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Land Tenancy and Rural Factor Market Imperfections Revisited*

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The hypothesis that differential household endowments of nontradeable (or imperfectly tradeable) factors of production (such as male, female, and child labor, draft animals) are significant determinants of variations in the net amount of land leased-in is tested while controlling for the role of unobservable heterogeneity. Longitudinal household level data from a nine year panel of 290 households from ten villages in Semi-Arid Tropic rural India is utilized. The results obtained confirm the hypothesis and indicate a significant bias arising from the correlation of time invairant unobservables and nontradeable factor endowments.

I. Introduction

Recent studies analyzing the extent of leased land by rural households in developing countries have emphasized the role of nontradeability of certain key factors of production; technical know-how (Reid, 1976), managerial ability (Bell and Zusman, 1976), draft animal power (Bell, 1977, and Bliss and Stern, 1982), credit (Jaynes, 1982) and family labor (Pant, 1983, and Jodha, 1984). The underlying idea in all these models is that in a context of imperfectly functioning markets (for bullocks, labor, credit etc.) households enter the land tenancy market in an attempt to achieve the optimal operational size necessitated by their farm production technology and their factor endowments (see Binswanger and Rosenzweig, 1986). More recently, this underlying framework has been extended to anlayze the choice among contractual arrangements in the market for land tenancies (see Eswaran and Kotwal, 1985 and the survey by Otsuka and Hayami, 1988).

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The failure of certain factor markets in the rural sectors of developing countries are typically attributed to factors such as moral hazard, monitoring costs, economies of scale arising from factor indivisibilities and risk aversion (Newbery and Stiglitz, 1979, and Binswanger and Rosenzweig, 1986). For example, severe moral hazard is involved in the renting of draft animal services in is difficult to tell by inspection whether the animals were fed or treated properly by the renter. Similarly, the use of hired labor paid a fixed wage per hour involves considerable supervision costs; on the other hand the use of family labor involves relatively lower supervision costs since family members are residual claimants and their effort is directly linked to their consumption. Along the same line of reasoning, a competitive market for farm management skills with market determined wages does not provide sufficient incentives for the hired manager to use his skills optimally either, since there is no extra reward — besides the fixed wage payment — for increased effort; such an arrangement is dominated by one where the manager's reward is tied to profitability (e.g. self-cultivation). Related explanations for the absence of factor markets are based on the riskiness of agricultural production, the ensuing seasonality and synchronic timing of operations which lead to all agents being on the same side of the market (e.g. everybody wants to rent-in (out) or buy (sell) at the same time) that result in no trade.

Under these circumstances then, transactions in other markets are interpreted as Pareto improving attempts to circumvent the constraints imposed by the absence of markets for factors of production. For example, Binswanger and Rosenzweig (1986) attribute the prevalence of self-cultivating households under various land tenancy settings in the rural sector of developing countries to the informational and incentive advantages of the institution of the family; see also Pollack (1985). Similarly, the Eswaran and Kotwal's (1985) model, each contractual form (fixed rental, sharecropping or wage labor) is rationalized as an arrangement that entils a different type of agent (e.g. landlord or tenant) providing non-marketable factor inputs (management and supervisory skills). In this setting, the alternative opportunities available to each agent given their endowments determine the contract that will be chosen by both parties.

In spite the intense research activity at the theoretical level there is a dearth of empirical evidence supporting the arguments outlined above. One of the few studies is by Jodha (1984) who provided tabular evidence from six villages in Semi-Arid Tropicl India, that confirmed the proposition that the principal reason for leasing land in most of the villages was resource adjustment. In each one of the villages studied, tenancy tended to reduce the large gap between landowner and tenant in abailability of land per bullock. However, tenancy did not tend to equalize the land/firmly labor ratios. Using multivariate regression analysis, Bliss and Stern (1982) also found strong support for the role of nontradeable factor endowments in the net amount of land leased-in by households in one village in India. The same result was also obtained by Nabi (1985) who tested the Bliss and Stern model using data from four villages in Pakistan, and also Pant (1983).

A shortcoming of these cross-sectional studies, however, is the possible presence of biases in their estimates arising from correlation of the error term with the independent variables. As Bliss and Stern (1982) acknowledge, one source of such correlation may be due to specification error arising from the unobservability of certain key variables determining the incidence of tenancy. For example, suppose, in line with the arguments outlined in the preceding paragraphs, that farmers with higher (unobservable) managerial abilities would like to cultivate a larger amount of land than that they own. If there is a positive correlation between this unobservable component of the error term and the independent variables in a regression (e.g., farmers with higher managerial skills also have more factors of production such as larger families, more bullocks, farm implements etc.) than the ordinary least squares estimates of the included variables are going to be upward biased.

