Off-farm Employment and Farming Efficiency in Modern Agriculture: A Dynamic Panel Analysis

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Abstract

Most of the empirical literature in this area tends to analyze labor allocation decisions of economic agents using cross-sectional data. But such methods implicitly assume that model parameters are stable (constant) across firms and over time. The use of cross-sectional methods is therefore glaringly at odds with the firm-specific aspects of the theoretical models employed in labor economics. Using a large panel data this study investigates the simultaneous relationship between farming efficiency and the off-farm labor supply decisions of both farm operators and their spouses. We also account for unobserved heterogeneity and correcting for simultaneity bias in such estimation. Results reveal several interesting findings. First, farming efficiency (ratio of farm revenue to total variable cost) has a positive and negative impact on hours of off-farm work by farm operators and spouses, respectively. Second, agricultural subsidy has a negative and positive effect on off-farm work hours of farm operators and spouses, respectively. Finally, we find a dynamic relationship between off-farm labor supply and farming efficiency. Specifically, in the case of the farm operator, off-farm work first increases farming efficiency in the first period and then decreases it in the second period. On the other hand, we observe a positive correlation, for both periods, between off-farm hours worked by spouses and farming efficiency.

Keywords: Panel, heterogeneity, simultaneity, off-farm labor, farm households, agricultural subsidy

JEL Codes: D13, J18, J22, Q12, Q18

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Introduction

During the past two decades considerable research has focused on agricultural policy and its unintended consequences on factor markets, including labor. Notably among them is the labor allocation decision of farm families. off-farm activities have provided a critical income source to a majority of farm households not only in the United States and Western European countries, but also in developing countries (e.g., Mishra et al. 2002; Ahearn, El-Osta & Dewbre, 2006; Benjamin and Kimhi, 2006; Glauben et al. 2008). Off-farm provision has been largely responsible for several documented improvements in the standard of living: (1) closing the income gap between farm and nonfarm households (Mishra et al. 2002; Holden et al. 2004; Gardner, 2002); (2) manage risk through diversification of income (Barrett et al. 2001; Mishra and Goodwin, 1997; Schultz, 1990); (3) food consumption and nutrition (Chang and Mishra, 2008; Babatunde and Qaim, 2010); and (4) farm input usage (Phimister and Roberts, 2006; Chang, Mishra & Livingston, 2011). One can easily conclude that off-farm income can contribute to significant improvements in the welfare of agricultural households (Hill, 2000).

A plethora of literature has evolved that investigates the determinants of farm household involvement in off-farm labor markets. However, with increased reliance on off-farm income may have resulted in less time devoted to farming related human capital. In particular, Smith (2002) argues that increased reliance on off-farm employment may have resulted in a reduction in on-farm efficiency. For example, Smith argues that increased off-farm work may have implied less attention to issues "smart farming"—which could lead to a reduction in variable costs.[2] While Goodwin and Mishra (2004), the only study, have shown that greater involvement in off-farm work creases farming efficiency (defined as the ratio of gross cash farm income to total variable costs), their study has several weaknesses. First, the study ignores the role of spouse in the joint determination of labor supply decision by farm operators and spouses. It is not surprising since most existing analyses of off-farm labor supply in Western economies usually include the decisions of the farm operator and not the joint decision of the spouse (Ahearn et al., 2006; Phimister and Roberts, 2006; Weiss, 1997). Second, the data is cross-sectional. Most of the empirical literature in this area tends to analyze labor allocation decisions of economic agents using cross-sectional data and limited dependent variable models (Maddala, 1986; Tobin, 1958). But such methods implicitly assume that model parameters are stable (constant) across firms and over time.

The use of cross-sectional methods is therefore glaringly at odds with the firm-specific aspects of the theoretical models employed in labor economics. Even though Sumner (1991) identified the utilization of longitudinal data to analyze farmers off-farm work decision as one of the most useful and promising extensions of the early literature, the use of such data is still rare (Ahituv and Kimhi, 2006). Panel data studies of the labor allocation effect generally control for this endogeneity through fixed effects or alternative instrumental variables estimators (Baltagi, 2008; Robinson, 1989). These procedures are inflexible in their treatment of worker heterogeneity as they generally assume the endogeneity is individual-specific and fixed. A preferable approach would decompose the endogeneity underlying union status into an individual-specific component and an individual or time-specific effect. Finally, Goodwin and Mishra (2004) use a static framework which may be rejected in favor of a dynamic model, which may give important insights when it comes to life-cycle labor supply or transitions between

[2] Smart farming term was used to reflect the use of best management practices, integrated pest management, and precision farming.

states (retirement, multiple-job holding, etc.). To our knowledge, the complications introduced by simultaneous estimation of farm operators' and spouses off-farm labor allocation decisions and farming efficiency has not been investigated.

Herein lies the objective of this study. We consider the simultaneous relationships between farming efficiency and the off-farm labor supply decisions of *both* farm operators and their spouses. Our goals are twofold. First, we consider the determinants of the off-farm labor supply of farm operators and spouses. Further, unlike existing analyses of off-farm labor supply that only model the decisions of the farmer; we estimate the joint decisions of the farm couple (husband and spouse). Finally, we estimate jointly a Tobit model of work activity and an endogenous dynamic panel estimation of farming efficiency, accounting for unobserved heterogeneity and correcting for simultaneity bias (Vella and Verbeek, 1994; Kalwij, 2003). A second important objective of our analysis is to evaluate the relationship between off-farm work, by farm couples and farming efficiency. Here we pay close attention to unobserved heterogeneity through the use of continuous unbalanced panel data (1989-2008) of Norwegian farm households.

Model Framework and Estimation

The following theoretical model of the farm household model illustrates the dependence between the off-farm labor allocation decision of the operator and spouse (Ahearn, El-Osta, and Dewbre 2006; Singh, Squire, and Strauss 1986). The farm household follows a utility maximization framework where it's assumed that the farm operator (*O*) and spouse (*S*) comprise the farm household and utility (*U*) is a function of leisure $\left(L_i\left(F,M\right)\right)$ and income $\left(Y\right)$. Utility is assumed to be affected by human capital characteristics $\left(K^O,K^S\right)$, and other household and regional characteristics $\left(Z_H\right)$ that are considered exogenous to current decisions, as well as unobserved heterogeneity in preferences, α_U. Finally, both farm household income and the time spent on leisure are a function of the time devoted to farming activities (F) and off-farm activities (M).

