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An econometric analysis of family labor supply decisions and household incomes: U.S. rural farm and nonfarm households, 1978-82

by

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CHAPTER I. INTRODUCTION

Rapidly changing economic and technological conditions have complicated the decision-making process of rural households. The economic environment is likely to remain unstable in the future, due to changes in business cycles, world markets, and government programs. Rural residents, who have limited job opportunities in thin labor markets, face special challenges under unstable economic conditions.

Nonmetropolitan America contains almost 25% of the nation's population and 33% of its labor force (Briggs, 1986). Rural residents are more likely to experience subemployment or poverty than their urban counterparts, and nonmetropolitan areas have lower wage rates and family income than urban areas. Rural labor markets have been affected by changes in industrial and occupational structure; the exodus of farmers out of agriculture over the last 50 years, along with the increasing participation of farmers in off-farm employment, provides evidence of changes. The strength of the rural economy is, however, vital to the welfare of millions of people.

Characteristics of Rural Labor Markets

Rural labor markets provide different job opportunities than urban labor markets, although industrial restructuring and technological change have affected both. The structure of rural labor markets has implications for labor force participation, wage rates, and household income of farm operators and rural residents.

Comparison of rural and urban labor markets

Rural labor markets do not provide the same job opportunities for potential participants as do urban labor markets. The industrial growth and composition of jobs have changed at different rates in rural and urban communities, although similar patterns have been observed. Between 1970-1977, 40% of newly-created jobs were in nonmetropolitan areas; these areas held 35% of the U.S. population. For a given occupation, rural and urban areas have experienced different rates of growth. The long-term shift from manual to nonmanual occupations took place faster in urban areas, and professional positions grew relatively faster in metropolitan areas (Tienda, 1985). Most rural job growth in the last ten years has occurred in areas close to metropolitan areas, and may be described as "spillover" from urban growth (Kale, 1986). And since 1979, nonmetropolitan unemployment rates have been higher than metropolitan rates (Pollack and Pendleton, 1986).

The industrial distribution of employed rural and urban residents is different. Table I.1 shows the industrial distribution--agricultural and nonagricultural industries--for urban, nonmetropolitan, and rural farm residents. As expected, the farm population is heavily represented in agricultural industries, but over 54% of the farm population is primarily employed in nonagricultural industries. In nonagricultural industries, urban, nonmetropolitan, and rural farm people are employed in different proportions. Over 72% of the urban population employed in nonagricultural industries in 1980 worked in service jobs, while only 64% of the rural population worked there. A larger proportion of the rural

Table I.1. Industry shares of the population, 1980[a]

	Percent distribution		
	Urban	Nonmetropolitan	Rural farm
Total	100.0	100.0	100.0
Agricultural	0.8	7.2	45.8
Nonagricultural	99.2	92.8	54.2
Total Nonagricultural	100.0	100.0	100.0
Forestry & fisheries	0.1	0.4	0.3
Mining	0.5	2.7	1.6
Construction	5.2	7.4	8.3
Manufacturing	21.5	25.4	24.7
Total Nonservice	27.3	35.9	34.9
Transportation & public utilities	7.7	6.9	7.3
Wholesale & retail trade	21.4	20.4	19.9
Finance, insurance & real estate	7.2	4.0	4.6
Service	30.7	27.8	28.3
Public administration	5.7	5.0	5.0
Total Service	72.7	64.1	65.1

[a]Source: U.S. Dept. of Commerce, Dec. 1983.

population is represented in forestry and fisheries, mining, construction, and manufacturing.

Unemployment rates in all industries climbed over the period 1978-1982 (see Table I.2). Unemployment rates have been higher in nonservice industries than in service industries. Rural residents, however, are more likely than urban residents to be employed in nonservice industries and may be harder hit by increases in the general unemployment rate.

Table I.2. Unemployment rates by industry of last job[a]

Industry	1978	1979	1980	1981	1982
Agriculture	8.8	9.1	11.0	12.2	14.7
Mining	4.1	4.9	6.4	6.0	13.4
Construction	10.6	10.3	14.1	15.6	20.0
Manufacturing	5.5	5.6	8.5	8.3	12.3
Transportation & public utilities	3.7	3.7	4.9	5.2	6.8
Wholesale & retail trade	6.9	6.5	7.4	8.1	10.0
Finance, insurance, & real estate	3.1	3.0	3.4	3.5	4.7
Service	5.7	5.5	5.9	6.6	7.6
Government	3.9	3.7	4.1	4.7	4.9

[a]Source: Statistical Abstract of the United States.

By 1982, the unemployment rate in agricultural industries was 14.7%, second only to construction. Some rural areas suffered more than others, for example farming dependent counties. Farming dependent areas are ones in which 20% or more of the work force is employed in farming and agribusiness. Farming dependent counties lost population more rapidly than other counties during 1980-1984. Also, 60% of the farming dependent counties lost population, while only 9% of other rural counties did (Petrulis et al., 1987). Five states, Idaho, Iowa, Nebraska, South Dakota and North Dakota, are especially farming dependent.

Effects of technological change

Technological change has affected rural labor markets. Highway and motor vehicle improvements have reduced transportation costs and facilitated access to more areas (Jordan and Hady, 1979). However, this

has not necessarily helped small communities. Small trade centers in rural areas can be likened to neighborhood stores in large cities--small purchases are made there, but major purchases are made elsewhere.

Deregulation of telephone, bus, trucking, banking, and airline services has had detrimental impacts on some rural areas. Telephone companies can now charge urban and rural consumers different prices for service; basic phone service in rural areas has increased because the costs of providing services are higher. Unprofitable bus and airline routes to remote rural areas have been dropped, leaving rural residents without public transportation services (Richards, 1987). Thus, deregulation has lowered the costs of services relatively more for urban than for rural areas and in some cases raised the cost in rural areas (Murphy and Watkins, 1986).

High technology industries may have mixed effects on rural areas. The research and development phase of computers and telecommunications is human capital intensive. Scientists, engineers and other highly-skilled technicians are more likely to be located in urban centers. Even if rural areas accounted for the same proportion of high technology jobs as urban areas, the 72,000 jobs created would be nowhere near the 3 million jobs needed for full employment in rural areas. Furthermore, with strong international competition in high technology fields, robotics and automation may replace U.S. low-wage rural workers who previously held a comparative advantage over U.S. urban workers. Because high technology may improve the efficiency of the adopting industries, the net impact is difficult to determine (Tweeten, 1987).

Rural economies might try to diversify away from high unemployment industries such as agriculture, mining, and other resource based activities. But rural residents would need to acquire other skills than agricultural ones. Different opportunities exist for individuals depending on the location of their residence. People living near metropolitan areas can get higher paid, higher skilled employment than those who live in remote areas where opportunities are limited.

A review of the literature on rural and local labor markets, spatial aspects of labor markets, off-farm labor supply, and human capital is presented in the following section.

Literature Review

The research presented here combines elements of off-farm labor supply and local labor market studies. In addition, joint decision-making within family units is an important aspect of the research; human variables are also included.

Human capital

Most off-farm labor supply studies rely on human capital variables to explain labor supply choices. Human capital refers to investments that individuals make in themselves to increase their welfare and productivity. These investments include schooling, on-the-job training, health care, and migration.

The notion of human capital gained recognition when T. W. Schultz (1961) stated that investments in skill, knowledge, and health were responsible for the rise in earnings per worker over time. About four-

fifths of the economic growth in the U.S. during the 20th century was unexplained by additional inputs; "nontraditional" inputs (improvements in the quality of the labor force) had to be considered as the reason. Gary Becker (1975) formalized the theory of human capital. Later, Schultz (1975) hypothesized that education enhances the ability to deal with economic disequilibria.

Off-farm labor supply

Off-farm labor supply studies focusing on the effect of human capital on work decisions of farmers include those by Huffman (1980), Sumner (1982), Lopez (1984), and Rosenzweig (1980).

Huffman (1980) used agricultural research and extension and farmers' education as human capital variables to estimate the off-farm labor supply of farmers. Human capital affects work decisions through efficiency effects: agricultural research and extension enhance the efficiency of farm production, and education affects household production as well. However, the net effect is ambiguous. Increased efficiency may shift the demand for farm work up or down. If leisure is a normal good, an increase in net farm income will increase leisure and decrease hours of farm work. But since the production process is more efficient, the net effect on hours worked is uncertain. Empirical results indicate that farmers with more education tend to reallocate their time from farm work to off-farm work faster than lesser-educated farmers.

Sumner (1982) examined off-farm wages, labor force participation, and hours worked of Illinois farmers. Human capital variables included

education, job experience, nonfarm training, and dummy variables for farm and other experience. He found education and experience to be the most important human capital variables in determining the off-farm wage. Job experience added to wages but vocational training had a negative impact. Farmers in northern and southern Illinois had proportionally more off-farm employment than those in central Illinois, reflecting local labor market conditions.

In a slightly different approach to estimating off-farm labor supply, Lopez (1984) used specific functional forms in a system of supply and demand equations to estimate farm and off-farm labor supply. He used a Gorman Polar Form for the indirect utility function and a Generalized Leontief for the profit function. Lopez found that education affected both farm and off-farm labor supply, but had a substantially larger effect on off-farm work. His findings are in general agreement with Huffman's (1980) study. Furthermore, the independence of utility and profit maximization was rejected at the 1% level of significance; strong gains in explanatory power are achieved when production and consumption are estimated jointly. The number of family dependents are also an important determinant of labor supply decisions. Lopez used full information maximum likelihood to jointly estimate demand equations for land structures, hired labor, animal stocks, and farm capital; output supply; and demand for hours of farm and off-farm work by the family and for consumption goods.

The labor-supply behavior of two-person households in developing countries was investigated by Rosenzweig (1980). The schooling levels of

husband and wife comprised the human capital variables. Labor supply for landholding and landless households was predicted assuming competitive market behavior and tested with a sample of rural households in India.

Joint decision making

Households are the decision units in this study, and the question arises of how to model time allocation decisions in the context of a household unit. Barnum and Squire (1979) used one household time endowment in their farm household model. But husband's time and wife's time are heterogeneous and should be indexed separately (Huffman and Lange, 1986; Becker, 1981). Husbands and wives allocate their time to different activities and may have dissimilar levels of productivity in market and home time. For this reason, time allocation and work decisions of husbands and wives are estimated separately in most studies. However, a subtle correlation may be present between the decisions of husbands and wives. Shocks to households are a common factor in married couples and probably other household members' decisions on time allocation. Because these households take account of differences in the effects of these shocks across their members, our econometric model will have better statistical properties if it also takes account of these common shocks.

Rosenzweig (1980) estimated wage equations for males and females in rural Indian households but did not jointly estimate off-farm work by members of the household. Huffman and Lange (1986) develop a farm household model where time allocation decisions of two household members

are made jointly. The econometric model has four structural equations, a labor demand and an off-farm labor supply function for each member. Market labor demand functions should not be affected by a spouse's decision on off-farm work and can be fitted by least squares estimation. They estimated conditional and unconditional off-farm labor supply curves with ridge regression. Ransom (1987a,b) considered the interdependent nature of family labor supply decisions in two recent papers.

Rural and local labor markets

Few studies have considered the workings of local labor markets. Topel (1986) looked at wage and employment dynamics in local markets, the local markets being state units. He found that wage rates are quite sensitive to inter-area differences in market conditions, and that the largest wage adjustments occur among the least geographically mobile workers. Variables included were employment growth rates, current and predicted state employment disturbances, probability of unemployment, UI replacement rate, education, experience, and race; CPS data were used for analysis. Rosen (1979) imputed an index of "quality of life" among metro areas using micro data from the CPS; he specified the real wage as a function of personal productivity variables such as education and experience, and city attributes such as the unemployment rate, the crime rate, population density, and others.

Characteristics of local labor markets matter when households are geographically immobile. In a competitive equilibrium where economic agents are perfectly mobile, wage rates will equilize over geographical

areas because areas with relatively high wages will attract migrants until the eventual surplus of labor pushes wages down to competitive levels. According to Topel (1986), "When costs of moving are negligible, arbitrage will eliminate any implications of purely local conditions for the determination of wages." But if agents are immobile, because of land ownership, family ties, or mobility costs, wage equilization over geographical areas may not occur.

Rosenzweig (1980) found geographical variables to be the main source of wage rate variability in rural Indian labor markets and were more important than personal characteristics once gender was taken into account. Local labor market variables included distance of the household from the village; size of the village; and dummy variables for adverse weather conditions, presence of factories or small scale industry, and the availability of government credit and assistance programs.

Bils (1985) examined the movement of real wages over the business cycle using disaggregate data. The key variable of interest was the change in the average national unemployment rate between year t and year t-1. Other important variables were the change in the size of the local labor force, education, experience, a dummy variable for residence in the southern United States, and a dummy variable indicating whether the individual had been unemployed three weeks or more in the past year. Bils found that real wages are procyclical, and are even more procyclical for individuals who move between employers or in and out of the work force. Mitchell et al. (1985) also found procyclical movements in real wages.

The unemployment rate is fairly useful as an economic indicator in metro areas. Although rural labor markets have some different characteristics, e.g., mix of jobs and job growth, than urban markets, the rural and urban labor markets of a state are expected to respond to general business cycle changes in similar ways. Nonmetropolitan areas have a higher proportion of low earning occupations than do metro areas; the incidence of poverty is higher in rural areas; and more people involuntarily work part-time in rural labor markets (Tweeten, 1978; Briggs, 1981). Some economists argue that "sub-employment" is a better indicator of labor market conditions in rural areas than the unemployment rate. An index of sub-employment, developed for urban areas, included discouraged workers, people involuntarily working part-time, and family heads or unrelated individuals working full time that had earnings below the poverty level, in addition to unemployed people. Despite agreement that better measures of rural labor market conditions are needed, little has been done in this direction (Briggs, 1981).

Spatial aspects of labor markets

Although it is useful to analyze "rural" as opposed to "urban" labor markets, it is important to consider spatial aspects of labor markets and their implications. We speak of labor markets as though they were self-contained entities. They are not. A considerable amount of activity occurs between labor markets, and in the long run, migration between markets lessens regional differences.

In The Changing Shape of Metropolitan America, Berry and Gillard

(1977) examine commuting patterns around metropolitan areas. Figures 4.1 and 5.1 (pages 46 and 100) clearly show that rural residents drive considerable distances to work in metropolitan centers, and such activity increased between 1960-1970.

Gillard (1977) discussed a number of studies asserting the importance of commuting. He cites, for example, Logan (1970), who states that "high incomes can occur in areas of low agricultural productivity near cities because of work force interaction with urban areas." A study by Nichols (1969) suggested that the small towns of western Georgia grew much faster than similar sized metropolitan and nonmetropolitan growth patterns for 1960-1970. He says: "One of the clearly discernible patterns that emerged was that which was related to the diffusion and decentralization of metropolitan growth, both in terms of population and median family income, beyond existing SMSA boundaries. The identification of extra-metropolitan growth areas suggests the expansion and/or intensification of commuting to the prospective metropolitan labor markets." A study by Greenwood (1981) states: "What seems to have happened recently . . . is that diseconomies associated with dense urban locations, in combination with the declining relative economic importance of distance, have improved the competitive position of dispersed spatial arrangements. The result is that relative productivity differentials between the urban and rural labor forces have diminished. . . ." Regional hierarchical structures lose their relevance when transport, communication, and technology combine to lessen the importance of distance (Berry, 1977). Nonmetropolitan growth may be tied directly to

increased population density in the more distant areas of urban commuting fields (McCarthy and Morrison, 1977).

Other studies

Lange (1979) developed a farm household model to examine determinants of demand for household labor of husbands and wives, for household capital services, and for the capital-labor ratio in household production. He estimated demand equations for various types of time use (farm labor, household labor, off-farm labor, and leisure). The study showed that the household's demand for the wife's labor is determined by household size and composition.

A number of studies discuss the adoption of innovations by farmers and how their ability to adapt to change is influenced by education. Huffman (1974, 1977) investigated the role of education in decision making, finding that farmers adjusted their use of nitrogen fertilizer to the optimal level faster as their schooling level increased. Nelson and Phelps (1966) hypothesized that education speeds the process of technical diffusion, as educated people make good innovators. Rahm and Huffman (1984) present evidence on the role of human capital variables in the adoption of reduced tillage, and Wozniak (1984) developed a model of the decision to adopt interrelated innovations using a human capital approach. Other studies include those by Petzel (1978), Khaldi (1975) and Fane (1975).

Research Objectives

The objectives of this study are: 1) to develop a general model of farm and rural nonfarm household behavior that explains how labor supply and labor participation decisions are made by married couple households; and 2) to develop econometric models of labor demand, labor supply, labor participation and household income for rural nonfarm households and of off-farm labor participation and household income for farm households. The models are fitted by least squares and maximum likelihood estimation procedures to micro data sets obtained from the Current Population Survey for 1978, 1979, 1981, and 1982 with state-level labor market and farm-profitability variables matched to households by state of residence.

In order to develop an improved specification of wage-labor participation decisions of married-couple households, the work-not work decisions of husbands and wives are hypothesized to be jointly determined, and are examined with the aid of a bivariate probit model. The bivariate probit procedure takes into account the correlation of the disturbance terms across husband's and wife's participation equations. Predicted probabilities from the bivariate labor participation equations are used to construct estimates of sample selection terms for labor demand and labor supply equations of the husband and wife.

Sample selection terms must be added to the labor demand and labor supply equations because these equations are fitted only to the selected sample observations where individuals work for a wage. The sample selection terms are the conditional means of the disturbance terms of the labor demand and labor supply equations, given that an individual works

in the market, and they are included in the regression equations to satisfy the assumptions of the classical linear regression model.

Rural households have received little attention in the labor supply literature although their labor market decision-making process may differ from urban households. Industrial and occupational restructuring have also occurred at different rates in rural and urban America, and relatively less economic activity takes place in nonmetropolitan labor markets. The research presented here examines how changes in job growth, unemployment rates, state-level wages and other measures of local labor market conditions have impacted on household decision-making in rural areas.

CHAPTER II. ECONOMIC MODEL

Work decisions are similar for farm and nonfarm households. Farm households operate a self-employed business and may participate in wage work. Nonfarm household members are assumed to engage primarily in wage work, with no self-employed income. Nonfarm households are a subset of the more general case of farm households with wage work, and as a result their work decisions are not as complex as the decisions of farm households. The following analysis considers the general case of a farm household with wage work.

Although the work decisions of farm and nonfarm households are similar, farm household decisions are more complex. Farm households have three activities to which time is allocated--farm work, off-farm (wage) work, and leisure; nonfarm households allocate time to wage work and leisure. Farm households make more complex resource (nontime) allocation decisions--they must choose optimal levels of inputs and outputs in farm production in addition to optimal levels of consumption. Nonfarm households need to choose only optimal levels of consumption. The time and resource allocation decisions are made simultaneously. Both time and resource allocation decisions are more complex for farm households than for nonfarm households. Furthermore, variables such as climatic conditions, farm profitability, and research and extension activities are expected to influence farm household decisions but have no impact on decisions of nonfarm households.

Farm and off-farm labor supply decisions can be viewed as the result

of constrained utility maximization. Households are assumed to maximize utility subject to constraints on time, income, and farm production. Labor supply together with labor demand determine the decision to participate in labor markets. A husband-wife farm household model with both partners participating in off-farm work is presented here. Models for wage earners and farm operators without wage work are presented in Appendix A.

Labor Supply

Labor supply is found by solving for optimal time allocation in constrained utility maximization. All households are assumed to derive utility from household members' leisure (L_1 = householder's leisure, L_2 = spouse's leisure), purchased goods (X_1), and factors that are exogenous to consumption decisions such as age, education, and the number of children in the household (X_2). The utility function is specified as $U = U(L_1, L_2, X_1; X_2)$.

Constraints

Individuals are constrained by time and income, and farm operators face a production constraint.

The time constraint is

$$T_i = T_{if} + T_{iw} + L_i, \qquad i = 1, 2$$

where: T_i = total time,

T_{if} = time spent in farm work,

T_{iw} = time spent in wage work, and

L_i = leisure time.

The implicit production constraint is:

$$G(Q, T_{1f}, T_{2f}, Y_2, K, Y_1) = 0,$$

where: Q = farm output,

T_{if} = time spent in farm work,

Y_2 = variable inputs in farm production,

Y_1 = exogenous factors in production (e.g., weather), and

K = capital inputs in farm production.

The resulting full income constraint is:

$$w_1T_1 + w_2T_2 + V + P_qQ = rK + P_1X_1 + P_2Y_2$$

$$+ w_1(T_{1f} + L_1) + w_2(T_{2f} + L_2),$$

the left-hand side being full income received and the right-hand side being income spent.

Here, P_q = price of farm output,

V = unearned income,

r = cost of capital,

P_1 = price of purchased consumption goods,

P_2 = price of variable farm inputs, and

w_i = wage of individual i.

The Lagrangian expression for a constrained utility maximization is:

$$\ell = U(L_1,L_2,X_1;X_2) - \lambda_1[G(Q,T_{if},Y_2,K;Y_1)]$$
$$- \lambda_2[w_1T_1 + w_2T_2 + V + P_qQ - rK - P_1X_1 - P_2Y_2$$
$$- w_1T_{1f} - w_1L_1 - w_2T_{2f} - w_2L_2] .$$

First-order conditions

The resulting first-order conditions are:

(1) $U_1 \quad - \lambda_2 w_1 = 0,$

(2) $U_X \quad - \lambda_2 P_1 = 0,$

(3) $-\lambda_1 G_{Tf} \quad + \lambda_2 w_i = 0,$

(4) $-\lambda_1 G_{y2} \quad + \lambda_2 P_2 = 0,$

(5) $-\lambda_1 G_K \quad + \lambda_2 r = 0,$

(6) $-\lambda_1 G_Q \quad + \lambda_2 P_q = 0,$

(7) $G(T_{1f},T_{2f},Y_2,K,Q) = 0,$

(8) $w_iT_i + V + P_qQ - rK - P_1X_1 - P_2Y_2 - w_iT_{if} - w_iL_i = 0.$

Appendix B contains a total differentiation of the first-order conditions. Comparative static results are obtained by solving simultaneously the equations in Appendix B.

