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Farmers' time allocation between farm work and off-farm work and the importance of unobserved group effects: evidence from Israeli cooperatives

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Abstract

Agricultural cooperatives differ in many attributes which affect members' farm and off-farm earnings in different ways. As a result, time allocation patterns between farm work and off-farm work will vary significantly across cooperatives. Participation equations in farm work and off-farm work of farmers who are members of Israeli moshavim are estimated jointly, including a cooperative-specific factor in each equation. The fixed effects are found to be significant and important, but can be only partly explained by observed cooperative attributes. This provides another support to the Monte Carlo results of Borjas and Sueyoshi, that controlling for group effects is superior to alternative models. The results also imply that unobserved factors have considerable effects on farmers' time allocation, and hence should be controlled for whenever possible.

1. Introduction

The tendency of farmers to work off the farm is an issue with important policy implications (Shucksmith and Smith, 1991). The economic literature is rich in studies of the time allocation of farmers (see Hallberg et al., 1991, and the references therein), starting with the pioneering work of Huffman (1980). Most studies emphasized the effects of observed personal characteristics, farm attributes, and local labor market conditions on the time allocation decisions. The absence of appropriate panel data has prevented researchers from accounting for unobserved factors which affect farmers' time allocation. However, the importance of the unobserved factors can be studied using data on farmers who are grouped according to a factor which is related to time allocation.

An example for a suitable data set is data on farmers who are members of moshavim (semi-cooperative villages) in Israel. The importance of the institutional environment to farmers' time allocation has been established empirically, mainly by using location-specific data (e.g. Tokle and Huffman, 1991). The moshav association itself creates a distinct institutional environment in which its members have to operate. The implications of this for time allocation has been studied by Kimhi (1991a).

The purpose of this paper is to investigate the importance of moshav-specific group effects in a model which explains the joint decisions of a farmer to engage in farm as well as off-farm work (Kimhi, 1994a). The empirical model used is based on the Borjas and Sueyoshi (1994) two-stage probit estimation technique, with an extension to a bivariate model. In particular, a cooperative-specific fixed effect is

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added to a bivariate probit model of farm and off-farm participation equations. This is the first stage, and its results are compared to the results of a pooled model in order to assess the contribution of the group effects. In the second stage, the estimated fixed effects are regressed on observed cooperative attributes, including geographic location and economic and demographic variables, using seemingly unrelated regression. This enables the evaluation of the relative importance of observed and unobserved group effects.

The results indicate that cooperative-specific factors have a statistically significant effect on work participation patterns of Israeli farmers, and that ignoring the cooperative effects could result in biased coefficients. Moreover, only about a quarter of the cooperative-specific effect can be explained by observed cooperative attributes. Therefore, using observed attributes such as location-specific variables is insufficient. This may be true for other data sets as well.

Section 2 of this paper discusses the special properties of the moshav cooperatives which give rise to estimation conditioned on cooperative attributes. The theoretical model of farm and off-farm participation is presented in Section 3, as well as the two-stage estimation procedure and its justification. The data set is described in Section 4. The results of the participation equations are reported in Section 5, and those of the cooperative effects in Section 6. Section 7 summarizes and concludes.

2. The importance of group effects in Israeli moshavim

The population of interest comprises family farms in Israeli moshavim. Moshavim is the Hebrew name (in plural form) for cooperative (or semi-cooperative) villages. In each moshav, membership is by family, each of which maintains its own household, farms its own allocation of land and earns its income from what it produces. The moshav takes advantage of economies of scale by collectively handling matters of mutual concern such as purchasing, marketing, investments and credit, and operating as a member of regional and national organizations. It is also responsible for education and social activities, and acts as a municipal entity. This structure is different from that of the better-known kibbutz collectives. In the latter, the family has no economic or social role; membership is individual, all property is community-owned, and work and consumption are equally shared by all members. [1]

In general, every economic decision taken by an agent is affected by its economic environment, in this case the cooperative, and every such decision indirectly affects other members. Zusman (1988) classified all these as transaction costs (p. 68):
"... the model of the Moshav in operation comprises both individual discretionary decisions and group choices regarding allocation rules, collective actions and restrictive regulations, all of which interact under conditions of uncertainty, imperfect information structures and complexity, thereby giving rise to a variety of 'transaction costs'... "

