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Is Child Work Necessary?*

SONIA BHALOTRA

Department of Economics, University of Bristol, Bristol, UK
(e-mail: s.bhalotra@bristol.ac.uk)

Abstract

This article investigates the hypothesis that child labour is compelled by poverty. It shows that a testable implication of this hypothesis is that the wage elasticity of child labour supply is negative. Using a large household survey for rural Pakistan, labour supply models for boys and girls in wage work are estimated. Conditioning on non-labour income and a range of demographic variables, the article finds a negative wage elasticity for boys and an elasticity that is insignificantly different from zero for girls. Thus, while boys appear to work on account of poverty compulsions, the evidence for girls is ambiguous.

I. Introduction

Why do children work? A common but not undisputed perception is that child work is compelled by the constraints of household poverty (e.g. see Basu and Van, 1998; US Department of Labor, 2000). Although the geographical distribution of child workers today and the economic history of specific regions demonstrate a negative association of child work and aggregate income (e.g. Krueger, 1996; Basu, 1999), looking at aggregate data does not establish a clear causal role for household poverty. Growth in aggregate income may be unequally distributed, with little increase in the incomes of households that supply child labour. Growth may nevertheless be associated with a reduction in the incidence of child labour

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because it is associated with the development of new technologies, the expansion of legal and political infrastructure or the evolution of social norms. Microdata are needed to disentangle household living standards (a microeconomic variable, which differs across households) from other factors like new technology, new laws or changed norms, which apply across households. Further, in order to test the hypothesis that it is binding poverty constraints that compel child labour, we need to consider not just income effects but also price effects.

A seminal theoretical paper that captures the role of poverty is Basu and Van (1998). These authors assume that children work only when subsistence constraints bind and then focus on how this sort of supply behaviour can result in multiple equilibria in the labour market, with striking policy implications. Although some previous studies have found negative income effects on child labour, this only establishes the normality of child leisure, or confirms that households are credit constrained. No previous research has directly tested the assumption that children work because of subsistence poverty.[1]

The hypothesis of interest in this paper, namely that extreme poverty compels child labour, has important implications for policy design. For instance, under this hypothesis, trade sanctions or bans on child labour will tend to impoverish the already very poor households supplying child labour.[2] Also, the force of any interventions in the education sector is likely to be limited unless they also lower the opportunity cost of sending a child to school. Since the marginal utility of consumption increases very rapidly as people get close to subsistence, creating matching increases in the marginal return to education may not be in the scope of policy. Thus, if subsistence poverty drives child labour then reducing school fees or improving school quality (as in the Back-to-School Programme in Indonesia) may have little impact, while policies that compensate families for taking their children out of work and putting them in school will tend to work. Many actual interventions are consistent with a model of binding poverty, e.g. the Food-for-Education Programme in Bangladesh (Ravallion and Wodon, 2000), Progresa in Mexico (Skoufias and Parker, 2001), Bolsa Escola and PETI in Brazil (World Bank, 2001).

This article proposes a test of poverty compulsions, and investigates it for data on children in wage work in rural Pakistan. The basic idea is straightforward. Suppose that children work because their households are very poor in the specific sense that income exclusive of child earnings falls below subsistence requirements, so that child work is *necessary*.[3] Then children will appear to work towards a target income, which is the shortfall between subsistence needs and other income. In this case, an increase in the wage will induce a reduction in child labour. In section III this intuition is

[1]Ray (2000) sets out to test the hypothesis of poverty compulsions but the test relies upon picking up an income effect on child labour, albeit at a threshold level given by a selected poverty line.

[2]Trade sanctions and the setting of international labour standards, issues that have provoked considerable public interest (see Basu, 1999; Bhalotra, 1999).

[3]This is precisely what is meant by the (lay use of the) term *necessary* in the title of the article. It is not to be confused with the definition of goods with income elasticities less than one as necessities.

formalized and it is shown that if child labour is compelled by poverty then the child wage elasticity is negative. The test is easily generalized to the more prevalent scenario of children working on household-run farms (or enterprises); see Bhalotra and Heady, 2003; Dumas, 2004. In these cases, the wage may be unobserved but can be proxied by the marginal product of labour. As marginal labour productivity is increasing in land acreage, children working under poverty compulsions will tend to work fewer hours on larger plots of land.

The data used are from a household survey for rural Pakistan, a region where child labour participation is high, child wage labour is unusually prevalent and there is a striking gender differential in education and work. Labour supply equations are estimated separately for boys and girls, conditioning on a rich set of demographic variables and a life cycle-consistent measure of the child's non-labour income. The main result is that the wage elasticity of hours is significantly negative for boys and insignificantly different from zero for girls. So it seems that boys work in order to help their households meet subsistence needs, but that girls work even when not 'necessary'. Consistent with this, I also find that the effect of household income on child work hours is much smaller for girls than for boys. The results stand up to a number of robustness checks.

This article is organized as follows. Section II describes the data. The theoretical model is developed in section III. Section IV describes the translation of the theory into a fairly general empirical model. The main results are presented in section 5.1, where the gender difference is discussed. Section 5.2 presents a range of specification checks, and section VI concludes.

II. Data and non-parametric statistics

Data, definition of children, definition of work

The data used are from the Pakistan Integrated Household Survey (PIHS) gathered by the World Bank in conjunction with the Government of Pakistan in 1991. Employment questions are put to all individuals 10 years or older. The data show that the proportion in school falls gradually after the age of 11 and exhibits a sharp drop from 31% at age 17 to 17% at age 18. For this reason, age 17 is a data-consistent cut-off. The analysis is therefore carried out for *10–17 year old* children. The 3,373 children in this age group come from 1,543 households, in 151 clusters. The main conclusions are not altered if the definition is narrowed to 10–14 year old children: results using this definition are available in an earlier version of this paper (Bhalotra, 2000).

By International Labour Organisation (ILO) conventions, work is defined as effort that results in a marketable output. This is reported in the survey under two subcategories: wage work (for which wage earnings are recorded) and work on household-run farms and enterprises (for which there is no explicit remuneration). Individuals are classified as participating in work if they report having worked at least

1 hour in the week preceding the survey. Hours of work are recorded for this preceding week. The survey also provides an estimate of the annual average of weekly hours of work, which smooths over seasonal fluctuations. The latter is the definition of hours adopted in this analysis, although the results are robust to using the other definition (section 5.2).

A profile of child labour in Pakistan

The data show a high prevalence of child labour, a remarkable gender gap in schooling and 'inactivity', and a substantial fraction of children engaged in (market) wage work (see Table 1).[4] The sample probabilities of participating in wagework are 8% for boys and 7% for girls.[5] Boys in wage employment work an average of 31 hours a week, the average for girls being much smaller at 9.5 hours a week. There is considerable variation around the mean values (see Figure A1), which I exploit in estimating the wage elasticity.

Choice of sample

The analysis is conducted separately for boys and girls in wage work in rural areas, where child labour and poverty are most prevalent. Wage work involves longer hours than other sorts of work, and virtually rules out school attendance. In contrast, own-farm labour is more compatible with schooling (see Table 1), and provides a relatively secure return to experience to the extent that children work on land that they are likely to inherit. Isolating wage work allows me to concentrate on the hypothesis of poverty compulsions, and to avoid the confounding influence of substitution effects that arise in own-farm work.[6] It is also convenient to look at explicitly waged work, although, as discussed in section I, the basic principle underlying the test can be extended to the case of children in farm work. The available sample size is further restricted by analysing data on hours conditional on participation. But this is, of course, essential to the question at hand since the wage elasticity can only be negative at positive hours of work. Many previous studies of child labour estimate participation equations pooling urban and rural data for boys and girls, and aggregating wage and non-wage work, but my investigations suggest that the implied pooling restrictions are invalid.

