

How Much of the Gender Difference in Child School Enrolment Can Be Explained? Evidence from Rural India

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Abstract: There are significant gender differences in child schooling in the Indian states though very few studies explain this gender difference. Unlike most existing studies we take account of the implicit and explicit opportunity costs of schooling and use a bivariate probit model to jointly determine child's participation in school and market jobs. Results obtained from the WIDER villages in West Bengal suggest that indicators of household resources, parental preferences, returns to and opportunity costs of domestic work significantly affect child school enrolment. While household resources have similar effects on enrolment of boys and girls, other arguments tend to explain a part of the observed gender difference. Even after taking account of all possible arguments, there remains a large variation in gender differences in child schooling that cannot be explained by differences in male and female characteristics in our sample.

Key words: Gender differences, Child schooling and child labour, Opportunity costs of schooling, Parental Preferences, Bivariate Probit, Oaxaca Decomposition.

JEL Classification: I21, O15

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1. INTRODUCTION

Aggregate data points to pronounced gender differences in child school enrolment in India. In 1991 Census the literacy rate for children aged 7 or more is 25% higher for boys than for girls. This figure, however, conceals the considerable inter-state variation: for example, the gender difference is about 7% in Kerala (the state with the highest literacy rate) while it is around 30% or more in the north Indian states of Bihar, Uttar Pradesh (UP) and Rajasthan. We use the data from West Bengal where the gender difference is 21% among children aged 7 or more. This gender difference persists even among 12-14 years old rural children in the state where 46% of girls as against 35% of boys were never enrolled in 1986-87. Recent research suggests that female schooling has important externality in that it plays a significant beneficial role on fertility (Pal and Makepeace, 2003) and child health outcomes (Pal, 1999) in low income countries like India. Thus boosting female literacy is necessary not only for itself but also for the wider social benefit.

There is a substantial literature on child schooling¹ in low-income countries. This literature identifies both demand (household income and parental education, e.g., see Behrman and Knowles, 1999; Duraisamy 2000; Kambhampati and Pal, 2001) and supply (variables reflecting quantity and quality of schools as in Drèze and Kingdon, 2001) factors as explanations of low educational achievement in these countries. Many of these studies find evidence of gender differences in schooling (e.g., Behrman and Knowles, 1999; Duraisamy, 2000) though there has been relatively much fewer attempts to *explain* gender differences in child schooling as such. In this context, we investigate the possible causes of gender differences in school enrolment among 5-15 year old boys and girls in rural West Bengal and also how much of the observed gender difference is explained by the characteristics of the sample children.

One can draw evidence from related studies to provide explanations of the observed gender differences in child school enrolment. Differential returns to boys' and girls' education seem to be the most common explanation in these studies. For example, using earnings function Kingdon (2002) argues that a significant proportion of the gender difference in child schooling in urban UP can be explained by gender differences in the returns to schooling.

¹ Different indicators of schooling have been used including completed years (Birdsall, 1985), current enrolment (Singh, 1992), ever attended (Cochrane, Mehra and Osheba, 1986), grades attained or grades failed (Drèze and Kingdon, 2001) and delayed enrolment (Glewwe and Jacoby, 1994).

Becker and Lewis (1965) argue that investment in the quality of children increases at higher levels of income. There is also some evidence that the gender gap closes at higher levels of income, especially if households are resource constrained. However, income seems to affect schooling choices of both boys and girls in rural India (Kambhampati and Pal, 2001).

Parental preferences may also be important in explaining gender differences. While Behrman (1988) has argued that parents are generally averse to inequality among children², there is evidence of 'son preference' among resource constrained parents in India (Sen and Sengupta, 1983; Kishor, 1993; Kingdon, 2002). Parents may prefer to invest in sons because they act as old-age security while girls leave the parents' house after marriage. It is, however, difficult to have a direct measure of parental preferences and thus most existing evidence in this respect is of indirect nature. For example, Garg and Morduch (1998) suggest that children (irrespective of their gender) are better off on measured health indicators if they have sisters and no brothers³ because parents tend to allocate less for girls. Dasgupta (1987) finds that, in rural Punjab, girls with older sisters suffer most. Kingdon (2002) used a variable relating to parental opinion about gender equality in education and finds that girls whose parents believed in gender equality attained significantly more education than other girls.

Parental preferences may not always be aligned; for example, mothers may have more empathy for daughters and fathers for sons. Lillard and Willis (1994) found that in Malaysia the mother's education has a far larger effect on the daughters' education (than on sons') and the father's education seems to have greater impact on sons. Arguing that each parent's education may be taken as indicator of his/her individual preference, Kambhampati and Pal (2001) also suggest that higher women's literacy encourages female education in rural Bengal. There may also prevail some complex inter-relation between household resource constraint and parental preferences in intra-household allocation of resources. This is highlighted in Quisumbing (1993) who argues that families with different land constraints have significantly different patterns of schooling investments resulting in inequality among siblings.

Thus the few existing studies of gender differences in child schooling in India tend to focus on a particular explanation of gender difference, e.g., differential returns to schooling (Kingdon, 1998) or parental preferences for sons (Kambhampati and Pal, 2001; Kingdon, 2002).⁴ There are

² This is indirectly supported by Butcher and Case (1994) who argued that girls raised with brothers have higher schooling.

³ However, the effect of sibling composition on child schooling is difficult to analyse because it may affect schooling choices in a number of ways. First sibling composition could affect the value of income per capita (with the arrival of new household members or demise of some existing ones). It may also affect the allocation of household tasks among various family members and thus the opportunity cost of attending schools. Assuming that parents are averse to inequality in child incomes, one also needs to account for the parental efforts to distribute resources equitably.

