NEOCLASSICAL THEORY AND THE OPTIMIZING PEASANT: AN ECONOMETRIC ANALYSIS OF MARKET FAMILY LABOR SUPPLY IN A DEVELOPING COUNTRY*

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Few attempts have been made to test empirically the multitude of models formulated to describe household labor supply behavior in the context of rural labor markets in developing countries. In this paper refutable predictions are derived from a neoclassical multi-person household model based on competitive assumptions modified to take into account differences in landholding status. A national sample survey of rural households from India is used to estimate the parameters of the model for male and female agricultural workers from farm and nonfarm households. The estimates generally conform to the implications of the neoclassical-competitive framework.

I. INTRODUCTION

A considerable body of literature concerned with the process of economic development has characterized rural labor markets in developing countries as uncompetitive—rural wages are presumed to be institutionally set at levels above the "market" equilibrium and significant under- and unemployment of labor is assumed to exist (see, for example, Lewis [1954], Ranis and Fei [1961], Reynolds [1965], and Sen [1966]. These characterizations, however, have rarely been subjected to rigorous empirical examination, nor has the noncompetitive distribution of market (paid) employment among rural households been well specified. Among studies using rural labor market data, Rodgers [1975], ignoring the identification problem, concludes that the competitive model is inapplicable, based on a gross negative correlation between wage rates and aggregate employment across seven Indian villages. In a more richly detailed study, however, Hansen [1969] presents descriptive evidence that household members in rural Egypt are employed for a considerable number of days during the year and other data that would appear consistent with a com-

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petitive framework. Hansen also finds a strong positive correlation between rural wages and hours worked per day during the year for males, females, and children. Given that the seasonal pattern of wages is fully anticipated by workers, this result can be interpreted as evidence of the positive compensated substitution effect implied in neoclassical labor supply models (see Ashenfelter and Heckman [1974]). Hansen does not, however, attempt to explain the cross-sectional variation in annual employment among families.

In this paper a neoclassical framework based on competitive assumptions is utilized to describe market (for pay) labor supply behavior in two-person households in developing countries and is tested in micro data from India. While the implicit assumption underlying most of the development literature is that this framework is inappropriate in such a context, many characteristics of rural areas of developing nations may make the application of the neoclassical labor supply model more appealing than in developed country labor markets—labor is less heterogeneous (but wage rates within narrowly defined occupations vary greatly because of geographical immobility), nonpecuniary differences in wage-jobs are likely to be fewer, taxation of savings may be ignored,[1] and time worked may be more flexible. Unfortunately, the standard neoclassical family labor supply model, designed to explain behavior in developed country labor markets, as presented in Kosters [1966], Ashenfelter and Heckman [1974], and Kniesner [1976] provides few predictions that are testable without high quality data on non-earnings income, which are particularly difficult to obtain in developing countries.[2] Moreover, the set of variables implicated by the neoclassical model as determinants of labor supply do not appear to differ from that generated by the "labor surplus" approach so that distinctions cannot be easily drawn between the two frameworks based on which labor market variables are empirically important.

It is shown here, however, that the extension of the theory to households owning land, who make up a major portion of rural households in India, and the comparison of landless and landholding household market supply relationships yields an array of refutable predictions that do not require the estimation of compensated effects and that do not appear to be readily derived from the surplus labor hypothesis. For instance, it is demonstrated that the gross own wage

1. Problems involved in taking account of the income tax in U. S. labor supply studies are discussed in Rosen [1976] and Wales [1973].
2. Such data are required to obtain accurate estimates of "pure" income effects on labor supply in order to test for the income-compensated wage effects implied by the neoclassical model.

effect on labor supplied to the market should be algebraically less in landless than in landowning households and that if schooling augments the allocative ability or technical efficiency of farm managers (or their wives), the labor supply-education relationships should be more negative in landholding households. Thus, as a byproduct of the theoretical analysis, a framework is established for testing for the marginal efficiency role of schooling in agriculture based on labor supply behavior.

A limitation of the analysis is that it is both a test of the competitive framework—in which an individual's employment within a labor market, given the market wage, is determined only by supply behavior—and the neoclassical model. Thus, it is possible that the predictions derived from the theory may be contradicted empirically not because rural labor markets are noncompetitive but because the neoclassical model of "peasant" behavior specified is wrong or incomplete. Alternatively, of course, peasants may be "neoclassical" but institutional restrictions on employment not taken into account in the analysis may foil attempts to test for such behavior. The empirical results obtained, while in some cases open to alternative interpretations, are, however, supportive of the behavioral implications of the neoclassical-competitive model and appear to reject a number of alternative labor surplus hypotheses.

In Section II the model of landless household labor supply in which the husband and wife are earners is briefly reviewed. A corresponding model for landholding households is formulated and the relevant comparative statics are derived and compared to those of the landless model. Data from a rural household survey from India are then used to test the set of predictions pertaining to the market labor supply of males and females in landless and landholding households derived from the models in Section III. Section IV contains a brief summary and conclusion.

II. THEORETICAL ANALYSIS

Landless and Landholding Households

The model of the landless household corresponds to the standard model applied to developed country data, as in Kosters [1966], Ashenfelter and Heckman [1974], and Kniesner [1976], and will be briefly set out here.

The household is assumed to act as if it maximized a monotonic twice-continuously differentiable, strictly concave household utility function, as in (1):

(1) $U^N = g\ (X^N, M^N, F^N; E_M^N, E_F^N)$,

where U^N is the utility of the household without land, X^N is the amount of the homogeneous market good consumed, and M^N, F^N represent the nonmarket time of each household member (husband and wife). E_M^N and E_F^N are the schooling levels of the husband and wife, which are assumed to influence the demand for nonmarket time.

The full-income constraint for the landless household is given by (2):

(2) $\Omega(W_F + W_M) + I^N = W_M M^N + W_F F^N + X^N$,

where Ω is the total time available to each family member, W_M and W_F are the market wage rates of male and female laborers, I^N is asset income, and X is the numeraire. Implicit in (2) is the assumption that each family member can work for any amount of time without affecting his (her) wage;[3] thus family employment, occurring only in the market, is determined solely by supply factors. It is assumed that the husband and wife spend some time in the market[4] so that $\lambda_M^N = \Omega - M^N$, $\lambda_F^N = \Omega - F^N$, and (2) can be rewritten in terms of market time as

(3) $\lambda_M^N\ W_M + \lambda_F^N\ W_F + I^N - X^N = 0.$

The appropriate Lagrangean equation is thus

(4) $V^N = g(X^N, M^N, F^N; E_M^N, E_F^N)$
                 $+ \mu^N[\lambda_M^N W_M + \lambda_F^N W_F + I^N - X^N],$

where μ^N is the Lagrangean multiplier. If only interior solutions are considered, first-order conditions for a utility maximum are

(5) $g_X - \mu^N = 0$

(6) $g_M - \mu^N W_M = 0$

(7) $g_F - \mu^N W_F = 0$

(8) $\lambda_M^N W_M + \lambda_F^N W_F + I^N - X^N = 0.$

As is well-known, without data of sufficient quality to allow relatively precise estimates of "pure" income effects (and thus of compensated substitution effects), neoclassical labor supply theory for-

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3. This assumption is generally employed in U. S. labor studies; see Kniesner [1976], Kosters [1966], and Schultz [1975], but is modified in Rosen [1976]. Indirect empirical evidence of the independence of the wage rate and labor supply in rural India is presented in Section III.