[4]Inactivity is likely to include unreported domestic work. Consistent with this, rates of 'inactivity' recorded for men and women over the age of 17 are 18% and 54% respectively.

[5]These participation rates are high for a rural economy where self-employment dominates wage employment. Comparing these figures for children with the corresponding figures for adults puts them in perspective. Amongst adults 18 years and older, only 36% of men and 15% of women are wage workers.

[6]Unlike cash, land has a dual role. It denotes wealth which can be used to buy child leisure (or education), but it is *also* a productive asset that can be used to employ child labour. In several developing country settings, this produces the remarkable result that children from richer households (i.e. households with big farms) work more. This apparent paradox can be resolved by incorporating into the model imperfections in land and labour markets. See Bhalotra and Heady (2003).

TABLE 1

A profile of child activities

	Boys	*Girls*
Total participation rates		
Wage work	8.0 (31.4)	7.1 (9.5)
Household farm work	24.5 (18.0)	25.0 (8.4)
Household enterprise work	3.4 (31.2)	1.4 (19.0)
Work (any of the above)	33.4	31.0
School	65.6	25.9
None of the above activities	11.8	44.5
Per cent that combine types of work		
Household farm and enterprise work	0.95	0.12
Household farm and wage work	1.4	2.3
Household enterprise and wage work	0.17	0.06
Per cent that combine work and school		
Household farm work and school	10.7	1.9
Household enterprise work and school	0.73	0
Wage work and school	0.39	0.13
Number of children	1,786	1,587

Notes: Rural Pakistan, 10–17 year old children. All values are percentages except values in parentheses which refer to the annual average of weekly hours of work conditional on participation.

Non-parametric relation of hours and wages

A simple locally weighted regression (a lowess smooth) of hours on the wage rate is in Figure A2. The graph reveals a monotonically negative relation for boys. For girls, the curve is flatter and non-monotonic. The results that emerge from the more structural analysis to follow are consistent with these data.

III. A theoretical framework

In deriving a test of the hypothesis that poverty compulsions drive children into work, I abstract from other reasons why children may work (for a useful discussion, see Basu and Van, 1998 who make a similar abstraction). Assume, in line with the literature on human capital and child labour that children do not bargain with their parents because they do not have a valid fallback option. The problem is for parents to select the optimal level of child labour.[7] For simplicity, assume the household has one parent and one child. Section 3.1 works with the Stone–Geary utility function, which is used in Basu and Van (1998). Section 3.2 shows that a similar testable prediction flows from the less restrictive and more commonly used constant elasticity

[7]This raises a potential agency issue. Parent altruism has been challenged in some historical and anthropological accounts of child labour (e.g. Parsons and Goldin, 1989). Altruism is investigated for the sample of households analysed in this article in Bhalotra (2004), and the data found consistent with altruism.

of substitution (CES) function. Indeed the prediction that is taken to the data is intuitive and flows from even more general models.

Although the models presented here are static, they may be thought to correspond to the second stage of a two-stage budgeting problem in an intertemporal model (e.g. Blundell and Walker, 1986). The implication of this for the empirical model is that the measure of non-labour income used is life cycle consistent (see section 4.1). In common with previous research on child labour, parent labour is not explicitly modelled. However, in the empirical model, I allow for the endogeneity of parental earnings.[8]

3.1. The Stone–Geary case

Let household preferences be represented by the Stone–Geary utility function:

$$\begin{aligned} u(C, L_i) &= \{(C-S)(L_i), \quad \text{if } C \geq S \\ &= \{(C-S), \qquad \text{if } C < S \end{aligned} \tag{1}$$

where C is joint consumption, S is the subsistence level of consumption and L_i denotes the non-work time of the child, which includes leisure and time spent at school. It is assumed that $C \geq 0$, $0 \leq L_i \leq 1$, and $S > 0$ is a parameter. The Stone–Geary function implies a larger 'weight' on child leisure (L_i) in richer households [which have a larger value of $(C - S)$]. It is assumed that parents (denoted by subscript j) always work. A convenient normalization is to set the time endowment to unity so that we can define child hours of work as $H_i = (1 - L_i)$ and write $H_j = 1$.

The budget constraint is

$$C = N + w_j + w_i(1 - L_i) \equiv Y + w_i(1 - L_i), \tag{2}$$

where the price of consumption is normalized to unity, w is the real wage, subscripts i and j denote child and parent, respectively, N is household non-labour income and the *child's non-labour income* is defined as $Y \equiv N + w_j$ (recall that $H_j = 1$).

Consider the case where $(Y + w_i) < S$. Then equation (2) implies that even if the child works the maximum possible hours so that $H_i = 1$, consumption, $C = Y + w_i$, falls below subsistence, S. Hence, the second segment of the utility function (1) applies, and the optimal solution is $H_i = 1$ and $C = Y + w_i$. If, however $(Y + w_i) \geq S$, then $C \geq S$ is achievable and the maximum utility is attained within the first segment, where the utility function is $U = (C - S)L_i$. Using equation (2) to eliminate C in this function, we can maximize $(Y + w_iH_i - S)(1 - H_i)$ to get optimal child hours of work:

$$\begin{aligned} H_i \equiv (1 - L_i) = \left(\frac{(S-Y) + w_i}{2w_i}\right) &\equiv \theta \quad \text{if } 0 \leq \theta \leq 1 \\ &= 0 \quad \text{if } \theta < 0 \\ &= 1 \quad \text{if } \theta > 1. \end{aligned} \tag{3}$$

[8]Child labour is explicitly modelled as a function of parental labour in Bhalotra (2002).

If we rewrite θ as $[(1/2) + (S - Y)/2w_i]$, we can see that if $S > Y$, then $H_i > 0$ or, as expected, if the child's non-labour income does not cover subsistence needs then the child works. There is some Y greater than S at which the child stops working or $H_i = 0$.

The wage elasticity of labour supply implied by equation (3) for the case $0 < \theta < 1$ is

$$\varepsilon_{Hw} \equiv \frac{\partial H_i}{\partial w_i}\frac{w_i}{H_i} = \frac{-(S - Y)}{w_i + (S - Y)}. \tag{4}$$

Using equation (3), it is clear that if desired hours, $\theta > 0$, then $w_i + S - Y > 0$ and the sign of equation (4) is the same as the sign of the numerator. Thus, the Stone–Geary case yields the strong prediction that the wage elasticity of child hours of work is negative *if and only if* $S > Y$, i.e. the child's non-labour income does not cover household subsistence needs. In other words, subsistence poverty implies that the child's wage elasticity of work hours will be negative *and*, also that observing a negative wage elasticity indicates subsistence poverty (i.e. compelling poverty).

3.2. The CES case

Consider the CES utility function

$$U = [\alpha(C - S)^\rho + (1 - \alpha)L_i^\rho]^{1/\rho}, \tag{5}$$

where $0 < \alpha < 1$, $\sigma = 1/(1 - \rho) > 0$ is the elasticity of substitution between consumption (C) and child leisure (L_i). In the special case where $\sigma = 1$ (i.e. $\rho = 0$), utility is given by the Cobb–Douglas function: $(C - S)^\alpha L_i^{(1-\alpha)}$. The Stone–Geary function derives from equation (5) when $\sigma = 1$ and $\alpha = (1 - \alpha)$. Thus the CES function allows the elasticity of substitution a greater range and it allows the weight on child leisure in the utility function to differ from the weight on consumption.[9]

Optimization of equation (5) subject to the budget constraint, equation (2), gives the marginal rate of substitution condition:

$$w_i = \frac{1 - \alpha}{\alpha}\left(\frac{C - S}{L_i}\right)^{1/\sigma}. \tag{6}$$

Using equation (2) in equation (6) to eliminate C, optimal hours of child work, $H_i \equiv (1 - L_i)$, can be written as[10]

[9]The relative weight on child leisure (or child consumption) in the utility function of the parent indicates the degree of parent altruism.