⁴ Kingdon examines the nature of gender difference in school *attainment* for *urban* children UP. Our study differs from Kingdon in that we jointly determine school and work participation in *rural* West Bengal.

moreover methodological issues which require a careful interpretation of these results. For example, use of earnings function (Kingdon, 2002) to estimate returns to schooling may yield biased estimates if one does not control for occupational differences or women's participation in non-market activities relating to pregnancy or child care. Secondly, in the absence of a better indicator to quantify parental preferences, use of an attitudinal variable (Kingdon, 2002) to measure parental attitude towards gender equality in education raises serious problems with the validity of these subjective responses. Also none of these studies take account of the important opportunity costs of schooling in terms of child's participation in domestic or market work.

We contribute to the existing literature in several ways. We take account of *several* possible causes of gender differences in child schooling and attempt to resolve some of the ambiguities/difficulties mentioned above. The paper is novel in a number of respects. Firstly there are important opportunity costs of schooling as reflected in the market job participation of some sample children (who may or may not participate in schools). Since participation in school and that in market work are both endogenous, we jointly estimate child's school enrolment and market participation, using a bivariate probit model.⁵ There may also be implicit opportunity costs of schooling in that a large number of children in the sample neither go to school nor take part in any market work. These children may be engaged in family farm/non-farm activities, which are not observed in our data set. It is argued that the included sibling composition variables included may indirectly capture a part of these implicit opportunity costs. Secondly since household expenditure is considered to be endogenous in household decision models, we use the predicted value of household expenditure per capita in stead. Thirdly gender difference in returns to schooling is an important explanation of gender difference in schooling. Returns to schooling are usually estimated by considering the effects of participation or wage rates on schooling. Since individual participation or wage rates are endogenous to household decisions in schooling, we use average *village-level* adult male and female participation rates and daily wage rates in stead. Finally we use a variant of the Oaxaca method to decompose gender difference in school enrolment into an explained and an unexplained variation. While the standard Oaxaca decomposition relates to wage earnings, we modify the technique and apply it to the correlated bivariate probit estimates of child school enrolment and wage employment. Our results suggest that there is a significantly large unexplained variation, often labelled 'discrimination', in gender differences in child school enrolment. The size of this unexplained variation alone, however, cannot constitute a proof of gender discrimination

⁵ Using the NCAER data from 16 major states in India, Duraisamy (2000) used a multinomial logit model to determine household decisions involving child schooling and child labour. In particular, she classifies children into three categories, namely, children going to school, children working and children involved in other activities. However, we find that children in our sample may combine schooling with or without work, the possibility that has not been accounted for by Duraisamy. Thus we consider a bivariate probit model to jointly determine school enrolment and work

hypothesis. Because use of some unobservable or imperfectly observable⁶ variables in our analysis may influence gender differences in school enrolment yet may not necessarily constitute discrimination.

The paper is organised as follows. Section 2 describes data and methodology while section 3 analyses the bivariate probit results and also those obtained from the Oaxaca-type gender decomposition exercise. Section 4 concludes.

2. DATA AND METHODOLOGY

The empirical analysis of gender differences in child schooling in this paper is based on the data from six villages⁷ in West Bengal for the period 1987-89 (for further details see Gazdar, 1992). The survey covered 749 households and 3972 individuals. The members were questioned for information about educational achievement, earnings, and employment experiences. In this paper, we make use of household and child characteristics data as well as school attendance data among 5-15 year old boys and girls in six study villages. The distinguishing feature of this survey is that many of the social and economic data were based on a *complete enumeration* of all households in these villages.

2.1. Data Description

These six villages taken together capture a good deal of the diversity present in rural West Bengal. The study villages are drawn from different agroclimatic regions of West Bengal - four villages from southern Bengal and two from North Bengal. Being located in different districts, they display interesting regional variations even within the state (see Pal, 1999 for a more detailed description of these villages). Bhagabandasan, in the Medinipore district of southern Bengal, is the most prosperous of the study villages while Simtuni is the poorest (Table 1A). Kalmandasguri is the only village with a significant Muslim population.⁸ All villages except Magurmari (which is close to some centres of traditional industry such as indigenous cigarettes making) are predominantly agricultural. Though there are primary schools in all the study villages, access to high schools is difficult in

participation and derive implications for gender differences in child school enrolment for the Indian state of West Bengal.

⁶ Please note that this equally applies to many existing studies, justifying their focus on some arguments of gender differences.

⁷ The survey was undertaken by Amartya Sen and Sunil Sengupta and funded by the World Institute of Development Economics Research (WIDER); that is why we refer to this data-set as the WIDER data-set.

⁸ Religion in WIDER classification includes two broad categories, namely, Hindus and Muslims. The caste variable, however, takes into account of the further division among the Hindus including the upper caste Hindus, scheduled castes and scheduled tribes while there is no caste division among the Muslims .

some villages like Kalmandasguri, Simtuni and Kuchly (Table 1B). There are also significant differences in the adult (aged above 15 years) male and female participation⁹ and wage rates among the study villages (Table 1C). Except the tribal dominated village of Simtuni, the male labour force participation rate is always significantly higher than the corresponding female participation rate. As with participation rates, female wage rates are generally lower than the male wage rates in most study villages though the extent varies: the difference is maximum in the north Bengal village of Magurmari and the minimum in the most prosperous southern village of Bhagabandasan.

Focusing on children aged 5-15 years, there are 548 male and 493 female children in our sample. Among these children, as high as 43% of the boys and 53% of the girls have never attended schools. Even when we consider the children aged 10-15 years, still about 32% of the boys and 45% of the girls have never been enrolled in schools. Thus there is some evidence of late school enrolment, though the gender difference in enrolment widens with age.

Considering a child's current market participation¹⁰ in relation to their current school participation, one can classify children into four categories: (a) only school participation (b) only market participation; (c) both school and market participation and (d) neither. Explicit participation in market jobs is rather limited in our sample. Most children fall into categories (a) and (d). In particular there are 529 children in category (a) and 445 children in category (d). Thus there are only 67 children in our sample who participate in some form of farm/non-farm work and may or may not go to school (i.e., categories (b) and (c)). Among these four categories of activity, an interesting case is (d) where children neither go to school nor explicitly participate in any market jobs. However, this does not rule out their informal participation in farm/non-farm work in family homes, which is not observed in our sample. This makes it difficult to take account of the role of the opportunity cost of schooling, which is not reflected in their market participation. In our analysis we shall capture this in terms of the sibling composition variables (see further discussion in section 3).

