Is Child Work Necessary?

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Abstract

This paper investigates the hypothesis that child labour is compelled by poverty or that the child's income contribution is needed by the household in order to meet subsistence expenditures. We show 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, we identify a negative wage elasticity for boys and an elasticity that is insignificantly different from zero for girls. Thus while the evidence is consistent with boys working on account of poverty compulsions, the evidence is ambiguous in the case of girls. The results are argued to be of interest to recent theoretical and policy developments in this area.

Keywords: child labour, education, poverty, gender, labour supply.

JEL Classification: J22, J13, D12, O12

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1. Introduction

Why do children work? A common but not undisputed perception is that child work is compelled by the constraints of household poverty (see Basu and Van 1998, US Department of Labor 2000, for example). Indeed, both the geographic distribution of child workers today and the economic history of specific regions demonstrate a negative association of child work and aggregate income (see Basu 1999, Dehejia and Gatti 2002, Krueger 1996). However, it is unclear that it was the rise in household incomes that eliminated child labour by dispelling the need for it. This is because growth in aggregate income may be unequally distributed, with little increase in the incomes of those households that supply child labour. In such cases, it may not be household poverty but, rather, the development of new technologies, expansion of legal and political infrastructure, or the evolving social norms associated with growth that are instrumental in the reduction in child labour. In the last decade and a half, micro-data for developing countries have become widely available, and these make it possible to disentangle household living standards (a microeconomic variable, which differs across households) from factors like new technology, new laws or changed norms, which apply across households.

Recent analyses of these micro-data have produced estimates of the effect of household income on child labour (for a review, see Bhalotra 2003: section 5.2). However, a negative income effect does little more than affirm the plausible belief that child leisure is a normal good. Where the effective choice is between child labour and schooling rather than labour and leisure, a negative income effect signifies credit constraints (see, for example, Becker and Tomes 1986). In either case, a negative income effect may be expected even in

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households that are not subsistence-poor. To say that the children of the poor are more likely to work than the children of the rich is not the same as saying that poverty compels child labour. The latter is a sharper question, investigating which has clearly relevant policy implications (described below). This question has not been addressed before[1].

This paper argues that the question of poverty compulsions can be addressed by studying the sign of the wage elasticity. Suppose that children work because their households are very poor in the specific sense that income exclusive of child earnings falls below subsistence requirements, or that child work is *necessary*. Then children will appear to work towards a target income, set as the shortfall between subsistence needs and other income. Section 3 formalises this intuition by incorporating subsistence constraints into a model of labour supply. We show that, for the Stone-Geary and Cobb-Douglas utility functions, the model predicts that the wage elasticity of child hours will be negative *if and only if* household subsistence needs are not met by non-child income. 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 positive. The more general CES utility function yields similar predictions (see Section 3.2).

To the extent that household poverty forces children into work, child work signifies the absence of equality of opportunity in human capital acquisition, and is a mechanism for the intergenerational transmission of poverty. Important theoretical models of child labour rest upon the assumption that poverty compulsions drive child labour, and it underlies several of the policy programmes currently in place to address this problem. For example, it is a critical axiom in the seminal theoretical paper of Basu and Van (1998), with implications for the analysis of the effects on child labour of trade sanctions, bans and fertility control

[1] Ray (2000) purports to test the hypothesis of poverty compulsions but the test again relies upon

policies. It is similarly critical in Basu (2000), where a framework for analysing the impact of adult minimum wages on child labour is presented. Plausible as it may seem, raw micro-data do not always offer overwhelming support for this hypothesis (see Boozer and Suri 2001, Bhalotra and Heady 2003, Bhalotra and Tzannatos 2002, Brown, Deardorff and Stern 2003:section 3.1).

To motivate the analysis further, consider some of the policy implications that may derive from it. First, if poverty compels child work then trade sanctions or bans on child labour will tend to impoverish the already very poor households supplying child labour[2]. Second, 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 investigating the hypothesis that poverty compels child labour is essential to determining whether public money committed to reducing child labour should be directed at reducing poverty or at raising the returns to education. Current interventions reflect some diversity of strategy. For instance, the first strategy has been adopted by the Food-for-Education Program in Bangladesh (see Ravallion and Wodon, 2000), Progresa in Mexico (see Skoufias and Parker, 2001), and Bolsa Escola and PETI in Brazil (World Bank, 2001), all of which offer subsidies to households that send children to school to compensate them for the opportunity cost. The second strategy underlies the Back-to-School Program in Indonesia that has offered block grants to poor schools and scholarships to poor children that lower the direct cost of schooling (see Sayed, 2000).

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picking up an income effect on child labour.

[2] The recent surge in public interest in child labour has provoked debates on trade sanctions and the setting of international labour standards (e.g. Golub (1997), Fields (1995), Basu (1999), Bhalotra (1999)).

The data we use are from a representative 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 lifecycle-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. Thus, while the evidence is ambiguous for girls, the analysis offers considerable support for the view that boys work in order to help their households meet subsistence needs.

The paper is organised as follows. Section 2 describes the data. The theoretical model is developed in Section 3. Section 4 describes the translation of the theory into an empirical model. The results are presented in Section 5. Section 6 concludes.

2. Data and Non-Parametric Statistics

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. To avoid pooling rural and urban data, the sample is restricted to rural areas where child labour and poverty are more prevalent. There are 18382 individuals in the 2400 rural households, spread over 151 clusters (communities). Employment questions are put to all individuals ten 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. It is also the cut-off used in a number of other studies (e.g. Boozer and Suri 2001, Rosati and Rossi 2001). We therefore model the labour supply of *10-17 year olds,* distinguishing *boys and girls.* The 3373 children in this age group come from 1543 households, in 151 clusters. Work is defined, consistent with ILO conventions, as work that results in a marketable output. This is reported in the survey under two sub-categories: 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 one 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.

A Profile of Child Labour in Pakistan

The data show a high prevalence of child labour, a remarkable gender gap and a substantial fraction of children engaged in (market) wage work. See Table 1. Overall, about one in three children in the sample work. The sample probabilities of being engaged in *wage* work are 8% for boys and 7% for girls[3]. Participation rates of boys and girls are also fairly similar for work on household farms and enterprises. However, there is an alarming differential in school attendance, with only 26% of girls in school as compared with 67% of boys. This may be partly explained by a higher rate of "non-activity" amongst girls: 44.5% as compared with 11.8% for boys[4]. Boys in wage employment work an average of 31 hours a week, the average for girls being smaller at 9.5 hours a week. There is considerable variation around this mean (see Figure A1), which is exploited in estimating the wage elasticity. Given the high commitment of time that wage work appears to require, it is unsurprising that very few children combine it with school attendance. Work on household farms is relatively easy to combine with schooling (Table 1).