[10]When H_i lies outside [0,1] it is restricted to 0 or 1. Consider $H_i > 1$. In equation (7) this implies numerator > denominator or $S > Y + w$. As pointed out in the Stone–Geary case, this is uninteresting because even with the child working to a maximum ($H_i = 1$), the family does not survive. For a sufficiently large Y, desired $H_i < 0$ and then of course we observe $H_i = 0$.

$$H_i = \frac{\alpha^\sigma w_i^\sigma - (1-\alpha)^\sigma (Y-S)}{\alpha^\sigma w_i^\sigma + (1-\alpha)^\sigma w_i}. \tag{7}$$

It is straightforward to confirm, as a check, that setting $\sigma = 1$ and $\alpha = (1 - \alpha)$ gives θ in equation (3). Without imposing these restrictions, however, the wage elasticity implied by equation (7) is given by:

$$\varepsilon_{Hw} = \frac{(1-H_i)}{w_i} \frac{[\sigma(Y-S) + (\sigma-1)w_i H_i]}{w_i - (S-Y)}. \tag{8}$$

The denominator of the second term is non-negative as survival is only feasible if the maximum possible consumption exceeds subsistence needs or $Y + w_i \geq S$. As $(1 - H_i)$ and w_i are both positive, the sign of the wage elasticity is the sign of $\sigma(Y - S) + (\sigma - 1)w_i H_i$. This depends upon whether subsistence poverty binds ($S > Y$), and on the value of σ (see Figure 1, adapted from Barzel and McDonald, 1973[11]). Consider the case of interest, $S > Y$. If $\sigma \leq 1$, the elasticity is negative. If $\sigma > 1$, the labour supply curve is *forward-falling*, that is, it is negatively sloped at *low wages* and positively sloped at high wages [note that w_i weights the second term in the numerator of equation (8); see graph IX in Figure 1]. As the high-wage case is implausible by context, we can conclude that, *if* $S > Y$ *then* $\varepsilon_{Hw} < 0$. This establishes the hypothesis.

To summarize, this section has shown that if subsistence constraints bind, then the wage elasticity is negative. Thus, if we observe a negative wage elasticity, we can conclude that the data are consistent with the hypothesis that working children come from households in which subsistence needs exceed the child's non-labour income, or that the child's income is necessary for survival. Note that these models do not preclude children from working even in households whose subsistence needs are met without the child's income contribution; however, in such households, the wage elasticity of hours will be non-negative.

IV. The empirical model

This section discusses translation of the theoretical model into an estimable model, subject to constraints imposed by the data. Mean values and standard deviations of the variables used are given in Table A1. Section 4.1 defines the variables and 4.2 describes the estimation techniques used.

4.1. The variables

Labour supply and wages

The dependent variable, H_i, is hours in wage work conditional on participation. Since the decisions to participate in waged and other (self-employed) work are

[11]While their analysis is different from that in this article in many respects and it is not about child labour, the article by Barzel and McDonald has in common with this article that it is concerned with characterizing the manner in which the level of non-labour income influences the shape of the labour supply curve (the sign of the wage elasticity) in a model with subsistence constraints.

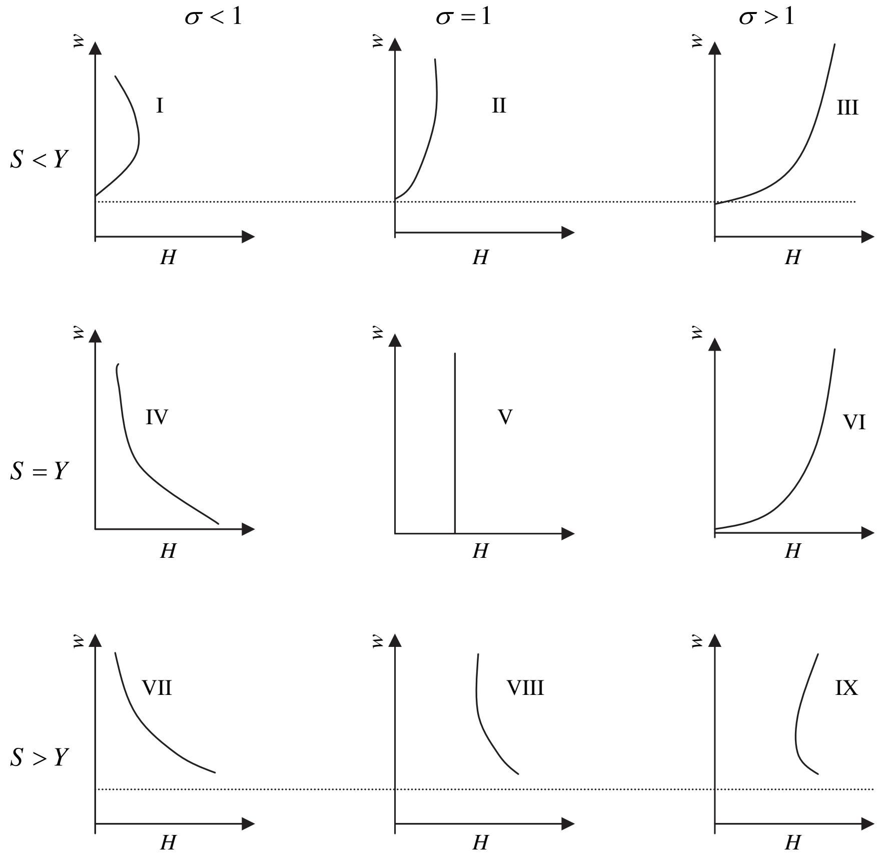

Figure 1. Labour supply curves. H is hours of work, w is the wage, Y is the child's non-labour income, S is the subsistence level of consumption, σ is the elasticity of substitution of consumption with respect to leisure. These graphs are adapted from Barzel and McDonald (1973).

probably taken simultaneously, hours in wage work can be written as a function of hours in other work.[12] As other (own-farm or enterprise) work is endogenous, it is substituted out by its exogenous determinants – land ownership, acreage, tenancy-type and demographics.[13]

[12]See the conditional time allocation model set out in Datt and Ravallion (1994). They model the individual's decision to work in a public employment programme jointly with their decision to participate in other work. The structure of my problem is similar.

[13]An alternative approach might, in principle, be to include selection–correction terms estimated from a multinomial logit model. However, the assumption in this model that the alternatives are independent is not appropriate in the current analysis since children combine waged work with work on household farms or enterprises.

I condition on participation because it is only at the intensive margin that the wage elasticity can be negative.[14] As children who participate may have unobservable characteristics that are correlated with the unobservables in the hours model, generating a potential sample selection bias, the inverse Mills ratio (λ) estimated from the participation equation is included as a regressor in the hours equation (see Heckman, 1974). To increase the robustness of this procedure to the assumed parametric distribution of the unobserved error terms, λ^2 is also included in the model. This approach rests on the semiparametric series estimator principle of Newey, Powell and Walker (1990). Identification in this context often relies upon functional form alone. Here, I further exploit the fact that children who participate in wage work do not, at the same time, participate in school (see Table 1). Thus dummy variables for the presence of a primary and secondary school respectively are used as instruments in the work-participation equation. Conditional upon participation, these are expected to have no effect on hours of work. Identification is potentially strengthened by virtue of there being a number of exogenous regressors that are significant in the participation equation and insignificant in the hours equation (compare Table A2 with Tables 2 and 3).