Comparative statics

A number of comparative statics results with respect to farm time are of interest.

Wage elasticity of demand The wage elasticity of demand for the householder's farm time can be written as follows:

$$\frac{d \ln T_{1f}}{d \ln w_1} = k_{T1w}\, \eta_{T_{1f}R} + k_{L_1}\sigma_{T_{1f}L_1} + k_{T_{1f}}\, \sigma_{T_{1f}T_{1f}} .$$

The householder's full real income elasticity of demand for farm time is

$\eta_{T_{1f}R}$; the proportion of total cost spent on factor r is denoted k_r. The σ_{ab}'s are the Allen partial elasticities of substitution, where a, b = (L, T_{1f}) (Allen, 1938). The own Allen partial elasticities of substitution, σ_{aa}, are required to be negative for utility maximization. For substitute inputs, $\sigma_{ab} > 0$; for complements, $\sigma_{ab} < 0$. However, the signs of σ_{ab} are indeterminant.

Farm price elasticity of demand The farm price elasticity of demand with respect to the householder's farm time is:

$$\frac{d \ln T_{1f}}{d \ln P_q} = k_Q \eta_{T_{1f}R} - k_Q \sigma_{T_{1f}Q} .$$

The expressions $\eta_{T_{1f}R}$ and k_Q were defined previously. The second term, $k_Q \sigma_{T_{1f}Q}$, is the percentage change in T_{1f} caused by a 1% change in P_1 holding real income constant.

Interpretation Interpretation of the comparative static results for the farm household with wage work model is more complicated than for the wage earner model. For example, the wage elasticity of demand for the householder's leisure time in the dual model is:

$$\frac{d \ln L_1}{d \ln w_1} = k_{T1w} \eta_{L1R} + k_{L1} \sigma_{L1L1} + k_{T_{1f}} \sigma_{L1T_{1f}} ,$$

a combination of the real income elasticity of demand for leisure and the elasticity of substitution between leisure and farm time. The sign is indeterminant. In comparison, the wage elasticity of demand for leisure time in the wage earner model is:

$$\varepsilon_{LW} = \varepsilon^c_{LW} + k_{Tw} \eta_{LF} .$$

The first part of the expression, ε^c_{LW}, is always negative because the substitution effect is always negative. The sign of the second part depends on the sign of η_{LF}. If leisure is a normal good, η_{LF} is > 0, and the sign of ε_{LW} depends on the relative magnitudes of the income and substitution effects. A detailed analysis of farm and nonfarm labor supply response is given by Huffman (1980).

Utility maximization conditions

For utility maximization, the marginal utilities of leisure and time in farm work must equal the off-farm wage. Rearranging equations (1) and (3) clearly shows that

$$\frac{U_{Li}}{\lambda_2} = G_{T1f}\frac{\lambda_1}{\lambda_2} = w_i \ .$$

Optimal (nontime) resource allocation is found by rearranging equations (4) and (5) to obtain

$$\frac{\lambda_1 G_y}{\lambda_2} = P_2 \qquad \text{and} \qquad \frac{\lambda_1 G_k}{\lambda_2} = r \ .$$

The marginal cost of the input must equal its marginal value in production.

Supply and demand functions

The first-order conditions are solved simultaneously to obtain the general supply and demand functions of the inputs. In general functional form, they are:

$$Y_2^* = Y_2(P_q, P_2; Y_1),$$
$$K^* = K(P_q, r; Y_1),$$
$$T_{iw}^* = T_w(P_q, P_1, P_2, r, V; Y_1, X_2, w),$$
$$L^* = L(P_q, P_1, P_2, r, V; Y_1, X_2, w),$$
$$X_1^* = X_1(P_q, P_1, P_2, r, V; Y_1, X_2, w), \text{ and}$$
$$T_{if}^* = T_f(w, P_q, Y_1).$$

The optimal values of Y_2^* and K^* are profit-maximizing quantities of inputs and so do not depend on the full range of variables that T_{iw}^*, L^*, and X_1^* do.

Relation to farm operators and wage earners

The models for farm operators and wage earners (Appendix A) are subsets of the farm operator with off-farm work model. In the wage earner model, the utility function is the same, but the time constraint does not contain farm time and there is no production function. Similarly, the farm operator model contains no wage work time.

Differences in the models lead to different determinants of optimal time use. Optimal farm time for a farm operator with no off-farm work is a function of input and output prices, the cost of capital, the price of consumption goods, and unearned income. Farm operators with off-farm work, on the other hand, determine farm time as a function of the wage rate and the price of farm output, and wage work time as a function of all of the above-mentioned variables.

As shown above, elasticities are different for each type of household. Wage earner households have elasticities with the most

straightforward interpretations: income and substitution effects are clearly shown. Elasticities become more complicated, and less easily signed, for farm operators with off-farm work.

Labor Demand

The labor demand or wage offer function depends on human capital (E) and labor market characteristics (M). Local labor market conditions, such as unemployment, affect demand when workers and firms are immobile; and labor demand is assumed to be independent of the number of hours worked. Because labor demand functions are different for men and women, wage offers are considered separately.

Participation in Off-Farm Work

An individual's decision to participate in the labor force can be determined in the following way. The "reservation wage," or wage at which a person will enter the labor force, is found by setting hours of work in the labor supply equation equal to zero and solving for $\ln w^R$. The individual will participate in the labor force if $\ln w^R < \ln w^D$, where $\ln w^D$ is the wage from the labor demand equation. The probability of being in the labor force is then

$$\Pr(\ln w^R < \ln w^D) \ .$$

The same framework can be extended to a farm couple's joint decision to take off-farm work. Both the husband and wife are assumed to spend some time working on the farm. Four combinations of off-farm work choices are possible:

$$P_{00} = Pr(\ln w_H^R > \ln w_H^D,\ \ln w_W^R > \ln w_W^D)',$$
$$P_{10} = Pr(\ln w_H^R < \ln w_H^D,\ \ln w_W^R > \ln w_W^D)',$$
$$P_{02} = Pr(\ln w_H^R > \ln w_H^D,\ \ln w_W^R < \ln w_W^D)', \text{ and}$$
$$P_{12} = Pr(\ln w_H^R < \ln w_H^D,\ \ln w_W^R < \ln w_W^D)',$$

where P_{00} is the probability that neither adult works off the farm, P_{10} is the probability that only the husband works off the farm, P_{02} is the probability that only the wife works off the farm, and P_{12} is the probability that both work off the farm. These decisions are assumed to be jointly determined (Huffman and Lange, 1986).

CHAPTER III. DATA AND VARIABLES

The following chapter describes the data and variables used in the empirical analysis.

Data

Data were obtained from the Bureau of the Census, the USDA, the Statistical Abstract of the United States, and Yale University.

Current Population Survey

The main source of data for this research is the Current Population Survey (CPS) March supplement. CPS data for the years 1979, 1980, 1982, and 1983 are used to pick up business cycle effects on employment decisions, and contain income information from the years 1978, 1979, 1981, and 1982. Data from the four years specified will be pooled to pick up variation over time as well as cross-sectional variation. The variables on households from the CPS are augmented by state-level labor market and farm-profitability variables matched to households by state of residence.

Description of CPS data The CPS is conducted monthly by the Bureau of the Census to provide estimates of employment, unemployment, and characteristics of the general labor force (CPS, 1980). In each survey, information is obtained from over 60,000 households. The March supplement contains information on income that is not found in the other monthly surveys, and it has been collected on a yearly basis since 1968.

Probability samples are used to select housing units for interview.

Eligible participants include the civilian noninstitutional population living in housing units and male members of the Armed Forces living in civilian housing.

Weights are prepared to compute monthly labor force status estimates. The basic weight is the inverse of the probability of selection; it specifies how many people are represented by a given person in the survey. The probability of selection is based on characteristics such as age, sex, race and labor force status as determined by the most recent decennial census counts (U.S. Department of Commerce, 1978).

The CPS uses a hierarchical file structure. Household, family, and individual information make up each household (or housing unit) record. Households may contain more than one family and may include unrelated individuals. The research in this paper utilizes information from the household head's family and spouse, plus a small amount of information on other adults living in the same household.

Farm households are represented in the CPS. Due to the large size of the survey, the farm sample for each year may number over 1500 households. For example, in 1979, the CPS numbered 55,000 households, of which 2,000 (about 3.6%) were farm households and 14,000 (about 25.5%) were nonmetropolitan nonfarm households.

Problems with CPS data When local labor and commodity markets are defined as state units, as they are here, a considerable amount of data is available. Information on units smaller than states is not available on a regular basis. Furthermore, the CPS identifies state of

residence for all households and SMSA of residence if the household is metropolitan. Information about smaller geographic units, such as counties or areas within states, is not reported. Also, Topel (1986) used states as units in his study of local labor markets because he used the CPS as his main data source. A few other problems arise in using the CPS data for rural labor market studies.

First, some variable definitions have been changed or added over the years. Many improvements have been made, but with a loss of consistency in some cases. One variable that is relevant to labor market research is the identification of states or state groups. Prior to 1977, many states were identified as being part of a group of states; identification of each specific state was impossible, limiting usefulness of the data for local labor market studies. Identification of each state is essential for determining effects of local labor markets. Defining labor markets as state units has the advantage of permitting us to use an abundance of data that are available at the state level. State-level labor markets are the largest unit of interest--regions or groups of states are too large to isolate local effects. Thus, the requirement of state identification of each household meant that only surveys later than 1977 could be used.

A second problem with the CPS data is the definition of hours worked by an individual. Hours worked include hours from all jobs, not just from the primary occupation. To obtain a wage, we must divide total income from wages and salaries by total hours worked. Wage earners with two jobs will have an average wage from the two jobs, but farmers with

off-farm employment and other self-employed persons will have total wages divided by total hours (farm and off-farm) worked. Furthermore, we do not know how farm operators with off-farm employment will determine their primary occupation for the survey.

Third, the CPS definition of a farm is not precise. A residence is considered to be a farm on the basis of farm income. The interviewer decides the farm status.

Fourth, before 1986 the CPS did not distinguish between rural and nonmetropolitan areas. Rural and nonmetropolitan are not really interchangeable terms. The Census Bureau defines rural areas as areas having less than 1,000 persons per square mile or towns having a population less than 2,500. Metropolitan areas are counties containing or within commuting distance of an SMSA; nonmetropolitan areas are the remaining counties. Land areas classified as nonmetropolitan greatly exceed those classified as rural (Briggs, 1986, p. 161).

Finally, there are disadvantages as well as advantages to using states as geographical units, a necessity when using the CPS. Much more data are available at the state level on a regular basis. An obvious advantage is that state measures provide more detail of local conditions than do groups of states or national measures. But state-level measures do not provide as much detail of local conditions as do counties or regions within states. If a manufacturing plant closes in a small town with few alternative job opportunities for its residents, many people may lose their jobs and the higher local unemployment rate is not adequately reflected in the state unemployment rate. Many rural communities are

dependent on one employer or industry (Beale, 1978). And, as the preceding discussion implies, conditions within a state can vary widely. Rural areas near cities may be less dependent on the resource-based industries which have experienced economic decline in recent years.

There is considerable activity taking place between labor markets, and migration lessens regional differences. Shocks to national economic conditions are felt throughout the country. A study by Bednarzik and Tiller (1982) examined regional sensitivity of unemployment to short- and long-run fluctuations in national unemployment over the period 1967-80. They state: "With few exceptions, independent regional cycles, although in evidence, contributed very little to regional fluctuations" although "there was some evidence that California may have a cycle of its own." They conclude that "national aggregate supply and demand disturbances are quickly transmitted throughout the economy, and both short- and long-run changes in regional labor market conditions conform closely to national developments. Regions do differ, however, in the degree of their sensitivity to changing national conditions." Wright (1987) discussed the integration of Southern labor markets into national labor markets starting in the late 1940s. However, after the Civil War but before World War II, a Southern labor market seemed to operate independently of national markets.

Other data sources

Data for the state-level labor market and farm-profitability variables were obtained from several sources. Weather data were obtained

from Weather in U.S. Agriculture (Weiss et al., 1985). The book contains monthly precipitation and temperatures weighted by geographical area and by harvested cropland, for each state over the period 1950-1984. Growing degree days for each state were estimated from information published by the U.S. Department of Commerce (1971, 1981).

Unemployment rates, total employment and other employment data were obtained from various editions of the Statistical Abstract of the United States. The unemployment rate is the percentage of total unemployed persons, aged 16 and over, based on the CPS.

Detailed farm input (hired labor and other inputs) and output (crop and livestock) price indices were calculated for each state by researchers at Yale University. The crop price index was based on expected prices of feed grains (corn, oats, barley and sorghum), food grains (wheat and rice), fruits (apples, grapes, oranges and grapefruit), vegetables (onions, lettuce, tomatoes and potatoes), oil crops (soybeans and peanuts), hay, cotton, tobacco, sugar and dry edible beans. Poultry and dairy products (milk, eggs, broilers, and turkeys) and meat animals (cattle and calves, lambs and sheep, and hogs and pigs) were included in the livestock price index. The hired farm labor index was based on expenditures on labor, including cash wages, noncash perquisites and payroll taxes. Other inputs included capital, feed, fertilizer, land, seed and miscellaneous expenditures.

Variables

The following section describes the construction and selection of dependent and independent variables.

Dependent variables

The study encompasses several endogenous variables. They include the real wage, hours worked per year, household income, and an indicator for labor force participation and off-farm work (dummy variable). The means of the dependent variables are given in Table III.1.

Table III.1. Dependent variables--farm and nonfarm

Symbol	Mean--nonfarm	Mean--farm	Name
H1	2045.9[a]	--	Annual hours worked--men
H2	1444.8[b]	--	Annual hours worked--women
W1	$3.48[a,c]	--	Wage--men
W2	$2.00[b,c]	--	Wage--women
THI	$8881.17[c]	$8380.81[c]	Total household income

[a]N = 24571.
[b]N = 17508.
[c]Wages and incomes are deflated to 1967 levels.

Wage Wage is the dependent variable in the individual's labor demand equation. For reasons given in the previous section, this variable can be defined only for wage-earning nonfarm persons. For the nonfarm nonself-employed population, "wage" is derived by dividing income

from "wages and salaries" by the product of "hours worked per week last year" and "weeks worked last year," which are all from the CPS. Nominal wages are deflated by the CPI to obtain the real wage rate. The natural logarithm of the real wage is used in the models. The income reported in the CPS pertains to the previous year. After 1981, "hourly earnings" are reported for wage earners; however, the wage variable is the wage in the survey year while income figures are for the previous year, and only about 20% of the households are asked for this information. To be consistent, the first method is applied to all years.

Because the real wage is determined by the forces of supply and demand, it is a random variable. Therefore, predicted wages will be used as an explanatory variable in labor supply equations. When the correct standard errors are employed, this procedure is equivalent to two stage least squares.

Hours of work Annual labor supply (to wage labor) is derived as "hours worked per week last year" times "weeks worked last year" from the CPS data. This variable is applicable only to wage earning nonfarm household persons.

Household income "Total household income" is derived as wage and nonwage income, adjusted by the CPI, from all household members in each CPS record. The following sources of income are included in total household income: wage and salary income; child support payments; dividends, interest and income from rental property; farm income; social security, railroad retirement, and supplemental security payments; public assistance and welfare; retirement payments; self-employment income;

veterans' payments; and workmen's compensation.

Labor force participation dummy variable The dependent variable for labor force participation is a 1-0 dummy variable that is equal to one if the individual participates in the labor force and is equal to zero otherwise. It is determined by the presence of "wage and salary" income in the CPS.

Independent variables

Two types of variables, personal characteristics and labor market characteristics, are of interest. The means of the independent variables are given in Table III.2.

Personal characteristics Information on all personal characteristics is found in the CPS data set. The following variables were constructed from CPS data. The names in parentheses are the symbol for each variable in the regressions.

(1) Education--Education is defined as the highest grade of schooling completed for husbands and wives (EDUCH and EDUCW).

(2) Age--Age in years of the husband and wife (AGEH and AGEW) is used as a proxy for experience and/or life cycle effects.

(3) Age squared (AGE2).

(4) Race (RACE) is specified as a dummy variable which is equal to one if the individual is nonwhite and equal to zero otherwise.

(5) Nonwage income--The natural log of interest, dividends, and income from rental properties, deflated by the CPI, comprise nonwage income (V). A constant was added to make all observations positive.

Table III.2. Independent variables--farm and nonfarm

Symbol	Mean--nonfarm[a]	Mean--farm[b]	Name
Educh	11.5	11.3	Husband's education
Educw	11.6	11.8	Wife's education
Ageh	47.0	50.5	Age of husband
Agew	43.9	47.2	Age of wife
Race	.07	.03	Race
V	$468.69	$851.97	Nonwage income
Kid6	.31	.27	Number of children under age 6
Kid618	.66	.69	Number of children 6-18
Urate	7.08	6.67	State unemployment rate
Jobgr	.048	.036	Growth in employment
Stw	$2.94	$3.06	State average manufacturing wage
Time	3.0	3.0	Time trend
DNC	.282	.456	North central region
DS	.513	.378	Southern region
DW	.059	.118	Western region
Dif3	-.226	-.268	Unemployment deviations
Service	1.000	.911	Change in percentage of a state's jobs in service occupations
PC	--	.486	Price index of crops
PL	--	.544	Price index of livestock
PI	--	.536	Price index of inputs
PW	--	.507	Price index hired farm labor
Rain	--	35.7	Average annual rainfall
GDD		3335.6	Growing degree days

[a]N = 32,662.
[b]N = 5866.

(6) Children--The number of children in the family is broken into two groups: children under age 6 (KID6) and children aged 6 to 18 (KID618).

The above variables are commonly used to explain labor supply choices. In addition, labor market characteristics are included.

Local market characteristics Local labor and farm output markets are defined as state units. Labor market characteristics are obtained from the CPS and other sources. All individuals who have the same state of residence are assigned the same local market variables.

The state-level labor market characteristics include the following variables.

(1) Unemployment rate of the entire adult population (URATE).

(2) Farm input (hired labor and other inputs) and output (crop and livestock) price indices; the input prices were lagged one year (PW, PI, PC, PL).

(3) Growth in total employment--The log of total employment in year t minus the log of total employment in year t-2 (JOBGR) measures employment growth.

(4) Weather/climate conditions--Two measures of climatic conditions were used: average total rainfall over a 24-year period (RAIN), and mean growing degree days between 10% freeze probability dates (GDD). In addition, an interactive term of growing degree days and rain was specified (GDDRAIN).

(5) Change in job mix--The change in the percentage of service jobs over a two-year period (SERVICE) attempts to capture changes in the mix of jobs. Service jobs include those in service, transportation, government, finance and wholesale and retail trade; total nonagricultural employment includes mining, manufacturing, and construction as well.

(6) State average wages--The state average real wage rate, defined as the average hourly earnings of production workers in manufacturing

industries, lagged two years (STW) is used as a price indicator across states.

(7) Time trend--The time trend is equal to 1 in 1978, 2 in 1979, 4 in 1981, and 5 in 1982 (TIME).

(8) Regional dummy variables are specified for the north central, southern, and western census regions of the U.S. (DNC, DS, DW).

(9) Deviations of actual from predicted unemployment rates, lagged one year (DIF3). Predicted rates for each state were estimated as:

$$E(U_t) = a + b*time + c*U_{t-1} + d*U_{t-2} .$$

The state equations were fit with OLS as the Durbin h test showed no evidence of autocorrelation (Johnston, 1984, p. 318). Predicted unemployment rate equations by state are given in Appendix C.

Given the differences in the industrial distribution of employed rural and urban residents (see Chapter I, Table I.1), it may seem more appropriate to measure the growth in employment, unemployment rates, or change in job mix for rural areas only; for example, the growth in nonmetropolitan employment may be a more precise measure than growth in total employment for estimating impacts of labor market variables on labor market outcomes of rural residents. One of the problems with such an approach is that state-level data needed to construct these variables are available only in census years (every ten years). And, there are several other reasons why total employment growth is a suitable measure.

First, rural-urban labor markets are not self-contained entities. Many rural residents who live near metropolitan areas and some who live

sizable distances away commute to urban centers for employment. Then, although their residence is rural, their place of employment is urban. As businesses move from central cities to outlying suburbs, nonmetropolitan residents gain closer access to job opportunities. Spatial relationships in labor markets are discussed briefly in the literature review.

Second, if total and nonmetropolitan employment growth are related, we may infer the effects of nonmetropolitan growth based on the effects of total job growth. The relationship of employment growth in nonmetropolitan areas in state i to total employment growth in state i, can be modeled as follows:

$$E_i^{NM} = \beta_0 + \beta_1 E_i^T + e_i \ .$$

A regression of E_i^{NM} on E_i^T and a constant showed that E_i^T explained 40% of the variation in E_i^{NM} for the growth in employment between 1970 and 1980.[1] There was a positive relationship between nonmetropolitan and total employment growth over the 10-year period.

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[1] $E_i^{NM} = \underset{(5.70)}{.142} + \underset{(5.33)}{.414}\, E_i^T$, $R^2 = .398$, $N = 45$.

CHAPTER IV. ECONOMETRIC MODELS

A number of empirical models were tested, including wage functions and labor supply functions for nonmetropolitan-nonfarm men and women, joint labor participation for nonmetropolitan-nonfarm husbands and wives and for farm husbands and wives, and income equations for farm and nonfarm households. The following chapter discusses the various econometric models.

Labor Demand

Reduced form wage functions for nonmetropolitan-nonfarm men and women can be fitted to pooled CPS data by least squares. In addition, the equations are refitted with a term to correct for sample selection bias using a Heckman-type two-step procedure (1979).

