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Enrollment: Evidence from Rural India

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Intrahousehold Resource Allocation, Child Labor, and School Enrollment: Evidence from Rural India

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Abstract

This paper empirically analyzes the determinants of child labor and school enrollment in rural Andhra Pradesh, India, where a high prevalence of child labor poses serious problems. For this analysis, we collected detailed time-use data, information to judge whether the household is under a binding credit constraint, and information on extra-household linkages covering pre-marital and parental generation. The last information provides us exogenous sources of variation in bargaining powers within a household. We derive empirical models from a theoretical framework of intrahousehold resource allocation and apply them to this dataset. The results based on a reduced-form approach show that the incidence of child labor or school enrollment is affected by the child's age and sex, the household size, land assets, social classes, and grandparental characteristics. We then estimate more structural models in order to investigate the effect of maternal labor and credit constraints on intrahousehold inequality of labor allocation, paying due consideration to the fact that both maternal labor and credit constraints are endogenous. The results based on a conditional labor supply approach show that when a mother works outside, her domestic labor is more likely to be replaced by daughters than by sons, but the sex contrast in market work by children is not affected much, resulting in a higher work burden on daughters. The estimation results based on an endogenous switching model show that access to credit is a major determinant of the child's time use: children of credit-constrained households spend less hours in school and for leisure but spend more hours working in the house. The regression results from all empirical models support collective household models against unitary models

Keywords: child labor, schooling, gender bias, credit constraint, household models.

1. Introduction

High incidence of child labor and staggering school enrollment of children continue to be a serious problem facing India. According to ILO-IPEC (2005), about 127 million children in the age group 5-14 are engaged in work in the Asia-Pacific Region, of

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which more than 100 million are attributed to India alone.¹ There have been a number of attempts to eradicate child labor and to send children to school in India, such as legal actions against child labor, trade sanctions, enlightenment of parents, education subsidies, etc. Nevertheless, the impact of these policies has been limited, judging from the current estimates of the number of working children. Similarly to the limited policy success, our understanding of the determinants of child labor is also limited, though the number of theoretical and empirical work on child labor has been increased rapidly in recent years (Basu 1999; Basu and Tzannatos 2003).

On child labor in India, there exist a few microeconomic studies that empirically analyze the determinants of child labor (e.g., Aggarwal 2004; Basu et al. 2003; Deb and Rosati 2002; Edmonds et al. 2005; Sakamoto 2006). These studies have one thing in common: they employ datasets collected by a large-scale sample household survey conducted by national or international agencies. In other words, the analysts of these papers were seldom involved in the design and implementation of the survey. This limits their analysis into the one using information available in the standardized household questionnaire only. Since the incidence of child labor and school enrollment is a result of households' decision making, we believe that it is critically important to analyze the incidence using detailed information on intrahousehold resource allocation within a household.

With this belief, we conducted a special household survey in rural Andhra Pradesh to collect detailed time use data and we went back in time (pre-marital and parental generation), recording information on extra-household linkages (Fuwa et al. 2006a). The information on time use at the individual level enables us to directly analyze the intrahousehold allocation of time among its members; the information of pre-marital and parental generation provides us more exogenous sources of variation in bargaining powers within a household than those usually available (Quisumbing and Maluccio 2000). The rich information thus collected and the authors' close involvement in the survey distinguishes the analysis of this paper from those by others mentioned above.

Methodologically, we derive empirical models from a unified framework of collective/unitary household models and apply them to this dataset. We first estimate a reduced-form model with the time use of children in various activities as the dependent variable, and then we estimate more structural models in order to investigate the effect of maternal labor and credit constraint on intrahousehold inequality of labor allocation. In the latter analysis, we pay due consideration to the fact that both maternal labor and credit constraints are endogenous. The female labor force participation rate tends to decline in the early stages of economic development and then increases. This is the U hypothesis of female labor market participation in the process of economic development (Durand 1975; Goldin 1995). In the earlier stages of economic development when agriculture is the dominant mode of production, female labor participation in family farms, household businesses, and domestic workshop tends to be high. While the forms of female labor vary depending on society, wage labor participation rate in India is unusually high for the poor families. By using Indian household data, Mammen and Paxson (2000) found that while Indian female labor force participation rates never exceed 60 percent, more than

¹ Lieten (2002) also estimated the number of working children in India as more than 100 million and commented that this number is 10 times more than the official figures available from census and NSS reports.

60 percent of both rural and urban working women at the lowest levels of per capita expenditure work for wages. In this case, women's high labor participation rate should be understood as a response against poverty and resource constraints under limited market accessibility. We thus need to carefully consider the constraints households face and the consequences of maternal labor participation on intrahousehold labor allocation because it often leads to an increase in child labor, especially daughter's labor, for domestic chores.

One reason why girls might fare badly within households is that in poor credit-constrained families, siblings must compete for resources and boys have the advantage, possibly because investments in boys yield higher future returns to parents than do investments in girls (Mammen and Paxson 2000; Garg and Morduch 1998). In developing countries, poor households, especially landless farm households, frequently cannot borrow against future income. Indeed, the lack of credit access is identified as one of the most serious sources of generating poverty (Deaton 1997; Morduch 1994). The binding credit constraint will have two grave consequences. First, mothers' labor income from outside will generate high value to the family. Accordingly, their labor participation will be induced significantly. Second, since the opportunity costs of child education are quite high in terms of a loss of marginal utility, the poor may choose optimally not to educate their children despite high rates of return on education. For these reasons, we investigate in this paper the determinants of simultaneous decision making of mother and child labor allocation under credit constraints. We believe our study bridge the existing gap in the two separate literatures, one on child labor and schooling (e.g., Jacoby 1994; Jacoby and Skoufias 1997; Beegle et al. 2003; Baland and Robinson 2000) and the other on labor supply (e.g., Kochar 1999; Skoufias 1996; Ito and Kurosaki 2006), both under credit constraints.

The remainder of this paper is organized as follows. Section 2 presents a theoretical model of intrahousehold resource allocation, from which we derive empirical specifications adopted in this paper. Section 3 describes the data used in our econometric estimation, characterizing children's activities observed in our dataset. Section 4 presents regression results for the covariates of children's time use. In Section 5, we empirically investigate the impacts of maternal employment on child time allocation in the study area. In Section 6, we delve deeper into the mechanism of children's time use by focusing on the role of credit constraints, as well as the determinants of those constraints. Section 7 concludes the paper.

2. Theoretical and Empirical Models

As a point of departure, we propose a two-period model for a household consisting of a child, his/her father, and his/her mother. Endogenous variables are C_t , F_t , M_t , L_t ($t=1, 2$), and S , where C_t is the household's consumption, F_t the father's labor supply, M_t the mother's labor supply, L_t the child's labor supply, and S the net saving in period 1. The first period budget constraint is given by

$$C_1 + S = Y_1 + w_F F_1 + w_M M_1 + w_C L_1, \quad (1)$$

And the second period budget constraint is

$$C_2 = Y_2 + (1+r)S + w_F F_2 + w_M M_2 + w_C f(T-L_1)L_2, \quad (2)$$

where r is the market interest rate, Y_t is unearned income, w_k is wage, $f(\cdot)$ is a function showing gross returns to human capital investment into the child ($f' > 0, f'' < 0$), and T is the time endowment. Endogenous variables are also subject to time and credit constraints

$$0 \leq F_t, M_t, L_t \leq T, \quad t=1,2 \quad (3)$$

$$S \geq -K, \quad (4)$$

where K is the exogenous borrowing limit imposed on the household. For simplicity, labor supply in the second period is assumed to be fixed and we focus on the 1st period decision making.

2.1. Labor Allocation under Unitary Household Models

Assuming that the household is a collective unit that behaves as a single economic agent (the unitary household assumption), the household's objective is to

$$\max U(C_1, F_1, M_1, L_1) + U(C_2, F_2, M_2, L_2)/(1+\delta) \quad (5)$$

with respect to C_1, F_1, M_1, L_1 , and S , subject to (1)-(4), where δ is the subjective discount rate. Inserting (1) and (2) into (5), letting λ denote the Lagrange multiplier on (4), and assuming an interior solution for (3), the first order conditions necessary for the optimality include

$$S: -U_{C1} + U_{C2}(1+r)/(1+\delta) + \lambda = 0, \quad (6)$$

$$F_1: U_{F1} + w_F U_{C1} = 0, \quad (7)$$

$$M_1: U_{M1} + w_M U_{C1} = 0, \quad (8)$$

$$L_1: U_{L1} + w_C U_{C1} - w_C f'(\cdot)L_2 U_{C2}/(1+\delta) = 0, \quad (9)$$

$$\lambda: \lambda(S+K) = 0, \quad (10)$$

where subscripts on U show partial derivatives.