[3] These participation rates are high for a rural economy where self-employment still dominates wage employment. Comparing these figures for children with the corresponding figures for adults puts them in perspective. Amongst adults 18 years and older, 36% of men and 15% of women are wage workers. Note also that girls in wage employment are more likely to be doing agricultural work on somebody else's farm while boys in wage employment are most likely to be engaged in non-agricultural activity. This may contribute to the fewer hours in the average week that girls are seen to perform.

[4] Inactivity may in fact be unreported domestic work. Consistent with this, rates of "inactivity" recorded for men and women over the age of 17 are 18% and 54% respectively.

The analysis here is restricted to wage work for the following reasons. It occupies children for longer hours and virtually rules out school attendance, making it a form of child labour that is of particular concern for policy. It is also convenient for our current purposes since it is only for wage work that wage data are available. Isolating wage work allows us to concentrate on the hypothesis of poverty compulsions without the confounding influence of substitution effects arising from ownership of land[5]. The determinants of wage work and work on household farms and enterprises are so different that restricting them to have common slope coefficients can result in biased estimates. Child labour on household farms in Pakistan (and Ghana) is investigated in Bhalotra and Heady (2003).

A simple locally weighted regression (a lowess smooth) of hours on the wage rate for children who participate in wage work is in Figure A2. The graph reveals a monotonically negative relation for boys. For girls, the curve is flatter and non-monotonic. The plot is, of course, only indicative of the sign of the wage elasticity because non-labour income and other variables have not been held constant in producing it.

3. A Theoretical Framework

We are interested in deriving a test of the hypothesis that poverty compulsions drive children into work. In order to focus on this question, we abstract from consideration of other reasons why children may work (for a discussion, see Basu and Van (1998) who make a similar abstraction)[6]. Assume, in line with the literature on human capital and that on child labour,

[5] Child labour participation on household-run farms is, in many countries, increasing in the size of land farmed. To the extent that households with big farms are richer, this produces the remarkable result that children from richer households work more. This apparent paradox can be resolved by incorporating into the model imperfections in land and labour markets. See Bhalotra and Heady (2003) who model this phenomenon and provide estimates for Pakistan and Ghana.

[6] The consequences of imperfect credit markets for human capital accumulation and child labour are discussed in Ranjan (1999), Jafarey and Lahiri (2002) and Gatti, Beegle and Dehejia (2003), while Baland and Robinson (2000) focus upon problems of intergenerational contracting. The role of imperfect labour and land markets in encouraging child farm labour is analysed in Bhalotra and Heady (2003). Intergenerational persistence in child labour is investigated in Emerson and D'Souza (2000) and Ilahi et al (2000). The widely used assumption of parent altruism is investigated in

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. We start, in Section 3.1, with the Stone-Geary utility function, which is used (to illustrate the possibility of multiple equilibria in the labour market) in Basu and Van (1998), Basu (1999) and Basu (2000). We then investigate if a similar testable prediction flows from the less restrictive and more commonly used CES function (Section 3.2).

3.1. The Stone-Geary case

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

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

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. We assume $C \geq 0$, $0 \leq L_i \leq 1$, 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[8]. A convenient normalisation 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 $H_j=1$.

Bhalotra (2002a). Bhalotra (2002b) investigates separability of parent and child labour supply. The possibility that child labour depends on the bargaining power that children can exert in household decision making is considered in Moehling (2003) and, implicitly, in Bhalotra and Attfield (1998). The impact of preference heterogeneity between fathers and mothers is considered in Galasso (2000) and Basu (2001).

[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), Nardinelli (1990), Khan (2001)). Parent altruism is investigated for the sample of Pakistani households analysed in this paper in Bhalotra (2002a), and the data are found consistent with altruism.

[8] This is a common assumption in studies of child labour including the paper by Basu and Van (1998), which is closely related to this paper. It is a natural way to model the hypothesis of interest. The empirical model is not quite so strict. The earnings of parents are included on the right hand side and their potential endogeneity is addressed.

The budget constraint is

$$(2) \qquad C = N + w_j + w_i(1 - L_i) \equiv Y + w_i(1 - L_i)$$

where the price of consumption is normalised to unity, w is the real wage, subscripts i and j denote child and parent respectively, N is household non-labour income and we have defined the *child's non-labour income* 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 (2) to eliminate C in this function, we can maximise $(Y+w_iH_i-S)(1-H_i)$ to get optimal child hours of work :

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

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 (3) for the case $0 < \theta < 1$ is

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

Using (3), it is clear that if desired hours, $\theta>0,$ then $w_i+S-Y>0$ and the sign of (4) is the same as the sign of the numerator. Thus, this model yields the clean 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, observing a negative wage elasticity indicates subsistence poverty (i.e. compelling poverty).

<u>3.2. The CES case</u>

Consider the popular constant elasticity of substitution (CES) utility function

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

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: $(C-S)^{\alpha}L_i^{(1-\alpha)}$. Comparing this with equation (1), it is clear that the Stone-Geary function derives from (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].

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

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

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

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

It is straightforward to confirm that setting $\sigma=1$ and $\alpha=(1-\alpha)$ gives θ in (3). The wage elasticity implied by (7) is given by:

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

where the denominator is non-negative since survival is only feasible if the maximum possible consumption exceeds subsistence needs or $Y+w_i \geq S$. Hence, the sign of ε_{Hw} is the sign of $\sigma(Y-S)+(\sigma-1)w_iH_i$ (got by taking $(1-H_i)/H_i$ out as a common factor in the numerator of (8)). Although this expression involves the endogenous variable, H_i, we can proceed to analyse the sign of the wage elasticity without knowledge of where H_i is set, given the information that $H_i \geq 0$, $L_i \equiv (1-H_i) \geq 0$ and $w_i \geq 0$.

The question of interest is: If subsistence poverty binds (S>Y) and children work ($H_i>0$) then what is the sign of the child wage elasticity (ε_{Hw})? This depends upon the value

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[9] The relative weight on child leisure (or child consumption) in the utility function of the parent indicates the degree of parent altruism: see Bhalotra (2002a), for example.

