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THE RELEVANCE OF THE EXTENT OF FARM WORK TO THE ANALYSIS OF OFF-FARM LABOR SUPPLY OF FARMERS

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ABSTRACT

Farm labor supply of farm operators is important for the analysis of their off-farm labor supply. We use a unique data set which includes such information to demonstrate its importance. Other studies had to use implicit assumptions in order to proxy the marginal product of farm labor with observable farm attributes. We find that these assumptions are too strong. We do it by estimating the off-farm participation equation over different subsamples defined according to the level of farm work. We correct for selection by estimating an endogenous switching regression model of the off-farm participation decision, in which the selection criterion is farm participation. Selection is found to be significant for farm workers only. The qualitative conclusions are unaffected by controlling for selection.

THE RELEVANCE OF THE EXTENT OF FARM WORK TO THE ANALYSIS OF OFF-FARM LABOR SUPPLY OF FARMERS

Estimation of off-farm labor supply functions of farmers may be seriously biased by ignoring the extent to which farmers work on the farm, and especially whether they work on the farm or not.[1] Farm and off-farm work are jointly determined, and should be jointly estimated. However, most surveys of farmers' economic behavior do not include farm labor supply information, presumably because of the objective difficulty of obtaining credible answers to questions of this type.

Economists often use farm attributes as proxies for farm work, when estimating off-farm participation equations or labor supply functions of farmers. This procedure utilizes the concept of conditional variable profit function described by Lopez. Following this line, among others, are the studies of Sumner, Huffman & Lange, Lass et al. and Tokle & Huffman. This approach is based on two implicit assumptions: (1) All "farmers" really work on their farm; (2) Farm attributes are good proxies for the marginal product of labor on the farm.

This paper suggests an alternative approach, based on the assumption that off-farm work decisions of farmers depend on their farm work decisions. One model based on this approach is the endogenous switching regression model (Maddala, p. 223), in which different behavioral equations are estimated for different subsets of the population, as well as a selection equation. This model is applied here in estimating off-farm participation equations which are conditioned on farm participation, using Israeli census data. A two-stage estimation strategy described in Kimhi (1991b) is used.

We discuss the implicit assumptions used in the literature and their implications, using the Kuhn-Tucker necessary conditions for the farmer's optimization problem. We present the alternative model and suggest empirical tests of the assumptions. We estimate separate off-farm participation models in different subsets of the data defined according to farm participation and according to the extent of farm work. Next we correct for selection and estimate the endogenous switching regression model. The final conclusion is that the extent of farm work, and especially participation, are valuable in modelling off-farm participation. However, we cannot reach a clear conclusion about the role of farm attributes as proxies for the marginal product of farm labor.

Relaxing the Assumptions Regarding Farm Work

Many studies of farmers' time allocation deal with farm operators only (e.g. Sumner, Simpson & Kapitani), and assume that they all work on farm by definition. This assumption is challenged by findings from the Israeli data set: about one out of ten farm operators is not working on his farm (Table 1). This may be due to reporting errors. However, it seems likely that the error is in the identity of the farm operator and not in the particular answer regarding farm work. This is because the farm household often includes two or more persons who are capable of answering the questionnaire. The identity of a single farm operator is not always clear, and the respondents lack an incentive to follow the formal definition. There is no reason to believe that this kind of error is specific to this data set.

The assumption is even more objectionable in studies of farm women's off-farm work (Godwin & Marlowe, Kimhi 1991b), or joint work decisions of farm operators and spouses (Huffman & Lange, Tokle & Huffman, Lass et al.). This is because specialization within the family often causes some household members not to work on farm. The effect of the assumption on empirical results is an empirical question itself, which we intend to examine here.

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 1991a). Farmers can spend time in farm and/or off-farm work. Formally, the optimization problem is:

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

$$\begin{aligned} \text{s.t.} \quad & 1.\quad C \le \pi(Tf) + W\cdot Tm + I \\ & 2.\quad Th + Tf + Tm \le T \\ & 3.\quad Tf \ge 0 \\ & 4.\quad Tm \ge 0 \end{aligned}$$

where *Th*, *Tf* and *Tm* are time spent on home activities, farm work and off-farm work, respectively, C is consumption, I is non-earned income, W is the off-farm wage and π is Lopez's conditional variable profit function.

