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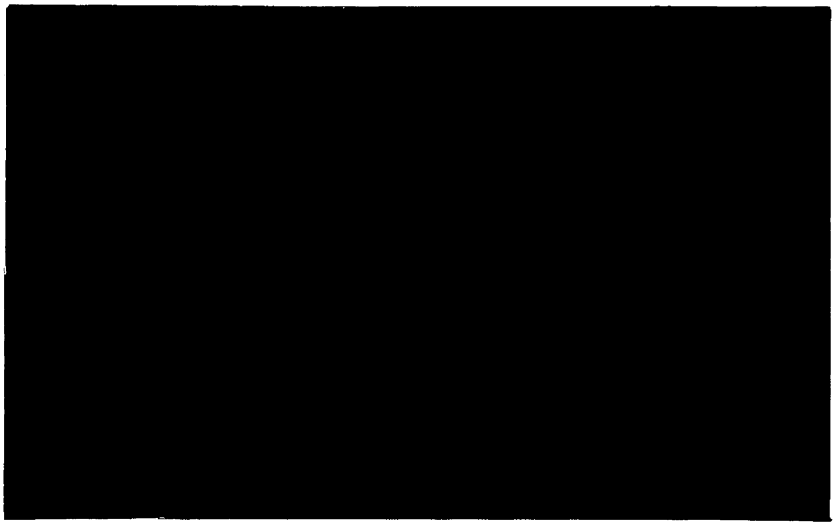

ESTIMATING OFF-FARM WORK PARTICIPATION EQUATIONS OF FARMERS CONTROLLING FOR FARM WORK PARTICIPATION STATUS: AN ENDOGENOUS SWITCHING REGRESSION APPLICATION

by

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ABSTRACT

This article claims that coefficients of farmers' off-farm work participation equations might be estimated inconsistently if selectivity based on farm participation is ignored. Participation equations will be different for farm residents who do not work on farm, especially because their reservation wages are independent of farm attributes. We estimate the off-farm participation model separately for those who work on farm and those who don't, correcting for selectivity bias, using Israeli data. We reject the hypothesis of insignificant selection bias, and the hypothesis of equal coefficients in the two subsamples.

Estimating Off-Farm Work Participation Equations of Farmers Controlling for Farm Work Participation Status: an Endogenous Switching Regression Application

1. INTRODUCTION

Empirical off-farm work participation equations of farmers are often conditioned on farm attributes. This is because farm attributes affect the marginal value of labor on the farm, which is the relevant reservation wage for off-farm work decisions of farmers. The studies of Godwin and Marlowe (1990), Lass, Findeis and Hallberg (1989), Simpson and Kapitani (1983), and Sumner (1982) follow this line, among others.

However, in most of these and other studies, farmers are indeed a self-selected group from the population of farm residents: those who chose to work on the farm. This fact may indicate that the estimated coefficients of the off-farm participation equation are inconsistent because of selection bias. For farm residents who wouldn't supply labor to the farm even if off-farm work was not available, farm attributes should not be included in the off-farm participation equation. These include farm residents who have other family members working on the farm, and those whose farm is highly non-profitable and hold it for other reasons.

In this paper we suggest an alternative approach. We estimate off-farm work participation equations separately for those who work on farm and for those who don't.

The sample separation is defined by the farm work participation dummy variable. We recognize that farm work participation and off-farm work participation are jointly determined. Therefore, we use an endogenous switching regression model in order to correct the bias caused by using an endogenous variable as a sample selection criterion.

We estimate the model for a sample of Israeli farm residents, using a two-stage estimation strategy described by Kimhi (1991a). We test three hypotheses in order to establish the advantage of our model over the commonly used approach of estimating a single off-farm participation equation using the farmers sample only. The first is the hypothesis that selection bias is unimportant for the farm workers' off-farm participation decision. The second is that the coefficients of the participation equations in the two subsamples are equal. The third hypothesis is that the coefficients of farm attributes are zero in the off-farm participation equation of those who don't work on the farm. Based on our theory, we expect to reject the first two hypotheses but not the third.

Section 2 of this paper describes the theoretical framework that leads to the off-farm participation equations and their dependence on farm attributes. Section 3 develops the empirical model and the two-stage estimation procedure used in this analysis. The data is described in section 4, and the results of estimating the empirical model and testing the hypotheses are presented in section 5. Section 6 concludes the paper.

