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OFF-FARM LABOUR PARTICIPATION OF ITALIAN FARMERS, STATE DEPENDENCE AND THE CAP REFORM

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Abstract

We analyse the determinants of off-farm labour participation of farmers. For estimation, we use different dynamic models, accounting for both heterogeneity and state dependence, as well as for the initial conditions. Our results suggest that, when keeping into account all these features, the present work state is almost totally explained by the previous state and by idiosyncratic characteristics, which implies a strong persistence. Variables concerning personal characteristics are not found to be significant in the dynamic setting. Finally, the variables related to the reform of the Common Agricultural Policy have little or no statistically significant effect.

Keywords: off-farm work, farm household, state dependence, panel data, CAP reform.
JEL Classification codes: J220, J430, Q120, Q180

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1. Introduction

The analysis of off-farm labour participation is relevant for the evolution of agriculture, but also for the issue of rural development. Though agriculture is often no more the principal source of employment in rural areas, it has undoubtedly a strong importance in shaping the territory and the patterns of population –or abandonment- of those areas.

The analysis of the determinants of off-farm labour participation already forms a large body of literature. Starting from the seminal paper by Huffman (1980), labour choices in farm households have been widely analysed, initially considering off-farm labour participation of farm operators (Sumner, 1982), later considering the joint decision making of off-farm labour participation of husbands and wives (e.g. Huffman and Lange, 1989; Tokle and Huffman, 1991; Lass and Gampesaw, 1992), also including the use of waged labour (Benjamin *et al.*, 1996, Blanc *et al.*, 2006), on- and off-farm labour participation of operators, spouses and other household members (Kimhi, 1994 and 2004, Benjamin and Kimhi, 2006; Bjørnsen and Biørn E, 2010; Corsi and Salvioni, 2012). Most of the research concerned the U.S., but several analyses are devoted to the determinants of off-farm labour participation in the EU (Corsi, 1994; Benjamin *et al.*, 1996; Weiss, 1997 and 1999; Woldehanna *et al.*, 2000; Benjamin and Kimhi, 2006; Salvioni *et al.*, 2008; Hennessy and Rehman, 2008; Bjørnsen and Biørn, 2010; Corsi and Salvioni, 2012).

A common feature of this stream of research, with few exceptions, is that the analyses are based on cross-sectional samples. This prevents analysing the dynamic nature of off-farm labour participation choices, and disregards the persistence of the phenomenon. Persistence was actually observed in the few cases in which panel data were available (Gould and Saupe, 1989; Weiss, 1997; Corsi and Findeis, 2000; Bjørnsen and Biørn, 2010). However, reason for persistence can be true state dependence and/or heterogeneity. True state dependence exists when the previous state modifies the attitudes, constraints or parameters so that the probability of present state is affected by the previous state. Heterogeneity represents unobservable time-invariant idiosyncratic characteristics that affect the choices across time periods.

This article analyses off-farm labour participation of farm operators based on a dynamic model allowing for both true state dependence and heterogeneity. It is based on a panel sample of Italian for the period 2003-2007. Since in this period a major reform of the Common Agricultural Policy (CAP), the so-called Fischler reform, took place, it also account for the relevant changes.

The structure of the paper is as follows. Section 2 will present the theoretical framework and the econometric strategy. In Section 3 the data on which the analysis is based will be presented. The estimation results will be presented and commented in Section 4. Some considerations will conclude.

2. Theoretical Model and Econometric Strategy

The well-known farm-household model (Nakajima, 1986; Singh *et al.* 1986; Huffmann, 1991) is usually used in the static analysis of off-farm labour participation. The reduced form usually employed is based on the comparison between the reservation wage w^* and the market wage w , which is a function of personal characteristics and of the

conditions of the local labour market. Typically, the reduced-form equations are probit or logit equations estimating the influence of these variables on the probability of $(w-w^*)$ being greater than zero. Four categories of explanatory variables are usually included: individual, household, farm and local market characteristics.

Few studies of off-farm employment decisions have used longitudinal data. Exceptions are Gould and Saupe (1989), Weiss (1997), and Corsi and Findeis (2000). Though, these studies have used two-wave panels, so that a truly dynamic modelling was not possible.

In a dynamic setting, expected utility and expected income streams are to be considered. Moreover, costs for shifting from one to the other work condition are to be taken into account. These costs are to a large extent sunk costs. For instance, taking on an off-farm job might imply search costs, and costs related to the adaptation of the farm operation to the new situation (e.g., a higher mechanization, or a change in the type of farming or in the farming intensity to cope with lower family labour input).