Two of the Kuhn-Tucker necessary conditions for this maximization problem (Waldman) are:

$$(1) \qquad \pi_1 + \delta/U_2 = U_1/U_2$$

$$(2) \qquad W + \phi/U_2 = U_1/U_2$$

where δ and ϕ are positive if and only if farm work and off-farm work, respectively, are zero, and subscripts denote partial derivatives. (2) implies that off-farm participation occurs if:

$$(3) \qquad W > U_1(I+\pi(Tf^*),T-Tf^*)/U_2(I+\pi(Tf^*),T-Tf^*)$$

assuming that the second order conditions are met, where Tf^* denotes optimal farm labor supply given no off-farm work.

This leads to the following participation index function:

$$(4) \qquad Y^* = W - RHS$$

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

where RHS is the right hand side of (3). It is clear that when specifying Y^* as a function of observable variables, this function will depend on Tf^*. In practice Tf^* is not observed and researchers use farm attributes as proxies. This is right under two conditions:

(1) farm attributes are not endogenous; and (2) if farmers were not allowed to work off-farm, they would all work on their farms. We want to concentrate on the second condition, assuming that the first one holds. In our data set, about half of the farm operators who don't work on the farm, don't work off-farm either. For these farmers, a constraint of not working off-farm would not be binding, and their behavior would not change. Al least for these farm operators, $Tf^{*}=0$ and farm attributes should not be included in the off-farm participation equation.[2] In our sample, therefore, there are between 5 and 10 percent of farmers whose reservation wages are not affected by farm attributes (the number is much larger for farm spouses). The question whether this affects empirical results is itself empirical.

We assume that the set of farmers who don't work on the farm is identical to the set of farmers for which $Tf^{*}=0$ (which is better than assuming that the latter is null). Under this assumption, we can test our conclusion by estimating the off-farm participation equation (4) separately for those who work on the farm and those who don't, and test the hypothesis that the two sets of parameters are equal. In particular, the model predicts that the coefficients of farm attributes will be zero in the non-participants equation, and this can also be tested. These tests will be performed in the following sections.

A second assumption that we want to challenge relates to the validity of farm attributes as proxies for the marginal product of labor on the farm. Since researchers assume that $Tf^{*}>0$ for all farmers, they can use (1) to write the participation equation (4) as $Y^{*}=W-\pi'(Tf^{*})$. They further use a set of farm attributes to proxy for $\pi'(\cdot)$. These proxies are not valid if farm labor supply is not sufficiently flexible and free of short run constraints in the time unit used for modelling off-farm labor supply decisions. Farm production activities stretch over relatively long periods of time, from the decision to engage in a certain activity to the realization of proceeds. During that period, farm labor supply is to some extent a fixed obligation, that has to be fulfilled even when short run considerations favor allocation of time to other activities such as off-farm work.[3] In the extreme case in which farm work is exogenous to the off-farm labor supply decision, we have to include it in the set of explanatory variables (Oliveira). Otherwise, there is an omitted variable problem.

Actual farm labor supply constitutes a combination of fixed time obligations and a variable component. Neither component can be isolated in the data. The fixed component solves the omitted variable problem when including farm work as an explanatory variable, but the variable component is probably correlated with the stochastic component of the off-farm participation equation. The choice between including farm work as an explanatory variable or not is indeed between two second-best solutions.

In the Israeli data set, farmers reported farm work as a four-level ordered qualitative variable. Hence, it is conceivable that it mostly reflects the fixed component, and not as much the stochastic element of the variable component of farm labor supply.