In general terms, these transaction costs affect potential farm profits and perhaps off-farm earnings, and thus have an effect on individual participation decisions. Moreover, in the presence of heterogeneity among moshavim, participation patterns will be different across moshavim. Moshavim are heterogeneous in location (land quality, access to external labor markets, etc.), resource allocation (land, water, quotas) and population characteristics (ethnic origin, education). Haruvi and Kislev (1984), who study the heterogeneity in the level of cooperation among moshavim, write (p. 54):
"Despite common structural traits, moshavim differ widely in degree and nature of cooperation. Some maintain joint cash management, central planning and direction, and strong public services. Others are loosely organized farm communities... "

The level of cooperation has perhaps the most obvious effect on the participation patterns. If the level of cooperation affects farm profits positively because of scale advantages, then it will affect farm work participation positively and off-farm work participation negatively. This is an example of a 'management factor': in cooperatives that are better managed, farmers will have a higher tendency to participate in farm work and a lower tendency to

[1] More information on time allocation in moshavim and their institutional structure is in Kimhi (1991a).

participate in off-farm work, other things being equal. Since the unobserved (unlike in the work of Haruvi and Kislev) level of cooperation is common to all farmers in the same moshav, it is included in a moshav-specific effect.

3. Theory and empirical model

Optimal time allocation is generally characterized by equal returns to time in different uses (Gronau, 1977). However, this is only true when the returns are a diminishing function of time in each possible use. Otherwise, corner solutions may occur. Corner solutions are possible in the presence of earning functions which are not nicely behaved in the spirit of Bator (1971). An example is the presence of fixed costs (Cogan, 1980). In modelling participation decisions, one concentrates on these corner solutions. A corner solution with respect to a certain time use means that the person is worse off by devoting any amount of time to that use.

In standard labor force participation models, a person participates if the market wage is higher than the reservation wage. The reservation wage is the marginal rate of substitution between consumption and leisure, measured at zero hours of work (Killingsworth, 1983). Similarly, for the case with fixed costs, reservation utility related to a certain time use is defined as the maximum value of utility attainable without devoting time to that use. Hence, a person participates if the utility he attains is larger than the relevant reservation utility.

More formally, denote utility by $U(c,t_1,\dots,t_n)$, where c is consumption and t_n is the time devoted to activity n. This notation allows time to affect utility directly in all its uses. The indirect utility function can be denoted as $V(X)$, where X includes all the determinants of utility and income other than time and consumption. For example, X may include farm attributes, socio-economic characteristics, labor market conditions, and market prices. Specifically, it includes the determinants of the fixed costs. The reservation utility with respect to activity n can be denoted as $V_n(X|t_n=0)$, and hence the relevant participation equation is $Y_n(X)=V(X)-V_n(X|t_n=0)$. The person participates only if $Y_n(X)>0$.

In empirical analyses, an approximation is used for modelling the participation equation as a function of the observable explanatory variables. Denote the approximation of Y by y and the approximation error by u. The resulting empirical participation equation is $y_n(x)+u_n$, where x is the observable subset of X. Suppose that there are data on I_j farmers in each of J cooperatives, and that the vector of observable covariates x can be partitioned to individual-specific covariates x' and cooperative-specific covariates x''. Also, it is natural to assume that the approximation error u includes a cooperative-specific element m. In this case, the participation equation for activity n becomes:

$$y'_{nij}(x')+y''_{nj}(x'')+m_{nj}+\epsilon_{nij};$$
$$i=1,\dots,I_j;\qquad j=1,\dots,J \tag{1}$$

where the function y is partitioned into an individual-specific component y' and a cooperative-specific component y'' according to the partitioning of the x vector, and $\epsilon_{nij}=u_{nij}-m_{nj}$.

This model was discussed by Borjas and Sueyoshi (1994). They proposed a two-stage estimation procedure which was found to be superior to alternative procedures in Monte Carlo simulations. According to their procedure, $y''_{nj}(x'')+m_{nj}$ is treated as a cooperative-specific fixed effect. The whole set of fixed effects is estimated in the first stage together with the coefficients of the y' function. This means estimating Eq. (1) by probit, assuming that the ϵ's are normally distributed. The estimated fixed effects, d_{nj}, should then be regressed on the cooperative-specific observed attributes x'' in the second stage, using least squares regression.

