Local Economic Conditions and Wage Labor Decisions of Farm and Rural Nonfarm Couples

J. G. Tokle and Wallace E. Huffman

Effects of geographical differences in local economic conditions on wage labor demand and wage labor participation decisions of rural couples are examined for Current Population Survey households 1978–82. Wage premiums are shown to exist for localities anticipating labor demand growth, higher unemployment rates, larger share of employment in services, and higher costs of living. These effects are stronger for males than females. Effects of local economic conditions on the probability of wage work are consistent with expected market wage and reservation wage effects, and for farm households the probability of wage work increases when expected farm output prices decline or the wage increases.

Key words: farm couples, human capital, local business cycles, local labor markets, rural, wage labor participation, wage rates.

The objective of this paper is to examine the effects of geographical differences in local economic conditions on wage labor demand and labor force participation decisions of U.S. farm and rural nonfarm couples. One hypothesis is that localities with higher anticipated employment growth and unemployment rates pay a wage premium to attract workers from other localities. Thus, when the local anticipated employment growth rate declines (relative to the national rate) real wage rates fall in these localities and rise in others. Another hypothesis is that localities experiencing unanticipated negative labor market shocks show a decline of real wage rates; an example is the case of the upper Midwest during 1981–82. Other geographical wage differences are the result of cost of living and locational amenities.

A considerable amount of research on labor supply of married males and females has utilized a single-worker model (DaVanzo, DeTray, and Greenberg; Mroz, table 1). Furthermore, Mroz has continued this tradition in a recent analysis of the sensitivity of female labor supply to an array of economic and statistical assumptions. However, we believe that new insights can be gained by considering labor supply decisions in a two-worker, husband-wife model, e.g., see Huffman and Lange. In particular, a spouse's wage has substitution and income effects rather than just income effects alone. A spouse's education may affect a partner's labor supply decision by changing the reservation wage through efficiency effects on family wealth and household production (and for farm households on farm production) or taste. A couple's labor supply decisions are affected by the same economic shocks, so a joint estimation procedure should increase the statistical efficiency of the parameter estimates.

The paper has the following organization: The empirical setting for the study is presented first. Second, the economic models of labor demand and labor supply are developed, including local labor market effects. Third, the data and econometric model are summarized. Fourth, the econometric results are reported. Finally, some conclusions and implications are presented.

The authors are, respectively, an assistant professor, Idaho State University, and a professor, Iowa State University.

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The Empirical Setting

This section presents summary measures of differences across localities in labor market characteristics, primarily unemployment rates, employment growth, and shocks to labor demand. The basic geographic unit is a state because states are the smallest unit for which annual data are available. Statistics are reported for the twenty-three states having the largest rural population in 1980.

Equilibrium unemployment rates differ geographically and have not tended to converge over time. Table 1 presents the ratio of predicted state unemployment rates to the predicted national unemployment rate for 1970, 1974, 1978, and 1982.[1] Only four years are reported for each state to conserve space; these years lead up to and include the 1978–82 period covered in this study. At any point in time important regional differences in unemployment rates exist and tend to persist over time (e.g., Hall, Abowd and Ashenfelter, Adams, Topel). For example, the unemployment rates are uniformly higher than average for Michigan, West Virginia, Mississippi, and Louisiana. The unemployment rates are uniformly below average for Pennsylvania, Minnesota, Iowa, Virginia, North Carolina, Georgia, and Texas. For other states, no simple relationship exists during 1970–82 between their unemployment rate and the national average rate.

Table 1 also presents the ratio of predicted state employment growth rates to the predicted national employment growth rate.[2] The table shows that the employment growth rates in Georgia, Florida, Louisiana, Texas, and Cali-

[1] Each state's annual unemployment rate, 1968–82, and the national unemployment rate were regressed on quadratic trend. The predicted unemployment rates are forecasts from these regressions.

[2] The natural logarithm of total private sector employment, 1968–82, was regressed on quadratic trend. Predicted employment growth rates are the first differences of the predicted employment values from these regressions.

Table 1. Geographical Distribution of Unemployment and Employment, Twenty-three Selected States, 1970–82

	Unemployment[a]				Employment[b]				Rural population
States	1970	1974	1978	1982	1970	1974	1978	1982	1980 (1,000)
New York	1.09	1.25	1.17	.94	−.32	−.04	.22	.46	2,700
Pennsylvania	.89	.98	1.08	.90	.45	.35	.26	.17	3,643
Ohio	.89	.92	1.07	1.27	.86	.58	.30	.25	2,879
Indiana	.87	.85	1.04	1.33	1.09	.71	.30	−.04	1,965
Illinois	.83	.82	.97	1.20	.41	.40	.39	.38	1,908
Michigan	1.28	1.26	1.42	1.63	1.09	.66	.26	−.13	2,711
Wisconsin	.85	.74	.85	1.10	1.23	1.02	.82	.63	1,685
Minnesota	.83	.74	.75	.78	1.41	1.28	1.12	.97	1,351
Iowa	.63	.51	.64	.94	1.64	1.06	.52	−.00	1,206
Missouri	.87	.79	.86	1.00	.77	.71	.65	.59	1,567
Virginia	.72	.74	.78	.80	2.09	1.68	1.29	.92	1,817
W. Virginia	1.35	1.05	1.15	1.44	1.32	.97	.60	.29	1,244
N. Carolina	.80	.80	.85	.90	1.59	1.33	1.08	.84	3,059
S. Carolina	1.02	.92	.97	1.11	2.18	1.64	1.08	.55	1,433
Georgia	.83	.92	.92	.84	1.59	1.42	1.29	1.13	2,054
Florida	.85	1.08	1.06	.83	2.45	2.17	1.94	1.72	1,533
Kentucky	.96	.79	.89	1.15	2.09	1.42	.73	.13	1,799
Tennessee	.87	.80	.96	1.23	1.82	1.28	.78	.28	1,818
Oklahoma	.98	.87	1.10	1.52	1.73	1.33	.95	.59	1,556
Mississippi	2.11	1.57	1.25	1.00	2.09	1.64	1.08	.55	1,328
Louisiana	1.30	1.11	1.06	1.05	1.32	1.55	1.77	2.02	1,319
Texas	.78	.75	.74	.70	1.68	1.90	2.16	2.39	2,896
California	1.35	1.28	1.15	.93	1.23	1.37	1.51	1.60	2,060
United States	4.6	6.1	7.2	8.1	2.20	2.26	2.32	2.38	59,495

[a] Numbers are ratios of the predicted state unemployment rate to the predicted national unemployment rate. Annual unemployment rates for each state and for the U.S. were regressed on quadratic trend, 1968–82, to obtain the predictions.
[b] Numbers are ratios of the predicted state employment growth rate to the predicted national employment growth rate. The logarithm of annual employment in the private sector for each state and the U.S. were regressed on quadratic trend, 1968–82. The predicted employment growth is the difference between the predicted logarithm of employment in two adjacent years.

fornia are uniformly larger than average, and for New York, Pennsylvania, Ohio, Illinois, and Missouri, employment growth is uniformly below the national average. During this period, some states (i.e., Iowa, West Virginia, South Carolina, Kentucky, Tennessee, Alabama, and Mississippi) went from having far above- to far below-average employment growth rates. In general, these data show a very unequal geographical distribution of employment growth 1970–82.

The geographical distribution of labor demand "shocks" is also unequal. Table 2 presents indexes of local labor demand shocks for the subperiod 1978–82. These measures were constructed as follows. First, the natural logarithm of annual private employment (1968–82) was regressed on quadratic trend. The residuals from these regressions, ϵ_t^s, are indexes of time-varying local demand conditions in state s and year t. Second, the natural logarithm of the national aggregate employment was regressed on quadratic trend. The residuals from this regression, ϵ_t, measure the aggregate labor demand disturbance in year t. Relative local labor demand disturbances in year t and state s are then defined as $\eta_t^s = \epsilon_t^s - \epsilon_t$, which expresses the current local labor demand "shock" as a deviation from the aggregate labor demand "shock." (This measure is used by Topel, p. S129).

