Effect of Farm Income and Off-Farm Wage Variability on Off-Farm Labor Supply

Ashok K. Mishra and Duncan M. Holthausen

This study models the effects of variability in farm income and off-farm wages on farm operators' labor allocation decisions. A simple theoretical model is employed to develop hypotheses, which are then tested empirically. Variability in farm income and off-farm wages is predicted to have a positive and negative effect, respectively, on off-farm hours worked. The empirical results confirm these predictions.

Key Words: farm income variability, off-farm labor supply, off-farm wage, time allocation

Agriculture in the United States changed dramatically during the 20th century. At one time, the farm was the sole source of income for most U.S. farmers and their families. For example, the 1939 Census of Agriculture estimated that about 29% of all farm operators held off-farm employment. By 1997, this number had risen to approximately 58% (table 1). It is interesting to note that the percentage of farm operators working less than 100 days off the farm declined over the 1929–1997 time period. In contrast, the percentage of farm operators reporting off-farm employment of 200 days or more increased from about 6% in 1929 to about 39% in 1997.

Traditionally, off-farm employment was considered to be a temporary condition for those involved; i.e., off-farm work was thought to be reserved for those trying to acquire skills and capital for entrance into farming on a full-time basis, or as a mechanism for easing the exit of retiring or marginal producers from agriculture. However, those motivations no longer appear to be the primary reasons for working off the farm.

The increased reliance on off-farm income by farm operators has been well documented in a number of studies (e.g., Ahearn, 1986; Perry and Hoppe, 1993; Sumner, 1982; Gunter and McNamara, 1990; Mishra and Goodwin, 1997; Mishra, 1996; and Hallberg, Findeis, and Lass, 1991). Current data from the U.S. Department of Agriculture (USDA) indicate almost 90% of U.S. farm operator households receive some off-farm income from either earned or unearned sources (Ahearn, Perry, and El-Osta, 1993; Perry and Hoppe, 1993; Sommer et al., 1997).

Table 2 reports average farm operator household income, from both farm and off-farm sources, over the period 1988–2000. As observed from these data, the average farm household earns much more in off-farm income than it earns on the farm. Also, since 1996, average total farm household income has been greater than average U.S. household income, with income levels in 2000 of $61,947 and $57,045, respectively.

It is well known that farm income is much more variable than nonfarm income because of the riskiness of the farming business (Mishra and Goodwin, 1997). To reduce income risks and raise total income, farm families have turned to off-farm work to supplement farm household income (Ahearn and Lee, 1991; Fuller, 1991; Barlett, 1991; Spitze and Mahoney, 1991). Mishra and Goodwin (1997), and Mishra (1996) concluded that as farm income variability increases, farm families seek off-farm employment (as a source of income) to reduce the variance in their household income. One drawback of the Mishra and Goodwin (1997) study is the assumption that off-farm wages are fixed.

Ashok K. Mishra is an agricultural economist, Resource Economics Division, Economic Research Service, U.S. Department of Agriculture, Washington, DC. Duncan M. Holthausen is a professor in the Department of Economics, North Carolina State University, Raleigh. The views expressed here are the authors' and do not necessarily represent those of the Economic Research Service or the U.S. Department of Agriculture.

Agricultural and Resource Economics Review 31/2 (October 2002): 187–199

Table 1. Farm Operators Working On and Off the Farm in the United States, Selected Census Years, 1929–1997

	No. Operators Working Off-Farm					
Census Year	1–99 Days (000s)	100–199 Days (000s)	> 200 Days (000s)	Total (000s)	No. Operators Not Working Off-Farm (000s)	Total Farm Operators (000s)
1929[a]	1,179	327	397	1,903	4,386	6,289
1934[a]	1,317	348	413	2,078	4,735	6,813
1939[a]	804	379	565	1,748	4,350	6,098
1944[a]	491	244	835	1,570	4,289	5,859
1949[a]	836	313	944	2,093	3,293	5,386
1954	820	306	1,027	2,154	2,629	4,783
1959	556	230	878	1,664	2,044	3,708
1964	449	189	824	1,462	1,696	3,158
1969	392	219	871	1,482	1,248	2,730
1974	197	157	658	1,012	1,268	2,280
1978	252	181	770	1,203	1,055	2,258
1982	224	189	775	1,188	1,054	2,242
1987	200	178	737	1,115	844	1,959
1992	165	162	666	993	802	1,795
1997	165	168	709	1,042	755	1,797
	Percent (%)					
1929	18.8	5.2	6.3	30.3	67.7	100.0
1934	19.3	5.1	6.1	30.5	69.5	100.0
1939	13.2	6.2	9.3	28.7	71.3	100.0
1944	8.4	4.2	14.2	26.8	73.2	100.0
1949	15.5	5.8	17.5	38.9	61.1	100.0
1954	17.1	6.4	21.5	45.0	55.0	100.0
1959	15.0	6.2	23.7	44.9	55.1	100.0
1964	14.2	6.0	26.1	46.3	53.7	100.0
1969	14.3	8.0	31.9	54.3	45.7	100.0
1974	8.6	6.9	28.9	44.4	55.6	100.0
1978	11.2	8.0	34.1	53.3	46.7	100.0
1982	10.0	8.4	34.6	53.0	47.0	100.0
1987	10.2	9.1	37.6	57.0	43.0	100.0
1992	9.2	9.0	37.1	55.3	44.7	100.0
1997	9.2	9.3	39.4	58.0	42.0	100.0

Sources: *Census of Agriculture*, U.S. Department of Commerce (various years cited above), and U.S. Department of Agriculture (1997).

[a] Information about off-farm work was collected for each of these years during the census of the following year (for example, 1929 information was collected in 1930).

Table 2. Average Income of Farm Operator Households in the United States, 1988–2000

Description	1988	1989	1990	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000
	$ / Farm Operator Household												
Farm Income to the Household[a]	4,201	5,796	5,742	5,810	7,180	4,815	4,376	4,720	7,906	6,205	7,106	6,359	2,598
Self-employment	3,837	4,723	4,973	4,458	5,172	3,623	3,407	4,059	6,009	4,971	5,941	5,415	2,260
Other farm income	364	1,073	768	1,352	2,008	1,192	970	661	1,898	1,234	1,165	944	339
+ Tot. Off-Farm Income	28,829	26,223	33,265	31,638	35,731	35,408	38,092	39,190	42,455	46,358	52,628	57,988	59,349
= Avg. Farm Operator Household Income	33,030	32,019	39,007	37,447	42,911	40,223	42,469	44,392	50,361	52,562	59,734	64,347	61,947
U.S. Average Household Income	34,017	36,520	37,103	37,922	38,840	41,428	43,133	44,938	47,123	49,692	51,855	54,842	57,045

Source: *Agricultural Outlook*, Economic Research Service, USDA (selected issues).