One possible solution to this problem is to include a measure of managerial ability in the analysis. This is a rather difficult task, however, given that the term managerial ability is typically used in place of a variety of attributes which are difficult or even impossible to quantify (e.g., drive, spunk etc.). The typical approach in most empirical studies is to use easily observable variables as proxies for managerial ability such as age and education. A more satisfying solution is offered when repeated observations (over time) for the same unit are available. Panel data permit the researcher to use the household as its own control in a simulated experiment. Assuming that the true stock of managerial ability of each farmer is unobsevable and time invarant, one can use the repeated observations on the same farmer to control for the role of managerial ability as well as for the role of all other time invariant unobservables. In this manner one is able to address the question at hand more rigorously.

The purpose of this paper is to subject a simple model of leasing behavior to scrutiny by explicitly testing for an controlling for the role of unobservable heterogeneity in managerial abilities. The study utilizes a unique longitudinal sample of 290 agricultural households from 10 villages in semi-arid tropical India. Detailed information concerning the value of assets owned by each household each year is combined with household demographic and socioeconomic information as well as measures of the risk attitude of the head of the household in order to explain the net amount of land leased-in each year. The empirical analysis finds strong evidence of the role of unobservables in the decision to lease land which are correlated with the observable factor endowments of each household. Estimation methods are used to control for these fixed effects. Once unobservable heterogeneity is controlled for, it is confirmed that differential household endowments of nontradeable (or imperfectly tradeable) factors of production play a significant role in explaining variations in the amount of land leased. In addition the adjustment of cultivated area towards desired cultivated area is incomplete.

The paper is structured as follows. Section II contains a brief description of the model, section III describes the data and contains the empirical analysis. The paper concludes with further suggestions for research.

II. Model

This section presents a version of the Bliss and Stern (1982) model by accounting for the role of household specific unobservables in the decision to lease land. Following their arguments, farmers[1] are assumed to have a desired demand for cultivated area which is derived based on their draft animals, family resources, managerial ability and attitudes toward risk at the beginning of the crop-cycle. Thus one can specify the following relation

$$(1) \qquad DCA_{it} = f(B_{it},\ F_{it},\ X_{it},\ Z_i,\ M_i)$$

where the subscript i indexes households, t indexes years, DCA_{it} stands for desired cultivated area of farmer i in year t, which is assumed to be determined by the stock of bullocks owned (B_{it}), the number of working age family members in the household (F_{it}), other variables (X_{it}) describing the asset position of the household such as the value of farm implements and the farming experience of the head proxied by age and other observable and time invariant characteristics of the head of the household (Z_i) such as caste, education and attitudes toward risk. The variable M_i is assumed to be time invariant and captures all the (farmer) household characteristics or endowments that are known to the household but are not observed by the econometrician such as managerial ability, etc.

1 The terms farmer, household and individual will be used interchangeably in this paper depending on the context. Typically, all decisions concerning farm production are made by the male head of the household.

Assuming a somewhat incomplete adjustment toward DCA one may then pose the relation

$$(2) \qquad NLI_{it} = h(DCA_{it} - OWN_{it})$$

where NLI denotes net land leased-in (either by sharecropping and/or fixed rental arrangements) and is a continuous variable taking positive (negative) values when land is leased-in (out). Thus the function $h(\cdot)$ relates the difference between desired cultivated area and owned area to the net amount of land leased-in (NLI). The latter variable is constructed by adding the total area of land leased-in through fixed rent and sharecropping arrangements and netting out the total area leased out (either through fixed rent or sharecropping) in each given year by each household. As Bliss and Stern point out, some reasonable restructions on the function $h(\cdot)$ may be: a) that it is non-decreasing in its arguments ($h'(x) \geqq 0$ for both positive and negative values of x), b) goes through the origin ($h(0) = 0$) and c) has a slope of less than or equal to one.[2] Combining (1) and (2) yields

$$(3) \qquad NLI_{it} = h(f(B_{it}, F_{it}, X_{it}, Z_{it}, M_i) - OWN_{it})$$

Taking a first oder Taylor series approximation of (1) and (3) (i.e. linearizing) yields the version of the equation to be estimated:

$$(4) \qquad NLI_{it} = \beta_0 + \beta_1 B_{it} + \beta_2 F_{it} + \beta_3 OWN_{it} + \beta_4 X_{it} + \beta_5 Z_i + \mu_i + \varepsilon_{it}$$

where β_0 is a constant, $\beta_1 = (h'f_B)$, $\beta_2 = (h'f_F)$, $\beta_3 = -h'$, $\beta_4 = (h'f_X)$, $\mu_i = (h'f_M) M_i$, $h' = dh/dDCA$, $f_K = \partial f/\partial k$ where $k = B, F, X, M$ and ε_{it} is an i.i.d. error term ($\varepsilon_{it} \sim N(O, \sigma_\varepsilon^2)$). By simple inspection of the estimated coefficients, their signs and significance one can easily test the validity of this simple model as well as the extent to which adjustment towards the desired cultivated area is complete (i.e. $\beta_3 = h' = -1$).

Inspection of expression (4) reveals that omission of the variable M_i from the regression and hence inclusion of the variable μ_i in the error term of a regression, is likely to result in biased estimates of the coefficients of

[2] An alternative specification may be the friction model of Rosett (1959) which permits differential adjustment when land is leased-in than leased-out. Such a specification, however, does not take advantage of the longitudinal nature of the data and does not control for unobservable heterogeneity (Skoufias, 1990).

the included variables in the regression. One way of reducing the omitted variables bias arising from the fact that μ_i is unobservable, is to include unobservable variables that are believed to be correlated with μ_i. For example, Bliss and Stern (1982) and Nabi (1985), acknowledge the potential role of ability in biasing their empirical results but cannot go further than including variables that are expected to be correlated with ability, such as education, age and caste. Similarly, Pant (1983) includes the value of farm implements as an explanatory variable as a proxy for the unobservable managerial ability of the farmer. However, none of these approaches is satisfactory. On the one hand, it is very likely that age, education and caste have distinct roles in the incidence of tenancy totally independent of the managerial ability of the farmer. In that case their inclusion in a regression would not control for the role managerial ability. On the other hand, these variables may also interact with the unobservable managerial ability in which case application of least squares would result in biased estimates.

The estimation method used in this paper treats the unobservable variable μ_i for each farmer as an unobservable individual (household) effect.[3] One underlying hypothesis to be tested is that this unobservable individual effect plays no significant role in explaining the variation of the dependent variable. Rejection of this null hypothesis would provide evidence in favor of the role of unobservable individual specific variables in the decision to lease land. However, rejection of the null hypothesis does not necessarily imply that the nontradeable stock of managerial ability determines the amount of land leased-in, since μ_i captures all time invariant characteristics of the farmer managerial ability being only one of them. To the extent, however, one is able to control for the role of other observable farmer, household and village specific time invarient variables, acceptance of the null hypothesis would provide some evidence against the proposition that the heterogeneity in nontradeable stocks of managerial abilities influences the net amount of land leased-in. For this purpose additional variables such as education, age, caste and an indicator of risk aversion of the head of the household are included.

In the econometric literature there are two approaches that differ in terms of the statistical properties attributed to unobservable individual effects. The *random effects* model treats unobservable individual effects as realizations of a random variable with a known distribution (e.g. $\mu_i \sim N(0, \sigma_\mu^2)$) that is uncorrelated with the included explanatory variables (e.g. B, F, OWN etc.). The *fixed effects* model treats the unobservable in-

[3] A sufficient condition that permits treatment of the μ_i term as time invariant is that both the f(·) and h(·) functions are linear in their arguments.

dividual effects as parameters that differ across individuals but does not impose any restrictions on their correlation with the included explanatory variables. In general the choice of treatment of unobservable individual effects is not very obvious and tests have been developed to test the specification appropriate for a given sample; Hausman (1978), Mundlak (1978), Hsiao (1986). A priori, one would tend to argue that farmers with higher managerial ability are likely to have a larger number of bullocks and larger amount of owned area, since the reward for higher managerial ability is higher profitability which may allow a faster accumulation of assets over time. Thus the fixed effects specification may appear at first sight to be more appropriate in this setting. However, rather than favoring one specification over the other, we explicitly test between the fixed and random effect specification using Hausman's specification test. For purposes of comparison with Bliss and Stern (1982) and Nabi (1985) the ordinary least squares estimates will be presented also.