4. None of the implications of the models are altered when women are not earners. Additional predictions (which are similar to those developed by Kniesner, [1976]) can be derived when this assumption is relaxed, but because they are of little empirical importance, they are not presented here.

mulated in terms of wage earners provides no testable "predictions" and thus cannot be readily used as a framework against which to contrast empirically alternative theories of wage-employment relationships. Not all participants in rural labor markets are members of landless households, however; thus the standard (landless) model must be modified to take into account family labor activities.

Landholding households are distinguished from landless households, for the purposes here, by the feature that in the former at least one household member combines part of his (her) time with other productive assets (chiefly land) owned by the household for the purpose of generating (farm) income from the production of the homogeneous (numeraire) commodity. For simplicity, it is assumed that both family members spend time in farm production. Households owning land or other productive assets are assumed to maximize a utility function identical to that of landless households:

(9) $$U^L = g(X^L, M^L, F^L; E_M^L, E_F^L).$$

The schooling levels of the husband and wife in landholding households are also assumed to affect the demand for household time in the same way as in landless households.

The production of farm output Q, derived from the production inputs (including labor) of the landholding family, is described by a twice-differentiable, strictly concave production function (10):

(10) $$Q = \Gamma(m, f, \kappa; e),$$

where m and f are the quantities of male and female labor used in farm production, κ is a vector of the prices and quantities of other farm inputs, including land, irrigation facilities, weather, etc., which are assumed to be exogenous.[5] For simplicity, family and hired labor of each type (sex) are assumed to be perfect substitutes,[6] but male and female labor are imperfectly substitutable. At least part of both m and f thus represent family labor.

e, a conditioning variable that represents the stock of managerial ability of the household, such that $\delta \Gamma_m/\delta e$, $\delta \Gamma_f/\delta e$, $\delta \Gamma_\kappa/\delta e > 0$, is hypothesized to be a function of both general and specific human capital—the schooling of the two family members and their work experience on their own farm; i.e.,

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5. It is assumed, as in almost all studies of India, that the land market is imperfect; that is, land is not readily bought or sold, and access to leased land is restricted. Bell and Zusman [1976] cite evidence that almost no households not *owning* any land are tenants in India and provide data that suggest that landholding status is exogenous.

6. Bardhan [1973] could not reject the null hypothesis that family and hired labor were perfect substitutes in agricultural production in five of the seven Indian farm surveys he analyzed. No attempt was made to distinguish between male and female (and child) labor, however.

(11) $$e = \Psi \ (E_M^L, E_F^L, A_M^L, A_F^L),$$

where

$$\Psi_1, \Psi_2, \Psi_3, \Psi_4 > 0.$$

It is further assumed that the level of specific experience amassed in off-farm jobs is minimal such that managerial proficiency cannot be hired out.[7] It is also assumed that there are no direct, i.e., worker effects, of schooling—schooling and work experience do not directly augment the productivity of workers in such farm tasks as weeding, plowing, reaping, etc.

The budget constraint for landholding households can be written as

(12) $\quad \Omega(W_M + W_F) + \Gamma(m,f,\kappa;e) - mW_M - fW_F + I^L$
$$= X^L + M^L W_M + F^L W_F,$$

or noting that $\lambda_M^L = \Omega - M - m$ and $\lambda_F^L = \Omega - F - f$:

(13) $$\Gamma(m,f,\kappa;e) + \lambda_M^L W_M + \lambda_F^L W_F^L + I^L - X^L = 0.$$

λ_M^L and λ_F^L represent *net* labor supply and need not be positive; on farms with productive capacity (κ) above some point, family labor will not be sufficient for profit (utility) maximization and the family will hire labor so that $\lambda_M^L, \lambda_F^L < 0$. W_M and W_F are thus the wages paid to hired workers by the landholding households and the wage rates received by family members if they work off the farm ($\lambda_M^L, \lambda_F^L > 0$). Consistent with the competitive assumption, there are no constraints on the quantities of labor hired or on market labor supplied.

The Lagrangean equation for the landholding household is thus

(14) $\quad V^L = g(X^L, M^L, F^L; E_M^L, E_F^L)$
$$+ \mu^L[\Gamma(m,f,\kappa;e) + \lambda_M^L W_M + \lambda_F^L W_F + I^L - X^L].$$

Assuming interior solutions for all control variables, first-order conditions are

(15) $$g_x - \mu^L = 0$$

(16) $$g_M - \mu^L W_M = 0$$

(17) $$g_F - \mu^L W_F = 0$$

(18) $$\Gamma_m - W_M = 0$$

(19) $$\Gamma_F - W_F = 0$$

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7. The nontradability of managerial skill is emphasized in Bell and Zusman [1976] as an important factor in determining the demand for leased land.

(20) $\Gamma(m,f,\kappa;e) + \lambda_M^L W_M + \lambda_F^L W_F + I^L - X^L = 0.$

The first three conditions are identical to those pertaining to landless households; the marginal value of each household member's time equals the relevant wage rate irrespective of whether work is performed *off* the farm; the standard "landless" model is nested in the landholding model. Conditions (18) and (19) are the profit-maximizing conditions for variable input use, implying that the level of farm profits is independent of or exogenous to the household's consumption preferences and levels of non-earnings income, since the quantities of m and f used will always be those corresponding to profit maximization. The left-hand side of (12) thus represents maximum potential income and corresponds to the concept of full income in the standard (landless) model. Given this independence between consumption and production, it is possible to compare the behavior of landless and landholding families in identical consumption equilibria, since if we can assume that all households in the same labor market face the same wage rates and prices, i.e., markets are competitive, we can set $[\Gamma(m,f,\kappa;e) - mW_M - fW_F]_{max} + I^L = I^N$. This latter assumption is tested in the next section.