Note: Farm operator households are defined as the households of operators with farms organized as individual operations, partnerships, and family corporations. The definition excludes households with farms organized as nonfamily corporations or cooperatives, as well as households where the operator is a hired manager. The national average is 1.1 households per farm sharing the income of a farm business.

[a] Farm income to the household equals self-employment income plus amounts that operators pay themselves and family members to work on the farm, income from renting acreage, and net income from a farm business other than the one being surveyed. Data for 1988–90 are based on surveys that did not fully account for small farms. Data for 1991 include an additional 350,000 farms, many with gross sales under $10,000 and negative farm incomes.

In this study, we assume variability exists both in net farm income and in off-farm wages. We hypothesize a risk-averse farm operator takes the riskiness of both farming income and off-farm wages into consideration when making the decision to work off the farm.

The objective of this study is to investigate the effect of farm income and off-farm wage variability on off-farm labor force participation by farm operators. First, we develop a model of off-farm work by farm operators facing variability in farm income and off-farm wages. We then test the predictions from the model using data from Kansas and North Carolina. The empirical results support our hypothesis. In particular, our findings reveal greater off-farm labor force participation as the variability of farm income increases and as the variability of off-farm wages decreases.

Theoretical Model

Consider a farm operator who has income-generating options in farming and in off-farm work. It is assumed perfect competition exists in the labor market, and therefore farm operators' labor allocation decisions have no effect on aggregate demand, supply, and price of labor. In a framework that recognizes risk, the farmer is assumed to maximize the expected value of a von Neumann-Morgenstern utility function subject to production and time constraints. In general, utility is a function of income and leisure:

$$(1)\quad U = U(\pi, L),$$

where π represents net income and L is leisure time. It is assumed that $U = U(\pi, L)$ is increasing in π and L, and is strictly concave.

Farm income may be uncertain for a variety of reasons, but to keep things simple, we assume the uncertainty is due solely to price fluctuations. Thus, product price (P) and off-farm wage (W) are random in our model (represented, respectively, as $\tilde{P}$ and $\tilde{W}$, where the tilde indicates a random variable). The farm operator's net income is given as:

$$(2)\quad \tilde{\pi} = \tilde{P}Q(F) - C(Q(F)) + I + \tilde{W}H,$$

where F is time allocated to farm work, H is time allocated to off-farm work, $Q(F)$ is total farm output as a function of F, $C(Q)$ is the cost of producing Q units of output, and I represents nonearned income. We assume the production function is concave in F, and the amount of capital used in farm production is fixed in the short run.

To further simplify the model, we assume leisure is fixed[1] so that $F + H = T$, where T is total time allocated to work. Substituting $T - F$ for H in (2), and dropping the argument F of the output function $Q(F)$, net income is specified as:

$$(3)\quad \tilde{\pi} = \tilde{P}Q - C(Q) + I + \tilde{W}(T - F).$$

Since leisure is fixed, utility in (1) is a function of net income alone, and thus the farmer wants to maximize the expected utility of net income, $E[U(\pi)]$. The first-order condition for maximization with respect to time worked on the farm is designated by:

$$(4)\quad \frac{\partial EU(\pi)}{\partial F} = E\left[U'(\pi)(Q_F(\tilde{P} - C') - \tilde{W})\right] = 0,$$

where Q_F is the derivative of Q with respect to F (i.e., the marginal product of time spent on farm production), and C' is the derivative of C with respect to Q (i.e., the marginal cost of production excluding the cost of the farm operator's time).

Equation (4) can be rewritten as:

$$E\left[U'(\pi)Q_F(\tilde{P} - C')\right] = E\left[U'(\pi)\tilde{W}\right].$$

Since each side of the equation is the expectation of the product of two random variables, we can again rewrite this expression as:

$$EU'(\pi)Q_F(E(\tilde{P}) - C') + Q_F\operatorname{cov}\left[U'(\pi), \tilde{P}\right] = EU'(\pi)E(\tilde{W}) + \operatorname{cov}\left[U'(\pi), \tilde{W}\right].$$

Dividing both sides by $EU'(\pi)$ yields:

$$(5)\quad Q_F(E(\tilde{P}) - C') + \frac{Q_F\operatorname{cov}\left[U'(\pi), \tilde{P}\right]}{EU'(\pi)} = E(\tilde{W}) + \frac{\operatorname{cov}\left[U'(\pi), \tilde{W}\right]}{EU'(\pi)}.$$

The left-hand side of (5) is the risk-adjusted expected marginal farm profit $E[MFP_{RA}]$, and the right-hand side is the risk-adjusted expected wage EW_{RA}.

[1] We have no data on farm operators' leisure, and so have chosen to simplify the model by assuming leisure is fixed. However, the same qualitative results can be derived if the utility function is separable in income and leisure. In particular, if utility is given by $U(\pi, L) = V(\pi)\Psi(L)$, the first-order conditions with respect to F and H imply the condition given by equation (4) in the text.

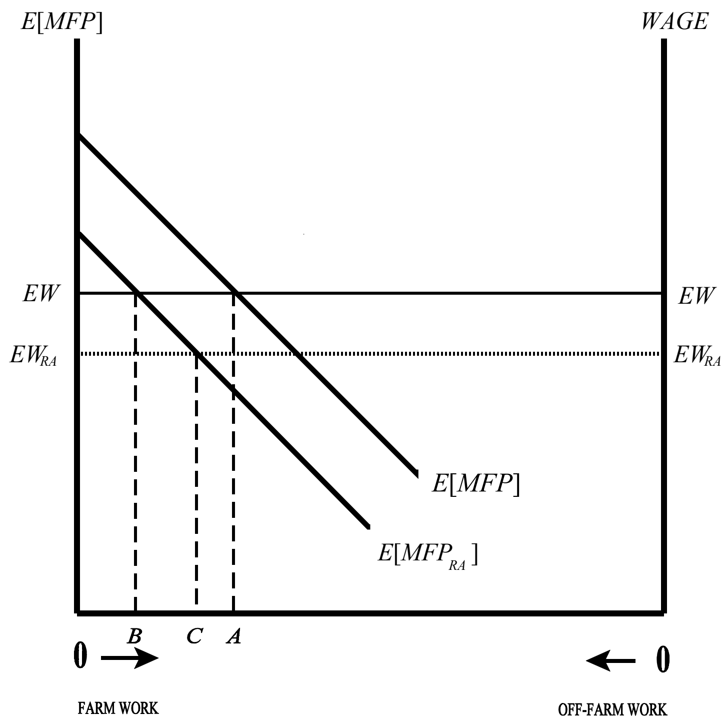

Figure 1. Risk and time allocation

Since both covariances are negative,[2] the risk-adjusted values are less than the expected values.