In common with most studies of labour supply, the wage rate, W_i, is measured as earnings divided by hours of work; division bias is discussed in section 4.2.[15]

Non-labour income

The *child's non-labour income*, *Y*, was defined in equation (2) as $Y = N + \sum_j w_j H_j$, where N is household non-labour income and $\sum_j w_j H_j$ denotes adult earnings. The measure of N that is used here has attractive features in terms of both interpretation and measurement. It is a life cycle-consistent measure (see Blundell and Walker, 1986). To see this, note that the intertemporal budget constraint that defines the time path of assets is: $A_{t+1} = (1 + r)A_t + \sum_k w_{tk} H_{tk} - C_t$, where $k = i$ (child), j (adults), A is assets, r is the interest rate, and t denotes time. This can be written as $rA_t - \Delta A_{t+1} = C_t - \sum_k w_{tk} H_{tk,} = N$ [using equation (2)]. In a one-period model, $N = rA_t$. The asset change term, ΔA_{t+1}, allows N to capture savings and dissavings between periods. The static measure is only valid if agents are myopic or if there

[14]See equation (3), which gives θ, the desired hours of work. It shows that $H > 0$ if and only if $w > Y - S$, from which it follows that the wage elasticity of participation is positive. A tobit specification is avoided because: (i) we cannot assume that the parameters – including the wage coefficient – are common to the participation and hours decisions; and (ii) estimation of the wage elasticity from a tobit would require predicting the wage for the approximately 90% of children who are not in wage work using the relatively thin sample of those who are.

[15]Measurement of earnings is complicated by some payments being made in kind and by earnings being reported for different payment frequencies. With guidance from Living Standards Measurement Survey (LSMS) staff at the World Bank, these were brought to a common denominator and payments in kind were incorporated using cluster-level grain prices and information on quantities of grain received. A dummy variable was included in the estimation to indicate observations for which the wage was imputed. This was insignificant and so was not retained.

exist no capital markets,[16] and it tends to generate wage elasticities that confound the effects of shifts in wage profiles with movements along them (see Blundell and MaCurdy, 1999). The lifecycle-consistent measure, Y, can be obtained from cross-sectional data as the difference between consumption expenditure (C) and total labour earnings.[17] It follows that the child's non-labour income is measured as the difference between C and child earnings:

$$Y = N + \sum_j w_j H_j = C - w_i H_i, \tag{9}$$

where j denotes adults and i denotes child.

Thus, household poverty is defined to include adult earnings, non-labour income and any saving or dissaving performed by the household. The logarithm of per capita non-labour income is included in the model (after dropping the one household for which Y takes a negative value). Rather than worry about equivalence scales, I take the more general approach of including household demographics as independent regressors in the equation. These are specified as household size and the proportion of household members in a full set of age–gender categories.[18] Province dummies are included to capture cross-sectional variation in prices.

Exogenous variables

Since an alternative use of child time is in work on the household farm or enterprise, the vector of exogenous conditioning variables [X in equation (10) below] includes *acres of land owned* by the household which, at given household size, reflects the marginal productivity of farm work.[19] To allow for a possible non-linearity at zero acres (62% of sample households own no land), a *dummy for land ownership* is included. To capture any effect of land tenancy type, dummies for whether the household *rents or sharecrops land* are included. A dummy for *female headship* is included to allow for heterogeneity in preferences between men and women, and for any vulnerability of the female-headed household that is not reflected in income. The *education level of both parents* is included to capture preferences for education and the efficiency of household production of human capital (e.g. Behrman *et al.*, 1999). The cluster-level *unemployment rate* (calculated by aggregation of individual responses) is included to allow for disequilibrium in the labour market (see Ham, 1986).

[16]While formal capital markets are underdeveloped in the rural areas of most low-income countries, there is considerable evidence of informal means of saving and dissaving. In the data used in this paper almost 50% of households reported borrowing or lending money.

[17]Consumption expenditures reported in the survey include imputed values for home-produced consumption.

[18]The average household size is 8 of whom, on average, 4.4 are over the age of 14, 2.6 are under the age of 10, 1.85 are aged 10–17 and one is in the age range 10–14.

[19]Rosenzweig (1980) presents formal models of labour supply in landholding and landless rural households and underlines the importance of conditioning on farm size when analysing wage labour.

I also include a quadratic in child *age*, dummies for *religion* of the household head and an *indicator for whether the child is the child of the household head*. This allows for differential treatment of nephews, siblings or other relations of the head. As this was insignificant in the equation for boys and girls, it was dropped. Child illness was similarly investigated but dropped because it was insignificant (and potentially endogenous).

The estimated equation

The semilog-linear functional form is chosen and expected to provide a local linear approximation to a range of more complex functions:

$$H_i = \alpha + \beta \ln W_i + \gamma \ln Y + \delta X_i + e_i, \tag{10}$$

where H is hours in wage work, ln W is the log of the wage earned by the child, ln Y is the log of the non-labour income of the child, X is a vector of exogenous variables and e denotes unobservables and measurement errors. Quadratic terms in ln W and ln Y were investigated. The log–log form was also investigated but the results were similar to those obtained with the semilog form.

4.2. Identification and estimation

Identification

Wage. Calculation of the wage as the ratio of earnings to hours tends to introduce 'division bias', a spurious *negative* correlation between the wage and hours, which is bigger, the bigger the measurement error in hours. However, my estimates show that the wage coefficient becomes *more negative* when instrumented, indicating that division bias is limited. There remains the possibility that the coefficient on the child wage is subject to endogeneity bias if unobservables like tastes for work are (positively) correlated with the wage. It is unlikely that this bias is sizeable given that the child's preferences will not influence the outcome when the decision is made by the parent. Parental preferences for school vs. child labour are likely to matter. If these are sufficiently captured by the included indicators of parental education, then again, the endogeneity issue is naturally limited. I nevertheless, test exogeneity and present both ordinary least squares (OLS) and instrumental variables (IV) estimates.

The individual child wage (W_i) is instrumented using the agricultural wage rate for children at the community level, its square and its inverse (these data are available from the village file). To allow for within-community variation, an interaction term between the community wage and the child's age is also included in the instrument set. In order to rule out the possibility that the community level wage for children is high in areas where relatively few children are willing to work (i.e. that it reflects supply rather than opportunities or demand), I investigated a simple community-level regression relating the child wage to the participation rate.

The estimates indicate an insignificant relation. I also investigated the use of the child's completed schooling as an instrument for the child wage as, in the wider literature on (adult) labour supply, education is a commonly used instrument for the wage. I found that schooling is rejected by a test of over-identifying restrictions, casting doubt on its frequent use in the literature (consistent with Pencavel, 1986), who argues that an individual's education is likely to be correlated with their tastes for work).