On an average the children who did not attend schools were from poorer families. There is also a gender difference noted here: average per capita expenditure for the female sample was Rs. 968 as opposed to Rs. 863 for the male. Similarly, among the children attending schools the averages were Rs. 1546 and Rs. 1398 respectively for the female and male children. Interestingly, average household income is higher for the female sample, irrespective of whether they are going to school or not. Also differences in parental literacy make a difference for boys and girls in the sample. For example, 71% of male children (as against 59% of girls) who were enrolled in primary

⁹ This is defined as the total days worked in a year for a wage in/outside the village. Since these figures are averages for the village male/female members, it would average out the gender difference in market participation attributable to female participation in non-market activities.

schools had *literate father*; in contrast, 69% of *female* children (as against 45% of boys) enrolled in schools had *literate mother*. Sibling composition too seems to be important such that boys with older brothers (as opposed to older sisters) were more likely to go to schools. This perhaps suggests that older brothers can relieve younger siblings of some family responsibility, say supplementing family earnings.

2.2. Methodology

Traditionally, the demand for schooling is derived from the Neo-Classical ‘common preference’ model of household behaviour where the household maximises the joint utility function of all its members (e.g., Becker and Lewis, 1965). This determines the quantity and quality of children, consumption of leisure and other market goods as well as household labour market participation decisions. An important indicator of child quality is child schooling which is main focus of this paper. Child schooling is however closely related to child’s participation in labour markets and in our simplified framework¹¹ decisions regarding child’s participation in schools and market jobs are determined jointly by the maximisation of the present discounted value of the family’s expected income net of costs of child schooling. Given that siblings born to same parents are expected to be of equal ability, investment in child schooling will depend on parental preferences, parental resources, returns to and opportunity costs of schooling.

The indicators of child schooling and labour market participation in our analysis are SCH and WORK respectively. The variable SCH (WORK) is one if the i -th child, $i = 1, \dots, n$ is currently attending a primary school (currently participating in farm or non-farm market jobs, full/part time) and zero otherwise. Suppose the following relationships hold:

$$\text{SCH} = 1 \text{ if } Y_1 = \beta_1 X_1 + \varepsilon_1 > 0 \text{ and } \text{SCH} = 0 \text{ otherwise}$$

$$\text{WORK} = 1 \text{ if } Y_2 = \beta_2 X_2 + \varepsilon_2 > 0 \text{ and } \text{WORK} = 0 \text{ otherwise}$$

where Y_1 , Y_2 are the latent variables for SCH and WORK respectively. Since both these variables, SCH and WORK, are dummy dependent variables, one may use univariate probit models to individually estimate them. However, since school enrolment and work participation are both endogenous and correlated, we jointly estimate these variables using a bivariate probit model, where ε_1 and ε_2 are jointly normally distributed with zero means, unit variances and a correlation coefficient

¹⁰ While we observe market participation of the sample children, we do not observe their participation in the family (farm or household work).

¹¹ Using a single cross-section data we assume, without much loss of generality, that the quantity of children, their birth order and parental labour market participation decisions are predetermined. Thus, we ignore the dynamics of fertility, consumption and labour market choices of parents and directly focus on household decisions regarding child

ρ . The set of covariates X_1 and X_2 will include similar characteristics of the child, his/her parents, household and the community s/he belongs to; but ideally there are also some different variables in the two equations.¹² Given the constraints posed by the available information, we choose these explanatory variables carefully so as to best reflect the hypotheses of our interest as indicated in the introductory section.

It is generally argued that benefits of education are lower for women in India (e.g., Kingdon, 1998). However, these estimates of male-female earnings differences are likely to be biased if one does not control for differential occupational choice of men and women and women's participation in other non-market activities. That is why Schultz (1993) suggests that a better variable to use would be the relevant wage rate. Since individual wage rates are endogenous to child schooling and labour, we include the village-specific average adult daily male (MDWAGE) and female (FDWAGE) wage rates. In an alternative specification, we also use the village-specific adult male (MPARTN) and female (FPARTN) participation rates and compare these two sets of estimates. Note, however, that these village level variables (participation or wage rates) may also reflect the relative prosperity of these villages and hence it may be difficult to disentangle the pure effect of returns to schooling in this context. Our attempt to include average village-level expenditure to control for the village-level prosperity was however unsuccessful because of the high degree of correlation between village-level wage/participation rates and average expenditure levels.

Since much of primary schooling is free in India, there is no significant gender difference in direct costs of attending schools. But the difference, if any, would reflect the differences in *opportunity costs of attending schools* for boys and girls. We have taken account of the most important component of opportunity cost in terms of child's market participation (and determine this jointly with child's school participation). However a large number of sample children neither go to school nor participate in any market jobs. These children are likely to participate in domestic farm/non-farm work though we do not observe that in our sample¹³. So following the general practice in this respect (e.g., Garg and Morduch, 1998), we argue that the sibling composition variables would take account of these implicit opportunity costs of schooling. In particular, we include two variables: if the child has any older brother (OLDB) and if the child has older sisters (OLDG). While more siblings may mean less parental resources per head, older siblings may supplement family resources and thus offer a greater opportunity of schooling for younger siblings.

school and work participation.

¹² These are the identifying variables that are present in one equation and not in the other. This is essential for the bivariate probit likelihood function to converge. This is further discussed later in the section.

¹³ Labelling this group of children as 'OTHER', we ran univariate probit regression of OTHER in terms of the same explanatory variables as in the WORK equation (see Tables 2, 3). Unlike the group of children participating in market

Moreover having older brothers may entail different benefits for schooling than having older sisters and these benefits may also differ between boys and girls. Among other things, this can be attributed to (a) allocation of family tasks between boys and girls (e.g., see Newman and Gertler, 1994) and (b) differential returns to boys' and girls' schooling. Thus, these sibling composition variables would take account of the interaction between household resources, parental preferences and opportunity costs of schooling.