The models with sample selection terms are to be estimated by least squares. The sample selection terms are necessary because the wage equations are estimated only for individuals who work. Those who do not work have no wage to report. When only individuals who work are included as observations, the expected value of the disturbance in the behavioral functions is no longer equal to zero. This violates a critical assumption in regression models. The sample selection terms are the conditional means of the disturbance terms in the behavioral relationships.

Sample selection terms are constructed in the following way. We assume that individual i works in the market if $w_i^r < w_i^d$, where w^r is the reservation wage and w^d is the market wage rate. The probability

that individual i will work is equal to

$$Pr(\mu_i^{\,r} - \mu_i^{\,d} < X_{i2}B_2 - X_{i1}B_1) = F(XB) \ ,$$

where

$$w_i^{\,r} = X_{i1}B_1 + u_i^{\,r}, \quad \text{and} \quad w_i^{\,d} = X_{i2}B_2 + \mu_i^{\,d} \ .$$

The expected value of the disturbance term is

$$E(\mu_i^{\,d}/\mu_i^{\,r} - \mu_i^{\,d} < X_iB) = -\phi_i/\Phi_i \ ,$$

where ϕ_i is the standard normal probability density function and Φ_i is the standard normal cumulative density function. The normal distribution is truncated from above (see Maddala, 1983, p. 365). The conditional values of the disturbance terms become more complicated in the case where the work decision of one individual is jointly determined with that of their spouse, a situation that arises in the labor supply models.

The real wage for the ith individual of the jth sex is modeled as follows:

$$\begin{aligned} \ln w_{ij} = {} & a_0 + a_1{*}educ_{ij} + a_2{*}age_{ij} + a_3{*}age_{ij}^{\,2} + a_4{*}race_{ij} \\ & + a_5{*}urate_i + a_6{*}jobgr_i + a_7{*}(\ln stw)_i + a_8{*}service_i + \\ & + a_9{*}dif3_i + a_{10}{*}time_i + a_{11}{*}DNC_i + a_{12}d{*}DS_i + a_{13}{*}DW_i + e_{ij} \ . \end{aligned} \qquad \text{(IV.1)}$$

Separate wage equations are fitted for men and women, and only nonmetropolitan-nonfarm residents with no self-employed income are included for reasons explained in a previous section.

A number of hypotheses are to be tested regarding the importance of

human capital and labor market variables on wages. The effect of human capital on wages is well-established in the literature (Willis, 1986). The results of previous studies lead us to hypothesize that education and experience have positive effects on wages. The coefficient of education, a_1, is expected to have a positive sign, which indicates that persons with higher levels of schooling receive higher wage offers in the labor market than persons with less schooling. Likewise, persons with more experience (another form of human capital) are expected to receive higher wage offers than persons who have less experience. Here, age is used to proxy experience and other life cycle effects, as it is not based on choice variables as experience may be. The coefficient of age, a_2, is expected to have a positive sign, while the coefficient of age^2, a_3, is expected to have a negative sign. Age-earnings profiles show earnings increasing most rapidly early in life, gradually slowing and finally decreasing near retirement.

Gerner and Zick (1983) analyzed labor demand, labor supply and labor participation by women in two-parent, two-children families. Their wage equation showed that a one-year increase in education increased wages by 7 percent. Age had a positive and diminishing effect on wages.

The sign of a_4, the coefficient of the race dummy variable, is expected to be negative. The level of significance may differ for men and women as the earnings gap between blacks and whites differs by sex. Hamermesh and Rees (1984, p. 308) report that in 1984, mean earnings of white women were only $140 above those of black women, while mean

earnings of white men were $5652 above those of black men. Their figures suggest that race may be a more important factor in determining wages for men than for women. Smith and Welch (1977) found race to have a negative impact on earnings.

We hypothesize that labor market characteristics effect the wage an individual receives in the labor market. The statistical significance and signs of coefficients a_5 through a_{13} should confirm or reject our hypothesis. An increase in the unemployment rate is expected to put downward pressure on labor demand and real wage rates. Higher unemployment implies more people available for work, and the increased supply of workers should result in lower wages. Therefore, a_5 is expected to have a negative sign. However, Gerner and Zick (1983) found the state unemployment rate to have a positive, but not statistically significant, effect on women's wages.

Other studies have used a variety of specifications of the unemployment rate. Molho (1986), in his wage equations for married couples, used the rate of growth of unemployment ($\Delta \ln U_t$) and found its coefficient to be negative for both men and women but not statistically significant for men. Bils (1985) estimated an equation for $\ln(w_t/w_{t-1})$, where w is the real wage, as a function of U_t-U_{t-1}, where U_t is the national average unemployment rate in year t. The coefficient of U_t-U_{t-1} was negative and statistically significant, although the regressors explained less than 2% of the variation in $\ln(w_t-w_{t-1})$. Bils also used the change in the size of the labor force over a one-year period and found its coefficient to be positive but not significant.

Rosen (1979) points out that the unemployment rate may have two opposing effects. It may reflect a temporary excess of labor which would push wages down. But to the extent that inter-area differences in the unemployment rate reflect long-run economic conditions, positive wage premiums in high unemployment areas are required as "compensating differentials" to make up for the higher risk of being unemployed. The latter effect dominated Rosen's data on urban residents.

Growth of total employment (jobgr) is expected to increase wages. If more workers are demanded in the economy, competing firms should have to pay higher wages to obtain their labor services. However, total employment is a result of the forces of demand and supply in the labor market, and we observe only the equilibrium of these forces. If supply side effects are greater than demand side effects, growth of total employment would decrease wages.

The state wage variable is included as a general wage or price indicator across states. Its coefficient, a_7, is expected to have a positive sign. A similar variable used by Rosenzweig (1980) in his study of rural Indian labor markets had a positive and statistically significant coefficient. Rosenzweig used a district-level daily wage as a proxy for aggregate market conditions, as important characteristics of local labor markets that influence wage rates may not have been captured by the other variables. A positive sign on the coefficient indicates that individuals residing in states with higher average wages are expected to earn higher wages, all else equal.

We would like to have some indication of how changes in occupational

structure affect wages; we would expect to see wages increase if a greater share of jobs were in higher-paying occupations. The variable "service," which measures the change in the percentage of service jobs over a two-year period, is hypothesized to have a positive coefficient.

In addition to the absolute level of the unemployment rate, deviations of predicted from actual unemployment rates are expected to affect wages. Specifically, positive deviations (which occur when the actual rate is higher than the predicted rate) are expected to put downward pressure on real wages. Therefore, we hypothesize that the sign of the coefficient of dif3 (a_9) should be negative.

A time trend was also included in the wage equations to capture long-term movements in real wages not due to other labor market variables. Real wages have been trending downward in recent years; therefore, a_{10} is expected to have a negative sign. Bils (1985) used a time trend in his study of real wages over the business cycle, and found its coefficient to be negative and statistically significant.

Dummy variables for the north central, southern, and western census regions are included to pick up unmeasured effects associated with regions. Expected signs of their coefficients are unknown. Bils (1985) used a dummy variable for the south in his wage equation and found its coefficient to be negative and statistically significant.

Labor Supply

Labor supply equations for nonmetropolitan-nonfarm men and women are fitted by two stage least squares to four subgroups with sample selection

terms for each sex. The sample selectivity terms are required to avoid non-participation bias. The labor supply equations are estimated first without the sample selection terms to determine which variables to include in the labor participation (probit) equation, from which the sample selection terms are derived. The procedure described by Heckman (1979), but expanded by Huffman-Lange (1986) advances for two persons, was used for estimation. The two-step procedure takes account of the two relationships in labor supply--the work-not work decision and the hours of work decision--and is generally preferred over the Tobit specification (Killingsworth and Heckman, 1986; Killingsworth, 1983).

An important point to remember in interpreting models of labor supply is that labor supply decisions are actually two decisions. The first decision is whether or not to work at all, and the second decision is how many hours to work given that the individual has decided to work. Work-not work decisions can be examined using discrete-choice probit models, and hours of work is a continuous variable that is truncated.

The Tobit model combines both the discrete choice and continuous choice aspects of labor supply. The regression is fitted to all observations; if an individual does not work, their hours of work equals zero. The Tobit procedure has been used frequently in the labor supply literature.

Selection bias corrected regression is similar to the Tobit model, but it is done in two steps rather than one. The first step is to fit a probit model to obtain the probability of being in the labor force. The second step is to use the predicted probabilities from the probit model

to construct sample selection terms to correct for possible sample selection bias. Because the sample selection terms are equal to the conditional means of the disturbance term in the conditional labor supply functions, their inclusion in the hours worked equations for individuals who work insures that the selection bias corrected regression will have a zero mean disturbance term.

Killingsworth (1983) points out that the selection bias corrected regression seems to involve an extra set of parameters (one for the probit and one for the least squares regression), while the Tobit procedure estimates one set of parameters. He concludes, however, that the selection bias corrected regression allows for discontinuities in the labor supply curve that are ignored in Tobit analysis. Furthermore, Mroz (1987) found that the Tobit procedure exaggerates income and wage effects for labor supply of married women.

Labor supply functions of men and women are fit separately. (For a discussion of the literature on male and female labor supply, see Ashenfelter and Layard, 1986.) Many econometric studies of labor supply have found own wage elasticities of married men to be quite small and sometimes negative. In contrast, the own wage elasticity of labor has been found to be relatively large for married women. Boskin (1973) argues that this occurs because their representation in the labor force is not as great as their husbands'. Enormous aggregation bias has been shown to occur in labor supply functions that ignore noneconomic effects such as sex, race, age, and family position (Boskin, 1973).

In this study, labor supply functions are fit separately for men and

women based on the labor participation status of their spouses. Ransom (1987a) showed that men with working wives have different labor supply functions than "would otherwise identical men whose wives do not work since additional wife's leisure cannot be purchased by the latter group." Like Huffman and Lange (1986), he used a model with an endogenous switching rule.

The four subgroups for which labor supply functions are fitted in this study are: 1) working men whose wives do not work; 2) working men whose wives work; 3) working women whose husbands do not work; and 4) working women whose husbands work. The derivation of the sample selection terms for this model are shown in Appendix D.

Annual hours worked of the ith individual in subgroup 4 can be modeled as follows:

$$\begin{aligned} T_{iw} = {} & b_0 + b_1{*}educ_{ih} + b_2{*}educ_{iw} + b_3{*}kid6_i + b_4{*}kid618_i \\ & + b_5{*}age_{ih} + b_6{*}age_{ih}^{\ 2} + b_7{*}race_{ih} + b_8{*}(\ln v)_i \\ & + b_9{*}(\ln w_1)_i + b_{10}{*}(\ln w_2)_i + b_{11}{*}time_i + b_{12}{*}DNC_i \\ & + b_{13}{*}DS_i + b_{14}{*}DW_i + e_i \ . \end{aligned} \qquad \text{(IV.2)}$$

Group 3 husbands do not work, so b_9, the coefficient of the husbands' wage, is zero. Likewise, b_{10} is equal to zero for group 1.

Each individual allocates time between three activities: work in the marketplace, leisure, and work at home. The time allocation decision depends in part on tastes and preferences of the individual. Education may influence time allocation decisions not only through wages and productivity, but also through tastes and preferences.

The signs and magnitudes of b_1 and b_2 depend not only on market productivity, but also on efficiency effects in household production. Increases in education may make nonmarket time more productive, resulting in less time being spent in the market and more at home. Therefore, the signs of the education coefficients are unknown a priori, although they are likely to be positive for men. A recent study showed that men with higher levels of education tend to complete longer work hours than men with lower levels of education (Pencavel, 1986). Likewise, the signs of the coefficients of age, age squared, and race are unknown a priori. Age may influence labor supply decisions in various ways. Including age as an exogenous variable controls for the problem of observing individuals at different points in their life-cycle. Asset accumulation is highly correlated with age; savings usually increase as earnings increase later in the life cycle. Childbearing and child rearing decisions are usually made early in the life-cycle.

Gerner and Zick (1983) found the wife's education to have a positive, but not significant, effect on her own hours of work. They also found that age had a negative but not significant effect on the wife's labor supply.

The signs of b_3 and b_4 are hypothesized to be less than zero for women. Studies have shown that the presence of children in a household tend to discourage women from working but have little effect on the labor supply of men. The presence of small children (under age 6) especially may decrease the likelihood that a woman will take wage work (Huffman and Lange, 1986; Gronau, 1977). Some economists argue that the number of

children should not be used to explain hours of work because couples jointly choose the number of children to have and hours of work; it is a choice variable that reflects tastes of the parents and household productivity. However, Mroz (1987) found no evidence that children or nonwife income were endogenous variables in married women's labor supply functions.

The coefficient of $\ln \hat{w}_1$, the husband's predicted wage, is expected to be negative for men. Most studies report wage elasticities for men that are small and negative. Its sign is expected to be negative for women whose husbands work, and is restricted to zero for women whose husbands do not work. Schultz (1980) and Gerner and Zick (1983) found the husband's wage to have negative and significant effects on the wife's labor supply.

The coefficient of the wife's predicted wage, $\ln \hat{w}_2$, is expected to be positive for women, because many studies have shown wage elasticities for women to be positive and much larger than those of men. Some recent studies, however, have found female wage elasticities to be smaller than those in previous studies (Killingsworth and Heckman, 1986). The coefficient of $\ln \hat{w}_2$ is restricted to be zero for men whose wives do not work, and its sign is unknown a priori for men whose wives work. Schultz (1980) reported positive own-wage effects for married women. One study found the wife's predicted wage to have a positive, but not significant, effect on her own hours of work (Gerner and Zick, 1983).

Increases in nonwage income are expected to decrease labor supply for both men and women. A time trend is included, although the sign of

its coefficient is unknown a prior for the four years examined. Similarly, the signs of the regional dummy variable coefficients are unknown.

Labor Participation

Labor participation is determined by labor demand and labor supply. An individual will participate if ln w^d (from labor demand) is greater than ln w^r (the reservation wage, or the minimum wage required to enter the labor force). Reduced form joint labor participation equations will be fit for husbands and wives in the nonmetropolitan-nonfarm and farm subgroups. Nonfarm self-employed households will be excluded. The dependent variable is a dummy variable, D_{ij}, which is equal to 1 if the ith individual of the jth sex participates in wage work. Labor participation is estimated by a maximum likelihood probit procedure.

The participation decisions of husbands and wives may be correlated; therefore, their decisions should be estimated jointly using a bivariate probit procedure. The bivariate probit procedure was used by Catsiapis and Robinson (1982) to estimate financial aids grants available to students using two selection rules, enrollment and net grants. To test the hypothesis of joint decision-making, we test whether ρ, the correlation coefficient, is significantly different than zero.

Nonfarm

The probability of participation in the labor market by the ith person of the jth sex is modeled as follows:

$$
\begin{aligned}
PR(D_{ij}=1) = {} & c_0 + c_1{*}educ_{ih} + c_2{*}educ_{iw} + c_3{*}race_{ih} + c_4 age_{ih} \\
& + c_5{*}age_{ih}^{\ 2} + c_6{*}kid6_i + c_7{*}kid618_i + c_8{*}urate_i \\
& + c_9{*}jobgr_i + c_{10}{*}(\ln\ stw)_i + c_{11}{*}service_i + c_{12}{*}dif3_i \\
& + c_{13}{*}time_i + c_{14}{*}(\ln\ v)_i + c_{15}{*}DNC_i + c_{16}{*}DS_i + c_{17}{*}DW_i + e_{ij}\ .
\end{aligned}
\tag{IV.3}
$$

All of these variables are from the labor demand and labor supply equations. Because the coefficients are combinations of those used in labor demand and labor supply, many of the signs are unknown a priori.

The signs of most of the human capital variable coefficients are unknown. They enter the labor participation decision through labor supply; therefore, their signs may depend on efficiency effects in household production. A woman will not participate in the labor market unless her market productivity exceeds her home productivity (Gerner and Zick, 1983). And, the characteristics that affect market productivity, such as education and experience, also affect home productivity.

The number of children is expected to reduce the probability of women participating in the labor market because their home productivity is raised. The number of children under age 6 had a negative impact on the probability of Japanese women working in the labor market (Hill, 1983). Gerner and Zick (1983) found that the older the oldest child in the family, the higher the probability that the woman will participate in the labor market.

An increase in the unemployment rate is expected to decrease the probability of entering the labor force; therefore, the sign of c_8 should be negative. The unemployment rate enters the probit equation through

the labor demand function. A study by Bowen and Finegan (1969) found that the unemployment rate had a negative, significant effect on aggregate labor force participation rates of rural nonfarm men aged 25-54. The coefficient of jobgr, c_9, may be either positive or negative, depending on the sign it takes in the labor demand function. The state unemployment rate had a negative and statistically significant effect on the probability of a woman working in the labor market in a study of two-parent, two-children households by Gerner and Zick (1983).

Coefficients of both stw (c_{10}) and service (c_{11}) are expected to be positive as they were hypothesized to be in the labor demand equation. Likewise, the coefficient of dif3 (c_{12}) is hypothesized to be negative.

The coefficient of the time trend is a combination of the time coefficients in the labor demand and supply equations. It was expected to be negative in the demand equation, but its sign was unknown in the supply equation. Therefore, its sign is indeterminant a priori in the participation equation.

An increase in nonwage income is expected to decrease the probability of participating in the labor market, just as it is expected to decrease hours worked in the labor supply equation. The coefficient of ln v (c_{14}) is hypothesized to be negative.

The signs of the coefficients of the regional dummy variables are unknown a priori. Their signs were indeterminant in the labor supply and labor demand equations as well.

Farm

The probability of a farm operator taking off-farm wage work is estimated in the same way as the probability of participating in the labor market for a nonfarm resident. Therefore, the off-farm participation equation for the farm group is the same as for the nonfarm group with the addition of variables that indicate farm profitability. The probability of the ith individual of the jth sex taking off-farm work can be modeled as:

$$\begin{aligned} Pr(D_{ij}=1) = {} & c_0 + c_1{*}educ_{ih} + \ldots + c_{17}{*}DW_i + c_{18}{*}(\ln pc)_i \\ & + c_{19}{*}(\ln pl)_i + c_{20}{*}(\ln pi)_i + c_{21}{*}(\ln pw)_i \\ & + c_{22}{*}rain_i + c_{23}{*}GDD_i + c_{24}{*}GDDrain_i + e_{ij} \ . \end{aligned} \qquad \text{(IV.4)}$$

Increases in the price indices of crops and livestock are hypothesized to decrease the probability of working off the farm. Higher output prices should make farming more profitable and increase individual's reservation wages; therefore, they should be less likely to work off the farm if output prices rise. Increases in input prices, however, make farming less profitable and farmers more likely to work off the farm. Farms located in areas with favorable weather conditions, such as abundant precipitation and a long growing season, are expected to be more profitable than farms in other areas. Therefore, c_{22} and c_{23} are expected to have negative signs. Huffman and Lange (1986) found that a longer growing season reduced the probability of the husband working off the farm in their sample of Iowa farm households. The sign of c_{24}, the coefficient of the interactive term between growing degree days and rain,

is unknown. The above hypotheses hold provided leisure is a normal good.

Previous empirical studies (Huffman, 1980; Huffman and Lange, 1986; Sumner, 1982; Lopez, 1984) provide insights as to expected signs of the human capital variable coefficients. Human capital investments clearly influence off-farm work decisions. Individuals with higher levels of education are more likely to take off-farm work than individuals who have less education; an increase in education of the wife decreases the probability of the husband taking off-farm work (Huffman and Lange, 1986). The husband's age has a positive and diminishing effect on the probability of working off the farm. Huffman and Lange (1986) also found that the presence of children under age 6 had a statistically significant negative impact on the probabilities of both men and women taking off-farm work.

Household Income

Household income is a function of employment and investment decisions by the members of the household. Because all of the variables that constitute decisions by members make up household income, it can be expressed as a reduced form of the exogenous variables in the members' behavioral equations. Studies of family or household earnings or income inequality typically include personal characteristic variables such as education, race, experience, and gender (Chiswick, 1983; Blau, 1984; Gardner, 1969; Podgursky, 1983).

Income for farm and nonfarm households will be estimated separately as more variables are necessary to explain farm household income.

Household income equations are estimated by ordinary least squares.

Nonfarm

The reduced form household income equation for the ith household is:

$$\begin{aligned} \ln(\text{total household income})_i = {} & d_0 + d_1{*}educ_{ih} + d_2{*}educ_{iw} + d_3{*}age_{ih} \\ & + d_4{*}age_{ih}^{\ 2} + d_5{*}race_{ih} + d_6{*}(\ln v)_i + d_7{*}kid6_i + d_8{*}kid618_i \\ & + d_9{*}time_i + d_{10}{*}service_i + d_{11}{*}(\ln stw)_i + d_{12}{*}urate_i \\ & + d_{13}{*}jobgr_i + d_{14}{*}dif3_i + d_{15}{*}DNC_i + d_{16}{*}DS_i + d_{17}{*}DW_i + u_i \ . \end{aligned} \tag{IV.5}$$

The signs of d_1 and d_2 are hypothesized to be greater than zero. Persons with more education can earn higher incomes than persons with less education. An increase in age (experience) is hypothesized to increase household income. Age may have many effects on household income, reflecting the fact that people face different decisions at different points in the life-cycle. Choices on asset accumulation and children affect household income and they are related to age. Households tend to accumulate assets over the life cycle until retirement. For farm households, the number of acres owned is likely to increase with age. This effect is expected to diminish at older ages so d_3 should be positive and d_4 should be negative. The same life cycle effects are evident in wage equations.

Since nonwhite individuals are hypothesized to earn lower wages, they are also expected to have lower total income. Therefore, d_5 should be negative.

Nonwage income is a component of total income; hence, its coefficient should have a positive sign. We may consider dropping this variable if we believe it is not truly exogenous; that is, current nonwage income is a result of past decisions by households. Past savings depend on labor supply (Keeley, 1981) and may depend on household income.