Therefore, using this measure in the estimation of farmers' off-farm participation could improve the quality of the results. We test this conclusion in two ways. First, we divide the data set into subsamples according to the extent of farm work, estimate the off-farm participation equation in each subsample and test the hypothesis that all the sets of parameters are equal. Second, we include dummy variables for the extent of farm work in the set of explanatory variables, estimate the model over the whole sample, and check the significance of these dummies and the effect of their inclusion on other coefficients.

Data and Empirical Results

We use data from the 1981 Census of Agriculture in Israel. Originally, it included 28526 observations of farms in moshavim (Kimhi 1991a). We eliminated those who explicitly defined themselves as "non-farming families" (6281), "private" (as opposed to "family") farms (2808), partnerships (341), landless families, and incomplete observations. The final data set includes 16818 observations, and its descriptive statistics appear in table 1.

Figure 1 illustrates the relation between farm and off-farm labor supply in the raw data. We observe 50% off-farm participation among those who don't report any farm work, while only 37% of farm workers work off the farm. The variation by the extent of farm work is more dramatic. While more than 70% of those who work part-time than 70% of those who work part-time on the farm participate in off-farm employment, only 6% of full-time farm workers do so.[4]

We now turn to the econometric model. Let Y^* of equation (4) be specified as a linear function of personal, family and farm variables, plus an i.i.d. standard normal stochastic component. In this case, we can estimate the parameters of Y^* by probit. We do it separately in the different subsamples defined according to the extent of farm work. The results are summarized in Table 2a.

Comparing the results for the farm workers and non-workers subsamples, we can reject the hypothesis of equal coefficients in very low significance levels. In particular, the coefficients of farm attributes are extremely different. While all of them have significantly negative coefficients in the workers' equation, only the land coefficient is significant in the equation of non-workers, and its magnitude is much smaller. The coefficients of capital stock and the dairy farm dummy are positive and non-significant.[5]

Moreover, excluding farm attributes as a group doesn't have any considerable effect on the other coefficients in the farm-nonworkers subsample, and we cannot reject the joint hypothesis that the coefficients of farm attributes equal zero at the 1% significance level. On the other hand, excluding them from the model using the whole sample does have an effect, especially on family variables (the formal exclusion hypothesis is rejected at very low significance levels). The prediction power of the model in the non-workers sample is not affected by the exclusion of farm attributes.

Other coefficients differ as well. Age profiles are much more concave in the workers' subsample, while the effect of schooling is much stronger in the non-workers subsample. The effects of adult family members is also much stronger in the non-workers' sample.

The results are clearly in support of the idea that reservation wages of farm workers and non-workers have different functional forms. The difference is especially notable in the coefficients of farm attributes. The separation of the sample according to farm work participation improves the prediction power of the model, as measured by the percent of correct predictions. The coefficients of the whole sample are significantly different from those of the farm workers sample. The major contribution to the difference comes from the number of adult family members. Therefore, imposing equal coefficients by using one probit equation for the whole sample results in inconsistent estimators.

Next, we examine the changes in the results when the sample is divided according to the positive levels of farm work. Figure 1 indicates that the probability of working off-farm is decreasing with the extent of farm work, and that the extent of off-farm work is decreasing in the extent of farm work. This is not surprising, since it is a direct implication of the binding time constraint.

We estimate (4) separately in each subsample defined according to the extent of farm work (table 2b), and test the hypothesis of equal coefficients across subsamples. The hypothesis is rejected in all reasonable significance levels, and the percent of correct predictions is remarkably higher in the three separate subsamples (average 88.8%) than in the combined farm workers sample (66.5%).

Looking at single coefficients across subsamples, we observe several noticeable trends. For example, the effect of the number of family members on off-farm participation is positive for the 1/3 group, still positive but smaller and non-significant for the 2/3 group, and becomes negative for full-time farmers. The schooling coefficient is always positive and significant, but decreases as the extent of farm work increases. The effect of capital stock is negligible for the 1/3 group, slightly positive for the 2/3 group, and significantly positive for full-time farmers.