These measures of local labor demand disturbances have two important features. First, for a given year, substantial geographical variation in employment disturbances occurs. They also are largely unrelated to the aggregate business cycle. Second, for a given state the successive disturbances occur as local cycles. Shocks with the same sign tend to persist for a couple of years then to reverse themselves.

When workers and firms are immobile, local economic conditions will affect real wage rates. Although workers are largely immobile in the short run, they are geographically mobile in the long run. However, differences among localities in their net advantages to firms (industries) and in their amenities and cost of living characteristics to workers contribute to permanent differences in labor markets of different localities.

Table 2. Relative Employment Disturbance, Twenty-three Selected States, 1978–82

States	1978	1979	1980	1981	1982
New York	−2.25	−2.24	−0.97	0.69	2.94
Pennsylvania	−0.40	−0.35	−0.31	0.25	0.32
Ohio	0.67	0.98	−0.34	−0.13	−0.31
Indiana	2.58	2.22	−1.21	−0.49	−1.28
Illinois	0.90	0.71	0.81	−1.05	−0.90
Michigan	4.12	3.05	−0.57	−1.25	−2.16
Wisconsin	0.60	1.42	0.33	−0.62	−0.77
Minnesota	0.66	1.48	1.00	−0.28	−1.56
Iowa	1.89	0.95	0.11	−0.53	−1.87
Missouri	1.22	1.46	−0.37	−0.98	−0.37
Virginia	−0.52	−1.41	0.49	−0.15	0.34
W. Virginia	0.41	1.98	0.62	−1.40	−1.23
N. Carolina	−0.17	0.42	0.12	0.03	−0.17
S. Carolina	−0.72	−0.67	0.06	0.57	0.47
Georgia	−0.43	−0.80	−0.49	0.11	1.65
Florida	−3.33	−2.85	0.14	1.88	2.55
Kentucky	1.01	1.29	−0.86	−1.13	−0.01
Tennessee	0.87	0.44	−0.77	0.39	−0.18
Oklahoma	−1.50	−2.56	−1.15	0.75	2.00
Mississippi	1.95	1.49	0.26	−0.97	−1.34
Louisiana	0.92	−1.03	0.23	0.40	−1.29
Texas	−1.09	−1.42	−0.72	0.76	0.68
California	0.67	0.85	0.84	0.12	−1.31
United States	2.39	3.59	1.88	0.34	−3.77

Note: The numbers are measures of $\eta_t^s = \epsilon_t^s - \epsilon_t$ except for the last row. See the text for details.

The Economic Model

In a labor market workers simultaneously sell their services and buy the attributes of jobs, including location. Employers also buy the services and characteristics of workers and sell job attributes. Therefore, the theory of equalizing wage differentials stresses both supply and demand for labor, and the market equilibrium process allocates or assigns specific workers to specific firms and locations (Rosen).

Labor Demand

The actual wage can be viewed as the summation of two distinct transactions, one for labor services and worker characteristics and another for job and locational attributes (Rosen). Localities are assumed heterogenous over characteristics that matter for optimal firm-industry (employer) and household (worker) decisions. This environment contains spatially related competitive labor markets with equilibrium wage differentials. A wage premium is paid by employers to induce workers to undertake undesirable tasks or work at an undesirable location. Unequal transitory shocks to labor demand in different localities are an added source of geographical differences in wage rates.

The wage elasticity of aggregate labor demand for a locality is negative, but individuals face a perfectly elastic demand for their labor. Our labor demand or wage rates at a given location and for a given sex are assumed to depend on skill or human capital (ζ) and job or locational characteristics—permanent or anticipated labor market conditions (Ω), transitory or unanticipated labor market conditions (ω), local cost of living (ψ), and locational amenities (Δ). This relationship is summarized as

$$W^j = W^j(\zeta^j, \Omega, \omega, \psi, \Delta), \quad (1)$$

where $j = M$(males), F(females).

Permanent or anticipated labor market conditions refer to important conditions that local firms and workers know and include in rational decision making. These variables include the anticipated or long-run unemployment rate, anticipated rate of job growth, and anticipated change in the occupational composition of employment. When anticipated unemployment harms workers (households) more than firms (employers), firms or localities with higher anticipated unemployment rates will have higher equilibrium wage rates than other firms or localities (Hall, Abowd and Ashenfelter, Adams, Topel). That is, higher wage rates compensate for higher probabilities of unemployment. When households (workers) bear most of the cost of geographical and occupational mobility, geographic differences in the anticipated rate of employment growth and in change of occupational composition of employment are also the source of compensating geographical wage differentials (Lilien, Adams, Topel). Thus, positive wage differences provide economic returns to households for geographical or occupational mobility.

Locational differences in the cost of living can be decomposed into effects of prices of goods traded among locations and of nontraded goods and services (Tolley, Kenny and Denslow). With competitive markets, the prices of traded goods in two areas differ by transport costs so that they likely are an insignificant source of cost of living differences. However, prices of nontraded goods and services differ more between two localities. The price of housing-plus-access is a good example for the nontraded good. Furthermore, because housing costs are about 14% to 15% of workers' household expenditures, locational differences in these prices are a significant source of interlocality wage differences.

Areas differ in the quantity and quality of locational amenities. Normal climatic conditions, e.g., average January and July temperature, are characteristics of the local environment that likely matter to households (workers) and possibly to firms. Other studies (e.g., Israeli, Hock and Drake, Kenny and Denslow) have found significant effects of climatic conditions on wage rates.

Labor Supply and Labor Force Participation

Households make labor supply and labor force participation decisions of their members. In this analysis, the focus is on single-family, husband-wife households, where the decisions are for one period, and households are risk neutral about uncertain outcomes including employment or unemployment. This means the expected wage is used in resource allocation decisions.

Participation decisions of farm and rural nonfarm households are analyzed. The nonfarm households are assumed to have only wage and asset income; farm households have asset income, self-employment income from their farm business, and perhaps off-farm wage income. Thus, the wage-labor participation decisions of farm households are more complex than for rural nonfarm households (Huffman and Lange, Strauss).

The economic decision-making framework of these households is summarized in equations (2)–(5).

$$(2)\qquad U = U(T_h^M, T_h^F, Y; \zeta^M, \zeta^F, \Delta, \tau)$$

$$(3n)\qquad \bar{T} = T_m^j + T_h^j, T_m^j \geq 0, j = M, F$$

$$(3a)\qquad \bar{T} = T_f^j + T_m^j + T_h^j, T_m^j \geq 0, j = M, F$$

$$(4n)\qquad (1 - u^M)\, W_0^M T_m^M + (1 - u^F)\, W_0^F T_m^F + V = P_y Y$$

$$(4a)\qquad (1 - u^M)\, W_0^M T_m^M + (1 - u^F) W_0^F T_m^F + V + P_Q Q - W_x X + V = P_y Y$$

$$(5a)\qquad Q = Q(T_f^M, T_f^F, X; \zeta^M, \zeta^F, \Delta).$$

Notation "*n*" and "*a*" after the equation number refer to nonfarm and farm households, respectively; other equations refer to both types of households.

Farm and nonfarm households derive utility from the leisure time of the husband and wife (T_h^j) and from goods purchased in the market (Y). Household utility also depends on husband's and wife's human capital (ζ^j), local climate (Δ), and other household characteristics (τ), e.g., number of children in the household and commuting distance to service centers.