The number of children is likely to influence total household income in different ways depending on the age of the children. Women are less likely to work if they have small children; if the wife does not work, the family will have less total cash income. Children aged 6 to 18, however, require less of their parents' time and their presence is less likely than the presence of younger children to influence the woman's work decision. In addition, teenage children may work for wages and add to total household income. Therefore, the coefficient of kid6 is expected to be negative, while the coefficient of kid618 may be of either sign.

The signs of the labor market variable coefficients are expected to be as follows: d_{10} and d_{11} are expected to be greater than zero; d_{12} and d_{14} are expected to be less than zero; and d_{13} may be of either sign. Regional effects are unknown a priori.

Farm

In addition to the variables specified for nonfarm household income, variables that reflect farm profitability are used to estimate farm household income. Farm household income for the ith household can be modeled as:

$$\ln(\text{total household income})_i = d_0 + d_1{*}\text{educ}_{ih} + \ldots + d_{17}{*}\text{DW}_i$$
$$+ d_{18}{*}(\ln \text{pc})_i + d_{19}{*}(\ln \text{pl})_i + d_{20}{*}(\ln \text{pi})_i + d_{21}{*}(\ln \text{pw})_i \quad \text{(IV.6)}$$
$$+ d_{22}{*}\text{rain}_i + d_{23}{*}\text{GDD}_i + d_{24}{*}\text{GDDrain}_i + e_i \ .$$

We would expect much the same response from the farm and nonfarm groups to the variables they have in common. As farm output prices increase, farm household income is expected to increase accordingly. However, as livestock prices rise, farmers may hold back young livestock to use as breeding stock in anticipation of even higher prices in the future--and the immediate effect is for income to decrease rather than increase. Because of this "inventory effect," the coefficient of livestock prices, d_{19}, is expected to be negative. The opposite effect is expected for input prices. As input prices rise, farming becomes less profitable and household income is expected to fall; d_{20} should be negative. Hired labor (d_{21}) is also an input, but if the wages of hired farm labor represent an opportunity cost of farmers' time, its coefficient may be positive.

Plentiful precipitation and a long growing season are expected to make farming more profitable and increase farm household income. Therefore, both rain and GDD are expected to have positive coefficients. The interaction term between rain and growing degree days is expected to be negative if they are complementary inputs.

CHAPTER V. WAGE WORK DECISIONS OF MARRIED COUPLES

Chapter V presents empirical estimates of the wage-labor participation, labor demand, and labor supply equations for the nonmetropolitan-nonfarm population, and of the off-farm labor participation equations for the farm population.

Labor Market Decisions of Nonmetropolitan-Nonfarm Couples

This section is concerned with several labor market outcomes of nonmetropolitan-nonfarm households. Labor participation, labor demand and labor supply equations are discussed separately.

Labor participation

Over the period 1978-82, participation rates of men have fallen, while participation rates of women remained basically the same over the period 1978-82. The participation rates for nonmetropolitan-nonfarm men and women during 1978-1982 were 75.2% and 53.6%, respectively (see Table V.1). Wage work participation of nonmetropolitan-nonfarm married men fell slightly between 1978 and 1982 in almost every age group except men aged 50-59. The participation rate for men of all ages declined from 76.5% in 1978 to 73.9% in 1982, a drop of 2.6% in five years. Participation rates of women of all ages remained basically the same over the five-year period at about 53.5%. Participation rates for women aged 30-39 and 40-49 actually increased, from 65.1% to 68.2% and from 60.7% to 65.1%, respectively.

Determinants of wage-labor participation decisions of

Table V.1. Participation rates of U.S. nonmetropolitan-nonfarm married men and women, by age (nonself-employed), 1978-82

	Age							
	<20	20-29	30-39	40-49	50-59	60-60	>70	All
	%							
	Men							
Total	94.3	93.4	93.7	92.4	82.5	45.4	12.5	75.2
1978	100.0	94.9	94.0	92.4	82.8	48.6	12.0	76.5
1979	97.8	94.3	94.1	93.4	81.6	46.3	13.9	76.2
1981	84.8	93.1	93.5	92.0	82.8	43.2	12.5	74.3
1982	90.9	91.0	93.3	91.8	82.7	43.5	11.4	73.9
	Women							
Total	64.9	69.2	67.2	64.0	49.1	22.0	4.0	53.6
1978	71.8	69.7	65.1	60.7	49.1	22.5	5.0	53.3
1979	65.8	70.3	65.9	64.8	50.0	20.9	4.1	54.0
1981	56.5	67.5	69.4	65.1	49.5	21.2	3.8	53.5
1982	63.6	69.1	68.2	65.1	47.7	23.7	3.3	53.6

nonmetropolitan-nonfarm married couples are the focus of this section. Two major hypotheses about wage work participation are tested: 1) participation decisions of married couples are made jointly (versus independently); and 2) participation decisions are significantly affected by state labor market conditions.

The first issue to be examined is the joint estimation of the equations explaining the probability that a husband and wife participate in the wage labor force. One way to test this hypothesis is to test whether the disturbance terms of the wage work participation equations of husbands and wives are correlated. The null and alternative

hypotheses can be formally stated as follows:

$$H_0: \rho = 0$$
$$\text{vs. } H_A: \rho \neq 0,$$

where ρ is the correlation between the disturbance terms in the husband's and wife's wage work participation equations.

When the wage-labor participation equations of husbands and wives were fitted jointly to the 32,662 observations on nonmetropolitan-nonfarm husband-wife households, the estimated value of ρ was 0.192.[1] The correlation of disturbance terms across equations is positive, and the correlation coefficient, which has a t-ratio of 15.8, is significantly different from zero at the 1% level. Thus, we reject the null hypothesis that disturbance terms are uncorrelated across wage-work participation equations of husbands and wives. The two participation equations should be fitted jointly. The work-not work decisions of husbands and wives respond similarly to common economic shocks, a result which suggests that the participation equations fitted by the bivariate procedure "fit" better than ones fitted by the univariate probit procedure.

The predicted probabilities of wage work for individuals from the

---

[1] If $\rho = 0$, the value of the log-likelihood functions for husbands and wives can be added to obtain an estimate of the joint log-likelihood function. This summation can then be compared with the value of the log-likelihood function of the bivariate probit equation. The value of the summation of the two log-likelihood functions is -29,672 and the value for the bivariate probit is -29,554. The sample value of the χ^2 statistic is 236 and the tabled value of χ^2 with one degree of freedom at the 1% level of significance is 3.84.

two procedures are, however, very similar. To see this, the predicted values of the univariate probabilities of an individual participating in wage work was regressed on the predicted probabilities for an individual from the bivariate model. The R^2 for these regressions was 99.6% for married men and 99.7% for married women (based on a 10% subsample of the data). The participation equations fitted by the bivariate procedure "fit" better, but in cases where the predicted probabilities are needed the univariate predicted probabilities would be a reasonably accurate proxy of the more expensively obtained predicted probabilities from bivariate estimation.

Results The participation equations explain the probability of an individual participating in wage work. The empirical specification of the participation equations for husbands and wives in nonfarm households (equation IV.3) were fitted to the whole sample (32,662 observations). The estimated coefficients of the univariate and bivariate probit equations are reported in Table V.2.[2] Because the models are nonlinear, the estimated coefficients are not a direct estimate of the marginal effect of a regressor on the probability of wage work. The marginal effect of X_j is $\partial P_i/\partial X_{ij} = f(X_i\beta)\beta^*$ where $f(x_i\beta)$ is the density function, and estimates of these effects arie reported in Table V.3.

The human capital variables, education and age, have coefficients that are highly statistically significant in both the univariate and

[2]The variable age squared had to be rescaled so the bivariate maximum likelihood function would converge. The univariate probit coefficients were used as starting values for the bivariate maximum likelihood function.

Table V.2. Probit estimates of wage labor participation equations for U.S. nonmetropolitan-nonfarm husband-wife couples, 1978-82[a]

	Men		Women	
	Univariate	Bivariate	Univariate	Bivariate
Intercept	3.03 (4.04)	3.03 (4.12)	8.73 (11.85)	8.68 (12.08)
Age_h	.106 (25.02)	.106 (27.16)	.025 (7.23)	.026 (7.62)
$Age_h^2/100$	-.159 (37.11)	-.158 (41.53)	-.067 (18.78)	-.067 (19.36)
$Educ_h$	.042 (11.57)	.041 (11.44)	-.018 (5.64)	-.017 (5.50)
$Educ_w$	.006 (1.26)	.007 (1.63)	.094 (24.31)	.095 (25.73)
Race	.125 (3.43)	.106 (3.22)	.331 (10.79)	.309 (10.42)
Kid6	-.034 (1.73)	-.027 (1.38)	-.485 (34.16)	-.495 (35.43)
Kid618	-.008 (.73)	-.010 (1.00)	-.090 (11.37)	-.083 (10.90)
Ln V	-.335 (4.14)	-.359 (4.53)	-.842 (10.57)	-.867 (11.16)
Ln stw	-.033 (.37)	.124 (1.53)	-.820 (11.03)	-.617 (9.21)
Urate	-.050 (7.82)	-.050 (8.04)	-.035 (6.86)	-.042 (8.50)
Dif3	-.024 (1.91)	-.007 (.57)	.003 (.25)	.010 (.96)

[a]Asymptotic t-ratios are in parentheses.

Table V.2. (continued)

	Men		Women	
	Univariate	Bivariate	Univariate	Bivariate
Jobgr	-.193	-1.27	-.522	-.202
	(7.13)	(4.54)	(2.27)	(.84)
Service	.005	.006	-.003	-.004
	(.77)	(1.00)	(.59)	(.75)
DNC	-.103	-.125	.206	.185
	(3.04)	(3.49)	(2.81)	(6.69)
DS	-.215	-.214	-.109	-.105
	(6.48)	(6.68)	(4.18)	(4.20)
DW	-.137	-.163	.027	-.026
	(2.81)	(3.98)	(.69)	(.79)
Time	-.014	.001	.026	.040
	(1.55)	(.16)	(3.53)	(5.61)
ln L	-11,346	-29,554	-18,326	-29,554
ρ	.192			
	(15.8)			

Table V.3. Estimates of marginal effects of explanatory variables on the probability of wage labor participation: U.S. non-metropolitan-nonfarm, 1978-82

	Men		Women	
	Univariate	Bivariate	Univariate	Bivariate
Age_h	-.012	-.012	-.015	-.015
$Educ_h$	.012	.012	-.007	-.007
$Educ_w$	.002	.002	.037	.038
Race	.035	.030	.132	.123
Kid6	-.101	-.008	-.193	-.197
Kid618	-.002	-.003	-.036	-.033
Ln V	-.004	-.005	-.015	-.015
Ln stw	-.003	.012	-.111	-.084
Urate	-.014	-.014	-.014	-.017
Dif3	-.007	-.002	.001	.004
Jobgr	-.545	-.356	-.208	-.081
Service	.001	.002	-.001	-.001
DNS	-.029	-.035	.082	.073
DS	-.061	-.060	-.043	-.042
DW	-.039	-.046	.011	-.010
Time	-.004	.000	.010	.016

bivariate estimates of wage-work participation equations. The signs of the human capital variables were unknown a priori (see equation IV.3) because characteristics that affect market productivity also affect home productivity. All age effects are captured by husband's age, and a one year increment to his age has a positive but diminishing marginal effect on the probability that he and his wife participate in wage work. At the sample mean, the marginal effect of age on the husband's probability of participating is -1.2% and for wives is -1.5%.

An increase of an individual's own schooling raises the probability that he (she) participates in wage work, other things equal. These results suggest that on average an increment to an individual's schooling raises his (her) wage offer by more than it raises the reservation wage. These results are strong statistically, and at the sample mean, they imply that the marginal effect of a one year increment to husband's schooling increases his probability of wage work by 1.2% and to his wife's schooling increases her probability of wage work by 3.8%. The results suggest that an increment to a wife's schooling causes the difference between her wage offer and reservation wage to increase by more than for her husband, and are consistent with results reported in a number of other studies (e.g., DaVanzo, Detray, & Greenberg, 1973).

Increments to schooling also have cross person effects. An increase in the husband's schooling causes a reduction in the probability that his wife participates in wage work. This can occur only if there is a simultaneous increase of her reservation wage. An increase in wife's schooling has a positive but not significantly different from zero effect

on her husband's probability of wage work. Thus, a wife's level of schooling does not appear to affect the reservation wage of her husband. Men are likely to work for a wage regardless of their wife's schooling level; e.g., over 90% of men under age 50 work for a wage (see Table V.1). Women are less likely to work for a wage than men, and many factors influence their decision, especially family characteristics.

The coefficient of the race variable is positive and statistically significant in the participation equation for both husbands and wives. Husbands and wives who are nonwhite are more likely to participate in wage labor than white husbands and wives, other things equal. At the sample mean, being nonwhite increases the probability of wage work by 3 percent for men and 13 percent for women. Thus, nonwhite women are much more likely to participate in wage work than white women, but nonwhite men are only slightly more likely to participate in wage work than white men.

When a couple has children less than 19 years of age, the probability of the wife participating in wage labor is reduced. The reduction in probability of participation occurs for both children 6 years of age or younger and children ages 6-18. The size of the reduction is 20 percent for each child in the youngest age group and 3.3 percent for each child in the older age group. Thus, the presence in a household of children age 6 years or less or age 6 to 18 raises the reservation wage of wife's time, but the increase in her reservation wage is larger per child for children 6 years of age or less. Although the coefficients on the KID6 and KID618 variables are negative in the

equation explaining the husband's probability of wage work, the coefficients are not significantly different from zero. These results are consistent with other studies.

These results do not conflict with the research of Gronau (1977). Comparative static results of his household production model suggest that the introduction of children into a household increases the productivity of home time, especially of a wife's home time. If leisure (not housework) is a normal good, more leisure is consumed, more time is spent working at home, and less time is spent working in the market. When home productivity increases, the opportunity cost of home time increases relative to the market wage. Gronau also argued that a comparative advantage in marriage--one person relatively more skilled in wage work and the other person relatively more skilled in home production--allows couples to collect gains from specialization. The strong marginal effect of children on women's participation and weak marginal effect on men's participation, along with the observation that women are the primary care-givers to children in most families, suggest that specialization in home and work time of couples results in the responsibility of child care being assumed to a large extent by women.

An increase of a household's asset income causes a reduction in the probability of wage work participation of a husband and wife. At the sample mean, a 1 percent rise in V causes a 0.5 percent reduction in the probability of a husband participating in wage work and a 1.5 percent reduction in the probability of a wife participating in wage work. Thus, an increase of asset income seems to cause a relatively larger rise in

the reservation wage of married women than of married men.

In the bivariate probit estimates of the participation equations, the coefficient of the state manufacturing wage is positive in the husband's equation but negative in the wife's equation. We believe that a rise in the manufacturing wage is an indication of generally higher wage rates. Furthermore, we believe that an increase of the wage offer of an individual increases their probability of wage work. In this case, a rise in the manufacturing wage can be interpreted as increasing the wage offers of both husbands and wives. If the rise in the husband's wage raises the reservation wage of the wife by more than her wage offer rises (a cross person effect), it would be possible for the coefficient of the manufacturing wage to be negative in the wife's labor participation equation. This outcome seems more likely when we recognize that a much larger share of employed men than employed women work in manufacturing.[3] Thus, a rise in the manufacturing wage can reasonably raise the expected wage of men by more than for women.

A higher unemployment rate causes a reduction in the probability that both married men and married women work for a wage, as expected. At the sample mean, a 1 percentage point rise in the state unemployment rate causes a 1.4 percent decline in the probability of wage work by married males and a 1.7 percent decline for married women. These effects are as expected. A higher than normal unemployment rate, however, has no

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[3]Approximately 21% of nonmetropolitan women and 29% of the men employed in nonagricultural industries work in manufacturing.

statistically significant effect on the probability of wage work, other things equal.

A more rapid growth in the total number of jobs in a state causes a reduction in the probability of married males being employed at wage work and tends to reduce the probability of employment of married women. Rapid employment growth, holding the manufacturing wage and unemployment rate constant, seems likely to indicate increases in the supply of labor to a state. Also, the occupations and industries that have experienced the most rapid growth during 1976-80 have been ones where a larger share of women than men are employed. Thus, the negative and significant effect of JOBGR on the probability of married males participating in wage work and negative but not significantly different from zero effect on probability of married women participating may be reasonable effects. The change in the share of a state's jobs in the service sector, however, did not have a statistically significant effect on the probability of participation of men or women.

The coefficients of the regional dummy variables provide estimates of broad regional effects that are not captured in the other regressors. These effects include regional differences in costs of living, commuting costs, and occupational-industrial mix of jobs. The results show that there are statistically significant regional differences in the probability of wage work. In the south and west, the probability of wage work is 16 to 20 percent lower for men than in the northeast and 11 to 18 percent lower for women. In the north central region, the probability of wage work is 12.5 percent lower for married men than in the northeast and

is 18.5 percent higher for married women. Of these regional effects, only the coefficient of DW in the participation equation for women is not significantly different from zero.

A time trend is included to pick up unmeasured effects that are correlated with time. All else equal, the probability that women participate in wage labor increased during 1978-82--the marginal effect is 1.6 percent per year. Time has no significant effect on the probability of male participation when men's and women's participation equations are estimated jointly.

Summary Because the disturbance terms of the wage-labor participation equations of husbands and wives were positively correlated, we may conclude that labor participation decisions of married couples are not independent and the correct way to model them is to estimate the equations for husbands and wives jointly. Human capital variables, race, and nonwage income were strongly significant in the probit equations. Additional children, both pre-school and school-age, decreased the probability of the wife's participation but had no statistically significant effect on the husband's work decision. This result is consistent with the findings of previous studies. Unemployment rates had the strongest effect on labor participation of the labor market variables.

Labor demand

This section focuses on wage labor demand functions of nonmetropolitan-nonfarm men and women. The empirical specification of

the labor demand equations for nonmetropolitan-nonfarm men and women (equation IV.1) were fitted to 24,571 observations on men and 17,508 observations on women. Human capital variables play a major role in determining wage offers (labor demand). Several state labor market variables are significant in determining labor demand, and the importance of the variables differs for men and women. The results of the wage equations are reported in Table V.4.

Results The positive but diminishing effect of an individual's age on wage offers of men and women has been reported in many studies. The age-log wage function of men is higher at every age than the age-log wage function of women. This result is expected because women are more likely than men to spend periods of time out of the labor force, and as a result they earn lower returns to experience (age) on average than men. Wages peak slightly later for men than for women, at ages 47.9 and 46.2, respectively. The marginal effect of age is .18% for women and .12% for men at mean age.[4]

An increment to a woman's schooling raises her wage offer by a larger percentage than an increment to a man's schooling--6.7 and 5.1 percent, respectively. This result is surprising until we recall that men earn higher wages on average than women ($3.48 vs. $2.00), so women actually experience a smaller absolute (dollar value) increase in wages from a one-year increment to schooling. A one year increment to men's education translates into a 17.7 cent increase in men's wages; a

[4]Mean age is 47 years for men and 43.9 years for women.

Table V.4. Econometric estimates of labor demand functions for married men and women in U.S. nonmetropolitan-nonfarm households, 1978-82[a,b]

	Men	Women
Intercept	-1.35 (15.17)	-.943 (11.49)
Age_j	.067 (17.49)	.036 (14.20)
$Age_j^2/100$	-.070 (13.48)	-.039 (12.19)
$Educ_j$	.051 (34.82)	.067 (27.63)
Race	-.202 (13.74)	-.073 (3.92)
Ln stw	.462 (12.66)	.004 (.07)
Urate	.002 (.83)	-.001 (.27)
Dif3	-.006 (1.12)	-.013 (2.00)
Jobgr	.130 (1.09)	-.127 (.84)
Service	.009 (3.66)	.008 (2.34)
DNC	-.114 (8.85)	-.051 (2.93)
DS	-.062 (4.74)	-.088 (5.08)

[a]Asymptotic t-ratios in parentheses.

[b]Dependent variable: ln (real wage).

Table V.4. (continued)

	Men	Women
DW	.036 (1.82)	-.031 (1.23)
Time	-.032 (9.08)	-.027 (5.65)
$\hat{\lambda}_j$	.075 (1.37)	-.039 (1.49)
R^2	.1648	.0774
N	24,571	17,508

one year increment to women's education increases women's wages by 13.4 cents. Topel (1986) reported a similar return to education for men (6-7%), and Gerner and Zick (1983) estimated a 7.4% return to education for married women.

Nonwhite men earn about 20% less than white men, all else equal, and nonwhite women earn about 7.3% less than white women. The race dummy variable was significant for men and women. Previous studies have shown large gaps in the wages of white and black men but little or no gap in white and black women's wages (Hamermesh and Rees, 1984). Here, we find a statistically significant difference in the wage rates of rural nonfarm white and nonwhite women, but the gap is much smaller for women than for men. Topel (1986) found an 18% difference in the wages of white and nonwhite men, a magnitude which is similar to the one reported here.

The state average manufacturing wage is used in the wage equations as a price indicator that controls for differences in general wage levels across states. Its sign was expected to be positive (see equation IV.1). A 10% rise in the manufacturing causes a 4.6% rise in the wage offer of men, but its effect on the wage offer of women is not significantly different from zero. This result lends support to the unexpected finding reported in the previous section that a rise in STW reduces the probability of women's labor force participation. It seems likely that the decreased probability of a women's labor participation is a result of the strong effect of STW on the husbands' wage offers.

Increases in a state's unemployment rate have little effect on labor demand for men or women; the coefficient of URATE is positive for men and

negative for women but neither is significantly different from zero. This is in contrast to the effects of the state unemployment rate on the probability of men and women participating in the labor force (see previous section). Our results are similar to those of Gerner and Zick (1983), who found that increases in state unemployment rates decrease the probability of married women participating in the labor force, but had no statistically significant effect on labor demand. They also reported that being in a rural area decreased wage rates but had no significant effect on the probability of being in the labor force. Unemployment rates appear to have more effect on wage-labor participation decisions than on wage offers.