The set of differential equations obtained by totally differentiating equations (15) through (20), which can be used to solve for the response of sex-specific net labor supply to changes in wage rates and other exogenous variables in landholding households, is given by (21):

(21)

$$
\begin{bmatrix}
g_{XX} & g_{XM} & g_{XF} & 0 & 0 & -1 \\
g_{MX} & g_{MM} & g_{MF} & 0 & 0 & -W_M \\
g_{FX} & g_{FM} & g_{FF} & 0 & 0 & -W_F \\
0 & 0 & 0 & \Gamma_{mm} & \Gamma_{mf} & 0 \\
0 & 0 & 0 & \Gamma_{fm} & \Gamma_{ff} & 0 \\
-1 & -W_M & -W_F & 0 & 0 & 0
\end{bmatrix}
$$

$$
\times
\begin{bmatrix}
dX^L \\
dM^L \\
dF^L \\
dm \\
df \\
d\mu^L
\end{bmatrix}
=
\begin{bmatrix}
0 \\
\mu^L \, dW_M \\
\mu^L \, dW_F \\
dW_M \qquad \Gamma_{m\kappa} \, d_\kappa \\
dW_F \qquad \Gamma_{f\kappa} \, d_\kappa \\
(-\lambda_M^L \, dW_M - \lambda_F^L \, dW_F - \Gamma_\kappa \, d\kappa - dI^L)
\end{bmatrix}
$$

The partial derivatives of male and female market labor supply with respect to the wage rates, obtained by solving the relevant equations in (21), can be written as (22) and (23):

$$(22) \quad \frac{\delta \lambda_M^L}{\delta W_K} = - \frac{\phi_{i2}^L}{\phi^L} + \lambda_K^L \frac{\phi_{62}}{\phi^L} - \frac{\phi_{j4}}{\phi^L} \qquad \begin{matrix} K = M, i = 2, j = 4 \\ K = F, i = 3, j = 5 \end{matrix}$$

$$(23) \quad \frac{\delta \lambda_F^L}{\delta W_K} = - \frac{\phi_{i3}^L}{\phi^L} + \lambda_K^L \frac{\phi_{63}}{\phi^L} - \frac{\phi_{j5}}{\phi^L},$$

where ϕ^L is the bordered Hessian determinant in (21) and ϕ_{rc}^L the cofactor of row r and column c in ϕ^L. However, it can be easily shown that

$$(24) \quad \frac{\delta \lambda_M^L}{\delta W_K} = - \left(\frac{\delta M}{\delta W_K} \right)_U - \lambda_K^L \left(\frac{\delta M}{\delta I} \right) - \frac{\Gamma_{mx}}{\Delta}$$

$$(25) \quad \frac{\delta \lambda_F^L}{\delta W_K} = - \left(\frac{\delta F}{\delta W_K} \right)_U - \lambda_K^L \left(\frac{\delta F}{\delta I} \right) - \frac{\Gamma_{fx}}{\Delta},$$

where $\Delta = \Gamma_{mm} \Gamma_{ff} - (\Gamma_{mf})_>^2 \, 0$ and $x = m, K = M; x = f, K = F$.

The first two terms in (24) and (25) are identical to those of the standard landless-household Slutsky equations, such as derived in Kneiser, except that the income effect is weighted by net labor supply λ_K^L, the difference between total family labor supply of member $k(\Omega - M, \Omega - F)$ and labor of type k used in farm production. The third term is the response of labor use to a change in the wage, which must be negative in the own case and positive otherwise if male and female labor in farm production are competitive inputs (see Allen [1964]). Because λ_K^L will be positive for households supplying labor to the market, the gross wage-net supply relationships are thus ambiguous for landholding households, as in the landless model. However, the sign of the *differential* between the uncompensated own wage effects on market labor supply in landholding and landless households must be positive. Subtracting the relevant "landless" Slutsky relations from (24) and (25) yields

$$(26) \quad \frac{\delta \lambda_M^L}{\delta W_M} - \frac{\delta \lambda_M^N}{\delta W_M} = - \frac{\Gamma_{mm}}{\Delta} + m \left(\frac{\delta M}{\delta I} \right) = - \frac{\delta m}{\delta W_M} + m \left(\frac{\delta M}{\delta I} \right) > 0$$

$$(27) \quad \frac{\delta \lambda_F^L}{\delta W_F} - \frac{\delta \lambda_F^N}{\delta W_F} = - \frac{\Gamma_{ff}}{\Delta} + f \left(\frac{\delta F}{\delta I} \right) = - \frac{\delta f}{\delta W_F} + f \left(\frac{\delta F}{\delta I} \right) > 0.$$

Expressions (26) and (27) indicate that if "peasant" households behave in a "neoclassical" manner and if labor markets are competitive, the own net market supply response to a wage change in landed households will be algebraically greater than that in landless households. This differential arises because an increase in the own wage

leads to a reduction in family labor time spent on the land owned by the landholding family, Γ_{mm}/Δ, $\Gamma_{ff}/\Delta < 0$, and because the rise in income associated with the wage increase is attenuated in landholding households (relative to that in landless households supplying the same total amount of labor) by the relevant labor input (m, f) becoming more expensive.[8] This "neoclassical" prediction is not an obvious implication of the labor surplus hypothesis.

The juxtaposition of landless and landholding market labor supply responses also provides a framework for testing for the existence of the hypothesized linkage between education (experience) and managerial efficiency. Let $\delta M/\delta E_K$ and $\delta F/\delta E_K$ be the unknown relationships between the demand for nonmarket time and schooling, identical for both landless and landholding households. From (11) and (21), the relationship between market labor supply and schooling in landholding households is thus given by

$$(28) \quad \frac{\delta \lambda_M^l}{\delta E_K} = -\frac{\delta M}{\delta E_K} - \left[\frac{\Psi_i (\Gamma_{fe}\Gamma_{mf} - \Gamma_{me}\Gamma_{ff})}{\Delta} \right] = -\frac{\delta M}{\delta E_K} - \frac{\delta m}{\delta E_K}$$

$$K = M, i = 1$$
$$K = F, i = 2$$

$$(29) \quad \frac{\delta \lambda_M^l}{\delta E_K} = -\frac{\delta F}{\delta E_K} - \left[\frac{\Psi_i (\Gamma_{me}\Gamma_{mf} - \Gamma_{fe}\Gamma_{mm})}{\Delta} \right] = -\frac{\delta F}{\delta E_K} - \frac{\delta f}{\delta E_K}.$$

