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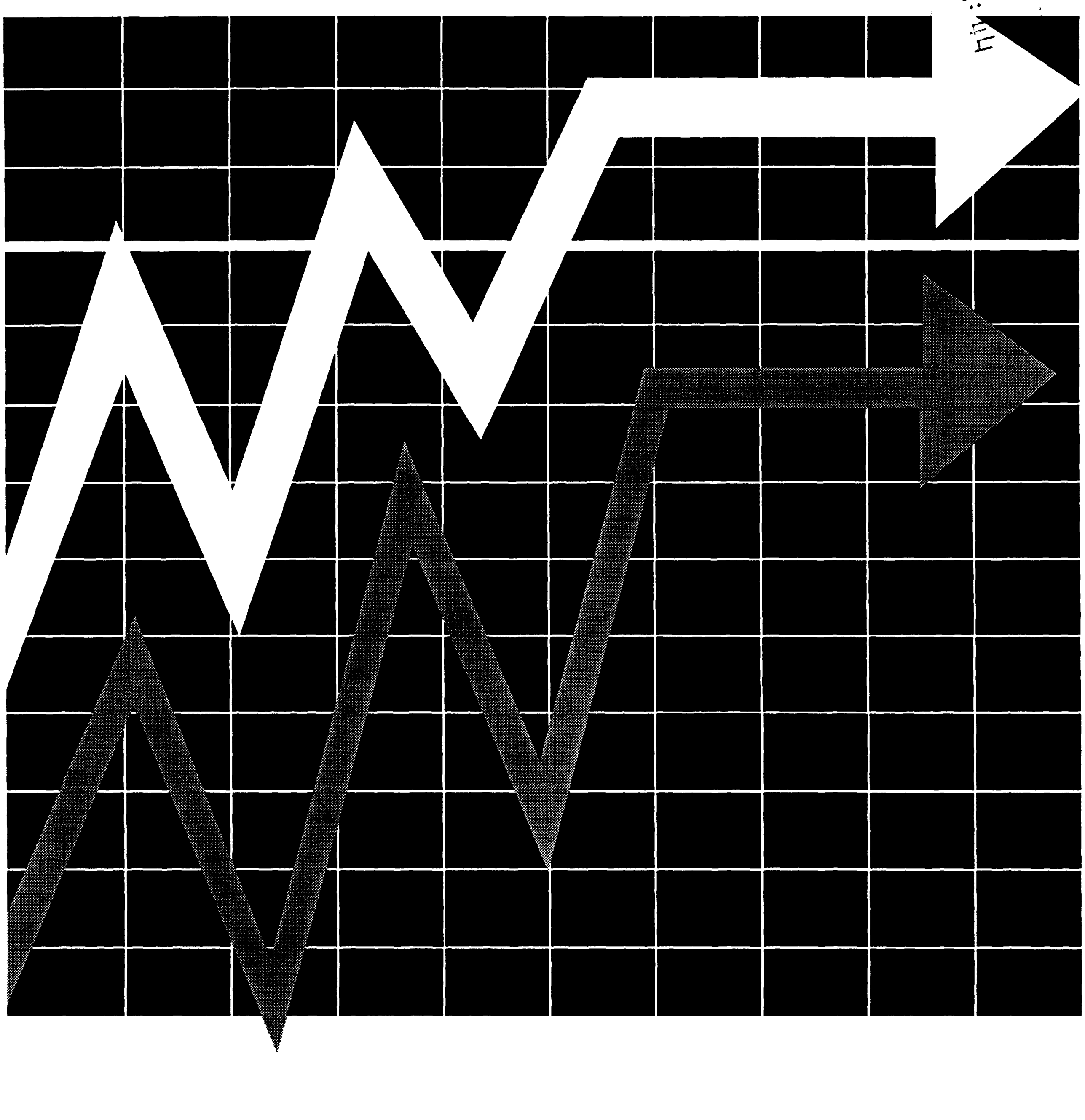
Estimating the Opportunity Cost of Unpaid Farm Labor for U.S. Farm Operators

Hisham S. El-Osta and Mary C. Ahearn



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Abstract

A farm operator household time-allocation model and data from the Farm Costs and Returns Survey and from other sources were used to derive off-farm wage equations for U.S. farm operators based on production region, size of farm, and farm type. These equations, in turn, were used to impute opportunity costs for farm operators' unpaid labor. Using this new method, instead of the ad hoc method of valuing labor at hired labor wage rates, should improve the structure of commodity cost and return statements, including those developed by the U.S. Department of Agriculture.

Keywords: Farm operator household, unpaid farm labor, off-farm labor participation, human capital, labor market area, off-farm wage rates.

Acknowledgments

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Contents

Summary

The U.S. Department of Agriculture's current economic cost statements for agricultural commodities account for operators' unpaid labor hours in an ad hoc fashion, namely, by multiplying hours of work reported in the Farm Costs and Returns Survey (FCRS) devoted to the enterprise by the State wage rate for agricultural workers. The AAEA (American Agricultural Economics Association) Taskforce on the Standardization of Commodity Costs and Returns Methods has reviewed alternative approaches on how to value unpaid labor more accurately. This report explores one of these alternatives by valuing unpaid labor using an opportunity cost approach based on off-farm labor markets.

To achieve this, a farm operator household time-allocation model is utilized in conjunction with data from the 1988 FCRS and from other sources. The analysis is carried out by imputing predicted off-farm wage rates to serve as proxies for operators' opportunity cost of unpaid labor for the entire United States, by region, by size of farm, and by farm type, with possible selection bias being accounted for using a two-stage correction procedure. In the first stage, logistic regression is used to estimate a sample selectivity variable. This variable, in turn, is employed in the second stage to estimate, by means of weighted least squares, unbiased and consistent estimates of off-farm wage rates, or alternatively, of operators' cost of unpaid labor. Because of the complex sample design of the underlying data, a specialized regression algorithm is utilized for all of the empirical estimation.

The results show that the predicted opportunity cost of farm operators' unpaid labor in 1988 averaged $9.68 per hour. Differentials in the opportunity costs based on production region, on size of farm, and on type of farm specialization are shown to exist.

The method outlined in this report regarding how to value farm operators' unpaid labor will be incorporated by ERS in its future estimation of commodity costs and returns. The results arrived at here can be used as a benchmark against which comparison of future costing of unpaid labor can be made.

Estimating the Opportunity Cost of Unpaid Farm Labor for U.S. Farm Operators

Hisham S. El-Osta and Mary C. Ahearn

Introduction

Farm labor is a critical input in agricultural production. Farm operators provide just about half of all labor hours in agriculture, most of which are unpaid. An additional 18 percent of farm labor hours are unpaid hours provided by farm household members, partners, and others. Consequently, two-thirds of the labor hours worked on U.S. farms are not valued directly in the marketplace. Since labor is generally considered to be a variable cost, which indirect method is used to value the unpaid labor in agriculture has important effects on the marginal cost of production. The indirect methods used in research and in the construction of statistics vary significantly.

The AAEA Taskforce on the Standardization of Commodity Costs and Returns Methods has reviewed alternative approaches to the issue of valuing unpaid labor. The intent of this paper is to explore one of the alternative methods to valuing unpaid labor in agriculture, that of an opportunity cost (i.e., the net value of time spent in the next best activity) in off-farm labor markets.

The theoretical foundation of this method of valuing human time has been established and demonstrated by Gronau, Heckman (1974), and Kniesner in U.S. labor supply studies, and has been applied by Rosenzweig in a study of rural labor supply in India. Kislev and Peterson (1982) and Huffman (1992) considered the wage rate for off-farm work as the opportunity cost of farmers' own labor. Huffman (1992) used an agricultural household production model to derive predicted wage rates that can serve as the cost of operators' farm labor.

This report derives predicted opportunity costs of labor for the entire United States using an approach closely related to Huffman's. The report also derives predicted opportunity costs of labor by region, by size of farm, and by farm type. This level of disaggregation is relevant since onfarm and off-farm labor allocation decisions of farm operators, along with the structure of their off-farm wages (i.e., the relative returns to different levels of skills and/or educational attainment), tend to differ based on the farm's regional location, its economic size, and its type of specialization.

Current Methods for Valuing Labor

In the U.S. Department of Agriculture's (USDA) methods for estimating costs and returns for individual agricultural commodities, all physical labor and management are not accounted for as a single labor input. Currently, USDA presents two different types of cost and return statements for each agricultural commodity. How labor is accounted for is, in part, a function of the purposes of these statements.

The first type of statement is strictly a cash statement. Therefore, only paid labor is included as a variable cash expense. If a household member or hired manager were paid a wage, those expenses would be included in the cash labor expenses of the cash statement. This implicitly assumes that all paid labor (both physical and management components) is a variable cost.

The second type of USDA statement represents almost all costs of production, cash and noncash, including opportunity costs for owned inputs. These are termed "economic costs." Only costs to risk and unpaid management are not accounted for in this alternative statement, except as the residual income.