Actual unemployment rates that are higher than "normal" have a dampening effect on wage rates, as expected, which is statistically significant for women but not for men.

Growth in a state's number of jobs tends to increase wage rates for men but not for women. Because the growth in employment may affect wages in several ways that move wages in opposite directions, the weak statistical result is reasonable. First, the number of jobs (or employment) observed in a state can be viewed as the result of the equilibrium condition between labor supply and labor demand. If labor supply shifts to the right faster than labor demand, then downward pressure on wage rates will occur. Second, because different sectors of the economy grow at different rates, different growth rates for low-paying and high-paying jobs in various industries adds further uncertainty to the effect of job growth on labor demand functions faced

by individuals. Third, the increase in relatively low-paid part-time jobs over the last decade depresses average wages, especially for groups most highly represented in part-time employment (i.e., women).

Increases in the share of a state's jobs that are in service industries increase wages for men and women, and the effect is statistically significant for both. The coefficients of SERVICE in the men's and women's demand functions are nearly the same magnitude--the marginal effect is .9% for men and .8% for women. Two factors associated with the increase in service employment, the fast growth in demand for services and the shift of labor into activities requiring specialized skills (Ott, 1987), may put upward pressure on wages in service industries.

Statistically significant unmeasured regional differences in wage offers relative to the northeast are present in the north central, south, and west regions for men and in the north central and south regions for women. Unmeasured regional effects depress men's wages by 11.4% and 6.2% in the north central and south regions, respectively, and raise men's wages by 3.6% in the west; women's wages are 5.1% and 8.8% lower in the north central and south regions than in the northeast, respectively, and the unmeasured effects on their wages in the west are not significantly different from zero.

Real wage rates for both men and women have a negative trend over the period 1978-1982--the real wage has declined 3.2% per year for men and 2.7% per year for women. Bils (1985) also reports a negative trend in the growth of the real wage over the period 1966-80, along with

negative unmeasured regional effects associated with the south.

The inverse mills ratio, $\hat{\lambda}$, was constructed from predicted values of the probability of wage-work participation obtained from bivariate probit equations (see Chapter IV). The coefficients of the sample selection terms are not significantly different from zero. Gerner and Zick (1983) also found that for a similar specification of the labor demand function for married women, the coefficient of the sample selection variable was not significantly different from zero.

Other specifications of the models were estimated by breaking a 10% subsample of the data into two groups by region, north central and south vs. east and west, and by breaking it into two groups by time, 1978-79 and 1981-82. No new insights were gained by breaking the sample into groups by time. Breaking the sample into regional groups showed that the unemployment rate was a relatively strong predictor of wages for the east-west group, and deviations of predicted from actual unemployment rates were important for the north central and southern group.

Summary Human capital variables have strong effects on the real wages of nonmetropolitan-nonfarm married men and women. Wage rates are strongly affected by an individual's age which is positively correlated with labor market experience (but is exogenous to that experience). Women experienced a higher rate of return to education then men, but because women earn lower wage rates than men, the actual (dollar value) increase in women's wages from one year of schooling is smaller than the increase in men's wages. Both nonwhite men and women earned lower wages than white men and women, all else equal.

Labor demand for both men and women have benefited from increases in the percentage of service jobs in the local (state) economy. Other labor market conditions have difference impacts on wage offers of men and women; for example, high state average manufacturing wages are associated with higher wages for men, and higher than normal unemployment rates tend to decrease wages of women. Real wages of both men and women have fallen over time due to trend effects.

Labor supply

The results of the labor supply equations demonstrate that the determinants of hours worked by nonmetropolitan-nonfarm men and women differ substantially by type--whether the husband or wife both work--of household. For example, race has significant effects on hours worked by both men and women in two-wage earner households, but little effect on the wage labor supply in single-earner households. Wage and income effects on labor supply for men and women are about the same across household types. Sample selection effects were significant in almost all of the labor supply functions.

The characteristics of husbands and wives and of their households differ by the type of wage work decision that they have reached. Couples in which only the wife works for a wage are substantially older, which suggests that retirement is the primary reason the husband does not work for a wage, have less education, fewer children, and more nonwage income than couples in which both spouses work for a wage. Husbands whose wives work for a wage work more hours than husbands whose wives do not

work; and wives who have working husbands work more annual hours in the market than wives whose husbands do not work for a wage. The average wage rates for women in the set whose husbands do not work for a wage and for the set whose husbands work for a wage are approximately the same, but men whose wives do not work earn substantially higher wage rates than men whose wives work ($3.92 vs. $3.23). A summary of the mean values of selected variables for each household type is given in Table V.5.

For reasons explained in Chapter IV, two estimates of wage labor supply equations were fitted for nonmetropolitan-nonfarm married men and women: (1) men (a) whose wives do not participate in wage work, and (b) whose wives participate in wage work; and (2) women (a) whose husbands do not participate in wage work, and (b) whose husbands participate in wage work.

The empirical labor supply equations (IV.2) are reported in Table V.6.[5,6] The equations for single earner households were fitted by two stage least squares with bivariate sample selection terms included. The equations for two wage-earner households were fitted by seemingly unrelated regression equations. Equations for two wage-earner households fitted by two stage least squares are presented in Appendix Table E.1. For husbands, 21 percent of the variation in ln hours was explained for

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[5]The two stage least squares estimates are consistent but not fully efficient (Greene, 1981).

[6]The independent variables of the labor supply equations were checked for multicollinearity. Age explained a great deal of the variation in age squared, but none of the other variables showed any evidence of high collinearity.

Table V.5. Descriptive statistics of family characteristics of married couples with one or more wage earners in U.S. nonmetropolitan-nonfarm households, by household type, 1978-82

	Household-type		
	Both spouses report wage work	Husband only reports wage work	Wife only reports wage work
Age - husband	39.3	44.8	53.4
Age - wife	36.6	42.2	48.3
Education - husband	12.4	11.7	10.4
Education - wife	12.4	11.4	11.5
Children <6	.34	.47	.15
Children 6-18	.80	.80	.46
Nonwage income	$257.52	$426.49	$420.42
Husband - hours	2073.7	1997.4	--
Husband - wage	$3.23	$3.92	--
Hours - wife	1448.7	--	1413.9
Wage - wife	$2.00	--	$1.99
N	15,502	9,069	2,006

Table V.6. Econometric estimates of wage labor supply functions for married couples in U.S. nonmetropolitan-nonfarm households, 1978-82[a,b]

	Husbands		Wives	
	Wife does not work for a wage[c]	Wife works for a wage[d]	Husband does not work for a wage[c]	Husband works for a wage[d]
Intercept	5.30 (8.45)	4.85 (10.19)	23.80 (9.87)	10.76 (10.80)
Ln V	.076 (1.11)	.170 (3.27)	-1.767 (6.71)	-0.441 (4.08)
$\widehat{\text{Ln } w_1}$	.411 (2.84)	.093 (1.25)		-.242 (3.02)
$\widehat{\text{Ln } w_2}$		-.242 (3.02)	3.369 (7.78)	1.299 (6.69)
Educh	.003 (0.40)	.009 (2.23)	-.019 (2.21)	.001 (0.16)
Educw	-.013 (2.40)	.011 (1.87)	-.177 (5.78)	-.048 (3.30)
Ageh	.056 (5.34)	.056 (8.61)	-.011 (0.80)	.035 (3.23)
$\text{Ageh}^2/100$	-.070 (6.03)	-.067 (8.66)	.003 (0.21)	-.047 (3.49)
Race	.015 (0.37)	-.131 (5.97)	.042 (0.49)	.138 (3.59)
Kid6	.123 (5.35)	.025 (3.01)	-.094 (1.36)	-.306 (18.11)
Kid618	-.008 (0.92)	-.010 (2.38)	-.114 (4.00)	-.121 (14.22)

[a]Asymptotic t-ratios in parentheses.
[b]Dependent variable: ln (annual hours worked).
[c]Fitted by two stage least squares.
[d]Fitted by seemingly unrelated regression equations.

Table V.6. (continued)

	Husbands		Wives	
	Wife does not work for a wage[c]	Wife works for a wage[d]	Husband does not work for a wage[c]	Husband works for a wage[d]
$\hat{\lambda}_1$	.152 (6.82)	-.203 (6.81)	.179 (3.52)	.002 (0.03)
$\hat{\lambda}_2$	-.352 (1.80)	1.480 (4.81)	-1.069 (2.44)	-.006 (-.01)
DNC	-.040 (1.80)	-.034 (2.63)	.150 (1.75)	.093 (3.46)
DS	-.002 (0.08)	.001 (0.09)	.211 (2.48)	.184 (5.58)
DW	-.089 (2.70)	-.050 (3.15)	.082 (0.69)	-.003 (0.09)
Time	-.022 (3.98)	-.030 (8.75)	.095 (5.77)	.008 (1.19)
R^2	.2118	.0926	.1111	.0700
N	9,069	15,502	2,006	15,502

the group that did not have working wives and 9 percent for husbands who had wives that participated in wage work. For wives, 11 percent of the variation in ln hours was explained for women whose husbands did not work for a wage and 7 percent for wives whose husbands worked for a wage.

Results The estimates of the real income constant own-wage elasticity of labor supply are positive as expected for both men and women.[7] For married men, the estimate of the compensated own-wage elasticity of labor supply is 0.411 when their wives do not participate in wage work and 0.093 when their wives participate in wage work. The compensated own-wage elasticity of labor supply for married women is much larger than for married men. For women whose husbands do not participate in wage work, the compensated own wage elasticity of labor supply is 3.369 and for other married women it is 1.299.

When husbands and wives both participate in wage work, the compensated cross-wage elasticities of labor supply provide evidence on

[7]The Slutsky equation for labor supply is

$$\partial H/\partial W = \partial H/\partial W \,|_{RIC} + H\ \partial H/\partial Y\ ,$$

where the term on the left side is the uncompensated wage effect on hours worked, the first term on the right side is the real income constant effect of an increase in the wage rate on hours of wage work which is restricted by theory to be positive, and the second term on the right side is the income effect of a wage increase on hours of wage work (Ashenfelter and Heckman, 1974; Pencavel, 1986). The Slutsky equation can be restated in terms of elasticities as:

$$\varepsilon_{HW} = \varepsilon^{RIC} + k_H\ \eta_{HY}$$

Because the empirical labor supply equations are in natural log functional form, the income and compensated wage elasticities are simply the coefficients of ln V and $\widehat{\ln w_1}$ ($\widehat{\ln w_2}$), respectively. The uncompensated wage elasticity is not directly observed but can be inferred using the Slutsky elasticity equation given above.

the substitute-complementariness of husband's and wife's nonwage work time. The labor supply equations were fitted with cross-equation symmetry conditions imposed on predicted wages. The two labor supply equations can give contradictory information about cross-wage effects if cross-equation restrictions are not imposed. In column 2, the coefficient of $\widehat{\ln w_2}$ is -0.242 which implies that husband's and wife's nonwage work time are substitutes.

The estimate of the income elasticity of labor supply is positive for the men which suggests that husband's nonwage work time is an inferior good in household consumption. For women, the estimate of the income elasticity of their labor supply is negative, as expected, which implies that their nonwage hours are a normal household consumer good. These results are consistent with a number of other studies.

An increase in the husband's schooling tends to increase his hours of work, other things equal including the predicted wage, whether or not his wife works. An increase in wife's schooling reduces her hours of wage work, other things equal. The own effects of schooling are statistically stronger for married women than for married men. There are cross-person effects of schooling, too. The effects of husband's schooling on his wife's wage labor supply are positive when he also participates in wage work. In contrast, when the husband does not participate in wage work, schooling has a negative and significantly different from zero effect on his wife's wage labor supply. A wife's higher schooling level tends to reduce the wage labor supply of her husband when she does not participate in the labor market. In contrast,

the effects of wife's schooling on her husband's wage labor supply are positive when she also participates in wage work.

The husband's age controls for many important life-cycle or life stage effects. Also, an older age is positively correlated with more wage-work experience for men and women. In these nonmetropolitan-nonfarm households, we obtain different age effects depending on the spouse's wage-labor participation status, which is not surprising because the signs of the human capital variables were unknown a priori (see equation IV.2). For men, age has a positive and diminishing marginal effect on labor supply irrespective of whether his wife is a wage work participant or not. At the sample mean both age effects are negative.

For women, age has a positive and diminishing marginal effect on labor supply when her husband is a wage work participant. When the husband does not work for a wage, an additional year of age has a negative but increasing marginal effect on her wage labor supply. At the sample mean, the age effect is negative irrespective of whether the woman's husband works for a wage or not.

If only the husband or wife participates in wage work, there is no significant difference in wage labor supply due to race (white vs. nonwhite). However, in households where the husband and wife both participate in wage work, nonwhite husbands work significantly fewer hours at wage work than white husbands (-13%) and nonwhite wives work significantly more at wage work than white wives (14%). The signs of the race variables were ambiguous a priori (see equation IV.2).

Additional children less than age 6 or age 6 to 18 cause an

economically and statistically different from zero reduction in wife's hours of wage work, as expected. In contrast, an additional child under age 6 increases husband's hours of wage work. For a married woman, the reduction in her annual hours of wage work per child under age 6 is much larger when her husband also participates in wage work, -30.6 percent versus -9.4 percent. For a married man, children under age 6 cause a larger increase in annual wage work when his wife does not participate in wage work--12.3 percent versus 2.5 percent. The effect on wife's annual hours of wage work due to children age 6 to 18 causes approximately the same size reduction in a wife's annual hours irrespective of her husband's wage work participation status. The reduction is about 12 percent. If a wife works for a wage, then the addition of children aged 6 to 18 causes a small (1.0%) but significantly different from zero decrease in husband's hours of wage work.

These results, together with the labor participation equation results, clearly demonstrate the enormous impact of children on wage work decisions of women. Not only do additional children reduce the probability that a woman participates in the labor market, it also reduces the annual hours of work for women who have chosen to work for wage. In contrast, additional children increase hours worked of husbands who participate in the labor force and have no significant effect on their labor force participation decision.

Regional differences in labor supply are larger and more significant for women than for men. Nonmetropolitan-nonfarm married men who reside in the north central and western regions work significantly fewer hours

at wage work than married men who reside in other regions of the U.S. The decrease is 4 to 9% if their wives do not participate in wage work and 3 to 5% when they do participate. Nonmetropolitan-nonfarm married women who reside in the north central and southern regions work significantly more hours than married women who reside in the northeast. For women whose husbands are also wage work participants, the increase is 18.4 percent for residence in the south and 9.3 percent for residence in the north central region. If a wife's husband is not a wage work participant, then the increase for each region is larger than the above estimates; e.g., for a married nonmetropolitan-nonfarm woman residing in the south and having a husband who does not participate in wage work, her annual hours of wage work are on average 21 percent larger than if she lives in the northeast.

Sample selection bias was statistically significant for the men and for women whose husbands do not participate in wage work. Selection bias of the spouse's labor participation decision was significant in each of the four groups.

Summary Husband's nonwage work time is an inferior good in household consumption and wife's nonwage work time is a normal good. Human capital effects are more highly significant for women than for men, and race plays a significant role in determining annual hours worked by husbands and wives in household with two wage earners. In households with children, men spend more hours and women spend fewer hours in wage work. Sample selection bias is significant for own and spouse participation decisions.

Wage Work Decisions of Farm Couples

In this section, estimated coefficients and the corresponding marginal probabilities of off-farm labor participation are presented and discussed. Personal characteristic variables, i.e., education, race, and the number of children in the household, performed well in explaining the probability of off-farm work while state-level labor market and farm profitability variables had statistically weaker effects overall.

Maximum likelihood estimates of the two univariate and one bivariate participation equations are reported in Table V.7, and Table V.8 contains the marginal probabilities.[8] In the CPS farm sample, 42.7% of the married men and 38.7% of the married women participated in off-farm wage work; 40.8% of the farm households reported no off-farm wage and salary income.

The hypothesis that the off-farm wage participation equations are independent for couples in the CPS farm sample is rejected. The value of p, which measures the correlation between the disturbance terms, is .264 for the jointly estimated off-farm labor participation equations of the husbands and wives. The t-ratio for the null hypothesis that $\rho = 0$, i.e., the two participation equations are independent, has a sample t-ratio of 12.3. The critical value of the t-statistic with 5815 degrees

[8]The age squared, growing degree days, and growing degree days times rain variables were rescaled so the bivariate maximum likelihood function would converge.

Table V.7. Probit estimates of off-farm labor participation equations for U.S. farm husband-wife couples, 1978-82[a]

	Men		Women	
	Univariate	Bivariate	Univariate	Bivariate
Intercept	3.89	2.23	4.25	4.08
	(2.95)	(1.77)	(3.16)	(3.12)
Age_h	.036	.033	.001	.002
	(4.13)	(3.18)	(0.13)	(0.20)
$Age_h^2/100$	-.065	-.063	-.034	-.035
	(7.41)	(7.29)	(3.80)	(3.71)
$Educ_h$	.023	.010	-.004	-.009
	(3.04)	(1.31)	(0.49)	(1.16)
$Educ_w$	-.026	-.027	.081	.080
	(2.96)	(3.09)	(9.03)	(9.22)
Race	.300	.320	.560	.406
	(2.86)	(2.94)	(5.34)	(3.82)
Kid6	-.055	-.033	-.403	-.391
	(1.66)	(1.03)	(11.49)	(11.58)
Kid618	-.024	-.040	-.074	-.073
	(1.42)	(2.40)	(4.22)	(4.43)
Ln v	-.515	-.344	-.568	-.535
	(3.82)	(2.66)	(4.13)	(4.01)
Ln stw	-.075	-.050	-.291	-.176
	(1.49)	(1.36)	(0.13)	(0.13)
Urate	.061	.045	.018	.009
	(3.83)	(2.70)	(1.11)	(0.57)
Dif3	-.008	-.007	-.046	-.042
	(0.31)	(0.25)	(1.71)	(1.50)
Jobgr	-.423	-.082	.383	.644
	(1.10)	(0.20)	(0.95)	(1.52)

[a]Asymptotic t-ratios in parentheses.

Table V.7. (continued)

	Men		Women	
	Univariate	Bivariate	Univariate	Bivariate
Service	-.021	-.020	.002	-.002
	(1.49)	(1.36)	(0.13)	(0.13)
Ln pc	.211	.384	-.020	.144
	(1.03)	(1.90)	(0.09)	(0.69)
Ln pl	.038	.037	.184	.038
	(0.18)	(0.18)	(0.86)	(0.18)
Ln pi	.223	-.685	-.420	-.331
	(0.44)	(1.24)	(0.82)	(0.58)
Ln pw	.791	.782	-.489	-.234
	(2.89)	(3.09)	(1.76)	(0.93)
Rain	.036	.035	.019	.017
	(4.70)	(4.59)	(2.38)	(2.21)
Gdd/1000	.269	.312	.167	.185
	(4.87)	(6.08)	(3.01)	(3.68)
Gddrain/10,000	-.080	-.079	-.042	-.037
	(4.59)	(4.54)	(2.41)	(2.16)
DNC	.023	-.125	-.089	-.141
	(0.25)	(1.04)	(0.93)	(1.22)
DS	.272	.134	-.205	-.265
	(2.54)	(1.02)	(1.91)	(2.08)
DW	.225	.226	.129	.008
	(1.79)	(1.58)	(1.02)	(0.05)
Time	-.020	-.009	-.035	-.022
	(0.75)	(0.34)	(1.32)	(0.09)
ln L	-3,594.7	-6,963.2	-3,485.4	-6,963.2
ρ	.264			
	(12.3)			

Table V.8. Estimates of marginal effects of explanatory variables on the probability of off-farm labor participation: U.S. farm, 1978-82

	Men	Women
Age_h	-0.0117	-0.0125
$Educ_h$	0.0038	-0.0034
$Educ_w$	-0.0103	0.0301
Race	0.1224	0.1526
Kid6	-0.0126	-0.1470
Kid618	-0.0153	-0.0274
V	-0.0103	-0.0158
Stw	-0.0063	-0.0216
Urate	0.0172	0.0034
Dif3	-0.0027	-0.0158
Jobgr	-0.0314	0.2421
Service	-0.0077	-0.0008
PC	0.3021	0.1113
PL	0.0260	0.0263
PI	-0.4885	-0.2319
PW	0.5897	-0.1733
Rain	0.0033	0.0063
Gdd	0.00001	0.00002
DNC	-0.0478	-0.0530
DS	0.0513	-0.0996
DW	0.0865	0.0030
Time	-0.0034	-0.0083

of freedom at the 1 percent significance level is 2.576. Thus, the null hypothesis is rejected; the disturbance terms are positively correlated.

A regression of the predicted values of the probability of off-farm wage work from the univariate probits on the predicted probabilities from the bivariate probits and a constant resulted in an R^2 of .97 for men and .99 for women. Although the bivariate model is statistically better than two univariate models of off-farm work participation, the predictions of the two procedures are very highly correlated.

Husband's age has a positive but diminishing marginal effect on the probability of off-farm wage work of husbands and wives. At the sample mean, the point estimate for a one year increase of age is -1.2 percent for husband's off-farm participation and -1.3 percent for the wife's. The marginal effect of age at the sample mean is to raise the opportunity cost of off-farm wage work by more than it raises the wage offer.