Parental preferences are also important here. Since these are difficult to quantify, most empirical evidence in this respect is derived from the birth order and sibling composition variables (e.g., Butcher and Case, 1994; Garg and Morduch, 1998). Evidence from India tends to suggest that parents prefer boys over girls. Dasgupta (1987) uses variables indicating birth order and presence of older sisters while Kingdon (2002) uses a parental attitudinal variable towards gender equality in education. Since we do not observe similar parental attitudinal attribute, and also because these subjective measures are subject to serious biases, we argue that both sibling composition variables (OLDB, OLDG) and parental educational status variables (HEADLIT, HWIFELIT) would reflect parental attitudinal attributes. HEADLIT and HWIFELIT will also take account of differences in parental preferences, if any, in child schooling.

The household resource constraint arguably plays an important role in child schooling where children from better off households are more likely to obtain more and better schooling. In societies with pro-male bias this may also result in higher schooling for boys (relative to girls) from resource constrained households since it would maximise the benefit from investment in child schooling. We include per capita household expenditure (PCEXP)¹⁴ as an indicator of household long-term income. Since expenditure may be non-linearly related to schooling, we include the natural logarithm of PCEXP as the relevant income variable. Since household expenditure is endogenous to household decisions regarding child schooling and child labour, we use the value of expenditure per capita predicted (LNPCEXP) by characteristics of household head, household demographic composition, household assets and relevant village-level characteristics. In addition, we include the household's head's occupation and caste/religion as instruments of household's economic status. For example, we include if the household head is an agricultural labourer (HEADLAB) considered to be the poorest occupational group in these villages (Pal and Kynch, 2000). Since there is a close correspondence between caste and ownership of resources in rural India, it is expected that children from upper caste households (e.g., Hindu) will have higher schooling; in this respect, we include the variable if the household belongs to a Hindu family

jobs (WORK=1), those aged below ten years and from agricultural labourers' (poorest occupational group) family are more likely to belong to this group (other parameter estimates being similar to those of the WORK equation).

¹⁴ We experimented with three related variables, namely, household current income, expenditure and landholding per capita and obtained similar results. Here we present estimates using current expenditure since it is regarded as a better measure of long-term income in a rural setting with seasonal income fluctuations.

(HINDU). We also include if the household is headed by a female member (FHEAD). Female-headed households are often poorer since they may not have any adult male earning member, thus inducing children from these families to participate in market jobs. If, however, one believes that mothers have higher preferences for the well-being of their offspring, these female headed households may discourage children's participation in market jobs.

We control for age and gender of the child. Given that age may be non-linearly related to child schooling, we include a number of dummies to represent different ages of the sample children. In particular, we include AGE6, AGE7, AGE8, AGE9, AGE10, and AGE10 in the school enrolment equation (where children aged 5 years act as the reference group). Inclusion of these age variables not only reflects discrete non-linearity (and performs better than including age and age square), but also allows us to account for evidence of late enrolment, if any. However, for child's work participation equation we find that usually children aged 10 or more participated in farm/non-farm work and other age categories, even if included, were not significant in alternative specification. Hence we include only AGE10 in the WORK equation

so that children aged ten or less form our reference group. Gender dummy and gender interaction terms with other individual and household characteristics are included while doing the pooled regressions. These gender variables are naturally dropped when we run gender-specific regressions.

Finally, there are important inter-village variations that need to be accounted for (see Table 1A and Table 1B). In our final specification, this village-level variation will also be accounted for by the village-specific participation rates (or wage rates in the alternative specification) since other village-level characteristics turn out to be consistently insignificant¹⁵, when included along with participation or wage rates variables.

Thus we include very similar variables in the equations for participation in school and market work. But there are also some identifying variables. While we include only one age variable, namely AGE10 in the work equation, we include different age variables, namely AGE6, AGE7, AGE8, AGE9 and AGE10, in the school participation equation. This is because we find that children who are older than ten years are more likely to work in our sample. We have also experimented with other possible identifying variables, e.g., effect of household ownership of farm or other non-farm business on child's WORK. We tried two possible indicators of household ownership of farm, namely, landholding and if the head is an owner cultivator. Ownership of landholding in our data is however very closely correlated to the household expenditure per capita (in fact the predicted value of expenditure is derived using own land holding as one of the most important variables) and thus landholding does not have any significant impact on WORK when we also include predicted

¹⁵ Insignificance of many village-level variables may be attributed to the high and significant degree of correlation between /among them. For example, village-specific participation and wage rates are closely correlated with the

expenditure among other variables. Even when we include if the household head is an owner cultivator or not, it is never significant. We however do not have any direct information about the household ownership of non-farm business. We tried using some indirect measures derived from the household head's occupational codes, e.g., if the head is an agricultural labourer (HEADLAB), craftsman, trader/transporter or someone in agriculture-allied activities (e.g., fishery, poultry etc.). But none of these variables were significant in the work equation. Among all different occupational codes of the household head, only HEADLAB turns out to be significant in the work equation of male children in the bivariate probit specification. Insignificance of many of these variables may be attributed to the fact that there are very few children participating in market jobs in our sample.

2.3. Gender Decomposition in Enrolment

The Oaxaca-type decomposition method (Oaxaca, 1973) is normally used to analyse gender differences in the returns to schooling. We modify the standard Oaxaca decomposition method and apply it to the bivariate probit joint estimates of participation in school and market jobs. This is described below.