Therefore, in USDA's current economic cost statement, labor is accounted for in three places in a piecemeal fashion: (1) paid expenses for hired labor and management are included in variable cash expenses, (2) unpaid physical work hours are imputed an opportunity cost equivalent to hired labor wage rates, exclusive of management, and (3) the value of unpaid management is not directly costed, so that it is included in the net returns (gross value of production less economic costs) as a residual along with the returns to risk.[1] This residual is commodity price (market or ex-post) determined and may be mainly the effects of random events, such as weather. The current structure of the USDA costs and returns statements does not provide an estimate of the total, cash and noncash, variable labor costs, and in effect it has differing treatment of management services depending on whether they are hired or not.

The practice of not treating operator labor and management as a single input appears to be prevalent standard among those who estimate commodity costs and returns in the United States. Klonsky reports that estimated labor hours for crops typically are based on a measure of machine time, which may or may not be adjusted for the time to move the equipment in or out of the field or for set-up time. These hours are valued at a hired wage rate (Klonsky, p. 155).

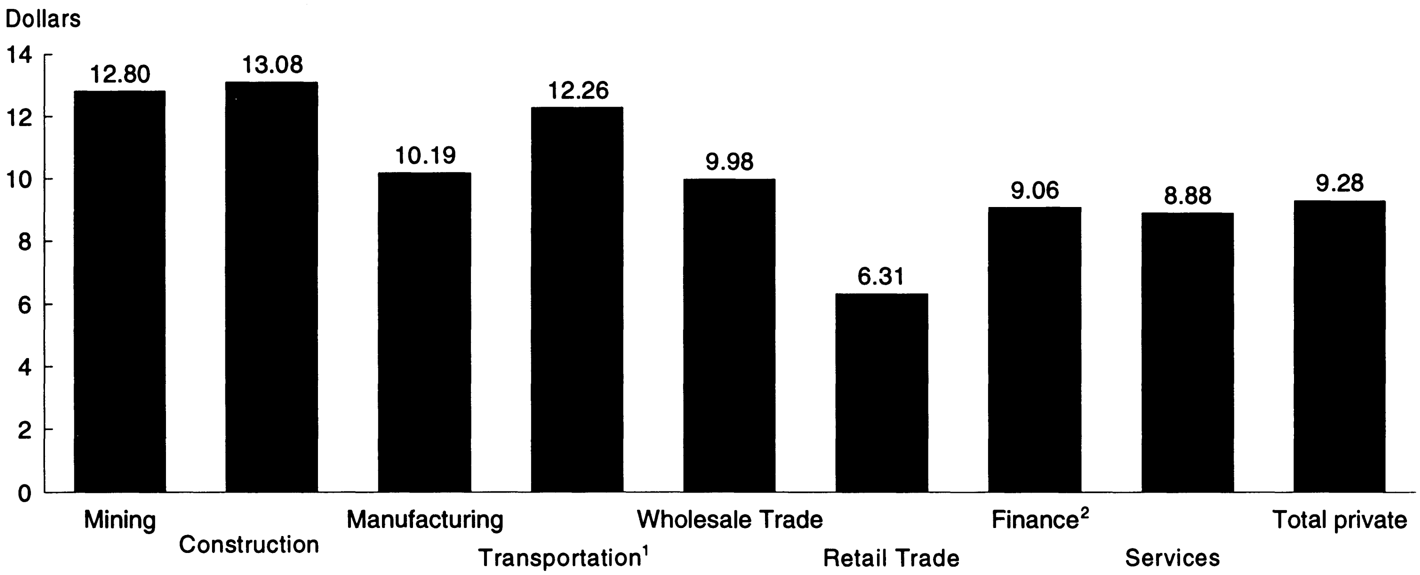
Sixteen of the 41 States which provided Klonsky with information on their measurement practices reported that they do not include any cost for management. When management costs are estimated, they are done so as a separate entry in a costs and returns statement. A variety of approaches are employed to measure the cost of management, including the methods used by the USDA in its economic costs and returns statement. Klonsky reported that 10 of the 41 States that provided information on their accounting systems charged for management as part of the residual return above total costs. The 15 States that did account for management explicitly used three basic approaches: (1) based on a percentage of gross receipts, (2) based on a percentage of costs, and (3) by multiplying hours by a specified wage rate, for example, by the wage rate for hired farm managers.

The USDA also produces estimates of the productivity of U.S. agriculture that require the valuation of labor in agriculture. Currently, the USDA approach is to value unpaid labor at the State average wage rate for hired workers (Hauver). In this series, labor is meant to include management, as well. Therefore, the use of a hired wage rate for farmworkers is especially troublesome in this series.

In the returns to equity and the returns to asset series produced by USDA, another approach is used to value unpaid labor and management in agriculture. To calculate the returns to assets and equity, part of the residual net income must be allocated to labor and management. The process for valuing labor and management is to value the unpaid hours worked at the State average wage

[1]While the USDA, in its economic costs and returns statement, values unpaid labor at the agricultural wage rate (USDA, 1990, p. 10), this valuation method is not unique to the United States. For example, although neither Canada nor EEC member countries have official programs to estimate annual costs of production, such estimation is usually performed in Canada by universities or by provincial governments for commodities of importance in their areas, and is undertaken in France by private associations or by banks owned and financed by growers. Studies on the cost of producing wheat in these countries have shown that unpaid labor is valued at hired labor rates (Ahearn, Culver, and Schoney; Le Stum and Camaret).

Figure 1
Average hourly earnings for production or nonsupervisory workers on nonfarm payrolls by major industry, 1988



[1]Includes also public utilities.
[2]Includes also insurance and real estate.
Source: Employment and Earnings, U.S. Dept. of Labor, Bureau of Labor Statistics, Dec. 1993, pp. 97-99.

rate for hired farmworkers plus to include a management charge of 5 percent of the gross income.

In whole-farm accounting, one common approach is to use the farm family living withdrawals as an estimate of management costs. Such an approach was included in the recent recommendations for whole-farm financial reporting, Financial Guidelines for Agricultural Producers (Farm Financial Standards Taskforce). The criticism of this approach is that family living withdrawals are affected by many factors other than the value of the unpaid labor in the farming operation, for example, the need for strictly household purposes.

Where unpaid farm labor hours are valued at hired labor wage rates, a serious consequence is the underestimation of the cost of human time spent on the farm. Schultz (1972) and Huffman (1992) note that farm operators and members of their households have on average a much larger stock of human capital (education, experience) than their hired farmworkers do. Pearson asserts that hired farmworkers have lower educational levels than the total population, and that, in 1979, 44 percent of hired workers had no schooling beyond the eighth grade.

Wolfson points out that education, for example, has effects on the alternative opportunities which a laborer may be aware of, or may consider. That is, the higher the level of schooling reached by the person, the broader the social and cultural vistas that are opened to the person. In 1988, the average U.S. hourly earnings for off-farm work was $9.28, compared with an average wage rate of $5.02 for hired workers, a difference of 84.9 percent (see figs. 1 and 2).[2]

[2]While it is true that hourly wage rates for hired farmworkers tend to be less than those of nonfarm workers, it should be noted that averages of hourly earnings and wage rates are not strictly comparable. That is because earnings are the actual return to the worker for a stated period of time and rates are the amount stipulated for a given unit of work or time (U.S. Dept. Labor, p. 166).

Earlier data (1986) for southwestern Wisconsin show similar patterns as the average hourly wage rate for off-farm work for male farm operators was \$10.09, versus \$4.12 for hired farm labor in the Lake States (Wisconsin, Minnesota, and Michigan), a difference of 145 percent (Huffman, 1992).

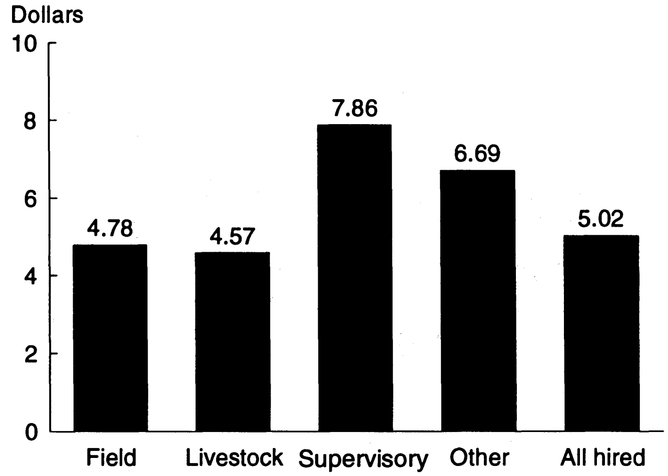
The Wisconsin data also show that the opportunity cost of farm work for farm operators and their spouses increases as they accumulate more years of formal education. The importance of human capital characteristics in explaining the off-farm wage rates of farm operators has been documented, among others, by Scott, Smith, and Rungeling, by Sumner (1981 and 1982), and by Huffman and Lange.