The second terms in (28) and (29), the effects of schooling on the demand for farm labor inputs, must be positive if schooling enhances the productivity of inputs. Thus, whatever the signs of $\delta M/\delta E_K$, $\delta F/\delta E_K$, the response of market labor supply to educational levels in landholding households will be algebraically less than that in landless households if schooling augments efficiency, the magnitude of the differential being the effect of the schooling attainment of family members on the demand for labor on the farm; i.e.,

$$(30) \quad \frac{\delta \lambda_M^l}{\delta E_K} - \frac{\delta \lambda_M^N}{\delta E_K} = -\frac{\delta m}{\delta E_K} < 0$$

$$(31) \quad \frac{\delta \lambda_F^l}{\delta E_K} - \frac{\delta \lambda_F^N}{\delta E_K} = -\frac{\delta f}{\delta E_K} < 0.$$

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8. The differential in gross *cross* effects, given by (26') and (27'), cannot be signed, since the smaller income effect on landholding households may be wholly offset by the production substitution effect, the increase in family labor time of males (females) in response to an increase in the wage rate of females (males).

$$(26') \quad \frac{\delta \lambda_M^l}{\delta W_F} - \frac{\delta \lambda_M^N}{\delta W_F} = -\frac{\Gamma_{mf}}{\delta} + f\left(\frac{\delta M}{\delta I}\right) = -\frac{\delta m}{\delta W_F} + f\left(\frac{\delta M}{\delta I}\right)$$

$$(27') \quad \frac{\delta \lambda_F^l}{\delta W_m} - \frac{\delta \lambda_F^N}{\delta W_M} = -\frac{\Gamma_{fm}}{\Delta} + m\left(\frac{\delta F}{\delta I}\right) = -\frac{\delta f}{\delta W_M} + m\left(\frac{\delta F}{\delta I}\right).$$

Similar results would obtain for differential experience effects (as indicated by age $= A_K^L, A_K^N$) if such experience is relevant to managerial efficiency only on a household's own land.

Refutable predictions can also be derived directly from the landholding model with respect to the relationship between nonlabor farm inputs and market labor supply:

$$(32) \quad \frac{\delta \lambda_M^L}{\delta \kappa} = - \Gamma_\kappa \frac{\phi_{62}^L}{\phi^L} - \frac{\Gamma_{f\kappa} \Gamma_{mf} - \Gamma_{m\kappa} \Gamma_{ff}}{\Delta} = - \Gamma_\kappa \left(\frac{\delta M}{\delta I} \right) - \frac{\delta m}{\delta \kappa} < 0$$

$$(33) \quad \frac{\delta \lambda_F^L}{\delta \kappa} = - \Gamma_\kappa \frac{\phi_{63}^L}{\phi^L} - \frac{\Gamma_{m\kappa} \Gamma_{mf} - \Gamma_{f\kappa} \Gamma_{mm}}{\Delta} = - \Gamma_\kappa \left(\frac{\delta F}{\delta I} \right) - \frac{\delta f}{\delta \kappa} < 0.$$

Since an increase in the level of inputs κ both raises the demand for labor time spent in farm production and, through the income effect, increases the demand for leisure (normality assumed), (32) and (33) must be negative. The magnitude of the farm asset effect is proportional to the marginal product of the factor input, the own leisure-income effect, and the response of labor time to other input changes. Household members on farms more endowed with production assets will thus participate less in the labor market.

III. EMPIRICAL ANALYSIS

The Data and Estimation Techniques

In this section the labor supply predictions derived from the landless and landholding household models formulated under the assumption of competitive labor markets are tested using data from a national sample survey of rural households in India collected in three rounds, 1968–1969, 1969–1970, and 1970–1971, and coded by the National Council of Applied Economic Research (NCAER). This survey provides information on a wide variety of household and farm characteristics, including the number of annual days worked for pay in agricultural and nonagricultural activities, and earnings from those activities for each household member. The sample used, stratified into landless and landholding households, is based on information collected in the third round of the survey, 1970–1971. Households in which either the head or spouse were absent, or were government employees, or salaried workers were excluded so that the data are restricted to cultivators and "casual" workers employed on a monthly or daily basis.

The market (for pay) labor supply equations to be estimated for heads of households and their wives in the two subsamples are given by (34) and (35):

$$(34) \quad \lambda_K^N = \alpha_K^N + \beta_{1K}^N W_M + \beta_{2K}^N W_F + \beta_{3K}^N I^N + \beta_{4K}^N E_M^N + \beta_{5K}^N E_F^N$$
$$+ \beta_{6K}^N A_M^N + \beta_{7K}^N A_F^N + \beta_{ZK}^N Z_K^N + u_K^N \qquad K = M,F$$

$$(35) \quad \lambda_K^L = \alpha_K^L + \beta_{1K}^L W_M + \beta_{2K}^L W_F + \beta_{3K}^L I^L + \beta_{4K}^L E_M^L + \beta_{5K}^L E_K^L$$
$$+ \beta_{6K}^L A_M^L + \beta_{7K}^L A_F^L + \sum_{i=8}^{11} \beta_{ik}^L \kappa_i + \beta_{ZK}^L Z_K^L + u_K^L,$$

where the β_{jK}^N, β_{jK}^L are the relevant coefficients for the landless and landholding households, the Z^N, Z^L are vectors of control variables, to be discussed below, and the u_K^N, u_K^L are stochastic error terms. The theoretical analysis implies the following coefficient or coefficient differential signs:

1. $\beta_{1M}^L - \beta_{1M}^N > 0$

2. $\beta_{2F}^L - \beta_{2F}^N > 0$

3. $\beta_{3K}^N < 0$

4. $\beta_{3K}^L < 0$

5. $\beta_{4K}^L - \beta_{4K}^N < 0$

6. $\beta_{5K}^L - \beta_{5K}^N < 0$

7. $\beta_{6K}^L - \beta_{6K}^N < 0$

8. $\beta_{7K}^L - \beta_{7K}^N < 0$

9. $\beta_{ik\bullet}^L < 0$
   $\quad i = 8, \ldots, 11 \qquad K=M,F.$

Sign relations 1 and 2 reflect the differential own gross wage effects in landless and landholding households for the two sexes, from (26) and (27);[9] 3 and 4 are consistent with the assumption that leisure is a normal good; coefficient restrictions 5 through 8 embody the hypothesis that schooling and experience (A_K^N, A_K^L) augment the managerial ability of the husband and wife in agricultural production, from (20) and (31); and the four sign predictions in 9 correspond to the predicted farm production asset effects on net labor supply, from (32) and (33).