The case with non-binding credit constraint

Since $\lambda = 0$, equation (6) becomes

$$U_{C1} = U_{C2}(1+r)/(1+\delta), \quad (6')$$

which is the standard Euler equation. In other words, marginal rate of substitutions are equalized to the interest rate. Equation (9) becomes

$$U_{L1} + w_C \{1 - f'(\cdot)L_2/(1+r)\} U_{C1} = 0, \quad (9')$$

which shows that $w_C \{1 - f'(\cdot)L_2/(1+r)\}$ is the shadow wage of child labor in period 1. When the child works in period 1, the household obtains wage income at the rate of w_C but the child labor reduces human capital investment and thereby reduces the wage rate in period 2. The household allocates child labor in period 1 considering this effect. If $U_{L1} = 0$ (the household does not obtain utility or disutility from child work), equation (9') becomes $f'(\cdot)L_2 = 1+r$, which is the standard condition for a pure investment. When $U_{L1} < 0$ (the household directly obtains disutility from child labor), the optimal child labor in period 1 decreases (human capital investment increases) than the pure investment case. Since equation (10) is automatically satisfied due to $\lambda = 0$, we can solve four equations (6'), (7), (8), and (9') for four endogenous variables S , F_1 , M_1 , and L_1 .

The case with binding credit constraint

Since $\lambda > 0$, equation (6) becomes

$$U_{C1} = U_{C2}(1+r)/(1+\delta) + \lambda > U_{C2}(1+r)/(1+\delta). \quad (6'')$$

Since the marginal utility of consumption in period 1 is higher than that in period 2, the household wants to increase the first period consumption but it cannot, because the credit binding is already binding ($S = -K$). Equation (9) becomes

$$U_{L1} + w_C \{1 - (1 - \mu)f'(\cdot)L_2/(1+r)\} U_{C1} = 0, \quad (9'')$$

where $\mu = \lambda/U_{C1}$, $0 < \mu < 1$. Equation (9'') shows that the difference of the shadow wage from the child's nominal wage is reduced. In other words, because of the binding credit constraint, the subjective interest rate goes up and the optimal human capital investment therefore decreases (the optimal child labor in period 1 increases).

By solving the system of five equations comprising (6''), (7), (8), (9), and the relation $S+K=0$ ((10'') from equation (10)), we obtain the optimal solution for five endogenous variables S , F_1 , M_1 , L_1 , and λ . The system can be divided into (10'') which includes only S , a sub-system of equations (7)-(9) which does not include λ , and (6'') which includes F_1 , M_1 , L_1 , and λ . Therefore, the system can be solved in a sequential way: first, determine S from (10''); second, solve the subsystem (7)-(9) to obtain F_1 , M_1 , L_1 ; third, insert these four into (6'') to obtain λ . This sequential solution is utilized in below to derive an endogenous switching model for the credit constraint.

2.2. A Non-Unitary Extension

Instead of the unitary household assumption, we can extend the theoretical model by

incorporating the bargaining process between the father and the mother, *a la* Browning and Chiappori (1998). Assuming that the intrahousehold resource allocation is Pareto-efficient and the parents have altruistic preferences, the father's utility is denoted by $U^F(C, F, M, L)$, and the mother's utility is denoted by $U^M(C, F, M, L)$, the household's objective is now to

$$\max \omega(EEP) \{U^F(C_1, F_1, M_1, L_1) + U^F(C_2, F_2, M_2, L_2)/(1+\delta^F)\} \\ + (1-\omega(EEP)) \{U^M(C_1, F_1, M_1, L_1) + U^M(C_2, F_2, M_2, L_2)/(1+\delta^M)\} \quad (5')$$

subject to the same constraints (1)-(4). *EEP* are called “extra-household environmental parameters” that shift the bargaining rule ω . Examples of *EEP*'s include local sex ratios, divorce law legislation, and the degree of prohibition on market work by gender. When the function $\omega(EEP)$ is a constant or the functional forms of U^F and U^M are exactly the same, the optimization problem (5') becomes equivalent to the optimization problem (5), or the unitary household model.

When $\omega(EEP)$ shifts with *EEP* and the functional forms of U^F and U^M are not exactly the same, the optimization problem (5') becomes a non-unitary household model. Under the non-unitary assumption, *EEP* variables should affect the optimal choice of S , F_1 , M_1 , L_1 , and λ . Therefore, we can test the relevance of the non-unitary assumption by testing exclusion restrictions on *EEP* variables.²

2.3. Determinants of Children's Time Use: Reduced-form approach

From the theoretical model above, we can derive our first empirical model, which is a completely reduced-form approach (in the followings, time subscripts are dropped and subscript h for a household and subscript i for a child are used). The reduced-form model is

$$L_{hi} = L(X_i, X_h, EEP_h) + u_{hi}, \quad (11)$$

where X is the shifter of market returns of child labor and schooling, the interest rate and credit constraints faced by the household, and the preferences of the household. More concretely, we include the child's age and sex in X_i , household head's characteristics, household's demographic characteristics, and household assets in X_h , and the non-resident grandparents' characteristics in EEP_h . u_{hi} is an error term with a mean of zero. To control for differences in local market conditions and preferences, we also include village fixed effects and community fixed effects.

We test the unitary assumption by testing the exclusion restrictions on EEP_h . Since this test may not be valid if our proxies for *EEP* variables also affect the preferences and investment technology directly, we have to be careful in interpreting the results, as pointed out by Kurosaki et al. (2006) and Fuwa et al. (2006b).

² Wage rates of the father and the mother may also shift ω . However, since wage rates also changes comparative advantages of market work and leisure within a household, they should also affect the marginal rate of substitution between consumption and leisure directly. Therefore, exclusion tests for wage rates cannot be used to test unitary against non-unitary household models (Alderman et al. 1995; Doss 1996).

As dependent variables for (11), we include not only the time used in remunerative works but also the time used in child care and domestic chores since these are also important work activities. We also examine the determinants of schooling time and leisure time. It is of great interest to test whether their determinants are the same as those for child labor (with the opposite sign).

2.4. Effects of Mother's Work on Child Labor: Conditional labor supply approach

Among the empirical mechanisms that lead to the intrahousehold resource allocation described by the reduced-form model of (11), we focus on the relationship between maternal employment and child time allocation. This is because the maternal employment should affect children's time use more than the paternal employment since mothers are usually responsible for domestic chores and child care in Indian households. For this purpose, we adopt a conditional demand function approach proposed by Pitt and Rosenzweig (1990) in the context of child time allocation and endogenous infant health. A child time demand function (or child labor supply function conditioned on mother's time use) for each activity can be derived from a household utility maximization model as follows:

$$L_{hij} = X_i \alpha_{1j} + X_h \alpha_{2j} + \alpha_{3j} M_h + u_{hij}, \quad (11')$$

where j denotes child type (as defined over sex and/or age group). M is the endogenous maternal employment, and X_h is a vector of household specific exogenous variables. While the activities of all household members affect child i 's time allocation, we are primarily interested in the impact of maternal employment on children's time allocation to various activities. Therefore, we allow maternal employment to appear as an argument in (11'), and to be determined endogenously through the process described in Subsections 2.1-2.2. Note that the parameter vector α is allowed to depend on child type, reflecting the differences across sex and age groups of the impact of maternal employment and individual/household characteristics.

Estimating equation (11') involves several issues. First, maternal employment is endogenous and may be correlated with unobservables (e.g. household preference) embodied in the error term of the child time use equation. Secondly, since we are looking at a sample of children including siblings belonging to the same households, it will be desirable to control for household-specific effects such as unobserved preference for education.

In order to address these issues, we utilize the method proposed by Pitt and Rosenzweig (1990). Rather than estimate equation (11') directly for each child, we take differences between two children from the same household. Doing so creates two advantages: First, it would allow us to remove any household-specific effects that enter into the time use equations. Moreover, taking differences between children will allow us to justify the use of certain instrument variables for maternal employment (Pitt and Rosenzweig 1990: 974-975). For example, while irrigated acreage affects

maternal employment, it may also affect child time use. However, if we assume that irrigated acreage enter in an identical manner the time use equations of different child-types, then taking differences cancels out the term containing it. Thus irrigated acreage becomes a candidate instrument variable for maternal employment. To see this, suppose that for a given activity, children of types b (boys) and g (girls) have the following time use equations:

$$L_{hib} = \alpha_0 + \alpha_1 X_{ib} + \alpha_{2b} X_{2h} + \alpha_2 X_{3h} + \alpha_{3b} M_h + u_{hib}, \quad (12)$$

$$L_{hi'g} = \alpha_0 + \alpha_1 X_{i'g} + \alpha_{2g} X_{2h} + \alpha_2 X_{3h} + \alpha_{3g} M_h + u_{hi'g}, \quad (13)$$

Subtracting (12) from (13) yields

$$L_{hi'g} - L_{hib} = \alpha_1 (X_{i'g} - X_{ib}) + (\alpha_{2g} - \alpha_{2b}) X_{2h} + (\alpha_{3g} - \alpha_{3b}) M_h + (u_{hi'g} - u_{hib}), \quad (14)$$

which no longer contains the regressors in X_{3h} . If X_{3h} forms part of the exogenous regressors in the maternal employment equation, then the variables therein can be used as instrument variables. Above all, we will be interested in estimating the parameters $(\alpha_{3g} - \alpha_{3b})$, which measures the difference between child-types in the effect of maternal employment M on time allocation. Therefore, we simultaneously estimate the system comprising equation (14) and a reduced-form regression model for the maternal labor supply equation

$$M_h = \beta_0 + \beta_2 X_{2h} + \beta_3 X_{3h} + u_h. \quad (15)$$

Since this is a system of two endogenous variables with a triangle structure, we can estimate (15) in the first stage, followed by the second stage IV estimation of equation (14) with X_{3h} serving as the identifying instrumental variables.