III. Data

The empirical analysis in this paper is conducted using a panel of 290 farming households in the semi-arid tropic region of India, observed for four to nine agricultural years (starting July 1975 and ending June 1985). The panel was extracted from the Village Level Studies (VLS) files of the International Crops Research Institute for the Semi-Arid Tropics (ICRISAT). The VLS contain detailed information collected for approximately 400 households in 10 villages on a monthly basis over a period of four to nine years. A stratified random sample of a total of 40 households was constructed in ech of the ten villages to ensure representation of all categories of households — labor households (10) and cultivator households (30) (i.e. small farm, medium and large farm). The information collected includes details on household resource endowments (e.g. land, family labor, farm machinery and irrigation equipment), production inputs and outputs at the plot level as well as demographics and transactions undertaken by the household.[4] Six of these villages (Aurepalle, Dokur, Shirapur, Kalman, Kanzara and Kinkheda) were observed for 9 agricultural years (June 1975 — July 1984) whereas four additional villages (Boriya, Rampura, Papda and Rampura Kalan) were included in the survey starting in 1980 and ending in 1985.

The ten villages were purposefully selected by ICRISAT to represent

4 For a more detailed description of the methodology and the sampling procedure of the VLS see Singh et al. (1985).

broadly the agroclimatic zones of semi-arid tropical India. The villages of Aurepalle and Dokur (Mahbubnagar District of Andhra Pradesh) were selected to represent the areas with uncertain rainfall and red soils, which have low moisture capacity. Shirapur and Kalman (Sholapur District of Mahrarshtra) represent areas of uncertain rainfall with deep and medium — deep black soils which have a high moisture capacity. The villages of Kanzara and Kinkheda (Akola District of Maharashtra) were chosen as typical of the relatively high and more assured rainfall areas with black soil. All of the preceding villages cultivate similar crops such as sorghum, pigeonpea, cotton, paddy and castor. The villages of Boriya and Rampura (Sabarkantha District of Gujarat) were added in the survey in 1980 because of the types of crops cultivated there (pearl millet, groundnut and chickpea) their sandy soils and uncertain rainfall. Finally, the villages of Papda and Rampura Kalan (Raisen District of Madhya Pradesh) were also included starting in 1981 given the similarity of cultivated crops to the Gujarat villages, and their characteristic black soils and assured high rainfall.

In addition to the data outlined above, in six of these villages, estimates of preferences toward risk were obtained based on an experimental study designed and conducted by ICRISAT investigators (Binswanger, 1980). Given the unique opportunity provided by this data set in analyzing the correlation of attitudes toward risk and the extent of land tenancy and the unbalanced nature of the panel it was decided to conduct the empirical analysis by splitting the panel into two different subpanels. Thus a panel of 179 households observed for a period of 9 years (1611 observations in total) which contained estimates of the risk aversion of the head of the household was formed (Sample A). The second panel with more cross sectional coverage (Sample B) was constructed by utilizing information from all 10 villages for the 3 common agricultural years 1981-1983. This yielded a panel of 290 households (870 observations in total).

Table 1 contain the description, means and standard deviations of the variables in each of the two samples used in the analysis. Land tenancy through fixed rental or sharecropping arrangements is very common in each one of the villages. The bulk of the land tenancy transactions is for cultivation purposes although some land is leased for pasture in Shirapur and kalman. All villages are characterized by active labor markets for both male and female workers although participation of higher castes in labor market activities outside their own farm is not very frequent (Binswanger et al., 1984). Additional employment opportunities are also available in nearby government sponsored projects although the frequency of and job availability in such projects varies from village to village. For example in