Because the NCAER data provide no information on labor input use on the land held by landholding households, net labor supply—the difference between sex-specific total labor supply ($\Omega - M, \Omega - F$) and

TABLE I

MEAN HOUSEHOLD CHARACTERISTICS BY SEX, MARKET PARTICIPATION, AND
LAND OWNERSHIP

	Males			Females		
	No-market	Market	Total	No-market	Market	Total
			Landless			
n	0	309	309	82	227	309
DAYS	—	247.7	247.7	0	195	143.4
EDH	—	1.04	1.04	1.39	0.92	1.04
EDW	—	0.48	0.48	0.32	0.53	0.48
AGE	—	43.3	43.3	40.4	35.6	36.9
KIDS	—	0.64	0.64	0.60	0.65	0.64
			Landed			
n	510	352	862	611	251	862
DAYS	0	166	68.0	0	171	49.9
LAND	13.22	4.66	9.72	11.87	4.49	9.72
EDH	2.39	1.53	2.04	2.29	1.43	2.04
EDW	0.48	0.28	0.40	0.46	0.25	0.40
AGE	50.2	44.1	47.7	42.7	36.2	40.8
KIDS	1.01	0.80	0.92	0.95	0.86	0.92

total farm labor usage (m,f)—is observed only for households in which the head or wife worked off the farm, i.e., for $\lambda_K^L > 0$. Table I, which gives household characteristics and days worked by sex and land ownership for the total sample, indicates that while all the heads of landless households and 73.5 percent of their wives worked at least one day for pay, only 40.8 percent of household heads with land and 29.1 percent of their wives supplied any market labor. The dependent variable used to represent net labor supply, days worked for pay D_K^L, is thus censored, bounded at zero and concentrated at that bound in the landholding subsample; i.e.,

$$D_K^L = 0, \qquad \lambda_K^L - u_K^L \leq 0$$
$$D_K^L = \lambda_K^L - u_K^L, \qquad \lambda_K^L - u_k^L > 0.$$

These properties of the dependent variable imply that if u_K^L is distributed $N(0,\sigma)$, the tobit estimation procedure would be more appropriate than classical least squares in the estimation of equations (35) (see Tobin [1958]), where λ_K^L would represent the tobit index and D_K^L the observed days worked off the farm. However, unlike the usual "corner solution" application of tobit in U. S. female labor supply studies (Rosen [1976], Schultz [1975]), all males in the landholding subsample are earners, and the "true" index λ_M^L may take on negative values (for net hirers of labor). The tobit index, or net supply, coef-

ficients for males are thus appropriately compared to the least squares landless male coefficients, estimated from equation (34) for which censoring is not a problem, $(D_M^N = \lambda_M^N)$ in verifying the restrictions of the neoclassical framework. Only for purposes of predicting the relationships between observed off-farm work and the independent variables are the "expected value" or observed days worked elasticities relevant. In the case of females, however, a proportion in both types of households devote all their time to household activities; thus for the landholding subsample the female days worked (for pay) D_F^L variable may not only be censored but may also be zero-valued because the wife does not participate in any earnings activities.

A second consequence of the lack of information on labor use in landholding households is that daily wage rates paid to laborers by households holding land but supplying no labor to the market, and thus the value of the time of family labor, are not available. The usual procedure employed in U. S. (female) labor supply studies, both to solve the missing wage problem and to eliminate the definitional relationship between the labor supply variable and the computed wage, is to impute a wage rate based on the personal characteristics of the relevant household member.[10] In Indian rural labor markets, however, the chief source of wage rate variability appears to be geographical rather than personal once sex has been taken into account—annual averages of daily agricultural wages computed within sharply defined categories such as weeding, reaping, plowing, etc., and stratified by sex and adult status vary significantly across Indian districts. Due presumably to the geographical immobility of rural households and the nature of rural occupations, individual wage rates thus may be determined by the interaction of aggregate labor demand and supply in individual labor markets, which is in turn a function of such factors as the distribution of landholdings, availability of water, and the existence of rural industry.[11]

Table II displays for heads and wives alternative specifications of wage equations in which the dependent variable is the natural logarithm of the computed (sex-specific) daily wage based on a combined sample of landless and landholding households in which either the head or the wife worked in the market. In specification 1, which corresponds to a human capital earnings function,[12] schooling attainment and the two age variables explain less than 3 percent of

the variation in male wages and *none* of the variance in the female wage rate (the critical F-value (500, 3) = 3.86 (5 percent level)), although the coefficient of the schooling of the male head is statistically significant. Specification 2 includes characteristics of the local labor market reported in the sample survey data that may affect daily wage rates—dummy variables taking on the value of one if crops are *not* adversely affected by weather conditions (*WEATHER*), if a factory is present in the village (*FACTRY*) or if there is any small scale industry (*SSIND*), and variables indicating the size of the village (*SIZEVLG*), the distance in kilometers between the household's residence and the village (*DSTNCE*), and whether or not the household resides in an agricultural development district subject to governmental technical assistance and credit programs (*IADP*). These variables, while adding significantly to the explanatory power of the wage equations for both males and females, do not, however, completely capture all the important characteristics of local labor markets that might influence wage levels. As a proxy for aggregative market conditions, therefore, the natural logarithm of the sex-specific district-level daily wage pertaining to the district in which the household resides (*LWAGE*) is added in specification 3.[13] The inclusion of this variable not only further improves the explanatory power of the wage equations but reduces the male schooling coefficient to insignificance; thus none of the human capital characteristics of the individual are significantly correlated with the wage received. The lack of significance of the schooling variables in the more fully specified equations explaining the wage rates of nonsalaried and nongovernment workers of both sexes should not, however, be interpreted as evidence that schooling does not increase earnings in India. Aside from the managerial efficiency effect for heads and wives in farm households, which is discussed below, schooling attainment appears to be positively correlated with the likelihood of being in a salaried or government job, where computed mean wage rates are higher than those observed in the sample of workers used.

The results in specification 3 are consistent with the hypotheses that labor is not perfectly mobile geographically in rural India and that wage rates are not importantly affected by human capital attributes in the nonsalaried, private-sector occupations characterizing

13. The correlation between the district-level male agricultural wage rates and a linear combination of such rural district characteristics as average landholding size, the population of households without land, a measure of the variance in the size-distribution of landholdings, the proportion of irrigated farms, and annual rainfall is 0.68, where the weights are least squares regression coefficients. The correlation for the female wage rate is 0.65.