Finally, we add dummy variables representing the extent of farm work to the model and estimate it using the whole sample. We find that off-farm participation probability is higher (lower) among those who work part-time (full-time) on the farm than among those who don't work on farm (See note 4). Among those who work part-time on the farm, off-farm participation decreases with the extent of farm work. This is consistent with figure 1. Farm attributes become smaller in absolute value and less significant, after controlling for the extent of farm work, but the hypothesis that they should be excluded is rejected at the 1% level. However, the coefficients of the farm work dummies and the other explanatory variables do not change after the exclusion of farm attributes. In terms of prediction power, the inclusion of the farm work dummies improves the performance of the model tremendously. The exclusion of farm attributes does not have an effect on the percent of correct predictions, in the presence of the farm work dummies.

In the last two exercises, we checked the performance of farm attributes as proxies for the marginal product of farm labor. First we divided the sample according to the extent of farm work, and then we included dummies representing the extent of farm work as explanatory variables. In both cases, the hypothesis that the change is unimportant was definitely rejected, the percent of correct predictions increased substantially, and the importance of farm attributes diminished considerably. The conclusion is that farm attributes are not good enough as proxies. However, the set of farm attributes remained statistically significant.

Controlling for Selection

In the previous section, we performed probit estimation on different subsets of the data without worrying about selection biases. In this section, we correct the selection bias caused by the possible endogeneity of the farm participation decision, and check the validity of our conclusions. We estimate an endogenous switching regression model (Maddala p. 223), adjusted to the fact that all dependent variables are discrete. We use a two-stage procedure in order to save computation time (Kimhi 1991b).

Formally, we can write the model that emerges from (4) as:

$$(5) \qquad Ym^* = \begin{cases} X_1 \cdot \beta_1 + u_1 & if\ Yf^* > 0 \\ X_2 \cdot \beta_2 + u_2 & otherwise \end{cases}$$

where Ym^* is off-farm labor supply, Yf^* is farm labor supply, and:

$$(6) \qquad Yf^* = X_f \cdot \beta_f + u_f.$$

In (5) and (6), X_i are row vectors of explanatory variables, β_i are conformable column vectors of associated parameters, and u_i are standard normal, possibly correlated (but independent across individuals), random variables. Estimating (5) using the subsample in which $Yf^* > 0$, as we did in the previous section, results in inconsistent estimators if the conditional expectation of u_1 is non-zero. Using the derivations of Johnson & Kotz (p. 112), we write (similar results can be derived for the other subsample):

$$(7) \qquad Ym^* = X_1 \cdot \beta_1 + \rho_1 \cdot E_1 + \varepsilon_1.$$

Where $E_1 \equiv E(u_1 | Yf^*>0) = \phi(-X_f \cdot \beta_f)/[1- \Phi(-X_f \cdot \beta_f)]$ and ρ_1 is the correlation coefficient between u_1 and u_f. One can show that:

$$(8) \qquad E(\varepsilon_1 | Yf^*>0) = 0$$

$$(9) \qquad V_1 \equiv Var(\varepsilon_1 | Yf^*>0) = 1 + \rho_1^2 \cdot E_1 \cdot (E_1 - X_f \cdot \beta_f).$$

The two-stage estimation strategy uses (6) and (7). First we estimate (6) by probit to get consistent estimates of β_f and therefore of E_1. We use these estimates in (7) and (9), and divide (7) by the square root of (9). Second, we estimate the resulting equation by probit to get consistent estimators of $\rho_1/V_1{}^{1/2}$ and

$\beta_1/V_1^{1/2}$, from which β_1 and ρ_1 can be identified. Finally, we calculate the correct standard errors of the estimators by the method suggested by Murphy & Topel. In Table 3 we compare the results of this estimation method to the previous results (those not corrected for selectivity). This enables us to test the hypothesis that selection bias is not important in this problem.

The comparison yields very different conclusions regarding the two subsamples. While for farm workers the correlation coefficient is -0.64 and highly significant, it is -0.25 and not significant in the farm nonworkers subsample, meaning that selection bias is not a problem in the latter subsample. In both subsamples, the only coefficient that changes qualitatively after correcting for selection bias, is the one related to the number of family members between the ages of 22 and 65. It changes signs from negative to positive in the workers subsample, and it increases substantially and becomes significant in the non-workers subsample. In the workers subsample, the coefficients of farm attributes are somewhat smaller in absolute value after correcting for selection bias. Overall, the qualitative conclusions regarding the importance of farm participation are not affected by selectivity considerations, and our previous discussion based on the raw results is valid.