Some implications of equation (5) are illustrated in figure 1. A risk-neutral farmer would divide his labor between farm and nonfarm work such that the expected marginal farm profit from an additional hour of farm work, $E[MFP]$, equals the expected off-farm wage, EW. In figure 1, this would lead the farmer to choose A hours of on-farm work.

If a producer were risk averse, her risk-adjusted expected marginal farm profit curve, $E[MFP_{RA}]$, would be lower than the corresponding curve for the risk-neutral farmer. The difference between the expected marginal profit of the risk-neutral farmer and the expected marginal profit of the risk-averse farmer is a compensating differential which reflects the riskiness of farm income and the risk aversion of the farmer. This differential is given by the second term on the left-hand side of (5).

If the wage for off-farm work were certain to be EW, the risk-averse farmer would choose to work B hours on the farm. This is less than the A hours chosen by the risk-neutral farmer. If the off-farm wage were also uncertain, the risk-averse farmer would discount the expected wage just as she does the expected marginal farm profit. If the risk-adjusted expected wage is EW_{RA}, the risk-averse farmer will choose to work C hours on the farm. This is still less than the A hours chosen by the risk-neutral farmer in figure 1, but if the EW_{RA} line were drawn somewhat lower, C could move to the right of A. Thus, when both farm income and off-farm wages are uncertain, a risk-averse farmer may choose to work more or less hours on the farm than an otherwise identical risk-neutral farmer.

Changes in the riskiness of the farmer's employment alternatives will bring about changes in the allocation of labor between farm and off-farm work for risk-averse producers. Specifically, as the riskiness of farm income rises, $E[MFP_{RA}]$ falls and the farmer will devote fewer hours to farm work. Conversely, as the riskiness of off-farm wages increases, EW_{RA} falls and the farmer will work fewer hours off the farm.

In order to derive a richer set of behavioral implications, we turn to a specific utility function and derive comparative static results. Assuming constant

[2] The $\text{cov}[U'(\pi), \tilde{P}]$ is negative because an increase in price leads to an increase in π, which in turn results in a decrease in $U'(\pi)$ because of the concavity of U. Thus, price and $U'(\pi)$ move in opposite directions, and hence their covariance is negative. A similar argument applies to the covariance between $U'(\pi)$ and the off-farm wage.

absolute risk aversion (CARA) and a joint normal probability density function for $\tilde{P}$ and $\tilde{W}$, the farm operator wants to maximize the following:

$$(6)\quad EU(\pi) = E(-e^{-\alpha\pi}), \quad \text{or}$$
$$EU(\pi) = \int_{-\infty}^{\infty} -e^{-\alpha[\tilde{P}Q(F) - C(Q) + I + \tilde{W}(T-F)]}\, k e^{-(\pi-\mu_\pi)/(2\sigma_\pi^2)}\, dP,$$

where μ_π and σ_π^2 are the mean and variance of net farm income, and $k = 1/\sigma_\pi\sqrt{2\pi}$ (note that π in this last term is the mathematical constant, not the farm operator's net income).[3]

It is well known that maximizing (6) is equivalent to maximizing $\Phi = E(\pi) - (\alpha/2)\sigma_\pi^2$, where α is the Arrow-Pratt degree of absolute risk aversion:

$$(7)\quad E(\pi) = E(\tilde{P})Q - C(Q) + I + E(\tilde{W})(T-F),$$

$$(8)\quad \sigma_\pi^2 = Q^2\sigma_P^2 + (T-F)^2\sigma_W^2 + 2Q(T-F)\sigma_{PW}.$$

In equation (8), σ_P^2 and σ_W^2 are the variances of P and W, and σ_{PW} is the covariance between P and W. The first-order condition for this problem is specified as:

$$(9)\quad \frac{\partial\Phi}{\partial F} = E(\tilde{P})Q_F - C'Q_F - E(\tilde{W}) - \alpha\big[QQ_F\sigma_P^2 - (T-F)\sigma_W^2 + (Q_F(T-F) - Q)\sigma_{PW}\big] = 0.$$

If P (output price) and W (off-farm wage) are unrelated (or nearly so), their covariance (σ_{PW}) will be zero (or close to it).[4] In this case, we can show:

$$\frac{\partial F}{\partial\sigma_P^2} < 0 \quad \text{and} \quad \frac{\partial F}{\partial\sigma_W^2} > 0.$$

Thus, the amount of time devoted to farm work decreases when farm price (and income) is more variable, and it increases when the off-farm wage is more variable. This result is what we would expect. When one source of income becomes more risky, a risk-averse farmer reallocates work time away from that income source.

To prove the first of these results, differentiate (9) and solve for $\partial F/\partial\sigma_P^2$, which gives:

$$\frac{\partial F}{\partial\sigma_P^2} = -\frac{\partial^2\Phi/\partial F\partial\sigma_P^2}{\partial^2\Phi/\partial F^2}.$$

Since $\partial^2\Phi/\partial F^2 < 0$ by the second-order condition for a maximum,[5] the sign of $\partial F/\partial\sigma_P^2$ is the same as the sign of $\partial^2\Phi/\partial F\partial\sigma_P^2$. Since $\partial^2\Phi/\partial F\partial\sigma_P^2 = -\alpha QQ_F < 0$, it follows that

$$(10)\quad \frac{\partial F}{\partial\sigma_P^2} < 0.$$

Similarly, the sign of $\partial F/\partial\sigma_W^2$ is the same as the sign of $\partial^2\Phi/\partial F\partial\sigma_W^2$. Since $\partial^2\Phi/\partial F\partial\sigma_W^2 = \alpha(T-F) > 0$,

$$(11)\quad \frac{\partial F}{\partial\sigma_W^2} > 0.$$

As shown by equation (10), an increase in the variability of farm output price (and therefore farm income) results in a decrease in the amount of time spent working on the farm by the farm operator.[6] Similarly, equation (11) implies an increase in off-farm wage variability results in an increase in the amount of time spent working on the farm.

In addition to variability, the farm operator also cares about the levels of expected price and expected wages. The effect of a change in expected output price is represented by:

$$\frac{\partial F}{\partial E(P)} = -\frac{\partial^2\Phi/\partial F\partial E(P)}{\partial^2\Phi/\partial F^2}.$$

Since the denominator is negative, $\partial F/\partial E(P)$ has the same sign as the numerator. The numerator equals Q_F, so $\partial F/\partial E(P) > 0$. The farm operator spends more time on farm work when the expected price of farm output increases. Similarly, it is relatively easy to show $\partial F/\partial E(W) < 0$. Thus the farm operator spends less time working on the farm when the expected wage rate increases.