TABLE II
SEX-SPECIFIC LN WAGE EQUATIONS, NONSALARY MARKET WORKERS

Independent variable	Male (1)	(2)	(3)	(4)	Female (1)	(2)	(3)	(4)
ED	0.060	0.039	0.009	0.010	0.007	0.022	0.021	0.020
	(4.12)	(2.45)	(0.75)	(0.83)	(0.61)	(1.44)	(1.40)	(1.30)
AGE	−0.007	0.009	0.008	0.008	−0.023	−0.016	−0.016	−0.016
	(0.57)	(1.00)	(1.00)	(0.01)	(1.55)	(1.10)	(1.07)	(1.10)
AGESQ	0.0001	−0.0001	−0.0001	−0.0001	−0.0003	0.0003	0.0003	0.0003
	(0.60)	(0.71)	(0.70)	(0.72)	(1.45)	(1.71)	(1.58)	(1.69)
WEATHER		0.166	0.166	0.168		0.091	0.099	0.095
		(3.46)	(3.70)	(3.73)		(1.75)	(1.90)	(1.79)
FACTRY		0.503	0.440	0.443		0.231	0.218	0.217
		(5.43)	(5.06)	(5.05)		(2.12)	(1.98)	(1.97)
SSIND		0.087	0.033	0.032		0.118	0.117	0.118
		(1.37)	(0.56)	(0.53)		(1.55)	(1.54)	(1.55)
SIZEVLG(×10⁻³)		0.062	0.042	0.043		0.076	0.069	0.068
		(9.98)	(6.56)	(8.17)		(7.67)	(6.27)	(6.18)
DSTNCE		0.224	0.256	0.250		1.274	1.338	1.322
		(0.34)	(0.42)	(0.41)		(2.00)	(2.10)	(2.08)
IADP		0.006	0.050	0.049		0.007	0.012	0.030
		(0.12)	(1.07)	(1.04)		(0.12)	(0.02)	(0.51)
LWAGE			0.419	0.423			0.119	0.116
			(7.75)	(7.69)			(2.37)	(2.05)
LAND				0.002				−0.003
				(0.55)				(0.61)
C	1.075	0.364	0.057	0.051	0.127	0.299	0.252	0.256
$\bar{R}^2$	0.029	0.246	0.338	0.337	−0.001	0.263	0.265	0.263
F	5.73	16.72	23.09	20.96	0.98	11.26	10.32	9.39
n	900	900	900	900	522	522	522	522

t-values in parentheses.

the rural labor market. It is possible, however, that if the wage labor market is noncompetitive, jobs may be rationed according to the status of the worker, with discrimination in wage offers related to size of landholdings, for example. In specification 4 the amount of land owned ($LAND$) by the worker is entered as an additional regressor. The insignificance of this variable in the two wage equations, however, indicates that large landowners are not able to exert market power to obtain higher wages. Moreover, stratification of the worker samples by landholding status (not reported) and application of the Chow test indicates that the hypothesis that the sets of coefficients (excluding $LAND$) in the wage equations for the two land status groups are identical cannot be rejected. The hypothesis that all farm workers in the same geographical labor market and of the same sex face the same wage thus appears to be supported by the data.

The relative unimportance of personal attributes in determining the wages received by market workers suggests as well that rural wages are not significantly affected by the number of days worked (which *is* a function of the personal characteristics of the individual worker). Thus, selectivity bias, inherent in a wage imputation procedure, based on specification 3 of Table II, may not be significant, since the error components in the wage equations, based on market conditions, are likely to be minimally correlated with the error terms in the individual supply (shadow wage) equations, consisting mainly of household variables.

The male and female wage rates used in equations (34) and (35) are thus estimated using the quasi-instrumental variables approach, based on a wage-predicting equation, including the variables of specification 3 of Table II but without schooling and age.[14] Of the other regressors in (34) and (35) requiring comment, the household's combined income from interest, dividends, and other personal (nonfarm) property income is used to represent non-earnings income ($NEARN$) and the age of the head and spouse ($AGEM/AGEF$) are included to capture life-cycle and cohort effects in the landless sample and to serve in addition as proxies for farm-specific work experience in landholding households. The variables representing nonlabor farm assets, κ_8, κ_9, κ_{10}, κ_{11}, consist of a three-year average of gross cropped area, in acres ($LAND$), and dummy variables representing farm irrigation ($IRR = 1$ if irrigated, 0 otherwise) weather conditions, and whether or not the farm household resides in an agricultural devel-

14. The labor supply results reported below are not significantly altered when age and schooling variables are used in the wage-predicting equations; however, significance levels decline.

opment district ($IADP$) and thus is exposed to governmental credit programs (increasing access to credit) and to the introduction of high-yielding grain varieties. Each of these farm assets variables should be positively correlated with farm labor productivity and thus negatively related to market (off-farm) labor supply.[15]

Included in the Z-vector are variables representing proximity to sources of nonagricultural employment—$FACTRY$, $SSIND$, $DSTNCE$—which will be significant determinants of annual days worked for geographically immobile laborers.

The number of children less than age five ($KIDS$) is also added to the market supply equations to test whether the presence of young children is importantly related to work decisions in rural areas of a developing country. However, because this demographic variable is likely to be endogenous (see Rosenzweig and Evenson [1977]), two specifications are used, one with the children variable omitted.

Male and Female Market Supply Function Parameter Estimates: Landless and Landholding Households

Tables III and IV report the coefficient estimates obtained for the market labor supply functions of males and females in landless and landholding households using ordinary least squares instrumental variables (OLS-IV) and tobit (TOBIT-IV). The overall results (which are not qualitatively altered by the further stratification of the subsamples according to the wife's participation in earning activities) are generally supportive of the neoclassical framework—of the twenty-two possible refutable sign restrictions only one, the differential in the male age coefficients in the female supply equations (Table IV) is wrong, although it is not statistically significant. Of the twenty-one correct coefficient signs, fourteen are statistically significant at (at least) the 10 percent level.

As was discussed, the estimates from the *landless* supply equations cannot be used by themselves as a test of neoclassical labor supply theory. Although the results for landless male workers are particularly weak in terms of the statistical significance of the individual coefficients (such is not the case for landless females), the hypothesis that days worked by landless males is *randomly* determined, one possible job allocation mechanism in a labor surplus economy, is rejected at the 1 percent level (critical F (15,200 +) = 2.05). Moreover, while the eleven independent variables explain only

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15. A dummy variable representing farm tenancy did not attain statistical significance in any of the equations and is thus omitted from the reported specifications.