2. THEORETICAL MODEL

The model that is used in this paper (as well as in most other studies) assumes utility maximization over consumption and leisure subject to time and budget constraints (Kimhi 1991b). Farm residents can spend time, other than home time, in farm and/or off-farm work. Formally, the optimization problem is:

$$\max_{Th,C,Tf,Tm} \quad U(Th,C;Z)$$

$$\text{s.t.} \quad 1.\ C \leq \pi(P;K,Tf) + W \cdot Tm + I$$

$$2.\ Th + Tf + Tm \leq T$$

$$3.\ Tf \geq 0$$

$$4.\ Tm \geq 0,$$

where *Th*, *Tf* and *Tm* are time spent on home activities, farm work and off-farm work, respectively, *C* is consumption, *Z* is a vector of taste shifters, *I* is non-earned income, and *W* is the off-farm wage rate. π is Lopez' (1984) conditional variable profit function. It describes farm profits as a function of market prices (*P*), conditional on farm fixed inputs (*K*, which include time inputs of other family members), and own farm labor input. The return to own farm labor ($\partial\pi/\partial Tf$) is assumed to be declining in *Tf*, while off-farm wages are assumed to be independent of *Tm*. These are standard postulates in studies of farmers' off-farm participation (e.g. Sumner 1982).

We characterize the optimal solution by the Kuhn-Tucker conditions, which are the first order conditions for maximizing the function

$$U(Th,C;Z) + \lambda[\pi(P;K,Tf)+W\cdot Tm+I-C] + \mu[T-Tf-Tm-Th] + \phi\cdot Tf + \delta\cdot Tm \qquad (1)$$

over $\{C,Tf,Tm,Th\}$ and minimizing it over $\{\lambda,\mu,\phi,\delta\}$. Given that the first derivatives of utility go to infinity when the respective arguments approach zero, the Kuhn-Tucker conditions are:

$$U_1 - \mu = 0 \qquad (2)$$

$$U_2 - \lambda = 0 \qquad (3)$$

$$\lambda\,\partial\pi/\partial Tf - \mu + \phi = 0 \qquad (4)$$

$$\lambda\cdot W - \mu + \delta = 0 \qquad (5)$$

$$\phi \geq 0 \;\; ; \;\; Tf \geq 0 \;\; ; \;\; \phi\cdot Tf = 0 \qquad (6)$$

$$\delta \geq 0 \;\; ; \;\; Tm \geq 0 \;\; ; \;\; \delta\cdot Tm = 0 \qquad (7)$$

$$\pi(P;K,Tf) + W\cdot Tm + I - C = 0 \qquad (8)$$

$$T - Th - Tf - Tm = 0, \qquad (9)$$

where U_1 and U_2 are the partial derivatives of utility with respect to home time and consumption, respectively.

Using (2) and (3), we can write (4) and (5) as:

$$\partial\pi/\partial Tf + \phi/U_2 = U_1/U_2 \tag{4'}$$

$$W + \delta/U_2 = U_1/U_2, \tag{5'}$$

where ϕ and δ are positive if and only if farm work and off-farm work, respectively, are zero. The right hand side of (5)' can serve as a reservation wage for off-farm work participation, when it is conditioned on $Tm=0$. Tf^*, optimal farm labor supply conditioned on $Tm=0$, can be derived from (4)' as the solution to

$$\partial\pi(P;K,Tf)/\partial Tf = U_1(T-Tf,\pi(P;K,Tf)+I;Z)/U_2(T-Tf,\pi(P;K,Tf)+I;Z) \tag{10}$$

as long as $Tf^*>0$. In this case, (10) implies that

$$Tf^* = Tf^*(P,K,T,I,Z), \tag{11}$$

and the reservation wage U_1/U_2 can be written as $R(P,K,T,I,Z)$. This is a function of farm attributes which are included in K. However, when $Tf^*=0$, the reservation wage is $R'(T,I,Z) = U_1(T,I;Z)/U_2(T,I;Z)$, assuming that $\pi(P;K,0)=0$. R' does not depend on farm attributes, and its functional form is different from that of R. Hence, the participation model based on (5)' should be formalized such that off-farm work participation occurs if $W>R^*$, where

$$R^* = \begin{cases} R(P,K,T,I,Z) & \text{if} \quad Tf^*>0 \\ R'(T,I,Z) & \text{if} \quad Tf^*=0. \end{cases} \tag{12}$$