Farm and nonfarm households receive human time endowments each year ($\bar{T}$) for the husband and wife, which are considered heterogenous. In nonfarm households, the time of each adult is allocated between work for a wage (T_m^j) and leisure (T_h^j). In farm households, time is allocated among work on their own farm (T_f^j), work for a wage (off farm) (T_m^j), and leisure (T_h^j). In farm and nonfarm households, optimal hours of wage work might be zero in any year. Hence, a nonnegativity constraint is imposed on wage work ($T_m^j \geq 0$).

The cash income of farm and rural nonfarm households is uncertain because of uncertain employment prospects. Thus, the wage rate, given employment, is adjusted for the expected probability of unemployment (u^j). These sex-specific unemployment rates are exogenously determined for the households. Furthermore, firms and households do not know the unanticipated parts of employment growth or unemployment rates. W_o^j is the anticipated wage, given unemployment, when unanticipated local labor disturbances are zero, i.e., $\omega = 0$ in equation (1). Expected household wage income is $(1 - u^M)\, W_o^M T_m^M + (1 - u^F) W_o^F T_m^F$. Household asset income is V. Farm households also have uncertain self-employment or net income from a farm business ($P_Q Q - W_x X$). Farm output (Q) is produced by inputs of husband's and wife's farm hours (T_f^j) and by purchased inputs (X). The efficiency of the production process is affected by human capital of the husband and wife (ζ^j) and climate (ϕ).[3]

Households are assumed to face a perfectly elastic supply of the consumption good (Y) at a price P_y. Farm households also face a perfectly elastic supply of inputs at price W_x and perfectly elastic demand for farm output. However, the price of farm output is uncertain when production plans are made, and P_Q denotes the expected price.

For rural nonfarm households, wage labor supply functions are obtained by maximizing (2) subject to (3n) and (4n). For farm households, wage labor supply functions are obtained by maximizing (2) subject to (3a), (4a), and (5a). These wage labor supply equations are:

$$(6n)\qquad T_m^j = S_m^j[(1 - u_o^j)W_o^j, (1 - u^k)W_o^k, P_y, V, \zeta^M, \zeta^F, \Delta, \tau], T_m^k > 0,$$
$$S_m^j[(1 - u^j)W_o^j, P_y, V, \zeta^M, \zeta^F, \Delta, \tau], T_m^k = 0, \quad j, k = M, F; j \neq k.$$

$$(6a)\qquad T_m^j = S_m[(1 - u^j)W_o^j, (1 - u^k)W^k, P_y, V, P_Q, W_X\, \zeta^M, \zeta^F, \Delta, \tau, \phi], T_m^k > 0;$$
$$S_m[(1 - u^k)\, W_o^j, P_X, V, P_Q, W_X, \zeta^M, \zeta^F, \Delta, \tau, \phi], T_m^k = 0, \quad j, k = M, F; j \neq k.$$

Equations (6n) and (6a) show that wage labor supply functions for a given individual have different structures (Huffman and Lange). When both married nonfarm individuals work for a wage, their wage labor supply is a function of their expected anticipated wage rates, price of consumption goods, asset income, their human capital stocks, the local climate, and tastes. For farm households, the expected price of farm output, the price of purchased farm inputs and a technology parameter also determine off-farm wage labor supply.

Wage-labor participation decisions of an in-

[3] The effect of a tax on income is excluded from the economic (and econometric) model. This simplification is unlikely to be of major consequence for the empirical results (Mroz).

dividual are modeled as a comparison of his (her) reservation wage and anticipated market wage. The reservation wage equation is derived from the wage-labor supply equation by setting wage work hours equal to zero:

$$(7n)\quad W_r^j = \left(\frac{1}{1-u^j}\right)$$
$$G_r^j[P_y, V, \zeta^M, \zeta^F, \Delta, \tau, (1-u^k), \Omega, \omega = 0, \psi], \quad j, k = M, F; j \neq k.$$

$$(7a)\quad W_r^j = \left(\frac{1}{1-u^j}\right)$$
$$G_r^j[P_y, V, P_Q, W_X, \zeta^M, \zeta^F, \Delta, \tau, \phi, (1-u^k), \Omega, \omega = 0, \psi], \quad j, k = M, F; j \neq k.$$

A nonfarm household member participates in wage work when his (her) reservation wage is less than the anticipated wage in the market.[4] A farm household member participates in nonfarm wage work when the marginal value of his (her) leisure and (or) farm work hours are less than the anticipated nonfarm wage.

The probability of wage work can be expressed as the probability that an individual's reservation wage is less than his (her) anticipated market wage. For the ith household and jth married individual, define

$$D_i^j = \begin{cases} 1 \text{ if } j\text{th individual works for a wage} \\ 0 \text{ otherwise} \end{cases}, \quad j = M, F;$$

then the probability of wage work for the ith individual is

$$(8)\quad P_r\{D_i^j = 1\} = F\{W_{ri}^j < W_{oi}^j]; j = M, F, \text{ or}$$

$$(8a)\quad P_r\{D_i^j = 1\} = F[P_y, V, \zeta^M, \zeta^F, \Delta, \tau, \Omega, (1-u^j), (1-u^k)], \quad j, k = M, F; j \neq k, \text{ and}$$

$$(8n)\quad P_r\{D = 1\} = F[P_y, V, P_0, W_X, \phi, \zeta^M, \zeta^F, \Delta, \tau, \Omega, (1-u^j), (1-u^k)], \quad j, k, = M, F; j \neq k,$$

where $F(\ )$ is a distribution function. Variables that explain the probability of wage work enter an individual's labor demand and labor supply functions, except for the individual's anticipated wage rate. When the labor supply schedule has a positive slope, variables that cause the labor supply curve to shift to the left will increase the reservation wage and reduce the probability of wage work. A change in a variable that raises the market wage—raises the labor demand curve—will increase the probability of wage work.

Selected variables are examined for their effects on the probability of wage work. An increase of an individual's schooling will increase the anticipated market wage and reservation wage. The net effect on the probability of wage work is a priori ambiguous, but other studies (e.g., Heckman and MaCurdy 1980, 1982; Huffman and Lange) have found a strong positive effect of an individual's schooling on the probability of wage work for married farm and nonfarm males and femalels in the United States.

When leisure is a normal good, a higher expected price of farm output reduces the probability of wage work by couples who operate a farm business. The reason is that the quantity demanded of husband's and wife's farm labor increases, provided their farm hours are a normal input and the quantity of their leisure demanded increases due to the increase of expected profit. Higher farm input prices change the probability of wage work in an a priori ambiguous direction. Hired (nonfamily) labor and family labor are heterogenous. If hired farm labor and husband's and wife's farm labor are gross substitutes, then a higher wage for hired labor will cause a rightward shift of the demand for husband's and wife's farm labor. A higher input price reduces expected farm profit and reduces the demand for leisure of the husband and wife. Thus, the net effect on the reservation wage and probability of wage work is a priori ambiguous. A change in the price of other inputs yields a similar conclusion.

The effects of local labor market conditions depend partially upon expectations. If the demand for leisure increases as the expected wage decreases, a higher expected local unemployment rate will raise an individuals' reservation wage. When anticipated unemployment hurts workers more than firms, firms or localities that have higher expected unemployment rates will pay higher wage rates. Thus, the net effect of anticipated unemployment on the probability of wage work depends on which of these changes is largest. Localities having more rapid anticipated job growth are expected to pay higher wage rates than other areas and hence will increase the probability of wage work. On the other hand,

[4] Alternatively, decisions can be made by comparing indirect utility functions associated with different outcomes.

unanticipated changes in local labor market conditions are not expected to affect the probability of wage work. These disturbances can at best be poorly forecasted and at worst cannot be forecasted. The anticipated wage rates are higher in these localities, while the reservation wage likely is unaffected.