TABLE III

OLS-IV AND TOBIT-IV MARKET SUPPLY EQUATIONS, ANNUAL DAYS
WORKED FOR PAY BY NONSALARIED MALES

| Independent variable | Landless | | Landholding | | | |
| | OLS-IV | | OLS-IV | | TOBIT-IV | |
	(1)	(2)	(1)	(2)	(1)	(2)
$PWAGEM$[†]	−16.29	−17.35	−11.52	−11.27	−7.10	−7.12
	(1.43)	(1.51)	(1.29)	(1.26)	(3.43)	(3.44)
$PWAGEF$[†]	11.66	13.91	4.68	3.62	62.03	61.02
	(0.69)	(0.82)	(0.29)	(0.22)	(1.72)	(1.69)
EDM	2.77	2.94	−4.18	−4.19	−9.04	−9.18
	(0.78)	(0.84)	(2.03)	(2.03)	(1.97)	(2.00)
EDF	−0.971	−0.968	−4.18	−4.27	−8.55	−8.77
	(0.43)	(0.43)	(2.00)	(2.04)	(1.82)	(1.86)
$NEARN$	−0.038	−0.041	−0.005	−0.005	−0.045	−0.045
	(1.19)	(1.27)	(0.64)	(0.65)	(1.48)	(1.47)
$LAND$			−2.20	−2.14	−12.58	−12.39
			(8.00)	(7.66)	(10.46)	(10.17)
IRR			−22.20	−22.58	−36.14	−36.63
			(3.70)	(3.76)	(2.70)	(2.74)
$WEATHER$			−1.83	−1.93	−15.14	−15.52
			(0.26)	(0.27)	(0.95)	(0.97)
$IADP$			−36.59	−36.70	−79.57	−79.66
			(5.43)	(5.45)	(5.82)	(5.83)
$FACTRY$	7.74	7.21	24.72	24.48	93.67	92.60
	(0.55)	(0.51)	(1.88)	(1.86)	(2.97)	(2.94)
$SSIND$	4.45	3.68	23.91	22.88	53.72	51.73
	(0.33)	(0.27)	(2.26)	(4.63)	(2.28)	(2.19)
$DSTNCE(\times 10^{-3})$	−40.43	−39.48	−68.90	−67.46	−485.96	−482.26
	(0.35)	(0.34)	(1.24)	(1.21)	(2.02)	(2.01)
$AGEM$	−1.10	−0.990	−0.327	−0.363	−1.24	−1.30
	(1.19)	(1.07)	(0.61)	(0.68)	(1.07)	(1.13)
$AGEF$	−0.499	−0.485	−1.44	−1.42	−2.34	−2.34
	(0.50)	(0.49)	(2.59)	(2.58)	(1.91)	(1.91)
$KIDS$		6.54		−3.06		−6.38
		(1.09)		(1.09)		(0.96)
C	332.29	321.70	212.33	216.58	374.87	384.49
					(8.15)	(8.16)
$\overline{R}^2$	0.054	0.054	0.257	0.257		
F/χ^2	2.75	2.60	22.21	20.82	145.9	145.9
n	309	309	862	862	862	862

Asymptotic t-values in parentheses.
† Instrumental variable.

5 percent of the variation in days worked by landless males, similar (landless) male supply equations estimated on micro data from the United States suffer also from low explanatory power [(Knieser,

1976)]. The landless male results may thus indicate more the limitations of the neoclassical theory of labor supply than the existence of a "surplus labor" economy. Interestingly, the own landless male supply elasticity estimate of -0.16 is consistent with estimated male supply elasticities obtained by Kneiser [1976] (dependent variable = weeks worked) and Finegan [1962] based on U. S. cross-sectional household and aggregate data. The negative signs of the non-earnings income coefficients in *all* equations are, moreover, in accord with the expectations that leisure is a normal good and thus are consistent with negative own wage effects on labor supply, but the estimates only approach statistical significance.

The tobit estimates for landholding households indicate that the *net* labor supply of farm males is also backward bending, with the own wage coefficient significantly less than zero at the 0.01 level; the observed off-farm days worked own wage elasticity is -0.18. Consistent with the theoretical framework, the coefficient of *NEARN* is negative, significant at the 0.10 level, and the male wage coefficient estimate is algebraically greater in the landholding than in the landless households, although the difference is not statistically significant. However, the negative differential in the male education coefficients between the two households is significant at the 0.05 level and supports the hypothesis that the schooling of male farm managers improves managerial efficiency. Thus, higher schooling levels of male heads of landholding households are associated with lower levels of (male) net labor supply, despite the small positive association between male schooling and male market work indicated in the landless equations. The more negative coefficient for female schooling in the landholding males equations, significant at the 0.10 level, additionally supports the hypothesis that the formal education of farm wives enhances the productivity of all farm inputs, including the husband's time in farm production. However, the coefficients of the age variables in the two households suggest that farming experience has only a minimal productivity effect; the age coefficient differentials have correct signs but are not statistically significant.

Another difference between the two subsamples is that the proximity of a factory or the presence of small-scale industry near the household is significantly and positively associated only with the market days worked of farm males, suggesting that males from farm households are significantly less geographically mobile than landless males. Such a result is consistent with the notion that there are strong imperfections in land and capital markets in India as suggested by Bardhan [1973] and Sen [1966].

TABLE IV

OLS-IV AND TOBIT-IV MARKET SUPPLY EQUATIONS, ANNUAL DAYS WORKED FOR PAY BY NONSALARIED FEMALES

Independent variable	Landless				Landholding			
	OLS-IV		TOBIT-IV		OLS-IV		TOBIT-IV	
	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
PWAGEF†	50.78	50.77	50.95	49.93	3.54	3.85	58.18	58.53
	(2.23)	(2.20)	(1.46)	(1.42)	(0.22)	(0.24)	(1.82)	(1.69)
PWAGEM†	-61.73	-61.73	-79.95	-79.51	-10.39	-10.46	-82.96	-82.64
	(3.99)	(3.97)	(3.62)	(3.59)	(1.18)	(1.18)	(2.78)	(2.77)
EDF	2.27	2.27	2.94	2.94	-2.87	-2.83	-7.90	-7.79
	(0.75)	(0.75)	(0.75)	(0.75)	(1.39)	(1.37)	(1.21)	(1.18)
EDM	3.17	3.17	2.50	2.40	-4.25	-4.25	-13.46	-13.32
	(0.67)	(0.66)	(0.40)	(0.38)	(2.09)	(2.09)	(2.04)	(2.02)
NEARN	-0.061	-0.061	-0.166	-0.165	-0.006	-0.006	-0.094	-0.094
	(1.41)	(1.41)	(1.79)	(1.78)	(0.75)	(0.74)	(1.48)	(1.48)
LAND					-1.42	-1.43	-12.20	-12.36
					(5.23)	(5.20)	(7.38)	(7.28)
IRR					-18.35	-18.24	-42.21	-42.01
					(3.14)	(3.12)	(2.28)	(2.22)
WEATHER					-13.49	-13.51	-29.28	-29.44
					(1.96)	(1.97)	(1.33)	(1.34)