[10] When H_i lies outside [0,1] it is restricted to 0 or 1. Consider $H_i>1$. In (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 $H_i=0$.

of σ. Figure 1, which is adapted from Barzel and McDonald (1973)[11], shows for the CES case how the shape of the child labour supply curve depends upon the values of σ and Y-S. As discussed, with $\sigma=1$, we get (4), in which case the elasticity is negative if and only if $S>Y$. For $\sigma<1$ and $S>Y$, we again get, from (8), that $\varepsilon_{Hw}<0$ everywhere. For $\sigma>1$ and S>Y, we have a forward-falling labour supply curve or a curve that is negatively sloped at low wages and positively sloped at high wages (note that w_i weights the second term in the numerator of (8); see graph VII in Figure 1). Setting aside the high-wage case as implausible by context, we can conclude that, for all values of σ, *if* S>Y *then* $\varepsilon_{Hw}<0$. This establishes the hypothesis.

The Stone-Geary case produced the even stronger result that $\varepsilon_{Hw}<0$ *if and only if* S>Y[12]. We now briefly consider whether the CES function offers this. In the case of $\sigma\geq1$, it does: equation (8) shows that $\varepsilon_{Hw}<0$ implies S>Y. However, if $\sigma<1$ and $\varepsilon_{Hw}<0$, the sign of Y-S is ambiguous- we can only say that $(Y-S)<[(1-\sigma)/\sigma]w_iH_i$. This case corresponds to graph I in Figure 1: the case of the backward bending labour supply curve. It seems unlikely that this is applicable to the current context of rural families with working children (also see section 5.1).

To summarise, if we observe a negative wage elasticity then 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

[11] While their analysis is different from ours in many respects and it is not about child labour, the paper by Barzel and McDonald has in common with this paper that it is concerned with characterising 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.

[12] The standard scientific procedure is, of course, to use theory to predict, say, that A$\Rightarrow$B. In this case, A corresponds to acute poverty defined as S>Y and B to the observable prediction that $\varepsilon_{Hw}<0$. If we observe not-B, we conclude not-A, i.e. we reject the hypothesis of poverty compulsions. If we observe B, we conclude that the data are consistent with A. Hypothesis testing often stops at this point, though there typically remains the possibility that A is just one of a number of scenarios compatible with B. Here, it is argued that the theory is able to generate the stronger prediction that B$\Rightarrow$A for the Stone-Geary case and, under specified plausible conditions, for the CES case too.

necessary for survival. This is sufficient for our purposes. However, it is interesting to note that, under certain fairly plausible restrictions, we can further argue that compelling poverty is implied by a negative wage elasticity.

4. The Empirical Model

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

4.1. The Variables

Since the hypothesis of interest is whether children in work have got there on account of poverty compulsions, we estimate a model of hours conditional on participation. While the participation wage elasticity is expected to be positive, the hours wage elasticity can be negative as shown in Section 3[13]. Since 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 a work 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)[14]. Identification in this context often

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[13] This is obvious enough since a negative income effect can only be generated when hours of work are positive. 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. Probit estimates of participation are presented in Table A2. A tobit specification is avoided because (a) we cannot assume that the parameters –including the wage coefficient- are common to the participation and hours decisions and (b) 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.

[14] Using US data, Newey, Powell and Walker (1990) investigate the robustness of estimates of women's labour supply to the normality assumption. They find no statistical difference between the conventional 2-step Heckman estimates and semi-parametric estimates obtained using the series

relies solely upon functional form. In this paper, we exploit the fact that children who participate in wage work do not, at the same time, participate in school (Table 1). We therefore use as instruments in the work-participation equation, (three) dummy variables for the presence of a primary, middle and secondary school. 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.

The dependent variable, H_i, is defined as the annual average of weekly hours in wage work. In common with most studies of (adult) labour supply, the (child) wage rate, W_i, is measured as earnings divided by hours of work[15]. The *child's non-labour income*, Y, is defined in equation (2) as $Y=N+\Sigma_j w_j H_j$, where N is household non-labour income and $\Sigma_j w_j H_j$ denotes adult earnings[16]. The measure of N that is used here has attractive features in terms of both interpretation and measurement. It is a lifecycle-consistent measure of non-labour income that is measurable using cross-sectional data (see Blundell and Walker 1986). To see this, note that the intertemporal budget constraint that defines the time path of assets is: $A_{t+1} =(1 + r)A_t + \Sigma_k w_{tk} H_{tk} - C_t$, where k=i (child), j (adults), A is assets, r is the interest rate, and t denotes time. This can be written as $rA_t - \Delta A_{t+1}= C_t - \Sigma_k w_{tk} H_{tk}, =N$ (using (2)). In a one-period model, N= rA_t. The asset change term, ΔA_{t+1}, allows N to reflect savings and dissavings

estimator. They also find that semiparametric estimates of the first stage participation equation are not significantly different from ML estimates. They conclude, in line with Mroz (1987), that the sensitivity of estimates of the hours of work model for women in the US depends more upon correct specification of the regression function and the choice of instrumental variables than on specification of the error distribution.

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 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.

16 In the theoretical model, it was assumed that parents always work or that H_j=1. This assumption is relaxed in the empirical model. Possible simultaneity of parent hours (H_j) and child hours (H_i) is allowed for in treating Y as endogenous (see below).

between periods. The static measure is valid if agents are myopic or if there exist no capital markets so that it is impossible to save and dissave but, in general, it is restrictive[17]. Of particular importance in this paper, a one-period measure of income tends to generates wage elasticities that confound the effects of shifts in wage profiles with movements along them (see Blundell and MaCurdy, 1999)[18]. In the absence of data on asset changes, N is measured as the difference between consumption expenditure (C) and total labour earnings[19]. It follows that the child's non-labour income is measured as the difference between C and child earnings:

(9) $Y = N + \Sigma_j w_j H_j = C - w_i H_i$, where j denotes adults and i denotes child.

Basu and Van denote household poverty by the adult wage rate in order to focus attention on an interesting configuration of labour market equilibria. We have generalised this by characterising household poverty as dependent on adult earnings, non-labour income and any saving or dissaving performed by the household[20]. This generalisation is of great practical importance because income from self-employment tends to dominate labour income in many

[17] 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 (see Besley, 1996). In the Pakistan data used in this paper almost 50% of households reported borrowing or lending money in 1991.