Other results are not as clear-cut. For example, it would be interesting to know how the amount of farm work depends on the degree of risk aversion. To examine this issue, we must determine the sign of $\partial F/\partial\alpha$, which will have the same sign as $\partial^2\Phi/\partial F\partial\alpha$. Taking this cross-derivative yields:

---

[3] As noted in footnote 1, we can derive the same qualitative results if the utility function is separable in income and leisure. In particular, if $U(\pi, L) = -e^{-\alpha\pi}e^{-\gamma L}$, the main results in this part of the article can be derived without holding leisure constant.

[4] Analysis of data from Kansas and North Carolina shows σ_{PW} is essentially zero.

[5] If the production function is concave in F, and the cost function is convex in Q, then the net income function, π, will be concave. Since the utility function is concave due to risk aversion, $U(\pi)$ will be concave. Thus, $EU(\pi)$ has a unique maximum at which the first derivative with respect to F is zero and the second derivative with respect to F is negative.

[6] Leisure is implicitly being held constant because we have assumed it is fixed in this model. However, as mentioned earlier, this result holds even if leisure may vary, as long as utility is separable in leisure and net income.

$$\frac{\partial^2 \Phi}{\partial F \partial \alpha} = -\left[QQ_F \sigma_P^2 - (T-F)\sigma_W^2\right],$$

which will be negative if

$$\frac{\sigma_P^2}{\sigma_W^2} > \frac{(T-F)}{QQ_F}.$$

Thus, if the variance of price is large relative to the variance of wages,[7] the more risk averse the farm operator is, the fewer hours the operator will work on the farm. We would expect the variability of farm income to be much greater than the variability of wages in most cases. Consequently, we would expect an increase in risk aversion would generally lead to an increase in hours worked off the farm.

A change in nonearned income, I, has no effect on the allocation of work hours in this CARA model, because the degree of risk aversion is independent of the amount of net income. However, the result in the previous paragraph suggests a farm operator with a decreasingly absolute risk averse (DARA) utility function would work more hours on the farm with an increase in nonearned income. We observe this result because the increase in nonearned income increases net income which, in turn, reduces the farm operator's degree of risk aversion. The decrease in risk aversion then leads the farmer to work more hours on the farm as long as the farm income risk is large relative to the wage risk, as established above.

Empirical Model

Strict concavity of the farm production function and the operators' utility function, along with a linear time constraint, guarantees the existence and uniqueness of a solution to the maximization problem in the previous section. Thus, it is legitimate to write the optimal hours worked on or off the farm as a function of the parameters of the problem.

Expanding on the parameter set used in the theoretical model, we write a reduced-form equation for off-farm labor supply as a function of several observable factors as follows:

$$(12)\quad H = f\left(W, W_F, E_o, T_o, I, F_C, O_C, \sigma_P^2, \sigma_W^2\right),$$

where W is the expected off-farm wage, W_F is the wage paid to hired farm workers, E_o represents the number of years the farm operator has been on the farm, T_o represents the tenure of the farm operator (full owner, part owner, and tenant), I is nonearned income (which we assume to be exogenous to the operator's consumption and leisure decisions), F_C and O_C represent farm and operator characteristics, and σ_P^2 and σ_W^2 are the variances of farm income and off-farm wages, respectively.

Ideally, we would like to use individual farm-level data to estimate (12). However, farm-level data which include information on variability of farm income and off-farm wages do not exist, so we estimate the empirical model using county-level data from the *Census of Agriculture* (USDA; U.S. Department of Commerce).[8] Therefore, each county is a unit of observation in our data set, and we use the number of farm operators reporting off-farm work[9] to measure participation in the off-farm labor market.

Specifically, the dependent variable, H, is measured by the participation rate, R_j, defined as:

$$(13)\quad R_j = \frac{N_j^O}{N_j},$$

where $N_j = N_j^O + N_j^F$, and N_j^O is the number of operators in county j reporting any off-farm workdays, N_j^F is the number of farm operators in county j reporting no off-farm workdays, and N_j is the total number of farm operators in county j.

The empirical model in (12) must therefore be recast as a logit model with the off-farm work participation rate as the dependent variable and a vector of explanatory variables, $\mathbf{X}_j$. The empirical model is then expressed as:

$$(14)\quad R_j = \frac{1}{\left[1 + e^{-(\beta_0 + \boldsymbol{\beta}_1 \ln(\mathbf{X}_j) + \varepsilon_j)}\right]},$$

where β_0 and the vector $\boldsymbol{\beta}_1$ are coefficients to be estimated, and ε_j is a random disturbance term. The value of R_j lies between 0 and 1, and thus violates the standard Generalized Linear Hypothesis assumption about the error term, which allows one to use ordinary least squares estimation (Fomby, Hill, and Johnson, 1984).

[7] Large in this context means at least as large as $(T - F)/QQ_F$.

[8] Huffman (1973a, b, 1977b, 1980), in his seminal work on off-farm labor supply of farm operators, used county-level data from three different states (Iowa, Oklahoma, and North Carolina). He acknowledged some aggregation bias might exist, but Sexton (1975) showed this bias may not be very significant.

[9] The *Census of Agriculture* classifies an operator as one participating in off-farm work if the operator spends at least four hours a day working off the farm.

Following Huffman (1980), and Slottje, Hayes, and Shackett (1992), we linearized the logistic specification[10] in equation (14) using a logarithmic transformation (see the appendix for details). This transformation gives:

$$(15)\quad \ln\left(\frac{R_j}{1-R_j}\right) \equiv \left[\ln(R_j) - \ln(1-R_j)\right] = \beta_0 + \boldsymbol{\beta}_1 \ln(\mathbf{X}_j) + \varepsilon_j .$$

Following Zellner and Lee (1965), assume the disturbances in (15) follow an independent distribution with mean zero and variance $1/[N_j R_j(1-R_j)]$. Then equation (15) may be estimated using weighted least squares.[11]

Interpretation of equation (15) is fairly straightforward. The left-hand side represents the difference between the natural logarithm of off-farm work participation and the natural logarithm of not participating in off-farm work. The difference reflects the amount by which operators' off-farm labor supply is favored in county j. Each coefficient in the $\boldsymbol{\beta}_1$ vector is the percentage change in the odds ratio (N_j^O/N_j^F) due to a 1% change in the corresponding explanatory variable in $\mathbf{X}_j$.[12]

Data Description

The data consist of 205 county-level observations in Kansas (105 counties) and North Carolina (100 counties). These states, in which agriculture continues to be an important industry, are quite different in climate, geography, and opportunities for off-farm work. In Kansas an average farm is 748 acres and typically specializes in production of wheat, sorghum grains, and beef cattle.