IADP					−28.78	−28.74	−94.83	−94.61
					(4.36)	(4.35)	(4.97)	(4.96)
FACTRY	27.40	27.40	34.7	34.91	16.40	16.47	74.33	74.81
	(1.43)	(1.43)	(1.37)	(1.38)	(1.28)	(1.28)	(1.63)	(1.64)
SSIND	39.49	39.49	48.12	48.52	21.36	21.66	68.68	69.86
	(2.17)	(2.16)	(2.02)	(2.03)	(2.05)	(2.07)	(2.04)	(2.06)
DSTNCE ($\times 10^{-3}$)	−149.74	−149.74	−210.58	−210.88	−28.21	−28.62	−354.67	−356.86
	(0.95)	(0.95)	(1.04)	(1.04)	(0.53)	(0.54)	(1.31)	(1.32)
AGEF	−0.577	−0.577	−0.727	−0.733	−1.56	−1.56	−4.07	−4.05
	(0.43)	(0.43)	(0.41)	(0.41)	(2.83)	(2.83)	(2.27)	(2.26)
AGEM	−1.08	−1.08	−1.55	−1.59	0.203	0.213	−0.590	−0.556
	(0.86)	(0.86)	(0.95)	(0.97)	(0.39)	(0.40)	(0.35)	(0.34)
KIDS		−0.015		−3.00		0.895		4.02
		(0.01)		(0.28)		(0.33)		(0.42)
C	270.80	270.83	324.53	329.38	156.46	155.22	359.82	353.54
			(7.67)	(7.20)			(5.46)	(5.25)
$\bar{R}^2$	0.144	0.141		0.166	0.166	0.165		
F/χ^2	6.19	5.61	3.05	3.05	13.26	12.37	4.22	4.22
n	309	309	309	309	862	862	862	862

Asymptotic t-values in parentheses.
† Instrumental variable.

Of the farm production asset variables, all the coefficients also display the theoretically correct (negative) signs, with those of *LAND*, *IRR*, and *IADP* statistically significant at the 0.01 level. The coefficient estimates suggest that a 10 percent increase in gross cropped area is associated with a 12 percent decline in the number of days worked off the farm by heads of landholding households and that the net supply of male labor on farms with irrigation facilities or in *IADP* districts is approximately thirty-six and fifty man-days less than that on unirrigated farms or on farms in non-*IADP* areas.

In the females equations of Table IV the qualitative results are similar to those obtained for males except that the market supply curves of women appear to be positively sloped, consistent with U. S. studies of female labor supply [Rosen, 1976; Rosen and Welch, 1971; and Schultz, 1975]. The tobit and OLS estimates of the female supply coefficients in the landless subsample are not significantly different, due to the high proportion of landless women participating in the market, except that the negative coefficient of *NEARN* increases in absolute value in the tobit equation. However, as expected, the OLS and tobit net female supply coefficients in the landholding subsample diverge significantly, with all coefficients increasing in absolute value in the tobit equation. The tobit estimates indicate that the observed days worked elasticity for women from landless households is 0.67, the observed female off-farm work elasticity is 0.72, and the net supply elasticity of farm women is 2.0.

The estimated gross male wage effects on female market supply in both landless and landholding households are negative and significant, consistent with the U. S. results cited above. Indeed, female market labor supply appears quite sensitive to movements in the male wage—a 10 percent rise in the wage rate of males is associated with a 14 percent reduction in the number of days worked by landless females and a 20 percent decrease in the number of days worked off the farm by wives of landholders, the latter in part due to the substitution of the wife's time for male labor in farm production, as suggested in equation (23) of the theoretical analysis.

Of the "predicted" coefficients, all but one conform to the implications of the neoclassical framework—the differential in the male age effect on female supply between the two households. All the theoretically correct (tobit) coefficient signs or sign differentials, except for the differential own gross wage effect, are statistically significant (0.10 level). Thus, as indicated by the theory and as found for rural males, less market work is supplied by women in households with higher levels of non-earnings income, greater landholdings, and

irrigated land, which are located in agricultural development districts and in areas experiencing good weather. Moreover, the schooling attainment of both household heads and their wives is associated significantly more negatively with the number of market days worked by wives in landholding than in landless households.

The presence of children less than five years of age appears to have no significant effect on the market labor supply of women in India, a result that contrasts with findings based on U. S. data (see Kniesner [1976], Leibowitz [1972], Rosen [1976], and Schultz [1975]), suggesting that *market* work and child rearing are not competitive activities in rural areas of developing countries. Thus, even if a part of fertility is "excess," in the sense that the number of children born to a family exceeds the number that would have been born if parents had more access to birth control information, the results suggest that the intensification of family planning programs in India may not have a significant impact on the quantities of labor supplied to the market by rural women (or men).

Finally, the results indicate, in contrast to those for males, that the proximity of small-scale industry, and to a lesser extent of a factory, is associated with higher amounts of market work by females in landless as well as landholding households, suggesting that females are significantly less geographically mobile than males in rural India, although female labor supply is not less responsive than male labor supply to changes in economic variables.

IV. CONCLUSION

Little empirical evidence exists on labor supply behavior in rural areas of developing countries and on the state of competitiveness of rural labor markets. Yet such information is crucial to any model of economic development formulated to serve as a useful policy-prescribing apparatus. In this paper refutable predictions were derived from the joint consideration of market labor supply behavior in neoclassical models of landless and landholding households to establish a test of the competitive framework in the context of rural labor markets in less developed countries. Empirical results based on micro data from rural India stratified by sex and landholding status were generally supportive of the neoclassical framework suggesting that the annual number of days wage of employment observed for individuals in rural India is *mainly* supply rather than demand determined, as implied by competitive models. The estimates also appeared to reject some simple labor surplus models of wage employment. Male

and female labor supply function estimates appeared similar in many respects to econometric labor supply findings based on U. S. data with the exception of the impact of fertility variables on labor supply, which was insignificant. The results also were consistent with the hypothesis that schooling, for both male and female members of landholding households, enhances agricultural production efficiency in India and thus tends to reduce the off-farm labor supply of cultivators (male and female), but indicate that geographical immobility is a marked characteristic of rural labor markets, particularly for males in land-holding households and women.

The evidence obtained thus points to the necessity of distinguishing empirically between the behavior of members of landless and landowning families in rural areas of developing countries and calls into question the implications of development models that assume exogenously fixed rural wage rates. While some of the results obtained admit to alternative interpretations and suffer from lack of precision, the further examination of the micro foundations of macro development models would appear to be a productive area of research.

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