$$Maximize\ U = U\left(L_i\left(F^O, M^O\right), L_e\left(F^S, M^S\right), K^O, K^S, Y, Z_H, \alpha_U\right) \quad (1)$$

subject to:

$$T^O = L_o + F^O + M^O\left(Z_M, M^S\right) \quad (2)$$

$$T^S = L_s + F^S + M^S\left(Z_M, M^O\right) \quad (3)$$

$$P_y Y = P_f Q\left(X, K^O, K^S, F^O, F^S, R, Z_F, \alpha_Q\right) + W^O M^O\left(M^S\right) + W^S M^S\left(M^O\right) + V - pX \quad (4)$$

$$L_o, F^O, M^O\left(M^S\right) \geq 0 \qquad L_s, F^S, M^S\left(M^O\right) \geq 0 \quad (5)$$

where P_y denotes the price of consumption good Y; L_o is home time (leisure) of the farm operator, O, and L_s is home time (leisure) for the spouse of the farm operator, S; T is the total time endowment, L is the time allocated to leisure, F^i is time allocated to farm work and M^i is time allocated to off-farm work by operator (O) and spouse (S); W^i is the off-farm wage rate. The off-farm wage rates[3] that the operators and spouse face depends on the their respective human capital $\left(K^O, K^S\right)$ characteristics, local labor market conditions Z_M, considered as exogenous, as well as unobserved individual indicated by heterogeneity, indicated by α_W^O and α_W^S. P_f denotes a vector of farm output prices, X denotes a vector of inputs used in the farm production and p denotes a vector of farm input prices, V signifies other nonlabor income including income from government payments/subsidies, K is human capital, and R describes location-specific attributes (e.g., local employment, farm characteristics and soil conditions). Equation 4 is the full income constraint and a non-negativity constraint is represented in

[3] $W^i = W^i\left(K^i, Z_M, \alpha_W^i\right) \quad (i = O, S)$. We assume flexibility in work schedule in off-arm activities, so that both operators and spouses are price takers and wages are determined independently of the number of hours worked.

equations 5. Note that M^O is a function of the off-farm hours worked by the spouse $\left(M^S\right)$. This allows for jointness in off-farm labor allocation decisions. The full income constraint is defined as the sum of income from the operator's off-farm labor $\left(W^O M^O\left(M^S\right)\right)$, spouse's off-farm labor $\left(W^S M^S\left(M^O\right)\right)$, farm profits $\left(P_f Q(.) - pX\right)$ and other household non-labor income (*V*) minus consumption expenditures $\left(P_y Y\right)$. Farm output, Q depends specifically, on the labor hours from operators and spouses $\left(F^O, F^S\right)$, which are assumed to be perfect substitutes, a vector of purchased farm inputs (X), human capital $\left(K^O, K^S\right)$ attributes, observed farm characteristics, Z_F (includes farming efficiency), as well as unobserved heterogeneity in the technology α_Q. Recall that the utility and the production functions are assumed to be concave, continuous and twice differentiable. The Lagrangian ($\mathcal{L}$) can be constructed for the outlined maximization problem with the following first order conditions for off-farm labor:

$$U\left(L_i\left(F^O, M^O\right), L_e\left(F^S, M^S\right), Y\right) \tag{6}$$

$$+\delta\left(P_y Y - \left(P_f Q\left(X, K^O, K^S, F^O, F^S, R\right) + W^O M^O\left(M^S\right) + W^S M^S\left(M^O\right) + V - pX\right)\right)$$

$$+\lambda_1\left(T^O - \left(L_o + F^O + M^O\left(Z_M, M^S\right)\right)\right)$$

$$+\lambda_2\left(T^S - \left(L_s + F^S + M^S\left(Z_M, M^O\right)\right)\right)$$

$$\frac{\partial L}{\partial M^O} \Rightarrow MRS^O_{L_o} - W^O = \left(\frac{\partial M^S}{\partial M_O}\right)\left\{W^S - MRS^S_{L_s,Y}\right\} \tag{7}$$

$$\frac{\partial L}{\partial M^S} \Rightarrow MRS^S_{L_s} - W^S = \left(\frac{\partial M^O}{\partial M^S}\right)\left\{W^O - MRS^O_{L_o,Y}\right\} \tag{8}$$

Per the cross-partial derivative in equations (7) and (8) if the off-farm labor allocation decision of the spouse is independent of the operator, then $\frac{\partial M^S}{\partial M^O} = \frac{\partial M^O}{\partial M^S} = 0$. For the operator, this implies that utility is maximized where the marginal rate of substitution $\left(MRS^{O}_{L_o,Y}\right)$ between leisure and the consumption goods is exactly equal to the off-farm wage, and $P_f Q'_{F^O} = W^O$ or that the value of the marginal product of farm labor is equal to the off-farm wage rate. Similarly, spouse's utility is maximized where the marginal rate of substitution $\left(MRS^{S}_{L_s,Y}\right)$ between leisure and the consumption goods is exactly equal to the off-farm wage, and $P_f Q'_{F^S} = W^S$ or that the value of the marginal product of farm labor is equal to the off-farm wage rate. Corner solutions are implied if either on-farm or off-farm labor supply is zero. Now if we hold the total amount of labor supplied constant, an increase in the price of output $\left(P_f\right)$ or an increase in on-farm labor productivity would yield more labor being supplied to the farm and less to off-farm activities. Similarly, an increase in off-farm wage rate would decrease the number of hours supplied to farming activities. However, it should be noted that the impact of such changes on overall labor supply depends on the competing income and substitution effects.[4]

Differentiating equation 6 with respect to $Y, L^O, L^S, M^O, M^S, F^O, F^S, X$ gives the first-order conditions. One can obtain the virtual on-farm labor supply, leisure, and off-farm labor supply functions. The goal of our analysis lies in providing estimates of descriptive off-farm labor supply decisions rather than explicit estimation of a structural model of labor supply. Thus we show a and relate off-farm labor supply decisions and on-farm efficiency (Φ) measures to observable and unobservable farm, operator, and spouse characteristics reflected in the determinants of wages, prices, and characteristics of production and utility functions. Specifically, we estimate the following model;

$$F^i = \Omega_{F^i}\left(P_f, p, W^O, W^S, K^O, K^S, Z_F, \alpha_Q\right), \qquad (9)$$

[4] The issue becomes more complicated as off-farm labor supply of farm operators also depends on how spouses react to changes in their off-farm wages. Notice that M^O, hours spent working off the farm is a function of the off-farm hours worked by the spouse $\left(M^S\right)$.

$$L^i = \Omega_{L^i}\left(P_f, p, W^O, W^S, K^O, K^S, Z_F, \alpha_Q, V, Z_H, \alpha_U, \alpha_Q\right), \tag{10}$$

$$M^i = \Omega_{M^i}\left(P_f, p, W^O, W^S, K^O, K^S, Z_F, \alpha_Q, V, Z_H, T^i, \alpha_U, \alpha_Q\right), \quad (i = O, S). \tag{11}$$

$$\Phi = \left(P_f, p, W^O, W^S, K^O, K^S, \alpha_Q, V, Z_H, T^i, \alpha_U, \alpha_Q\right), \quad (i = O, S). \tag{12}$$