2.5. Effects of Credit Constraint on Child Labor: Endogenous switching model approach

Another important determinant of children's time use, which is also endogenous, is the credit constraint. Although it is possible to estimate a conditional labor supply equations similar to equation (14), such an approach is unsatisfactory since it can elicit the difference effect only, not the level effect, of an endogenous variable. Since the ill effect of credit constraint is often discussed in the context of poverty reduction policies, it is more desirable to investigate the level effect of credit constraints. For this purpose, this paper proposes an endogenous switching model that can be derived from the sequential structure of the theoretical model discussed above.

Denoting a household under a binding credit constraint with subscript c and a household otherwise with subscript u , the theory suggests an empirical model

$$L_{hic} = \alpha_{0c} + \alpha_{1c}X_i + \alpha_{2c}X_{1h} + \alpha_{3c}K_h + u_{hic}, \quad (16a)$$

$$\text{if } \lambda_h = \lambda(X_{1h}, K_h) > 0,$$

$$L_{hiu} = \alpha_{0u} + \alpha_{1u}X_i + \alpha_{2u}X_{1h} + u_{hiu}, \quad (16b)$$

$$\text{if } \lambda_h = \lambda(X_{1h}, K_h) = 0,$$

where X_i and X_{1h} are vectors of exogenous variables that affect labor supply in the absence of credit constraints, such as the shifter of market returns of child labor, schooling, and the preferences of the household. In contrast, K_h is a vectors of exogenous variables that shift K in the theoretical model. K_h directly affect children's time use only when the credit constraint is binding.

Since λ_h is not observable, we replace this by a dichotomous variable of cc_h , which takes the value of one when the constraint is binding and zero otherwise (Sawada et al. 2006). We discuss in the next section on data how to observe cc_h . By combining (16a) and (16b), we obtain an endogenous switching model

$$L_{hi} = (\alpha_{0u} + \alpha'_{0c}cc_h) + (\alpha_{1u} + \alpha'_{1c}cc_h)X_i + (\alpha_{2u} + \alpha'_{2c}cc_h)X_{1h} + \alpha_{3c}cc_hK_h + u_{hi}, \quad (17)$$

$$cc_h = 1 \text{ if } \lambda_h^* \geq 0 \text{ and } cc_h = 0 \text{ if } \lambda_h^* \leq 0, \quad (18)$$

$$\lambda_h^* = \beta_1X_{1h} + \beta_2K_h + e_h, \quad (19)$$

where $\alpha'_{0c} = \alpha_{0c} - \alpha_{0u}$, $\alpha'_{1c} = \alpha_{1c} - \alpha_{1u}$, $\alpha'_{2c} = \alpha_{2c} - \alpha_{2u}$, λ_h^* is a latent variable associated with the credit constraint. We estimate the system comprising equations (17)-(19). When $\alpha'_{1c} = \alpha'_{2c} = \alpha_{3c} = 0$, equation (17) is reduced into

$$L_{hi} = \alpha_0 + \alpha'_{0c}cc_h + \alpha_1X_i + \alpha_2X_{1h} + u_{hi}, \quad (17')$$

which allows the intercept to shift depending on the regime. The system comprising (17'), (18), and (19) is an empirical model of endogenous program placement that is popular in the literature on program evaluation. Both types can be estimated as Type 5 Tobit Model of Amemiya (1985), with endogenous and observable regimes. In our empirical part, we first estimate the system with (17') by a maximum likelihood method and test the significance of α'_{0c} . Then the system with (17) is estimated by a two step method in which some of the coefficients are restricted to have the same coefficient under both regimes.

2.6. Effects of Credit Constraint on the Cross-effects between Mother's and Children's Time Use

Finally, we examine how (endogenous) credit constraint affects the pattern of the

cross-persons effects in time use between the mother and her children (as examined by the approach described in section 2.4). In effect, we combine the conditional labor supply approach to examine the (endogenous) effect of mothers' labor supply on children's time use (section 2.4) and the switching regression model to differentiate the 'determinants of time use' equations under two regimes—one for credit-constrained households and the other for unconstrained households (section 2.5). The conditional labor supply equations, (12) for boys and (13) for girls, can be separated into two 'regimes,' credit constrained and unconstrained, in the same way as the distinction between equations (16a) and (16b):

$$L_{hibc} = \alpha_{0c} + \alpha_{1c}X_{ib} + \alpha_{2bc}X_{2h} + \alpha_{2c}X_{3h} + \alpha_{3bc}M_h + u_{hibc}, \quad (12a)$$

if $\lambda_h = \lambda(X_{1h}, K_h) > 0$ (boys' time use in credit constrained households),

$$L_{hibu} = \alpha_{0u} + \alpha_{1u}X_{ib} + \alpha_{2bu}X_{2h} + \alpha_{2u}X_{3h} + \alpha_{3bu}M_h + u_{hibu}, \quad (12b)$$

if $\lambda_h = \lambda(X_{1h}, K_h) = 0$ (boys' time use in unconstrained households),

$$L_{hi'gc} = \alpha_{0c} + \alpha_{1c}X_{i'g} + \alpha_{2gc}X_{2h} + \alpha_{2c}X_{3h} + \alpha_{3gc}M_h + u_{hi'gc}, \quad (13a)$$

if $\lambda_h = \lambda(X_{1h}, K_h) > 0$ (girls' time use in credit constrained households),

and

$$L_{hi'gu} = \alpha_{0u} + \alpha_{1u}X_{i'g} + \alpha_{2gu}X_{2h} + \alpha_{2u}X_{3h} + \alpha_{3gu}M_h + u_{hi'gu}, \quad (13b)$$

if $\lambda_h = \lambda(X_{1h}, K_h) = 0$ (girls' time use in unconstrained households).

As discussed in section 2.5, whether a household is credit-constrained is determined endogenously. As a result, the expectation of the error terms (u 's) in the above equations are generally not equal to zero. However, assuming joint normality of error terms, we can rewrite each error term as follows (e.g., Maddala 1983);³

$$u_{hibc} = -\sigma_{bce} \frac{\phi(\beta_1 X_{1h} + \beta_2 K_h)}{\Phi(\beta_1 X_{1h} + \beta_2 K_h)} + \varepsilon_{hibc}, \text{ where } E(\varepsilon_{hibc} | X, M, K, cC_h = 1) = 0,$$

$$u_{higc} = -\sigma_{gce} \frac{\phi(\beta_1 X_{1h} + \beta_2 K_h)}{\Phi(\beta_1 X_{1h} + \beta_2 K_h)} + \varepsilon_{higc}, \text{ where } E(\varepsilon_{higc} | X, M, K, cC_h = 1) = 0,$$

³ We can elaborate an empirical framework without the assumption of multinormal distribution (Lee 1982). For example, Newey et al. (1990) proposed a procedure of nonparametric approximations for the sample selection correction terms. Yet, in general, these extensions provide comparable results to those derived under the joint normality assumption.

$$u_{hibu} = \sigma_{bu\epsilon} \frac{\phi(\beta_1 X_{1h} + \beta_2 K_h)}{1 - \Phi(\beta_1 X_{1h} + \beta_2 K_h)} + \varepsilon_{hibu}, \text{ where } E(\varepsilon_{hibu} | X, M, K, cc_h = 0) = 0,$$

and

$$u_{higu} = \sigma_{gu\epsilon} \frac{\phi(\beta_1 X_{1h} + \beta_2 K_h)}{1 - \Phi(\beta_1 X_{1h} + \beta_2 K_h)} + \varepsilon_{higu}, \text{ where } E(\varepsilon_{higu} | X, M, K, cc_h = 0) = 0.$$