[18] Conditioning labour supply on N, the current period allocation out of lifecycle wealth, is an alternative to the Frisch approach of Heckman and MaCurdy (1980) in which the conditioning variable that captures future anticipations and past decisions is λ, the marginal utility of wealth. It is a particularly attractive alternative when the data, as here, are limited to a cross section. This is because N is observable in a cross section if data on consumption and earnings are available (see (2)), while λ is not (refer Blundell and Walker, 1986).

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

[20] This definition allows, consistent with the actual data, that there are earners other than parents in the household. Although we are implicitly thinking of a representative child or, like most theoretical models, of a 1-child family, writing Y as $C-w_iH_i$ effectively includes any sibling earnings in Y too. We also investigate a specification in which the hours and wages of siblings are averaged (section 5.2).

developing country households and there is typically huge inequality in asset (especially land) distribution.

It is conventional to express income in per capita terms, but we 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[21]. There are some price data in the community-level survey but these contain lots of missing values. Rather than use these, we include province dummies to capture any cross-sectional variation in prices.

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[22]. 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 also 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 levels of both parents* are included to allow for educated mothers having greater power in household decision making (e.g., Thomas 1991). The education of parents and especially mothers may also increase the efficiency of allocating resources towards child schooling (e.g., Behrman, Foster, Rosenzweig and Vasishtha 1999, Lam and Anderson 2001) and, as argued in the following section, parental education may reflect underlying preference parameters: dynastic tastes for

[21] The average household size is 8 of whom, on average, 4.4 are over the age of fourteen, 2.6 are under the age of ten, 1.85 is in aged 10-17 and 1 is in the age range 10-14.
[22] 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, something that many existing empirical studies of child labour do not do.

education or low discount rates. The cluster-level *unemployment rate* (calculated by aggregation of individual responses) is included to allow for disequilibrium in the labour market. The common practice of excluding it from labour supply models results in mis-specification (see Ham (1986), Card (1988)). In addition, the vector X includes a quadratic in child *age* and dummies for *religion* of the household head. A further additional regressor is 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 (see Case, Lin and McLanahan 2000 and Case and Paxson 2001) for evidence that the biological relation of adult and child influences resource allocation within the household). As it was insignificant in the equation for boys and girls, it was dropped. Other variables investigated but not retained because insignificant include an indicator for child illness and quadratic terms in lnW and lnY.

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

$$(10)\ H_i = \alpha + \beta lnW_i + \gamma lnY + \delta X_i + e_i$$

where, as discussed earlier, H is hours in wage work, lnW is the log of the wage earned by the child, lnY 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 lnW and lnY are investigated[23]. We investigated the log-log form but as the results were similar to those obtained with the semi-log form, these are not discussed further. In view of large

[23] Amongst labour supply studies that derive from lifecycle models, some normalise Y as (Y/W). This is consistent with theoretical restrictions and convenient because Y can take negative values for which the logarithm is undefined (e.g., Blundell, Duncan and Meghir, 1994). A difficulty with adopting this specification in this paper is that it restricts the wage elasticity to be more negative at *higher* levels of income, contrary to what we would expect in rural Pakistan. In our sample of children, no boys have a

gender differences in participation and hours of child labour (Section 2), equation (10) is estimated separately for girls and boys aged 10-17.

4.2. Estimation

The coefficient on the child wage is subject to endogeneity bias if, for example, unobservables like tastes for work (laziness) are (negatively) correlated with both the wage and work hours of the child[24]. The going agricultural wage rate for children at the cluster (community) level –and its square-are used to instrument the individual child wage. Community-level wages are obtained from village leaders as part of the village questionairre. To allow for within-community variation, an additional instrument is defined by interacting the community wage with the child's age[25]. 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), we investigated a simple community-level regression relating the child wage to the participation rate. The estimates indicate an insignificant relation.

Income is typically correlated with parents' education which, in turn, may be higher in families (or dynasties) which have a taste for education or a lower discount rate. As this will tend to create an over-estimate of the effect of income on child labour, we control for parental education. If income includes child earnings (as in many previous studies of child labour) then there is scope for positive feedback from child hours to household income.

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negative value of Y and only two girls, both from the same household, do. These two girls are dropped from the data and we are then able to use the logarithm of Y.

[24] On the other hand, 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, as our estimates show that the wage coefficient becomes *more negative* when instrumented, division bias would appear to be small in these data.

[25] Finding a valid instrument for the wage in a labour supply equation is known to be a difficult problem. Estimates of adult labour supply equations in the literature are often obtained by instrumenting the wage with education (for example, Fortin and Lacroix (1997), Kooreman and Kapteyn (1986), Hernandez-Licona (1996)) on the arguable assumption that education does not affect preferences for work (see Pencavel, 1986). This assumption may be especially strong when the data refer to children, for whom education is a concurrent decision.

Although variations in child earnings are not directly reflected in our measure of Y (see (9)), Y is endogenous if parental and child labour supply are jointly determined or on account of it being defined by lifecycle consumption and leisure choices. Instrumenting Y for endogeneity also addresses the problem of measurement error that is known to afflict measures of income in rural economies[26]. A suitable instrument for income in a child labour model is, in general, difficult to find. Previous empirical studies of child labour have tended to use a static definition of income and to ignore its possible endogeneity (see Bhalotra and Tzannatos 2002 for a review). In this paper, we instrument Y with a quadratic in the community-level going wage rate for men. In order to preserve within-community variation in income, this is interacted with the school-years of the father and mother and with the acreage of farmland owned by the household. We also investigate overidentifying restrictions associated with a range of indicator variables describing the economic and infrastructural development of the community in which the household is located and retain an indicator for whether the community has a shop.

Estimation is by the two-step efficient generalised method of moments estimator (GMM) though, for comparison, we also report traditional instrumental variables (2SLS) and ordinary least squares (OLS) estimates. GMM is more efficient than 2SLS and robust to heteroskedasticity of unknown form, as well as to arbitrary intra-cluster correlation (see Wooldridge 2002: p.193). To the extent that the instruments are weak, it is important to offer the OLS estimator for comparison (see Bound, Jaeger and Baker, 1995). We use Cragg's

[26] Measurement error bias may be exacerbated by virtue of the differencing implicit in creating Y. Let C^* denote true consumption and E^* ($=W_iH_i$) denote true child labour earnings and let these variables be measured with random errors denoted u and e with variances σ_u and σ_e. The measured variables are then $C=C^*+u$ and $E=E^*+e$. The lifecycle measure of Y is defined as (C-E) which equals $(C^*-E^*)+(u-e)$, where $var(u-e)=\sigma_u+\sigma_e-2cov(u,e)$, which we suspect is larger than σ_u, the variance of the measurement error when household living standards are measured simply by consumption, C, rather than by Y. At the same time, the signal in (C-E) is diminished to the extent that C and E tend to be close to each other. As a result, the noise-signal ratio is expected to be larger for Y than for C.