The remainder of this paper is organized as follows. The next three sections describe the method used to estimate the opportunity cost of unpaid labor and the underlying conceptual relationships, empirical specifications, and data sources. Next, the empirical results are presented and discussed. The final section offers some summary comments on the major results, including a discussion on the limitation of the proposed alternative to valuing human time in agriculture.

Conceptual Relationships and Methodology

The literature on off-farm labor supply suggests that the optimal allocation of time and wage rate determination of household members is reached by solving the equivalent of the following optimization problem (Lee, 1965; Sumner, 1982; Huffman and Lang; Tokle and Huffman; Huffman, 1992; Furtan, Van Kooten, and Thompson; Jensen and Salant; among others):[3]

Figure 2
Average wage rates for hired workers, by type of work, 1988[1]



[1]Excludes agricultural service workers.
Source: Farm Labor, U.S. Dept. of Agr., National Agr. Stat. Service, Nov. 1988, p. 14.

$$Maximize\ U = (Y, L^i), \quad (i = 0, S) \quad (1)$$

subject to the constraints:

$$T^i = L^i + E^i + F^i, \text{ and} \quad (2)$$

$$Y = \Sigma_i\, \theta^i\, (E^i, H^i, M^i) + p_f\, \varphi\, (F^i, X) - p_x X, \quad (3)$$

where U in the objective function is utility, Y is total income of farm operator household and L^i are times allocated by the operator, O, and the spouse, S, to leisure. In equation 2, T is the total time endowment, L is the time allocated to leisure, E is time allocated to off-farm work, and F is time allocated to farmwork for both the operator and spouse. Total household income Y in equation 3 is a budget constraint where $\theta(.)$ is the net off-farm earnings function, H is human capital, M are labor market conditions, p_f is the price of farm output, p_x is a vector of farm input

[3]Because of unavailability of data, factors such as household production, commuting costs to off-farm jobs, and possible nonpecuniary income from farming are ignored. For a description of how commuting costs and nonpecuniary income can be augmented in models of labor force participation in general, see studies by Solberg and Wong, Axe and Golden, and Warner and Goldberg.

[4]When only human characteristics are used, the earnings function in equation 3 is sometimes interpreted as a hedonic price function. The theory, which was originated by Rosen (1974), reflects the equilibrium of the supply and demand for working at each level of schooling and experience (Willis).

prices, X is a vector of nonlabor farm inputs, and φ (.) is the farm production function.[4]

Under the assumption that the utility function and the nonstochastic farm production function of the household are concave, continuous, and twice differentiable, the optimal allocation of time by farm operators and their spouses (neutrality of preferences is maintained here) between leisure, onfarm activities, and off-farm work is obtained by solving the set of first-order conditions of the above model. Such optimal allocation is reached when the marginal values of time devoted to leisure, farmwork, and off-farm work are set equal to each other. Further, the marginal wage rates, which are considered as proxies for the opportunity cost of their unpaid farm labor, are determined as:

$$w^i = \frac{\partial \theta^i}{\partial E^i} = w^i(H^i, M^i),\ i = (0, S) \qquad (4)$$

This relationship assumes that the stock of human capital and labor market conditions affect the marginal wage rates of the farm operators and of their spouses. Equation 4 further assumes that the marginal wage rates do not depend on hours of off-farm work or any associated fixed costs (Sumner, 1982).[5]

Under the assumption that the opportunity cost of a farm operator's unpaid farm labor is equal to his/her marginal product function of off-farm work, equation 4 can be generalized for all farm operators who participated in off-farm employment. Using a variant of Mincer's human capital earnings function, equation 4 can be rewritten as:[6]

$$ln\ w_j = Z_j\beta + \varepsilon_j, \qquad (5)$$

$$(j = 1, \ldots, n) \text{ and } E(\varepsilon_j, \varepsilon_k) = \begin{pmatrix} \sigma^2 \text{ if } j = k \\ 0 \text{ otherwise} \end{pmatrix}$$

where ln is the natural logarithm, w_j is the hourly wage rate of operator j ($j=1,...,n$) computed as annual net earnings from off-farm wages and salaries and/or from off-farm business divided by off-farm work hours, Z_j is the j^{th} operator's vector of explanatory variables, β is the vector of parameters to be estimated including an intercept term, and εj is the random error term.[7]

Since not all farm operators participate in off-farm employment, the dependent variable in equation 5 is observed only in a limited range where the j^{th} operator's wage rate ($w^*{}_j$) is positive. These nonzero values of the dependent variable have the following expectation:

$$E(ln\, w_j | Z_j, w_j^* \succ 0) =$$
$$Z_j\beta + E(\varepsilon_j | w_j^* \succ 0) \qquad (6)$$

for $j=1,...,n$. Equivalently, equation 6 can be written as:

$$E(ln\ w_j \mid Z_j, w_j^* \succ 0) = Z_j\beta + E(\varepsilon_j \mid \varepsilon_j^* \succ -Z_j\beta) =$$
$$Z_j\beta + \sigma\lambda_j \qquad (7)$$

where $\varepsilon_j^* \sim N(0, \sigma^2)$ and where $\sigma\lambda_j$ is the conditional expectation of the truncated variable ε_j as defined in equation 5.[8] The term $\sigma\lambda_j$ is zero if operators with w_j^*0 are a random sample of all operators. In this case, ordinary least squares regression (OLS) on the participating sample will yield unbiased estimates of the parameters (Gunter and McNamara). However, because the decision of the operator to participate in off-farm employment is based on the relative marginal

[5]See Rosen (1976) for a discussion on how wages may depend upon the number of hours worked in models of labor supply. Hausman (1980) and Cogan (1981) demonstrate the effect of fixed consts on women's labor force participation.

[6]Since the subsequent discussion pertains to the farm operator only, subscript i used in equations 1-4 has been dropped, and instead, subscript j has been introduced.

[7]The definition of net earnings used in this report is that of the Bureau of the Census (U.S. Dept. Commerce, p. xvi).

[8]See Kmenta, p. 563. The reciprocal of the quantity λ_j is known in the literature as the "Mill's ratio."

utilities of farmwork, off-farm work, and leisure, and these relative marginal utilities are related to vector Z_j, participating operators are not a random sample of all operators, and $\sigma\lambda_j$ is not zero. As a result, OLS estimation of equation 5 without incorporating $\sigma\lambda_j$ in the regression will yield biased estimates. To test and to correct this problem, which is known in the literature as sample selectivity bias, it is common to use a two-stage estimation technique proposed by Heckman (e.g., Gunter and McNamara; Furtan et al.; Gould and Saupe; Tokle and Huffman; Lass, Findeis, and Hallberg; Findeis; Kiker and Oliveira).

In the context of this study, testing and correcting for the problem of sample selection bias starts by forming the following binary variable that tracks the decision of the farm operator to participate in off-farm employment:

$$I_j = \begin{pmatrix} 1, if\ w_j \succeq W_j, (j = 1,\ldots,n) \\ 0, if\ w_j \prec W_j, (j = n+1,\ldots,N) \end{pmatrix} \qquad (8)$$

where W_j is the j^{th} operator's reservation wage rate (i.e., the lowest wage rate considered acceptable), w_j is the j^{th} operator's observed wage rate as in equation 5, $j = 1$ indicates participation, and $j = 0$ indicates nonparticipation.[9] Using E to denote mathematical expectation, the probability that the farm operator works off the farm is measured as:

$$E(I_j) = P_j(I_j = 1) = P_j(u_j \prec Z_j\beta) =$$
$$F(Z_j\beta) = \int_{-\infty}^{Z_j\beta} f(u)d(u), \qquad (9)$$

where $F(.)$ is the cumulative distribution function and $f(.)$ is the probability density function of the random variable u_j, and where u_j is a composite term that approximates the random aspects of the farm operator's participation decision (see Aldrich and Nelson, pp. 35-36). In equation 9, the probability distribution of P_j can take many forms including

[9]For a discussion on the labor market effects of reservation wages, see Holzer.

normal or logistic distributions.[10] If the distribution is assumed normal, Heckman's two-stage method becomes the valid technique to address the issue of self-selectivity. This technique uses probit regression in its first stage to estimate the parameters ß of equation 9 which are used in the second stage to test and to correct for selection bias. However, if the distribution of P_j is assumed logistic, which is a valid assumption if some of the explanatory variables in equation 9 are dummy variables (Sinden and King) as in this report, testing and correcting for self-selection bias is carried out using Lee's two-stage procedure (Lee, 1979, 1982, and 1983).[11] This technique is similar to Heckman's since it computes consistently the selectivity variable M in a similar fashion, and since it is performed in an omitted variable framework. However, Lee's technique is more general than Heckman's since it allows for P_j to have either a normal or a logistic distribution, which in turn allows for the use of either the probit or the logit regression procedures. The first stage of this procedure requires the estimation of P_j as follows:

[10]In addition to the Normal and the Logistic distributions, Aldrich and Nelson (p. 33) suggest the Angular, Gompertz, Burr, Urban, and the Truncated Linear Probability distributions.