The alternatives to this two-stage procedure are: (a) to ignore the cooperative-specific error term m_{nj} altogether and estimate Eq. (1) in one stage including all the covariates; (b) to treat m_{nj} as random effects and use an appropriate maximum likelihood technique. While the first alternative is easy to implement and provides consistent estimators, the estimators are relatively inefficient because the procedure incorrectly assumes independence of observations which belong to the same cooperative. On the other hand, the second alternative yields estimators with superior asymptotic properties but is computationally not easy to implement, especially in our case of 90

Table 1
Descriptive statistics

Variable	Complete sample		Selected sample		Units
	Mean	SD	Mean	SD	
Quantitative					
Age	48.5	15.1	49.7	14.3	Years
In Israel [a]	33.0	10.6	34.5	10.1	Years
Tenure [b]	18.5	11.4	19.5	11.4	Years
Schooling	9.0	4.6	9.3	4.5	Years
Family [c]	4.7	2.5	4.8	2.5	Persons
Children [d]	1.5	1.5	1.4	1.5	Persons
Land	5.7	7.7	6.7	7.4	Acres
Old capital [e]	20.0	31.2	28.6	39.7	\$1000 (1981)
	Complete sample		Selected sample		
Qualitative					
Working on-farm	62.2%		65.0%		
Working off-farm	47.3%		45.4%		
Ethnic origin [f]	44.9%		53.6%		
Dairy farm [g]	5.7%		8.1%		
Number of observations	27534		6093		

[a] The number of years since immigrating to Israel.
[b] The number of years since starting to operate the farm.
[c] Number of family members (including the respondent).
[d] Number of children up to 14 years of age.
[e] Value of farm assets built or purchased at least 10 years ago.
[f] Being born in Europe or America, or to an immigrant from those countries, or having an unknown origin.
[g] Having a positive number of milk cows.

different cooperatives. In any case, the Monte-Carlo results of Borjas and Sueyoshi (1994) support the use of their proposed two-stage procedure.

In this study, three possible uses of time are assumed: farm work; off-farm work; home time (including leisure). Since home time is assumed to be always positive, two participation equations are jointly estimated: farm work and off-farm work (Kimhi, 1994b).[2] It is common to assume that ϵ is independent across sample observations, but it is expected that the ϵ's of different participation equations of the same individual will be correlated. This could be because the same unobserved personal characteristics are included in all of them, for example. In statistical terms, one can say that the ϵ's are independent across i and j but not across n. If the vector $(\epsilon_{fij},\epsilon_{oij})'$ is drawn from the same bivariate normal distribution for all the observations, where f stands for farm and o stands for off-farm, then the joint participation model can be estimated by the bivariate probit maximum likelihood method (Amemiya, 1985, p. 317).

In the second stage, the two sets of estimated dummy variables (each belongs to one participation equation), should be regressed on cooperative attributes. This is done jointly by using a seemingly unrelated regression, in order to maximize efficiency.[3]

[2] It should be noted that the possibility of a farm-household head who does not work on the family farm exists in the current data set. In these cases, generally another family member, such as a spouse of an adult child, operates the farm.

[3] However, corrections have not been made to the fact that the dependent variables in this stage (the dummies) are measured with error.

Table 2
Testing the importance of village dummies [a]

Model:	(1)		(2)	
Village dummies included:	Yes		No	
Log-likelihood:	−5836		−6483	
Number of observations:	6093		6093	
Correlation coefficient:	−0.5242		−0.5360	
Equation:	On-farm		Off-farm	
Variable	(1)	(2)	(1)	(2)
Intercept	−1.265 (−2.98) * * *	−0.6739 (−2.01) * *	−3.969 (−8.69) * * *	−2.158 (−5.93) * * *
Origin	0.0846 (1.17)	0.1272 (2.82) * * *	0.0735 (1.07)	−0.0763 (−1.73) * *
Age	0.0388 (2.56) * * *	0.0257 (1.86) * *	0.1633 (9.97) * * *	0.1429 (9.59) * * *
$(\text{Age})^2/100$	−0.0515 (−3.07) * * *	−0.0431 (−2.81) * * *	−0.2050 (−10.9) * * *	−0.1737 (−10.1) * * *
In Israel	0.0045 (0.67)	−0.0027 (−0.46)	−0.0053 (−0.65)	−0.0121 (−1.68) * *
Tenure	0.0082 (0.66)	0.0110 (0.97)	−0.0579 (−4.32) * * *	−0.0448 (−3.67) * * *
Age × Tenure [b]	−0.0024 (−0.11)	−0.0038 (−0.20)	0.0958 (3.99) * * *	0.0679 (3.08) * * *
Schooling	0.0179 (0.91)	−0.0053 (−0.30)	0.0357 (1.58) *	0.0206 (1.01)
Schooling × In Israel [c]	−0.0234 (−0.41)	0.0172 (0.34)	0.0633 (0.94)	0.0897 (1.45) *
Land	0.0272 (18.1) * * *	0.0159 (17.0) * * *	−0.0120 (−11.0) * * *	−0.0090 (−13.5) * * *
Dairy farm	0.6632 (7.07) * * *	0.4694 (5.48) * * *	−0.7988 (−9.37) * * *	−0.6577 (−8.54) * * *
Old capital [d]	0.1034 (13.2) * * *	0.1050 (15.9) * * *	−0.0536 (−7.26) * * *	−0.0549 (−8.71) * *
Child	0.0281 (1.34) *	0.0349 (1.80) * *	−0.0526 (−2.60) * * *	−0.0314 (−1.69) * *
Family	0.0039 (0.29)	0.0117 (0.98)	0.0503 (3.76) * * *	0.0291 (2.45) * * *