Kansas farms receive considerably larger government payments than North Carolina farms. Off-farm employment opportunities are limited and, as a result, fewer farm operators work off the farm in Kansas compared to North Carolina farm operators. North Carolina farms are smaller (185 acres on average) compared to their Kansas counterparts, and generally produce tobacco, turkeys, poultry, hogs, cotton, and peanuts.

The average farm in North Carolina receives more in rental income than through government payments. Because of North Carolina's geographic setting and more diversified and extensive manufacturing base, more farm operators work off the farm and for higher wages compared to Kansas farmers.

The primary data source for this analysis is the *Census of Agriculture* (USDA, 1997), which provides a rich source of data for constructing empirical measures of the variables used to estimate the supply of off-farm work by farm operators in Kansas and North Carolina.

The independent variables used in this study are as follows. *AGE* is the average age of farm operators in each county. This variable controls for possible differences in work time allocation due to human capital vintage. Age squared (*AGESQ*) is included to capture the possibility that human capital depreciates after some age. *OTHINC* denotes other income per farm.[13] *VALPROD* is the value of agricultural product sold per farm[14] in each county. This variable is a proxy for farm size, which we found to affect off-farm labor supply in earlier unpublished empirical work.

Our theoretical model does not include farm size, but in preliminary versions (not reported here), the theoretical effect of farm size on the labor allocation decision was ambiguous. The reason for this finding is that as farm size increases, the exposure to income from farming increases, but increasing farm size also increases the expected marginal product for on-farm work. These circumstances have opposite effects on the desired allocation of labor, and hence the effect of farm size is ambiguous.

Farm ownership (tenure of operator) influences decisions regarding off-farm work participation because operators in different tenure categories are likely to have different objectives and face different economic and resource constraints. For example, a part-owner operator is more likely to work off the farm because off-farm income can provide necessary capital for gaining full ownership.

In this study, farm tenure categories include full-owner operator (*FOWNER*), part-owner operator

---

[10] El-Osta, Bernat, and Ahearn (1995) used a similar transformation of the Gini coefficient to investigate the role of off-farm income in income inequality.

[11] Huffman (1980) used a similar model to estimate off-farm labor supply of farm operators in Iowa, Oklahoma, and North Carolina.

[12] Note that $\ln(N_j^O/N_j^F)$ is equivalent to the left-hand side of equation (15). It is the logarithm of the ratio of the probability of participating to not participating in off-farm work.

[13] Other income includes income from custom work, rental income, government payments, and other farm-related income (income from hunting leases, fishing fees, and other recreational activities, sales of farm by-products, and other businesses or income closely related to agricultural operations).

[14] The *Census of Agriculture* reports value of agricultural products removed from the farm. Thus it may include commodities produced under contracts.

Table 3. Definitions and Summary Statistics for Variables in the Empirical Model: Kansas and North Carolina Farm Operators, 1997

		KANSAS		NORTH CAROLINA	
Variable	Definition	Mean	Std. Dev.	Mean	Std. Dev.
AGE	Average age of farm operators in the county (years)	53.14	1.86	54.54	2.23
EXPER	Average number of years spent on the present farm	22.31	1.76	20.75	1.73
POWNER	Proportion of farms operated by part-owner operators	0.44	0.07	0.34	0.08
TENANT	Proportion of farms operated by tenants	0.16	0.06	0.14	0.014
VALPROD	Value of agricultural products sold per farm in the county ($000s)	181.30	252.58	95.01	79.74
OTHINC	Other income per farm in the county ($000s)	16.29	9.76	10.57	8.29
CVINCM	Coefficient of variation in net farm income per farm in the county	1.66	8.63	0.64	0.24
CVWAGE	Coefficient of variation in net nonfarm weekly wages in the county	0.14	0.02	0.05	0.02
FINCM	Net farm income per farm in the county, 1996 ($000s)	22.26	12.93	11.36	12.88
OFWAGE	Nonfarm average hourly wages in the county, 1996 ($/hour)	9.59	2.78	10.46	1.66
HIREEXP	Expenditures on hired farm labor per farm in the county ($000s)	12.83	12.93	19.46	15.54
No. of Observations		105		100	

Source: *Census of Agriculture*, USDA (1997).

(*POWNER*), and tenant (*TENANT*). The variables are expressed as proportions. For example, *FOWNER* is defined as the ratio of the number of full-owner operators to the total number of operators, which implies (*FOWNER* + *POWNER* + *TENANT*) = 1. Therefore, to avoid multicollinearity, one category (*FOWNER*) was dropped from the estimation procedure.

Net farm income per farm (*FINCM*) and nonfarm average hourly wage (*OFWAGE*) for each county were collected from the Regional Economic Information System (REIS) (1969–96), maintained by the Bureau of Economic Analysis, U.S. Department of Commerce. Farm income variability (*CVINCM*) and off-farm wage variability (*CVWAGE*), measured as coefficients of variation, were calculated from these data.

HIREEXP represents expenditures on hired farm labor per farm. Our original thought was that *HIREEXP* would be a proxy for cost of production, and we assumed higher values would result in less profitable farming, thus reducing the number of hours the farmer worked on the farm. However, there is a countervailing effect, because a farmer might substitute his/her own labor for hired labor as the cost of hired labor increases. Thus, we cannot say a priori how this variable may affect the farmer's labor allocation decision.

Finally, *EXPER* represents the average number of years spent by the farm operator on the present farm. Economists have evaluated the effects of farming experience on the off-farm labor supply of farm operators and have determined that farming experience affects the off-farm labor supply indirectly through farm production and directly through the labor supply function. Furtan, van Kooten, and Thompson (1985) estimated the direct impact of years of farming on the off-farm labor supply and found it to be negative and statistically significant. Additionally, Sumner (1982) observed that farm operators with "some farm training" supply fewer hours of off-farm work, although the estimated impact is not significant. Table 3 provides summary statistics for all the variables used in this study.

Results

The dependent variable for the participation model is the natural logarithm of the odds of participating in the off-farm labor market, and the independent variables are also expressed in natural logarithms. Therefore, equation (15) is estimated by the weighted least squares procedure, where the observed variables were multiplied by $\sqrt{N_j R_j (1 \& R_j)}$. Thus, counties with large N_j and with R_j equal to 1 ! R_j receive the greatest weight (Huffman, 1980).