The Data and Econometric Model

The econometric model incorporates effects of local economic conditions on wage labor demand and probability of wage work of married rural males and females. Focusing on males and females provides a better picture of the total effect on households and on differences resulting from gender. Farm and rural nonfarm households from the Current Population Surveys (CPS) and having economic activity in 1978, 1979, 1981, and 1982 are the decision units to be studied.

The Data

The CPS is a monthly survey containing information about the employment status of members of approximately 60,000 interviewed households residing in every state of the United States. The annual demographic file (March) contains information on labor force participation, employment, and earnings of household members during the calendar year before the survey. Starting in 1977, the state of residence and farm-nonfarm residence of each household are identified.

One sample consists of nonmetropolitan-nonfarm households where the husband and wife are present and no self-employment income was received. This is the closest definition to rural nonfarm because households having rural nonfarm residence are not identified in the CPS. This sample comprises "rural nonfarm" wage earning households. The other sample consists of husband-wife households that have a farm residence and self-employed income from farming. In both samples, households having a residence in the contiguous forty-eight states, except for New England, are included. Households in New England were excluded because of the lack of importance of agriculture in those states. These two samples from the forty-two states consist of about 8,115 rural nonfarm households and 1,466 farm households per year.

Local markets are defined as state units in this study. A state is the smallest geographic unit in which a CPS household can be identified. States are also the smallest political-economic-geographic unit for which annual data are collected on employment, unemployment, and agricultural prices. In addition, government programs frequently target state units. Other recent studies (e.g., Adams, Topel) have used state units as labor markets.

The years 1978–82 give variations in local economic conditions that affect wage and participation rates. This period was chosen for several reasons. The CPS first identified the farm-nonfarm residence of households in 1977. Second, the period 1975–79 is the trough-to-peak part of a national business-cycle expansion (Executive Office of the President). The national average unemployment declined from 8.3% in 1975, to 5.8% in 1979. Net farm income was relatively good in the late 1970s. The period starting in 1980 contained a business-cycle contraction. The national unemployment rate rose from 7% in 1980 to 9.5% in 1982–83. The sharp rise of real interest rates and fall in the value of the U.S. dollar contributed to the drop in net farm income during 1981 and 1982. Although the depression of the farm economy continued after 1982, extending the analysis through 1983 did not seem wise because the first large government payment-in-kind (PIK) programs occurred in 1983. Third, the data in table 2 show relatively large geographical variation in shocks to labor demand during this period. Fourth, Lilien has shown that employment growth was unequally distributed across U.S. industries in the 1970s. Significant shifts of the occupational-industrial mix of employment occurred, especially a rise in the share of employment in services, finance, insurance and real estate but a decrease in manufacturing.

When the four cross-sectional files are combined, the data on forty-two states and four years gives 168 potentially distinct observations on different local economic conditions. Although data for four years are used, it is not a panel consisting of the same households. About 25% of the CPS households in any year are replaced.

Empirical Definitions

Definitions and sample mean values of the variables are presented in table 3. More details are presented below on the derivation of selected variables.

Average hourly wage rates are derived for ru-

Table 3. Variable Names and Sample Means, Rural Married Couples 1978–79, 1981–82

Symbol	Variable Description	Sample Mean	
		Nonfarm	Farm
Individual/household			
AGEM	Husband's age (yrs)	47.0	50.5
AGEF	Wife's age (yrs)	43.9	47.2
EDM	Husband's schooling (yrs)	11.5	11.3
EDF	Wife's schooling (yrs)	11.6	11.8
RACE	1 if nonwhite; 0 otherwise	.07	.03
KIDS06	Number of children under age 6	.31	.27
KIDS618	Number of children ages 6–18	.66	.69
ASSETINC	Household real nonwage and nonfarm income (1967 prices)[a]	$10,386[a]	$10,701[a]
Local labor market conditions			
PJOBGR	Predicted state employment growth rate (see text)	2.17	2.06
PURATE	Predicted state unemployment rate (see text)	7.46	7.18
ΔSHRSER	Change of share of a state's jobs in serv. occupation (previous 2 yrs.)	1.02	.91
ESHOCK	Relative state employment growth shock (see text)	.08	.15
RURATE	Residual state unemployment rate (see text)	−.37	−.38
Cost of living and locational amenities			
PLAND	State average price of agricultural land in 1978 ($1,000/acres)	.91[a]	.79[a]
URBAN	Percentage of state population urban	.68	.66
JAN	Normal January ave. temperature (degree F)	34.1	29.7
JULY	Normal July ave. temperature (degree F)	75.9	75.1
Agricultural prices and climate			
PCROP	State real price index for crops (1967 prices)		.444[a]
PLIVE	State real price index for livestock (1967 prices)		.517[a]
FARMWAGE	State real wage rate for hired farm labor (1967 prices)		.504[a]
POTINP	State real price index for nonlabor farm input (1967 prices)		.536[a]
RAIN	State annual average precipitation (inches)		35.7
GDD	State average growing season length—growing degree days (1,0005)		3.336
Regional dummies and trend			
NC	1 for resident in North Central Region; 0 otherwise	.28	.46
SOUTH	1 for resident in South; 0 otherwise	.51	.38
WEST	1 for resident in West; 0 otherwise	.06	.12
TIME	Trend	3.00	3.00
Dependent variables			
WAGEM	Real married male nonfarm wage ($/hr, 1967 prices)	$2.87[a]	
WAGEF	Real married female nonfarm wage ($/hr, 1967 prices)	$1.66[a]	
D^M	1 if husband works for wage; 0 otherwise	.75	.43
D^F	1 if wife works for a wage; 0 otherwise	.54	.39

[a] Geometric mean. All other numbers are arithmetic means.

ral nonfarm married males and females. For individuals in the rural nonfarm wage work households, the average wage of an individual is his (her) wages and salaries for the year preceding the survey divided by the product of his (her) hours worked per week last year and weeks worked last year. An accurate measure of the average wage rate cannot be computed for individuals in a household with self-employment income because hours worked include hours at all jobs, both wage work and self-employment. Thus, average wage rates are not available for CPS farm household members. Nominal wage rates are deflated by the consumer price index.

Five measures of local labor market conditions are derived. The predicted state employment growth rate ($PJOBGR_t$) is the difference in forecasted values of the natural logarithm of a state's private sector employment in t and $t - 1$. The forecasts were obtained from a regression of the natural logarithm of employment, 1968–82, on a quadratic trend. A state's unemployment rate was measured for all private sector employees rather than having separate rates for males and females. The predicted state unemployment rate ($PURATE_t$) measures the anticipated local unemployment rate. It is obtained from a regression of a state's annual unemployment rate, 1968–82, on quadratic trend. The change in the share of a state's employment that is in the service sector ($\Delta SHRSER_t$) indicates changes in the occupational mix of local labor demand. It is defined as the share in t minus the share in $t - 2$. Service jobs include employment in services, transportation, government, finance, and wholesale and retail trade. The derivation of the relative shocks to labor demand in state labor markets ($ESHOCK_t$) is described above. In this study, these shocks are unanticipated by firms and workers. The residual or unanticipated unemployment rates ($RURATE_t$) is the actual unemployment rate in t minus the predicted unemployment rate for t.