In these error terms, $\phi(\bullet)$ and $\Phi(\bullet)$ represent the standard normal density function and distribution function, respectively. The indicator variable, cc_h , indicating whether a household is credit constrained, is defined as in (18) and (19) above. As was done in section 2.4, we can define the difference in time use between girls and boys under the alternative regimes based on the household status in terms of credit constraint; subtracting (12a) from (13a), and (12b) from (13b), respectively, yields (14a') and (14b') as follows:

$$\begin{aligned} L_{hi'gc} - L_{hibc} &= \alpha_1(X_{i'gc} - X_{ibc}) + (\alpha_{2gc} - \alpha_{2bc})X_{2h} + (\alpha_{3gc} - \alpha_{3bc})M_h + (u_{h'gc} - u_{hibc}), \\ &= \alpha_1(X_{i'gc} - X_{ibc}) + (\alpha_{2gc} - \alpha_{2bc})X_{2h} + (\alpha_{3gc} - \alpha_{3bc})M_h \\ &\quad - (\sigma_{gu\epsilon} - \sigma_{bu\epsilon}) \frac{\phi(\beta_1 X_{1h} + \beta_2 K_h)}{\Phi(\beta_1 X_{1h} + \beta_2 K_h)} + (\varepsilon_{h'gc} - \varepsilon_{hibc}), \end{aligned} \quad (14a')$$

if $\lambda_h = \lambda(X_{1h}, K_h) > 0$ (credit constrained), and

$$\begin{aligned} L_{hi'gu} - L_{hibu} &= \alpha_1(X_{i'gu} - X_{ibu}) + (\alpha_{2gu} - \alpha_{2bu})X_{2h} + (\alpha_{3gu} - \alpha_{3bu})M_h + (u_{h'gu} - u_{hibu}) \\ &= \alpha_1(X_{i'gu} - X_{ibu}) + (\alpha_{2gu} - \alpha_{2bu})X_{2h} + (\alpha_{3gu} - \alpha_{3bu})M_h \\ &\quad + (\sigma_{gu\epsilon} - \sigma_{bu\epsilon}) \frac{\phi(\beta_1 X_{1h} + \beta_2 K_h)}{1 - \Phi(\beta_1 X_{1h} + \beta_2 K_h)} + (\varepsilon_{h'gu} - \varepsilon_{hibu}), \end{aligned} \quad (14b')$$

if $\lambda_h = \lambda(X_{1h}, K_h) = 0$ (credit unconstrained),

where $E(\varepsilon_{higc} - \varepsilon_{hibc} | X, M, cc_h = 1) = E(\varepsilon_{higu} - \varepsilon_{hibu} | X, M, K, cc_h = 0) = 0$.

If we assume that the covariance between the error term in the time use equation (u 's) and the error term in the determination of credit access (e_h 's) is the same for boys and for girls (i.e., $\sigma_{gu\epsilon} = \sigma_{bu\epsilon}$ and $\sigma_{gu\epsilon} = \sigma_{bu\epsilon}$), then the inverse Mills ratios drop out from equations (14a') and (14b') (c.f. Pitt and Rosenzweig 1990, p. 978). Thus, equations (14a') and (14b'), together with the determination of mothers' labor supply equation (15), can be estimated by two stage least squares estimation, in the same way as equations (14) and (15) with the only difference that we now split the sample into two groups, credit constrained households and unconstrained households: equation (14a') is estimated using only the sample households with binding credit constraint while Equation (14b') is estimated using only the sample households with unconstrained credit access.

3. Data

3.1. The Household Survey in Rural Andhra Pradesh, India

We utilize data from a survey of approximately 400 rural households in Kurnool district of the southern Indian state of Andhra Pradesh. The survey, conducted during February-March 2005 in a cluster of 32 villages, was designed to obtain various household-level and individual-level information on production, consumption, asset holdings, time allocation, and several other economic aspects. Further details on survey design and objectives are found in Fuwa et al (2006a)⁴.

The study region belongs to the semi-arid tropics of Deccan Plateau, notorious for high risk in agricultural production (Walker and Ryan 1990). The survey period of February-March is usually characterized by abundant demand for agricultural labor but the specific period of February-March 2005 experienced a drought, resulting in lower demand for farm labor. Nevertheless, we observed numerous instances of child labor. In addition, the state of Andhra Pradesh is a region with higher incidence of child labor than in other states of India.

For these reasons, our dataset covers sample households with higher incidence of child labor than found in other data sources in India. For instance, at the all India level (NSS dataset 1999/2000), the child labor incidence ratio among children aged 10-14 was 12.5% when a wider definition of child labor including household chores is used (Edmonds et al. 2005); in UP and Bihar (LSMS dataset 1997/98), where income poverty is more severe than in other regions of India, the child labor incidence ratio was reported to be around 28.3 percent (Sakamoto 2006); in our sample, the corresponding figure is 54.2 percent (Kurosaki et al. 2006).

3.2. Time Allocation of Children

Our survey contained a “one week time use module” whose reference period was the 7 days immediately prior to the interview date. The questionnaire asked the respondent about his/her activity in each “half-day” (AM or PM) during the reference period, so that a total of 14 half-days were classified as belonging to the following categories:

- 1) Remunerated work (including labor on own farm/enterprise)
- 2) Non-remunerated work
- 3) Household chores
- 4) Child care
- 5) Schooling
- 6) Social activities
- 7) Leisure

⁴ We randomly chose sample households using a variable probability sampling method in order to collect a sufficient number of households containing child labor. In the statistical and econometric analyses of this paper, we corrected for the difference in the sampling probability by weighting. See Fuwa et al. (2006a) for details.

- 8) Sickness
- 9) Other, specify

Borrowing the classification in the ILO standards, we cover children in the age group 5-14 in this paper. Table 1 summarizes the one week time use data for about one thousand children aged 5-14 included in the sample households. Among the nine categories mentioned above, we list *schooling* (category 5), *chores* (category 3), *child_care* (category 4), *rem_work* (category 1), and *leisure* (category 7). Child time use in each activity is measured as the number of half-days spent on that activity during the reference period, so that each variable takes on integer values between 0 and 14.

The table shows that *schooling* and *leisure* time are longer among younger children than among older children, and longer among boys than among girls. The gender difference is statistically significant at 1 to 10 percent levels. Time spent on market-oriented work (*rem_work*) is similar among boys and girls. Older children spend more time working for the market than younger children. Among domestic works (*chores* and *child_care*), girls spend more time than boys but the gender difference is statistically insignificant for the child care among older children. This seems to indicate a tendency that girls spend more time in domestic works and both girls and boys spend more time as they grow, resulting in shorter time spent on schooling and leisure among older girls. We will investigate whether these observations based on the bivariate comparison hold even under the multivariate context, by estimating the reduced-form model described in Subsection 2.3.

Since the mothers of these children also provide domestic works, we investigate the time use of children according to the working status of mothers (Table 2). Maternal employment is described by an indicator variable equal to one if there was at least one half-day of remunerative work during the reference period by a female household member having one or more children, regardless of whether or not she is the wife of the household head. Otherwise, the indicator takes on a value of zero.

Table 2 shows that except for the number of infants in the household and girls' time at remunerative work, there is no statistically significant difference between households with maternal employment and those without maternal employment, in terms of the number of children and child time use patterns. Gender differences are apparent in Table 2, most notably that girls in households with maternal employment spend significantly more time in remunerative work, whereas such tendencies are weaker and insignificant for boys. Also, the gender gaps in schooling and leisure seem to be greater under maternal employment than under no maternal employment.

While these observations are interesting in themselves, they may mask the true relationship between maternal employment if there are individual or household characteristics, either observable or unobservable, that affect maternal employment and child time use. For this reason, we will estimate the conditional labor supply model described in Subsection 2.4.

3.3. Credit Constraints

The conventional empirical approach to incorporating credit constraints into an estimation model is to ignore the endogeneity of the constraints and to split the sample into those who are likely to be credit constrained and those who are not likely to be credit constrained exogenously (Zeldes 1989; Morduch 1990). The exogenous split approach, however, has two problems. First, it is unlikely that a single variable such as income-wealth ratio or land ownership is a good predictor of consumers' ability to borrow (Garcia et al. 1997, p.158; Jappelli 1990). Second, credit constraint is endogenously generated and thus should be treated as an endogenous variable. Otherwise, estimation results will suffer from endogeneity bias (Scott, 2000).

In order to overcome these problems, we designed the credit module carefully in our questionnaire so that we can identify credit-constrained households directly, as is suggested by Scott (2000). In identifying credit constraints, household heads were asked about members' experience with credit suppliers during the 12 months prior to the survey. To construct liquidity constraint indicators with sufficient variation across households, we decided to concentrate on formal credit sources. Our focus on formal credit derives from the following reasoning: First, a clear division between credit constrained and unconstrained households is likely to emerge in the context of bank or formal credit in the study region (Pender 1996), because access is often determined by the household's ability to provide collateral, which generally depends on ownership of land title; on the other hand, informal credit comes in numerous forms so that it is difficult to classify households according to credit constrained-ness, and the determinants of access are less clear cut. The second reason for focusing on bank credit is that over the last few decades, formal sources of finance have become more accessible and important to the village economy in the study area. Given the increasing importance of formal credit, its impact on household behavior is interesting in itself.