In a similar way, we can correct the selectivity bias in the model of those who work full-time on the farm. We estimate the farm labor supply equation (6) by ordered probit, using the qualitative farm work information. We assume that there exist unknown thresholds d_1 and d_2 such that:

farmer doesn't work on the farm iff: $Yf^* < 0$;

farmer works up to 1/3 on farm iff: $0 \leq Yf^* < d_1$;

farmer works up to 2/3 on farm iff: $d_1 \leq Yf^* < d_2$;

farmer works full-time on farm iff: $d_2 \leq Yf^*$.

Using the ordered probit estimates, we estimate a model similar to (7), using the subsample of farmers who work full-time on the farm, in which $(d_2-X_f \cdot \beta_f)$ substitutes for $(-X_f \cdot \beta_f)$.

The results are in the last column of table 2b. We see that the correlation coefficient is not significant, so that selection bias is not important in that model. The percent of correct predictions is unchanged. The importance of farm attributes is somewhat diminished after correcting for selection bias, and this supports our view that farm attributes are not affecting off-farm participation if one properly controls for farm work variations.[6]

Summary and Conclusions

This paper is challenging two assumptions that are implicit in many analyses of farmers' off-farm work decisions: that all farmers work on their farm, and that farm attributes are proper proxies for the marginal product of farm labor. These assumptions stem from the lack of farm labor supply information. We are able to relax them by using a data set that includes such information.

Separating those who work on farm from those who don't, we find significant differences in the coefficients of the two groups' off-farm participation equations. The separation improves the prediction power of the model. Farm attributes are not significant in the non-workers equation. This supports the view that the distinction between farm workers and non-workers is important, even when the latter group is relatively small, and that ignoring this distinction is likely to result in inconsistent parameter estimates. These conclusions are unchanged after correcting the selection biases caused by estimating the model on subsets of data.

Estimating the off-farm participation model in separate subsamples defined according to the extent of farm work, we find that the coefficients are significantly different across subsamples. The prediction power of the model increases. The coefficients of farm attributes as a group remain significant, though. When we included dummies for the extent of farm work as explanatory variables using the whole sample, the farm attributes coefficients became smaller in absolute value but remained significant. The exclusion of farm attributes did not change the other parameters, though, and did not affect the percent of correct predictions. We conclude that the extent of farm work is important to the explanation of off-farm participation, but farm attributes are not just proxies for it; they have some explanatory power even after controlling for the extent of farm work.

Overall, this paper has demonstrated the importance of having direct information about farm work to the analysis of off-farm participation. We found that the distinction between those who work on the farm and those who don't is extremely important, especially regarding the treatment of farm attributes. Controlling for the level of farm work is also important and improves the prediction power of the model. However, the results have been somewhat inconclusive regarding the role of farm attributes as proxies.[7]

Notes

1. One may say that someone who doesn't work on the farm is not a farmer. We use here a broader definition, which includes those who own farms and live there.

2. Farm attributes appear in the reservation wage of these individuals only as a result of the joint family budget constraint when other family members work on farm. Even in this case, the coefficients of farm attributes will be different.

3. It is recognized that off-farm work also involves long-run commitments, but most studies ignore this as well. In the Israeli case, it is likely that farm work is more constrained over time than in other developed economies, because of the institutional restrictions on resource transactions (Kimhi 1991a).

4. The fact that those who work part-time on the farm have a higher tendency to work off-farm than those who don't work on the farm at all might be an indication of a two-stage decision process: first, the farmer decides whether we works at all or not, and then he decides on the optimal time allocation between farming and off-farm work. This possibility is not addressed here.