Geographical differences in the cost of living and locational amenities are tied to cost of housing-plus-access and climate. The price of land is a major part of the cost of housing-plus-access. For households living in rural areas, the base price of land is represented by the average agricultural land price in 1978 (U.S. Dep. of Commerce 1980). In addition, the price of housing-plus-access increases as the percentage of the population living in urban areas increases because the cost of land plus commuting is larger in urban areas. When land prices and wage rates are positively associated and housing and husband's and wife's leisure are substitutes, an increase of *PLAND* will increase the reservation wage. This makes the effect of *PLAND* on the probability of wage work a priori uncertain. However, if housing and husband's and wife's leisure are complements, their reservation wage rates will decline. The percentage of the population living in urban areas in 1980 (U.S. Dep. of Commerce 1981) is used as a second proxy for the cost of housing-plus-access. For locational amenities, thirty-year (1950–80) normal average January and average July temperatures (Weiss, Whittington, and Teigen) were used. Kenny and Denslow and Hoch and Drake found nonlinear effects of these temperatures on $\log_e$ wage rates in earlier studies.

The profitability of local agriculture is represented by indexes of agricultural prices and agricultural climate. Indexes of crop prices, livestock prices, wage for farm labor, and prices of other inputs are derived and deflated by the consumer price index. Output price indexes for crops and livestock are constructed because the average labor intensity is significantly different for these output groups (Huffman and Evenson, chap. 10). The crop price index is composed of prices of twenty-six different commodities or commodity groups. The livestock price index is composed of seven commodity groups. The expected prices used to derive the Fisher-type output price indexes are primarily one-year lagged prices of the commodities (Huffman and Evenson, chap. 10).

Input prices are split into two groups, farm labor and other inputs. The price of farm labor is the hourly wage paid to employees working for cash wages only. The nonlabor input prices include fertilizer, feed, capital, seed, land, and miscellaneous inputs. Current prices are used to derive Fisher-type price indexes of these inputs (Huffman and Evenson, chap. 10). Both output and input price indexes are deflated by the consumer price index.

Normal annual precipitation and normal growing-degree-days are important climatic variables for agricultural production. Normal annual rainfall is a twenty-five-year average of annual precipitation (Weiss, Whittington, and Teigen). Natural precipitation is the primary source of water for much of U.S. agriculture. However, in low precipitation areas, irrigation is a costly substitute. Accumulated growing degree days, *GDD*, is a measure of accumulated heat units from a temperature range that is particularly favorable to corn production (U.S. Dep. Agriculture and U.S. Dep. Commerce; U.S. Dep.

Commerce 1971, 1981). Corn is grown in almost every state, but more generally the index is highly correlated with good growing conditions for warm-season crops.

Household asset income reflects earnings from interest, dividends, and rental property deflated by the consumer price index. Thus, all income- and work-conditioned transfers are excluded.

The Econometric Model

The econometric model consists of two labor demand equations and two wage-participation equations. The empirical specification of the labor demand equations is similar for married males and females:

$$(9) \quad \ln WAGE_i^j = \alpha_1^j + \alpha_2^j EXP_i + \alpha_3^j EXP_i^2 + \alpha_4^j ED_i + \alpha_5^j RACE_i + \alpha_6^j PJOBGR_i + \alpha_7^j PURATE_i + \alpha_8^j \Delta SHRSER_i + \alpha_9^j ESHOCK_i + \alpha_{10}^j RURATE_i + \alpha_{11}^j \ln PLAND_i + \alpha_{12}^j URBAN_i + \alpha_{13}^j JAN_i + \alpha_{14}^j JAN_i^2 + \alpha_{15}^j JULY_i + \alpha_{16}^j JULY_i^2 + \alpha_{17}^j NC_i + \alpha_{18}^j SOUTH_i + \alpha_{19}^j WEST_i + \alpha_{20}^j TIME_i + \alpha_{21}^j \lambda_i^j + \epsilon_i^j,\ i = 1, \ldots, n, j = M, F.$$

amenities proxy relatively well the sources of geographical wage differentials, then the regional dummy variables will not make a statistically significant contribution to real wage rates. When women spend more time out of the labor force and are more geographically immobile than men, women's wage rates likely will be less responsive to their measured human capital and to local economic conditions (Mincer). The coefficients of variables in the female wage equation then will be smaller than in the male equation.

Sample selectivity and autocorrelation are potential problems in the wage equation. If a sample selection variable was not included in equation (9), the disturbance term of the wage equation would have a nonzero mean because the equation is fitted to a nonrandom subset of the total population (Heckman). This is a potential source of statistical bias in estimated coefficients. With a selection variable included, the disturbance has a zero mean and a normal distribution.

The natural logarithm of an individual's real wage is expressed as a function of his (her) own human characteristics—experience, experience squared, education, race—and job/local conditions that are potential sources of geographical wage differentials. The last group of variables includes sets of variables for local labor market conditions, cost of living and locational amenities, and regional dummy variables. A time trend and sample selectivity variables (λ_i^j) are also included in each equation.

Equation (9) is quadratic in experience and in January and July temperatures but not in other variables. This choice was largely based on evidence reported in other studies (e.g., Adams, Topel, Kenny and Denslow) and preliminary fits of the wage equations. Experience is defined as an individuals' age minus years of schooling completed minus 6. This measure of experience is a reasonable proxy for useful work experience of males but less so for females engaged in household activities. It is less endogenous to current labor market decisions than actual experience (Heckman and MaCurdy). When a sample selection variable is included in the wage equation, sample selection bias is unlikely to be associated with the use of a work experience variable (Mroz).

The regional dummy variables may contain redundant information about labor demand. If the set of variables representing local labor market conditions and cost of living and locational

The data files consist of two adjacent-year cross sections (1978–79, 1981–82) that are separated by one year. Some autocorrelation might be expected in adjacent year observations on the same individual, but observations that are one or more years apart are less likely to be correlated. This short, disjointed time series, where the composition of the sample changes over time, is not conducive to correction for autocorrelation. Failure to correct for autocorrelation when it is present results in some loss of estimation efficiency, but the least-squares estimator is unbiased or consistent (Johnston). Contemporaneous correlation of disturbances in the two wage equations might occur. However, because the observations in the wage equations of the married males and females are not equal, the two wage equations of the married males and females are not equal, the two wage equations cannot be estimated jointly in the seemingly unrelated regression model. Ignoring contemporaneous cross-equation correlation of disturbances results in some loss of estimation efficiency (Johnston).

The empirical specification of the probability of the *j*th married individual in the *i*th farm household participating in wage work is

(10a)

$$\Pr(D_i^j = 1) = F[\beta_1^j + \beta_2^j AGEM_i + \beta_3^j AGEM_i^2 + \beta_4^j EDM_i + \beta_5^j EDF_i + \beta_6^j RACE_i + \beta_7^j KIDS06_i + \beta_8^j KIDS618_i + \beta_9^j \ln ASSETING_i + \beta_{10}^j PJOBGR_i + \beta_{11}^j PURATE_i + \beta_{12}^j \Delta SHRSER_i + \beta_{13}^j ESHOCK_i + \beta_{14}^j RURATE_i + \beta_{15}^j \ln PLAND_i + \beta_{16}^j URBAN_i + \beta_{17}^j JAN_i + \beta_{18}^j JAN_i^2 + \beta_{19}^j JULY_i + \beta_{20}^j JULY_i^2 + \beta_{21}^j \ln PCROP_i + \beta_{22}^j \ln PLIVE_i + \beta_{23}^j \ln FARMWAGE_i + \beta_{24}^j \ln POTINP_i + \beta_{25}^j GDD_i + \beta_{26}^j RAIN_i + \beta_{27}^j RAIN_i \bullet GDD_i + \beta_{28}^j NC_i + \beta_{29}^j SOUTH_i + \beta_{30}^j WEST_i + \beta_{31}^j TIME_i] \; j = M, F,$$

where $F(\;)$ is the normal distribution function. The participation equations for rural nonfarm household members are similar, except that prices of agricultural outputs and inputs, annual precipitation, and the length of growing season are excluded. This probability is a function of a set of variables representing an individual's own characteristics, spouse characteristics, household characteristics, and local conditions.[5] A time trend is also included.