Income. If income includes child earnings, there is scope for positive feedback from child hours to household income. This is not a problem here since income is the non-labour income of the child [see equation (9)]. An instrument for Y is nevertheless sought, so as to allow for the possible joint determination of parent and child labour supply, and for measurement error, known to afflict measures of income in rural economies.[20] The instruments are an indicator for whether the community has a shop,[21] a quadratic in the community-level wage rate for men, and interactions of this wage with the education of the father and mother, and the acreage of farmland that they own. The interactions allow within-community variation. In principle, local labour market conditions that determine community wages may also influence work-hours. This is unlikely in the estimated model since I control for the community-level unemployment rate. Nevertheless, to ascertain that the results are not driven by community-level variation, the community-level average of Y was included in the model. The main results were unchanged.[22] As income is often used as an instrument for expenditure (e.g. Blundell *et al.*, 1998, Attanasio and Lechene, 2002) and Y in this article is expenditure net of child earnings, I also investigated adding a quadratic in reported household income to the instrument set. However, as the validity of these additional instruments was only marginal using the C-statistic, they were not retained.[23]

Estimation

Estimation is by the two-step efficient generalized method of moments (GMM) estimator although, for comparison, traditional IV (2SLS) estimates are also presented. GMM is more efficient than 2SLS and robust to heteroskedasticity of unknown form, as well as to arbitrary intracluster correlation (see Wooldridge, 2002, p. 193). GMM also tends to perform better than 2SLS when the correlation between

[20]These considerations are often neglected. For example, Rosati and Rossi (2004), Emerson and Souza (2003) and Ray (2000) treat income net of child earnings as exogenous.

[21]A number of indicators of community infrastructure were investigated and an indicator for presence of a village shop was the only powerful variable in the first stage.

[22]Attanasio and Lechene (2002) also use the community-level wage as an instrument for household expenditure.

[23]The C-statistic or the difference in Sargan statistic allows a test of a subset of the orthogonality conditions.

the instruments and the explanatory variables is weak. Baseline OLS estimates using Cragg's estimator[24] are also reported. The standard errors of all reported estimates are adjusted for non-independence within sampling clusters and amongst siblings.

V. Results

Section 5.1 presents the main findings, section 5.2 discusses the gender differences that emerge, and section 5.3 investigates robustness of the results to alternative specifications.

5.1. The main results

For main results refer Tables 2 and 3. There is considerable variation in hours around the mean, a good deal more than is typically observed for adult hours of work in industrialized nations (Figure A1). The OLS estimates explain about 54% of this variation for boys and 32% for girls. The inverse mills ratio (λ), which corrects for selection into wage work, is significant and positive in the equation for boys, consistent with expectation. It is insignificant, although positive, in the equation for girls. The square of λ was insignificant in both equations, and so it was dropped.

The wage elasticity

The wage elasticity is significantly negative for boys. At the sample mean of hours of work, the GMM estimate is −0.53 and the OLS estimate is −0.33. Thus, the hypothesis that boys work on account of poverty compulsions cannot be rejected. The estimates suggest that if a boy's wage rate drops, he works longer hours to make up the loss in earnings. Conversely, if his wage rate increases, rather than exploit the higher marginal reward for effort on the wage labour market, he works less. Girls exhibit a wage elasticity that is insignificantly different from zero in each of columns 1–3. The GMM estimate is −0.13 and the OLS estimate is 0.00095. Although we cannot strictly reject the hypothesis that girls' work is compelled by poverty, the evidence is ambiguous. For example, a wage elasticity of zero is also consistent with the hypothesis that parents are selfish if selfishness is defined as sending a child to work the maximum feasible hours, irrespective of the marginal return (the wage). For corroboration of the result that household poverty explains child labour amongst boys but that other explanations may play a greater role for girls, we return to the raw data: we find that working boys come from poorer families than working girls (Table 4),[25] and a nonparametric graph of hours against wages (Figure A2) is consistent with the econometric results. Further discussion of the results is in the next section.

[24]This is more efficient than standard OLS in the presence of heteroskedasticity of unknown form (Davidson and MacKinnon, 1993, pp. 599–600).

[25]This comparison also makes it implausible that the negative wage elasticity observed for boys represents a backward-bending labour supply curve (see section 3.2).

TABLE 2

Hours equation for boys

Estimator	*OLS*	*2SLS*	*GMM*
Ln child wage	**−10.44**	**−18.93**	**−16.72**
	(5.31)**	(3.00)**	(3.05)**
Ln household income	**−7.44**	**−24.59**	**−19.80**
	(2.90)**	(1.33)	(1.24)
Inverse mills (participation)	38.96	48.50	48.92
	(2.13)*	(2.13)*	(2.37)*
Child age	34.90	49.94	48.53
	(2.35)*	(2.42)*	(2.49)*
Child age squared	−1.01	−1.51	−1.46
	(2.14)*	(2.20)*	(2.23)*
Ln household size	−0.89	−1.07	−1.52
	(0.74)	(0.74)	(1.19)
Proportion males <10	13.80	−28.87	−9.56
	(0.63)	(0.53)	(0.20)
Proportion females <10	−9.99	−53.10	−43.41
	(0.59)	(1.05)	(0.98)
Proportion males 10–17	−5.96	−17.89	−18.68
	(0.24)	(0.73)	(0.80)
Proportion females 10–17	16.80	11.71	7.29
	(0.72)	(0.45)	(0.28)
Proportion females 18–59	−38.03	−75.93	−77.09
	(1.40)	(1.42)	(1.56)
Proportion males 60+	130.79	120.67	145.70
	(3.09)**	(2.13)*	(2.85)**
Proportion females 60+	−159.64	−132.10	−111.99
	(2.26)*	(1.48)	(1.37)
Age of household head	−0.09	0.00	−0.12
	(0.51)	(0.01)	(0.53)
Father's age	−0.60	−0.76	−0.49
	(1.12)	(1.08)	(0.73)
Mother's age	1.00	1.17	0.93
	(1.70)	(1.48)	(1.22)
1 (female head)	17.06	10.31	15.80
	(1.86)	(0.97)	(1.66)
Father's school years	−1.94	−1.03	−1.18
	(1.34)	(0.54)	(0.66)
Mother's school years	10.49	13.20	13.13
	(4.60)**	(2.00)*	(2.16)*
1 (household owns land)	−16.66	−18.96	−20.64
	(2.63)**	(2.22)*	(2.55)*
1 (household rents land)	3.51	9.10	4.78
	(0.43)	(0.82)	(0.49)
1 (household sharecrops land)	−11.86	−15.33	−14.09
	(2.00)*	(2.19)*	(2.12)*

continued overleaf

TABLE 2

(continued)

Estimator	*OLS*	*2SLS*	*GMM*
Acres of land owned	1.46	1.84	2.22
	(1.14)	(0.68)	(0.87)
Acres squared	−0.03	−0.04	−0.05
	(1.20)	(0.71)	(0.91)
1 (non-muslim head)	3.45	9.88	12.19
	(0.33)	(0.66)	(0.86)
Cluster unemployment rate	−14.01	−16.03	−15.55
	(0.46)	(0.36)	(0.38)
R^2 (uncentred; centred)	0.83; 0.54	0.74; 0.29	0.79; 0.41

Notes: The dependent variable is hours of wage work conditional on participation. Province dummies are included (not shown). Omitted case is Punjab, only Sindh is significant, and positive. Robust *z*-statistics in parentheses, *significant at 5%; **significant at 1%. The number of observations is 130. The Hansen–Sargan J statistic is $\chi^2(7) = 5.8$ with a P-value of 0.56. This is a test of the joint null hypothesis that the excluded instruments are valid. It is distributed as χ^2 with degrees of freedom equal to the number of overidentifying restrictions (see Davidson and McKinnon, 1993, pp. 235–236). The first stage R^2 is 0.24 in the participation probit (pseudo-R^2), 0.43 in the wage equation and 0.41 in the income equation.

OLS, ordinary least squares; 2SLS, traditional instrumental variables; GMM, generalized method of moments.

Income effects

Household income has a negative impact on boys' work, the GMM elasticity being −0.63. The fact that the OLS estimate is significant but the IV estimate is not most likely reflects the low power of the instruments in a small sample. The impact of income on girls' hours of work is much smaller and insignificant, consistent with the finding that their labour is not compelled by poverty. These income effects are, recall, conditional upon parental age and education, land ownership and participation, although it is not uncommon in the literature on child labour to find small or insignificant income effects, even on participation (see Basu and Tzannatos, 2003; Bhalotra and Heady, 2003; Bhalotra and Tzannatos, 2003; Rogers and Swinnerton, 2004).