[a] Asymptotic t-values in parentheses: * coefficient significant at 10%; * * coefficient significant at 5%; * * * coefficient significant at 1%.
[b] Interaction of Age and Tenure, divided by 100.
[c] Interaction of Schooling and In Israel, divided by 100.
[d] This is in fact log(old capital + 1).

A final note relates to the identification of the cooperative-specific fixed effects. Each cooperative's fixed effect is identified by the observations of that cooperative only. Hence the fixed effects can be identified only for cooperatives in which a sufficient number of observations exist.

4. The data

The data set includes individual records on farm-household heads in moshavim from the 1981 census of agriculture in Israel. Information is available on personal characteristics, farming activities, and time allocation (in qualitative terms). The original data set included 27 534 observations from 393 moshavim. However, moshav-specific information (location and establishment year), which is necessary for the second part of the empirical procedure, was available only for 90 of the moshavim. These included 6093 observations on farm-family heads of household, after removing incomplete records and landless families. The reason is that the partly missing moshav-specific variables were imported from a data set related to a different survey. In that survey, the 90 moshavim were drawn randomly from the 393 moshavim in the country. Therefore, there is no a priori suspicion of selection bias related to this issue.

In addition, descriptive statistics of the complete data set and of the selected sample are compared in Table 1, which also gives definitions of variables. The comparison does not reveal a strong suspicion of selectivity bias. The only notable differences are the higher percentage of farmers of western origin in the subsample, and the higher levels of farm attributes such as land and capital. This last point indicates that the subsample includes fewer inactive farms. This does not necessarily correspond to the selection of the 90 moshavim, but rather to the exclusion of landless families. [4]

[4] I have tried to define inactive farms in several ways and exclude them from the analysis, but the results did not change qualitatively.

Table 3
Seemingly unrelated regression of village dummies

Variable	On-farm	Off-farm
Intercept	−2.0771 (−2.70) ***	3.8668 (4.62) ***
Establishment year	0.0194 (3.22) ***	−0.0111 (−1.85) *
Distance		−0.0157 (−2.16) **
Age	0.0456 (3.03) ***	−0.0684 (−4.12) ***
Tenure	−0.0319 (−1.98) *	0.0465 (2.76) **
Land	0.0052 (2.07) **	
Dairy farm	0.6809 (1.83) *	
Fraction of old capital		0.5220 (1.99) *
Family	−0.0819 (−2.23) **	
Number of observations	89	89
R^2	0.193	0.200

[a] Dependent variables are the two sets of village dummies estimated in the first stage.
Asymptotic t-values in parentheses: * coefficient significant at 10%; ** coefficient significant at 5%; *** coefficient significant at 1%.

5. Results of participation equations

The bivariate probit model was estimated with and without the cooperative-specific fixed effects. [5] The results are reported in Table 2. The main finding is that the village-specific dummies, whose estimates (89 in number) are not reported, are statistically significant as a group at the 1% level. The likelihood ratio statistic for the exclusion of these dummies is 1294 with 178 degrees of freedom, compared to a critical value of 225. Most parameters are qualitatively robust to the inclusion of these dummies, at least in terms of sign. However, several coefficients changed quite a bit after the inclusion of the fixed effects. For example, the age profile of farm work participation peaks later and becomes more concave. Also, the positive schooling effect on off-farm participation rises notably, the effects of land and the dairy dummy change in various ways, and the effects of children and other family members on farm (off-farm) participation become much smaller (larger) in absolute value.