Table 1

DESCRIPTION, MEAN AND STANDARD DEVIATION OF VARIABLES

Variable	Description	Sample A[a] (1611 obs.)	Sample B[b] (890 obs.)
Age:	Age in years of the head of the household	47.299[c] (12.27)[d]	46.987 (11.835)
Educated:	= 1 if head of the household had any education	.418 (.493)	.459 (.499)
High Caste:	= 1 if the household belongs in any one of the two castes ranked highest in the village.	.496 (.500)	.534 (.499)
Implements:	Value of farm implements owned by the household (1983 Rs.)x10^{-3}	2.166 (4.931)	2.314 (4.411)
Bullocks:	Number of bullocks owned by the household	1.327 (1.72)	1.453 (1.456)
Net land leased-in:	Net land leased-in by the household in hectares (ha) (1 hectare = 2.47 acres) (= area leased/sharecropped-in — area leased/sharecropped-out)	.120 (3.304)	.079 (2.15)
Other Assets:	Value of other assets owned by the household net of liabilities (= value of buildings + food stocks + financial assets + consumer durables-liabilities) (1983 Rs.)x10^{-3}	7.464 (12.77)	13.291 (20.09)
Owned Area:	Area (in ha) owned by the household at the beginning of the agricultural year	5.055 (5.79)	4.494 (5.026)
Number of Children:	Number of children in household less than 15 years of age.	2.431 (1.797)	2.378 (1.599)
Nmber of Adult Females:	Number of females in the household greater than 14 years of age	1.969 (1.955)	1.986 (.992)
Number of Adult Males:	Number of males in the household greater than 14 years of age	1.994 (1.132)	2.105 (1.099)
Risk Averse:	= 1 if head of the household is classified as risk averse (low, medium or extreme) (see Binswanger, 1980)	.719 (.499)	

Notes: a: Aurepalle, Dokur, Shirapur, Kalman, Kanzara and Kinkheda for 1975-1984

b: All ten villages for years 1981-1984

c: Mean

d: Standard deviation

the villages of Shirapur and Kanzara employment in government projects is easily available although in the villages of Kalman and Kinkheda such projects are not common. The hiring of draft animal services is limited in all ten villages. As farmers reported in all of the villages, serious constraints in the hiring of bullock services arise especialy during planting and harvesting due to the synchronic timing of operations. Another impediment to the hiring of bullock services also arises from the fact that typically the owners of bullocks belong to a higher caste than that of potential renters of bullock services. Under these circumstances the owners of the bullocks would refuse to hire out their services to lower caste farmers, since moral hazard arguments require that bullocks be accompanied by their owner as a driver. Similar arguments but to a lesser extent apply in the labor market. Workers from higher caste typically are not employed for wages by lower caste workers.

Table 2 contains the pooled ordinary least squares estimates of equation (4), ignoring problems arising from the presence of unobservable individual specific terms.[5] The least squares results are derived under the assumption that the "composite" error term, say $\eta_{it} = \mu_i + \varepsilon_{it}$, is uncorrelated with the included right hand side variables and also uncorrelated across time and farmers.[6] This method is comparable to that of Bliss and Stern (1982) and Nabi (1985). The coefficient of owned area which indicates whether or not adjustment towards the desired cultivated area is less than perfect, has a negative sign, as predicted by the simple model above and is less than one.[7] Thus adjustment toward the desired cultivated area appears to be seriously constrained. The number of bullocks owned by the household also appears to play a significant positive role. On the other hand, among the family composition variables only the number of adult males seems to play a role in the incidence of tenancy. The net area leased-in by households of higher castes is typically lower than that least-in by households of lower castes. Age (a proxy for farming experience) and education of the head of the household have no significant effect. Risk averse farmers, however, appear to lease in less land

[5] Identical regressions were run for each village separately without a significant change in the *qualitative* results. An F-test of the hypothesis that coefficients are equal across villages rejected the null, suggesting substantial differences across villages. For the sake of brevity, however, it was decided to report the results obtained from pooling the data across all villages and including village dummies. Village by village, least squares, fixed effects and random estimates are available upon request from the author.

[6] So η_{it} is an error term with $E(\eta_{it}) = 0$, $Var(\eta_{it}) = \sigma^2_\eta$, $\operatorname{plim}_{N\to\infty} (1/N)(X_{it}\eta_{it}) = O$, $Cov(\eta_{it}\ \eta_{js}) = 0 \quad i \neq j$ and $t \neq s$.

[7] A two-tailed t-test of the null hypothesis that the coefficient of owned area is equal to 1 was rejected in the conventional significance levels. For a study considering the role of rationing in the land tenancy market see Bell and Sussangkarn (1988).