Conditioning on participation, the variation in hours is not very sensitive to many of the other included variables. This is also the case for estimates of hours equations for high-income countries (e.g. Heckman, 1994).

5.2. Discussion of gender difference in results

The findings presented in the previous section are corroborated by the finding of similar gender-specific results for the other main form of child labour in the Pakistan data, namely, labour on household-run farms (see Bhalotra and Heady, 2003). In particular, substitution effects associated with land ownership dominate wealth effects for girls, but not for boys. We observe that girls in households that own relatively large plots of land are both more likely to work and less likely to attend

TABLE 3

Hours equation for girls

Estimator	*OLS*	*2SLS*	*GMM*
Ln child wage	**0.009**	**−2.65**	**−1.21**
	(0.01)	(0.84)	(0.42)
Ln household income	**1.16**	**4.68**	**1.92**
	(0.71)	(0.80)	(0.39)
Inverse mills (participation)	2.46	5.67	−3.97
	(0.39)	(0.80)	(0.65)
Child age	−9.44	−3.23	−7.56
	(2.06)*	(0.44)	(1.24)
Child age squared	0.35	0.13	0.27
	(2.01)*	(0.49)	(1.22)
Ln household size	0.386	0.149	0.894
	(0.60)	(0.17)	(1.36)
Proportion males <10	4.19	18.73	−22.87
	(0.21)	(0.55)	(0.97)
Proportion females <10	−11.93	−10.50	−24.91
	(1.00)	(0.68)	(2.14)*
Proportion males 10–17	−35.24	−21.34	−35.02
	(2.57)*	(1.03)	(1.92)
Proportion females 10–17	−17.32	−5.69	−27.33
	(1.11)	(0.20)	(1.23)
Proportion females 18–59	2.20	2.81	−27.71
	(0.10)	(0.10)	(1.39)
Proportion males 60+	−30.77	−12.56	−4.87
	(1.16)	(0.41)	(0.17)
Proportion females 60+	−41.99	−22.28	−51.49
	(1.52)	(0.52)	(1.33)
Age of household head	0.083	0.107	−0.010
	(0.65)	(0.62)	(0.08)
Father's age	0.040	0.118	−0.050
	(0.22)	(0.55)	(0.27)
Mother's age	0.33	0.15	0.20
	(1.58)	(0.61)	(0.88)
1 (female head)	5.65	−0.13	6.59
	(1.35)	(0.02)	(1.10)
Father's school years	−0.11	−0.08	0.20
	(0.36)	(0.22)	(0.68)
Mother's school years	18.42	6.68	17.69
	(1.93)	(0.48)	(1.47)
1 (household owns land)	−2.20	−4.75	−2.40
	(0.68)	(1.03)	(0.55)
1 (household rents land)	0.12	8.35	−0.42
	(0.04)	(0.93)	(0.06)
1 (household sharecrops land)	0.35	2.79	−0.44
	(0.17)	(0.68)	(0.14)

continued overleaf

TABLE 3

(continued)

Estimator	*OLS*	*2SLS*	*GMM*
Acres of land owned	−1.02	−1.24	−1.06
	(0.61)	(0.69)	(0.79)
Acres squared	0.006	0.046	0.030
	(0.06)	(0.45)	(0.41)
1 (non-muslim head)	6.61	7.37	14.43
	(0.78)	(0.96)	(2.14)*
Cluster unemployment rate	−1.64	−19.43	56.33
	(0.02)	(0.19)	(0.69)
R^2 (uncentred; centred)	0.66; 0.32	0.63; 0.26	0.61; 0.21

Notes: See notes to Table 2. The number of observations, is 106. Province dummies are included (not shown). They are insignificant. Hansen–Sargan J is $\chi^2(7) = 6.97$ with a P-value of 0.43. The first stage R^2 is 0.21 in the participation probit (pseudo-R^2), 0.32 in the wage equation and 0.57 in the income equation.

OLS, ordinary least squares; 2SLS, traditional instrumental variables; GMM, generalized method of moments.

TABLE 4

Gender comparison of mean household living standards for children participating in wage work

	Mean (boys)	*Mean (girls)*	*P-value of t-test*
Ln(child's non-labour income)	5.53	5.75	0.009
Per cent of households that own land	19.1	27.8	0.056
Years of schooling of father	1.18	1.74	0.036
Per cent of households with a female head	8.8	3.7	0.053

Notes: Mean values are over the sample of households with at least one boy or girl (respectively) engaged in wage work. The null hypothesis is that the gender difference in the means is zero. The *t*-test is against the alternative hypothesis that the mean for girls is 'better' than the mean for boys. In all cases except that of female headship, 'better' refers to larger. The tests confirm that households with boys in wage employment are poorer on average than households with girls in wage employment.

school than girls in households with smaller landholdings. This implies that the households using girl labour on their farms are not subsistence constrained.

Insofar as girls' work is not compelled by poverty, their participation in work and their very low participation in schooling may be explained by relatively low expected returns to schooling.[26] This changes the weight on non-labour time, L_i (which includes schooling) in the utility function (e.g. Basu and Tzannatos, 2003). A gender difference in the return to schooling may reflect not only unfairness and inflexibility in labour markets but also social norms relating to the age of marriage of girls, the propriety of women working, or their being unable to migrate for work. If

[26]Evidence of low market returns to schooling for girls in neighbouring India is provided in Kingdon (1998) and Behrman *et al.* (1999).

so, it is just as important to alter attitudes to girls' education, work and marriage, as it is to regulate employers. Regulation of employers is difficult in the informal sectors in which children work, but there is scope for change on the supply side (i.e. within households).

Although improving returns to school for girls is likely to raise girls' school attendance, its impact on the incidence of child labour is likely to be small at first. This is because additional enrolment is likely to derive initially from the substantial group of girls reporting inactivity (see Table 1). Indeed, this is what was found in analysis of a school subsidy programme in Bangladesh (Ravallion and Wodon, 2000).[27]

Although the relative returns argument is quite compelling, consider now an alternative explanation of the gender difference in terms of differences in the type of work that boys and girls engage in. Information in the PIHS indicates that girls in wage employment mostly work on a non-household farm, while boys in wage employment mostly report non-agricultural activity. Consistent with this, girls' work is more seasonal than that of boys. This is evident from comparison of hours worked in the week before the survey with the annual average of hours worked in a week, the two definitions referred to in section II. So the zero elasticity that girls exhibit may be determined by demand rather than supply conditions: they may work hard in peak agricultural seasons and have no option to increase their hours in response to lower wages in slack seasons. A natural way to investigate this possibility is to use the alternative definition, hours worked in the reference week. Results reported in Bhalotra (2000) show that we still find a negative wage elasticity for boys and an insignificant elasticity for girls. It therefore seems unlikely that seasonality in girls' work explains the main finding.

The different effects of income on boys and girls work are striking. Boys' leisure (or time at school) appears to be a normal good, while girls' non-work time is not. The same results were found for child labour on household-run farms: household income had a large and negative effect on boy labour, but no significant effect on girl labour (Bhalotra and Heady, 2003). These income effects are consistent with the wage effects, which suggest that girls' labour is not compelled by poverty. They are also consistent with parents in this sample being more altruistic towards boys than towards girls.[28]

Estimates for the other variables in the model are also quite different for girls and boys. For example, land ownership reduces the hours of boys but not girls, while belonging to a household where the head is not Muslim increases the hours of girls but not boys.