Table 4. Parameter Estimates of County-Level Off-Farm Work by Farm Operators in Kansas and North Carolina, 1997

	KANSAS	NORTH CAROLINA
Variable	Parameter Est. (Std. Dev.)	Parameter Est. (Std. Dev.)
Intercept	!329.319 (252.440)	!216.307 (105.968)
AGE	167.059** (84.930)	109.441** (53.433)
AGESQ	!21.057* (11.440)	!13.739** (6.724)
EXPER	!0.726* (0.431)	!0.021** (0.010)
POWNER	0.250 (0.284)	0.097 (0.082)
TENANT	!0.129** (0.061)	!0.173*** (0.037)
VALPROD	0.059 (0.090)	!0.061** (0.031)
OTHINC	!0.241** (0.121)	!0.136*** (0.045)
CVINCM	0.072* (0.041)	0.062** (0.031)
CVWAGE	!0.423** (0.201)	!0.096** (0.050)
FINCM	!0.027 (0.041)	0.001 (0.022)
OFWAGE	0.197** (0.091)	0.151* (0.081)
HIREEXP	0.063 (0.097)	!0.066* (0.036)
R^2 (adjusted R^2)	0.64 (0.62)	0.74 (0.70)
No. Observations	105	100

*, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively.

Parameter estimates for the Kansas and North Carolina participation models are reported in table 4. All parameters have signs consistent with the theoretical model. Since the variables are in natural logarithms, the estimated coefficients represent elasticities of the odds of off-farm work participation by farm operators, where the "odds" is a ratio of the probability of participating to not participating in off-farm work. The model explains 64% and 74% of variation in off-farm work participation by farm operators in Kansas and North Carolina, respectively.

The coefficient of age of farm operator (*AGE*) is positive and statistically significant at the 5% level for both Kansas and North Carolina. One possible explanation for this finding is that the present value of the returns to investing in the search for off-farm jobs will be greater for young farm operators, since the pay-off period for such activities is much longer. Conversely, the older the age at which a farm operator considers engaging in off-farm work activities, the shorter will be the recoupment period and the greater the costs of the job search relative to the benefits. The tendency for the returns from a job search to decline relative to the costs as age increases will therefore induce older farm operators to devote more of their time to nonmarket or household activities and less to off-farm work.

Yet another possible explanation is that productivity tends to rise with age, at least until depreciation significantly offsets accumulated experience. This phenomenon is evident from the coefficient of age squared (*AGESQ*). In both the Kansas and North Carolina cases, the sign of the coefficient is negative, and is statistically significant at 10% and 5%, respectively. Further, farm operators' participation in off-farm activities shows a definite inverted *U* shape—participation in off-farm work increases as age increases up to age 54 in the case of Kansas, and 49 in the case of North Carolina, and thereafter declines as age increases above those levels. These results are consistent with the findings of Sumner (1982) and Huffman (1973a).

Farm size, as measured by the value of agricultural product sold per farm (*VALPROD*), has a negative and statistically significant coefficient in the case of North Carolina but a positive and insignificant coefficient in the case of Kansas. Since the theoretical model's prediction is ambiguous in this case, the empirical results are of some interest. In particular, the empirical model suggests a 10% increase in the value of farm products sold decreases the odds of off-farm work participation by 0.6% in North Carolina. This result is consistent with the findings of Sumner (1982), Lass and Gempesaw (1992),[15] and Mishra and Goodwin (1997).

Among other factors, farm tenure influences the decision to participate in off-farm work. Tavernier, Temel, and Li (1997) point out the importance of tenure structure on land conversion and suggest implications for off-farm work. Ownership participation elasticities indicate farm tenure affects the decision to participate in off-farm work. The coefficient of *TENANT*, the ratio of number of tenants to the total number of farm operators, is negative and statistically significant at the 5% level in Kansas

[15] Sumner (1982), and Lass and Gempesaw (1992) used total acres as an indicator of farm size.

and the 1% level in North Carolina. In contrast, the coefficient of *POWNER*, the ratio of the number of part-owners to the total number of farm operators, is insignificant in both cases.

Specifically, our findings suggest a 10% increase in farms operated by tenants decreases the odds of participation in off-farm work by approximately 1.3% in Kansas and by 1.7% in North Carolina, but a similar increase in part-owners has no effect on off-farm work. The decreased tendency on the part of tenants to participate in off-farm employment may be because they have already committed themselves to farming by renting farmland. These results are consistent with the findings of Kimhi (1994); Kilkenny (1993); Tavernier, Temel, and Li (1997); Mishra and Goodwin (1997); and Mishra (1996).

Other income (*OTHINC*), which includes government payments, rental income, custom work, and other farm-related income per farm, is expected to have a negative effect on off-farm work participation under specific conditions, as detailed earlier. The estimated coefficients for *OTHINC* are negative and statistically significant for both Kansas and North Carolina. In Kansas, a 10% increase in other income reduces the odds of off-farm work participation of farm operators by approximately 2.4%, with a comparable estimate for North Carolina operators of approximately 1.4%. This finding suggests that, in part, other income substitutes for off-farm work.

These results are supported by the findings of Tavernier, Temel, and Li (1997), and Mishra and Goodwin (1997). The fact that the elasticity is lower in North Carolina may be due to the more diversified economic base in North Carolina. Thus, fine adjustments in off-farm work are easier to accomplish in North Carolina than in Kansas.

The theoretical model predicts that the riskiness of farm income and off-farm work are positively related, and this prediction is supported by the empirical results. Farm income variability as represented by the coefficient of variation in net farm income (*CVINCM*) has a positive effect on off-farm work participation by farm operators in Kansas and North Carolina. As seen from table 4, a 10% increase in the coefficient of variation in farm income (*CVINCM*) increases the odds of off-farm work participation by 0.7% among farm operators in Kansas and by 0.6% for operators in North Carolina. This result is consistent with the findings of Mishra and Goodwin (1997). Again, the greater economic base in North Carolina may account for the difference in elasticities.

The theoretical model also predicts a negative relationship between the riskiness of off-farm wages and off-farm work. The riskiness in off-farm income, as represented by the coefficient of variation in wages (*CVWAGE*), has a negative and statistically significant effect on off-farm participation by farm operators in Kansas and North Carolina. A 10% increase in the coefficient of variation in off-farm wages (*CVWAGE*) decreases the odds of off-farm work participation by 4.2% for farm operators in Kansas and by about 1% among North Carolina farm operators.