OLS estimator, which is more efficient than standard OLS in the presence of heteroskedasticity of unknown form (Davidson and McKinnon 1993: pp.599-600). The standard errors of all reported estimates are robust to heteroskedasticity (see White 1980) and, further, are adjusted for non-independence within sampling clusters (e.g. Deaton 1997: chapter 2).

The Hansen-Sargan J statistic, a version of the Sargan statistic that is robust to heteroskedasticity, is presented as 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.235-36, Hayashi 2000: pp.227-28). The C-statistic or the difference-in-Sargan statistic allows a test of a subset of the orthogonality conditions. Shea's partial R^2 is presented as a measure of instrument relevance (power) that takes account of correlations amongst instruments (Shea 1997). The more conventional F-test of the joint significance of the instruments is also presented.

5. Results

5.1. The Main Results

Refer Tables 2 and 3, which report OLS and IV (2SLS and GMM) estimates. Notes to the Tables report the tests of the validity and power of the instruments that were described above. Validity cannot be rejected but, in some cases, the instruments are quite weak (see Tables 2, 3). For this reason, the OLS estimates are considered together with the IV estimates. There is considerable variation in hours around the mean, a good deal more than is typically observed for adult hours of work in industrialised nations (see Figure A1). The OLS estimates explain about 54% of this variation for boys and 32% for girls (this is the centred R^2). 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.

The wage elasticity is significantly negative for boys. The GMM estimate is –0.53 and the OLS estimate is –0.33, both at the sample mean of hours of work. We therefore cannot reject the hypothesis that boys work on account of poverty compulsions. The estimates suggest that if a boy's wage rate drops, he works harder 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 col.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). What we can conclude is that household poverty is a more compelling factor in determining child labour amongst boys than amongst girls. This is corroborated by the raw data: see Table 4, which confirms that households with working boys are poorer on average than households with working girls[27].

Insofar as girls' work is not compelled by poverty, their participation in wage work and their very low participation in schooling may reflect that the expected returns to schooling are low relative to the returns to work. This will change the weight on non-labour time, L_i (which includes schooling), in the utility function (e.g. Basu and Tzannatos 2003).

Improving returns to school for girls may contribute to closing the gender gap in schooling apparent in Table 1. However, to the extent that the additional enrolment derives initially from the group of girls reporting inactivity, the *immediate* impact on the incidence of child labour may be small. Indeed, analysis of a school subsidy programme in Bangladesh

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

shows that the subsidy had a much bigger impact on school attendance than it did on child labour participation, given the large pool of inactive children (Ravallion and Wodon 2000)[28].

Further analyses of the Pakistan data used in this paper reinforce the finding that the determinants of work are quite different for boys and girls. Thus, for instance, analysis of child labour on household-run farms shows that the substitution effects associated with land ownership dominate wealth effects for girls, though not for boys. As a result, girls in households that own relatively large plots of land are both more likely to work and less likely to attend school than girls in households with smaller landholdings (Bhalotra and Heady 2003). This is consistent with consideration of relative returns to work and school having greater power in explaining girls' work than that of boys.

Household income has a negative impact on boys' work, the GMM elasticity being –0.63. This is fairly large. The fact that only the OLS estimate is significant reflects the relatively low power of the instruments. The impact of income on girls' hours of work is smaller and insignificant, consistent with the finding that their labour is not compelled by poverty. Analysis of data for other regions often find insignificant (or small) income effects on child labour (see Bhalotra and Tzannatos 2002, Brown *et al* 2003). Measurement error in the income variable may contribute to this but is unlikely to be a sufficient explanation since, for example, we typically find larger effects of income on schooling. It is likely that inclusion in the model of parental education and land ownership variables reduces the income effect. The income effect is further investigated by studying parental altruism (Bhalotra 2002a) and the relevance of the common assumption of separability of parent and child labour supply (Bhalotra 2002b). In this paper, the emphasis is on the size of the wage elasticity.

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[28] As discussed in Section 2, the high inactivity rates amongst girls and women may reflect full-time engagement in domestic work that is not counted as "work". Certainly, lay observation does not

Estimates for the other variables in the model are displayed in Tables 2 and 3. The results are again quite different for girls and boys. For example, land ownership reduces boys' hours, while belonging to a household where the head is not Muslim increases girls' hours. More of the control variables are significant in the participation equation (Appendix Table A1)[29].

5.2. Alternative Specifications

How robust are these estimates? Comparable estimates of the wage elasticity of child work hours are unavailable in the literature. Estimates of wage and income elasticities for adults in industrialised countries have exhibited a wide range, even when obtained using the same data. Some of this variation has been shown to arise on account of 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 paper was discussed in section 4. 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, for instance, Attanasio and MaCurdy (1997), Kniesner (1976) for the US, Kooreman and Kapteyn (1986) for the Netherlands), though 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 Mexico (Hernandez-Licona, 1996) and rural India (Rosenzweig, 1980) using data on adults. Looking at children here sharpens the question: If adult earnings constitute the larger share of household earnings and these contribute to the non-labour

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support the view that girls and women sit idle in the household. It is, for these reasons, often more useful to think in terms of increasing school enrolment rather than in terms of reducing child labour.

[29] Conditioning on participation, the variation in hours is not very sensitive to the included variables. This is also the case in estimates for high-income countries: see Heckman (1994), for example, who reports that wage and income elasticities of hours of work are close to zero in developed countries, particularly for men. The labour supply elasticities for hours of work of children in Pakistan that are identified here are, in comparison, not surprisingly small.

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[30].

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 unemployment 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 & 3 with column 1 in Table 2).

The estimates reported in Section 5.1 were obtained from 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 all-boys-in-the-household is –0.77 and significant at the 1% level. The income elasticity is –1.29, significant at the 10% level. Averaging over all-girls in the household does not change the results: both the wage and income coefficients are insignificant.