Table 2

NET LAND LEASED-IN
ORDINARY LEAST SQUARES ESTIMATES

Variable	(1)[a]	(2)[a]	(3)[b]	(4)[b]
Constant	−.260	−.267	−.309	−.288
	(.598)[c]	(.574)	(.894)	(.760)
Age	−.005	−.001	−.007	−.004
	(.737)	(.168)	(1.353)	(.764)
Number of Adult Males	.330	.316	.267	.258
	(4.122)	(3.943)	(4.082)	(3.949)
Number of Adult Females	−.128	−.096	−.121	−.096
	(1.585)	(1.173)	(1.845)	(1.468)
Number of Children	.052	.036	.044	.033
	(1.223)	(.848)	(1.247)	(.942)
Owned Area (ha)	−.430	−.435	−.419	−.423
	(21.086)	(21.312)	(23.703)	(23.915)
Number of Bullocks	1.154	1.141	1.046	1.039
	(18.883)	(18.683)	(20.405)	(20.272)
Value of Implements x 10^{-3}	−.008	−.004	.016	.019
	(.460)	(.216)	(1.179)	(1.355)
Value of Other Assets x 10^{-3}	.011	.014	.018	.019
	(1.560)	(2.065)	(3.998)	(4.343)
Fallow Area (ha)	.268	.278	.252	.262
	(5.361)	(5.570)	(5.689)	(5.901)
High Caste	−.661	−.682	−.481	−.494
	(3.816)	(3.948)	(3.490)	(3.593)
Educated	.174	.175	.049	.048
	(1.015)	(1.027)	(.358)	(.352)
Risk Averse	−.365	−.377		
	(2.077)	(2.148)		
Village Dummies	YES	YES	YES	YES
Year Dummies	NO	YES	NO	YES
Adjusted R^2:	.294	.298		
Breusch-Pagan $\chi^2(1)$:	712.95	740.00	1,015.71	1,049.57

Notes: a: Sample A, see text and Table 1 for details.
b: Sample B, see text and Table 1 for details.
c: Absolute t-values in parentheses.

than risk neutral farmers. Finally, as expected, farmers with larger parts of their owned land left fallow make-up for the land shortage by leasing in more land.[8]

As it was pointed out in the preceding paragraphs, the presence of the individual specific effects (μ_i) in the error term of the pooled regression model may, in the very least, introduce a correlation among the residuals of the same farm household, even though the residuals from different households are independent. In an effort to test for the presence of variance components, the Breusch-Pagan (1980) Lagrange Multiplier test was performed based on the OLS residuals derived from equation (4). The null hypothesis tested is that the variance of the individual time invariant components is zero ($\sigma_\mu^2 = 0$). The results of the test are reported in the last row of Table 2. The χ^2 statistics obtained reveal the presence of random components which indicates that signifcance tests based on estimates of the variances derived from ordinary least squares are inappropriate. The proper estimation method in this case is generalized least squares (GLS).

Given the evidence indicating the presence of random components estimation of equation (4) was carried out assuming the random effects specification. This amounts to decomposing the (composite) error term μ_{it} into two independent random components — one associated with the farmer (μ_i) and one purly random component varying across time and farmers (ε_{it}). More specifically, it is assumed that:

$$\begin{aligned} &E(\mu_i) = E(\varepsilon_{it}) = 0 \\ &\mathrm{Var}\ (\varepsilon_{it}) = \sigma_\varepsilon^2 \\ &\mathrm{Var}\ (\mu_i) = \sigma_\mu^2 \\ &\mathrm{Cov}(\varepsilon_{it}\mu_j) = 0 \text{ for all } i,\ t, \text{ and } j \\ &\mathrm{Cov}(\varepsilon_{it}\varepsilon_{js}) = 0 \text{ if } t \neq s \text{ or } i \neq j \\ &\mathrm{Cov}(\mu_i\mu_j) = 0 \text{ if } i \neq j \end{aligned} \tag{5}$$

Estimation of this model can be carried out by using the weighting scheme, suggested by Hausman and Taylor (1981), that transforms the covariance matrix to one that is proportional to the identity matrix. Thus, estimation of the random effects model is carried out by applying ordinary least squares to the following transformed model:

[8] There is a possible source of simultaneity bias arising from the possibly joint decision of the extent of land leasing and fallowing in any given year. An exogeneity test concerning the variable fallow using the amount of owned area cultivated in the last two years as instrumental variables did not reject the null hypothesis.

$$\text{(6)}\qquad NLI_{it}-(1-\theta)NLI_{i.} = \beta_0\theta + (B_{it}-(1-\theta)B_{i.})\beta_1 + (F_{it}-(1-\theta)F_{i.})\beta_2 + (OWN_{it}-(1-\theta)OWN_{i.})\beta_3 + (X_{it}-(1-\theta)X_{i.})\beta_4 + \theta Z_i\beta_5 + \theta\mu_i + (\varepsilon_{it}-(1-\theta)\,\varepsilon_{i.})$$

where $\theta = (\sigma_\varepsilon^2/(\sigma_\varepsilon^2 + T\sigma_\mu^2))^{1/2}$ and $Y_{i.}$ denotes the mean over time of the observations on variable Y for farmer i.[9]

The estimates obtained under the random effects specification are presented in Table 3. As the last row of Table 3 reveals, the presence of farmer specific random components in the error term results in a correlation among the residuals of the same household that ranges from .365 to .444. The GLS estimates in Table 3 are not very different from the least squares estimates although they are more efficient. However, when individual specific random components are accounted for, the sign of the coefficient of the number of adult females in the household changes from negative to positive, although it remains insignificant.