To identify credit constrained households, we asked first whether a household tried to obtain a loan in a particular period. Then, for those who tried to borrow money, we asked whether a household could borrow as much as they requested. If the answer was yes, we identified the household as unconstrained. On the other hand, we identify those households who had their loan applications rejected, or could not borrow sufficiently as being credit constrained.

Second, for those who did not try to borrow, the enumerators questioned the respondents about the reason for not availing of the bank loans. The answer choices were:

- 1) No need for credit
- 2) Not want to be in debt
- 3) Terms are not attractive (duration too short, interest rate too high, etc.)
- 4) Too much paperwork
- 5) Live too far from lender
- 6) Already have large amount of debt
- 7) Believed I would be refused by lender
- 8) Don't know how to get credit / not know lender

9) Don't know anyone who can be guarantor

10) Other, specify

Respondents who chose one of the options 3)-9) as the reason for not attempting to obtain bank credit were identified as households most likely to be credit-constrained with regard to formal sources. The remaining respondents who did not try to borrow were considered to be unconstrained. This definition is called the “broad definition of credit constraint” and denoted by a dummy variable cc_1 in this paper. Respondents who chose one of the options 3)-5) might not be credit-constrained in reality. Hence, we attempted a similar procedure to construct an indicator variable under the “narrow definition of credit constraint,” identifying as constrained households choosing one of the options 6)-9). We denote this status by a dummy variable cc_2 in this paper.

On the basis of these responses, we can identify the credit-constrained households who were not able to access credit. Since almost none of the existing multi-purpose household panel surveys include direct questions that identify credit constraints (Scott 2000), our data set provides us with valuable information directly to separate the constrained and unconstrained households.

Table 3 shows the descriptive statistics of all the 331 households used in this study. Among them, 205 households (or 61.9 percent of the total) are identified to be credit-constrained under the wider definition, and 164 households (or 49.5 percent) are identified to be credit-constrained under the narrower definition, indicating that a significant proportion of households are constrained. While the age and education profiles of the constrained and un-constrained households are similar, the other characteristics of the two groups are remarkably different. The value of land owned by un-constrained households is higher in comparison to that of constrained ones. The number of household members in constrained households is larger than that in un-constrained households. In summary, the descriptive statistics partly confirm the documented stylized facts regarding poor households in developing countries who are credit-constrained as far as the land asset is concerned, but not a clear contrast with respect to household education. We will investigate whether these observations based on the bivariate comparison hold when other factors are controlled for, by estimating the endogenous switching model described in Subsection 2.5.

4. Determinants of Children’s Time Use: Estimation results of the reduced-form model

Explanatory Variables: To estimate equation (11) in Subsection 2.3, the following independent variables are included in the basic model (Model 1): (1) Individual characteristics of a child: age , $agesq$ (defined as $(age-5)^2$, to capture non-linearity of the age effect), and sex (a dummy for a girl); (2) Household characteristics: hd_age (the age of the household head to control for the lifecycle effect), f_edud (the schooling dummy for the child’s father), m_edud (the schooling dummy for the child’s mother), their cross terms with sex to investigate the gender disparity among children, $hhsiz$ (the number of household members) and its composition by age and sex, $bplhold$ (a dummy variable for the ration card holder under the Public Distribution System of the Government of India), $landval$ (the total amount of land

owned by the household in 100,000 Rs., to control for the household's main asset), *bullock_q* (the number of bullocks owned by the household, to represent the livestock asset), and dummy variables for communities (religion and wider caste groupings: *SC* [scheduled castes], *ST* [scheduled tribes], *UMH* [upper and medium Hindu castes], and *Muslim* [Muslims], with the reference category as those households belonging to so-called “other backward classes” [OBC]); and (3) Village fixed effects, which collectively control for differences in market conditions, environments, and school qualities.

If mothers' education is more important to girls and fathers' education is more important to boys in reducing child labor and in increasing school enrollment, we expect coefficients on *m_edud*sex* and *f_edud*sex* to have the opposite signs (Thomas 1994; Quisumbing and Maluccio 2003). Variable *bplhold* is a proxy for the government's labeling that the household is not rich. Thus the variable is expected to capture the effect of poverty or the effect of households' interaction with poverty reduction policies. Variables *landval* and *bullock_q* are meant to capture the wealth effect, which is theoretically predicted to have a negative impact on child labor and a positive impact on school enrollment.

In India, it is often claimed that scheduled castes and tribes are backward strata with lower interests in education. If this is correct, we expect coefficients on *SC* and *ST* to be positive on child labor and negative on enrollment. We will examine whether this holds even when we control for other individual and household characteristics. We also expect that the inclusion of community dummies (or more detailed caste fixed effects) reduces the possible bias due to omitted variables at the household level.

In Model 2, additional variables are included, which characterize grandparents of the child. Many of them are dead already or live separately from the current household (note that we limit the analysis to children belonging to the standard family type with the child's mother and father present in the household). Therefore, a prediction of unitary household models is that the grandparental characteristics should not affect the child's working status if we sufficiently control for the returns of his/her labor and schooling and the household's wealth and credit access status. If these characteristics affect the child's working status, it may be a reflection of father's and mother's bargaining parameters, which should affect the child's working status under the assumptions of non-unitary household models. In other words, we use the characteristics of the parents' parents as EEPs to distinguish unitary and non-unitary household models.⁵ Concretely, the additional variables include: *hdf_lit* (literacy of the father's father), *hdm_lit* (literacy of the father's mother), *spf_lit* (literacy of the mother's father), *spm_lit* (literacy of the mother's mother), *hdp_adiff* (the age difference of the father's parents), and *spp_adiff* (the age difference of the mother's parents). We can test whether these proxies for EEPs affect child labor and schooling by a chi-squared test for the null hypothesis that all coefficients on these variables are zero.

⁵ This test may not be ideal, since our dataset is only a cross-section so that omitted-variable bias may be serious enough. We also acknowledge that the significance of coefficients on any grandparental variables does not rule out preference-based explanations consistent with unitary models if certain traits or preference may be transmitted through generations. For example, a mother whose mother is educated may reveal a preference for greater investments on her daughter and such preference is reflected in the household's unitary utility function. Thus, we need to be careful in the interpretation of the results.

Regression Results: Table 4 reports summary statistics of the empirical variables (the table also includes those of the empirical variables used in Sections 5-6) and Table 5 reports estimation results. Considering the censoring of the dependent variables (they are restricted to lie between zero and 14), two-sided tobit models are estimated with village and community dummies. Because two dependent variables of *chores* and *child_care* do not have a sufficient number of uncensored observations, we merge the two into a new dependent variable of *dom_work* in the tobit regressions. Coefficients on village fixed effects are not reported for brevity. The estimation results are qualitatively the same when linear models are estimated instead of tobit models.

Individual attributes of age and sex are important determinants of children's time use. Girls' *dom_work* is longer than boys' but the gender difference in *rem_work* is not statistically significant, resulting in less time for girls to spend on *schooling* and *leisure*. The age effect is linearly positive on *dom_work* and *rem_work* while it is concave on *schooling* and convex on *leisure*. The marginal effect of age on *schooling* is positive when the child is five years old but it declines continuously and becomes negative at the age of eight. The marginal effect of age on *leisure* is negative when the child is five years old and it is attenuated continuously but still remains negative at the age of 14. These results indicate that older children sacrifice their schooling and leisure time for working time, both inside and outside the household.

Second, the effect of parental education is in the direction of increasing schooling and decreasing child labor as documented in the existing empirical studies on India (Aggarwal 2004; Basu et al. 2003; Deb and Rosati 2002; Dr'eze and Kingdon 2001; Sakamoto 2006), but most of these effects are insignificant. Father's education is insignificant in all specifications, while mothers' education significantly increases *schooling* of girls and decreases *rem_work* on both boys and girls. The contrast between fathers' and mothers' education may seem readily consistent with the bargaining hypotheses under non-unitary household models: mothers prefer more education for children than fathers and mothers' bargaining power is increased by mothers' relative position in education. However, since education also affects market and reservation wages of mothers and fathers, our results can also be compatible with the unitary approach. For instance, better educated mothers may raise the returns to children's education more than better educated fathers, since, for example, (stay-home) mothers are arguably in a better position to facilitate children's learning (e.g., through helping their homework) than are fathers.