Stochastic specification and model estimation

The model framework contains seven specific equations in the seven endogenous variables (or variable vectors), $(F^O, F^S, L^O, L^S, M^O, M^S, \Phi)$ of which the six time allocation variables are subject to censoring as specified in equation (5). Before operationalizing the stochastic version of the model we also need to account for the incidental truncation of the off-farm wage rates of operator and spouse, (W^O, W^S). Wage rates can only be observed for individuals participating in off-farm work. Following the logic of the model we must assume that the reservation wage, or the marginal product of farm labor, exceeds the market wage rate when an individual chooses not to participate in off-farm work. The stochastic version of the model consequently consists of two wage equations (13) where the wage rate is only observed for individuals participating in off-farm work, six censored time allocation equations for operator's and spouse's farm, leisure and off-farm hours (14)-(16), one equation for efficiency in farm production (17) which includes pre-determined variables of endogenous variables, six observable random variables, (18)-(20), related to the latent time allocation variables (14-16), and two observable random variables (21) related to the truncated wage rates (13). In addition, the model imposes the restriction that any individual's time consumption$(f_{iht}^* + l_{iht}^* + m_{iht}^*)$ can not exceed total time endowment in any time period as specified in equations (2)-(3). With subscript $i, j = o, s$ representing operator and spouse respectively, $h = 1, \dots, H$ representing farm household, $t = 1, \dots, T$ representing time period, and $s = 0, \dots, p$ representing lags from time period t, the full simultaneous equations model system can be presented by the following 17 equations.

$$w_{iht}^{*} = \alpha_i^w + x_{iht}^w \beta_i^w + u_{ih}^w + v_{iht}^w \tag{13}$$

$$f_{iht}^{*} = \alpha_i^f + \sum_j w_{jht}^{*} \eta_{ij}^f + x_{iht}^f \beta_i^f + u_{ih}^f + v_{iht}^f \tag{14}$$

$$l_{iht}^{*} = \alpha_i^l + \sum_j w_{jht}^{*} \eta_{ij}^l + x_{iht}^l \beta_i^l + u_{ih}^l + v_{iht}^l \tag{15}$$

$$m_{iht}^{*} = \alpha_i^m + m_{jht,j\neq i}^{*} \delta_i^m + \varphi_{ht} \gamma_i^m + \sum_j w_{iht}^{*} \eta_{ij}^m + x_{iht}^m \beta_i^m + u_{ih}^m + v_{iht}^m \tag{16}$$

$$\varphi_{ht} = \alpha^{\varphi} + \sum_{s=1}^{p} \varphi_{ht-s} \gamma_s^{\varphi} + \sum_j \sum_{s=0}^{p} m_{jht-s}^{*} \delta_s^{\varphi} + \sum_j w_{jht}^{*} \eta_j^{\varphi} + x_{ht}^{\varphi} \beta^{\varphi} + u_h^{\varphi} + v_{ht}^{\varphi} \tag{17}$$

$$f_{iht} = \begin{cases} f_{iht}^{*} & if \;\; f_{iht}^{*} > 0 \\ 0 & otherwise \end{cases} \tag{18}$$

$$l_{iht} = \begin{cases} l_{iht}^{*} & if \;\; l_{iht}^{*} > 0 \\ 0 & otherwise \end{cases} \tag{19}$$

$$m_{iht} = \begin{cases} m_{iht}^{*} & if \;\; m_{iht}^{*} > 0 \\ 0 & otherwise \end{cases} \tag{20}$$

$$w_{iht} = \begin{cases} w_{iht}^{*} & if \;\; m_{iht}^{*} > 0 \\ missing & otherwise \end{cases} \tag{21}$$

where (w_{iht}^{*}) is the rate for operator or spouse in household *h* at time *t*, $(f_{iht}^{*}, l_{iht}^{*}, m_{iht}^{*})$ are hours spent at farm work, leisure, and off-farm work, respectively, (φ_{ht}) is efficiency in farm production defined as farm revenue divided by total variable costs, $(f_{iht}, l_{iht}, m_{iht}, w_{iht})$ are observable random variables related to the latent variables $(f_{iht}^{*}, l_{iht}^{*}, m_{iht}^{*}, w_{iht}^{*})$. $(x_{\circ}^{\circ})$ are vectors of observable exogenous variables of which some are time-invariant, some are household-invariant and most are time- and household-varying $x_{it}^{\circ} = (x_i^{\circ}, x_t^{\circ}, x_{it}^{\circ})$. The vectors of observable exogenous variables may vary between the different equations of the model system and also between operator and spouse. $(\alpha_{\circ}^{\circ}, \beta_{\circ}^{\circ}, \eta_{\circ}^{\circ}, \gamma_{\circ}^{\circ}, \delta_{\circ}^{\circ})$ are vectors of unknown parameters, $(u_{\circ}^{\circ})$ are random heterogeneity parameters and $(v_{\circ}^{\circ})$ are genuine random disturbances. (v_{ht}^{φ}) may be serially correlated and in a simultaneous equations system $(u_{\circ}^{\circ}, v_{\circ}^{\circ})$ may also be correlated with covariates across equations. The system of equation can be generalized further by allowing for lagged dependent variables in all structural equations (14)-(17).

Now turning our attention to model estimation; we find that estimation of the parameters in the above system is not straight forward. The model includes linear and non-linear equations, dynamics, and endogenous regressors. When we have endogenous variables among the regressors, we can apply 2 stage least squares (2SLS) estimation in linear models when there are appropriate instruments available. The joint determination of the variables in this model is in part recursive because the wage equations (13) are completely determined by exogenous factors, although possibly correlated with the off-farm labor supply equations through the truncation mechanism in equations (21). It is, however, reasonable to assume that off-farm wage rates are determined in the labor market irrespective of work effort and technology on the farm.

Let us assume that $(\varepsilon_{iht}^{w} = u_{ih}^{w} + v_{iht}^{w})$ are *i.i.d.* normal errors with zero mean and standard deviation $(\sigma_{iu}^{w} + \sigma_{iv}^{w})$ and that the two wage equations are uncorrelated, $Corr(\varepsilon_{oht}^{w}, \varepsilon_{sht}^{w}) = 0$. We can then separate the wage equations from the full system and apply a Heckman two-stage procedure (Heckman 1974, 1976, 1979) to predict wage rates for individuals not participating in off-farm work. The first step of Heckman's method is to estimate the probability of observing the wage rate, *i.e.*, the probability of participating in off-farm work, $\Pr(m_{iht}^{*} > 0)$. Let participation be represented by an index dummy variable (I_{iht}) which gives the selection mechanism indicated by

$$I_{iht} = \begin{cases} 1 \;\; if\; m_{iht}^{*} > 0 \\ 0 \; otherwise \end{cases}, \qquad (i = O, S\,,\; h = 1, \ldots, H\,,\; t = 1, \ldots, T) \qquad (22)$$

The wage equation and off-farm labor supply is still given by equations (13) and (16) where we for simplicity assume that $(\gamma_{i}^{m} = \delta_{i}^{m} = \eta_{ij}^{m} = 0)$ so that labor supply is a function of solely exogenous variables.