5. Land and to some extent the dairy dummy are exogenous even over the long run due to institutional constraints (Kimhi 1991a). Capital stock includes only capital assets at least ten years old.

6. It is also possible to correct for selection bias in the models of those who work part-time on the farm. This demands correcting for double selection rules, which is a natural extension of our model, but is not performed here.

7. One explanation to this puzzle can be the endogeneity of the extent of farm work. In subsequent work, we plan to deal with that by jointly estimating farm and off-farm participation equations.

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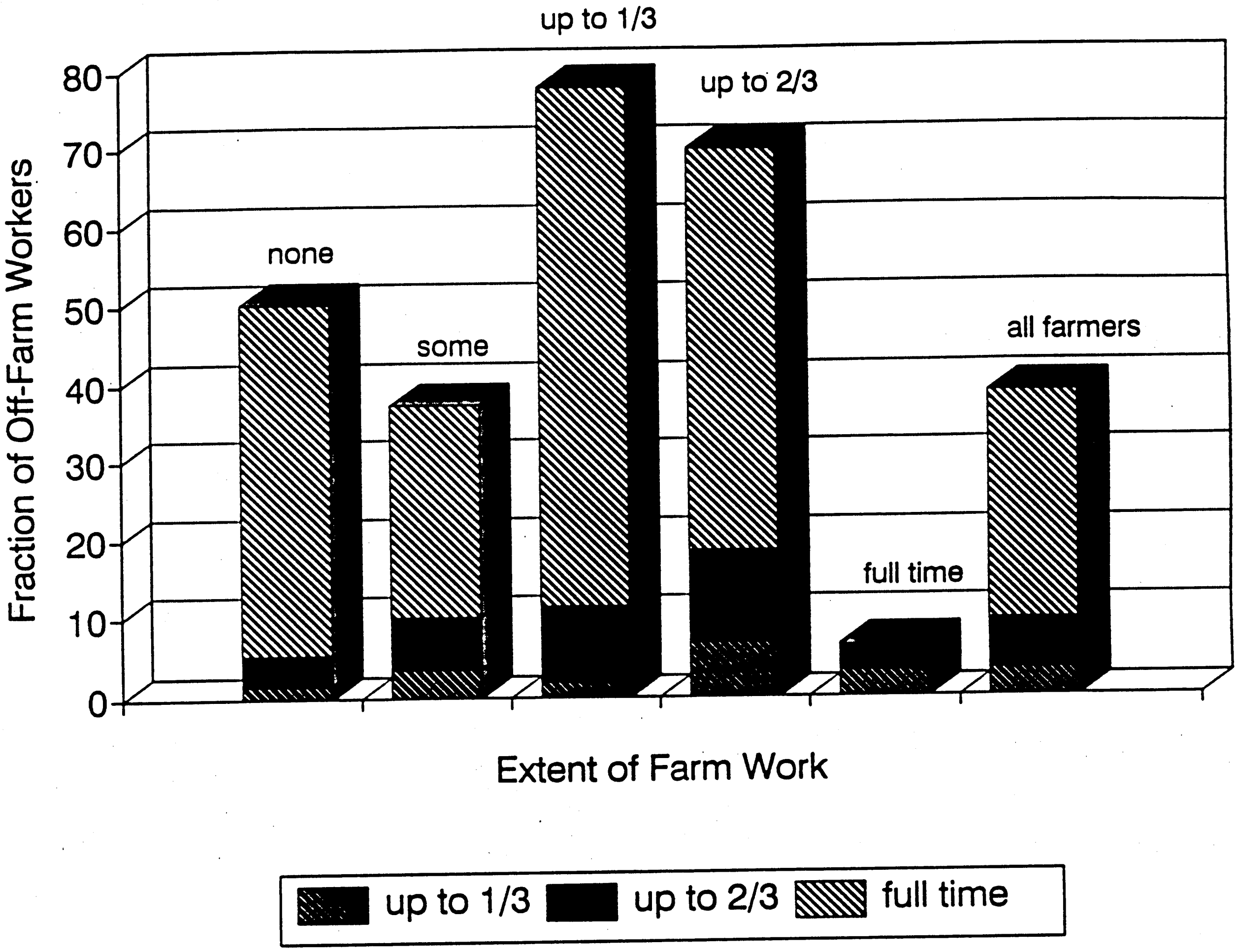

