 (.089)	–
Household education expenditure $\times 10^{-6}$	–	–	.254 (.413)
District sex ratio in 2000 at the father's age	1.029 (.111)	1.021 (.107)	1.015 (.117)
Mother's nonlabor income $\times 10^{-6}$	.003 (.063)	.005 (.079)	.004 (.080)
Father's nonlabor income $\times 10^{-6}$	.056 (.437)	.061 (.455)	.066 (.479)
Child's age	13.8 (4.5)	10.4 (2.2)	–
Child's sex (1 if female)	.484	.508	–
Number of children	3330	1951	–
Number of households	1627	1282	1763
Number of communities	118	118	118

Note: Standard deviations are in the parentheses.

† Computed using positive expenditures only.

Table 2: Estimated coefficient of sex ratio and nonlabor incomes: by equation

Variable	Enrollment	Individual expenditure (in million)♯	Household expenditure (in million)	Enrollment	Individual expenditure (in million)	Household expenditure (in million)
	(A) Without household earnings and asset variables			(B) With household earnings and asset variables		
	(a) Urban Areas			(a) Urban Areas		
Sex ratio	.216	.288*	1.558*	.088	.243*	1.300*
	(.410)	(.062)	(.370)	(.416)	(.060)	(.383)
Mother's nonlabor income$\times 10^{-6}$	.298*	.139*	.156	.241*	.158*	.045
	(.082)	(.040)	(.122)	(.087)	(.045)	(.153)
Father's nonlabor income$\times 10^{-6}$	.025	-.006	.049	.001	-.008	.036
	(.040)	(.006)	(.059)	(.044)	(.006)	(.062)
Income coefficient difference	.273*	.145*	.107	.241**	.166*	.009
	(.096)	(.042)	(.142)	(.107)	(.047)	(.169)
Joint test for sex ratio coeff. &	8.29†	17.40‡	10.42‡	5.11†	31.00†	6.09‡
income pooling hypothesis	(.016)	(.000)	(.004)	(.078)	(.000)	(.002)
No. of observations	2909	1511	1381	2765	1477	1305
	(b) Rural Areas			(b) Rural Areas		
Sex ratio	-.780**	.034	.033	-.728**	.043***	.037
	(.344)	(.022)	(.087)	(.351)	(.022)	(.084)
Mother's nonlabor income$\times 10^{-6}$	-.342	-.018	-.054	-.352	-.022	-.032
	(.807)	(.027)	(.051)	(.761)	(.028)	(.043)
Father's nonlabor income$\times 10^{-6}$	.119	.004	.012	.086	.005	-.011
	(.074)	(.008)	(.028)	(.071)	(.009)	(.016)
Income coefficient difference	-.461	-.021	-.066	-.439	-.027	-.021
	(.813)	(.031)	(.071)	(.767)	(.033)	(.048)
Joint test for sex ratio coeff. &	5.55†	1.41‡	0.48‡	4.72†	4.32†	0.21‡
income pooling hypothesis	(.062)	(.246)	(.618)	(.095)	(.115)	(.811)
No. of observations	3330	1951	1763	3169	2064	1686

Note: Other RHS variables are the parents' age, age squared, and education; number of hh members by age group; the number of schools the community uses; and total population. The enrollment and the individual expenditure equation also have the child's age, age squared, and sex dummy.

Robust standard errors corrected for within-household correlations are in the parenthesis

*$p < .01$; **$.01 \leq p < .05$; ***$.05 \leq p < .10$. (two-sided test); *$p < .01$; **$.01 \leq p < .05$; ***$.05 \leq p < .10$. (one-sided test)

† χ^2 Wald test statistic and the p-value in the parentheses; ‡ F test statistic and the p-value in the parentheses;

♯ The coefficients are estimated by OLS using uncensored observations because Heckit procedures fail to converge.

Table 3: Estimated coefficient of sex ratio and nonlabor incomes, controlling for the mother's desire for another child and the community-level average income and wealth: IV regression results

Variable	Enrollment	Individual expenditure (in million)	Household expenditure (in million)
(A) Urban Areas			
Sex ratio	.001	.189*	1.475*
	(.078)	(.065)	(.520)
Mother's nonlabor income$\times 10^{-6}$	.078*	.069*	.207
	(.022)	(.023)	(.153)
Father's nonlabor income$\times 10^{-6}$	.021**	-.002	.045
	(.009)	(.007)	(.086)
Mother wants another child (1 if yes)	-.198	-.034	-.628
	(.283)	(.152)	(.847)
Average household earnings$\times 10^{-6}$	.002	.003	.006
	(.002)	(.002)	(.018)
Average household assets$\times 10^{-9}$	.138	.594***	3.205**
	(.189)	(.307)	(1.287)
Income parameter difference	.058*	.072*	.162
	(.025)	(.023)	(.194)
Joint test for sex ratio coeff. & income pooling hypothesis	2.72‡	9.69‡	5.45‡
	(.066)	(.000)	(.004)
No. of observations	2677	1418	1179
(B) Rural Areas			
Sex ratio	-.176**	.037	-.011
	(.088)	(.031)	(.152)
Mother's nonlabor income$\times 10^{-6}$	-.025	-.023	.011
	(.050)	(.029)	(.054)
Father's nonlabor income$\times 10^{-6}$	.007	.019	-.034
	(.014)	(.020)	(.023)
Mother wants another child (1 if yes)	-.144	-.088**	-.095
	(.127)	(.041)	(.207)
Average household earnings$\times 10^{-6}$	.004**	-.001	-.001
	(.002)	(.001)	(.004)
Average household assets$\times 10^{-9}$	.844	.847*	1.568
	(.607)	(.261)	(1.040)
Income parameter difference	-.031	-.042	.045
	(.058)	(.047)	(.068)
Joint test for sex ratio coeff. & income pooling hypothesis	2.53‡	0.99‡	0.22‡
	(.080)	(.372)	(.800)
No. of observations	2966	1775	1455