The off-farm wage rate is also significant in explaining variation in off-farm work participation by operators in Kansas and North Carolina. The coefficient of the off-farm wage rate (*OFWAGE*) in both states is positive and statistically significant. This result gives us an upward-sloping off-farm labor supply curve of farm operators in both states. Results show if the nonfarm wage rate increases by 10%, the odds of off-farm work participation by operators would increase by approximately 2% in Kansas and 1.5% in North Carolina. This is consistent with findings reported in Huffman's (1973a) study using county-level data.[16]

Finally, the coefficient of expenditures on hired farm labor (*HIREEXP*) is negative and statistically significant in North Carolina but essentially zero and insignificant in Kansas. Specifically, results indicate a 10% increase in hired labor expense reduces the odds of off-farm work participation by approximately 0.7% in North Carolina, suggesting North Carolina farm operators may substitute their labor for paid farm labor when the cost of paid labor increases. It seems plausible this effect might be more important in North Carolina than in Kansas since North Carolina farms are smaller, and thus an operator's on-farm labor could substitute for a significant fraction of the work performed by hired laborers.

Farming experience is proxied by the variable *EXPER*, defined as the average number of years an operator has been operating the present farm. From table 4, the coefficient of *EXPER* is negative and statistically significant for both Kansas and North Carolina farm operators. However, the magnitude of the coefficient is larger in the case of Kansas. A coefficient of 0.726 in the Kansas equation implies that a 10% increase in the number of years of on-farm experience reduces the odds of off-farm

[16] Counties from Iowa, Oklahoma, and North Carolina were used in Huffman's study, with a combined total of 276 counties.

participation of farm operators in Kansas by approximately 7.3%. On the other hand, a 10% increase in the number of years of on-farm experience reduces the odds of off-farm participation of farm operators in North Carolina by only 2.1%. Thus, as farming experience increases, the likelihood a farm operator will participate in off-farm work decreases. Our results are consistent with the findings of Sumner (1982); Huffman (1977a, b, 1980); Mishra (1996); Mishra and Goodwin (1997); Tavernier, Temel, and Li (1997); and Furtan, van Kooten, and Thompson (1985).

A possible explanation for the negative relationship between farming experience and off-farm work is that human capital in the form of on-the-job experience raises earnings from time spent farming. Past experience determines the earnings in the current period, while experience gained in the current period raises future income to the extent that a person continues to work and therefore uses his or her capital. Both the effect from accumulated experience and the expectation of future work are important for explaining life cycle labor supply patterns of farmers.

Summary and Conclusions

A novel feature of this study is that we investigate the riskiness of off-farm wages as well as the riskiness of farm income on off-farm employment decisions made by farm operators. We employ a simple theoretical model to develop hypotheses, which are tested empirically. The empirical framework involves the estimation of a participation rate model. The estimation is performed by heteroskedasticity-consistent weighted least squares using per farm averages for counties in Kansas and North Carolina.

In summary, variability in both farm income and off-farm wages is an important determinant of off-farm employment decisions by farm operators. Our results show there is a positive and significant relationship between farm income variability, representing riskiness in farming, and off-farm work participation by farm operators in North Carolina and Kansas. Likewise, variability in off-farm wages, representing the inherent riskiness in the nonfarm sector, exhibits a negative and significant correlation with off-farm work decisions as predicted by the theoretical model.

In addition, the empirical evidence shows the following: (*a*) participation in off-farm work increases with age up to a point, but thereafter decreases with age; (*b*) experience in farming and participation in off-farm work are inversely related; (*c*) the level of other income reduces the likelihood that farm operators will participate in off-farm work; (*d*) as farm size increases, working off the farm decreases in North Carolina; (*e*) the level of the off-farm wage has a positive impact on operator participation in off-farm work; and (*f*) operator work on the farm is a substitute for hired labor in North Carolina.

These results reinforce findings of other studies where a link has been observed between farm income variability and farm operators' decisions regarding on- and off-farm work. It seems clear that farm income variability should be taken into consideration when formulating public policy for agriculture.

The results suggest policies such as agricultural price supports have a two-pronged effect on farm operators' labor allocation decisions. Price supports generally raise the average price of the target commodity as well as reduce the variability of price. Both effects will lead farm operators to work more on the farm and less off the farm. Whether this is a beneficial or detrimental effect of price support programs is beyond the scope of this investigation, but it is an effect which should be recognized by policy makers.

What is unique in this study is the finding that farming decisions are also closely linked to economic conditions in the nonfarm sector. Since many farm operators combine farming with nonfarm work, the level and variability of nonfarm wages are important. For example, to the extent that minimum wage laws increase average nonfarm wages and reduce their variability, they will induce farm operators to work more off the farm.

Also, during economic downturns when nonfarm wages and off-farm opportunities decrease or during periods of high nonfarm wage variability, we would expect farm operators to spend more work time on the farm. This would most likely increase farm production, resulting in a rightward shift of the market supply curve and causing commodity prices to decline. If an economic downturn also led to reduced demand, commodity prices could decline even further, dramatically reducing farm incomes and ultimately leading some farm operators to exit farming.

While some of these concluding remarks are a bit speculative, they do affirm the potential importance of the relationships examined in this study.

References

Ahearn, M. (1986). "Financial Well-Being of Farm Operators and Their Households." AE Report No. 563, Economic Research Service/U.S. Department of Agriculture, Washington, DC.

Ahearn, M., and J. E. Lee. (1991). "Multiple Job-Holding Among Farm Operator Households in the U.S." In M. C. Hallberg, J. L. Findeis, and D. A. Lass (eds.), *Multiple Job-Holding Among Farm Families* (pp. 3–30). Ames, IA: Iowa State University Press.

Ahearn, M., J. E. Perry, and H. S. El-Osta. (1993). "Economic Well-Being of Farm Operator Households, 1988–90." AE Report No. 666, Economic Research Service/U.S. Department of Agriculture, Washington DC.

Barlett, P. F. (1991). "Motivations for Part-Time Farmers." In M. C. Hallberg, J. L. Findeis, and D. A. Lass (eds.), *Multiple Job-Holding Among Farm Families* (pp. 45–70). Ames, IA: Iowa State University Press.

El-Osta, H. S., G. A. Bernat, and M. C. Ahearn. (1995). "Regional Differences in the Contribution of Off-Farm Work to Income Inequality." *Agricultural and Resource Economics Review* 24(1), 1–14.

Fomby, T., R. C. Hill, and S. R. Johnson. (1984). *Advanced Econometrics Methods.* New York: Springer-Verlag.

Fuller, A. (1991). "Multiple Job-Holding Among Farm Families in Canada." In M. C. Hallberg, J. L. Findeis, and D. A. Lass (eds.), *Multiple Job-Holding Among Farm Families* (pp. 31–44). Ames, IA: Iowa State University Press.