The most significant determinants of farm work participation are farm attributes. Land, capital and the dairy dummy are all associated with a higher farm participation probability. [6] Farm participation is higher in families with more children. This is similar to the results of a model which emphasizes family structure and controls for joint decisions of household heads and spouses (Kimhi, 1994c). Farmers of Western ethnic origin have a higher tendency to work on farm, and this tendency also increases with age, but in a decreasing rate.

Off-farm work participation is more sensitive to personal characteristics. The age effect has the same sign as in the case of farm participation, but is much stronger. Farm tenure (number of years on current farm) reduces off-farm participation, in a rate that is decreasing with age. The same is true for years in Israel, where the rate is decreasing with years of schooling. Farm and family attributes have exactly the opposite effect on off-farm participation than on farm participation, and, in addition, off-farm participation is also rising with the number of adult family members. These results are in line with previous research on the topic (Lass et al., 1991).

[5] Estimation was performed using the HotzTran program. I thank Joe Hotz and Jeff Smith for their help on this matter.

[6] Farm attributes are suspected of being endogenous to the time allocation decision. The analysis may be perceived as conditioned on farm attributes. However, the selected farm attributes are those least expected to be endogenous (Kimhi, 1991a).

6. Results of cooperative effects

In the second stage of estimation, the estimated cooperative-specific fixed effects are regressed on observable cooperative attributes according to the regression equation $d_{nj} = y''_{nj}(x'') + m_{nj}$. Assuming linearity of the regression equations, the two sets of dummies derived from the farm work participation equation and the off-farm participation equation were regressed jointly. The Seemingly Unrelated Regression method was used in order to allow for dependence between the two m_{nj} terms for each cooperative. The explanatory variables included two types of cooperative attributes. The first type includes the cooperative's establishment year, and several variables based on location information, which represent attributes of the local labor market (Tokle and Huffman, 1991; Findeis et al., 1991): employment of males in the surrounding subdistrict (in thousands of employees), distance (in miles) to the nearest town of at least 25 000 residents, and the population of the town to which the distance was measured. Of those three, only distance turned out significant. The second type of cooperative attributes are the within-cooperative means of the individual variables used in the participation equation: age, ethnic origin, years in Israel, tenure, land, capital, dairy farm, and family size. These averages proxy for the demographic characteristics and agricultural attributes of each cooperative. Variables which were not significant in any of the two equations were dropped entirely.

The results are reported in Table 3. [7] Overall, the explanatory variables account for less than one-fifth of the variation in the dummies. Establishment year had a significant coefficient in both equations. In moshavim that were established later, the cooperative effect works in favor of farm work and against off-farm work. This could be related to the life-cycle of the cooperatives (Kimhi, 1991b). Of the three location variables, only distance to town had a significant effect (negative, on the off-farm dummy). Average age has coefficients similar in sign to establishment year. Since these variables are negatively correlated in the sample, this could be due to multicollinearity. Average tenure has a negative (positive) effect on the cooperative dummy associated with farm (off-farm) work, contrary to intuition. This too could result from life-cycle effects and succession considerations. The other variables affect the dummies in a way similar to the way they affect participation patterns in the individual regressions.

Finally, the cooperative-specific observed attributes were included in the bivariate probit model instead of the fixed effects. The performance of this model was poorer than that of the fixed effects model, as expected. In particular, the coefficients of location variables were not statistically significant as a group. The estimates of the other coefficients were significant as a group (likelihood ratio statistic of 546 with 46 degrees of freedom, compared to a critical value of 72) at the 1% level, but seemed to suffer from the multicollinearity between the individual variables and the village means.

[7] Variable exclusions were determined by stepwise regression.

7. Summary and conclusions

This paper has shown that cooperative-specific factors have considerable effects on the participation patterns of their members in farm and off-farm work. Specifically, cooperative-specific fixed effects were significant in a joint participation model, after controlling for personal, family and farm characteristics. Moreover, several coefficients changed quite a bit after the inclusion of the fixed effects, which means that ignoring the cooperative-specific factors may lead to omitted variable bias and incorrect conclusions. Also, most of this effect (about 80% in this data set) is due to unobserved cooperative attributes. This means that the problem cannot be properly solved by using observed cooperative attributes instead of the fixed effects. This calls for using more detailed data sets in order to perform a finer decomposition of the unexplained variation in participation patterns. This will increase our understanding of the behavior of farmers and the way it is affected by their institutional environment.

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