Among other household characteristics, the household demographic size (*hsize*) has a negative effect on *rem_work*. Adding the household demographic composition shares did not improve regression results and they had jointly insignificant coefficients. Therefore, we report in Table 5 the results without these shares.

The coefficient on *landval* is significantly positive on *schooling* and *leisure*. This is consistent with the poverty effect hypothesis. However, the negative coefficients are mostly insignificant on child labor, while the positive effect on school enrollment is statistically significant. Thus the wealth impact is stronger on school enrollment but it is not discernible on reducing child labor. In other words, child labor is found to be almost constant over the (sampled) support of wealth distribution, holding other variables fixed. This indicates low wealth level, conditional on other covariates, is not a sufficient condition of child labor. These results were robust to the

inclusion of other controls or disaggregation of assets into its components such as land, building, livestock, etc. (Kurosaki et al. 2006). In South Asia, some authors find that farm land ownership is associated with more child labor because the productivity effect through family labor when wealth takes the form of land may dominate the usual wealth effect (so-called “wealth paradox”: see Bhalotra and Heady 2003). One reason for the absence of the wealth paradox could be that our survey was conducted in a drought year, resulting in a smaller productivity effect through family labor when wealth takes the form of land.

The estimation results of Model 2 show that some of the coefficients on the characteristics of grandparents of the child (proxy for EEPs) are statistically significant and they are jointly significant on *schooling*, *dom_work*, and *rem_work*. Variables that can be associated with a higher bargaining power of the mother such as *spp_adiff* and *spf_lit* are positively correlated with *schooling* and negatively correlated with *dom_work*. This seems to suggest the superiority of the non-unitary, bargaining model over the unitary household model discussed in Subsection 2.2. This argument is valid if these variables represent EEPs only. It is not impossible to interpret the same results under the unitary framework as well: these variables may capture the household’s access to quality education that is not sufficiently controlled by other variables. If this is the case, the variables included in Model 2 may not be good proxies for EEPs.

Finally, the effects of the community dummies remain even after controlling for individual and household characteristics and village fixed effects. In the *rem_work* regression, coefficients on *SC*, *ST*, *UMH*, and *Muslim* are negative with statistical significance, implying that households belonging to these groups are less likely to send children to remunerated work than households belonging to “other backward classes.” In the *schooling* regression, coefficients on *UMH* and *Muslim* are positive and statistically significant, implying that households belonging to upper and medium castes and Muslims are more likely to send children to school than households belonging to “other backward classes.” The effects of the upper and medium castes dummy are as expected since in the rural setting of India, these castes are regarded as socially advanced. However, the signs of the coefficients on *SC* and *Muslim* dummies are opposite to our expectation and to the findings of Deb and Rosati (2002), Dr’eze and Kingdon (2001), Aggarwal (2004), and Sakamoto (2006). Though not significant in several cases, the signs of the coefficients on *ST* also suggest that the welfare of scheduled-tribe children is better than that of children of “other backward classes.” These households are more enthusiastic about schooling and more averse to child labor than “other backward classes.” This may reflect the impact of civil movements in rural Andhra Pradesh to improve the social conditions of the scheduled castes, scheduled tribes, and Muslim households. The findings above were robustly found from different specifications using the indicator variables of children’s usual status, rather than the time use (Kurosaki et al. 2006).

5. Child Time Allocation and Mothers’ Work: Estimation results of the conditional labor supply model

Determinants of Mothers’ Work: We estimate a system of two equations: equation (14)

for the intra-household difference in children's time use and equation (15) for the determinants of mothers' work. To be consistent with the reduced-form approach, we use the same set of individual-level and household-level variables used in Section 4. As X_{2h} (household-level variables that appear both in equations (14) and (15), we include *infants* (the number of infants in the household). As X_{3h} (household-level variables that appear only in equation (15)), we use all other variables.

The first stage regression results for the determinants of mothers' work are shown in Table 6. While we consider liquidity constraint to be a crucial determinant of maternal employment, analysis of its effect on maternal employment is left for Sawada et al. (2006) and the results in Table 6 should be interpreted as the reduced-form with respect to the endogenous liquidity constraint. The education level of the child's father increases the time of maternal employment, while that of the mother has insignificant effect. This is against the findings in developed countries where more educated women are more likely to work outside the household. As Mammen and Paxson (2000) show using Indian household data, more than 60 percent of both rural and urban working women at the lowest levels of per capita expenditure work for wages. In our dataset, the most important jobs of *motwor* (the dependent variable of Table 6) are agricultural wagework and self-employment in agriculture. It is natural to observe that more educated women do not work in these sections.

In contrast to the results in Section 4, the household demographic compositions are significant predictors of mothers' time use. Here again, against the findings in developed countries where women with more school aged children are less likely to work outside the household, our dataset shows that mothers with higher number of such children are more likely to work outside the household (*s_cm* and *s_cf* have significantly positive coefficients). This reflects higher demand for money earned by mothers when the number of such children is large and such demand is facilitated by such children's help in sharing the work burden within the household. The evidence for the latter mechanism is shown immediately below.

The effect of *landval* is significantly positive, reflecting the tendency that mothers in more landed households spend more time working on their family farm. In other words, this effect shows that family labor and hired labor are incomplete substitutes in conducting farming operations.

The results of Model 2 show that some coefficients on the characteristics of the child's grandparents (proxy for EEPs) are statistically significant and they are jointly significant at the 5 percent level. This again seems to suggest the superiority of the non-unitary, bargaining model over the unitary household model discussed in Subsection 2.2. Since the EEP variables are jointly significant, Model 2 in Table 6 is used as the first stage regression for equation (14). Finally, coefficients on the community dummies show that mothers belonging to scheduled tribes (*ST*) work longer and mothers belonging to *Muslim* and *UMH* work shorter than mothers belonging to "other backward classes," which matches the general pattern of labor force participation in India.

Results for the Girl-Boy Subsample: We first estimate the system of (14) and (15) on the subsample containing only girl-boy child pairs. We focus on the impact of maternal employment on the following activities: schooling, household chores, child care, remunerated work, and leisure. In order to create the differenced dependent variable as in equation (14), we generate all possible girl-boy pairs within each household. Pairs are sorted so that the differences are taken in a specific order: namely,

“girl minus boy.”

The parameter on maternal employment is allowed to vary by sex, so that the coefficient on maternal employment, *motwor*, measures the difference between the coefficient for girls and the coefficient for boys. According to our ordering rule, a positive coefficient on *motwor* can be interpreted as the girls’ parameter on maternal employment being greater than the boys’ parameter.

In this specification, the only individual child characteristic that we include in the time use difference equation (14) is *d_age* (difference in age within the child pair), because difference in *sex* enters only through differences in the coefficients. The household-level variable, other than maternal employment, that is not canceled out in the differencing is assumed to be *infants*. This is justified by the observation in the literature that infants requiring child care tend to affect more strongly the time use pattern of girls than that of boys (Pitt and Rosenzweig 1990).

The first part of Table 7 presents the results from the girl-boy pair subsample. The second stage linear regressions present some significant differences between girls and boys in terms of the impact of maternal labor on time allocation. In the *schooling* regression, the impact of maternal employment on girls’ schooling time is smaller (or more negative) than on boys’ schooling, implying that girls shorten their schooling by more than boys in response to mother’s work outside the house. The coefficient on *d_age* is significantly negative, implying that the girl-boy schooling gap grows with the age difference between the two children, which in turn implies that older girls spend less time at school.

The result for the *chores* regression is in stark contrast to the result for *schooling*. It is seen that maternal employment significantly increases the difference between girls and boys in the amount of time spent on household chores. Specifically, girls spend more time on chores relative to boys when her mother takes up employment. In the same column, we see that the coefficient on age, which can be interpreted as the direct effect of age on chores time according to equation (14), is significantly positive, so that older children spend more time on chores.

The *child_care* equation shows that having infants in the same household significantly increases the difference between girl and boy time spent on child care. This is quite intuitive. In the *rem_work* equation, we do not find any significant effect of maternal employment on girl-boy differences. In the last column, we see that leisure follows a similar pattern as schooling: maternal employment affects negatively the difference “girl leisure time minus boy leisure time.” When this difference is negative to begin with, maternal employment exacerbates the difference.

Results for the Full Sample: In order to incorporate information from same-sex child pairs, we estimate the system defined by equations (14) and (15) on the full sample of child pairs. Here, we allow the parameter on maternal employment indicator to vary by relative birth order within pairs as well as by sex. If the parameter varied only by sex, then *motwor* would be canceled out in same-sex child pairs, defeating the purpose of pooling observations.

In constructing the differenced dependent variable, we generate all possible pairs of children within each household. Pairs are sorted so that the differences are taken according to the following general rules: “girl minus boy” and “older child minus younger child”. When there is a conflict between sex and age ordering, then sex is given priority, so that a girl-boy pair where the girl is younger than the boy

would be differenced according to the “girl minus boy” rule.