Call C_1 the costs for shifting from non-participation to participation, and C_2 the cost for the reverse change; T the time-horizon for work incomes; r the discount rate; w_t the wage the farmer could earn working off the farm in time t ; f_{1t} and f_{2t} the farm income the farmer could earn when working and not working off the farm, respectively, in time t . The probability of participation in year t for a farmer not already participating (for simplicity, we consider utility as only stemming from income, non-pecuniary benefits from different jobs can be easily incorporated) is:

$$prob(y_t = 1 | y_{t-1} = 0) = prob \left\{ E \left[\sum_{t=1}^T (w_t + f_{1t} - f_{2t}) \frac{1}{(1+r)^t} - C_1 \right] > 0 \right\} \quad (1)$$

The probability of continuing participating for a farmer already participating is:

$$prob(y_t = 1 | y_{t-1} = 1) = prob \left\{ E \left[\sum_{t=1}^T (w_t + f_{1t} - f_{2t}) \frac{1}{(1+r)^t} + C_2 \right] \geq 0 \right\} \quad (2)$$

Therefore in general, $prob(y_t=1|y_{t-1}=1) > prob(y_t=1|y_{t-1}=0)$. This requires including past off-farm work state as an explanatory variable in the participation equation, as a proxy for the costs for shifting from one condition to another, that create a state dependence. Unlike previous literature, we exploit the panel nature of our data to analyse the determinants of off-farm labour participation. The probability of a farm operator working off-farm is estimated by applying a dynamic non-linear probit random effects model accounting for both unobserved heterogeneity and true state dependence.

The non-linear dynamic random effect model we start from is:

$$Prob(y_{it}=1) = \Phi(\beta_0 x_{it} + \gamma y_{it-1} + u_i + \varepsilon_{it}) \quad (3)$$

where y_{it} is a dummy variable equal to 1 if the farm operator i participates in off-farm work in year t ($t=1, \dots, T$), else 0; \mathbf{x} is a vector of observable time-variant and time-invariant explanatory variables; β_0 and γ are parameters to be estimated; Φ is the normal cumulative density function; u_i is an individual idiosyncratic term representing time-constant unobservable individual characteristics, i.e., heterogeneity; and ε_{it} is a random component, uncorrelated across individuals and years. The \mathbf{x}_{it} vector includes all variables influencing the potential external wage w_t (personal characteristics of the operator and labour market conditions) and farm incomes f_{1t} and f_{2t} (farm characteristics and operator's skills, proxied by personal characteristics). Shifting costs C_1 and C_2 are assumed to be a function of the

previous state y_{t-1} , and possibly of labour market characteristics, of farm characteristics, of personal characteristics.

The standard random effects model assumes that the unobserved individual-specific components u_i are uncorrelated with the observed explanatory variables. In the real world, this assumption may not hold. A further issue, that may result in biased estimates of the parameter of the lagged variable, is that commonly referred to as the initial conditions problem – whether y_{i1} is independent of u_i . In recent years, several estimators for the nonlinear dynamic panel data model have been proposed (Heckman, 1981; Wooldridge, 2005; and Orme, 2001), and it has been shown (Arulampalam and Stewart, 2009), on the basis of simulation experiments, that none of the three estimators dominates the other two in all cases. In this paper, we allow for the possibility of the individual term to be correlated to the regressors and we take into account the initial conditions problem following the approaches by Heckman (1981), Wooldridge (2005) and Orme (2001).

Heckman's approach to the initial conditions problem involves assuming a specification of the initial value of the dependent variable including x_{i1} and other instrumental variables, and estimating the function for the initial time period jointly with the one of the remaining time periods, including the correlation term. In Wooldridge's approach, the individual term is assumed to be a function of the initial condition and of the means of the explanatory variables, and the model is estimated as a random effects probit. Orme's approach to the initial conditions problem is to insert a correction term in the equation. This term is the generalised error term from a probit estimated for the initial year, and is inserted as a regressor in a random effect probit model for the remaining years. We experiment all these approaches and confront them with the estimates based on cross-sectional models as well as among them.

3. Data

This study relies on data collected by the Italian Farm Accountancy Data Network (FADN) survey. The survey is conducted yearly on a random sample of more than 10,000 farms. The field of observation is the total of commercial farms, that is, farms with an economic size greater than 4 ESU (4,800 euro). In the absence of genuine panel data, we used the repeated cross-sectional data to construct a balanced panel in which each farm is tracked over time. The result is a 5 waves balanced panel of 3294 farms for which information were collected in all years from 2003 to 2007. We only kept family farms¹. The random effects probit models are estimated over the period 2004-2007, since year 2003 is used to define the initial condition. Table 1 presents the descriptive statistics for the dependent and explanatory variables over the 5 years.