3. EMPIRICAL MODEL AND ESTIMATION PROCEDURE

The theory above leads to the following empirical model. Let $Y^* = W\text{-}R^*$ be a latent variable describing the tendency to participate in off-farm work, based on (5)' and (12), where W is the off-farm wage rate and R^* is the relevant reservation wage. Let Y be the observed participation index, i.e.:

$$Y = \begin{cases} 1 & \text{if} \quad Y^* > 0 \\ 0 & \text{otherwise.} \end{cases} \tag{13}$$

Specifying Y^* as a linear function of observable variables and a stochastic component, $Y^*=X\beta+u$, where u is identically and independently distributed standard normal random variable, the coefficients β can be estimated by probit Maximum Likelihood. This means maximizing the expression $Y \cdot Prob(u>-X\beta) + (1\text{-}Y) \cdot Prob(u\leq -X\beta)$ over β.

To the extent that $X\beta$ is a first order approximation of $W\text{-}R^*$ and u is the approximation error, it follows from (12) that

$$Y^* = \begin{cases} X_1\beta_1+u_1 & \text{if} \quad Tf^*>0 \\ X_2\beta_2+u_2 & \text{if} \quad Tf^*=0, \end{cases} \tag{14}$$

where X_1 includes $\{P,K,T,I,Z\}$ and X_2 includes $\{T,I,Z\}$. Thus, the model has to be estimated separately in each of the subsamples, defined by $Tf^*>0$ and $Tf^*=0$, respectively.

If u is normally distributed, u_1 and u_2 are truncated normal random variables, since there is no reason to believe that u and Tf^* are independent. Hence, u_1 and u_2 have to be transformed to standard normal before the separate probit equations can be estimated. Clearly, (14) resembles an endogenous switching regression model (Maddala 1983, p. 223). Next, we adjust this model to the fact that all dependent variables are discrete, and use a two-stage estimation procedure (Kimhi 1991a) rather than maximum likelihood in order to save computation time.

We start by formalizing the farm labor supply function as:

$$Tf^* = X_f \beta_f + u_f, \tag{15}$$

where X_f is a row vector of explanatory variables, β_f is a conformable column vector of associated parameters, and u_f is a standard normal, possibly correlated with u (but independent across individuals), random variable.

We continue by concentrating on the subsample of those who work on the farm. Similar results can be derived for the other subsample. We can write (14) for this subsample as:

$$Y^* = X_1 \beta_1 + \rho_1 E_1 + \varepsilon_1, \tag{16}$$

where $E_1 \equiv E(u_1/Tf^*>0) = \phi(-X_f\beta_f)/[1-\Phi(-X_f\beta_f)]$, ρ_1 is the correlation coefficient between

u_l and u_f, ϕ is the standard normal density function, and Φ is its cumulative distribution function. One can show that (Johnson and Kotz 1972, p. 112):

$$E(\varepsilon_l/Tf^*>0) = 0 \tag{17}$$

$$V_l \equiv Var(\varepsilon_l/Tf^*>0) = 1 + \rho_l^2 E_l (E_l - X_f \beta_f). \tag{18}$$

The two-stage estimation procedure uses (15), (16) and (18) as follows. First we estimate (15) to get consistent estimates of β_f and hence of E_l. We use these estimates in (16) and (18), and divide (16) by the square root of V_l as defined in (18). As a result, $\varepsilon_l/V_l^{1/2}$ is conditionally standard normal. Second, we estimate the resulting equation by probit to get consistent estimators of $\beta_l/V_l^{1/2}$ and $\rho_l/V_l^{1/2}$, from which β_l and ρ_l can be identified. Finally, we calculate the correct standard errors of the estimators by the method suggested by Murphy and Topel (1985).

4. DATA

We use data from the 1981 Census of Agriculture in Israel. Originally, it included 28526 observations of farms in moshavim (these are semi-cooperative villages, consisting of privately- operated family farms; see Kimhi 1991b). We eliminated those who explicitly defined themselves as "non-farming families" (6281), "private" (as opposed to "family") farms (2808), and partnerships (341). These are all exceptional types of farms and we excluded them in order to minimize unnecessary noise. Landless families and incomplete observations were also excluded.

The final data set includes 16818 observations, and its descriptive statistics appear in table 1. Farm work and off-farm work are reported in qualitative terms. For each of the two sectors, the respondent had to indicate whether he/she is working up to 1/3 of the time in that sector, up to 2/3 of the time, full time, or not at all. The measures of 1/3, 2/3 and 'full time' were left for the respondent's discretion. The only limitation was that a respondent cannot report full time work on the farm *and* full time work off the farm.