Suppose $\Pr(\text{SCH}, X_i, \theta_i^*)$ is the probability of school attendance for a typical individual characterised by X_i where θ_i^* is the set of maximum likelihood estimates of SCH for the i -th sample, $i = m, f$ for male and female samples respectively. Given the discrete nature of the dependent variables SCH and WORK, we can distinguish between four cases (a), (b), (c) and (d) as indicated in section 2.1. We use bivariate normal distribution to calculate these probabilities as follows:

$$\Pr(\text{SCH}=1, \text{WORK}=1) = \Pr [\varepsilon_1 > -X_1\beta_1, \varepsilon_2 > -X_2\beta_2] \quad (1a)$$

$$\Pr(\text{SCH}=1, \text{WORK}=0) = \Pr [\varepsilon_1 > -X_1\beta_1, \varepsilon_2 < -X_2\beta_2] \quad (1b)$$

$$\Pr(\text{SCH}=0, \text{WORK}=1) = \Pr [\varepsilon_1 < -X_1\beta_1, \varepsilon_2 > -X_2\beta_2] \quad (1c)$$

$$\Pr(\text{SCH}=0, \text{WORK}=0) = \Pr [\varepsilon_1 < -X_1\beta_1, \varepsilon_2 < -X_2\beta_2] \quad (1d)$$

Summing (1a) and (1b), we obtain the probability of enrolment:

$$\Pr(\text{SCH}=1) = \Pr(\text{SCH}=1, \text{WORK}=1) + \Pr(\text{SCH}=1, \text{WORK}=0) \quad (2a)$$

Similarly, summing up (1c) and (1d) we obtain the probability of non-enrolment:

$$\Pr(\text{SCH}=0) = \Pr(\text{SCH}=0, \text{WORK}=1) + \Pr(\text{SCH}=0, \text{WORK}=0) \quad (2b)$$

The expected probability of going to school is then given by:

distance of the village from the railway station, health centre, market centre as well as the secondary school. There is also a high degree of correlation between wage rates, participation rates and average expenditure at the village level.

$$S_m^* = \sum_{j=0}^1 \Pr(SCH = 1, WORK = j / X_m, \mathbf{q}_m^*) \quad (3a)$$

$$S_f^* = \sum_{j=0}^1 \Pr(SCH = 1, WORK = j / X_f, \mathbf{q}_f^*) \quad (3b)$$

Using these expected grades for male and female samples respectively, one can decompose the male-female differential in degree performance as follows:

$$\begin{aligned} S_m^* - S_f^* &= \sum_{j=0}^1 [\Pr(SCH = 1, WORK = j / X_m, \mathbf{q}_m^*) - \Pr(SCH = 1, WORK = j / X_f, \mathbf{q}_m^*)] + \\ &\quad \sum_{j=0}^1 [\Pr(SCH = 1, WORK = j / X_f, \mathbf{q}_m^*) - \Pr(SCH = 1, WORK = j / X_f, \mathbf{q}_f^*)] \\ &= \textit{Explained variation} + \textit{unexplained variation} \end{aligned} \quad (4a)$$

$$\begin{aligned} S_m^* - S_f^* &= \sum_{j=0}^1 [\Pr(SCH = 1, WORK = j / X_m, \mathbf{q}_f^*) - \Pr(SCH = 1, WORK = j / X_f, \mathbf{q}_f^*)] + \\ &\quad \sum_{j=0}^1 [\Pr(SCH = 1, WORK = j / X_m, \mathbf{q}_m^*) - \Pr(SCH = 1, WORK = j / X_m, \mathbf{q}_f^*)] \\ &= \textit{Explained variation} + \textit{unexplained variation} \end{aligned} \quad (4b)$$

Equations (4a) and (4b) are two alternative ways of decomposing the total variation in school enrolment into explained and unexplained components. In both equations (4a) and (4b), the explained variation (terms in the first summation in (4a) and (4b) respectively) holds the estimated parameters constant but allows gender-specific characteristics to vary. In other words, the explained variation, alternatively labelled as the 'endowment gap' by Cameron and Heckman (2001), is that part of the total variation attributable to the different characteristics of male and female children. The unexplained variation (terms in the second summation of (4a) and (4b)), however, holds sample-specific covariate characteristics constant, but allow the parameters to vary. This is the conventional 'discrimination' component or 'behaviour gap' (Cameron and Heckman, 2001), attributable to the different treatment of male and female children in the households. Generally, the size of the unexplained variation is taken to be a measure of gender discrimination. However, the whole of the unexplained variation cannot be attributed to gender discrimination alone, as the inclusion of some unobserved or imperfectly observed variables in the regression equations may also contribute to the unexplained variation but may not necessarily be related to discrimination between boys and girl's schooling (see further discussion in section 3.2).

3. EMPIRICAL RESULTS

We start our analysis by considering the univariate and bivariate probit results for the pooled regressions with gender interaction terms. Since the bivariate correlation coefficient is significantly different from zero, bivariate estimates are preferred to univariate estimates. These correlated bivariate estimates show evidence of significant gender difference in schooling with respect to individual's age (AGE6, AGE9, AGE10, AGE10), and parental literacy (HEADLIT and HWIFELIT). There is also significant gender difference in work participation with respect to AGE10 and HEADLIT. Hence we proceed to estimate separate male/female bivariate probit regressions for primary participation in school and market work.

3.1. Joint Estimates of Child Schooling and Child's Market Participation

The rest of the paper focuses on the correlated bivariate probit estimates of SCH and WORK for boys and girls (see Table 2 and Table 3). The correlation coefficient between the unobserved residual terms in the two equations is significant for both male and female children in our sample. As expected, the relationship is such that higher market participation entails lower school participation among both boys and girls in our sample.

Child characteristics: Relative to children aged five years, the likelihood of school enrolment is significantly more among boys aged 6 years and above and girls aged 7 years and above. Thus there is some gender difference with respect to AGE6 in that girls are likely to be enrolled about a year later than boys. However, both boys and girls aged more than ten years are more likely to participate in work.

Sibling characteristics: Sibling characteristics are important although their effects differ between male and female children in our sample. For example, having older brothers enhances the probability of schooling and lowers the probability of work among boys though it does not significantly affect the probability of school enrolment or market work participation among girls. The effect of having older sisters is however insignificant for both boys and girls in our sample.