$$w_{iht}^{*} = \alpha_{i}^{w} + x_{iht}^{w}\beta_{i}^{w} + u_{ih}^{w} + v_{iht}^{w} \qquad (13)$$

$$m_{iht}^{*} = \alpha_{i}^{m} + x_{iht}^{m}\beta_{i}^{m} + u_{ih}^{m} + v_{iht}^{m} \qquad (16')$$

In the above equations $(x_{iht}^{w}, x_{iht}^{m})$ are different vectors of selected and observed personal, farm, and labor market characteristics, $(\alpha_i^w, \alpha_i^m, \beta_i^w, \beta_i^m)$ are random parameters, (u_{ih}^{w}, u_{ih}^{m}) are normally distributed household-specific heterogeneity $N{\sim}(0, \sigma_{iu^{\circ}}^{2})$ and correlation coefficient θ_i, and $(v_{iht}^{w}, v_{iht}^{m})$ are genuine binormal disturbances with zero means, variances $(\sigma_{iv^w}^2, \sigma_{iv^m}^2 = 1)$, and correlation coefficient ρ_i.

A probit estimation of the participation decision, represented by I_{iht}, provides estimates of Heckman's lambda λ_{ih}, $(\hat{\lambda}_{ih} = \phi(\hat{\alpha}_i^m + x_{iht}^m \hat{\beta}_i^m + \hat{u}_{ih}^m)/\Phi(\hat{\alpha}_i^m + x_{iht}^m \hat{\beta}_i^m + \hat{u}_{ih}^m))$, which subsequently are used to produce starting values for $(\hat{\alpha}_i^w, \hat{\beta}_i^w, \hat{u}_{ih}^w)$ to be applied in a final maximum likelihood estimation of the wage equations to predict wage rates for those not participating in off-farm work. The log likelihood function then becomes (2002 Econometric software, p. E23-3).

$$ln\, L_{ih} = \sum_{I_{iht}=0} log\Phi(-\alpha_i^w - x_{iht}^w \beta_i^w - u_{ih}^w) + \sum_{I_{iht}=1} \frac{-log2\pi}{2} + log\kappa_i - \left(w_{iht}^{*}\kappa_i - \tilde{\theta}_{ih} - x_{iht}^w \tilde{\beta}_i^w\right)^2 + log\Phi\left[\left(\sqrt{1-\tilde{\rho}_i^2}\right)(\alpha_i^w + x_{iht}^w \beta_i^w + u_{ih}^w) + \tilde{\rho}_i\left(w_{iht}^{*}\kappa_i - \tilde{\theta}_{ih} - x_{iht}^w \tilde{\beta}_i^w\right)\right] \quad (23)$$

where $\kappa_i = 1/\sigma_{iv^w}^2$, $\tilde{\beta}_i^w = \left(1/\sigma_{iv^w}^2\right)\beta_i^w$, $\tilde{\theta}_{ih} = \theta_i/\sigma_{iv^w}^2$, $\tilde{\rho}_i = \rho_i/\sqrt{1-\rho_i^2}$.

Even with predetermined (predicted) wage rates, the remaining model system still contains thirteen linear and non-linear equations with elements of censoring, lagged dependent variables, and heterogeneity. We need to simplify further and we do this by delimiting the model to the equations relevant for our research questions. Our main interest is to investigate the relationship between the off-farm labor supply of both farm operator and spouse and farming efficiency. We choose to disregard equations (14), (15), (18) and (19) and only operationalize equations (16), (17) and (20). This implies that we assume zero correlation between on- and off-farm labor decisions, something which intuitively seems unrealistic although we can argue that off-farm labor supply is mainly determined in the market. The simplification leaves us with a system of five-equations to be estimated simultaneously. The system consists of the two incidentally censored off-farm labor equations (16) with observable random variables (20) and a dynamic farm efficiency equation (17). We solve the estimation problem by use of the procedure proposed by Nelson and Olson (1978). A general model (Hsiao,1996) of this type containing G structural equations can be expressed as

$$\mathbf{\Gamma}\boldsymbol{y}_{ht}^{*} = \boldsymbol{\alpha} + \mathbf{B}\boldsymbol{x}_{ht} + \boldsymbol{\varepsilon}_{ht}, \quad h = 1, \dots, H \ , t = 1, \dots, T \tag{22}$$

$$\mathbf{\Gamma}\boldsymbol{y}_{ht} = \boldsymbol{h}(\boldsymbol{y}_{ht}^{*}) \tag{23}$$

where $\mathbf{\Gamma}$ and $\mathbf{B}$ are $G \times G$ and $G \times K$ matrices of coefficients; $\boldsymbol{y}_{ht}^{*}$ is a $G \times 1$ vector of latent dependent variables with corresponding observed variables $\boldsymbol{y}_{ht}$ defined by some $G \times 1$ vector function $\boldsymbol{h}(\boldsymbol{y}_{ht}^{*})$. $\boldsymbol{x}_{ht}$ is a $K \times 1$ vector of exogenous variables, including the wage rates, $\boldsymbol{\alpha}$ is a $G \times 1$ vector of intercepts, and $\boldsymbol{\varepsilon}_{ht} = \boldsymbol{u}_h + \boldsymbol{v}_{ht}$ is a $G \times 1$ vector of unobserved household effects and true disturbances and are multivariate normally distributed with zero mean and covariance matrices $Eu_h u_i = \mathbf{\Omega}_u$ if $h = i$ and $Ev_{ht}v_{is} = \mathbf{\Omega}_v$ if $h = i, t = s$. Multiplying (22) and (23) by $\mathbf{\Gamma}^{-1}$ on both sides gives the reduced form equations. In his 1979 paper, Amemiya derived the asymptotic variance-covariance matrix for the Nelson-Olson estimator when only one equation is subject to censoring and Schmidt (*see* Manski and McFadden, 1981) considers the constraints on the parameters in simultaneous Tobit models when some or all endogenous variables are truncated. Identification of the gth structural equation requires that the number of independent restrictions on the $1 + K + G - 1$ coefficients $(\boldsymbol{\alpha}, \mathbf{B}, \mathbf{\Gamma})$ exceeds the number of equations minus one, $(G - 1)$.