Furtan, W. H., G. C. van Kooten., and S. J. Thompson. (1985). "The Estimation of Off-Farm Supply Functions in Saskatchewan." *Journal of Agricultural Economics* 36(2), 211–220.

Gunter, L., and K. T. McNamara. (1990). "The Impact of Local Labor Market Conditions on the Off-Farm Earnings of Farm Operators." *Southern Journal of Agricultural Economics* 22(2), 155–165.

Hallberg, M. C., J. L. Findeis, and D. A. Lass, eds. (1991). *Multiple Job-Holding Among Farm Families*. Ames, IA: Iowa State University Press.

Huffman, W. E. (1973a). "A Cross-Sectional Analysis of Nonfarm Work of Farm Operators." Working Paper No. 23, College of Business Administration, Oklahoma State University, Stillwater.

———. (1973b). "Household Behavior: Hours of Work." Working Paper No. 22, College of Business Administration, Oklahoma State University, Stillwater.

———. (1977a). "Allocative Efficiency: The Role of Human Capital." *Quarterly Journal of Economics* 91(1), 59–79.

———. (1977b). "Interactions Between Farm and Nonfarm Labor Markets." *American Journal of Agricultural Economics* 59(5), 1054–1061.

———. (1980). "Farm and Off-Farm Work Decisions: The Role of Human Capital." *Review of Economics and Statistics* 62(1), 14–23.

Kilkenny, M. (1993). "Rural/Urban Effects of Terminating Farm Subsidies." *American Journal of Agricultural Economics* 75(4), 968–980.

Kimhi, A. (1994). "Participation of Farm Owners in Farm and Off-Farm Work Including the Option of Full-Time Off-Farm Work." *Journal of Agricultural Economics* 45(2), 232–239.

Lass, D. J., and C. M. Gempesaw. (1992). "The Supply of Off-Farm Labor: A Random Coefficient Approach." *American Journal of Agricultural Economics* 74(4), 400–411.

Mishra, A. K. (1996). "Farm Income Variability and the Off-Farm Labor Supply of Farmers and Their Spouses." Unpublished Ph.D. dissertation, Department of Economics, North Carolina State University, Raleigh.

Mishra, A. K., and B. K. Goodwin. (1997). "Farm Income Variability and the Off-Farm Labor Supply." *American Journal of Agricultural Economics* 79(3), 880–887.

Perry, J. E., and R. Hoppe. (1993, November). "Off–Farm Income Plays Pivotal Role." *Agricultural Outlook*, pp. 3–5. Economic Research Service/U.S. Department of Agriculture, Washington, DC.

Sexton, R. N. (1975). "Determinants of Multiple Job-Holding by Farm Operators." Unpublished Ph.D. dissertation, Department of Economics and Business, North Carolina State University, Raleigh.

Slottje, D. J., K. J. Hayes, and J. Shackett. (1992). "Labor Force Participation, Race, and Human Capital: Influence on Earnings Distributions Across States." *Review of Income and Wealth* 38(1), 27–37.

Sommer J. E., D. E. Banker, R. C. Green, J. Z. Kalbacher, N. R. Peterson, and T. Y. Sun. (1997). *Structural and Financial Characteristics of U.S. Farms, 1994.* Nineteenth Annual Family Farm Report to the U.S. Congress. Pub. No. AIB 735, Economic Research Service/U.S. Department of Agriculture, Washington, DC.

Spitze, G. F. R., and R. K. Mahoney. (1991). "Evolving Dimensions of Dual Employment of Illinois Farm Families." In M. C. Hallberg, J. L. Findeis, and D. A. Lass (eds.), *Multiple Job-Holding Among Farm Families* (pp. 163–180). Ames, IA: Iowa State University Press.

Sumner, D. A. (1982). "The Off-Farm Labor Supply of Farmers." *American Journal of Agricultural Economics* 64(4), 499–509.

Tavernier, E. M., T. T. Temel, and F. Li. (1997). "The Role of Farm Ownership in Off-Farm Work Participation." *Agricultural and Resource Economics Review* 26(1), 67–81.

U.S. Department of Agriculture. (1988–1998). *Agricultural Outlook.* Economic Research Service/USDA, Washington, DC. Selected issues.

———. (1997). "State and County Data." *Census of Agriculture.* USDA, Washington, DC.

U.S. Department of Commerce, Bureau of the Census. (1969–1996). "State and County Data." *Census of Agriculture.* Washington, DC. Various years.

U.S. Department of Commerce, Bureau of Economic Analysis. (1969–1996). Regional Economic Information System (REIS) data. Regional Economic Measurement Division, Washington, DC.

Zellner, A., and H. T. Lee. (1965). "Joint Estimation of Relationship Involving Discrete Random Variables." *Econometrica* 33, 382–394.

Appendix: Logarithmic Transformation of the Empirical Model, Text Equation (14)

For the logarithmic transformation of the empirical model, we take text equation (14), now denoted appendix equation (A1), as follows:

$$\text{(A1)} \qquad R_j = \frac{1}{\left[1 + e^{-(\beta_0 + \boldsymbol{\beta}_1 \ln(\mathbf{X}_j) + \varepsilon_j)}\right]}.$$

For simplicity, let us denote $\beta_0 + \boldsymbol{\beta}_1 \ln(\mathbf{X}_j) + \varepsilon_j$ as Z_j, and equation (A1) can be rewritten as:

$$\text{(A1.1)} \qquad R_j = \frac{1}{\left[1 + e^{-(z_j)}\right]}.$$

Multiplying both sides of equation (A1.1) by

$$\frac{1}{\left[1 + e^{-(z_j)}\right]}$$

and rearranging the terms results in the following:

$$\text{(A1.2)} \qquad 1 + e^{-(z_j)} = \frac{1}{R_j}.$$

Rearranging the terms in equation (A1.2) yields:

$$\text{(A1.3)} \qquad e^{-(z_j)} = \frac{1 - R_j}{R_j} \quad \text{and} \quad \frac{1}{e^{(z_j)}} = \frac{1 - R_j}{R_j}.$$

Inverting equation (A1.3) and taking natural logarithms on both sides results in the following:

$$\text{(A1.4)} \qquad z_j = \ln\left[\frac{R_j}{1 - R_j}\right].$$

Rearranging (A1.4) results in text equation (15):

$$\text{(A1.5)} \qquad \ln\left(\frac{R_j}{1 - R_j}\right) \equiv \left[\ln(R_j) - \ln(1 - R_j)\right] = \beta_0 + \boldsymbol{\beta}_1 \ln(\mathbf{X}_j) + \varepsilon_j.$$