Equation (10) is a reduced-form specification, including only variables in the labor supply or labor demand functions (Huffman and Lange). The husband's age (and age squared) controls for nonlinear life-cycle and work-experience effects when education is held constant. Age and experience are highly correlated, so experience is not included as a separate variable. Because husband's and wife's ages are also highly correlated, only the husband's age is used for explaining both the husband's and wife's probability of wage work.

Given that participation decisions of a husband and wife in farm and nonfarm households are assumed to be joint within a household-optimizing framework, the probability of a married individual participating in wage work is affected by some of their spouse's characteristics. Also, these decisions are affected by random or unmeasured shocks to labor supply (reservation wage) and labor demand. These shocks likely are across spouses (Huffman and Lange). Thus, the estimation procedure for equation (10) is bivariate probit. If a husband and wife are affected similarly by a given shock, the correlation between these disturbances will be positive, and vice versa.

The Results

Results from testing hypotheses about the effects of local economic conditions and other variables on labor demand for rural nonfarm married males and females and on the probability of wage labor participation of married males and females in farm and rural nonfarm households are reported in this section.

Labor Demand

Labor demand equations are fitted to the data for 24,571 married rural nonfarm males and 17,508 married rural nonfarm females. These equations (9) are fitted by least squares with an instrumental variable included to control for sample selectivity. Conclusions reached about labor demand for married rural nonfarm males and females likely are applicable to married farm males and females.

Estimates of four labor demand functions are reported in table 4; two for males and two for females. The second equation excludes the regional dummy variables. A test of the null hypothesis that the coefficients of the three regional dummy variables are jointly equal to zero is rejected for males but not for females. The sample value of the F-statistic is 5.08 for males and 1.90 for females, and the critical value of the F-statistic with 3 and infinite degrees of freedom is 3.79 at the 1% significance level. Thus, for males the eleven variables representing local labor market conditions, cost of living differences, and location amenities do not capture all of the geographical differences in wage rates. Furthermore, the long history of wages being lower in the South in no longer supported. Male wages are higher in the South than in other regions, and female wages are not lower.

In the wage equations, all of the coefficients of the human characteristics have expected signs and are significantly different from zero at the 1% level. An increase of an individual's experience has first a positive but diminishing marginal effect on the real wage. The maximum effect occurs at 36.9 years for males and 31.5 years for females. This pattern has been reported in many studies. However, the $\log_e$ wage-experience relationship is more convex for males than

[5] Although the number of children at home and the amount of household asset income is endogenous in a lifetime planning horizon, Mroz found that these variables were not endogenous in his tests. They are exogenous in our study.

Table 4. Labor Demand Equations: Rural Nonfarm Married Males and Females, 1978–79, 1981–82

	ln Wage			
	Males		Females	
Variables	(1)	(2)	(3)	(4)
Human capital				
EXP (AGE-ED-6)	.031	.031	.017	.017
	(20.04)	(19.95)	(12.88)	(12.93)
$EXP^2/100$	−.042	−.041	−.027	−.027
	(10.30)	(10.20)	(9.35)	(9.38)
ED	.055	.055	.071	.071
	(40.21)	(40.12)	(30.12)	(30.19)
RACE	−.204	−.203	−.065	−.064
	(13.72)	(13.69)	(3.48)	(3.43)
Labor market conditions				
PJOBGR	.016	.024	.009	.011
	(2.68)	(4.89)	(1.19)	(1.73)
PURATE	.012	.011	.004	.003
	(4.21)	(4.13)	(1.16)	(0.76)
ΔSHRSER	.005	.005	.002	.002
	(1.91)	(1.97)	(0.72)	(0.67)
ESHOCK	.005	.004	.005	.004
	(1.79)	(1.45)	(1.31)	(1.07)
RURATE	−.006	−.004	−.011	−.010
	(1.15)	(0.74)	(1.68)	(1.51)
Cost of living and locational amenities				
ln *PLAND*	.073	.060	.053	.056
	(5.46)	(5.47)	(3.02)	(3.87)
URBAN	.255	.180	.011	.012
	(5.77)	(5.33)	(0.20)	(0.26)
JAN	.003	.008	−.002	.002
	(1.15)	(3.96)	(0.55)	(0.74)
$JAN^2/100$	−.001	−.014	.002	−.001
	(3.91)	(5.29)	(0.57)	(0.32)
JULY	−.087	−.059	.197	.166
	(1.43)	(1.25)	(2.41)	(2.59)
$JULY^2/100$	.057	.036	−.136	−.117
	(1.41)	(1.14)	(2.49)	(2.77)
Regional dummies and trend				
NC	−.023		−.038	
	(1.37)		(1.67)	
SOUTH	.057		−.015	
	(2.65)		(0.51)	
WEST	.064		−.002	
	(2.09)		(0.05)	
TIME	−.026	−.027	−.012	−.013
	(6.29)	(6.42)	(2.22)	(2.27)
$\hat{\lambda}$	.279	.286	−.020	−.021
	(5.53)	(5.68)	(0.79)	(0.82)
Intercept	3.175	2.260	−7.625	−6.404
	(1.41)	(1.27)	(2.50)	(2.66)
R^2	.1619	.1614	.0781	.0780
N	24,571	24,571	17,508	17,508

Note: The t-ratios are conditioned on the sample selection variables.

for females. This result is consistent with married females having on average less actual labor market experience for any given measured experience than for married males (Mincer and Ofek). Males may also make larger investments in experience during their early work-life than females.

A one-year increase in schooling causes a larger percentage increase of the female than the male wage, 7.1% versus 5.5%. These relative magnitudes are consistent with some results reported in other studies, e.g., Topel for males and Gerner and Zick for females. Although the average wage for married rural nonfarm working females in the CPS sample is 57.8% of the wage for males, the marginal increase of the real wage is larger for males than for females ($.16 versus $.12 per hour).

Nonwhite rural nonfarm males earn 20% less than rural nonfarm white males, other measured variables equal, and nonwhite females earn 6.5% less than white females. Topel found an 18% difference in wages of white and nonwhite males for 1976–79. Other studies have also shown large gaps in the wages of white and black males on average but little or no gap in white and black womens' wage ratios (e.g., Hammermesh and Rees).

All of the signs of coefficients on variables representing local labor market conditions are positive, except for the unanticipated unemployment rate, in both the male and female wage equations. They are all consistent with hypotheses developed earlier. Most coefficients in the male wage equation are significantly different from zero at the 5% level. For females, the effects of the local labor market variables are smaller for anticipated variables than for males but are as large or larger for unanticipated variables. Except for the unanticipated unemployment rate, the local labor market variables are statistically weaker in the female than in the male wage equations.

The real wage rates of married males incorporate compensation for local market conditions that can be anticipated by employers and workers. The evidence, however, is weaker for females. This outcome is expected when married females spend a smaller share of their time in the labor force and are tied to their husband's locational choice (Mincer). For married males, wage rates are higher in localities with higher expected rates of employment growth and higher expected unemployment rates.

The wage premium in localities having higher expected growth of labor demand compensates males for the costs of geographical (and possibly occupational) mobility. Topel found a similar effect. Our results show that a 1% increase of the expected growth rate of local employment increases the male real wage by 1.5% to 2.5%. For females, the marginal effect is about 1%. These results imply that married rural males and females in the upper Midwest experienced significant reductions in real wage rates during the early 1980s when employment growth rates fell far below (about 2 percentage points) the national average (see table 1).

Localities with higher anticipated unemployment rates also pay higher rural wage rates for males. This result is similar to the findings of Abowd and Ashenfelter, of Adams, and of Topel. Localities (firms) with higher expected unemployment rates pay a premium to compensate male workers for bearing this risk; the evidence is much weaker for females. Again, this finding is consistent with greater immobility of married females than males. These results imply the wage premium associated with higher anticipated unemployment rates is large enough for males to keep expected real wage rates approximately unaffected.