Note: For other RHS variables, look at the note of Table 2.

Robust standard errors corrected for within-household correlations are in the parentheses.

*$p < .01$; **$.01 \le p < .05$; ***$.05 \le p < .10$. (two-sided test)

‡ F test statistic and the associated p-value in the parentheses.

Table 4: Estimated coefficient of sex ratio and assets at marriage, controlling for the current household income and assets and the year of marriage

Variable	Enrollment	Individual expenditure (in million)♯	Household expenditure (in million)
(A) Urban Areas			
Sex ratio	.088	.238*	1.229*
	(.430)	(.062)	(.389)
Mother's assets at marriage $\times 10^{-9}$	-.539	.228*	.359**
	(.691)	(.022)	(.139)
Father's assets at marriage $\times 10^{-9}$	.856*	-.071*	.513
	(.140)	(.022)	(.547)
Joint significance of asset coefficients	37.21†	59.55‡	3.52‡
	(.000)	(.000)	(.030)
Joint significance of sex ratio and asset coefficients	45.55†	40.88‡	4.14‡
	(.000)	(.000)	(.006)
No. of observations	2654	1379	1242
(B) Rural Areas			
Sex ratio	-.746**	.037	-.011
	(.376)	(.022)	(.092)
Mother's assets at marriage $\times 10^{-9}$	55.254	.809	-4.313
	(50.881)	(.864)	(5.253)
Father's assets at marriage $\times 10^{-9}$	-4.754	-.118	-.369***
	(3.798)	(.119)	(.189)
Joint significance of asset coefficients	2.77†	0.98‡	2.09‡
	(.250)	(.377)	(.124)
Joint significance of sex ratio and asset coefficients	6.58†	2.00‡	1.40‡
	(.087)	(.112)	(.241)
No. of observations	3031	1786	1608

Note: Robust standard errors corrected for within-household correlations are in the parentheses.

$^{*}p < .01$; $^{**}.01 \leq p < .05$; $^{***}.05 \leq p < .10$. (two-sided test)

$^{\star}p < .01$; $^{\star\star}.01 \leq p < .05$; $^{\star\star\star}.05 \leq p < .10$. (one-sided test)

† χ^2 Wald test statistic and the associated p-value in the parentheses.

‡ F test statistic and the associated p-value in the parentheses.

♯ The coefficients are estimated by OLS using uncensored observations because Heckit procedures fail to converge.

Table 5: Estimated coefficient of sex ratio and nonlabor incomes interacted with the child's sex

Variable	Enrollment	Individual expenditure (in million)
(A) Urban areas		
Sex ratio	.112	.304*
	(.545)	(.074)
Sex ratio	-.097	-.112
× female dummy	(.656)	(.083)
Mother's nonlabor income$\times 10^{-6}$	.402***	.175*
	(.233)	(.030)
Mother's nonlabor income$\times 10^{-6}$	-.802*	-.018
× female dummy	(.286)	(.064)
Father's nonlabor income$\times 10^{-6}$	.130**	-.015
	(.058)	(.010)
Father's nonlabor income$\times 10^{-6}$	-.235*	.011
× female dummy	(.071)	(.012)
Female dummy	.282	.120
	(.690)	(.084)
No. of observations	2765	1477
(B) Rural areas		
Sex ratio	-.419	.072**
	(.442)	(.033)
Sex ratio	-.718	-.060***
× female dummy	(.574)	(.035)
Mother's nonlabor income$\times 10^{-6}$	-.446	.0003
	(.656)	(.024)
Mother's nonlabor income$\times 10^{-6}$	.314	-.047
× female dummy	(.560)	(.064)
Father's nonlabor income$\times 10^{-6}$	.102	-.009
	(.083)	(.007)
Father's nonlabor income$\times 10^{-6}$	-.045	.019
× female dummy	(.107)	(.014)
Female dummy	.754	.064
	(.598)	(.035)
No. of observations	3169	2064

Note: The variables listed in the note of Table 2, the household income, and the household assets are included in the RHS.

Robust standard errors corrected for within-household correlation are in the parenthesis.

$^{*}p < .01$; $^{**}.01 \leq p < .05$; $^{***}.05 \leq p < .10$.