The joint density of the random effects and true errors becomes messy in a system of five simultaneous equations and the number of multiple integrals in the log likelihood function consequently rendering an unsuitable maximum likelihood. Had the model been recursive, or the idiosyncratic errors been independently identically distributed over h and t, three least squares techniques (3SLS) would have been feasible. Because this is not the case, the least squares estimators would be inconsistent. One alternative approach to the estimation problem is to calculate the reduced-form equations, where we initially ignore the dynamics in equation (17) in order to avoid dynamic panel bias and make estimations more straightforward in the first stages. By excluding lagged values of the left hand side variable we are able to predict values for all years and thus base the second stage estimations on the whole sample. Let the reduced form representations of equations (16), (17) and (19) be given by

$$m_{oht}^* = \pi_0^o + x_{ht}'\pi_1^o + \epsilon_{ht}^o \quad (24)$$

$$m_{sht}^* = \pi_0^s + x_{ht}'\pi_1^s + \epsilon_{ht}^s \quad (25)$$

$$\varphi_{ht} = \pi_0^\varphi + x_{ht}'\pi_1^\varphi + \epsilon_{ht}^\varphi \quad (26)$$

$$m_{iht} = \begin{cases} m_{iht}^* & if \;\; m_{iht}^* > 0 \\ 0 & otherwise \end{cases}, \; i = o, s \quad (27)$$

where $m_{iht}^*, m_{iht}, \; i = o, s$ is the latent and observed off-farm labor supply and $\varphi_{ht}$is observed farm efficiency. We assume that the covariance matrices of $(\epsilon_{ht}^o, \epsilon_{ht}^s, \epsilon_{ht}^\varphi)$ are unrestricted. While estimation of the structural system with two latent and one observed variable is computationally difficult, even in the simplified version with no dynamics, estimation of the reduced form equations is relatively simple. The reduced form coefficients are estimated by maximum likelihood applied to each of the off-farm labor supply equations (24) and (25) separately and GLS on equation (26) to form instruments for the latent and observed dependent variables $(\hat{m}_{oht}^*, \hat{m}_{sht}^*, \hat{\varphi}_{ht})$ to be used in a second stage estimation of the structural equations. The instruments $(\hat{m}_{iht}^* = \hat{\pi}_0^i + x_{ht}'\hat{\pi}_1^i, \; i = o, s)$ and $(\hat{\varphi}_{ht} = \hat{\pi}_0^\varphi + x_{ht}'\hat{\pi}_1^\varphi)$ are asymptotically uncorrelated with the disturbances. In a second step we replace the endogenous right-hand side variables with their instruments and treat the instruments as fixed regressors to estimate the structural parameters in equations (16) by maximum likelihood applied to each equation separately. The two-stage estimates are consistent and asymptotically normal and uncorrelated with the error terms (Nelson and Olson 1978). The efficiency function (17) is estimated by two-stage GLS both with and without AR(1)-disturbance. Because ordinary GLS estimation gives inconsistent estimators in the dynamic panel data model, we compare the two-stage estimations with instrumental variables estimation using both exogenous and lagged variables as instruments for the endogenous regressors.

Data and sample selection

The panel data for this study comes from Norwegian farm households (1989-2008). The data is collected by Norwegian Agricultural Research Institute (Norsk institutt for landbruksforskning, NILF). This is a yearly survey amongst approximately 1,000 farm households representing different regions and agricultural produce. Most farm household report between 1,800 and 3,100 on-farm hours per year while a standard man-labor year in agriculture is set to 1,875 hours. On-farm hours are normally distributed with mean and median of approximately 2,500 hours. On average, operator work 2,000 hours and spouses 450 hours annually. Hired help is on average 400 hours but with standard deviation of more than 500 hours and 30 per cent of the farm households report no hired help on the farm.

The original panel data set, as well as our final sample, is unbalanced, and some 5-10 per cent of the respondents are replaced each year. The extracted sample covers 20 years from 1989 to 2008 and 17,605 observations (total 19,972) were used in the analysis. The attrition of almost 2,400 observations is due to single adult households. The unbalanced panel includes 1,791 unique households which are represented for anything from three to 20 years, and on average for ten years. We find no evidence of potential endogenous sample selection to influence the results but some self-selection may be present because of voluntary participation in the survey and attrition bias from excluding farm units where there is no spouse present.

Variable definitions and descriptive statistics

The variables of particular interest to our study is off-farm work hours and farm efficiency Multi-employment is common amongst both farm operators and spouses and many farm operators even have their main income from outside farming. Most farm operators report off-farm work in at least some years, but many only supply a marginal number of hours. As many as 20 per cent work between 0 and 37.5 off-farm hours annually. This finding is not surprising because it is well known that many farm operators take on small commissions, e.g., for neighbors (road mending, snow clearing, holiday relief etc.). For this reason, we define working off the farm as having more than 37.5 annual working hours outside the farm. This threshold equals one standard labor week. Operators working less than this are defined as not working off the farm and we rationalize this truncation by the effect on possible measurement error and misclassification in the data. Although the problem of few reported off-farm hours occurs less frequently for spouses, we choose the same definition for both partners. We see from table 1 that spouses, on average, work more off-farm hours annually than operators but the variance is greater. Both for operator and spouse there are a few extreme observations in the upper tail of the distribution with reported off-farm hours exceeding two standard man labor years and for this reason we choose to include upper censoring in the off-farm hours estimations at one standard deviation above a standard labor year.

Efficiency is defined as farm revenue divided by total variable cost. On average, revenue is almost four times the total variable cost and median efficiency is 3.5. Approximately 80 per cent of the observations lie within the range between 2 and 5. All three endogenous variables show greater between than within variance. Because several of the variables in the explanatory models are time-invariant or nearly time-invariant, this suggests random effects modeling. Summary statistics and definition of the variables are provided in table 1.

Results and Discussion

(a) Labor supply functions

We start with presenting the results for the factors affecting *off-farm labor supply* of farm operators and spouses. Table 2 presents the results of the off-farm labor participation model for both farm operators and spouses. We use simultaneous equation two stage Tobit estimation. A statistical test devised by Hausman (1978) tests for orthogonality of the random effects (REM) and the regressors. Under the null hypothesis the REM model is the correct specification and we fail to reject the null hypothesis in our study. The two stage Tobit model has a much higher value of the likelihood function than the single equation estimation and also a higher random effects contribution to total variance with $\rho=0.835$ for operators (column 3, table 2) and $\rho=0.826$ for spouses (column 3, table 3).

There are several interesting finding from this study. First, we find that the off-farm labor supply of farm operators in negatively correlated with the spouse's off-farm labor supply. Results indicate that an additional hour of off-farm work by the spouse decreases off-farm labor supply of the farm operator modestly by 4 hours. However, the opposite is true when your look at the correlation between spouses' and farm operator's labor supply. For example, an additional hour of off-farm work by the farm operator increases off-farm labor supply of spouse's by approximately 17 hours (table 2).

Farming efficiency (ratio of farm revenue to total variable cost) has an interesting impact on off-farm labor supply of farm operators and spouses. Because of endogeneity we use the predicted values of farming efficiency in the regression. Results, in table 2, show that farming efficiency has a positive and negative impact on hour of off-farm work by farm operators and spouses, respectively. For example, a 1% increase in farm efficiency increases off-farm hours of farm operators by approximately 7% and decreases off-farm work by spouses about 2%. Human capital variables like age, education, and farming experience have expected signs and impact on off-farm labor supply. The direction of all of the relationships between human capital, farming experience and off-farm work hours was as expected. These expected relationships have been consistently found in the empirical literature (e.g., Mishra et al. 2002; Ahearn, El-Osta & Dewbre, 2006; Benjamin and Kimhi, 2006; Glauben et al. 2008; Mishra and Goodwin , 1997; 1998). However, it should be noted that farming experience of farm operators has a positive and significant impact on the off-farm work hours of spouses. Finding here may suggest that increase farming experience may indicate comparative advantage and in farming, as a result spouses tend to increase their off-farm work hours. Finally, number of children under the age of six has a negative impact on the hours worked of the farm for both farm operators and spouses. Findings reported here are consistent with those obtained Mishra and Goodwin, 1997; Ahearn, El-Osta & Dewbre, 2006; El-Osta, Mishra and Ahearn, 2004.