To the extent that fertility is a choice variable correlated with investments in child human capital, household size – though almost always on the right hand side of equations describing child labour- is a potentially endogenous regressor. Its exogeneity was therefore investigated using the C-statistic (described in Section 4.2), 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). We also confirmed that the

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[30] Only the key results are reported here. A full set of results for this and other specifications described in this section is available from the author on request).

wage and income elasticities for boys and girls are not significantly altered when household size is dropped from the model.

Since some of the control variables are likely to be correlated with the key variables of interest (e.g., child age with child wage, acres of land owned with household income), an equation is estimated with just the child wage and income as regressors. There is no statistically 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.

6. Conclusions

This paper finds support for the assumption that poverty compels boys to work. In the case of girls, the evidence is ambiguous. While we cannot strictly reject the hypothesis that poverty compels girls to work, the results in this paper and in related research (summarised in section 5.1) invite consideration of other factors such as the relative return to school driving girls' work. The results have potent implications for further developments in theory and policy: no previous research has effectively tested for poverty compulsions, even though most theoretical models of child labour and a vast array of actual policy interventions rest upon this assumption (see section 1). The results also suggest that research and policy design in this area should be gender specific.

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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
Percent that combine types of work		
Household farm & enterprise work	0.95	0.12
Household farm & wage work	1.4	2.3
Household enterprise & wage work	0.17	0.06
Percent that combine work & school		
Household farm work & school	10.7	1.9
Household enterprise work & school	0.73	0
Wage work & school	0.39	0.13
Number of children	1786	1587

Notes: Rural Pakistan, 10-17 year-olds. All figures are percentages except figures in brackets, which refer to the annual average of weekly hours of work conditional on participation.

Table 2
Hours Equation for Boys

	1	2	3
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]*
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	-1.94	-1.03	-1.18
	[1.34]	[0.54]	[0.66]
mother's school	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]*
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]
1(sindh province)	16.61	19.71	17.36
	[2.51]*	[2.16]*	[2.15]*
1(baluchistan province)	-9.45	-17.14	-20.05
	[0.38]	[0.51]	[0.62]
1(NWFP province)	-4.66	2.44	-0.32
	[0.87]	[0.25]	[0.04]
R^2 (uncentered; centred)	0.83; 0.54	0.74; 0.29	0.79; 0.41

Notes: The dependent variable is hours of work conditional on participation. Robust z-statistics in brackets, * significant at 5%; ** significant at 1%. The number of observations, N, is 130. Elasticities reported in the text are calculated at sample means. Mean ln(household income) is 5.53 and mean hours of work are 31.4. The Hansen-Sargan J statistic is χ^2 (7)=5.8 with a p-value of 0.56. Instrument validity, therefore, cannot be rejected. The power of the instruments is weak in the participation and income equations, while being good in the wage equation: The first stage R^2 is 0.19 in the participation probit (pseudo R^2), 0.43 in the wage equation and 0.41 in the income equation. In the participation equation, the test of joint significance of the school dummies is χ^2 (3)=4.29, p>χ^2 =0.23. In the wage equation, Shea's partial R^2 for the instrument set is 0.09 and F(9,93)=1.98, p>F=0.05. For the income equation, Shea's partial R^2 is 0.04 and F(9,93)=0.41, p>F=0.92.

Table 3
Hours Equation for Girls

	1	2	3
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]
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	-0.11	-0.08	0.20
	[0.36]	[0.22]	[0.68]
mother's school	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]
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]
1(sindh province)	-1.66	2.53	-0.86
	[0.68]	[0.68]	[0.29]
1(NWFP province)	0.79	-1.18	6.81
	[0.11]	[0.14]	[1.01]
R^2 (uncentered; centred)	0.66; 0.32	0.63; 0.26	0.61; 0.21

Notes: See Notes to Table 2. The number of observations, N, is 106. Elasticities are calculated at sample means. Mean hours of work are 9.5 per week. The Hansen-Sargan J statistic is χ^2 (7)=6.97 with a p-value of 0.43. Instrument validity, therefore, cannot be rejected. The power of the instruments is reasonable in the participation and income equations but weak in the wage equation: The first stage R^2 is 0.18 in the participation probit (pseudo R^2), 0.32 in the wage equation and 0.57 in the income equation. In the participation equation, the test of joint significance of the school dummies is χ^2 (3)=5.97, p>χ^2 =0.11. In the wage equation, Shea's partial R^2 for the instrument set is 0.11 and (F(9,69)=0.91), p>F=0.52). For the income equation, Shea's partial R^2 is 0.12 and F(9,69)=1.88, p>F=0.07.