[11]Amemiya (p. 1487) and Kmenta (p. 555) note that the binomial logistic and cumulative normal functions are very close in the midrange, but the logistic function has slightly heavier tails than the cumulative normal. Accordingly, it makes no difference which function is used except in cases where the sample size is large and where the data are heavily concentrated in the tails due to the characteristics of the problem being analyzed. Because the size of the underlying sample is large, as will be evident in the data section, it is argued here that the presence of dummy variables in the model may cause the data to be concentrated at tails, hence warranting the assumption of logistic function. This asumption is convenient here for computational purposes, since, as far as the authors know, there exists no regression package that can accommodate probit analysis where the underlying data are based on surveys with complex sample design as in this report.

$$E(I_j) = P_j(I_j = 1) = \frac{1}{1 + e^{-lnw_j}} = \frac{1}{1 + e^{-Z_j\beta}} = \frac{e^{Z_j\beta}}{1 + e^{Z_j\beta}} \quad (10)$$

The selectivity variable λ is then estimated as:

$$\lambda_j = \frac{\phi(q_j)}{\Phi(q_j)}, \quad (11)$$

where

$$q_j = \Phi^{-1}(P_j). \quad (12)$$

In equation 11, $\phi(.)$ and $\Phi(.)$ are the standard normal probability density function and cumulative standard normal density function, respectively, evaluated at the argument.

The second stage of Lee's method requires the inclusion of λ in the OLS estimation of equation 5. Hence, conditional on I_j=1, the new censored off-farm wage rate equation becomes:

$$ln\,w_j = Z_j\beta + \sigma\lambda_j + \xi_j, \quad (j = 1,\dots,n), \quad (13)$$

where(

$$E(\xi_j | w_j^* \succ 0, Z_j, \lambda_j) = 0. \quad (14)$$

Selection bias is indicated if the coefficient of λ in equation 13 is significant. The inclusion of the selectivity variable in the estimation of farm operators' expected off-farm wage rates gives unbiased and consistent estimates of the model's parameters.[12] This, in turn, allows for a more precise estimation of the opportunity cost of unpaid labor regardless of the operator's off-farm work status. Specifically, the predicted opportunity cost of unpaid farm labor for operators who worked off the farm ($j=1,\dots,n$) and for operators who did not ($j=n+1,\dots,N$) can be computed as:

$$w_j = E(e^{lnw_j}) = e^{Z_j\hat{\beta} + \hat{\sigma}\lambda_j}, \quad (j = 1, \dots, n, n + 1, \dots, N), \quad (15)$$

where e is the natural exponential function and where $\hat{\beta}$ and $\hat{\sigma}$ are the estimated parameters from equation 13.[13]

Empirical Specification

The variables used in modeling off-farm labor participation decisions by farm operators are mostly those found in the literature (table 1). The first group of variables describe the characteristics of the operator and include such variables as the age of the operator, age-squared, and the operator's education.[14] Previous studies (e.g., Sumner, 1982; Gunter and McNamara; Tokle and Huffman) have found that the

[12] A reviewer correctly questions the accuracy in the estimation of the standard errors of the model's parameters under Heckman's (see Greene) and under Lee's method. Under Lee's method, the problem arises due to the restrictive assumption of the normality of the disturbances ε_j^* ($j = 1,\dots,n$) in equation 7. Lee (1982) proposes the construction of parameters' standard errors using a version of White's (1980) heteroskedasticity-consistent covariance matrix estimate as a way of relaxing this restrictive assumption—a proposition that was adopted by McMillen and McDonald in their testing of the presence of selection bias in urban land value functions. However, such a proposition is not practical when the underlying data are based on surveys with complex design, as in this study, because of the added complexity in the estimation of the variance of the model's parameters. A consequence of this is that it may be difficult to ascertain the presence of the selectivity bias in the regression equation, especially when regression is performed on small data sets where the possibility of misspecifying the distribution of ε_j^* as normal increases.

[13] In estimating equations 11, 13, and 15, λ_j is programmed as: $(1 - I_j)\{\phi(q_j) / (1 - \Phi(q_j))\} - I_j\{\phi(q_j) / \Phi(q_j)\}$.

[14] The 1988 Farm Operator Resource (FOR) version of the Farm Costs and Returns Survey (FCRS), one of the underlying data sources in this report, did not collect information regarding the gender of the operator. More recent FCRS data revealed that 94 percent of farm operators are male (Ahearn, Perry, and El-Osta, p. 9).

likelihood of off-farm participation by farm operators peaks with age, but then declines as farm operators grow older, *ceteris paribus*. The effect of educational attainment on the participation decision is ambiguous (Huffman, 1980; Sumner, 1982) as an increase in education can increase the marginal value of time both on and off the farm. Lass and Gempesaw found a positive, but insignificant, correlation between the educational level of the farm operator and the operator's decision to participate in off-farm employment. Miller asserts that education in general provides credentials necessary to enter high-paying occupations in addition to reflecting the acquisition of skills, and as such is likely to increase individuals' full-time participation in labor market activities. While most off-farm jobs positively weigh formal education in the hiring process, the credential aspect of educational attainment alone does not hold value. This is because many farmers who do not have the educational credentials may still possess the needed set of skills to enable them to get hired.

Five household characteristics are hypothesized to affect how farm operators allocate their time to off-farm work. The first of these variables describes the spouse's decision to participate in off-farm employment. The hypothesized relationship of this variable is not clear *a priori*. For example, Findeis and Lass found that when both the operator and the spouse work off the farm, the farm operator will allocate more time to off-farm employment and less time to onfarm work if the farm hires farm labor. However, particularly in urban localities, the opposite occurs when no supplemental labor is hired. Findings by LeClere show, only in nonmetro areas however, that the participation of wives in the off-farm labor market has significantly increased the probability of participation of their husbands. The author suggests, as in Huffman and Lange (1989) and Lass *et al.* (1989), that labor allocation decisions by farm operators and their spouses are not jointly determined by farm and family labor and income needs.

The next variable considered in the analysis is the size of the household. This variable is expected to be positively correlated with the operator's decision to work off the farm. This is due to the greater demands for cash which would be required to support a larger household, and the greater possibility for unpaid labor on the farm.

The third and fourth of the household characteristics' variables are farm operator household's income from other sources (other than from farming, off-farm earned income, and interest dividends) and household's net worth. These variables, which are included to capture the household's financial characteristics are expected to reduce the likelihood of off-farm employment by the farm operators as they tend to increase the marginal value of time for leisure activity and/or for work on the farm (see, e.g., Robinson, McMahon, and Quiggin; Jensen and Salant; and Gould and Saupe).

The last of the household characteristics' variables is one that describes the life-cycle in farming of farm operators. As designed, the variable is intended to show how the participation decision is affected when the farm operator is an established operator. Simpson and Kapitany note that because off-farm work skills depreciate over time, established farmers are less likely to engage in off-farm work. To the extent that this may be true, the life-cycle-in-farming variable is expected to be positively correlated with farm operators' off-farm labor.

Three farm characteristic variables are hypothesized to affect the decision of farm operators to allocate time to off-farm employment: machinery value per acre operated, participation in government commodity programs, and whether the farm specializes in dairy production.[15] Machinery

[15]A farm operator is considered a participant in government commodity programs if any payment (in cash or certificate) was received for enrolling in any State or Federal farm programs. The payments received are for set-aside or Acreage Reduction Programs, 10-year Conservation Reserve Program rental payments, incentive or cost share payments for conservation practices and improvements, and for any other programs including, e.g., dairy buyout, Federal emergency feed, etc.

value per acre operated is expected to be negatively related to participation because, as Simpson and Kapitany point out, this variable increases farm productivity which in turn increases the marginal returns to farming. Robinson et al. find negative correlation between wealth, as proxied by the amount of resources committed to the farm, and the hours of labor supplied to off-farm employment.

A contrasting view is offered by Albrecht and Murdock, who hypothesize that for types of production where more mechanization can be used, farm operators will have more time to participate in off-farm employment. However, results of their study show weak positive relationships between off-farm work and farm mechanization.

Participation in government commodity programs is expected to be negatively related to off-farm work because direct payments reduce the risk of farming and may, therefore, make the less risky off-farm employment relatively less attractive. Because labor requirements in dairy production remain high and inflexible, despite increased mechanization of dairy farms (Oliveira, p. 8), the correlation between dairy production and operators' off-farm labor participation is expected to be negative, as has been found in previous studies (e.g., Lass, Findeis, and Hallberg; Gould and Saupe).

This report included in the participation model a set of variables that attempt to capture the conditions of the labor market area (LMA). These variables are: the unemployment rate, the percent of LMA's income from agriculture, the change in employment during a recent period (1984-88), and the percent of LMA's employment in various industries.