The estimated off-farm wages[5] for both the farm operator and the spouse have an impact on off-farm labor supply. Results indicate that higher off-farm wage rate of operators tends to increase the off-farm hours of both farm operators and spouses, whereas higher off-farm wage rate of spouses tend to decrease off-farm hours of both operators and spouses. This finding needs further investigation. The positive impact of hired farm workers (hired labor hours) on operator's hours of off-farm work suggests that hired labor and farm work by operators may be complementary inputs in agricultural production. On the other hand, a negative impact of hired farm workers (hired labor hours) on spouse's hours of off-farm work suggests that hired labor and spouse farm

[5] Hourly wage rate is determined by age, the probability of having attained higher education, work hours devoted to farming, the centrality of the region, and the regional employment rate. The off-farm wage rates are used to predict wage rates for individuals not working off the farm.

work may be viewed as substitutable inputs in agricultural production. Our findings here are consistent with the finding in the literature (Benajmin, Corsi, and Guyomard, 1996).

The variable of special interest in this study is agricultural subsidy. Results in table 2 indicate that the coefficient of agricultural subsidy is negative and statistically significant at the 1% level of significance in the farm operator case—suggesting that agricultural subsidy reduces off-farm work hours of farm operators. On the other hand, the coefficient of agricultural subsidy is positive and statistically significant at the 1% level of significance in the spouse case. A plausible explanation is that agricultural subsidy makes off-farm job (off-farm wage rate) more attractive and perhaps spouses have extra time that they can easily devote to off-farm work. Findings reported here are consistent with those obtained Mishra and Goodwin, 1997, 1998; Ahearn, El-Osta & Dewbre, 2006; Dewbre and Mishra 2007; El-Osta, Mishra and Ahearn, 2004.

Agricultural income was significant and positive for farm operators and negative and significant for spouses. The positive sign indicates that as agriculture income increases, the number hours working off the farms increases. At first this may seem to be counterintuitive, but take the case of Norway where farms cannot sell agricultural land for other uses, but can only rent it out for agricultural production, this result makes sense. This notion is reinforced further, by a positive and significant coefficient on farm size (total farming acres). On other hand, the coefficient on agricultural income is negative and statistically significant in the case of spouse—indicating a wealth effect.

The coefficient of other income, an indicator of nonlabor, was negative and statically significant at the 1% level of significance for both farm operators and spouses—suggesting that nonlabor income decreases hours of off-farm work—wealth effect. Our findings are consistent with Mishra and Goodwin, 1997, 1998; Ahearn, El-Osta & Dewbre, 2006; Dewbre and Mishra 2007; El-Osta, Mishra and Ahearn, 2004. Finally, we found that local area variables, such as regional unemployment rate and location of the farming household, are important in explaining off-farm work hours by farm couples. For example, if the share of regional employment rate (regional employment/total population in the region) is high—indicating a healthy off-farm labor market and increased opportunity cost of farm work—then farm operator and spouses are more likely to find off-farm employment. Our findings are consistent with Mishra and Goodwin, 1998. The coefficient on location of the farm and/or farm household (farms located in metro area with population 10,000 and driving distance of 45 minutes or more) variable is negative and significant at the 1% level of significance. As expected these results show that lack of opportunity for off-farm work—driving distance and demand for work—reduce the number of hours worked off the farm by farm operators and spouses.

(b) Labor supply and farming efficiency

Let us now turn our attention to the main objective of this study: the impact of off-farm labor supply on farming efficiency. We apply a two-stage Generalized Least Squares (GLS) model on the predicted values of the endogenous variables, $(\hat{m}^{*}_{oht}, \hat{m}^{*}_{sht}, \hat{\varphi}_{ht})$, from the reduced form estimation. The lagged variables are assumed predetermined in the two-stage GLS. Initially, we assume that household heterogeneity is represented by time-invariant household effects, uncorrelated with predetermined and exogenous variables but we estimate efficiency both with the assumption of random and fixed effects and compare the two specifications by use of the Hausman test statistic. The efficiency equation includes regressors which are nearly, but not fully time-invariant. The random effects estimators are supposedly more efficient, but because the random effects GLS-estimator is biased in the dynamic panel data models (Baltagi, 2008), the consistent fixed effects estimator may be preferred. The Hausman test for comparing fixed effects against random effects confirms the preference for fixed effects modeling even though table 3 show that the results of the random and fixed effects GLS produce similar parameter estimates both in magnitude and efficiency (see table 3). We also see that both the within and between samples R^2 is large in both the random and fixed effects estimations.

We repeat the same random and fixed effects GLS estimation procedure, but we now include an AR(1)-process for the true errors. Results show a systematic difference between the fixed and random effects estimator (table 3, column 4 and 5). The random and fixed effects coefficients of the autoregressive model are not similar, at least for the lagged regressors, and the random effects seems to be more efficient (table 2). When we compare across models with and without autoregressive errors, we find that the two random effects estimations produce nearly identical result, same as the two fixed effects estimations. The Baltagi-Wu locally best invariant (lbi) test of autocorrelation (H_0: $\rho=0$) rejects the presence of a positive serial correlation in the data, implying that ordinary 2S GLS is preferred over the AR 2S GLS. In both the fixed and random effects model, the test statistic equals 1.9.

Taking GLS fixed effects model as the appropriate model (column 3, table 3) results in the farming efficiency model, table 3 the lagged dependent variable is significant for both t-1 and t-2 time period, suggesting the possible existence of dynamics; it should be noted that the coefficient in t-1 period is twice as large (0.18 compared to 0.09) as the coefficient in the t-2 time period. This finding suggests that farming efficiency, after controlling for off-farm labor supply of farm operators and spouses, increases with a decreasing rate—a time adjustment period is warranted. Finally, it should be pointed out that Goodwin and Mishra (2004) did not use lagged farming efficiency variable in their model. Failure to account for such variables may lead to biased results.