Farm attributes included land size broken down by crops and by irrigation status, and livestock by type. Normative values of sales and of value added for each type of farm output were also included (physical quantities were multiplied by average values calculated from more detailed surveys). Normative values of capital assets were reported by type of assets and by type of product for which they were used (see the note above).

The problem with including such farm attributes as explanatory variables in an off-farm participation equation is that some or all of them may be endogenous. This is true to the extent that off-farm participation and levels of farming activities are determined jointly. The literature has been mixed about this point (Lass, Findeis and Hallberg 1991). While several researchers used many farm attributes (e.g. Lass et al. 1989), others did not use them at all (e.g. Huffman and Lange 1989). We follow the approach of Sumner (1982), and try to use those farm attributes which are more likely to be exogenous to the off-farm participation decision (it is important to include at least several farm attributes since input and output price data is not available).

In particular, we use the farm's original land allotment, the value of capital assets which were purchased or built at least ten years prior to the survey, and whether the farm includes a dairy operation. The land variable is appropriate because of the unique institutional arrangements governing the behavior of these farm residents (Kimhi 1991b). Specifically, land was equally distributed among residents in each village at the time of establishment of the village. Farm owners are not allowed to buy or sell land (with the exception of selling the whole farm and moving out of the village). Land rentals are illegal, although short time rentals exist in practice. However, land rentals are not included in the land variable used in this analysis. Therefore, current time allocation decisions cannot affect the land variable. Obviously, the same is true for the old capital stock variable (which is highly correlated with total capital stock).

The dairy dummy was chosen because of two reasons. First, labor requirements of a dairy operation are much different from those of other farm activities. Hence, dairy farming activity information is important for studies of off-farm participation more than information on other farm activities. Second, entries into and exits from dairy farming have been relatively rare in Israel. This was in part because of agricultural policy (milk production has been heavily subsidized over the years, and hence was subject to strict quotas). Also, dairy farming involves prohibitively large capital investments, especially for farmers without sufficient collateral for raising debt (land is nationally owned).

A problem with applying the empirical model (section 2) to this data is that our sample separation criterion is based upon Tf^*, which is the optimal level of farm work conditional on not working off the farm. This is not observed. Of course, for those who don't work off the farm, Tf^* is equal to Tf, observed farm labor supply level. But for those who work off the farm, it is likely that $Tf^* > Tf$.

We try to diminish the impact of this problem in several ways. First, we estimate (15) using farm work participation information only, rather than using the complete labor supply information. This results in a probit model, which yields a consistent estimator of β_f. It is conceivable that the noise involved in using Tf instead of Tf^* is reduced when only using the information about Tf crossing a threshold. Second, we also try to estimate β_f using only those who don't work off the farm rather than the whole sample. This estimator may be closer to the true parameter vector of the Tf^* function.

Finally, note that actual farm participation is a conservative measure for Tf^* crossing a threshold, in the sense that it is harder to reject the hypothesis $\beta_1=\beta_2$ and it is easier to reject the hypothesis that the coefficients of farm attributes included in X_2 equal zero. This is because the subsample used for estimating β_2 includes observations that really belong to the other subsample (in our sample, 10% of farm operators report $Tf=0$, while 5% report $Tf=0$ and $Tm=0$. The true size of the subsample for which $Tf^*=0$ is between 5% and 10%, whereas we use all the 10%). Hence, if we are able to reject the second hypothesis but not the third one, as we expect according to the theory, then we can be confident that we would have had the same conclusions had we used data on Tf^*.

5. RESULTS

We estimated the farm work participation equation using the whole sample and alternatively, only those who don't work off the farm. While the results were somewhat different, the impact of the difference on the second stage results was minimal. Hence, we only report the results for the case in which β_f is estimated using the whole sample.

We estimated the coefficients β_1 and β_2 of (14) using the procedure described at the end of section 3. Each equation was estimate twice: once with X_1 (including farm attributes) and once with X_2 (excluding them). This way we can test the hypothesis that $\beta_1=\beta_2$. The fact that $X_2 \subset X_1$ enables us to test the significance of farm attributes in each equation.