Own schooling effects are positive and significantly different from zero for farm women. An increase in an individual's schooling raises his (or her) market wage by more than it increases the opportunity cost of off-farm wage work. For husbands, the statistical significance of their schooling is, however, much stronger in the univariate than in the bivariate estimates. Cross-person effects of schooling are negative, as expected (see equation IV.4). An increase in a wife's (husband's) schooling causes an increase in her husband's (his wife's) opportunity cost of off-farm work. The negative cross person effects are statistically stronger for the effect of wife's schooling on husband's probability of wage work than vice versa. The coefficient of race is

positive and significantly different from zero in the off-farm participation equations of husbands and wives.[9] At the sample mean, nonwhite married males have a 12 percent higher probability of off-farm wage work than white married males and nonwhite married women have a 15 percent higher probability of off-farm work than white married women. Not only are nonmetropolitan-nonfarm nonwhite married couples more likely to participate in wage work than white couples, but nonwhite farm residents are more likely to take off-farm wage work than their white counterparts.

Additional children less than age 6 or age 6 to 18 reduce the probability of off-farm wage work of married men and married women. These results imply that children raise the opportunity cost of wage work for husbands and wives. The largest effect, however, is from an increase in the number of children under age 6; the marginal effect of a child under age 6 is to reduce the off-farm participation rate of wives by 14.7 percent. The marginal effects of older children are -1.5 and -2.7 percent per child for husbands and wives, respectively.

The coefficient of nonfarm asset income is negative and significantly different from zero in the participation equations of husbands and wives. Thus, an increase of nonfarm asset income is to raise the opportunity cost of wage work time. This is consistent with positive income elasticities of nonwage time.

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[9] A regression of the regional dummy variables on race explained only 2.8% of the variation in race.

The state labor market variables were weaker statistically than expected, but the null hypothesis that the coefficients of STW, URATE, DIF3, JOBGR, and SERVICE are jointly equal to zero was rejected at the 5 percent significance level.[10] The sample value of the χ^2 is 22.8 for male participation and 5.6 for female participation. The critical value of the χ^2 with 5 degrees of freedom at the 5 percent significance level is 11.1. Hence, the null hypothesis is rejected for male participation but not for female participation. An increase of the state manufacturing wage in expected to indicate a general rise of off-farm wage rates. Thus, we expected the coefficient of STW to be positive. The negative (although not significantly different from zero coefficients) are opposite expectations. This anomaly may be tied to the coefficient of the farm wage.

A higher state unemployment rate is associated with a higher probability of off-farm work for married farm males. The coefficient of URATE in the participation equation for married women is also positive, but it is not significantly different from zero. These effects of the unemployment rate are consistent with structural aspects of unemployment if married farm males are more likely to be employed in industries that have relatively high structural unemployment. When unemployment rates are higher than normal, suggesting a decline in expected off-farm wage

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[10]The labor market and farm profitability variables were checked for multicollinearity. Regressions of the independent variables on each other did not suggest that an extraordinary amount of the variation in any one variable was explained by the other explanatory variables.

offers, the probability of off-farm work of husbands and wives tend to be reduced. These results are, however, weak statistically.[11]

An increase in the rate of state job growth tends to cause a reduction in the probability of off-farm wage work of married males and to increase the probability of off-farm work of married females. These results are statistically stronger for female participation than for male participation. This would occur if the state job growth during 1976-80 was primarily in occupations-industries that employed women. An increase in the share of a state's jobs that are in the service sector tends to cause a reduction in the probability of off-farm work of married men and women.

The farm input and output price variables did not perform as well as expected. The estimated coefficients for price of crop output and price of livestock output and of purchased nonlabor inputs are opposite expectations. The fact that none of these estimated coefficients is significantly different from zero suggests that these prices do not affect the probability of off-farm work of farm couples. The coefficient of the wage rate for hired labor is positive in the husband's participation equation and negative in the wife's participation equation. A positive sign was expected because a rise in the wage for hired labor is

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[11]The joint hypothesis that the coefficients of the farm output and input prices were jointly equal to zero was tested. The sample value of the x^2 statistic was 10.2 for the husband's participation equation and 4.4 for the wife's participation equation. The critical value of x^2 with 4 degrees of freedom at the 5% level of significance is 9.5. Thus, the null hypothesis was rejected for husbands' off-farm participation decisions but not for wive's participation.

expected to reduce the profitability of farming. These results suggest that the opportunity cost of husband's time for off-farm work is reduced when PW increases. For wives, the results suggest that their opportunity cost of time for wage work increases when PW increases. This may be reasonable when the cross person effects of a change in PW are considered. When PW increases, the probability of wage work for husbands increases but for wives it decreases. The latter effect is, however, weak statistically. Hired farm labor appears to be a complement for the husband's farm labor and a substitute for farm labor of other household members.

The coefficients of agroclimatic variables are opposite from initial expectations. Within a census region, larger amounts of rainfall or a longer length of growing season is expected to increase the profitability of farming--at least of crop production. Crop production is not very labor intensive relative to livestock production. Thus, what these results seem to be telling us is that better agroclimatic conditions make crop production relatively more attractive compared to livestock production and accommodate off-farm wage work. At the sample mean, the marginal effect of an additional 3.6 inches of normal rainfall is 1.18% and of 336 growing degree days is 39%.

The probability of a married woman who resides in the south participating in off-farm work is 10% lower (statistically significant at the 5% level) than for other regions. Other regional effects represented by the regional dummy variables are not significantly different from zero. Over the period 1978-82, the probability of off-farm work has a slight negative trend, other things equal. The marginal effect of time

on the probability of wage work was -.3% for men and -.8% for women.

The absence of unmeasured regional effects on off-farm labor participation may be due to the presence of other variables that "capture" regional differences. In this case, normal rain and growing degree days are able to explain regional climatic differences. Regressions of normal rainfall (growing degree days) on the regional dummy variables resulted in 51% (63%) of the variation in rain (growing degree days) being explained. Climatic conditions are one major source of regional differences, but there are undoubtedly others.

The density of state population does not seem to affect off-farm labor participation (see Table V.9 below).[12] Population density could be correlated with unmeasured regional effects because most of the sparsely populated states were in the north central and western regions. However, higher population density does not seem to induce higher levels of off-farm participation, as one might expect, and our model correctly did not find strong regional differences that may have been due to that factor.

A strong positive correlation was present in the disturbance terms of the husband's and wife's labor participation equations which suggests that joint decision making is a correct assumption to make for farm

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[12]Sparsely populated states included Minnesota, Iowa, Kansas, North Dakota, South Dakota, Montana, Nevada, Utah, Colorado, New Mexico, an Arizona and contained 2,920 observations. Densely populated states included New York, New Jersey, Pennsylvania, Ohio, Indiana, Illinois, Michigan, Wisconsin, Maryland, Delaware, Virginia, West Virginia, North Carolina, south Carolina, Georgia, Florida, Kentucky, Tennessee, Alabama, Mississippi, Arkansas, Louisiana, Oklahoma, Texas, Washington, Oregon, and California and contained 2946 observations.

Table V.9. Proportion of married men and women reporting off-farm wage and salary income by population density of state, 1978-82

	Densely populated states		Sparsely populated states	
	Men	Women	Men	Women
1978	.51	.40	.50	.39
1979	.44	.40	.43	.40
1981	.46	.40	.40	.38
1982	.46	.38	.51	.42

couples. Human capital variables, as well as children and nonwage income, performed well in explaining off-farm labor participation decisions of husbands and wives. Labor market variables and farm profitability variables had weak effects, with few exceptions.

CHAPTER VI. AN ECONOMETRIC EXPLANATION FOR HOUSEHOLD CASH INCOME

The question of ultimate economic importance in this study is how human capital, general business cycles, local labor market conditions, and agricultural conditions affect household real cash income. This is one, but not the only, restriction on households' attempts to attain higher levels of welfare or utility.

The econometric results show that human capital variables are important determinants of household cash income for farm and nonmetropolitan-nonfarm households. Also, state labor market variables are shown to be important determinants of nonmetropolitan-nonfarm household cash income but not of farm household cash income. Farm input and output prices and climatic conditions are shown to be important determinants of cash income of farm households.

Determinants of Household Cash Income

The reduced form equations for household cash income are fitted to the CPS observations on husband-wife households in farm and nonmetropolitan-nonfarm populations and are reported in Tables VI.1 and VI.2, respectively. Because farm input and output prices and climatic conditions were hypothesized to affect farm household cash income but not nonfarm household cash income, more variables are included in the farm household income equation than in the nonfarm.

For the farm households, the CPS measure of farm income has some significant deficiencies. It fails to account for the value of farm

Table VI.1. Econometric estimates of reduced-form equations for U.S. farm household cash income, 1978-82[a]

	Regression equation			
	1	2	3	4
Intercept	5.12 (28.97)	5.14 (29.91)	5.11 (29.98)	5.14 (29.91)
Age	.014 (12.13)	.014 (12.50)	.014 (12.13)	.014 (12.49)
$Age^2/100$	-.015 (13.84)	-.016 (14.47)	-.015 (13.84)	-.016 (14.48)
$Educ_h$	.101 (9.16)	.010 (9.39)	.009 (17.15)	.009 (17.51)
$Educ_w$	.008 (6.90)	.008 (6.83)	.009 (17.15)	.009 (17.51)
Race	-.008 (0.54)		-.008 (0.55)	
Ln v	.490 (27.31)	.490 (27.38)	.490 (27.35)	.490 (27.43)
Kid6	-.035 (7.02)	-.035 (7.03)	-.035 (7.01)	-.035 (7.02)
Kid618	.001 (0.53)		.001 (0.55)	
Ln stw	.080 (2.81)	.104 (4.16)	.080 (2.81)	.104 (4.18)
Urate	.002 (0.82)		.002 (0.85)	
Dif3	-.005 (1.22)	-.004 (1.14)	-.005 (1.22)	-.004 (1.14)

[a]Asymptotic t-ratios in parentheses.
[b]Regression (1) all variables included; (2) best fit; (3) all variables, $\beta_{educh} = \beta_{educw}$; (4) best fit, $\beta_{educh} = \beta_{educw}$.

Table VI.1. (continued)

	Regression equation			
	1	2	3	4
Jobgr	-.024 (0.43)		-.024 (0.42)	
Service	.001 (0.39)		.001 (0.38)	
Rain	.004 (3.86)	.004 (4.17)	.004 (3.89)	.004 (4.17)
Gdd/1,000	.038 (4.79)	.033 (4.56)	.038 (4.84)	.033 (4.62)
Gddrain/10,000	-.0082 (3.26)	-.0064 (2.96)	-.0083 (3.30)	-.0065 (3.00)
Ln pc	.067 (2.24)	.088 (3.36)	.066 (2.23)	.088 (3.37)
Ln pl	-.091 (2.98)	-.109 (3.91)	-.091 (2.98)	-.109 (3.92)
Ln pi	.074 (1.01)	.135 (2.27)	.075 (1.02)	.137 (2.29)
Ln pw	.159 (4.03)	.159 (4.38)	.160 (4.06)	.160 (4.41)
Time	-.014 (3.66)	-.010 (3.73)	-.014 (3.66)	-.010 (3.70)
DNC	-.008 (0.57)		-.008 (0.58)	
DS	-.004 (0.26)		-.004 (0.27)	
DW	.014 (0.77)		.014 (0.77)	
R^2	.2468	.2458	.2468	.2457

Table VI.2. Econometric estimates of reduced-form equations for household cash income of U.S. nonmetropolitan-nonfarm households, 1978-82[a]

	Regression equation[b]			
	1	2	3	4
Intercept	-6.93	-6.93	-6.93	-6.92
	(24.72)	(24.72)	(24.72)	(24.72)
Age	.062	.062	.062	.062
	(50.59)	(50.59)	(50.60)	(50.60)
$Age^2/100$	-.069	-.069	-.069	-.069
	(57.00)	(57.01)	(57.00)	(57.01)
$Educ_h$	.041	.041	.041	.041
	(31.21)	(31.28)	(64.54)	(64.71)
$Educ_w$	.042	.042	.041	.041
	(26.69)	(26.69)	(64.54)	(64.71)
Race	-.130	-.130	-.130	-.130
	(10.31)	(10.32)	(10.31)	(10.31)
Ln v	1.490	1.488	1.490	1.490
	(48.90)	(48.91)	(48.90)	(48.91)
Kid6	-.076	-.076	-.076	-.076
	(12.84)	(12.85)	(12.86)	(12.86)
Kid618	.017	.017	.017	.017
	(5.05)	(5.05)	(5.04)	(5.04)
Ln stw	.169	.172	.169	.172
	(5.44)	(5.59)	(5.43)	(5.59)
Urate	-.011	-.011	-.011	-.011
	(4.88)	(5.65)	(4.89)	(5.65)

[a]Asymptotic t-ratios in parentheses.

[b]Regression (1) all variables; (2) best fit; (3) all variables, $\beta_{educh} = \beta_{educw}$; (4) best fit, $\beta_{educh} = \beta_{educw}$.

Table VI.2. (continued)

	Regression equation			
	1	2	3	4
Dif3	-.003		-.003	
	(0.70)		(0.71)	
Jobgr	.057		.056	
	(0.60)		(0.59)	
Service	.010	.010	.010	.010
	(4.70)	(4.65)	(4.71)	(4.66)
DNC	-.103	-.103	-.103	-.103
	(9.32)	(9.39)	(9.32)	(9.38)
DS	-.114	-.111	-.114	-.111
	(10.37)	(10.60)	(10.36)	(10.60)
DW	-.088	-.084	-.088	-.084
	(5.34)	(5.37)	(5.34)	(5.37)
Time	-.031	-.030	-.031	0.030
	(10.25)	(10.83)	(10.25)	(10.82)
R^2	.3493	.3490	.3493	.3490

products consumed at the site of production or to adjust for changes in farm inventories of livestock and grain (Harrington and Carlin, 1987). The estimate of mean farm family income obtained from the CPS is about 70 to 80 percent as large as mean farm family income estimated by the USDA.[1]

Farm household income

The fitted equation for farm household income performs relatively well (see Table VI.1).[2] Husband's age is shown to have a positive but diminishing marginal effect on the percentage increase in household cash income, as expected (see equation IV.6). Farm household cash income peaks where the husband's age is 43.8, and at the sample mean, a one year increment to husband's age results in a -.3 percent decrease in household income. These results are consistent with the many effects that are associated with age--experience, energy, hours of work, farm size, and off-farm wage rates.

Husband's and wife's schooling are shown to have similar effects on

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[1]CPS estimates of mean farm family income are from Farm Population of the United States, 1984, Current Population Report-Farm Population, Series P-27, No. 58, U.S. Dept. of Commerce, Bureau of the Census. USDA estimates of mean farm family income are from Income and Balance Sheet Statistics, 1983, USDA ERS ECIFS 3-3, September 1984. The ratio of CPS to USDA estimates was .715 in 1980, .769 in 1981, .809 in 1982, and .790 in 1983.

[2]Because farm total household income was negative for 280 households (5% of the sample), a constant was added to the dependent variable so logarithms could be used as the functional form. The dependent variable was ln(total household cash income + 20,000). Twenty thousand was added so the value for household income of every observation was much larger than zero. All of the independent variables listed in Table III.2 were included.

farm household cash income. In particular, the estimated coefficients are not significantly different. The sample F-value is .44 and the critical F-value with 1 degree of freedom at the 5% significance level is 3.54. Thus, for this sample, an increment of 1 year to husband's or wife's schooling increases farm household cash income by 3 percent. This equality of effects on cash income is somewhat surprising, given the evidence that women's schooling is generally valued at a lower level in wage work employment than men's schooling. In the case of farm households, the complexity of household and farm decisions and the opportunities for farm work as well as wage work for men and women seems to result in an increment to education of husbands and wives being equally valuable in generating household cash income. The equations in which the education coefficients of the husbands and wife are restricted to being equal are given in columns 3 and 4 of Table VI.1.[3]

For farm households, the coefficient of the dummy variable for race is negative, small in size, and not significantly different from zero. Thus, in these data, differences in farm household cash income that are associated with race are due to the nonrace variables; i.e., education, number of children, asset income, geographic region, etc. The large migration of blacks out of U.S. agriculture during the 1950s and 1960s may help explain why the nonwhite households that remain in agriculture are not any "different" from the white households for generating cash

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[3]The advantage of imposing the restriction is that the variance of the coefficients decreases when restrictions are imposed.

income.

Household nonfarm asset income is a direct component of household cash income. It also is one of the variables believed to drive labor supply decisions. Thus, the positive and significantly different from zero effect of V on farm household cash income is expected. The elasticity of total household cash income with respect to nonfarm asset income is .127.

Young children--age six or less--cause a reduction in farm household cash income. The reduction is 11.6 percent per child. Older children, however, do not have a statistically significant effect on farm household cash income. Although a larger number of older children cause a reduction in the probability that the husband and wife participate in off-farm work, older children can be expected to work themselves and make a direct contribution to household cash income. This includes both farm work and off-farm work. Supplementary results reported in Appendix Table F.1 show that the number of older children in a farm household has a positive and statistically strong effect on the probability that a farm household has off-farm wage or salary income earned by household members other than the husband or wife.

The primary effect of state labor market conditions on farm household cash income is through the wage rate in manufacturing and higher than normal unemployment rates. The coefficient of the state manufacturing wage is positive which suggests that higher state nonfarm wage rates do result in larger farm household cash income. These effects are believed to be derived primarily from the effect of STW on the

probability of off-farm work or the wage rate when off-farm work occurs, especially for married farm males. A higher than normal state unemployment rate tends to reduce farm household income. This effect is believed to be a result of a reduced probability of off-farm wage work. This latter effect is, however, significantly different from zero at the 15% level. The other state labor market variables--the state unemployment rate and the change in the share of the jobs that are in the service industries--do not have a significant effect on farm household asset income.

Geoclimatic conditions and farm input and output prices have statistically significant effects on household cash income. The estimated coefficient for normal rainfall is positive; and an increase of 1 inch per year in normal rainfall increases farm household cash income by .42 percent. Extending the length of the crop growing season as represented by GDD has a positive but diminishing marginal effect on household cash income. At the sample mean, a 334 unit (10%) increase in annual normal GDD increases farm household cash income by .29 percent. The coefficients of both rain and gdd were expected to be positive (see equation IV.6).

The coefficient of the crop output price is positive and of the livestock output price is negative in the farm household cash income equation. Both are significantly different from zero at the 5% level. The negative coefficient on the livestock output price can be explained by the behavior of inventories when the price rises (Rosen, 1987). An increase in demand and market price for livestock causes farmers to

initially withhold female animals from the market so that the inventory of breeding stock can be built up. Thus, a rise in the livestock output price can actually cause a reduction in household cash income in the short run.

The coefficients of the prices of farm input are also positive in the household cash income equation. These coefficients were expected to be negative. The positive sign of the coefficient for the farm wage can be reconciled if it represents a good proxy for the marginal value of the time of some household members--e.g., older children, the wife, or the husband. If this was the case, then an increase in PW would represent an increase in the marginal value of farm household members' time and an increase in the cost of one of the farm inputs. The net effect of these opposing forces on farm household cash income could very well be positive. We, however, do not have a good explanation for the positive coefficient on the price of other variable inputs (PI).

Although three regional dummy variables were included in the farm household cash income equations, none of the coefficients was significantly different from zero at the 5% significance level. Thus, these results show that there is no pure regional effect on farm household cash income that is not represented by other variables that are included in the cash income equation. These results have the somewhat surprising implication that farm household cash incomes are equally in equilibrium across all four geographic regions. Low farm household income is the result on average of other real characteristics.

The results do show a strong negative time trend in real cash income

during 1978-82. The decline has been at a rate of about 1 percent per year on average.

A few other specifications should be noted here. Although some economists argue that nonwage income and the number of children are choice variables, the hypothesis of exogeneity of these variables could not be rejected at the 5% level of significance (see Appendix G). Therefore, they were included as regressors. The coefficients of the other exogenous variables changed very little, however, when nonwage income and children were excluded as regressors. The inclusion of these variables increased the explanatory power of the equations substantially. (See Appendix Tables G.1 and G.2.)

In addition, the farm group was divided into two subsets based on population density of the states. More of the variation in total household income was explained for households in the densely populated states (28% vs. 19% in the sparsely populated states). The average value of the residuals was -.0011 for the densely populated states and .0047 for the sparsely populated states. These equations are reported in Appendix Table H.1.

Nonmetropolitan-nonfarm household income

The set of variables representing household-individual and state labor market characteristics explains a relatively large share of variation in nonmetropolitan-nonfarm household cash income for 1978-82. This sample consists of all 32,588 nonmetropolitan-nonfarm husband-wife households in the CPS for 1978, 1979, 1981, and 1982. The dependent

variable is the natural logarithm of annual total household cash income deflated by the CPI (1967-100). The econometric estimates of the nonmetropolitan-nonfarm household cash income equations are reported in Table VI.2.

Husband's age is a very important indicator of experience and life stage effects in nonfarm as well as in farm households. A one year increase in husband's age has a positive but diminishing marginal effect on the percentage increase in household cash income. Nonfarm household cash income peaks at a husband's age of 44.9 years. Farm household cash income peaked at a slightly younger age (43.8 years). At the sample mean age of husbands (47 years), an increment to his age is negative, -.4%. These age effects are very strong statistically.

The estimated coefficient of husband's and wife's schooling are surprisingly similar in size. Furthermore, a statistical test of the null hypothesis that they are equal could not be rejected at the 1 percent significance level. The sample value of the F-statistic was 0.16 and the critical value with 1 degree of freedom at the 1 percent significance level was 6.63. When the household cash income equation was refitted with the equality restriction imposed, the standard error of the estimated coefficient decreased substantially.[4] This finding of equal effects of husband's and wife's schooling on household cash income is somewhat surprising, given the relatively low value placed on women's

[4]The regressions with the restriction that imposes equality of the husband's and wife's education coefficient are given in columns 3 and 4 of Table VI.2.

schooling in the labor market. These results suggest that wife's schooling has very important effects on household cash income that go far beyond its effect on market wage offers of women. A one-year increase in husband's or wife's schooling increases household cash income by 4.1 percent.