The next step in the empirical analysis is to address the question of whether a fixed effects or random effects formulation is more appropriate for equation (4). For this purpose the model is re-estimated under the fixed effect specification. This amounts to applying ordinary least squares to expression (6) after setting $\theta = 0$;[10] which transforms each variable as a deviation from the farmer mean and thus eliminates all farmer specific and time invariant observable and unobservable variables from the regression (e.g., education, risk preferences, caste and μ_i). By performing this transformation, one can be certain that the estimated coefficients of the variables that very across time and households will be unbiased and consistent, irrespective of whether μ_i is corrected or not with the right hand side variables (Mundlak, 1978).

The fixed effects estimates are presented in Table 4. The explanatory power of the model increase considerably under all specifications. An F-test of the null hypothesis that all farmer specific intercepts are equal rejected the null in all cases. The coefficient of owned area remains negative

9 A consistent estimate of σ_ε^2 can be obtained from the residuals of the within (fixed effects) regression while a consistent estimate of σ_μ^2 is obtained from the residuals of the between regression (i.e. a regression containing the mean values of all the variables in the model for each farmer). T denotes the length (number of years) of the panel; (see Hausman and Taylor, 1981).

10 An equivalent method is to apply OLS on equation (4) including n-1 dummy variables for the n farmers in the sample.

Table 3

NET LAND LEASED-IN
RANDOM EFFECTS ESTIMATES

Variable	(1)[a]	(2)[a]	(3)[b]	(4)[b]
Constant	.328	−.490	−.086	−.045
	(.422)[c]	(.608)	(.105)	(.055)
Age	−.024	−.004	−.013	−.013
	(2.148)	(.330)	(1.139)	(1.156)
Number of Adult Males	.334	.317	.317	.315
	(3.317)	(3.065)	(2.832)	(2.809)
Number of Adult Females	.061	.113	−.024	−.025
	(.659)	(1.212)	(.216)	(.230)
Number of Children	.072	.041	−.078	−.076
	(1.364)	(.781)	(1.246)	(1.219)
Owned Area (ha)	−.388	−.397	−.555	−.555
	(13.647)	(13.966)	(15.792)	(15.792)
Number of Bullocks	.896	.885	.926	.929
	(13.025)	(12.920)	(11.036)	(11.043)
Value of Implements x 10^{-3}	.000	.007	.066	.066
	(.006)	(.408)	(3.099)	(3.105)
Value of Other Assets x 10^{-3}	.007	.012	.022	.022
	(1.029)	(1.642)	(3.352)	(3.352)
Fallow Area (ha)	.386	.399	.568	.567
	(7.063)	(7.322)	(7.447)	(7.419)
High Caste	−.732	−.782	−.456	−.459
	(2.155)	(2.311)	(1.558)	(1.568)
Educated	.294	.345	−.087	−.090
	(.868)	(1.020)	(.301)	(.312)
Risk Averse	−.339	−.374		
	(.978)	(1.080)		
Village Dummies	YES	YES	YES	YES
Year Dummies	NO	YES	NO	YES
Adjusted R^2:	.158	.168	.288	.288
$\sigma_\mu^2/(\sigma_\mu^2+\sigma_\varepsilon^2)$:	.365	.366	.444	.443

Notes: a: Sample A, see text and Table 1 for details.
b: Sample B, see text and Table 1 for details.
c: Absolute t-values in parentheses.