The unemployment rate is expected to be negatively related to farm operators' off-farm employment since increased unemployment means a greater supply of workers in the labor market area competing for job openings. This negative relationship has been found in previous studies (e.g., McNamara and Gunter, Gunter and McNamara). In contrast, studies by Findeis and by Findeis and Lass have found a positive relationship to exist between unemployment rate and farmers' off-farm work, albeit insignificant.

The percent of the area's labor and proprietor income from agriculture is an indicator of the dependence of the local economy on agriculture. This variable is expected to be negatively related to farmers' off-farm work because of the lower availability of nonfarm jobs and the greater supply of labor of farm people to nonfarm jobs with a similar set of human capital characteristics.[16] Robinson *et al.* found such a negative relationship in their study of off-farm work by Australian farmers. The authors attributed this negative relationship to the likelihood that higher costs of labor market entry (both time and money costs) incur in areas where a greater proportion of the labor force is employed in agriculture.

An increase in total employment during the immediate 5-year historical period is expected to be positively related to participation in farmers' off-farm employment because more jobs will have opened up to new entrants. An increase in total employment, furthermore, is generally an indicator of a healthy, thriving local economy.

The final set of variables considered in the analysis are those that measure the percent of employment in various industries. The availability of jobs that demand human capital characteristics that are relatively abundant in farm people will affect off-farm participation. A study by Hearn, McNamara, and Gunter, for example, found that farm operators are more likely to participate in off-farm employment as the percentage of employment in construction, manufacturing, professional services, and sales increases.

[16]Even farm and farm-related jobs are in most cases located far away from farming areas. Majchrowicz (p. 32), for example, points out that nearly 71 percent of all these jobs were in metro counties in 1989.

In terms of the off-farm wage rate model, the human capital and the LMA variables are expected to influence the wage rates in the same manner they influence the probability of farm operators working off the farm. For example, Robinson et al. note that since human capital is an indicator of an individual's productivity, it can be expected to exert a positive influence on the demand of the operator's off-farm labor by increasing the remuneration obtained from off-farm employment and by increasing the likelihood of obtaining a job. Browne (p. 37) suggests that the higher the unemployment rate, the more time a job seeker is likely to spend in job search and, as a result, the lower are the earnings that can be expected from a decision to seek employment.

Data Considerations

The farm household data used in this report are from the 1988 Farm Operator Resource (FOR) version of the Farm Costs and Returns Survey (FCRS), USDA. The FCRS, which has a complex stratified, multiframe design, is a national, annual survey of farms conducted by the Economic Research Service (ERS) and the National Agricultural Statistics Service (NASS) every February-March since 1985. Each year, the FCRS is composed of multiple versions, all of which collect detailed and consistent information on the farm business. Each of the versions has two sets of expansion factors, or weights, one of which allows the observations to be used in conjunction with the other versions for common data items and the other of which allows for the version to be expanded to the national level using only the observations in that version.[17] The FOR version is dedicated to collect special data on farm operator households regarding the off-farm income and hours worked by the operator and the spouse, their hours worked on the farm, hired employees, and other household details, as well as the standard details on the farm business.

The 1988 version, which represented 1,755,991 farms (based on a sample size of 2,985 farms), was the first year of the FOR version. Not included in these 1.8 million farms, due to FCRS sampling methods, were many of the farms that were in the sales classes of less than $10,000. Because of these sampling methods, the FCRS was known to undercount farms by about 350,000-400,000 based on the official USDA number of farms.[18] This study does not adjust for this undercount, because in order to do so, certain assumptions that are beyond the scope of the analysis about the characteristics of those farms and households that the FCRS sampling methods had missed would have to be made. However, beginning with the 1991 survey, improvements to survey design allowed ERS and NASS to systematically adjust for the undercount of farms and to enhance the adjustments for nonresponse.

For the purpose of this study, some operator households were excluded from the analysis—those where the farms are organized as nonfamily corporations or cooperatives, those whose operators do not receive any of the net income of the farm business, and those whose operators do not have spouses. The final sample count was 2,535, which statistically represents 1,547,414 farm operator households in the 48 contiguous States.

Data on the local area characteristics needed in the analysis are based on county-level data from the Bureau of Economic Analysis income files for 1988, the Bureau of Economic Analysis employment files for 1984 and 1988, and the Bureau of Labor Statistics. The labor market area in this study is defined at the commuting zone level. As in Killian and Tolbert, a

[17]Each observation in the FCRS represents a number of similar farms, the particular number being the survey expansion factor. Each expansion factor, which is the inverse of the probability of the particular farm's being selected, is used by survey users to expand the FCRS sample to represent the population of all farms.

[18]See Ahearn, Perry, and El-Osta (p. 174) for additional information on response rates, interview times, and a discussion of data reliability.

commuting zone is defined as a grouping of counties and is structured to capture local trading areas.

In analyzing the data, the complex sample design of the FCRS imposes significant restrictions to the scope of econometric techniques that can be employed. As Kott notes, standard regression packages not designed to accommodate stratified samples yield biased standard errors, although parameter coefficients are unbiased. Some software packages, such as PC CARP (Fuller et al.) and SUDAAN (Babubhai *et al.*), have been developed to correctly account for the sample design in the computation of variances. However, these are limited to the simplest of models. In the context of off-farm labor participation models, for example, software does not exist that allows for a joint estimation of operator and spouse labor allocations, e.g., a bivariate logit or probit model.

The model of the decision of the farm operator to participate in off-farm employment as described in equation 10 was estimated using the logit analysis for data with complex sample designs in PC CARP. Similarly, the weighted least squares procedure needed for the estimation of farm operators' off-farm wage rates (equation 13) was also estimated using PC CARP.[19]

Empirical Results

Table 2 provides the means of the variables used in the estimation of farm operators' off-farm work as described in equation 10 for the United States, and for data disaggregated by region, by economic size of farm (less than $50,000; $50,000-$99,999; and $100,000 or more), and by type of farm specialization (cash grains; other crops; beef, sheep, hogs; and other livestock including dairy).[20] The results of the estimations, which were performed using weighted logistic regression, are presented in table 3.

The off-farm labor participation model for the United States showed a reasonable fit based on the computed value of pseudo-R^2 (0.31) and had 72.8 percent success in correctly predicting off-farm participation by farm operators.[21] Further, the model was significant at the 1-percent level based on an F-statistic value (not shown because of space limitation) of 8.34 with 18 and infinite degrees of freedom. Variables age and age-squared had the expected sign and were found significant at the 1-percent level. These coefficients indicate, ceteris paribus, that the likelihood of a farm operator participating in off-farm employment increases with age, peaks at age 38, then declines as the operator grows older.

Other variables that were shown to increase significantly the likelihood of participation in off-farm employment were farm operator's

[19]PC CARP was also used in estimating the standard errrors for all reported means and other relevant results.

[20]The regions considered in this study are those defined by the Bureau of the Census. The South includes the East South Central (Alabama, Kentucky, Mississippi, and Tennessee), West South Central (Arkansas, Louisiana, Oklahoma, and Texas), and South Atlantic (Delaware, Florida, Georgia, Maryland, North Carolina, South Carolina, Virginia, and West Virginia) divisions. The West includes the Mountain (Arizona, Colorado, Idaho, Montana, Nevada, New Mexico, Utah, and Wyoming) and Pacific (California, Oregon, and Washington) divisions. The Northeast region includes the New England (Connecticut, Maine, Massachusetts, New Hampshire, Rhode Island, and Vermont) and the Middle Atlantic (New Jersey, New York, and Pennsylvania) divisions. The Midwest region includes the East North Central (Illinois, Indiana, Michigan, Ohio, and Wisconsin) and the West North Central (Iowa, Kansas, Minnesota, Missouri, Nebraska, North Dakota, and South Dakota) divisions.

[21]The pseudo-R^2 (also known as McFadden's) is defined by $R^2 = [1 - L(\beta)/L(0)]$ where L(0) is the maximum of the log likelihood function L subject to the constraint that all the regression coefficients except the intercept are zero, and β is the vector of parameters estimated using the maximum likelihood procedure (Amemiya, p. 1505; Maddala, p. 39). This R^2 will equal 0 (indicating poor fit) if the model predicts operators' off-farm labor participation no better than a simple flip of a coin, and will be equal to 1 if the model predicts off-farm labor participation perfectly.

education, participation of the spouse in the off-farm labor force, and nonfarm net worth of the household. All these variables had the expected sign with the exception of wealth and spouse's participation in off-farm work.