Figure 1: Fractions of Off-Farm Workers by Extent of Farm Work

Table 1. Descriptive Statistics

a. 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	heads
Family Members 15-21	.89	1.3	0-8	heads
Family Members 22-65	1.5	1.1	0-9	heads
Family Members 66+	.12	.37	0-2	heads
Total Land[c]	30.2	38.2	1-3030	dunam[d]
Old Capital[e]	11.1	26.7	0-1049	$1000[f]

b. Qualitative

	Number	Percent
Dairy Farm		
Working On Farm		
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
Working Off-Farm		
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

a.	Subsample (according to farm participation)			
	All Sample	Work on Farm	Don't Work Farm	
Intercept	-2.458 (14.9)	-2.286 (13.5)**	-0.263 (0.32)	-0.550 (0.71)
Age	0.121 (17.1)	0.116 (15.5)**	0.058 (1.91)*	0.054 (2.37)**
$(\text{Age})^2/50$	-0.072 (20.6)**	-0.068 (18.1)**	-0.058 (4.29)**	-0.057 (5.36)**
In Israel	0.009 (5.65)**	0.009 (4.93)**	0.014 (2.17)*	0.014 (2.17)*
Years on Farm	0.001 (0.63)	0.000 (0.03)	-0.011 (1.85)*	-0.013 (2.24)*
Schooling	0.042 (14.3)**	0.037 (11.5)**	0.080 (8.72)**	0.078 (8.70)**
Family -14	0.007 (0.94)	0.007 (0.89)	-0.006 (0.25)	-0.005 (0.22)
Family 15-21	0.042 (4.68)**	0.039 (4.27)**	0.026 (0.94)	0.022 (0.91)
Family 22-65	0.056 (5.64)**	0.022 (1.91)*	0.115 (4.70)**	0.110 (3.84)**
Family 66+	0.061 (2.19)*	0.039 (1.29)	0.159 (1.65)*	0.151 (1.62)
Total Land	-0.306 (25.2)**	-0.309 (24.6)**	-0.141 (3.11)**	
Old Capital	-0.018 (4.19)**	-0.019 (4.22)**	0.002 (0.15)	
Dairy Farm	-0.576 (14.5)**	-0.602 (13.3)**	-0.037 (0.24)	
No. of Cases	16818	15074	1744	1744
% Correct	66.2	66.5	79.6	79.5
Log Likelihood	-10132	-9090	-792	-797

Continued on next page

Table 2. ***(Continued)***

b.	Subsample (according to extent of farm work)			
	Up to 1/3	Up to 2/3	Full Time	
Intercept	-2.549 (5.79)**	-3.450 (7.53)**	-2.980 (8.02)**	-3.140 (7.69)**
Age	0.146 (8.19)**	0.192 (9.89)**	0.081 (4.86)**	0.080 (4.35)**
$(Age)^2/50$	-0.096 (11.4)**	-0.119 (12.3)**	-0.046 (5.38)**	-0.047 (4.98)**
In Israel	0.007 (1.52)	0.008 (1.79)*	0.001 (0.21)	0.001 (0.17)
Years on Farm	-0.001 (0.12)	-0.012 (2.44)**	0.001 (0.26)	0.001 (0.24)
Schooling	0.081 (11.6)**	0.048 (6.23)**	0.026 (3.52)**	0.026 (3.61)**
Family -14	0.041 (2.30)*	0.007 (0.37)	-0.015 (1.03)	-0.017 (0.98)
Family 15-21	0.050 (2.22)*	0.032 (1.53)	0.026 (1.22)	0.025 (1.01)
Family 22-65	0.083 (3.53)**	0.029 (1.09)	-0.047 (1.93)*	-0.060 (2.10)*
Family 66+	0.135 (1.92)*	0.085 (0.98)	-0.028 (0.45)	-0.032 (0.47)
Total Land	-0.045 (1.45)	0.005 (0.15)	-0.058 (1.84)*	-0.030 (0.78)
Old Capital	-0.001 (0.06)	0.006 (0.52)	0.015 (1.50)	0.017 (1.84)*
Dairy Farm	-0.226 (1.74)*	-0.296 (2.58)**	-0.243 (3.27)**	-0.179 (1.79)*
ρ				-0.199 (1.14)
No. of Cases	4040	2797	8237	
% Correct	86.0	79.1	93.5	93.5
Log Likelihood	-1409	-1319	-1920	-1919