Time spent in off-farm work, both my farm operators and spouses, have an impact on farming efficiency. Using predicted values of off-farm work hours of farm operators results in table 3 (column 3) show that hours worked off-farm in *t* and *t-1* period has a positive and negative impact on farming efficiency, indicating that off-farm work increases farming efficiency in the current period, however, in the second period—consistently working off the farm—decreases farming efficiency. A plausible explanation is that if the operator is devoting more and/or consistently devoting hours to off-farm work it is likely that he/she may be interested in off-farm work as a permanent source of income. Further, it may also signal that the farm operator is losing touch with production agriculture and may be relying on hired hands to perform tasks related to production agriculture, thus resulting is loss of efficiency. The results obtained here are in contrast to Goodwin and Mishra (2004) who, using a cross-sectional data found that off-farm work by farm operators decreased on-farm efficiency. Findings are also in disagreement with Smith's (2002) conjecture that off-farm work reduces farming efficiency. However, our study finds that Smith's argument may not valid if one takes into account the dynamics of labor allocation decision (panel data). Findings from our study indicate that the impact of off-farm work on farming efficiency is dynamic in nature. In particular, results show that off-farm work by farm operators first increases on-farm efficiency (*t* period) and then decreases it in the second period (*t-1*).

Turning our attention to the spouse—which was ignored in the Goodwin and Mishra's (2004) study—results in table 3 (column 3) shows that spouse's off-farm hour's increase farming efficiency. This finding may signal specialization in some sense. Spouses' hours worked off-farm in *t* and *t-1* period is positive and the impact is five times smaller in the *t-1* period. This finding may lend credence to the fact that spouses are more likely to work off the farm and in permanent jobs for fringe benefits (Mishra et al. 2002).

In what seems to be counterintuitive finding, greater farming experience appears to be negatively correlated with farming efficiency. An explanation for this finding could be that older farmers may be less to adopt new technologies and thus may fail to realize certain efficiency advantages that come with technological advances. Thus, our finding of a negative correlation between farming experience and farming efficiency is not unexpected. An interesting find to report here is that agricultural subsidies increase farming efficiency. The coefficient of agricultural subsidy is positive and statistically significant at the 1% level of significance. Farms receiving agricultural subsidy may specialize in crop or commodities that are supported through agricultural support payments and thus encouraging investments in farming business that increase production and efficiency of production agriculture.

Conclusions

A plethora of literature has evolved that investigates the determinants of farm household involvement in off-farm labor markets. Most of the empirical literature in this area tends to analyze labor allocation decisions of economic agents using cross-sectional data and limited dependent variable models. But such methods implicitly assume that model parameters are stable (constant) across firms and over time. The present study considers the simultaneous relationships between farming efficiency and the off-farm labor supply decisions of *both* farm operators and their spouses, which has been ignored in the literature. We investigate this issue while correcting for unobserved heterogeneity through the use of continuous unbalanced panel data (1989-2008) of Norwegian farm households.

Results from the panel data show that that the off-farm labor supply of farm operators in negatively correlated with the spouse's off-farm labor supply. Farming efficiency (ratio of farm revenue to total variable cost) has a positive and negative impact on hours of off-farm work by farm operators and spouses, respectively. Agricultural subsidy reduces off-farm work hours of farm operators; in the case of spouses, agricultural subsidy hours of off-farm work are positively correlated. Finally, we found that local area variables, such as regional unemployment rate and location of the farming household, are important in explaining off-farm work hours by farm couples.

The two-stage GLS fixed effects model is preferred when modeling the relationship between farming efficiency and off-farm labor supply of farm couples. Results from this study indicate that the correlation between farming efficiency and off-farm labor supply of farm couples is dynamic in nature. In the case of the farm operator, off-farm work first increases farming efficiency in the first period and then decreases it in the second period. On the other hand, we observe a positive correlation, for both periods, between off-farm hours worked by spouses and farming efficiency. This study highlights the importance of heterogeneity and correcting for simultaneity bias when investigating labor allocation decisions of farm couples. The present study also contributes to the literature by investigating the impact of joint labor allocation decision—farm operator and spouse—on farming efficiency.

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Table 1: Summary statistics and definition of variables

Variable	Mean	Std. Dev.	Min.	Max.
Operator Off-farm hours (annual hours)	491.562	625.658	0	3751
Spouse Off-farm hours (annual hours)	722.45	710.854	0	4400
Farming efficiency[1]	3.821	1.962	0.634	72.933
Operator age (*years*)	46.062	9.696	12	79
Spouse age (*years*)	44.08	9.914	16	78
Number of Children (under the age of 6)	0.229	0.567	0	4
Operator—probability of having a higher education	0.088	0.079	0	0.982
Spouse—probability of having a higher education	0.224	0.1	0.033	0.996
Operator—probability of having an agricultural education	0.664	0.087	0	0.855
Operator wage rate (*Kr/hr*)	130.284	41.487	60.074	605.396
Spouse wage rate (*Kr/hr*)	129.894	35.583	60.557	534.63
Operator farming experience (*years*)	16.076	9.74	-13	60
Operator annual farming hours	2,025.99	798.206	2	7650
Spouse annual farming hours	449.50	565.843	0	4970
Hired farm labor hours	397.44	539.074	0	11000
Milk yield (*1,000 litres*)	57.516	58.872	0	710.312
Agricultural subsidies (*1,000 kroners*)	200.091	103.173	0	1,268.724
Agricultural income, gross farm income (*1,000 kroners*)	735.435	457.242	2.37	7926.377
Other income (*1,000 kroners*)	45.534	98.915	-239.163	3,604.47
Investments/capital[2]	0.084	0.125	0	1.306
Total farming area (*hectares*)	22.436	14.489	0	173.8
Rented land (*hectares*)	0.24	0.232	0	2.44
Fallow land (*hectares*)	0.077	0.149	0	1.835
Livestock units (AEU)[3]	40.638	62.171	0	1440.4
Regional employment/total population in the region (Share of regional employment)	0.747	0.081	0.41	1.07
Location Metro 1 area with 50,000 population(=1 if the farm is located in the Metro 1 area, with 45 minute drive, 0 otherwise)	0.243	0.429	0	1
Location Metro 2 area with 10,000 population (=1 if the farm is located in the Metro 2 area, with 45 minute drive, 0 otherwise)	0.455	0.498	0	1
Location Highland (=1 if the farm is located in the highland region, 0 otherwise)	0.653	0.476	0	1
Number of Observations		17,605		

[1] Defined as ratio of farm revenue to total variable cost.
[2] Total farm investments in a year/total farm capital stock.
[3] AEU (animal equivalent unit): One AEU equals 1000 pounds of animal weight. As an example, a calf that weighs 500 pounds is 0.5 AEUs.