Table 4

NET LAND LEASED-IN
FIXED EFFECTS ESTIMATES

Variable	(1)[a]	(2)[a]	(3)[b]	(4)[b]
Age	−.087	−.044	−.059	−.013
	(4.230)[c]	(1.088)	(.469)	(.204)
Number of Adult Males	.316	.303	.02	.017
	(2.509)	(2.400)	(.104)	(.091)
Number of Adult Females	.204	.219	.439	.447
	(1.942)	(2.078)	(2.346)	(2.390)
Number of Children	.091	.070	.111	.109
	(1.467)	(1.132)	(.953)	(.944)
Owned Area (ha)	−.295	−.298	−.333	−.337
	(6.175)	(6.247)	(2.410)	(2.450)
Number of Bullocks	.725	.722	.184	.178
	(9.329)	(9.310)	(1.789)	(1.742)
Value of Implements x 10^{-3}	−.003	.002	−.019	−.019
	(.133)	(.079)	(.404)	(.407)
Value of Other Assets x 10^{-3}	.008	.007	.008	.008
	(.976)	(.945)	(.783)	(.809)
Fallow Area (ha)	.472	.479	.893	.891
	(7.878)	(7.971)	(8.192)	(8.217)
Village Dummies	NO	NO	NO	NO
Year Dummies	NO	YES	NO	YES
Adjusted R^2:	.604	.608	.839	.840
H_0: Farmer constants equal (F-test):	7.283	7.286	6.744	6.726
Hausman's $\chi^2(9)$:	48.66	34.62	281.88	273.84

Notes: a: Sample A, see text and Table 1 for details.
b: Sample B, see text and Table 1 for details.
c: Absolute t-values in parentheses.

and is generally smaller in magnitude compared to the least squares estimates. The underlying reason for the decreased coefficent of owned land under fixed effects rests on the fact that the fixed effects estimation method utilizes variation within households only while ignoring any variation between households. For the reasons detailed in Rosenzweig and

Wolpin (1985), the sales market for land in rural areas is very thin because the children of the farm head have the specific experience in cultivating the family land that makes them the highest "bidders" for that land. Under these circumstances, the changes in land owned by any given household (through market sales or purchases) would be very small.

The main effect of the use of the fixed effects estimation method is on the significance of the family composition variables. The number of adult makes workers remain significant in specifications (1) and (2) whereas the number of adult female workers becomes significantly positive in all specifications. This result is consistent with the notion that family male and female labor resources are nonmarketed family resources that play a significant role on the extent of land leased-in. Similarly, the coefficient of the number of bullocks owned by the household is also significantly positive, albeit less than that in Table 2. Households with more bullocks, lease in more land.

Since application of the GLS estimator only leads to gains in efficiency and GLS yields biased and inconsistent estimates in case the individual unobservable components are correlated with the included right hand side variables, a Hausman (1978) specification test is also conducted. The last row of Table 4 contains the results of the hypotheses tests concerning possible correlation of the individual unobservable components with the right hand side variables. Hausman's chi-squared test statistics suggest that the fixed effects specification is more appropriate. As a second check, the version of the specification test suggested by Mundlak (1978) and discussed by Hsiao (1986) and Hausman (1978), was also constructed and identical results were obtained.

IV. Concluding Remarks

This paper provided empirical evidence in favor of the hypothesis that the land tenancy can be rationalized as a response to imperfections in markets for family labor, bullock services and managerial abilities. The leasing of land serves as a means of bringing land to the nontradeable factors of production (managerial ability, family labor, and bullocks). Due to the presence of monitoring costs for hired labor, indivisibilities and moral hazard in bullock labor services and difficulties in the market for managerial skills, one observes an active land lease market for the factor (i.e. land) that is subject to these problems to a much lesser degree.

In short, the evidence emerging from this study can be summarized by two points. First, unobservable heterogeneity in farmer endowments such as managerial ability, drive, etc., has a significant role in the decision to lease land. Second, ordinary least squares estimates should be handled with care. The extent of the bias caused by unobservable heterogeneity will vary from sample to sample. In this study correcting for heterogeneity had only a marginal effect on the size of the majority of coefficients-except for adult female labor were the coefficient turned from negative to significantly positive. Nevertheless, it was only after applying the fixed effect estimation method that a key element of the proposed hypothesis was confirmed.

Furthermore, the results concerning the role of family labor resources in the decision to lease land (i.e. self-cultivate) have important implications for the specification of agricultural labor services in the recent empirical literature of the farm household (see Singh, Squire and Strauss, 1986). A common practice in this line of research is the assumption of perfect substitutability between male and female labor services as well as family and hired labor services in each sex category. If family labor of each sex type were perfectly substitutable with hired labor then one should anticipate an insignificant role of family resources in the decision to lease land. The results of this study suggest that the assumption of perfect substitutability may be inappropriate. Further research concerning this issue is called for.

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