The dependent variable is a dummy variable indicating whether the farm operator works off the farm. The share of farms operators having an off-farm job is 6.6 percent over the whole 2003-2007 period. It increases from 6.3 percent in 2003 to 7.3 percent in 2007, with a drop to 6 percent in 2004.

¹ Farms that were not sole ownership or private partnership were not considered in this analysis.

Table 1 - Descriptive statistics of the variables (2004-7 unless otherwise stated)

	Mean	Std.Dev.
share of off-farm labour participants 2003-2007	0.066	0.066
share of off-farm labour participants 2003	0.063	0.013
share of off-farm labour participants 2004	0.06	0.012
share of off-farm labour participants 2005	0.067	0.013
share of off-farm labour participants 2006	0.07	0.014
share of off-farm labour participants 2007	0.073	0.015
<i>Personal characteristics</i>		
Operator's age	54.19	13.6
Operator's age squared	3121.84	1513.44
Operator's gender (1=M, 0=F)	0.83	0.38
<i>Farm characteristics</i>		
UAA (ha)	33.7	69.74
Share of land in property (%)	0.67	0.39
Working capital (1000 Euro)	128.04	357.75
Total Debts (1000 Euro)	11.87	106.42
Hp/UAA	19.03	44.27
Types of Farming labour intensive all year round (0,1)	0.36	0.48
Types of Farming labour intensive seasonally (0,1)	0.33	0.47
Mountain (0,1)	0.19	0.39
Hills (0,1)	0.45	0.5
Plains (0,1)	0.37	0.48
Direct sales (0,1)	0.21	0.41
Organic farming (0,1)	0.04	0.19
Agro-tourism (0,1)	0.03	0.16
<i>Household characteristics</i>		
Pension income (0,1)	0.23	0.42
Capital income (0,1)	0.01	0.11
<i>CAP reform</i>		
Coupled Payments (1000 Euro)	3.48	33.79
Single Farm Payment (1000 Euro)	7.84	33.21
<i>Economic environment</i>		
Agricultural to total employment (%)	6.35	4.03
VA per inhabitant (1000 Euro)	21.72	5.23

Following previous research, we use four categories of explanatory variables: individual characteristics (age, age squared and gender of the operator); household characteristics (household non-labour incomes, i.e., capital income and pensions, used to explore the existence of a wealth effect) farm characteristics (size; location; total debts; degree of mechanization; working capital; specialization in labour intensive, seasonally and all year round, Types of Farming; presence of direct selling; dummies for organic farming and agro-tourism) and local economic characteristics (VA per inhabitant, as a

proxy for the potential wage; agricultural to total employment as a proxy for off-farm job opportunities). We introduced the amount of Single Farm Payments (SFP) annually received by individual farms to detect the effect of the CAP reform. It is both an indicator of the policy structural change (since it is zero before the reform implementation in 2005) and of the intensity of the intervention. Prior to the introduction of the SFP, eligible farms received partially coupled payments (per hectare or per animal head) represented by the relevant variable. The *ex ante* overall effect of these latter payments is ambiguous, since they may have both an income and a substitution effect.

4. Results

Table 2 reports the results of the different models. We will present the results in sequence, so to highlight the changes deriving from tackling the different econometric issues. We start with the estimation of a static pooled probit model that assumes no heterogeneity and no state dependence. The second model is a random effects dynamic model, including a past state and an initial state condition variable. Then we pass to dynamic settings accounting for heterogeneity and state dependence, as well as for the initial condition problem. All models are estimated for the 2004-2007 period, so to allow for comparisons. Heckman's, Wooldridge's and Orme's models also use information from year 2003 for the dynamic setting and for dealing with the initial year issue². All models are overall highly significant.

Column [1] gives the standard pooled probit estimates. This is tantamount to having cross-sectional estimates, so that it is in a sense a benchmark for "traditional" analyses of off-farm labour participation. Actually, the results are largely similar to the ones typically found in cross-sectional analyses. The probability that the farm operator participates in off-farm work is significantly affected by the idiosyncratic characteristics of the farmer (age, sex) and of the household (presence of pensions and of capital income) while no statistically significant influence is found for the socio-economic conditions at the provincial level. As for the farm characteristics, higher mechanization rates, higher percentages of owned land, the farm being located in mountainous areas and making use of direct selling, and a higher single farm payment increase the probability of working off the farm. On the contrary, a larger working capital significantly decreases the probability of the operator working off-farm. The same significant effect is found when the farm adopts organic farming techniques, as well as if it is specialized in the production of all-year-round labour-intensive agricultural products. We then move to dynamic settings

² The initial conditions equation in the Heckman's model is estimated with the explanatory variables set including the exogenous instrumental variable Value Added per employed in agriculture as a percentage of overall Value Added per employed.