As with the girl-boy subsample, we include *age* as an individual characteristic, and *infants* as the household specific exogenous variable that is not cancelled out in the differencing. Covariates in the first-stage regression are the same as in the girl-boy subsample.

In order to allow the coefficient on *motwor* and *infants* to vary across child-types, we include several cross-terms involving the endogenous variable. For example, the interaction term *motwor*difsex*oldyoung* will pick up the difference in the maternal employment parameter between a girl and her younger brother.

The second half of Table 7 shows the result from the full sample. The *schooling* equation shows that the girl-boy difference in the effect of maternal employment on schooling is more significant when the girl is older than the boy (see coefficient on *motwor*difsex*oldyoung*). On the other hand, in the *child_care* equation we find that the difference in the effect of infants on time spent on child care is highest between girl-boy pairs where the girl is younger. Girls spend more time, relative to older boys, on child care in response to an additional infant. This is the flipside of what we see for the *chores* equation, where infants exacerbate the difference in schooling between girls and older boys (significantly negative coefficient on *infants*difsex*youngold* in the *chores* equation). The *chores* equation also shows that the impact of maternal employment on girls’ chores is higher than the impact on boys’ chores, *even when the girls are younger*.

In the *leisure* equation, we find that younger girls are cutting back their leisure (relative to boys) in response to maternal employment. This stands in contrast to our finding from the column for *schooling*, where older girls were cutting back their schooling time. However, it should be noted that the coefficient on *d_age* is larger in the leisure specification so that older girls’ leisure time is shorter than that of younger girls. Finally, we see that age has a positive impact on remunerative work, but that the impact of maternal labor or infants on remunerative work is not as clear cut as in the case of other activities.

6. Child Time Allocation and Credit Constraint: Estimation results of the endogenous switching model

Because of the completely-reduced-form or conditional approaches adopted in the previous sections, the effect of credit constraints on the level of children’s time use was not investigated explicitly so far. However, since the limited access to credit is one of the most important determinants of child labor and maternal employment, this section focuses on the effect of credit constraints on child time allocation (see Sawada et al. 2006 for the effect of credit constraints on maternal employment using the same dataset).

We estimate a system of three equations: equation (17) or (17') for children's time use and equations (18) and (19) for the incidence of binding credit constraints. To be consistent with the reduced-form approach, we use the same list of individual-level and household-level variables used in Sections 4 and 5. As X_{1h} and K_h in equation (19) (household-level variables that determine the household status with

respect to credit constraint), we include all household-level variables that were used in estimating the mother's labor supply equation (Table 6). As X_i and X_{1h} (individual- and household-level variables that determine children's time use), we include all explanatory variables that were used in estimating the children's time use equations (Table 5), except for *landval*, which was excluded to identify the effect of credit constraint. In other words, we treat *landval* as the variable that is included in K_h , not in X_{1h} .

Determinants of Credit Constraint: The first stage regression results for the determinants of credit constraint are shown in Table 8. Among the explanatory variables, the market value of currently owned land (*landval*) has a significantly negative coefficient and the number of household members (*hhsz*) has a significantly positive coefficient. Therefore, households with fewer land assets and more household members are more likely to be under a binding credit constraint. This finding confirms in the multivariate context the casual finding in Table 3. Among the education variables, the mother's education (*m_edud*) has a significantly negative coefficient when the wider definition of credit constraint is used. This suggests that higher human capital weakens credit constraints, although the evidence is not strong.

Household demographic compositions also have significant coefficients. Coefficients on the share of children and infants (*s_cm*, *s_cf*, *s_inf*) are positive, implying that younger households are more likely to be credit-constrained. This is as expected since adult household members who are already in the labor force can work more or borrow more easily when liquidity is required than infant members can. The coefficient on *bplhold* is significantly positive in the cc_2 regression, suggesting that the targeting in the PDS policy through distributing ration cards is toward households with inferior access to credit.

Looking at proxies for EEPs, the results based on the wider definition (cc_1) show that some coefficients on the characteristics of the child's grandparents are statistically significant and they are jointly significant at the 1 percent level. This is consistent with the superiority of the non-unitary, bargaining model over the unitary household model discussed in Subsection 2.2. However, when we use the narrower definition of credit constraint (cc_2), none of these variables are individually significant and they are jointly insignificant. To be on the safer side, models with EEPs are used as the first stage regression for equation (17) or (17').

Among the community fixed effects, *Muslim* and *ST* (scheduled tribes) dummies have positive coefficients, both of which are statistically significant under both definitions of credit constraints. This suggests that households belonging to these communities face systematically higher probabilities of credit constraints than other groups. This may reflect the social isolation of tribal and Muslim populations in rural regions with Hindu majority.

Intercept Effects of Credit Constraint on Child Time Allocation: The first part of Table 9 shows the intercept effects of credit constraint on child time allocation. Equation (17') is estimated together with equations (18) and (19) by a full information maximum likelihood method, so that the coefficients for equation (19) are slightly different from those reported in Table 8. However, they are very close. Moreover, when Equation (17') is estimated using Heckman=Lee's two-step estimation procedure, the results are very similar to those reported in the first part of Table 9.

The estimated model includes all explanatory variables that were used in estimating the children's time use equations (Table 5), except for *landval*. Since the

majority of them have similar coefficients in equation (17'), their coefficients are dropped for brevity. One interesting change is that the ill effects of *ST* and *Muslim* dummies on *dom_work* and *leisure* in Table 5 are attenuated when the endogenous credit constraint is included as a regressor. This implies that a substantial part of the reasons why *Muslim* and *ST* children spend more time on *dom_work* and less time on *leisure* can be attributed to differences in credit access, not to differences in preferences or in labor productivity.

Coefficients on the endogenous credit constraint are significant in five of eight specifications reported in the first part of Table 9. When they are significant, the null hypothesis that the two equations are independent is also rejected. Therefore, ignoring the endogeneity of credit constraint may lead to biased inferences.

The effect of credit constraint on *schooling* and *leisure* is negative with a very large coefficient (*schooling* or *leisure* time is reduced by about 3 days per week). Its statistical significance is also high. Therefore, children in credit-constrained households are substantially worse off, suffering from the current low level of leisure and the future low level of income due to low schooling. The effect of credit constraint on *rem_work* is insignificant while that on *dom_work* is positive and marginally significant (at the 20 percent level when cc_1 is used and at the 5 percent level when cc_2 is used). Combining with this finding with the finding by Sawada et al. (2006) that mothers in credit-constrained households work more outside the household, we can derive the second route in which credit constraint defers schooling. The first route is that human capital investment becomes less attractive when the household's subjective equilibrium interest rate is increased due to the binding credit constraint. The second route is that mothers in credit-constrained households work more outside, leaving more domestic work for children. Our analysis in Section 5 showed that the increase of children's working time in chores and child care due to maternal employment outside the household falls more on girls than on boys. Therefore, it is more likely that the ill effect of the second route is larger on girls than on boys.

Slope Effects of Credit Constraint on Child Time Allocation: Given the significance of credit constraints on child time allocation clarified through the estimation results of equation (17'), it is of great interest to examine whether the binding credit constraint also affects the marginal impact of other exogenous factors on child time allocation. We therefore estimated equation (17) using Heckman=Lee's two-step estimation procedure, adopting the specification proposed by Maddala (1983, pp.223-228).

First, we estimated a model in which all coefficients on the household and individual-level explanatory variables are allowed to vary between the two regimes. The null hypothesis that the slope difference is zero for all variables ($\alpha'_{1c} = \alpha'_{2c} = 0$ in equation (17) for all variables) was rejected for three dependent variables (*schooling*, *dom_work*, *rem_work*). Since this model has too many coefficients to be interpreted, we reduced the number of coefficients by deleting insignificant cross-term coefficients (i.e., some of α'_{1c} and α'_{2c} in equation (17) were set zero), and re-estimated the equation. Since we are interested in the gender contrast, the cross term of *sex*cc* was retained even when it was not significant in the least restricted specification.

The results thus obtained are reported in the second part of Table 9. In this table, all coefficients reported, except the one on *sex*cc₂* in the *leisure* regression are

statistically significant at the 20 percent level. Regarding the *schooling* regression, the additional age effect under the credit constraint almost offsets the concave age effect reported in Table 5. Under the credit constraint, sending children to school becomes more difficult so that the age profile of schooling time becomes very flat. The coefficient on *sex*cc* is positive but its statistical significance level is not high (20 percent when the wider definition is used and 10 percent when the narrower definition is used). The positive coefficient suggests that the sex difference in schooling is attenuated under credit constraint, because the ill effect of credit constraint is larger on boys than on girls who do not go to school even without the credit constraint.