One possible explanation for the unexpected positive relationship between nonfarm wealth and participation may be that, since the nonfarm wealth includes the nonfarm business wealth of farmers who are self-employed at nonfarm jobs, it is reasonable to expect that the likelihood of off-farm participation will increase as the nonfarm wealth increases. The unexpected positive relationship shown to exist between operators and their spouses in terms of off-farm labor participation may reflect a joint purpose, namely, to use earned income from working off the farm to improve the financial position of the farming operation, or to save cash to expand the farm business to a more profitable size. Another plausible explanation is that farm operators derive utility from working off the farm regardless of their spouses' off-farm work status. Reality suggests, as Long and Jones point out in a study of labor force entry and exit by married women, that camaraderie, learning experiences, and satisfaction are valuable aspects of labor force participation.

As expected, participation in government commodity programs was negatively related to off-farm labor participation by farm operators, since direct payments reduce the risk of farming. Similarly, and as one would expect, specialization in dairy production was found to have a significant negative effect on farm operators' participation in off-farm work.

In terms of LMA variables, only one variable exhibited any effect on the likelihood of off-farm labor participation by farm operators—change in LMA's total employment. This variable showed a moderately significant negative influence on the likelihood of off-farm work by farm operators. This unexpected result may suggest that the positive economic prospects associated with an increase in LMA's total employment may discourage current participation just as negative or uncertain prospects may encourage participation.

Table 3 shows generally similar patterns in terms of the explanatory power of the specified off-farm labor participation model as the U.S. sample was disaggregated by region, by size, or by type of farm. The results show that all the models were significant at the 1-percent level, and had around 70 percent success in correctly predicting off-farm participation by farm operators. As for pseudo-R^2, a wider variation in the results is shown to exist, as the R^2 ranged from a low of 0.1880 for farms with sales of $100,000 or more, to a high of 0.4524 for farms in the Northeast region.

In terms of significance of variables, and contrary to expectations, age and age-squared were not significant in the Northeast sample, in the intermediate economic size ($50,000-$99,999) sample, and in the other livestock sample. The education variable was significant only in the South sample, in the sample that contained operators of farms with sales of less than $50,000, and in the beef, sheep, and hog sample. The effect of a spouse's participation in off-farm employment on the operator's off-farm employment was significantly positive, similar to that for the Nation, in all of the disaggregated samples. Exceptions to this were samples in the South and Midwest regions, and in the cash grains and other livestock samples.

The coefficients of the remaining variables in all of the disaggregated samples were only sporadically significant, particularly those that describe the characteristics of LMA's. For example, an increase in the LMA's unemployment rate is shown to cause a decrease, as expected, in the likelihood of off-farm work only in the cash grains sample. Also, an increase in the percentage of LMA's jobs in construction is shown to increase significantly the probability of off-farm work by farm operators in the Midwest and the beef, sheep, hog samples only.

The farm operators' off-farm wage models for the United States and for the disaggregated samples were estimated using weighted least squares in accordance with equation 13. The results, shown in table 4, indicate that the selectivity variable M was significant only in the U.S. sample and in the South and other crops samples. These results further indicate that sample selectivity was present and parameter estimates of farm operators' wage equations in these samples would have been biased and inconsistent had M not been incorporated in the analysis.

For the United States, table 4 demonstrates the importance of age, which can be considered a proxy for experience, and education in explaining farm operators' off-farm wages. For example, an increase in the age of the operator by 1 year is shown to increase his/her hourly wage rate by almost 12 percent, while an additional year of schooling causes the wage rate to increase by almost 7 percent. None of the variables that describe local area characteristics were found to add significantly to off-farm wages.[22] When the data were disaggregated by region, by size of farm, and by type of farm, the age and education variables became insignificant in the samples that contained farms in the West and Northeast regions, in the middle ($50,000-$99,999) range of economic size, and in the other livestock group of farm types. The education variable was particularly important for farm operators in the $100,000 economic size group, as a 1-year increase in educational attainment correlates with a 22-percent increase in wage rates. Similar to the results in table 3 for disaggregated data, table 4 shows that the importance of the LMA variables can be characterized as sporadic.

Parameter estimates of farm operators' wage equations, as shown in table 4, were used to predict the off-farm wage rates for those who worked off the farm along with those who did not (see equation 15). These predicted wages serve as proxies for the opportunity cost of farm operators' unpaid farm labor. Table 5 presents the means, medians, and quantile distributions of these values for the United States, and by region of production, by economic size of farm, and by type of farm specialization.

Based on the U.S. sample, table 5 shows that in 1988, the average predicted opportunity cost of farm operators' unpaid labor was $9.68 per hour. Based on regional disaggregation of the data, the opportunity cost of farm operators' labor ranged from a low of $8.24 per hour in the South to a high of $15.74 in the West. In terms of economic size of farm, the results show that farm operators with farm sales of $50,000 or more have opportunity costs that stand at about twice the predicted level for those with sales of less than $50,000. Table 5 reveals that the opportunity cost of farm operators' unpaid labor is highest for cash grains farmers ($16.38 per hour), compared to other farmers with other types of farm specialization, while producers of beef, hogs, and sheep have the lowest costs ($8.58 per hour).

Concluding Remarks

This study employed a procedure to impute opportunity costs of farm operators' unpaid labor. The procedure is based on farmers' off-farm labor participation and their own levels of human capital. A two-stage procedure was used to, first, specify a logistic regression to estimate a sample selectivity variable and then, to estimate unbiased and consistent estimates of off-farm wage rates using weighted least squares regression on a human capital wage model. Regression results indicate that the level of education of farm operator is positively associated with imputed opportunity costs of unpaid labor. This was found to be true regardless of whether these costs were estimated by region, by size of farm, or by farm type. The effect of labor market

[22]Only variables with t-ratios larger than 1 are kept in the model. The purpose of doing this, a technique which was originated by Dhrymes, is to increase the explanatory power of the model (Pindyck and Rubinfeld, p. 78).

characteristics on the opportunity cost of unpaid labor was minimal.

The approach suffers from the same weakness that all attempts to value entrepreneurial skills do, that is, it is difficult to fully capture these skills by easily quantifiable human capital variables. This may be especially true in agriculture where credentials and skills can be very specific. Formal education may have limited practical application on a specific farm. Although this result could be considered as a deficiency of the valuation method, it is likely the best alternative available to agricultural economists at present. In particular, it avoids the ad hoc valuation approaches commonly employed by the profession.

Future work could include, subject to data availability, cost of living and locational amenities, as some studies have found that such variables may be related to the decision of farm operators to work off the farm. Also, contingent on the availability of newer regression software that would allow for the modeling of bivariate logit or bivariate probit when data are based on surveys with complex design as in this study, research on off-farm work participation and consequently research on off-farm wage rates should consider the decision of farm operators to participate in off-farm employment from a family perspective. This is relevant due to the possibility of the jointness of such decision by the operators and their spouses.

Nevertheless, until all of these limitations are overcome, ERS will utilize the method of valuing operator's unpaid labor outlined in this report in its future estimation of commodity costs and returns. This procedure, along with the implementation of the recommendations of the forthcoming report by the AAEA Taskforce on the Standardization of Commodity Costs and Returns Methods should improve the accuracy of the structure of commodity costs and returns accounts.

Table 1-- Definition of variables

Variable	Definition
OL	Operator off-farm labor force participation: coded 1 if the operator receives any income from off-farm wages and salaries and/or from an off-farm business; 0 otherwise
Operator characteristics:	
A	Age of the operator
AS	Age of the operator squared
E	Education of the operator (years)
Household characteristics:	
SL	Spouse off-farm labor force participation dummy: 1=spouse works off the farm
HS	Size of the household
OI[1]	Unearned income of the household ($1,000)
W	Nonfarm net worth of the household ($1,000)
LC	Life-cycle in farming dummy: 1=the operator expects to continue farming the size of farm over the next 5 years with or without participation in off-farm activity
Farm characteristics:	
VM	Value of machinery per acre ($1,000)
G	Government program dummy: 1=farm receives government payments
D	Farm specialty dummy: 1=dairy
Labor market area (LMA) characteristics:	
U	Unemployment rate (%)
AI	Percent of LMA's income from agriculture
EC	LMA's employment change, 1984-88 (%)
M	Percent of LMA's employment in manufacturing
C	Percent of LMA's employment in construction
S	Percent of LMA's employment in services
RT	Percent of LMA's employment in wholesale and retail trade

[1] Includes income from retirement (e.g., social security, private pensions), public assistance, disability , unemployment compensation, supplemental security income, and income from other sources (e.g., off-farm rental income, gifts, etc.)