The nonlinear probit equations based on (16) and (18) were estimated using the Gauss program (copies of the data and the program are available upon request). The results are in table 2. For two thirds of the farm workers subsample, off-farm participation was correctly predicted (a correct prediction means that the probability of participation is greater than one half for those who participate and less than one half for those who don't). The percent of correct predictions was close to 80% in the other subsample. The correction for sample selection bias was important only in the subsample of those who work on the farm. Therefore, the first hypothesis is rejected, and the conclusion is that ignoring the selection into farm work results in inconsistent estimation of off-farm participation equations.

The second hypothesis that we test is that the coefficients are equal in the two subsamples: $\beta_1=\beta_2$. If this hypothesis is not rejected, our method has no value added over the traditional one for this sample. It is clear from table 2 (first and third columns) that $\beta_1\neq\beta_2$, and a formal likelihood ratio test rejects the hypothesis in all reasonable significance levels (likelihood ratio statistic of 500 with 17 degrees of freedom). The conclusion is that pooling the two subsamples and estimating a single off-farm work participation equation over the whole sample results in incorrect estimators.

The third hypothesis is that the coefficients of farm attributes are zero in the subsample of those who don't work on the farm. If this hypothesis is rejected, then there is scope for estimating separate off-farm participation equations, but probably for a reason other than the one we stated. The likelihood ratio statistic for this hypothesis is 9.32, with 3 degrees of freedom. Therefore, the hypothesis can be rejected at the 5% significance level, but not at the 1% level (p-value $\cong$ 2.5%). However, comparing the last two columns of table 2, we see that the change in the other coefficients after excluding farm attributes is marginal, and so is the drop in the percent of correct predictions. On the other hand, farm attributes are highly significant in the equation of those who work on the farm (likelihood ratio statistic of 610 for a similar exclusion hypothesis).

Personal characteristics have the expected effects on off-farm participation (Lass et al. 1991). Note that these are reduced form effects, in the sense that the coefficients measure the effects of the explanatory variables on the difference between the off-farm wage rate and the reservation wage. Off-farm participation is first increasing with age and then decreasing. It peaks around the age of 43 for those who work on farm and around

25 for those who don't. Schooling and years in Israel increase off-farm work participation, more so for those who don't work on the farm. Both are measures of general human capital. Years on the farm decrease off-farm work participation only for those who don't work on the farm. We would have expected this effect to be stronger in the other subsample. However, farm work in the past is not known, and perhaps those who don't work on the farm had done so in the past. The equations included also dummy variables for ethnic origin (not reported in table 2). Among those who work on farm, those born in foreign countries were more likely to work off the farm than native Israelis. The origin dummies were not significant in the subsample of those who don't work on the farm.

The number of family members in different age groups increase off-farm participation, though only some of the effects are statistically significant. This is probably because other family members can substitute for operator's labor both on the farm and in house work. All three farm attributes included (land holdings, old capital stock and the dairy dummy) affect off-farm participation negatively. This is plausible since they all contribute to farm labor productivity.

There can be several explanations for our inability to strongly establish that farm attributes are not relevant to off-farm work participation decisions of those who don't work on the farm. The most plausible one is, of course, the fact that we used the group of farmers for whom $Tf=0$ rather than those for whom $Tf^*=0$. As said before, this means that off-farm work participation decisions of some of the farmers in this subsample are really affected by farm attributes, and that is why the hypothesis was marginally rejected.

Another explanation relates to our assumption that $\pi(P;K,0)=0$. This means that farm income is unimportant for economic decisions of those who don't work on the farm. However, if other family members are working on farm, and their farm labor supply (and hence farm profits) is affected by farm attributes, then the reservation wage of the farm operator will depend to some extent on farm attributes. In this case, the dependence reflects the income effect only. Therefore, we still expect the coefficients of the participation equations, and especially those of farm attributes, to be different in the two subsamples, as we really observe.

The analysis also ignores an important aspect of farm-household behavior, namely family-level decision making. Huffman and Lange (1989) and others have recently shown that this is relevant to off-farm participation and labor supply decisions. According to this argument, off-farm participation of the farm operator is determined jointly with farm work decisions of other family members, even if he doesn't work on the farm. As a result, it will be affected by farm attributes. There is some support for this view in our results: in both subsamples, when excluding farm attributes from the off-farm participation equations, the coefficient that changes the most is that of the number of prime-age family members. An application of this extension is left for future research.