For nonfarm households, those with a nonwhite head of household receive less cash income than white households. The point estimate of the difference is 13 percent, and it is significantly different from zero at the 1 percent significance level. In contrast, race had no significant effect on household cash income of farm households. Thus, holding household-individual and state labor market variables constant, there is a separate and significant effect of race on household cash income for nonmetropolitan-nonfarm households.

Household asset income has a positive effect on household cash income, with an elasticity of 0.067. This effect includes both the direct contribution that V has to cash income and the indirect contribution through effects on labor supply of husbands, wives, and other household members.

Young children cause a decrease in household cash income--7.6 percent per child age 6 or younger; older children increase household cash income--1.7 percent per child ages 6 to 18. The effects of young children on household cash income are registered primarily through labor supply decisions of husbands and wives. The older children affect household cash income in an additional way because they can work for a wage themselves. Supplementary regression results reported in Appendix

Table E.2 show that children ages 6 to 18 have a positive and statistically different from zero effect on the probability that nonmetropolitan-nonfarm households have wage and salary income that is earned by household members other than the husband and wife.

State labor market variable are shown to have statistically significant effects on household cash income of the nonmetropolitan nonfarm households. A 10 percent increase in the state average wage rate in manufacturing--which signals a generally higher wage structure--raises household cash income by 1.7 percent. A higher state unemployment rate causes nonmetropolitan-nonfarm household cash income to decrease. Higher than normal unemployment rates, however, do not seem to have any statistically significant effect on household cash income. The latter two results differ substantially from the ones obtained for farm households. An increase in the share of the state's jobs that are in service industries causes an increase in household cash income. The primary reason for this latter effect seems to be the statistically strong positive effect that SERVICE had in the labor demand functions for both husbands and wives in the nonmetropolitan nonfarm households (see Table V.4).

In contrast to the results for farm households, there are statistically significant regional effects on nonmetropolitan nonfarm household cash income. Other things equal, household cash income is largest in the northeast region. It is 10 to 11 percent lower for households that reside in north central and southern regions and 8 percent lower for households that reside in the western region. Purely

regional effects seem to exist for nonmetropolitan-nonfarm households. Some of the differences are undoubtedly due to cost of living differences, but there also may be other reasons.

In other specifications, the nonmetropolitan-nonfarm sample was broken down by regions (east and west vs. north central and south) for a 10 percent subsample of the data. Not much was gained by using this specification. Race has a larger impact on household income in the east and west than in the north central and south. Income decreased significantly due to time trends only in the north central and southern regions.

Comparison of Farm with Nonmetropolitan-Nonfarm Total Household Income

Differences in the farm and nonmetropolitan-nonfarm total household income equations may arise for the following reasons: (1) there are many more observations in the rural nonfarm sample than in the farm sample (32,588 vs. 5,866); (2) more variables are used to estimate farm total household income, such as the farm output and input price indices and climatic variables; and (3) there are real differences in how farm and rural nonfarm households respond to local labor market conditions and how they benefit from human capital investments. For purposes of discussion, the full model for each group will be used for comparison.

Different dependent variables are used in the farm and rural nonfarm household income equations, so the coefficients are not directly comparable. The dependent variable in the farm equation is ln(total

household cash income + 20,000) while it is ln(total household cash income) in the nonmetropolitan-nonfarm equation. To compare the results of the two types of households in a meaningful way, the percentage change in household income due to one unit increases in the independent variables was calculated and reported in Table VI.3.

A one-year increase in husband's schooling increases farm household income by 3.3% and rural nonfarm household income by 4.1%. Nonmetropolitan-nonfarm households appear to reap higher returns than farm households from their educational investments. Marginal effects of husband's age are surprisingly similar at mean age for farm and nonfarm households. Mean age is, however, higher for the farm population.

The income of nonwhite farm households is 2.7% less than that of white farm households (not significantly different from zero); the income of nonwhite nonmetropolitan-nonfarm households is 13% less than that of white nonfarm households. There appears to be a much larger discrepancy in income by race for the nonfarm sample.

The local labor market variables have much more explanatory power in the equations explaining nonfarm household than farm household cash income. Farm household income increases by 2.9%, and rural nonfarm household income increased by 1.8%, for every 10% increase in the state average manufacturing wage. A 10% increase in the wage is about 30 cents. A 1% increase in the unemployment rate increased farm household cash income by .7% but decreased nonmetropolitan-nonfarm household cash income by 1.1%. The negative time trend in real cash income of farm households was larger than for nonfarm households (4.6% per year vs. 3.1%

Table VI.3. Comparison of effects of independent variables on farm and rural nonfarm household incomes

Variable	Units	% change in income for a one unit change in the independent variable[a]	
		Farm	Rural nonfarm
Age	Year	-.003	-.004
Education	Year	.030	.041
Race	0-1	-.027	-.130
Nonwage income[b]	$	.127	.067
Kid6	1,2,...	-.116	-.076
Kid618	1,2,...	.033	.017
Stw[b]	$	.265	.169
Urate	%	.007	-.011
Dif3	Δ%	-.017	-.003
Jobgr	%	-.079	.057
Service	Δ%	.003	.010
Time	Year	-.046	-.031
DNC	0-1	-.026	-.103
DS	0-1	-.013	-.114
DW	0-1	.046	-.088

[a]Evaluated at mean values.

[b]Converted to 1% change in the independent variable.

per year).

Farm total household income decreases by 11.6% for every child in the family under age 6, and increases by 3.3% for every child aged 6 to 18. In contrast, the addition of young children in nonfarm households decreases household income by 7.6% per child, and the addition of older children increases nonfarm household income by a smaller percentage per child than in farm households (1.7%). Children appear to have larger impacts on cash income in farm households than in nonfarm households.

The econometric reduced form equations for household cash income had greater explanatory power for nonfarm than farm households--R^2 of .35 vs. .25. This difference is consistent with the evidence that net farm income has a relatively large transitory component in any given year (Friedman, 1957). Measurement error and differential underreporting may also be more serious in the CPS measure of household cash income for farm than nonfarm households.

Farm and nonmetropolitan-nonfarm households may differ in other ways. We tend to think of farm households being relatively more geographically immobile than nonfarm households. The reasons are land ownership and high costs of farm relocation. The econometric results did not, however, show significant regional differences in farm household cash income. Our measures of farm prices and state labor market conditions did not perform very well in explaining wage work participation or household cash income. The reasons for this are not obvious.

The results from this study suggest that the state labor market

variables do a fairly good job of capturing the effects of local and general business cycle effects on rural household decisions affecting household cash income.

CHAPTER VII. SUMMARY AND CONCLUSIONS

Chapter VII presents a summary of the important results reported in this study and discusses their policy implications. The summary focuses on the empirical equations for labor supply, labor demand, labor participation, and household income of married couple nonmetropolitan-nonfarm households, and for off-farm labor participation and household income of married couple farm households.

Summary

Nonmetropolitan-nonfarm households

An important issue addressed in this research concerned the joint estimation of wage-labor participation equations of husbands and wives. The positive and significant correlation of the disturbance terms in the equations showed clearly that wage work decisions of married couples respond similarly to common economic shocks and therefore are not independent.

Human capital variables have strong effects on the labor market outcomes of nonmetropolitan-nonfarm households. Schooling in particular was an important explanatory variable in each of the wage work outcomes of married couples. Because the effect of an increment to schooling is to raise an individual's wage offer, both men and women with higher levels of schooling are more likely to participate in the labor force and receive higher wage rates than persons with lower levels of schooling. Furthermore, increments to education have strong cross-person effects in labor participation equations--for example, an increase in husband's

schooling decreases the probability of wife's wage labor participation by raising her reservation wage. Age had a positive and diminishing effect on wage offers and labor participation decisions of husbands and wives.

Nonwhite nonmetropolitan households receive significantly less cash income than white households. Both nonwhite men and women earn lower wages, and nonwhite married men and women in two-wage earner households work fewer hours in the market than their white counterparts. The probability of participation in wage work by married men and women, however, is higher for nonwhites than for whites, all else equal.

Wage and income effects on labor supply of nonmetropolitan-nonfarm men and women were about the same across household types. Husband's leisure is an inferior consumption good to nonfarm households, as indicated by the positive income elasticity of husband's labor supply. The negative income elasticity of wife's labor supply suggests that wife's leisure is a normal good to nonfarm households.

Women who have children under age 6 are less likely to work for a wage, and work fewer hours, than women who do not have preschool age children. Although children do not have a statistically significant effect on the probability of husband's wage work, additional young children do reduce his annual hours of wage work. The presence of children under age 6 decreases total household cash income. The number of children aged 6 to 18 has effects on labor market behavior that are similar to those of preschool children, but they also cause household cash income to be larger. The reason is that older children participate in wage work.

State level labor market variables as a whole performed better in the household cash income equation than in the labor demand decision equations. For example, increases in the unemployment rate put downward pressure on household income, while wage offers were seemingly unaffected. Participation rates of both men and women, however, declined significantly when unemployment rates rose. Actual unemployment rates that are higher than normal put downward pressure on the wage offers of women, but had little effect in household income or wage labor participation.

Increases in state manufacturing wages raised wage offers for men but decreased female participation rates. The cross person effect of higher wage rates for men raising the reservation wages of their wives is a likely cause for the different response to state manufacturing wages by men and women. Wage offers of men and women, as well as total household cash income, increased when the percentage of jobs in the service sector rose.

Unmeasured regional effects were detected in the labor demand equations of both men and women and in the household income equations. Individuals who had a residence in the north central and south regions tended to receive lower wage offers and their household cash income was lower than residents of the northeast region. Labor supply was not affected by geographical location of married men and women.

Real household cash income and real wage rates of men and women fell during 1978-82. Real wage rates of men and women have fallen 3.2% and 2.7% per year, respectively, and household income declined 30% per year

over the period 1978-82.

Farm households

Time and resource allocation decisions are far more complex for farm households than for households comprised solely of wage earners. For this reason, the results of the off-farm wage work and farm household cash income equations are statistically weaker, and less easily interpreted, than the results of the various wage work decision models of nonfarm households. Efficiency effects of education, for example, enhance productivity of farm, household, and wage work time. Hence, it is difficult to identify separately the various efficiency effects.

Equations explaining off-farm wage work participation of married men and women should be estimated jointly. The disturbance terms of the off-farm labor participation equations of husbands and wives were positively correlated, as they were in the wage-labor participation equations of rural nonfarm married couples.

Variables representing effects of human capital performed well in explaining both household cash income and off-farm wage work decisions of married couples. Husband's age has a positive but diminishing effect on his and his wife's off-farm wage work participation and on household cash income. Increments to schooling increase the probability of off-farm work for husbands and wives, and schooling of both individuals raise household cash income. Cross-person effects of schooling on the probability of off-farm work are negative; for example, an increase in wife's schooling increases her husband's opportunity cost of wage work

and decreases the probability he will work off-farm.

Children raise the opportunity cost of wage work for husbands and wives and lower the probability that they participate in off-farm work. The effect is much larger for young children. Household cash income is reduced by the effects of preschool children and increased by older children, who may contribute directly to household cash income by working on the farm or at off-farm work.

Farm household decisions on wage work and household cash income were shown to be largely unaffected by state labor market variables. An increase of state unemployment rates is associated with increased probabilities of off-farm work by married men, and higher than normal unemployment rates tend to reduce household cash income. Higher state manufacturing wages result in larger farm household cash income.

Off-farm work participation is much less affected by farm output and input prices than expected. Furthermore, the effects of the farm wage rate seem to be more in line with the opportunity cost of farm household members' time at wage work than with the cost of farm labor. This would be reasonable if farm family and hired farm labor are homogeneous inputs in farm production. The effects of farm output and input prices on household cash income were also full of surprises. Climatic variables, rain and growing degree days, raise farm household cash income by enhancing the profitability of farming. They tend to encourage off-farm labor participation by making the relatively less labor intensive crop production more attractive relative to livestock production. Farm input and output prices had stronger effects on household cash income than on

off-farm labor participation.

Conclusions

A number of options are available to policy makers who are concerned with rural development issues. The strategies they choose in promoting rural development will depend on the relative importance and impact of human capital, business cycles, and labor market conditions on labor market decisions of rural residents. The results of this analysis offer some insights into the determinants of labor market outcomes in rural households.

First, wage-labor decisions of husbands and wives are affected by common economic shocks. To obtain a better econometric fit of labor participation or off-farm wage work equations of married couples, the participation equations of husbands and wives should be estimated jointly.

Second, in both nonfarm and farm households, higher levels of schooling of husbands and wives are associated with higher levels of household cash income. The effect is larger for nonfarm than for farm households. These results imply that human capital investments are very important factors in enhancing welfare of rural households, and public policies should encourage job training programs and better educational programs in rural areas.

Third, state labor market conditions are also important determinants of household welfare at least for nonmetropolitan-nonfarm households. Higher employment rates decrease the probability of labor participation

and reduce household income of nonmetropolitan-nonfarm residents. Household cash incomes increase when state manufacturing wages increase. Increases in the percentage of service jobs have benefited nonfarm households by increasing household income. Growth in service industries has increased income because the demand for services, many of which require specialized skills, has grown rapidly and pulled up wages. Employment growth has ambiguous effects on labor market behavior in rural households. These results suggest that plans to diversify rural economies away from manufacturing and extractive industries into service industries may be warranted.

In short, rural development planners will need policies targeted at both individual skill requirements and general economic growth, since both personal and market characteristics were shown to have significant impacts on household cash income and wage work decisions in rural areas.

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APPENDIX A. UTILITY MAXIMIZATION FOR WAGE EARNERS AND FARM OPERATORS

Wage Earners

The time constraint faced by wage earners can be represented as

$$T_i = L_i + T_{iw}, \qquad i = 1, 2,$$

where: T_i = individual i's total time,

T_{iw} = time spent in wage work, and

L_i = leisure time.

In addition, this household faces a full income constraint of

$$F = V + w_i T_i = P_1 X_1 + w_i L_i ,$$

where: V = unearned income,

P_1 = the price of market goods, and

w_i = individual i's wage rate.

Maximizing utility subject to the constraints,

$$\ell = U(L_i, X_1; X_2) + \lambda[V + w_i(T_i - L_i) - P_1 X_1] ,$$

and the resulting first-order conditions are:

(A.1) $U_L - \lambda w_i = 0,$

(A.2) $U_{x_1} - \lambda P_1 = 0,$ and

(A.3) $V + w_i(T_i - L_i) - P_1 X_1 = 0 .$

Equations (A.1)-(A.3) can be solved for the optimal levels of L_i,

X_1, and T_{iw}:

$$L_i^* = L_i(P_1, w_i, V),$$
$$X_1^* = X_1(P_1, w_i, V), \text{ and}$$
$$T_{iw}^* = T_i - L_i^* = T_{iw}^*(P_1, w_i, V) .$$

Comparative statics can be obtained by setting the farm production variables equal to zero in the general case presented in the theoretical section.

Farm Operator

The farm operator faces a time constraint of

$$T_i = L_i + T_{if}, \qquad i = 1, 2,$$

where: T_{if} = time spent on farm labor and management.

The farm operator's income constraint is:

$$F = V + P_q Q = P_1 X_1 + rK + P_2 Y_2 ,$$

where: P_q = price of farm output,

Q = farm output,

r = cost of capital,

K = capital, and

$P_2 Y_2$ = total variable cost of farm output.

In addition, the production function is:

$$Q = Q(T_{if}, Y_2, K, Y_1),$$

where: Y_1 = inputs that are exogenous to current production decisions.

Combining these constraints with the utility function results in:

$$\ell = U(L_i, X_1; X_2) + \lambda[V + P_q Q(T_{if}, Y_2, K; Y_1) - P_1X_1 - P_2Y_2 - rK] .$$

The first-order conditions for utility maximization are:

(A.4) $U_L + \lambda P_q(dQ/dT_{if})(dT_{if}/dL_i) = 0,$

(A.5) $U_{X_1} - \lambda P_1 = 0,$

(A.6) $\lambda[P_q f_K - r] = 0,$

(A.7) $\lambda[P_q f_{y2} - P_2] = 0,$ and

(A.8) $V + P_q Q - P_1X_1 - P_2Y_2 - rK = 0 .$

Resulting input demands are:

$$Y_2^* = Y_2(P_q, P_2, Y_1),$$
$$K^* = K(P_q, r, Y_1),$$
$$X_1^* = X_1(P_q, P_1, P_2, r, V, Y_1, X_2),$$
$$L_i^* = L(P_q, P_1, P_2, r, V, Y_1, X_2), \text{ and}$$
$$T_{if}^* = T_i - L_i^* .$$

Comparative statics results are similar to the general model presented in the theoretical section except that the wage work variables (T_{iw}'s) are equal to zero.

APPENDIX B. TOTAL DIFFERENTIATION OF THE FIRST-ORDER CONDITIONS

(1) $U_{L_1L_1}dL_1 + U_{L_1L_2}dL_2 + U_{L_1X}dX - \lambda_2 dw_1 - w_1 d\lambda_2 = 0$

(2) $U_{L_2L_1}dL_1 + U_{L_2L_2}dL_2 + U_{L_2X}dX - \lambda_2 dw_2 - w_2 d\,\lambda_2 = 0$

(3) $U_{XL_1}dL_1 + U_{XL_2}dL_2 + U_{XX}dX - \lambda_2 dP_1 - P_1 d\lambda_2 = 0$

(4) $\lambda_1[G_{QQ}dQ + G_{T_{1f}Q}dT_{1f} + G_{T_{2f}}dT_{2f} + G_{YQ}dY + G_{KQ}dK]$

$+ G_Q d\lambda_1 + P_q d\lambda_2 + \lambda_2 dP_q = 0$

(5) $\lambda_1[G_{QT_{1f}}dQ + G_{T_{1f}T_{1f}}dT_{1f} + G_{T_{2f}T_{1f}}dT_{2f} + G_{YT_{1f}}dY + G_{KT_{1f}}dK]$

$- w_1 d\lambda_2 - \lambda_2 dw_1 + G_{T_{1f}}d\lambda_1 = 0$

(6) $\lambda_1[G_{QT_{2f}}dQ + G_{T_{1f}T_{2f}}dT_{1f} + G_{T_{2f}T_{2f}}dT_{2f} + G_{YT_{2f}}dY + G_{KT_{2f}}dK]$
$- w_2 d\lambda_2 - \lambda_2 dw_2 + G_{T_{2f}}d\lambda_1 = 0$

(7) $\lambda_1[G_{QY}dQ + G_{T_{1f}Y}dT_{1f} + T_{T_{2f}Y}dT_{2f} + G_{YY}dY + G_{YK}dK]$

$- P_2 d\lambda_2 - \lambda_2 dP_2 + G_Y d\lambda_1 = 0$

(8) $\lambda_1[G_{QK}dQ + G_{T_{1f}K}dT_{1f} + G_{T_{2f}K}dT_{2f} + G_{YK}dY + G_{KK}dK]$
$- r d\lambda_2 - \lambda_2 dr + G_K d\lambda_1 = 0$

(9) $G_{T_{1f}}dT_{1f} + G_{T_{2f}}dT_{2f} + G_Y dY + G_K dK + G_Q dQ = 0$

(10) $w_1 dT_1 + T_1 dw_1 + w_2 dT_2 + T_2 dw_2 + dV + P_q dQ + Q dP_q$

$- r dK - K dr - P_1 dX - X dP_1 - P_2 dY - Y dP_2 - w_1 dT - w_1 dL_1$

$- w_2 dT - w_2 dL_2 - (T_{1f} + L_1)dw_1 - (T_{2f} + L_2)dw_2 = 0$

$$d\phi = (T_1 - T_{1f} - L_1)dw_1 + (T_2 - T_{2f} - L_2)dw_2 + dV$$
$$+ QdP_q - Kdr - XdP_1 - YdP_2$$

$$R = w_1(L_1 + T_{1f}) + w_2(L_2 + T_{2f}) + P_1X + rK + P_2Y$$
$$= w_1T_1 + w_2T_2 + V + P_qQ$$

APPENDIX C. PREDICTED UNEMPLOYMENT RATE EQUATIONS, BY STATE

$$U_t = a + b*time + c*U_{t-1} + d*U_{t-2} + \varepsilon_t$$

State	a	b	c	d	R^2	F-test
AL	2.461*	.570*	.768*	-.655*	.8559	***
AZ	6.503**	-.011	.529	-.425	.2608	
AR	4.788**	.462*	.240	-.432	.6291	**
CA	7.881***	-.025	.778**	-.739**	.5590	**
CO	4.212**	.187	.187	-.223	.3275	
CN	10.447***	-.255*	.557*	-.795***	.7250	**
DE	4.739**	-.008	.800**	-.415	.4561	
FL	5.407**	-.011	.762**	-.512	.4416	
GA	4.283**	.073	.749**	-.527	.4720	
ID	3.727*	.242	.380	-.221	.5886	*
IL	2.916*	.561**	.311	-.383	.7665	***
IN	3.367*	.512	.438	-.459	.6956	**
IA	1.099	.272*	.895**	-.569*	.9003	***
KS	2.491*	.148	.389	-.247	.5648	*
KY	2.002	.519**	.377	-.308	.7624	***
LA	3.988*	.213*	.872*	-.627	.7827	***
MD	5.402***	.280*	.621*	-.908**	.6745	**
MA	7.293**	-.202	.680*	-.465	.5299	*
MI	5.921**	.519	.508	-.530	.6071	**
MN	3.692**	.193	.524	-.501	.5481	*
MS	2.617	.616*	.290	-.335	.7623	***
MO	3.707**	.350*	.535	-.602	.6806	**
MT	4.532**	.065	1.048**	-.826*	.7001	**
NB	2.214*	.124	.455	-.271	.5073	
NV	5.817**	.009	.988***	-.996**	.6651	**
NJ	5.530**	-.034	.823**	-.506	.4745	
NM	6.936**	.111	.555	-.600	.4173	
NY	5.766**	.032	.914**	-.690**	.5980	*
NC	4.384*	.293	.291	-.412	.4326	
ND	3.104*	.072	.349	-.130	.3056	

Note: All equations were tested for autocorrelation with the Durbin h test. Only one, CN, showed signs of autocorrelation.