Table 2--Mean values of variables by region, by economic size of farm, and by type of farm, 1988

Category	OL	A	E	SL	HS	OI	W	LC	VM	G	D	U	AI	EC	M	C	S	RT	Sample	Population
United States	0.46	54	12.38	0.34	2.92	4.25	61.70	0.56	0.35	0.35	0.08	6.32	6.00	6.24	16.68	4.99	21.67	20.30	2,535	1,547,414
Region																				
South	0.51	55	12.22	0.37	2.63	4.57	60.24	0.57	0.30	0.21	0.03	7.04	4.67	6.93	18.81	5.32	20.16	19.59	1,099	693,913
West	0.43	54	12.98	0.29	2.95	5.23	90.38	0.58	0.56	0.15	0.03	7.09	6.87	8.59	11.88	5.08	23.67	20.66	484	197,430
Northeast	0.43	51	12.47	0.28	3.68	6.28	49.41	0.49	0.42	0.27	0.31	4.69	0.97	9.93	18.53	5.43	25.96	21.66	221	94,484
Midwest	0.42	52	12.35	0.32	3.13	3.18	55.48	0.56	0.33	0.60	0.13	5.43	8.20	3.97	15.41	4.47	22.12	20.81	731	561,586
Size ($1,000)[1]																				
Less than $50	0.56	55	12.22	0.36	2.77	5.13	53.11	0.56	0.36	0.23	0.02	6.52	4.82	7.29	17.44	5.08	21.59	20.14	1,307	1,123,874
$50-$99.99	0.27	49	12.52	0.26	3.31	2.35	40.56	0.55	0.35	0.57	0.27	5.81	8.78	4.04	14.98	4.76	21.62	20.57	325	153,485
$100 or more	0.16	48	12.93	0.28	3.30	1.67	109.46	0.57	0.33	0.72	0.23	5.78	9.35	3.18	14.47	4.71	22.04	20.78	903	270,055
Type of farm[2]																				
Cash grains	0.45	53	12.59	0.38	2.85	3.75	62.27	0.51	0.20	0.80	0.00	5.79	9.87	2.04	14.73	4.61	21.54	20.72	474	268,161
Other crops[3]	0.45	55	12.43	0.31	2.80	5.52	55.66	0.57	0.51	0.27	0.00	6.67	5.36	7.38	16.87	5.13	21.95	20.24	603	287,737
Beef, sheep, hogs	0.51	55	12.28	0.34	2.80	4.43	65.92	0.59	0.29	0.25	0.00	6.52	5.14	6.96	16.98	5.06	21.43	19.99	1,044	790,419
Other livestock[4]	0.30	49	12.40	0.32	3.62	2.41	52.96	0.52	0.59	0.28	0.63	5.74	5.19	7.45	17.82	5.01	22.41	21.02	414	201,096

Note: Estimates that are underlined have coefficients of variation in the range of 25 to 35 percent. All other estimates have coefficients of variation of less than 25 percent.

[1] Based on the operation's and landlord's value of agricultural sales and government payments.

[2] Based on the type of farm production that generates the largest portion of the gross income from the farming operation.

[3] Includes tobacco, cotton, other field crops, vegetables, fruits, nuts, and nursery or greenhouse.

[4] Includes animal specialties, poultry, and dairy.

Table 3--Weighted logit results of off-farm labor participation by region, by economic size of farm, and by type of farm, 1988

Category	Coefficients																			
	β_O	β_A	β_{AS}	β_E	β_{SL}	β_{HS}	β_{OI}	β_W	β_{LC}	β_{VM}	β_G	β_D	β_U	β_{AI}	β_{EC}	β_M	β_C	β_S	β_{RT}	R^2
United States	-2.904	0.150	-0.002	0.081	0.911	0.038	-0.042	0.001	0.116	-0.089	-1.413	-2.354	-0.029	-0.013	-0.016	0.019	0.099	-0.026	0.024	0.3111
	(-1.86[a])	(3.27[c])	(-4.58[c])	(2.17[b])	(5.84[c])	(0.66)	(-2.05[b])	(1.93[a])	(0.79)	(-1.90[a])	(-7.83[c])	(-7.05[c])	(-0.84)	(-0.86)	(-1.92[a])	(1.44)	(1.48)	(-0.88)	(0.65)	[72.8]
Region																				
South	-3.295	0.143	-0.002	0.166	1.035	0.104	-0.043	0.001	0.026	-0.170	-1.232	-3.950	-0.005	-0.029	-0.009	0.011	0.138	-0.039	-0.023	0.3217
	(-1.41)	(1.97[b])	(-2.75[c])	(2.98[c])	(4.34[c])	(0.98)	(-1.16)	(1.69[a])	(0.12)	(-1.02)	(-3.98[c])	(-6.26[c])	(-0.08)	(-1.19)	(-0.72)	(0.70)	(1.25)	(-0.82)	(-0.36)	[70.8]
West	-8.847	0.361	-0.004	0.051	0.218	-0.178	-0.073	-0.0004	0.383	-0.011	-2.739	-2.414	-0.024	-0.001	-0.032	-0.016	0.275	-0.090	0.187	0.3632
	(-1.99[b])	(3.02[c])	(-3.44[c])	(0.67)	(0.56)	(-1.52)	(-2.16[b])	(-0.44)	(1.12)	(-0.11)	(-5.10[c])	(-2.95[c])	(-0.27)	(-0.02)	(-1.73[a])	(-0.41)	(1.22)	(-1.32)	(1.77[a])	[78.9]
Northeast	2.713	-0.085	-0.0003	0.047	0.422	-0.073	-0.075	0.001	-0.537	-0.063	-0.903	-2.885	0.192	0.040	-0.044	-0.122	-0.233	-0.066	0.363	0.4524
	(0.27)	(-0.36)	(-0.11)	(0.35)	(0.72)	(-0.49)	(-1.35)	(0.30)	(-0.74)	(-0.21)	(-1.44)	(-3.44[c])	(0.54)	(0.05)	(-0.51)	(-1.25)	(-0.46)	(-0.50)	(1.24)	[76.9]
Midwest	-4.023	0.172	-0.002	0.027	1.120	0.072	-0.029	0.002	0.206	-0.087	-1.872	-2.618	-0.054	0.010	-0.004	0.036	0.476	0.009	-0.024	0.3383
	(1.41)	(2.36[b])	(-3.29[c])	(0.35)	(3.84[c])	(0.76)	(-1.29)	(2.34[b])	(0.79)	(-1.30)	(-5.72[c])	(-4.81[c])	(-0.75)	(0.36)	(-0.24)	(1.42)	(2.44[b])	(0.20)	(-0.38)	[71.7]
Size ($1,000)																				
Less than $50.00	-3.664	0.159	-0.002	0.158	0.696	0.151	-0.076	0.002	0.232	-0.010	-0.792	-1.319	-0.039	0.007	-0.016	0.017	0.068	-0.004	0.006	0.4017
	(-1.85[a])	(2.65[c])	(-3.89[c])	(3.20[c])	(2.94[c])	(1.81[a])	(-3.18[c])	(2.23)	(1.10)	(-0.16)	(-3.15[c])	(-2.47[b])	(-0.80)	(0.33)	(-1.48)	(1.12)	(0.72)	(-0.10)	(0.11)	[81.4]
$50.00-$99.99	-6.577	0.013	-0.001	0.110	1.074	0.031	-0.037	0.001	-0.488	-0.183	-0.860	-1.303	0.076	0.003	-0.043	0.041	0.395	-0.046	0.197	0.2446
	(-1.37)	(0.07)	(-0.40)	(0.98)	(2.26[b])	(0.28)	(-0.84)	(1.33)	(-0.91)	(-0.41)	(-1.52)	(-1.80[a])	(0.75)	(0.08)	(-1.34)	(0.95)	(1.89)	(-0.60)	(1.74[a])	[74.8]
$100.00 or more	-5.880	0.147	-0.002	-0.007	1.160	-0.072	0.006	0.002	-0.108	-0.029	-0.058	-0.756	-0.107	0.010	-0.039	0.045	0.038	-0.075	0.164	0.1880
	(-1.39)	(1.15)	(-1.56)	(-0.07)	(3.46[c])	(-0.49)	(1.09)	(2.50[b])	(-0.31)	(-0.32)	(-0.14)	(-1.62)	(-1.63)	(0.44)	(-2.32[b])	(1.80)	(0.23)	(-1.35)	(2.35[b])	[85.0]
Type of farm																				
Cash grains	-2.015	0.174	-0.002	0.049	1.215	0.255	0.003	0.001	0.749	0.587	-2.221	--	-0.193	0.001	-0.029	0.005	-0.037	-0.037	0.031	0.2932
	(-0.65)	(1.87[a])	(-2.44[b])	(0.57)	(3.85[c])	(1.98[b])	(0.58)	(1.26)	(2.28)	(0.85)	(-5.92[c])	--	(-3.34[c])	(0.03)	(-1.77[a])	(0.21)	(-0.26)	(-0.56)	(0.33)	[69.2]
Other crops	0.148	0.165	-0.002	0.100	0.288	-0.086	-0.068	0.001	0.061	-0.467	-1.433	--	0.093	-0.082	-0.016	-0.009	0.050	-0.067	-0.056	0.3362
	(0.04)	(1.74)[a])	(-2.48[b])	(1.39)	(0.84)	(-0.82)	(-1.67[a])	(0.71)	(0.18)	(-2.15[b])	(-3.72[c])	--	(1.18)	(-1.96[b])	(-0.82)	(-0.31)	(0.33)	(-1.13)	(-0.59)	[69.5]
Beef, sheep, hogs	-4.742	0.189	-0.003	0.113	0.997	0.082	-0.099	0.002	-0.158	0.091	-1.567	--	0.017	-0.031	-0.011	0.029	0.309	-0.016	-0.015	0.4123
	(-1.86[b])	(2.72[c])	(-3.69[c])	(1.86[a])	(3.91[c])	(0.77)	(-4.52[c])	(1.91[a])	(-0.63)	(1.11)	(-5.20[c])	--	(0.27)	(-1.14)	(-0.79)	(1.37)	(2.84[b])	(-0.31)	(-0.26)	[74.4]
Other livestock	5.774	-0.033	-0.0004	0.012	0.902	-0.052	0.006	-0.0005	0.887	-0.016	-0.400	-2.182	-0.085	-0.111	-0.003	-0.066	-0.278	-0.096	0.107	0.2790
	(1.08)	(-0.27)	(-0.35)	(0.10)	(1.61)	(-0.45)	(0.20)	(-0.48)	(1.89[a])	(-0.16)	(-0.74)	(-3.79[c])	(-0.67)	(-2.16[b])	-0.09	(-1.18)	(-1.38)	(-1.04)	(0.81)	[82.9]