Table 4

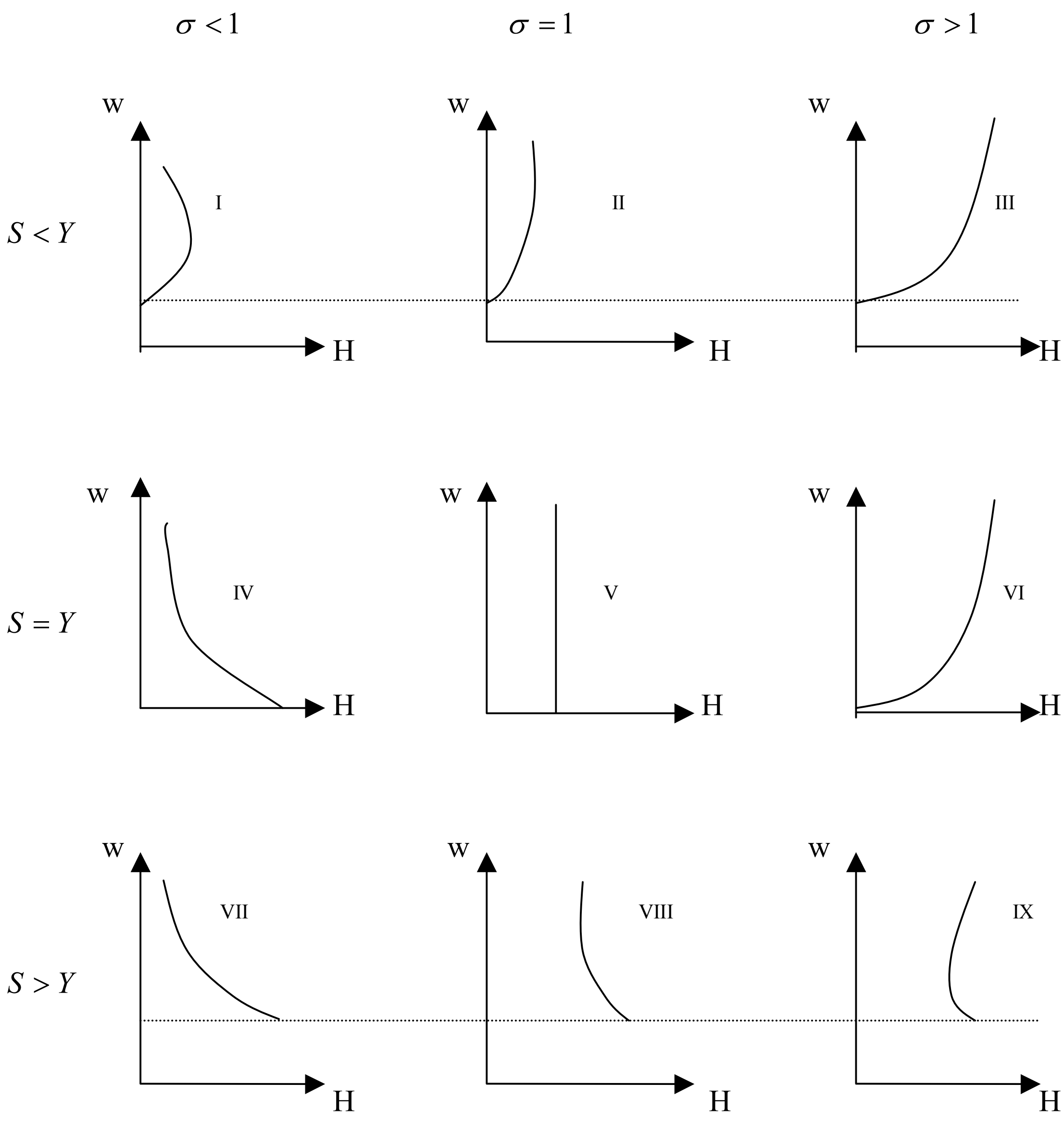
Gender Comparison of Mean Household Living Standards

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
percent of households that own land	19.1%	27.8%	0.056
years of schooling of father	1.18	1.74	0.036
percent of households with a female head	8.8%	3.7%	0.053

Notes: Means 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.

Figure 1
Labor Supply Curves



Notes: H is hours of work, w is the wage, Y is the child's non-labor 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).

APPENDIX

Table A1
Means and Standard Deviations of Variables
Sample of children participating in wage work

Variable	Boys mean	Boys s.d	Girls mean	Girls s.d.
hours in wage work	32.46	24.19	9.34	9.25
ln wage	1.31	0.97	0.87	0.91
ln nonlabour 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
prop males<10	0.15	0.12	0.15	0.12
prop females<10	0.12	0.12	0.13	0.12
prop males 10-17	0.22	0.11	0.09	0.09
prop females 10-17	0.08	0.10	0.23	0.10
prop females 18-59	0.19	0.10	0.18	0.09
prop males 18-59	0.19	0.11	0.17	0.09
prop males >59	0.04	0.07	0.02	0.06
prop 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
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
middle (c)	0.49	0.50	0.45	0.50
secondary school (c)	0.33	0.47	0.26	0.44
ln male wage (c)	3.68	0.28	3.74	0.28
ln child wage (c)	3.05	0.38	2.97	0.46
child wage	44.39	7.53	39.86	9.10
male wage(c)*father's educ	4.41	7.43	6.14	10.42
male wage(c)*mother's educ	0.53	2.55	0.14	0.72
male wage(c)*acres owned	4.08	19.09	5.81	13.04
1(shop present, c)	0.93	0.26	0.91	0.29

Notes: N=136 for boys, 108 for girls. 1(x) denotes a dummy variable for x, c denotes a community-level variable, prop is proportion, ln is logarithm.

Table A2
Reduced Form Participation Equations: Probit Estiamtes

	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 yrs	0.041	0.029
	[0.79]	[4.10]**
proportion females<10 yrs	-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]*
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]
cluster 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)	-0.035	0.002
	[1.49]	[1.17]
1(middle school, community)	-0.01	0.001
	[0.71]	[0.33]
1(sec school, community)	-0.006	-0.004
	[0.44]	[2.21]*
Number of observations	1744	1460

Notes: Figures are marginal effects, absolute t-statistics in parentheses. ** denotes significance at the 1% level and * at the 5% level. No girls in Baluchistan province work.

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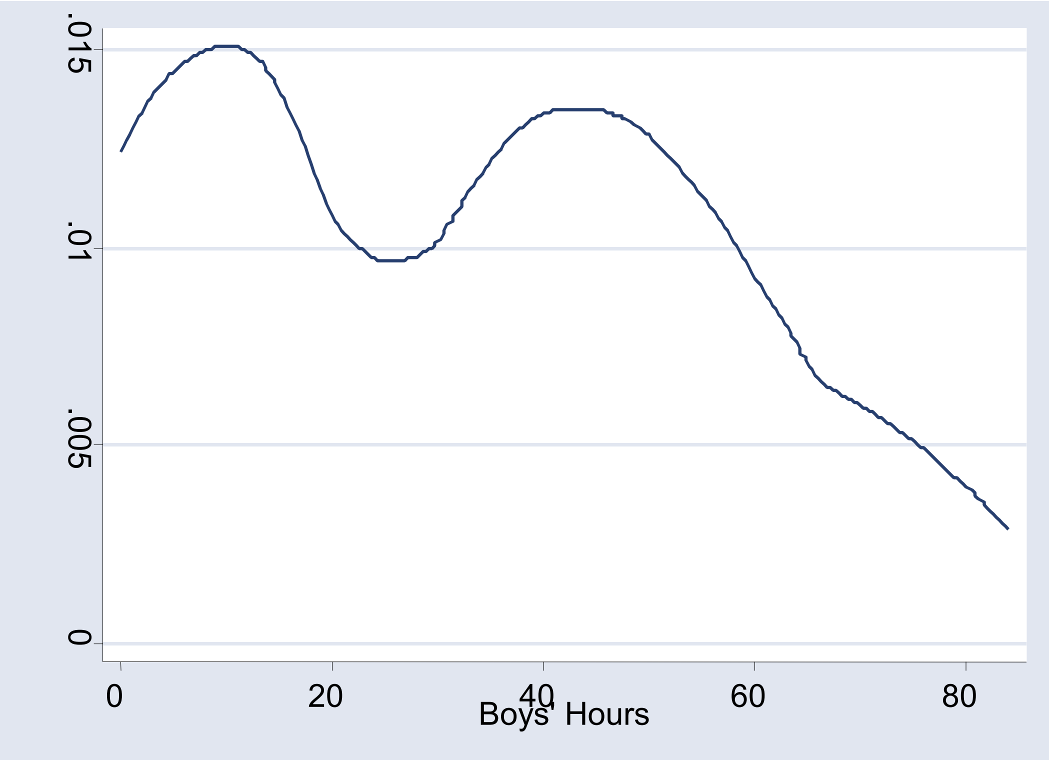
<u>Figure A1</u>
<u>Hours of Wage Work Conditional on Participation: Kernel Density Estimates</u>