An increase locally of the share of service jobs increases the real wage. The effect is significantly different from zero at the 5% level for males but not for females. Although service jobs span a wide range of skills from motel and restaurant staff to investment bankers, compensation is needed to create the incentives for males to invest in skills and change occupations.

Real wage rates respond to unanticipated changes in local labor market conditions in a way that is consistent with employers and employees sharing good and bad outcomes. Topel found similar effects for wage rates of nonfarm males. The coefficients of *ESHOCK* are significantly different from zero at the 7% level [table 4, equation (1)] in the male wage equation and at the 20% level in the females' wage equation. The coefficient of *RURATE* is significantly different from zero at the 10% level in the female wage equation but only at the 25% level in the male wage equation. Other studies have not considered the effect of unanticipated unemployment on the female wage. Heckman and MaCurdy found a negative effect of the current local unemployment rate on the female wage. However, Adams found a significant positive effect of unanticipated unemployment rate on the male wage. Thus, our results differ slightly from other studies for the effects of unanticipated unemployment on the female real wage.[6] This result is reasonable if married males on average have significantly larger investment in firm-specific human capital than married females. Workers having firm-specific human capital are less likely to be laid off (Becker).

Wage rates of males and females differ because of local differences in cost of living and locational amenities. Under the null hypothesis that the six coefficients of these variables are jointly equal to zero, the sample value of the F-statistic is 19.8 for males and 8.9 for females. The critical value of the F-statistic with 6 and infinite degrees of freedom at the 1% significance level is 2.1. Thus, the null hypothesis of no effect is rejected.

Cost of living differences matter. A larger land price increases significantly the wage rates of males and females. The elasticities are .060–.073 for males and .053–.056 for females. These magnitudes fall in the range predicted by Kenny and Denslow for wage adjustments to compendate for home-site cost differences when housing costs are 15% of consumption expenditures and the price of the home site accounts for 21% of the value of a home, including site. An increase of the urban share of the population increases significantly the wage rate of males. However, *URBAN* does not statistically significantly affect female wages.

Climate is a proxy for locational amenities, and the effects of *JAN* and *JULY* on wage rates are conditioned by whether the regional dummy variables are included. Because *JAN* and *JULY* are correlated with the regional dummy variables, the best measure of winter and summer temperatures on wage rates is obtained when the regional dummy variables are excluded. The effect of *JAN* on $\log_e$ male wage is quadratic and statistically significant. In the female wage equation, *JAN* and JAN^2 do not have statistically significant effects. The effects of *JULY* on $\log_e$ wage are quadratic and statistically significant for males and females. For males, an increase of *JULY* first causes a reduction of the wage but at a decreasing rate until it reaches the low point at 82°F which is the maximum observed value. This type of relationship was also found by Hoch and Drake. For females, the quadratic effect goes in the opposite direction. The peak occurs at 72°F which is near the bottom of the observed values and corresponds to average July temperatures in Pennsylvania and South Dakota. Thus, for most of the observed values of *JULY*, an increase of *JULY* reduces the wage rates of males and females.

Participation in Wage Work

The bivariate probit estimates of the equations explaining the probability of wage work for farm and rural nonfarm couples are reported in table 5. The equations were fitted to 32,662 observations on rural nonfarm households and 5,865 observations on farm households. Marginal effects of the regressors on the probability of wage work are evaluated at the sample mean and reported in table 6.

The first two columns of table 5 present results for rural nonfarm couples; the last two columns are for farm couples. The estimated cross-equation correlation coefficient of the disturbances in the participation equations is positive—0.26 for farm couples and 0.19 for rural nonfarm couples—and significantly different from zero at the 1% level. These results imply: (*a*) that the random disturbances in married male and female wage-work participation decisions are affected in the same direction by random shocks (or unmeasured effects), and (*b*) that the wage-work participation decisions of married males and females are not statistically independent.

For farm and nonfarm males and rural nonfarm females, the life-cycle effect on probability of wage work is quadratic. At young ages, a higher age increases the probability of wage work. The maximum effect occurs at age 26.2 and 33.2 for farm and rural nonfarm males, respectively, and at age 20.8 for nonfarm females. At older ages, the probability of wage work decreases as age increases. For farm females, the probability of wage work is largest at a young age and may be dominated by cohort effects. Labor force participation rates of married women in younger cohorts are significantly higher than in older cohorts (Killingsworth and Heckman).

A husband or wife who has more schooling has a higher probability of wage work. Additional schooling raises an individual's market wage by more than it raises his/her reservation wage. For nonfarm females these results for females are similar to those of Heckman and of Nakamura and Nakamura. For farm males, the

[6] The null hypothesis that the 5 coefficients of the local labor market variables in the labor demand equation for females are jointly equal to zero cannot be rejected at the 1% significance level. For equation (3), table 4, the sample value of the F-statistic is 1.14, and the critical value for 5 and infinite degrees of freedom is 3.02 at the 1% significance level.

Table 5. Bivariate Probit Estimates of Wage Labor Participation Equation for U.S. Farm and Rural Nonfarm Married Couples, 1978–79, 1981–82

	Wage Work			
	Rural Nonfarm		Farm	
Variables	Husband	Wife	Husband	Wife
Individual/household				
AGEM	.105	.027	.033	.003
	(26.82)	(7.92)	(3.78)	(0.30)
*AGEM*²/100	−.158	−.068	−.063	−.036
	(41.07)	(19.69)	(7.28)	(3.83)
EDM	.044	−.016	.010	−.010
	(11.70)	(5.43)	(1.31)	(1.38)
EDF	.007	.095	−.029	.079
	(1.71)	(25.76)	(3.31)	(9.05)
RACE	.117	.333	.350	.403
	(3.55)	(11.20)	(3.23)	(3.75)
*KIDS*06	−.026	−.497	−.028	−.386
	(1.31)	(35.55)	(0.87)	(11.41)
*KIDS*618	−.008	−.084	−.037	−.072
	(0.83)	(10.99)	(2.23)	(4.33)
ln *ASSETINC*	−.338	−.829	−.347	−.542
	(4.26)	(10.57)	(2.68)	(4.08)
Local labor market conditions				
PJOBGR	.041	.052	.075	.036
	(2.87)	(4.50)	(3.31)	(1.60)
PURATE	−.017	−.024	.050	.008
	(2.55)	(4.25)	(2.25)	(0.37)
Δ*SHRSER*	−.001	.003	−.020	−.007
	(0.21)	(0.55)	(1.23)	(0.40)
ESHOCK	.005	−.001	−.014	−.017
	(0.71)	(0.16)	(0.91)	(1.12)
RURATE	−.013	−.012	−.004	−.032
	(1.07)	(1.18)	(0.13)	(1.03)
Cost of living and locational amenities				
ln *PLAND*	.168	.040	−.066	.027
	(5.97)	(1.72)	(1.03)	(0.41)
URBAN	−.078	−.190	−.188	−.094
	(0.79)	(2.47)	(0.62)	(0.31)
JAN	.007	−.020	.053	.022
	(1.05)	(3.99)	(4.42)	(1.80)
*JAN*²/100	−.032	.019	−.077	−.010
	(4.35)	(3.08)	(3.59)	(0.45)
JULY	−.115	.498	.081	−.027
	(0.84)	(4.55)	(0.45)	(0.15)
*JULY*²/100	.082	−.321	−.091	.014
	(0.90)	(4.38)	(0.75)	(0.12)
Agricultural prices and climate				
ln *PCROP*			−.251	−.135
			(1.33)	(0.71)
ln *PLIVE*			−.488	−.054
			(1.78)	(0.19)
ln *FARMWAGE*			.890	−.291
			(3.04)	(0.98)
ln *POTINP*			−.520	−.589
			(0.84)	(0.93)
RAIN			.014	.004
			(1.19)	(0.30)
GDD/1,000			.323	−.014
			(2.82)	(0.12)
RAIN × *GDD*/1,000			−.038	−.013
			(1.61)	(0.55)