Parental and other household characteristics: There is some evidence of significant resource constraints in child schooling and labour market participation decisions in our sample. Controlling for all other factors, both male and female children from less wealthy households are less likely to be enrolled in primary schools. In contrast both boys and girls from less wealthy households are more likely to work.

Maternal (HWIFELIT) and paternal (HEADLIT) education however affects schooling of boys and girls differently. Mother's literacy is insignificant for boys though it significantly enhances the

probability of school enrolment among girls. Father's education however significantly encourages boys' schooling only and does not have any perceptible impact on girls.

Among other household characteristics, a child from an upper caste Hindu (HINDU) family will have a greater likelihood of going to school relative to those from lower caste Hindu or Muslim households. Thus caste cannot explain gender differences in school or work participation. However, whether the household is headed by a female member (FHEAD) or whether the household head is an agricultural labourer (HEADLAB) does not significantly affect school participation among boys or girls.

Village-level characteristics: We had included two sets of village-level variables in the two alternative specifications, namely, male-female participation rates (FPARTN and MPARTN, Table 2) and male-female daily wage rates (FDWAGE, MDWAGE, Table 3).¹⁶ The likelihood of girls' school participation increases if the female market participation rate in the local economy is higher. However, higher local male participation rates enhance the likelihood of boys' market participation and thus indirectly lower male schooling. Results of the alternative specification using village-level daily wage rates are shown in Table 3. Though compared to Table 2, we generally obtain similar results, there are some differences with respect to the indicators of returns to schooling. While girl's schooling does not respond to local female wage rates, higher male wage rates significantly encourage boys' work and this in turn lowers their school participation. Thus boys' schooling responds more to local wage rates while girls' schooling to local participation rates. While it is common for boys to participate in market work, girls' market participation is influenced more by the local socio-cultural practices as reflected in the local female participation in market work rather than local female wage rates. Either way there is some confirmation that indicators of returns to schooling affect gender differences in school participation in our sample.

3.2. Gender Decomposition in Enrolment

Here we examine the implications of the bivariate probit estimates for explaining gender differences in school participation. In particular, based on equations (1a) to (1d) on page 11 and using bivariate probit parameter estimates (for male and female children as shown in Table 2), we calculate the predicted probabilities of enrolment and non-enrolment for boys and girls, depending on whether they are also participating in some market jobs. This is done for the following cases. (i) Male students using estimated parameters obtained from the male equation, (ii) male students using estimated female parameters, (iii) female students using estimated female parameters and (iv)

¹⁶ We have also tried including other village level characteristics, e.g., distance of the village from the nearest railway station, market centre, health centre in both school and work equations. None of these other variables are significant in any specification. Also see footnote 15.

female students using estimated male parameters. These results are summarised in Table 4, which highlight the pronounced gender difference in school enrolment.

The probability of school enrolment and non-enrolment with market participation is rather low for both the male and female children in our sample. So we focus on the probability of school enrolment and non-enrolment *without* market participation. Let us first consider the children who neither go to school nor participate in any work. The actual probability for this category is 0.48 for female and 0.38 for male. The corresponding predicted probability for female using female parameters is 0.48 and 0.39 for male using male parameters. If, however, we use female parameters, the probability of non enrolment among boys increases to 0.52. On the other hand, the probability of female non-enrolment decreases to 0.36 if male parameters are used instead.

When we consider the children who go to school, but do not participate in work, the predicted probability is 0.47 for female using female parameters and 0.55 for male using male parameters. If however, we use female parameters to predict male probability for this case, the probability of no-schooling (and no work) increases from 0.39 to 0.52 and that of school enrolment falls from 0.55 to 0.43. Similarly, when we use male parameters to predict female probability, the distribution mirrors that of the male sample: the probability of school enrolment increases from 0.47 to 0.53 and that of no-schooling decreases from 0.48 to 0.36.

Next, based on equations (2a) and (2b), we calculate the probability of schooling and no schooling for these four cases as shown in rows (3) and (6) respectively of Table 4. In particular, entries in row (3) are obtained by summing up the corresponding entries in rows (1) and (2). Similarly entries in row (6) are obtained by summing up the corresponding entries in rows (4) and (5). Finally, following equations (4a) and (4b) of section 2.3, we calculate the explained and unexplained variation of the gender differences in school enrolment and find that both equations yield the same result. In each case, the explained variation is 0.03 and the unexplained variation is 0.07, making the total variation to be 0.10. In other words, the explained variation in child schooling is only 30% while the unexplained variation is 70% of the total variation. Thus, even after including most established arguments of gender differences in school enrolment, a significant proportion of the total variation remains unexplained. The latter can primarily be attributed to different treatment of male and female students in terms of different regression functions, usually labelled as 'discrimination'. The size of this unexplained variation cannot however constitute a test of the discrimination hypothesis since this large unexplained variation can also be attributed to many unobserved and imperfectly observed factors (e.g., variables to account for a child's participation in domestic work or parental preferences) and/or child and household specific unobserved heterogeneity, which may affect gender differences and yet may not necessarily constitute discrimination.

4. CONCLUDING COMMENTS

There are significant gender differences in child schooling in the Indian states though there are very few attempts to explain gender differences in child schooling in the country. Moreover, none of the existing studies take account of the important opportunity costs of schooling. This paper uses a correlated bivariate probit model to jointly determine child's participation in school and market work among 5-15 year old boys and girls in rural West Bengal in eastern India and examine the factors determining gender difference in child schooling. We consider several possible causes of gender differences including differential returns to schooling, household resource constraint, and nature of parental preferences and also child's implicit opportunity costs of domestic work. Finally, we use these estimates to decompose the total variation in observed gender differences in current school enrolment into explained and unexplained parts.