6. SUMMARY AND CONCLUSIONS

In this article we claim that estimating off-farm work participation equations of farmers may produce biased estimates if it does not control for selection on the basis of farm work participation. We argue that reservation wages of those who wouldn't work on the farm even if off-farm employment was not available, and of those who would, have different functional forms. In particular, the former group's reservation wages should not depend much on farm attributes. As a result, we suggest estimating off-farm participation equations separately for the two groups of farm operators.

We perform the separate estimation corrected for sample selection bias using Israeli data. We strongly reject the hypotheses that selection is unimportant and that the off-farm participation equations' coefficients are equal in the two groups of farmers. We can only marginally reject the hypothesis that farm attributes don't affect off-farm participation of those who don't work on the farm. Overall, this supports our suggestion that off-farm work participation equations should be conditioned on the farm work participation dummy in empirical applications, and that the two participation decisions should be jointly analyzed.

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Table 1. Descriptive Statistics

Quantitative

	Mean	*S.D.*	*Range*	*Units*
Age	47.3	13.4	16-80	years
Years in Israel[a]	32.0	8.9	1-80	years
Years on Farm	18.7	10.9	0-61	years
Schooling	8.7	4.5	0-20	years
Family Members 0-14[b]	1.6	1.6	0-11	persons
Family Members 15-21	.89	1.3	0-8	persons
Family Members 22-65	1.5	1.1	0-9	persons
Family Members 66+	.12	.37	0-2	persons
Total Land[c]	30.2	38.2	1-3030	dunams[d]
Old Capital[e]	11.1	26.7	0-1049	$1000[f]

Qualitative

	Number	*Percent*
Dairy Farm	1315	7.8
Extent of Farm Work		
None	1744	10.4
Up to 1/3	4040	24.0
Up to 2/3	2797	16.6
Full Time	8237	49.0
Total	16818	100
Extent of Off-Farm Work		
None	10297	61.2
Up to 1/3	549	3.3
Up to 2/3	1053	6.3
Full Time	4919	29.2
Total	16818	100

[a] For native Israelis, equal to age.
[b] Number of family members in each age group, excluding operator.
[c] Original land allotment.
[d] 1 dunam = 0.23 acre.
[e] Normative value of capital assets at least ten years old.
[f] In 1981 prices. Factor of exchange: 12.39.

Table 2. Probit Off-Farm Participation Results

	Work on Farm		*Don't Work on Farm*	
Intercept	-2.305 (11.3)**	-2.392 (11.4)**	0.290 (0.26)	-0.163 (0.16)
Age	0.112 (12.2)**	0.082 (9.47)**	0.056 (1.77)*	0.050 (1.58)
$(\text{Age})^2/100$	-0.136 (14.8)**	-0.099 (11.8)**	-0.118 (4.21)**	-0.112 (4.00)**
In Israel	0.008 (4.53)**	0.004 (2.32)**	0.014 (1.96)*	0.014 (1.90)*
Years on Farm	-0.001 (0.35)	-0.006 (2.91)**	-0.012 (1.87)*	-0.014 (2.16)*
Schooling	0.038 (10.7)**	0.032 (8.42)**	0.081 (8.66)**	0.081 (8.50)**
Family 0-14	0.002 (0.24)	-0.002 (0.24)	-0.015 (0.52)	-0.009 (0.32)
Family 15-21	0.037 (3.24)**	0.033 (2.78)**	0.028 (0.95)	0.024 (0.81)
Family 22-65	-0.024 (1.40)	-0.069 (5.78)**	0.072 (1.07)	0.096 (1.93)*
Family 66+	0.023 (0.65)	-0.016 (0.44)	0.138 (1.21)	0.144 (1.28)
Total Land	-0.277 (13.5)**		-0.121 (2.28)*	
Old Capital	-0.016 (3.24)**		0.005 (0.34)	
Dairy Dummy	-0.520 (9.13)**		0.051 (0.25)	
ρ	-0.641 (3.59)**	-1.113 (35.6)**	-0.248 (0.78)	-0.103 (0.44)
No. of Cases	15074	15074	1744	1744
% Correct Pred.	66.6	64.8	79.7	79.4
Log Likelihood	-9079	-9384	-792	-797

Notes: All models included a set of ethnic origin dummies.
Asymptotic t-statistics in parenthesis.
* - significant at the 5% level.
** - significant at the 1% level.
% of correct predictions: prediction=1 (0) if the probability of Y=1 is greater (smaller) than 1/2.