Continued on next page

Table 2. ***(Continued)***

c.	All Sample with Farm Work Dummies	
Intercept	-2.827 (13.0)**	-3.015 (14.8)**
Age	0.154 (17.1)**	0.155 (17.4)**
$(Age)^2/50$	-0.099 (22.5)**	-0.099 (22.7)**
In Israel	0.009 (4.18)**	0.008 (3.91)**
Years on Farm	-0.001 (0.66)	-0.002 (1.10)
Schooling	0.060 (16.0)**	0.060 (16.0)**
Farm Work 0-1/3	0.546 (13.7)**	0.549 (14.1)**
Farm Work 1/3-2/3	0.143 (3.44)**	0.141 (3.36)**
Farm Work 2/3-1	-2.172 (53.0)**	-2.203 (53.9)**
Total Land	-0.068 (4.13)**	
Old Capital	0.010 (1.84)*	
Dairy Farm	-0.191 (3.64)**	
% Correct	87.57	87.63
Log Likelihood	-5586	-5601

Notes: All models included a set of ethnic origin dummies.
t-statistics in parenthesis.
* - significant at the 5% level.
** - significant at the 1% level.
% of correct predictions: prediction 1 means probability of 1 is greater than 1/2.
part c: family variables are not reported; coefficients are very close to those in part a.

Table 3. Changes in the Results After Correcting for Selection Bias

	Work on Farm		Don't Work on Farm	
	Uncorrected	Corrected	Uncorrected	Corrected
Intercept	-2.286 (13.5)**	-2.305 (11.3)**	-0.263 (0.32)	0.290 (0.26)
Age	0.116 (15.5)**	0.112 (12.2)**	0.058 (1.91)*	0.056 (1.77)*
$(Age)^2/50$	-0.068 (18.1)**	-0.068 (14.8)**	-0.058 (4.29)**	-0.059 (4.21)**
In Israel	0.009 (4.93)**	0.008 (4.53)**	0.014 (2.17)*	0.014 (1.96)*
Years on Farm	0.000 (0.03)	-0.001 (0.35)	-0.011 (1.85)*	-0.012 (1.87)*
Schooling	0.037 (11.5)**	0.038 (10.7)**	0.080 (8.72)**	0.081 (8.66)**
Family 0-14	0.007 (0.89)	0.002 (0.24)	-0.006 (0.25)	-0.015 (0.52)
Family 15-21	0.039 (4.27)**	0.037 (3.24)**	0.026 (0.94)	0.028 (0.95)
Family 22-65	0.022 (1.91)*	-0.024 (1.40)	0.115 (4.70)**	0.072 (1.07)
Family 66+	0.039 (1.29)	0.023 (0.65)	0.159 (1.65)*	0.138 (1.21)
Total Land	-0.309 (24.6)**	-0.277 (13.5)**	-0.141 (3.11)**	-0.121 (2.28)*
Old Capital	-0.019 (4.22)**	-0.016 (3.24)**	0.002 (0.15)	0.005 (0.34)
Dairy Farm	-0.602 (13.3)**	-0.520 (9.13)**	-0.037 (0.24)	0.051 (0.25)
ρ		-0.641 (3.59)**		-0.248 (0.78)
No. of Cases	15074		1744	
% Correct	66.5	66.6	79.6	79.7
Log Likelihood	-9090	-9079	-792	-792

Notes: see notes to table 2.