The analysis is based on the WIDER data from six villages in West Bengal, which shows significant inter-village variation. Our results suggest that indicators of returns to schooling (instrumented by local participation and wage rates), opportunity costs of participation in domestic work (instrumented by sibling composition), parental preferences (instrumented by parental literacy levels), household expenditure and interaction between and among these arguments are important determinants of current school enrolment of boys and girls in our sample. Despite its robust significance on school enrolment, the predicted value of household expenditure has similar effect on enrolment of both boys and girls and hence cannot explain the gender differences in enrolment. However, sibling composition, parental education, local adult work participation and daily wage rates are found to explain a part of the observed gender differences in enrolment. First, girls are more likely to be enrolled if the local adult female participation rate is higher while boys are more likely to work if the local male participation rate is higher. If, however, we use local daily wage rates, likelihood of boys' market participation increases with higher local male wage rates while female participation does not respond to local female wage rate. Either way there is some support in favour of the gender differences in returns to schooling. Second, boys with older brothers are more likely to be enrolled though the variable is insignificant for girls. Third, paternal and maternal education significantly encourages boys' and girls' enrolment and in a differential manner: while father's education favourably affects boys' schooling, mother's education is essential for girls' schooling only.

Recent research unequivocally suggests the significant beneficial effects of women's education for fertility and child health outcomes. Thus unequal treatment of women in access to schooling is not only unfair for its own sake, but is also socially undesirable. Even after accounting for the major arguments of gender differences in child school enrolment, only about one third of total gender difference in schooling is explained by the characteristics of boys and girls in our sample.

Thus there remains a significant unexplained variation, which can partly be attributed to the different male-female regression functions, commonly labelled as 'gender discrimination'. A part of the observed unexplained variation is however likely to be attributable to the use of imperfect instruments ϕ household resources, opportunity costs of schooling in terms of participation in domestic work, parental preferences, and/or unobserved individual/household level heterogeneity, which may not constitute discrimination.

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TABLE 1A. SELECTED SOCIO-ECONOMIC CHARACTERISTICS OF THE STUDY VILLAGES

Variables	Study Village					
	Kuchly	Sahajpur	Bhagaban-dasan	Simtuni	Kalman-dasguri	Magurmari
District	Birbhum	Birbhum	Medinipur	Purulia	Kochbehar	Jalpaiguri
Household no.	142	227	134	75	89	99
Fsize[1]	6.90 (3.7)	6.70 (3.06)	5.48 (3.37)	6.55 (2.09)	7.04 (1.8)	6.04 (3.00)
Female(%)	55.6	49.7	38.9	51.3	57.5	57.8
Landless(%)	46	58	29	3	47	-
S. Caste (%)	38.4	37.5	15.5	1.3	33.4	2.6
S. Tribe (%)	11.7	22.3	11.8	86	8.3	1.13
Muslim (%)	-	-	-	-	40.8	-
Literacy [2]	0.38(0.29)	0.40 (0.30)	0.66 (0.55)	0.10 (0.01)	0.52 (0.39)	0.35 (0.23)
Landreform(%)	45	25	37	17	15	-
Pnfinc	0.21	0.53	0.35	0.32	0.31	0.65
PCINC [3]	1647 (881)	1545 (643)	2213 (1085)	1160 (292)	1212 (448)	1441 (669)
Modal wage	3.42	3.30	3.60	2.57	2.66	-
Poverty	40.4	52.3	16.5	62.5	72.7	56.6

Note: Fsize: family size; Female: Average proportion of female members; Landless: % of landless households; Land reform: % of household who have gained from the land redistribution programme. Pnfinc: proportion of income earned from non-farm activities while PCINC is mean income per head measured in rupees; Modal wage: Kilogram of rice per day in 1988; Poverty: % of households below poverty line. [1]: Numbers in the parentheses show the corresponding standard deviations. [2] Female literacy in the parenthesis.[3] Numbers in the parentheses show the corresponding standard deviations.

TABLE 1B. INTER-VILLAGE VARIATION IN PUBLIC ENVIRONMENT

Village	Distance from the nearest facility(in km.)				
	Railway Station	Pucca Road	Health centre	High School	Market Centre
Bhagabandasan	6	0	6	1	0
Magurmari	4.5	0	2.5	2.5	2.5
Kalmandasguri	9.5	3	3	33	3
Simtuni	68.0	0	3	30	2
Kuchly	18	3	8	18	18
Sahajapur	8	0	3	8	8

TABLE 1C. INTER-VILLAGE VARIATION IN LABOUR MARKET INDICATORS

	Adult wage rates		Adult participation rates		expenditure pc
	Mean (sd)		Mean (sd)		Mean (sd)
	Male	Female	Male	Female	All
Bhagabandasan	14.90 (0.18)	14.80 (1.7)	0.42 (0.19)	0.29 (0.12)	1628.1 (776.4)
Magurmari	11.00 (3.7)	5.44(1.9)	0.72 (0.16)	0.53 (0.19)	1036.5 (445.0)
Kalmandasguri	11.40 (2.5)	8.50 (1.6)	0.58 (0.23)	0.36 (0.21)	1273.2 (579.9)
Simtuni	13.82(2.4)	12.77 (3.5)	0.26 (0.18)	0.23 (0.11)	1137.0 (400.6)
Kuchly	13.56 (2.8)	13.25 (0.88)	0.57 (0.29)	0.18 (0.30)	1424.7(702.2)
Sahajapur	12.71 (3.2)	11.86 (3.7)	0.54 (0.25)	0.23 (0.20)	1368.6 (638.3)

Note: Per capita expenditure for Simtuni is not available and hence we use an per capita income in stead.