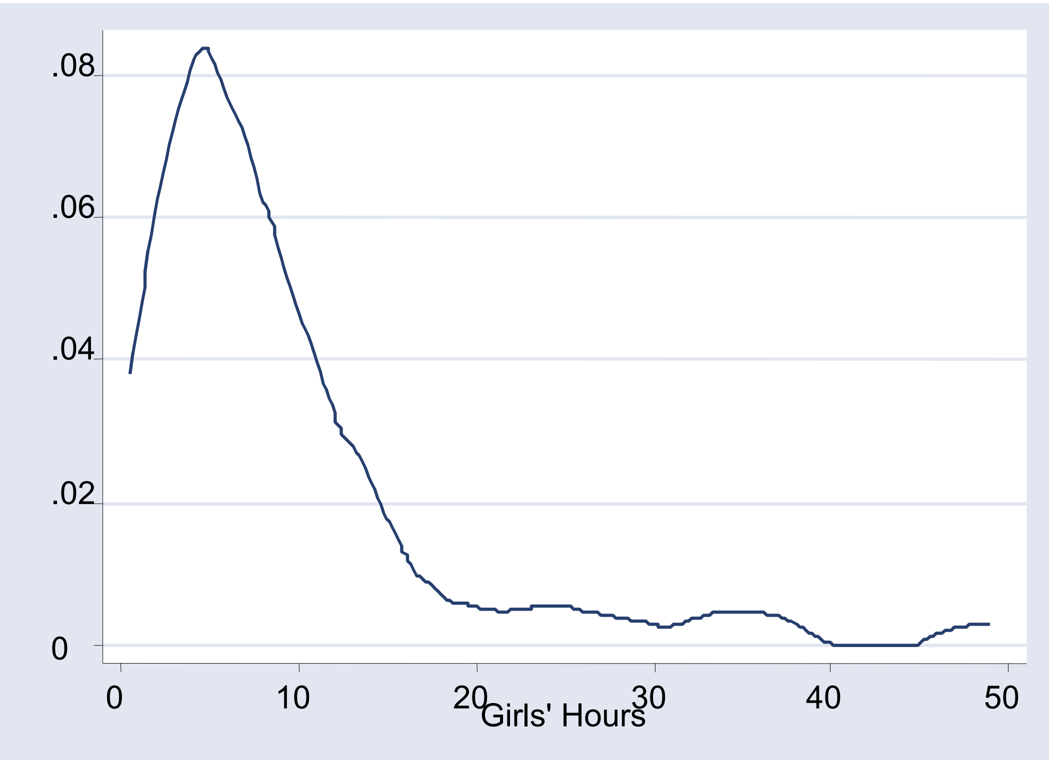
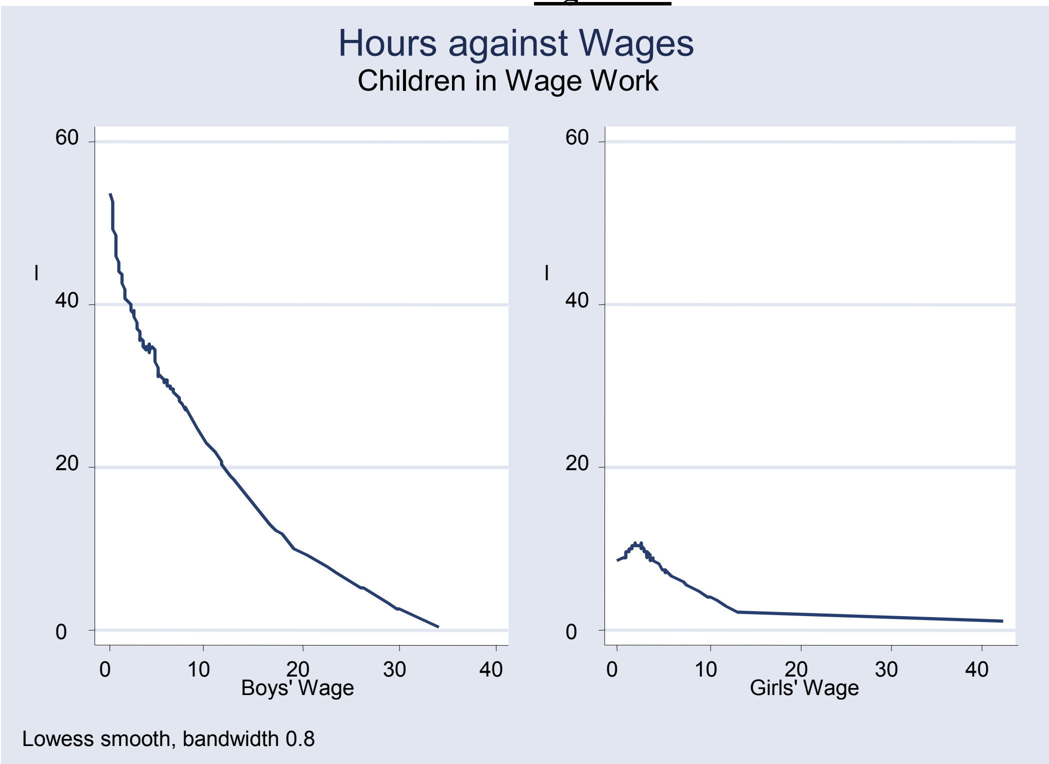


Figure A2



Hours against Wages
Children in Wage Work
60
40
20
0
0
10
20
30
40
Boys' Wage
60
40
20
0
0
10
20
30
40
Girls' Wage
Lowess smooth, bandwidth 0.8