Note: Figures inside parentheses are t-ratios. Figures inside brackets indicate the percent of the individual responses that are correctly classified.
[a] Significant at 90% level. [b] Significant at 95% level. [c] Significant at 99% level.

Table 4--Weighted least squares estimates of farm operators' off-farm wage equations by region, by economic size of farm, and by type of farm, 1988

	Coefficients[1/]															
Category	β_0	β_A	β_{AS}	β_E	β_U	β_{AI}	β_{EC}	β_M	β_C	β_S	β_{RT}	λ	F-statistic[2/]	R^2	Sample	Population
United States	-1.117	0.119	-0.001	0.071								-0.253	10.11	0.0889	912	711,775
	(-1.91)	(4.83[c])	(-4.79[c])	(3.12[c])								(-2.29[b])	[4, 528][c]			
Region																
South	-2.056	0.151	-0.002	0.080			-0.008					-0.468	9.44	0.1364	421	351,724
	(-2.61[c])	(4.56[c])	(-4.64[c])	(2.24[c])			(-1.96[b])					(-2.92)[c]	[5, 252][c]			
West	2.800					-0.018	0.007					-0.136	3.84	0.0503	164	84,855
	(2.10[b])					(-2.23[b])	(1.35)					(-0.55)	[3, 80][c]			
Northeast	0.672	0.098	-0.001									-0.506	0.50	0.0924	80	41,075
	(0.34)	(1.07)	(-1.11)									(-0.91)	[3, 49]			
Midwest	-0.865	0.098	-0.001	0.091			0.009					0.147	4.33	0.1405	247	234,121
	(-0.82)	(2.66[c])	(-2.34[b])	(2.61[c])			(1.23)					(0.92)	[5, 148][c]			
Size ($1,000)																
Less than $50	-1.277	0.123	-0.001	0.056		-0.001				0.015		-0.260	6.66	0.1021	708	625,627
	(-1.98)	(4.21[c])	(-3.92[c])	(2.42[b])		(-1.82[a])				(1.93[a])		(-1.50)	[6, 382][c]			
$50-$99.999	6.275										-0.168	-0.093	2.72	0.1356	86	41,790
	(3.98[c])										(-2.30[a])	(-0.33)	[2, 48][b]			
$100 or more	-0.946			0.224			-0.016		0.169			-0.073	2.76	0.1887	118	44,359
	(-0.75)			(2.69[c])			(-1.16[a])		(1.19)			(-0.20)	[4, 70][b]			
Type of farm																
Cash grains	-1.959	0.104	-0.001	0.124					0.127			0.229	5.74	0.1653	168	119,513
	(-1.64)	(3.14[c])	(-2.94[c])	(2.33[b])					(1.86[b])			(1.11)	[5, 100][c]			
Other crops	-1.242	0.114	-0.001	0.067				0.016		0.041	-0.051	-0.531	2.15	0.1381	199	128,677
	(-0.79)	(2.13[a])	(-2.20[b])	(2.31[b])				(1.51)		(1.87)	(-1.22)	(-2.81[b])	[7, 114][b]			
Beef, sheep, hogs	-1.221	0.131	-0.001	0.060								-0.290	7.16	0.0963	457	403,565
	(-1.68)	(3.83[c])	(-3.95[c])	(1.86[a])								(-1.55)	[4, 248][c]			
Other livestock	0.033	0.013			0.164	-0.054			0.253			-0.014	11.03	0.2698	88	60,022
	(0.04)	(1.37)			(2.54[b])	(-2.54[b])			(2.04[a])			(-0.06)	[5, 48][c]			

[1/] Figures inside parentheses are t-ratios.
[2/] This statistic tests whether all model parameters (except intercept) = 0 with first figure inside brackets denoting number of coefficients tested and the second figure denoting the number of segments minus the number of strata. [a] Significant at 90% level. [b] Significant at 95% level. [c] Significant at 99% level.

Table 5--Predicted opportunity costs of farm operators' unpaid labor by region, by economic size of farm, and by type of farm, 1988

Category	Quantiles ($)[1] 0.05	0.10	0.25	0.50	0.75	0.90	0.95	Mean ($)[1]	Interquartile range ($)[1]	Sample	Population	Share of operators
United States[2]	3.43	4.43	6.83	9.48	12.34	14.76	16.54	9.68	5.51	2,911	1,727,560	100.0
	(0.21)	(0.18)	(0.13)	(0.16)	(0.15)	(0.19)	(0.29)	(0.12)	(0.16)			
Region												
South	1.79	2.82	4.54	7.36	11.38	13.60	16.74	8.24	6.84	1,264	779,490	45.1
	(0.22)	(0.20)	(0.21)	(0.33)	(0.22)	(0.41)	(0.89)	(0.20)	(0.25)			
West	9.45	11.66	13.95	16.02	17.86	19.38	20.14	15.74	3.91	563	224,332	13.0
	(0.66)	(0.58)	(0.25)	(0.22)	(0.30)	(0.36)	(0.25)	(0.20)	(0.36)			
Northeast	4.46	5.17	6.50	11.22	14.75	17.73	20.02	11.53	8.25	257	109,309	6.3
	(0.34)	(0.45)	(0.48)	(0.96)	(0.50)	(0.94)	(1.78)	(0.55)	(0.59)			
Midwest	6.50	7.24	9.53	12.14	14.97	18.09	20.08	12.49	5.44	827	614,430	35.6
	(0.26)	(0.31)	(0.39)	(0.26)	(0.32)	(0.53)	(0.78)	(0.26)	(0.39)			
Size ($1,000)												
Less than $50	3.03	3.83	6.09	9.03	12.15	13.97	15.41	9.13	6.06	1,459	1,246,730	72.2
	(0.21)	(0.20)	(0.16)	(0.26)	(0.16)	(0.23)	(0.35)	(0.15)	(0.16)			
$50-$99.999	9.87	10.42	12.44	16.24	21.98	27.86	39.21	18.75	9.54	378	173,391	10.0
	(0.47)	(0.47)	(0.51)	(0.55)	(1.04)	(3.76)	(4.55)	(0.89)	(0.97)			
$100 or more	7.08	8.14	10.52	14.03	20.73	31.46	39.23	17.21	10.21	1,074	307,440	17.8
	(0.35)	(0.43)	(0.34)	(0.45)	(1.25)	(2.20)	(2.12)	(0.57)	(1.26)			
Type of farm												
Cash grains	7.45	9.17	11.28	15.15	19.51	24.37	29.34	16.38	8.22	551	299,476	17.3
	(0.52)	(0.43)	(0.39)	(0.51)	(0.50)	(0.72)	(2.30)	(0.41)	(0.61)			
Other crops	2.80	3.64	5.07	8.11	12.85	15.99	18.16	9.33	7.79	717	337,645	19.5
	(0.22)	(0.16)	(0.24)	(0.45)	(0.37)	(0.55)	(0.85)	(0.30)	(0.39)			
Beef, sheep, hogs	2.64	3.37	5.66	8.57	11.36	13.47	14.86	8.58	5.70	1,175	861,160	49.8
	(0.21)	(0.18)	(0.22)	(0.27)	(0.18)	(0.29)	(0.52)	(0.19)	(0.20)			
Other livestock	3.73	6.45	10.09	14.73	19.83	25.47	30.11	15.75	9.75	468	229,280	13.3
	(1.37)	(0.87)	(0.46)	(0.71)	(0.80)	(1.56)	(2.44)	(0.55)	(0.83)			

[1] Figures inside parentheses are standard errors.
[2] U.S. figures here are generated based on U.S. wage model (see table 4).

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