TABLE 2. BIVARIATE PROBIT ESTIMATES OF SCHOOLING AND WORK

Variable	Male				Female			
	SCH		WORK		SCH		WORK	
	Coefficient	T-ratio	Coef	T-ratio	Coef	T-ratio	Coef	T-ratio
Constant	12.74	5.258**	12.23	1.942*	-10.08	4.533**	7.63	1.234
Age 6	1.5	3.450**	-	-	0.43	1.121	-	-
Age 7	1.68	3.551**	-	-	1.68	3.958**	-	-
Age 8	2.45	5.394**	-	-	1.67	3.984**	-	-
Age 9	2.9	5.740**	-	-	1.42	3.194**	-	-
Age 10	2.57	5.549**	-	-	1.39	3.262**	-	-
Age greater than 10	2.62	5.725**	1.96	4.977**	1.51	3.856**	0.91	2.532*
Older brother	0.27	1.726*	-1.8	3.190**	0.05	0.259	-0.79	1.491
Older sister	-0.01	0.067	0.18	0.378	0.16	0.905	-0.26	0.507
Instrument of expenditure	1.38	4.063**	-2.12	2.181*	1.11	3.583**	-1.46	1.724*
Female head	0.20	0.576	-0.33	0.261	0.62	1.129	-0.22	0.322
Head literate	0.79	5.039**	-0.75	2.348*	0.20	1.202	0.26	0.745
Head's wife literate	-0.14	0.682	-0.21	0.374	0.36	1.934*	0.06	0.141
Head agricultural labour	-0.2	1.117	-0.48	1.665*	-0.32	1.521	0.18	0.474
Hindu	0.54	2.878**	-0.39	1.096	0.58	3.182**	-0.07	0.182
Male participation rate	0.55	0.736	4.39	2.925**	-0.3	0.354	1.19	0.759
Female participation rate	0.44	0.528	-5.9	4.220**	1.43	1.733*	-0.44	0.238
RHO(1,2)	-0.43	2.457*			-0.46	2.308*		
LOG-L	-311.7976				-302.1518			
OBS	548				493			

Note: ** denotes that the corresponding variable is significant at 10% or lower level while * denotes the same at 1%.

Dependent variable: SCH: 1 if the child has currently been attending a primary school; WORK: 1 if the child is currently participating in farm and/or non-farm work in the market Explanatory variables: AGE6 : 1 if the child is 6 years old; AGE7: 1 if the child is 7 years old; AGE8: 1 if the child is 8 years old; AGE9: 1 if the child is 9 years old; AGE10: 1 if the child is 10 years old; AGE10+: 1 if the child is more than 10 years old; OLDB: 1 if the child has older brothers; OLDG : 1 if the child has older sisters; LNPCEXP : Natural logarithm of per capita household expenditure (instrumented); FHEAD : 1 if the head of the household is female; HEADLIT : 1 if the head of the household is literate; HWIFELIT : 1 if the wife of the head of the household is literate; HEADLAB : 1 if the head of the household is an agricultural labour; HINDU : 1 if the household belongs to upper caste Hindu category; FPARTN: Village-level female participation rates ; MPARTN : Village-level adult male participation rates. FDWAGE: Average local daily female wage rate; MDWAGE: Average local daily male wage rate.

**TABLE 3. BIVARIATE PROBIT ESTIMATES OF SCHOOLING AND WORK
(ALTERNATIVE SPECIFICATION)**

Variable	Male				Female			
	SCH		WORK		SCH		WORK	
	Coefficient	T-ratio	Coefficient	T-ratio	Coefficient	T-ratio	Coefficient	T-ratio
Constant	-12.51	5.253**	5.33	1.015	-9.68	4.410**	-	-
Age 6	1.48	3.401**	-	-	0.39	1.005	-	-
Age 7	1.6	3.346**	-	-	1.62	3.927**	-	-
Age 8	2.42	5.311**	-	-	1.62	3.918**	-	-
Age 9	2.87	5.650**	-	-	1.35	3.050**	-	-
Age 10	2.53	5.432**	-	-	1.32	3.208**	-	-
Age > 10	2.59	5.681**	1.91	5.325**	1.45	3.809**	0.91	2.537*
Older brother	0.26	1.573	-1.72	2.909**	0.06	0.305	-0.83	1.659*
Older sister	-0.005	0.031	0.15	0.351	0.14	0.787	-0.19	0.373
Instrument of expenditure	1.38	4.073**	-1.3	1.689*	1.12	3.613**	-1.35	1.650*
Female head	0.20	0.600	0.15	0.133	0.68	1.212	-0.29	0.413
Head literate	0.8	5.046**	-0.85	2.792**	0.22	1.284	0.19	0.546
Head's wife literate	-0.087	0.419	-0.36	0.848	0.36	1.934*	0.10	0.224
Head agricultural labour	-0.26	1.438	-0.28	1.048	-0.38	-1.857*	0.31	0.815
Hindu	0.54	2.885**	-0.34	1.027	0.59	3.221**	-0.13	0.337
Male wage rate	0.23	1.434	0.59	1.679*	0.09	0.633	-0.46	1.143
Female wage rate	-0.23	1.590	-0.47	1.443*	-0.11	0.837	0.40	1.069
RHO(1,2)	-0.39	2.323*			-0.44	2.101*		
LOG-L	-317.5265				-301.8728			
OBS	548				493			

Note: ** denotes that the corresponding variable is significant at least at 10% while * denotes the same at 1%. Also see note to Table 2 on regression variables.

TABLE 4. ACTUAL AND PREDICTED PROBABILITY OF SCHOOL ENROLMENT

	Actual probability		Predicted probability from bivariate probit regression			
	Female	Male	Female using female equation	Male using male equation	Female using male equation	Male using female equation
(1) No schooling and no work	0.48	0.38	0.48	0.39	0.36	0.52
(2) No-schooling and work	0.01	0.03	0.04	0.05	0.06	0.04
(3) No-Schooling	0.49	0.41	0.52	0.44	0.42	0.56
(4) Schooling and no work	0.47	0.54	0.47	0.55	0.53	0.43
(5) Schooling and work	0.04	0.05	0.01	0.03	0.02	0.08
(6) Schooling	0.51	0.59	0.48	0.58	0.55	0.51

Note: These predicted probabilities are calculated using the parameter estimates shown in Table 2 and the underlying bivariate normal distribution.