[27]If the high rates of 'inactivity' (i.e. neither in school nor in work) reported for girls and women reflect full-time engagement in domestic work, then whether any increases in school attendance amongst girls are associated with reduction in domestic work or in market work will depend upon their relative payoffs, on the scope for substitution of other labour, and on attitudes to market work.

[28]Bhalotra (2004), for example, shows that if good A has a larger weight in the parent's utility function than good B, then the marginal effect of parental income on A is larger than on B. The basic idea follows from Becker (1993).

5.3. Robustness and alternative specifications

This section discusses robustness of the results to the instrumenting strategy, to the age definition of children, and to the definition of work hours. It also considers whether demand and supply effects might be confounded, and how the results change if we average across children within the household. Robustness to conditioning variables including household size is investigated, and adult wage elasticities are presented and argued to reinforce the plausibility of the estimated elasticities for children.

The Hausman test cannot reject exogeneity of wages or income, so the preferred results are OLS.[29] In the boys' equation, the t-statistic associated with the residual estimated from an auxiliary reduced form wage equation is 0.58, and that associated with the income residual is -0.76 (and the F-test of the joint significance of the two residuals is 0.82 ($P > F = 0.45$). In the girls' equation, the corresponding test statistics are 1.03, 1.31 and $F = 0.99$ ($P > F = 0.38$). We nevertheless report IV results, to allow for measurement error. Tests reported in the tables indicate that we cannot reject the over-identifying restrictions.

Instruments for wages and participation (selection) have sufficient power. Consistent with expectation, IV makes the wage elasticity more negative. The instruments for income are weak. A number of checks were performed to confirm that this does not affect the plausibility of the main results. In particular, the wage elasticity is almost identical if I drop income and allow it to be proxied by the age and education of the mother and father, land ownership, community-level unemployment and household demographics. This is the preferred strategy in many previous studies. For example, Ravallion and Wodon (2000) replace income with land and parental education, while Thomas, Strauss and Henriques (1991) use land and education as instruments for income. Also, as suggested in Bound, Jaeger and Baker (1995), the plausibility of IV estimates can be confirmed by comparison with OLS estimates when the instruments are weak. And the OLS estimates here offer the same broad conclusions.

The choice made in defining children as 10–17 year olds was justified in section II. The equations were re-estimated restricting the sample to 10–14 year old children. Similarly, section II explained that the reported equations use the annual average of weekly hours of work. The equations were re-estimated defining hours, instead, as hours worked in the week before the survey. The main results were unaltered and are available in an earlier version of this paper (Bhalotra, 2000). The wage elasticity for boys is -0.33 and that for girls is close to zero.

The negative relation of work hours and the wage is unlikely to reflect labour demand rather than labour supply because these are individual level data, and demand effects are captured by province dummies and the village-level unemploy-

[29]Exogeneity of the wage could not be rejected in any specification. Exogeneity of income was rejected in a specification in which the land variables were excluded from the model. However, conditional on land size and tenancy type, the income variable appears to be exogenous.

ment rate. Also, as we have seen, the result for boys persists and is even stronger when the actual wage is replaced by the offered wage (compare columns 2 and 3 with column 1 in Table 2).

The results reported so far were obtained using data on individual children. Household-level estimates are obtained for comparison, with the dependent variable defined as average hours per child in the household and the child wage as an average weighted by hours. The wage elasticity for the average boy is −0.77, significant at the 1% level. The income elasticity is −1.29, significant at the 10% level. Averaging over girls in the household, both the wage and income coefficients are insignificant, as before.

To the extent that fertility is a choice variable correlated with investments in child human capital, household size, although it regularly appears on the right-hand side of equations describing child labour, is a potentially endogenous regressor. Its exogeneity was investigated using the *C*-statistic, and could not be rejected [$\chi^2(1) = 1.13$, $P = 0.29$ for boys, $\chi^2(1) = 0.018$, $P = 0.89$ for girls]. I have also confirmed that the wage and income elasticities for boys and girls are not significantly altered when household size is dropped from the model.

As some of the control variables are likely to be correlated with the key variables of interest (e.g. child age with child wage and acres of land owned with household income), a parsimonious equation was estimated with just the child wage and income as regressors. There was no significant change in the wage elasticity. The GMM elasticity for boys is −0.52 and for girls it remains insignificantly different from zero. The income elasticity is smaller in both equations.

Comparable estimates of the wage elasticity of child work hours are unavailable in the literature. Estimates of wage and income elasticities for adults in industrialized countries have exhibited a wide range, even when obtained using the same data. This variation has been shown to be associated with differing assumptions regarding functional form, selection, simultaneity or measurement error (see Mroz, 1987; Heckman, 1994). The way in which these issues are addressed in this article was discussed in section IV. Alternative functional forms were investigated; tests of the instruments were provided and results reported for alternative estimators.

There is some evidence of negative wage elasticities for adult men (see, e.g. Attanasio and MaCurdy, 1997 for the USA, Kooreman and Kapteyn, 1986 for the Netherlands), although these are typically found at high wage levels (backward bending labour supply curves). Negative wage elasticities *at low wages* (forward falling labour supply curves) have been found for rural India using data on adults (Rosenzweig, 1980). Looking at children here sharpens the question: if adult earnings constitute the larger share of household earnings and these contribute to the non-labour income of children, then we may expect a negative wage elasticity to be *less* likely to be observed for children than for adults. In other words, having found a negative wage elasticity for boys, we may expect to find a negative wage elasticity for their parents. This is indeed what is found. The wage elasticity is −0.57 for men [−0.35 if OLS] (as it happens, very similar to that for boys!) and −0.83 [−0.19 if OLS] for women.

VI. Conclusions

This article finds support for the assumption that poverty compels boys to work. This suggests that cash transfers offered to households supplying child labour will be effective in reducing child labour amongst boys. In the case of girls, the evidence is ambiguous. While the hypothesis that poverty compels girls to work cannot strictly be rejected, the results in this article and in related research suggest that girls may work even when poverty is not compelling, possibly because the perceived return to their education is relatively low. I also found that variation in non-labour income, which includes parental earnings, had no effect on girls' hours of work, which is consistent with parents attaching little weight to the education (or leisure) of their daughters. In contrast, the income effect for boys was fairly large. These results suggest that interventions designed to reduce child labour need to be gender specific. While poverty reduction may reduce work and increase schooling amongst boys, to achieve similar results for girls, it may be necessary to also raise the returns to schooling for girls, to do which it may be necessary to first alter attitudes to girls' education.