Table 2: Estimates of Two Stage Random Effects Tobit model of off-farm labor supply of operators and spouses, Norway: 1989-2008

Variable	***Farm Operator*** 2 Stage Random Effects Tobit	***Spouse*** 2 Stage Random Effects Tobit
Ln Off-farm hours, spouse	-0.0397***	-
	(0.0034)	-
Ln Off-farm hours, operator	-	0.1712***
	-	(0.0166)
Ln farming efficiency	7.2702***	-1.7573***
	(0.0852)	(0.2173)
Operator age	0.2352***	-
	(0.0025)	-
Operator age squared	-0.0029***	-
	(0.0000)	-
Spouse age	-	0.2749***
	-	(0.0063)
Spouse age squared	-	-0.0034***
	-	(0.0001)
Number of children under 6	-0.1633***	-0.1078***
	(0.0035)	(0.0085)
Pr(higher education), operator	-0.3831***	-13.3074***
	(0.0773)	(0.1409)
Pr(higher education), spouse	4.1674***	30.8026***
	(0.1198)	(0.1321)
Wage rate, operator	0.0057***	0.0172***
	(0.0001)	(0.0003)
Wage rate, spouse	-0.0085***	-0.0833***
	(0.0003)	(0.0004)
Farming experience, operator	-0.0149***	0.0503***
	(0.0005)	(0.0013)
Hired labor hours	0.0002***	-0.0001***
	(0.0000)	(0.0000)
Agricultural subsidy (*1,000 Kroners*)	-0.0033***	0.0008***
	(0.0000)	(0.0001)
Agricultural income (*1,000 Kroners*)	0.0002***	-0.0015***
	(0.0000)	(0.0000)
Other income	-0.0019***	-0.0065***
	(0.0000)	(0.0001)
Total farming area (*hectares*)	0.0142***	0.0160***
	(0.0003)	(0.0008)
Livestock, units	0.0015***	0.0015***
	(0.0000)	(0.0001)
Share of regional employment	6.1322***	3.1287***
	(0.0522)	(0.1511)
Location Metro 2 area	-0.4163***	-1.4382***
	(0.0229)	(0.0497)
Constant	-14.4245***	0.3049
	(0.0974)	(0.3168)
Observations	17,603	17,603
LL	1734.142	-12196.936
Sigma_u	0.4056	0.8881
	(0.0074)	(0.0162)
Sigma_v	0.1801	0.4076
	(0.0010)	(0.0024)
rho	0.8353	0.8260
LR/Wald χ^2 (18)	40006.01	36856.79
p-value	0.0000	0.0000
LR-test of sigma_u=0, χ^2 (1)	1.9e+04	1.8e+04
p-value	0.000	0.000

Standard errors in parentheses
* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 3: Estimates of Two-Stage GLS model of farming efficiency, Norway: 1989-2008

Variable	**GLS Random Effects**	**GLS Fixed Effects**	**AR GLS Random Effects**	**AR GLS Fixed Effects**
Lag(Ln) farming efficiency	0.2556***	0.1845***	0.2205***	0.0816***
	(0.0061)	(0.0059)	(0.0059)	(0.0062)
Lag_2(Ln) farming efficiency	0.1371***	0.0958***	0.1961***	0.1035***
	(0.0056)	(0.0055)	(0.0056)	(0.0058)
Ln Off-farm hours, operator	0.0296***	0.0310***	0.0296***	0.0308***
	(0.0005)	(0.0006)	(0.0005)	(0.0006)
Lag(Ln) Off-farm hours, operator	-0.0051***	-0.0031***	-0.0052***	-0.0005
	(0.0005)	(0.0005)	(0.0005)	(0.0005)
Lag_2(Ln) Off-farm hours, operator	-0.0033***	-0.0009	-0.0054***	-0.0003
	(0.0005)	(0.0005)	(0.0005)	(0.0005)
Ln Off-farm hours, spouse	0.0023***	0.0020***	0.0024***	0.0019***
	(0.0002)	(0.0002)	(0.0002)	(0.0002)
Lag(Ln) Off-farm hours, spouse	0.0007***	0.0004*	0.0008***	-0.0001
	(0.0002)	(0.0002)	(0.0002)	(0.0002)
Lag_2(Ln) Off-farm hours, spouse	0.0001	-0.0002	0.0003	-0.0003
	(0.0002)	(0.0002)	(0.0002)	(0.0002)
Hired farm labor hours	-0.0000***	-0.0000***	-0.0000***	-0.0000***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Milk yield, (*1,000 liters*)	-0.0007***	-0.0009***	-0.0006***	-0.0009***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Farming experience (*years*)	-0.0008***	-0.0010***	-0.0008***	-0.0011***
	(0.0001)	(0.0001)	(0.0001)	(0.0001)
Agricultural subsidies (*1,000 Kr*)	0.0004***	0.0004***	0.0004***	0.0003***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Investments/capital	-0.0330***	-0.0326***	-0.0337***	-0.0314***
	(0.0015)	(0.0015)	(0.0016)	(0.0015)
Total farming area (*hectares*)	-0.0004***	-0.0000	-0.0005***	0.0001
	(0.0000)	(0.0000)	(0.0000)	(0.0001)
Share of rented land	-0.0344***	-0.0486***	-0.0335***	-0.0526***
	(0.0016)	(0.0020)	(0.0015)	(0.0023)
Share of fallows land	-0.0334***	-0.0369***	-0.0340***	-0.0402***
	(0.0019)	(0.0020)	(0.0020)	(0.0023)
Livestock, units (AEU)	-0.0010***	-0.0010***	-0.0010***	-0.0010***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Operator age	0.0033***	0.0033***	0.0036***	0.0038***
	(0.0003)	(0.0003)	(0.0003)	(0.0004)
Operator age squared	-0.0000***	-0.0000	-0.0000***	-0.0000
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Location-Highland	-0.0270***	-0.0315***	-0.0315***	-0.0312***
	(0.0016)	(0.0035)	(0.0014)	(0.0045)
Metro 1 area with 50,000 population	0.0099***	0.0060	0.0097***	0.0041
	(0.0013)	(0.0125)	(0.0011)	(0.0188)
Pr(higher education) of operator	-0.1491***	-0.1482***	-0.1566***	-0.1533***
	(0.0039)	(0.0039)	(0.0040)	(0.0044)
Pr(higher education) of spouse	-0.0918***	-0.0909***	-0.0952***	-0.0884***
	(0.0030)	(0.0029)	(0.0030)	(0.0031)
Pr(ag. education) of operator	-0.2708***	-0.2846***	-0.2794***	-0.2944***
	(0.0056)	(0.0054)	(0.0058)	(0.0057)
Constant	0.8282***	0.9558***	0.8109***	1.0639***
	(0.0089)	(0.0101)	(0.0094)	(0.0104)
Observations	14,022	14,022	14,022	12,232

Baltagi-Wu LBI			1.8950	1.8950
Sigma_u	0.0186	0.0394	0.0143	0.0396
Sigma_v	0.0174	0.0174	0.0199	0.0173
R^2_o	0.9335	0.8921	0.9379	0.8560
R^2_w	0.8370	0.8474	0.8302	0.7545
R^2_b	0.9426	0.9001	0.9480	0.8640
Wald χ^2 (24)	1.0e+05	2828.20	1.1e+05	1364.17
p	0.000	0.000	0.000	0.000
rho	0.5336	0.8373	0.3408	0.8408
rho_ar			0.2306	0.2306

Standard errors in parentheses

* p<0.05, ** p<0.01, *** p<0.001