Regarding *dom_work*, the coefficient on *sex*cc* is positive and significant, implying that the additional burden of domestic work under credit constraint (probably due to mothers working more outside due to credit constraint) falls dis-proportionally more on girls. Under the binding credit constraint, *hhsz* and *bplhold* becomes significant and a positive predictor of children's time spent on domestic works. In contrast, the coefficient on *sex*cc* on *rem_work* is significantly negative. However, this only shows that boys under credit constraint bear the burden of outside work more than girls do. The absolute effect of credit constraint on girls is not negative. Therefore, our results show that the additional work burden due to credit constraint has a very clear sex contrast: girls for more domestic work and boys for more outside work.

Finally, the credit constraint effect on the sex contrast in the *leisure* equation shows that the gender difference is attenuated under binding credit constraint, although its statistical significance level is only 20 percent. Its coefficient is large (-10) but evaluating at the average characteristics, the total effect of credit constraint on *leisure* is still negative on girls. The total effect on *leisure* is highly negative on boys, though.

7. Credit Constraint, Mothers' Work, and Children's Time Allocation: Conditional Labor Supply with Regime Switching

Results for the Girl-Boy Subsample: Finally, we examine the effects of credit constraint on the interaction between mother's and her children's time use. In particular, our focus is on the question of how the access to credit may alter the ways girls' and boys' time is used to substitute for mother's household chores (including childcare) when her labor supply (in remunerative work) changes. In order to address this issue, we re-estimate the conditional labor supply equations in difference form (as reported in Section 5), by splitting the sample households into credit constrained and unconstrained groups (i.e. estimating equations (14a), (14b) and (15) by two stage least squares).

Estimation results are reported in Table 10, for the case when the narrower definition of credit constraint (*cc*₂) is used.⁶ The upper panel reports the results based on the girl-boy sub-sample. Looking at the *schooling* equation, while an increase in

⁶ The results are qualitatively similar when the wider definition of credit constraint (*cc*₁) is used.

mothers' labor supply tends to decrease the schooling time of girls more than that of boys (as we saw in Section 5), the magnitude of such a gender disparity within credit constrained households (0.14 half-day) is about half the magnitude of gender disparity within unconstrained households (0.31 half day). Thus, the anti-girl bias in children's schooling time due to increased labor supply by their mothers tends to be attenuated by a credit constraint. This is consistent with the similar finding reported in the previous section based on the time use (level) analysis. Also in line with our earlier findings, the burden of child care falls disproportionately on girls as mothers' labor supply increases when the household is credit constrained. We also find that mother's labor supply significantly increases the gender disparity in child care within credit constrained households but not in unconstrained households. On the other hand, however, credit constraints appear to have little impact on the effect of mother's labor supply on the girl-boy disparities in terms of the time devoted to household chores (excluding childcare), remunerative work or leisure. Rather surprisingly, credit constraints appear to have no significant effect on the gender disparity in child labor (i.e., remunerative work), and, in addition, credit constraints appear to *lessen*, rather than exacerbate, the gender gap in schooling time.

Credit access has little effects on the time use differences in age gap for schooling, remunerative work or leisure, but we can see that the age effects on the gender disparity in household chores and childcare (as we saw in Section 5) are mainly driven by such effects observed within credit constrained households. Gender disparity in household chores is exacerbated as the age gap widens and the similar gender gap in childcare time is mitigated as the age gap widens *within credit constrained households but not in unconstrained households*.

Results for the Full Sample: Estimation results using the full sample are reported in the lower panel of Table 10. Consistent with the finding in the girl-boy sample, the gender disparity in the effect of mothers' labor supply on schooling is somewhat mitigated in credit constrained households than in unconstrained households (see the coefficients on *motwar*difsex*oldyoung* in the *schooling* equation). Similarly, while mothers' labor supply tends to increase the time spent for household chores more for girls than for boys, the magnitude of such gender disparities is smaller within credit constrained households than within unconstrained households among the pairs between elder girls and younger boys (see the coefficients on *motwar*difsex*youngold* in the *chores* equation). In addition, rather surprisingly again, mothers' labor supply significantly increases the gender gap in time spent for remunerative work (it increases girls' remunerable work more than it does boys', or it reduces boys' remunerative work more than it does girls') *only* within unconstrained households, but not within the credit constrained, among elder girl-younger boy pairs.

At the same time, however, among the pairs between younger girls and older boys, mothers' labor supply increases girls' time for household chores more than it increases boys' time for chores *only within credit constrained households* (while there is no significant effect among unconstrained households). Similarly among the pairs between younger girls and older boys, mothers' labor supply exacerbates the gender disparities in leisure (i.e., mothers' labor supply decreases girls' leisure more than boys'). As a result, it seems, younger girls within credit constrained households are arguably particularly disadvantaged, bearing disproportionate shares of time burden in the form of longer remunerative work time and shorter leisure time (though not in terms of shorter time in school) as their mothers increase their time in outside work. Apart from gender disparities, elder brothers tend to reduce schooling time more than

do their younger brothers in response to an increase in their mother's labor supply *only* within credit constrained households (see the coefficients on *motwar*boys*oldyoung* in the *schooling* equation).

As for the effects of additional infants on children's time use, there is some evidence that younger girls, relative to their older sisters, tend to bear more burden in terms of schooling time and household chores *within unconstrained households*, but not in the credit constrained (see the coefficients on *infant*girls*oldyoung* in the *schooling* and *chores* equations). However, older sisters tend to reduce their leisure time more than do younger sisters in both credit constrained and unconstrained households (see the coefficients on *infant*girls*oldyoung* in the *leisure* equation).

8. Conclusion

In this paper, we empirically investigated the determinants of child time allocation, including schooling time, domestic work, and remunerative work, in rural Andhra Pradesh, India. For this analysis, we collected detailed time-use data, information to judge whether the household is under a binding credit constraint, and information on extra-household linkages covering pre-marital and parental generation. The last information provides us exogenous sources of variation in bargaining powers within a household. We derive empirical models from a unified framework of collective/unitary household models and apply them to this dataset. There are three significant findings that emerged in our analysis.

First, the pre-marital and parental generation variables significantly affect the child's time use, which gives support to non-unitary household models. This finding was obtained both from a reduced-form and more structural approaches, suggesting that the unitary model of household is rejected and that collective models of household are supported.

Second, the results based on a conditional labor supply approach show that when a mother works outside, her domestic labor is more likely to be replaced by daughters than by sons, but the sex contrast in market work by children is not affected much, resulting in a higher work burden on daughters. The additional work burden is especially high on elder girls who tend to reduce their schooling time by the largest amount.

Third, in line with findings from the existing studies (e.g., Jacoby 1994, Edmonds 2005), the estimation results based on an endogenous switching model, based on the direct observation of the credit constrained households, show that access to credit is a major determinant of the child's time use: children of credit-constrained households spend less hours in school and for leisure but spend more hours working in the house. The child labor enhancing effect of credit constraint has a sharp gender component: on boys, the pressure is more on remunerative work outside the household; on girls, the pressure is more on domestic work such as household chores and child care. Rather surprisingly, however, we find some suggestive evidence that the gender disparity in schooling may be smaller in credit constrained households compared to unconstrained households. While girls' schooling time is shorter than boys' among unconstrained as well as credit constrained households, boys are likely to be pulled out of school disproportionately (probably because girls are already

(sufficiently) withdrawn from school) within credit constrained households, thereby reducing the gender gap (and aligning the schooling time to lower levels for both girls and boys). For a similar reason perhaps, increased labor supply by mothers tends to reduce the schooling time of boys more than that of girls within credit constrained households relative to such adjustments made within unconstrained households.

Combining these findings, we can derive two routes in which credit constraint defers schooling of children in rural India. The first route is that human capital investment becomes less attractive when the household's subjective interest rate is increased due to the binding credit constraint. The second route is that mothers in credit-constrained households work more outside, leaving more domestic work for children. Our analysis in this paper showed that the increase of children's working time due to maternal employment falls more on girls than on boys mainly through their increased domestic work. Therefore, it is more likely that the ill effect of the second route is larger on girls than on boys, the phenomenon not shown econometrically in the existing literature.

Addressing credit market failures and enhancing mothers' opportunity for (outside) work would require a careful design in policy interventions. For example, policies to increase opportunities for women's productive work (and thus to increase household income, as well as potentially increasing women's bargaining power within the household) could have significant implications for children's (especially girls') time burden in domestic work and schooling. Thus, such policies may need to be supplemented with additional measures to lessen mothers' domestic work (e.g., providing easy access to child care facilities). Furthermore, our findings imply a (potentially disturbing) possibility that an immediate/short-term effect of relieving credit constraints may be to increase the gender gap in schooling by encouraging boys to go to school but not girls to the same extent.

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