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Appendix

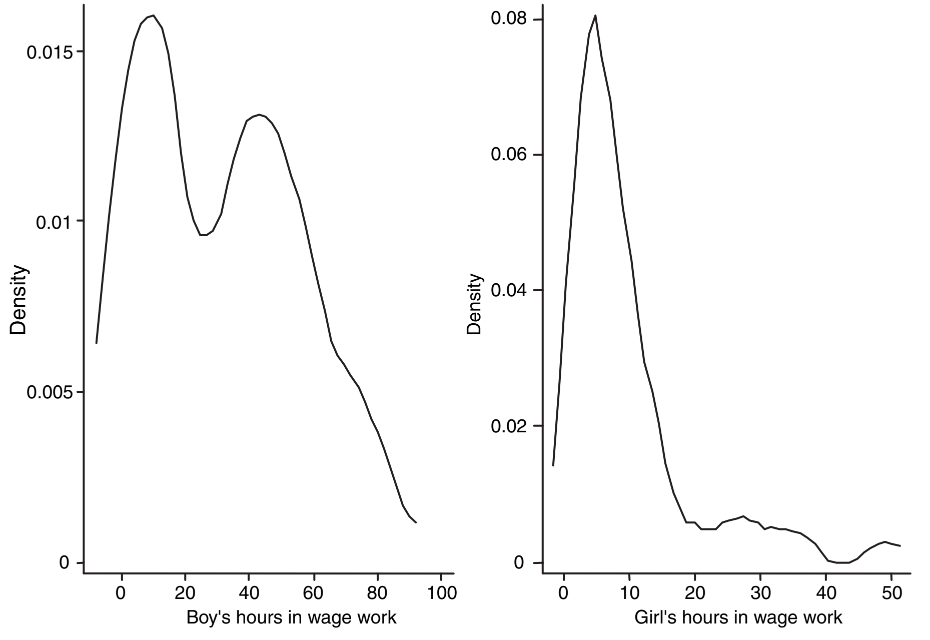

Figure A1. Hours of wage work conditional on participation: kernel density estimates

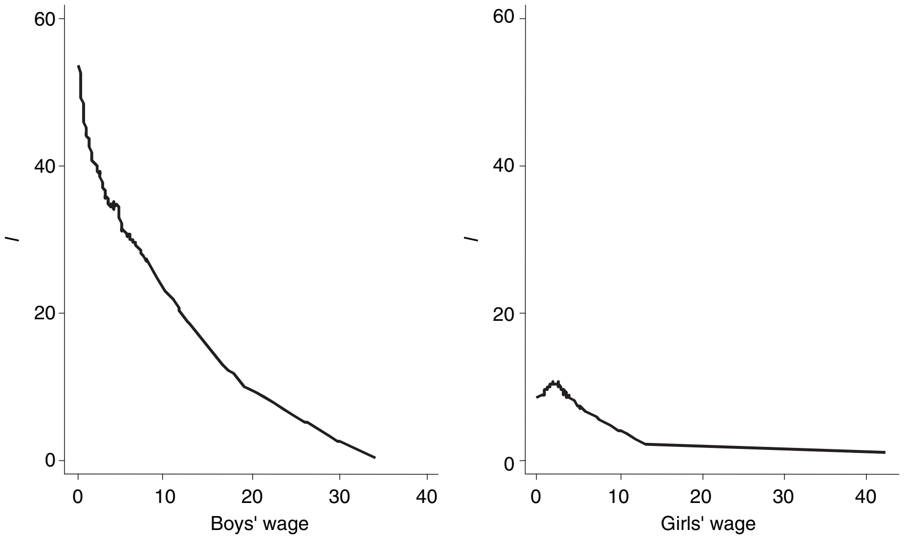

Figure A2. Hours against wages; children in wage work

TABLE A1

Means and SD of variables. Sample of children participating in wage work

Variable	*Boys, mean (SD)*	*Girls, mean (SD)*
Hours in wage work	32.46 (24.19)	9.34 (9.25)
Ln wage	1.31 (0.97	0.87 (0.91)
Ln non-labour income	5.53 (0.80)	5.74 (0.58)
Age	14.63 (1.92)	13.46 (2.21)
Household size	8.60 (2.92)	8.30 (2.65)
Proportion males <10	0.15 (0.12)	0.15 (0.12)
Proportion females <10	0.12 (0.12)	0.13 (0.12)
Proportion males 10–17	0.22 (0.11)	0.09 (0.09)
Proportion females 10–17	0.08 (0.10)	0.23 (0.10)
Proportion females 18–59	0.19 (0.10)	0.18 (0.09)
Proportion males 18–59	0.19 (0.11)	0.17 (0.09)
Proportion males >59	0.04 (0.07)	0.02 (0.06)
Proportion females >59	0.01 (0.03)	0.02 (0.05)
Age of head	46.71 (14.01)	46.82 (12.45)
Father's age	50.45 (9.81)	48.72 (8.89)
Mother's age	43.27 (8.43)	42.09 (7.42)
1 (female head)	0.09 (0.28)	0.04 (0.19)
Father's years of school	1.18 (1.96)	1.67 (2.83)
Mother's years of school	0.13 (0.62)	0.04 (0.19)
1 (non-Muslim)	0.08 (0.27)	0.03 (0.17)
Unemployment rate (*c*)	0.03 (0.05)	0.01 (0.02)
Sindh province	0.26 (0.44)	0.28 (0.45)

continued overleaf

TABLE A1

(continued)

Variable	*Boys, mean (SD)*	*Girls, mean (SD)*
Baluchistan province	0.03 (0.17)	0.00 (0.00)
NWFP province	0.15 (0.36)	0.03 (0.17)
1 (own land)	0.19 (0.39)	0.28 (0.45)
1 (rent)	0.06 (0.24)	0.11 (0.32)
1 (sharecrop)	0.13 (0.33)	0.25 (0.44)
Acres owned	1.08 (4.76)	1.56 (3.44)
Primary school (*c*)	0.88 (0.33)	0.94 (0.23)
Secondary school (*c*)	0.33 (0.47)	0.26 (0.44)

Notes: $n = 130$ for boys, 106 for girls. 1(x) denotes a dummy variable for x, c denotes a community-level variable.

TABLE A2

Reduced form participation equations: probit estimates

	Boys	*Girls*
Child age	0.059	0.003
	(2.17)*	(0.86)
Child age squared	−0.002	−0.000
	(1.57)	(0.68)
Household size	−0.002	−0.001
	(1.71)†	(2.70)**
Proportion males <10 years	0.041	0.029
	(0.79)	(4.10)**
Proportion females <10 years	−0.026	0.020
	(0.47)	(2.61)**
Proportion males 10–17	−0.007	0.018
	(0.10)	(1.89)†
Proportion females 10–17	−0.059	0.025
	(1.05)	(2.15)*
Proportion females 18–59	−0.073	0.026
	(1.02)	(2.37)*
Proportion males 60+	0.075	−0.012
	(0.87)	(0.98)
Proportion females 60+	−0.293	0.032
	(2.28)*	(2.43)*
Age of household head	−0.001	0.000
	(1.25)	(0.26)
Father's age	0.001	0.000
	(1.18)	(0.46)
Mother's age	−0.001	0.000
	(0.93)	(0.12)
1 (female head)	0.052	−0.001
	(2.08)*	(0.28)
Father's school years	−0.007	−0.000
	(3.90)**	(2.23)*

TABLE A2

(continued)

	Boys	*Girls*
Mother's school years	−0.002	−0.002
	(0.38)	(2.35)*
1 (household owns land)	−0.029	−0.003
	(2.14)*	(1.58)
1 (household rents land)	−0.003	0.006
	(0.24)	(1.81)†
1 (household sharecrops land)	−0.021	0.001
	(1.51)	(0.49)
Acres of land owned	−0.002	0.001
	(0.97)	(1.88)†
Acres squared	0	−0.000
	(0.90)	(2.05)*
1 (non-Muslim head)	0.059	−0.003
	(2.02)*	(1.34)
Community unemployment rate	0.087	−0.078
	(0.60)	(3.06)**
1 (Sindh province)	−0.016	−0.002
	(1.07)	(0.89)
1 (Baluchistan province)	−0.031	
	(2.28)*	
1 (NWFP province)	−0.013	−0.006
	(0.94)	(3.45)**
1 (primary school, community)	−0.045	0.032
	(1.91)†	(1.76)†
1 (secondary school, community)	−0.016	−0.004
	(1.89)†	(2.21)*
Number of observations	1,744	1,460

Notes: Figures are marginal effects, absolute *t*-statistics in parentheses. **Significance at 1% level, *at 5% level, †at 10% level. No girls in Baluchistan province work.