*Significant at the 10% level.
**Significant at the 5% level.
***Significant at the 1% level.

State	a	b	c	d	R^2	F-test
OH	3.831**	.472*	.638*	-.659*	.7493	***
OK	3.830	.161	.480	-.522	.4000	
OR	5.985***	.200	.895***	-.809***	.7676	***
PA	5.35**	.510**	.490	-.769*	.7712	***
SC	4.919**	.288	.664*	-.724**	.6464	**
SD	1.855*	.093	.984**	-.651	.7626	***
TN	3.246**	.495*	.553	-.601	.7314	**
TX	4.968**	.166	.683*	-.905*	.6747	**
UT	3.936	.060	.800	-.539	.4610	
VT	7.164**	-.139	.692**	-.596*	.5244	*
VA	4.481***	.186	.567*	-.703**	.5848	*
WA	5.870**	.116	.762**	-.540	.5886	*
WV	1.671	.523*	.794	-.453	.8202	***
WI	2.818*	.397	.527	-.505	.6582	**
WY	2.157	.148	.739	-.531	.6144	**

APPENDIX D. SAMPLE SELECTION TERMS FOR JOINTLY DETERMINED LABOR SUPPLY EQUATIONS

Sample selection terms are derived as the conditional means of the disturbance terms of behavioral equations. In the classical OLS model, the expected values of the error terms are equal to zero. But when sample selection is involved, the expected value of the errors are no longer equal to zero. To correct for this bias, sample selection terms are constructed.

In the problem of jointly determined labor supply, sample selection arises from having a sample consisting of only people who work. Furthermore, we hypothesize that the work decision of an individual is jointly determined with the work decision of his/her spouse.

To solve for the sample selection terms in the labor supply equations, we begin by specifying a labor demand equation for each sex:

$$w_1^d = X_1\beta_1 + V_1 \qquad \text{if} \qquad w_1^d > w_1^r$$
$$w_2^d = X_2\beta_2 + V_2 \qquad \text{if} \qquad w_2^d > w_2^r \ .$$

The labor supply equations are specified as:

$$\begin{aligned} H_1 &= w_1a_{111} + w_2a_{121} + z\alpha_{11} + \mu_{11} && \text{if both work,} \\ &= w_1a_{112} \qquad\qquad + z\alpha_{12} + \mu_{12} && \text{if person 1 works,} \\ \text{or} &= 0 && \text{if neither works.} \end{aligned}$$

Likewise,

$$\begin{aligned} H_2 &= w_1a_{211} + w_2a_{221} + z\alpha_{21} + \mu_{21} && \text{if both work,} \\ &= w_2a_{222} \qquad\qquad + z\alpha_{22} + \mu_{22} && \text{if person 2 works,} \end{aligned}$$

or = 0 if neither works.

The reservation wages are found by setting hours worked = 0 and solving for the wage.

$$H_1 = 0 = w_1 a_{11} + w_2 a_{12} + z\alpha_1 + \mu_1,$$

so $$w_1^r = -\frac{1}{a_{11}}(w_2 a_{12} + z\alpha_1 + \mu_1)$$

and

$$H_2 = 0 = w_1 a_{21} + w_2 a_{22} + z\alpha_2 + \mu_2,$$

so $$w_2^r = -\frac{1}{a_{22}}(w_1 a_{21} + z\alpha_2 + \mu_2) .$$

Note that the last subscript is dropped for simplification. Then, substitute in the labor demand equations to get

$$w_1^r = -\frac{1}{a_{11}}(X_2\beta_2 a_{12} + V_2 a_{12} + z\alpha_1 + \mu_1)$$

and

$$w_2^r = -\frac{1}{a_{22}}(X_1\beta_1 a_{21} + V_1 a_{21} + z\alpha_2 + \mu_2) .$$

Notice that the values of the error terms are conditional on which situation is being considered; that is, on the labor force participation status of each spouse.

The following example will illustrate how to find the conditional values of the error terms. Consider the situation where both spouses work. The probability that they both work is:

$$P_{12} = P_r(w_1^d > w_1^r,\ w_2^d > w_2^r)$$

$$= \Pr(X_1\beta_1 + V_1 > -\frac{1}{a_{11}}(X_2\beta_2 a_{12} + a_{12}V_2 + z\alpha_1 + \mu_1))'$$
$$X_2\beta_2 + V_2 > -\frac{1}{a_{22}}(X_1\beta_1 a_{21} + V_1 a_{21} + z\alpha_2 + \mu_2)'$$
$$= \Pr(V_1 + \frac{a_{12}}{a_{11}}V_2 + \frac{1}{a_{11}}\mu_1 > -X_1\beta_1 - \frac{a_{12}}{a_{11}}X_2\beta_2 - \frac{\alpha_{11}}{a_{11}}z,$$
$$V_2 + \frac{a_{21}}{a_{22}}V_1 + \frac{1}{a_{22}}\mu_2 > -X_2\beta_2 - \frac{a_{21}}{a_{22}}X_1\beta_1 - \frac{\alpha_2}{a_{22}}z)' .$$

For simplification, define

$$\xi_1 = V_1 + \frac{a_{12}}{a_{11}}V_2 + \frac{1}{a_{11}}\mu_1 ,$$

$$\xi_2 = V_2 + \frac{a_{21}}{a_{22}}V_1 + \frac{1}{a_{22}}\mu_2 ,$$

$$\Omega_1 = -X_1\beta_1 - \frac{a_{12}}{a_{11}}X_2\beta_2 - \frac{\alpha_1}{a_{11}}z ,$$

and

$$\Omega_2 = -X_2\beta_2 - \frac{a_{21}}{a_{22}}X_1\beta_1 - \frac{\alpha_2}{a_{22}}z .$$

The probability $P_{12} = \Pr(\xi_1 > \Omega_1, \xi_2 > \Omega_2)$. The following notation is needed before proceeding:

f = density function of bivariate $N(0, \sigma_1^2, \sigma_2^2, \sigma_{12})$

$P = \int_{-\alpha_1 X_t}^{\infty}\int_{-\alpha_2 X_t}^{\infty} f(V_1, V_2)dV_1dV_2$ (where $y_{1t}^* = \alpha_1 X_t + V_{1t}$ and $y_{2t}^* = \alpha_2 X_t + V_{2t}$)

f_1 = density of $N(0,\sigma_1^2)$ evaluated at $\alpha_1 X_t$

f_2 = density of $N(0,\sigma_2^2)$ evaluated at $\alpha_2 X_t$

F_1 = distribution function of $N(0, \sigma_1^{*2})$ (where $\sigma_1^{*2} = \sigma_1^2 - \sigma_{12}^2/\sigma_2^2$ and $\sigma_2^{*2} = \sigma_2^2 - \sigma_{12}^2/\sigma_1^2$)

and

F_2 = distribution of $N(0, \sigma_2^{*2})$ (Amemiya, 1974).

Amemiya then shows that when the dependent variable is truncated normal,

$$PE(V_1) = \sigma_1^2 f_1 F_2 + \sigma_{12} f_2 F_1$$

and

$$PE(V_2) = \sigma_2^2 f_2 F_1 + \sigma_{12} f_1 F_2$$

because in the general case,

$$E(V_i) = \frac{1}{P} \sum_{q=1}^{n} \sigma_{iq} f_1(a_1) F_{(q)}$$

where: f_q = marginal density of the qth variable,

$F_{(q)}$ = distribution of the remaining variables, and

$P = \int_{a_1}^{\infty} \cdots \int_{a_n}^{\infty} f(\lambda)\, d\lambda$ and $\frac{1}{P} f(V_1, \ldots, V_n)$
= density of V.

Furthermore, E is the expectations operator and $\rho_{ij.K}$ is the coefficient from the regression $V_j = \rho_{ij.2}\xi_1 + \rho_{2j.1}\xi_2 + \psi_j$ (Johnson and Kotz, 1972, p. 86).

Given the definitions listed above, the expected value of the disturbance term for individual 1 given that both work is:

$$E(\mu_{11}) = E(\mu_1 | \xi_1 > \Omega_2) = E[\rho_{\xi_1\mu_1.\xi_2}\xi_1 + \rho_{\xi_2\mu_1.\xi_1}\xi_2 |$$

$$\xi_1 > \Omega_1, \xi_2 > \Omega_2) = \rho_{\xi_1\mu_1.\xi_2} E[\xi_1 | \xi_1 > \Omega_1, \xi_2 > \Omega_2]$$

$$+ \rho_{\xi_2\mu_1.\xi_1} E[\xi_2 | \xi_1 > \Omega_1, \xi_2 > \Omega_2] ;$$

then recall that $PE(V_i) = \sum_{q=1}^{n} \sigma_{iq} f_q(a_q) F_{(q)}$ (Amemiya, 1974). Therefore,

$$\rho_{\xi_1\mu_1\cdot\xi_2}\, E[\xi_1|\xi_1 > \Omega_1,\ \xi_2 > \xi_2] + \rho_{\xi_2\mu_1\cdot\xi_1}\, E[\xi_2|\xi_1 > \Omega_1,\ \xi_2 > \Omega_2]$$

$$= \rho_{\xi_1\mu_1\cdot\xi_2} \frac{1}{P_{12}} [\sigma^2_{\xi_1} f_{\xi_1} F_{\xi_2} + \sigma_{\xi_1\xi_2} f_{\xi_2} F_{\xi_1}]$$

$$+ \rho_{\xi_2\mu_1\cdot\xi_1} \frac{1}{P_{12}} [\sigma^2_{\xi_2} f_{\xi_2} F_{\xi_1} + \sigma_{\xi_1\xi_2} f_{\xi_1} F_{\xi_2}]$$

Expected values for the remaining error terms are derived in a similar fashion.

APPENDIX E. WAGE LABOR SUPPLY EQUATIONS OF TWO WAGE-EARNER HOUSEHOLDS FITTED BY TWO STAGE LEAST SQUARES

Table E.1. Econometric estimates of wage labor supply functions for married couples in U.S. nonmetropolitan-nonfarm households with two wage earners, 1978-82[a,b]

	Husband	Wife
Intercept	5.33 (10.79)	10.91 (11.06)
Ln V	.158 (2.93)	-.558 (5.20)
$\widehat{\text{Ln } w_1}$	.005 (0.06)	-1.420 (8.69)
$\widehat{\text{Ln } w_2}$	.212 (2.20)	1.461 (7.60)
Educh	.014 (3.06)	.058 (6.42)
Educw	-.018 (2.44)	-.053 (3.68)
Ageh	.045 (6.34)	.111 (7.91)
$\text{Ageh}^2/100$	-.055 (6.65)	-.128 (7.81)
Race	-.115 (4.99)	-.083 (1.80)
Kid6	.023 (2.72)	-.345 (20.13)
Kid618	-.011 (2.52)	-.124 (14.56)

[a]Asymptotic t-ratios in parentheses.
[b]Dependent variable: ln (annual hours worked).

Table E.1. (continued)

	Husband	Wife
$\hat{\lambda}_1$	-.167 (5.41)	.174 (2.82)
$\hat{\lambda}_2$	1.144 (3.63)	-1.454 (2.31)
DNC	-.008 (0.57)	.057 (2.08)
DS	.046 (2.48)	.086 (2.32)
DW	-.520 (2.55)	.021 (0.52)
Time	-.023 (6.19)	-.016 (2.16)
R^2	.0926	.0700
N	15,502	15,502

APPENDIX F. PROBIT ESTIMATES OF WAGE LABOR PARTICIPATION EQUATION TABLES

Table F.1. Probit estimates of wage labor participation equation for U.S. farm household members other than the husband or wife, 1978-82[a]

Intercept	-5.31 (3.66)	Jobgr	.081 (.18)
Age	.287 (19.31)	Service	.015 (.92)
$Age^2/100$	-.278 (19.73)	Ln pc	.274 (1.17)
$Educ_h$	.007 (.78)	Ln pl	.013 (.05)
$Educ_w$	-.024 (2.42)	Ln pi	-.400 (.69)
Race	.351 (3.14)	Ln pw	.486 (1.55)
Kid6	-.406 (7.26)	Rain	-.003 (.28)
Kid618	.259 (12.91)	Gdd/1000	-.141 (2.20)
Ln V	-.235 (1.63)	Gddrain/10,000	.018 (.68)
Stw	.483 (2.15)	DNC	-.074 (.68)
Urate	-.001 (.04)	DS	-.023 (.19)
Dif3	.025 (.82)	DW	-.057 (.40)
		Time	.009 (.31)

[a]Asymptotic t-ratios in parentheses.

Table F.2. Probit estimates of wage labor participation equation for U.S. nonmetropolitan nonfarm household members other than the husband or wife, 1978-82[a]

Intercept	-5.34 (3.39)	Stw	.046 (.16)
Age	.231 (13.79)	Urate	-.008 (.40)
$Age^2/100$	-.221 (13.70)	Dif3	.053 (1.35)
$Educ_h$	-.014 (1.22)	Jobgr	.759 (.85)
$Educ_w$	-.021 (1.49)	Service	-.002 (.11)
Race	.138 (1.26)	DNC	-.002 (1.55)
Kid6	-.374 (4.44)	DS	-.269 (2.77)
Kid618	.208 (7.30)	DW	-.089 (.59)
Ln V	-.087 (.50)	Time	.004 (.13)

[a]Asymptotic t-ratios in parentheses.

APPENDIX G. EXOGENEITY TEST

Some economists hypothesize that children and nonwage income are choice variables and should not be treated as exogenous in labor supply equations. A recent study by Mroz (1987), however, found children and nonwife income to be exogenous in married women's labor supply functions.

The exogeneity of nonwage income and children in the household income equations were tested using a 10% subsample of the rural nonfarm group. An exogeneity test described by Luger and Stahl (1986) was used. They considered a model,

$$Y = X_1\beta + v\alpha + \varepsilon, \quad \varepsilon \sim N(0, \sigma^2),$$

where Y is a Txl vector, X_1 is a TxK matrix of explanatory variables, ε is an error term, and v is a Txl variable hypothesized to be uncorrelated with ε. The joint hypothesis that $Y=X_1\beta + \varepsilon$ is the true model and $E(v'\varepsilon)=0$ is tested by running a regression and testing $\alpha=0$. If $\alpha\neq0$, then either the true model includes v, or v is correlated with ε. Then, let Z_1 be a TxL matrix (L>K) which includes X_1 plus the variables to be tested for exogeneity. The variables to be tested (KID6, KID618, and ln V) were replaced by instrumental variables and the household income equation was estimated by two stage least squares. The instrumental variables were tested one at a time. If the hypothesis $\alpha=0$ is rejected for any one of the set of instrumental variables, they cannot be used as a set and a different combination of variables must be used.

The hypothesis $\alpha=0$ could not be rejected for the three variables

hypothesized to be exogenous (t-ratios were -1.26 for KID6, 1.30 for KID618, and -1.31 for ln V). We may assume that nonwage income and children can be used as exogenous variables in the household income equations.

Econometric estimates of the reduced-form farm and nonfarm household cash income equations that exclude KID6, KID618, and V as regressors are reported in Tables G.1 and G.2, respectively.

Table G.1. Econometric estimates of reduced-form equations for U.S. farm household cash income, 1978-82, with nonwage income, the number of children aged 6 or less, and the number of children aged 6 to 18 excluded[a]

	Regression[b]			
	1	2	3	4
Intercept	9.43 (127.69)	9.47 (154.52)	9.42 (127.73)	9.47 (154.56)
Age	.017 (15.27)	.017 (15.27)	.017 (15.25)	.017 (15.25)
$Age^2/100$	-.016 (14.66)	-.016 (14.66)	-.016 (14.66)	-.016 (14.66)
$Educ_h$	.013 (11.38)	.013 (11.41)	.012 (20.39)	.012 (20.44)
$Educ_w$	.010 (7.52)	.010 (7.52)	.012 (20.39)	.012 (20.44)
Race	-.030 (1.81)	-.030 (1.73)	-.030 (1.84)	-.028 (1.76)
Ln stw	.091 (3.00)	.095 (3.28)	.091 (3.00)	.095 (3.27)
Urate	.003 (1.37)	.005 (2.53)	.003 (1.42)	.005 (2.60)
Dif3	-.002 (0.50)	-.003 (0.79)	-.002 (0.49)	-.003 (0.80)
Jobgr	-.046 (0.77)		-.045 (0.75)	
Service	-.001 (0.21)		-.001 (0.23)	

[a]Asymptotic t-ratios in parentheses.
[b]Regression (1) all variables included; (2) best fit; (3) all variables, $\beta_{educh} = \beta_{educw}$; (4) best fit, $\beta_{educh} = \beta_{educw}$.

Table G.1. (continued)

	Regression[b]			
	1	2	3	4
Rain	.004 (3.45)	.003 (3.46)	.004 (3.51)	.003 (3.53)
Gdd/1000	.044 (5.25)	.041 (5.28)	.045 (5.35)	.041 (5.40)
Gddrain/10000	-.009 (3.35)	-.0075 (3.19)	-.0092 (3.45)	-.0077 (3.27)
Ln pc	.056 (1.77)	.050 (1.64)	.055 (1.74)	.049 (1.62)
Ln pl	-.113 (3.45)	-0.109 (3.43)	-.113 (3.46)	-.108 (3.42)
Ln pi	.104 (1.33)	.157 (2.47)	.105 (1.34)	.159 (2.50)
Ln pw	.159 (3.78)	.151 (3.91)	.161 (3.83)	.154 (3.97)
Time	-.013 (3.35)	-.015 (3.92)	-.013 (3.51)	-.015 (3.92)
DNC	.001 (0.05)		.0004 (0.03)	
DS	.006 (0.33)		.005 (0.32)	
DW	.020 (1.01)		.020 (1.01)	
R^2	.1439	.1437	.1436	.1434

Table G.2. Econometric estimates of reduced-form equations for household cash income of U.S. nonmetropolitan nonfarm households, 1978-82, with nonwage income, the number of children aged 6 or less, and the number of children aged 6 to 18 excluded[a]

	Regression[b]			
	1	2	3	4
Intercept	6.38 (126.92)	6.39 (130.23)	6.38 (127.22)	6.39 (130.52)
Age	.069 (58.04)	.069 (58.03)	.069 (58.04)	.069 (58.03)
$Age^2/100$	-.071 (60.47)	-.071 (60.46)	-.071 (60.47)	-.071 (60.46)
$Educ_h$	.048 (36.04)	.048 (36.16)	.048 (75.36)	.048 (75.65)
$Educ_w$	.048 (29.92)	.048 (29.93)	.048 (75.36)	.048 (75.65)
Race	-.163 (12.55)	-.163 (12.56)	-.163 (12.57)	-.164 (12.58)
Ln stw	.174 (5.39)	.174 (59.43)	.174 (5.39)	.174 (5.44)
Urate	-.012 (5.33)	-.013 (6.27)	-.012 (5.33)	-.013 (6.27)
Dif3	.0003 (0.08)		.0003 (0.08)	
Jobgr	.114 (1.15)		.114 (1.15)	
Service	.010 (4.33)	.009 (4.22)	.010 (4.33)	.009 (4.22)

[a]Asymptotic t-ratios in parentheses.
[b]Regression (1) all variables included; (2) best fit; (3) all variables, $\beta_{educh} = \beta_{educw}$; (4) best fit, $\beta_{educh} = \beta_{educw}$.

Table G.2. (continued)

	Regression[b]			
	1	2	3	4
DNC	-.094 (8.22)	-.096 (8.38)	-.094 (8.22)	-.096 (8.39)
DS	-.100 (8.74)	-.096 (8.82)	-.100 (8.74)	-.096 (8.82)
DW	-.064 (3.74)	-.058 (3.57)	-.064 (3.74)	-.058 (3.57)
Time	-.025 (7.97)	-.025 (8.71)	-.025 (7.97)	-.025 (8.71)
R^2	.2976	.2969	.2976	.2969

APPENDIX H. ECONOMETRIC ESTIMATES OF REDUCED FORM EQUATIONS TABLE

Table H.1. Econometric estimates of reduced form equations for farm household cash income of U.S. farm households, 1978-82, by population density of state of residence[a]

	Sparsely populated states	Densely populated states		Sparsely populated states	Densely populated states
Intercept	5.50	5.29	Service	-.002	.0005
	(20.67)	(20.08)		(.45)	(.17)
Age	.009	.016	Ln pc	.102	.073
	(5.55)	(10.05)		(2.26)	(1.67)
$Age^2/100$	-.010	-.018	Ln pl	-.109	-.185
	(6.24)	(11.59)		(1.94)	(3.71)
$Educ_h$	.008	.010	Ln pi	.335	.206
	(1.46)	(6.31)		(1.83)	(1.90)
$Educ_w$	.003	.010	Ln pw	.035	.105
	(1.46)	(6.31)		(.48)	(1.61)
Race	.048	-.010	Rain	.0006	.001
	(1.22)	(.55)		(.10)	(.61)
Kid6	-.035	-.034	Gdd	-.020	.024
	(5.29)	(4.70)		(.77)	(1.48)
Kid618	.010	-.003	Gddrain	.115	-.003
	(2.90)	(.77)		(.73)	(.70)
Ln V	.503	.477	DNC	-.058	.015
	(19.09)	(19.30)		(1.62)	(.79)
Ln stw	.048	.069	DS		-.017
	(.70)	(1.45)			(.90)
Urate	-.021	-.003	DW		-.005
	(3.24)	(.77)			(.22)
Dif3	-.009	-.003	Time	.009	-.009
	(1.00)	(.50)		(.97)	(1.63)
Jobgr	.232	-.047	R^2	.19	.28
	(1.25)	(.65)		(.97)	(1.63)

[a]Asymptotic t-ratios in parentheses.