Table 5. Continued

	Wage Work			
	Rural Nonfarm		Farm	
Variables	Husband	Wife	Husband	Wife
Regional dummies and trend				
NC	.069	.004	−.102	−.156
	(1.51)	(0.12)	(0.74)	(1.15)
SOUTH	−.019	−.115	.038	−.314
	(0.39)	(2.93)	(0.24)	(2.00)
WEST	.007	−.047	−.287	−.247
	(0.10)	(0.80)	(1.61)	(1.40)
TIME	−.008	.018	.000	.007
	(0.77)	(2.09)	(0.02)	(0.28)
Intercept	6.702	−11.190	−.030	5.058
	(1.30)	(2.70)	(0.00)	(0.75)
Cross-equation correlation coeff.	.187		.262	
	(15.32)		(12.16)	
Sample size	32,662		5,866	

Table 6. Marginal Effects (Percentage Point Changes) on the Probability of Wage Work

	Rural Nonfarm		Farm	
Regressors	Husband	Wife	Husband	Wife
Individual/household				
AGEM	−.42	−.36	−.46	−.49
EDM	.41	−.16	.15	−.15
EDF	.07	.91	−.43	1.16
RACE	1.12	3.20	5.16	5.93
*KIDS*06	−.25	−4.77	−.41	−5.70
*KIDS*618	−.08	−.81	−.55	−1.06
ln *ASSETINC*	−3.25	−7.96	−5.11	−7.98
Labor market conditions				
PJOBGR	.40	.50	1.10	.53
PURATE	−.17	−.24	.73	.12
ΔSHRSER	−.01	.03	−.30	−.10
ESHOCK	.05	−.01	−.20	−.25
RURATE	−.13	−.11	−.06	−.47
Cost of living and locational amenities				
ln *PLAND*	1.61	.38	−.98	.39
URBAN	−.75	−1.82	−2.77	−1.39
JAN	−.15	−.07	.11	.24
JULY	.10	−9.46	−.81	−.09
Agricultural prices and climate				
ln *PCROP*			−3.70	−1.99
ln *PLIVE*			−7.19	−.80
ln *FARMWAGE*			13.11	−4.28
ln *POTINP*			−7.67	−8.69
RAIN			.21	.05
GDD			4.76	−.21
Regional dummies and trend				
NC	.67	.04	−1.51	−2.30
SOUTH	−.18	−1.11	.56	−4.63
WEST	.07	−.45	−4.23	−3.64
TIME	−.08	.17	.01	.10

results are similar to those of Sumner and of Huffman and Lange. The marginal effect of a year of female schooling on the probability of wage work is larger than that of male schooling (.41 versus .91 for nonfarm males and females and .15 versus .16 for farm males and females). Thus, a year of schooling seems to increase the difference between a wife's reservation and market wage relatively more than for her husband. This is consistent with results reported by Huffman and Lange for a different data set. Also, the marginal effect of schooling on the probability of wage work for a given gender is larger for rural nonfarm than for farm adults.

Although cross-person effects of education are seldom included (e.g., Heckman; Nakamura and Nakamura; Heckman and MaCurdy 1980, 1982; Sumner), negative and statistically significant effects of a farm or rural nonfarm husband's schooling on his wife's participation and of farm wife's schooling on her husband's wage work participation occur. Huffman and Lange also found a similar negative effect of a farm wife's schooling on her husband's probability of wage work. For the rural nonfarm male, additional schooling of the wife tends to increase the probability of working for a wage.

Additional children at home under age 18 have well-known and statistically significant negative effects on the probability of wage work by married females. The largest reduction occurs for additional children under age six—about 5% per child. For older children, the negative marginal effect is larger for farm than for rural nonfarm married females. The coefficients of *KIDS06* and *KIDS618* are negative in both participation equations for married males, but they are generally weaker statistically. These results imply that additional children at home raise the reservation wage of married women relatively more than the reservation wage of married men.

Although no other study has examined the effects of local labor market conditions on the probability of wage work, the effects of these variables are largely as expected. Anticipated variables have statistically stronger effects than unanticipated ones. Higher local employment growth raises the market wage of males and females, and it increases the probability of wage work. These effects are significantly different from zero at the 5% level, except for farm wives. Recall that the labor demand equations showed that a 1% rise of the expected unemployment rate caused the male wage rate to rise by 1.1% to 1.2% but the female wage rate to rise by only 0.3% to 0.4%. Thus, if households are risk neutral, the positive effects of *PURATE* on the participation of farm and nonfarm males are consistent. There is no significant effect, however, on the probability of participation of farm females. A larger *ΔSHRSER* does not significantly affect the probability of wage work of males or females, although it does increase the wage rate of males. The relative employment shocks and unanticipated unemployment rate do not have statistically significant effects either.

The positive and statistically different from zero effect of *PLAND* on the probabilities of wage work implies that market wage effects are larger than reservation wage effects. The effects of January and July temperatures on the probability of wage work are mixtures of effects through the market wage and reservation wage. Statistically significant effects on participation occur for nonfarm females and farm males. For farm males, a higher *JAN* first increases the probability of wage work; the peak effect occurs at 26°F; and then it decreases.

For farm households, the farm output and input price indexes have consistent signs, and their marginal effects are larger and stronger in the husband's participation equation. A drop of crop and livestock output prices increases the probability of wage work of husbands and wives. A higher farm wage increases the probability of off-farm wage work by husbands. Other input prices have negative coefficients, but they are not significantly different from zero. Agricultural climatic effects are important for farm males. Larger *RAIN* or *GDD* has a negative impact on labor force participation when these effects are evaluated at the sample mean.

Except for women living in the South, the regional effects are not significantly different from zero. Married women living in the South, however, have a lower probability of wage work than women in the Northeast. Other things equal, the reduction is about 1% for rural nonfarm women and 5.5% for farm women. The probability of wage work for rural nonfarm women has a positive and statistically significant trend. None of the other coefficients of *TIME* is significant.

Conclusions and Implications

This study has shown that wage premiums exist for localities anticipating labor demand growth, higher unemployment rates, larger shares of employment in services, and higher costs of living. Unanticipated negative disturbances in local labor markets associated with employment growth and unemployment reduce wage rates. Male wage rates are more responsive to anticipated changes in local labor market conditions, and female wage rates are more responsive to unanticipated changes. The differences by gender for anticipated effects reflect less time spent by females in the labor force. Women's wage rates bear a greater burden of negative unanticipated labor market outcomes because they are more geographically and occupationally immobile.

Labor force participation decisions of households are also affected by changes in anticipated local economic conditions. For farm households, the probability of wage work increases when expected farm output prices decline and decreases when local labor demand grows. Given the general trend toward larger farms, the results show that existing agriculture is significantly easier for farm households that reside in localities where nonfarm employment is expected to grow rapidly. These effects, however, pull more strongly on males than females. However, participation by farm and nonfarm couples is unaffected by unanticipated labor market outcomes.

The main policy recommendations are that (*a*) public information be made available to firms and households about anticipated changes in all local labor markets, (*b*) regionally targeted stabilization policies be employed to moderate unanticipated changes in local labor demand, and (*c*) tax credits or other incentives be given to workers to hasten their movement from slow growing regions and occupations to rapidly growing ones. Future research could extend this analysis through the late 1980s when national economic conditions and local labor fluctuations were moderate, and examine the effects of local economic conditions on hours of wage work through effects on wage rates.

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