Table 2-Estimates of the models of off-farm labour participation

	[1]		[2]		[3]		[4]		[5]		[6]	
	Pooled		Dynamic pooled		Dynamic RE		Heckman		Orme		Wooldridge	
	Coeff.	Std.E rr.	Coeff.	Std.E rr.	Coeff.	Std.Err	Coeff.	Std.E rr.	Coeff.	Std.E rr.	Coeff.	Std.Err
y(t-1)	-		2.650***	0.055	1.143***	0.121	1.417***	0.115	1.223***	0.118	1.234***	0.127
y(0)	-		-		2.953***	0.302	-		-	0.142	2.724***	0.293
<i>Personal characteristics</i>												
Operator's age	0.011	0.010	-0.011	0.012	-0.031	0.022	-0.015	0.022	-0.019	0.021	0.029	0.040
Operator's age squared	-0.000***	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Operator's gender (1= male)	0.258***	0.050	0.091	0.065	0.092	0.130	0.164	0.123	0.203	0.128	0.045	0.321
<i>Farm characteristics</i>												
UAA (ha)	-0.001	0.000	-0.001	0.001	-0.002	0.002	-0.001	0.001	-0.002	0.001	-0.004	0.006
Share of land in property (%)	0.370***	0.051	0.267***	0.067	0.410***	0.137	0.515***	0.128	0.390***	0.134	0.140	0.350
Working capital (1000 Euro)	-0.001***	0.000	-0.000*	0.000	0.000	0.000	-0.001***	0.000	-0.000**	0.000	0.000	0.001
Total Debts (1000 Euro)	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.002
Hp/UAA	0.001**	0.000	0.000	0.000	0.001	0.001	0.001	0.000	0.001	0.001	0.001	0.005
TF labour int. all year round (0,1)	-0.380***	0.048	-0.255***	0.064	-0.392***	0.123	-0.298***	0.109	-0.496***	0.121	0.129	0.239
TF labour int. seasonally (0,1)	-0.091**	0.043	-0.032	0.057	0.042	0.109	0.047	0.098	-0.011	0.107	0.258	0.223
Mountain (0,1)	0.107**	0.050	0.139**	0.066	0.272*	0.143	0.202	0.139	0.272**	0.138	0.276**	0.137
Hills (0,1)	0.027	0.042	0.048	0.057	0.181	0.124	0.124	0.109	0.135	0.121	0.140	0.121
Direct sales (0,1)	0.175***	0.043	0.109*	0.057	0.096	0.097	0.132	0.090	0.156	0.096	-0.019	0.133
Organic farming (0,1)	-0.296***	0.101	-0.191	0.133	-0.226	0.255	-0.172	0.217	-0.254	0.250	-0.135	0.401
Agro-tourism (0,1)	-0.094	0.116	-0.208	0.156	-0.356	0.288	-0.142	0.272	-0.320	0.279	-0.906	0.623

Table 2 (continued) -Estimates of the models of off-farm labour participation

	[1]		[2]		[3]		[4]		[5]		[6]	
	Pooled		Dynamic pooled		Dynamic RE		Heckman		Orme		Wooldridge	
	Coeff.	Std.Err.	Coeff.	Std.Err.	Coeff.	Std.Err.	Coeff.	Std.Err.	Coeff.	Std.Err.	Coeff.	Std.Err.
<i>Household characteristics</i>												
Pension income (0,1)	-		-		-		-		-		-	
	0.271***	0.049	0.362***	0.067	0.781***	0.102	0.771***	0.108	0.818***	0.102	-1.384***	0.127
Capital income (0,1)	0.641***	0.129	0.536***	0.166	0.903***	0.219	0.735***	0.259	0.868***	0.209	0.695**	0.336
<i>CAP reform</i>												
Coupled Payments (1000 Euro)	-0.002	0.002	-0.002	0.002	-0.002	0.006	-0.001	0.003	-0.002	0.005	-0.001	0.005
Single Farm Payment (1000 Euro)	0.002***	0.001	0.001*	0.001	0.002	0.002	0.002*	0.001	0.002	0.002	0.002	0.002
<i>Economic environment</i>												
Ag. to total employment (%)	0.001	0.006	0.007	0.008	0.021	0.016	0.009	0.015	0.017	0.016	-0.005	0.071
VA per inhabitant (1000 Euro)	0.000	0.005	0.009	0.006	0.027**	0.013	0.013	0.012	0.018	0.012	0.085**	0.039
<i>2004-2007 averages¹</i>												
Pension income (0,1)	-		-		-		-		-		1.205***	0.197
TF labour int. all year round (0,1)	-		-		-		-		-		-0.607**	0.273
Constant	1.660***	0.290	1.871***	0.373	3.000***	0.726	-2.773	0.661	2.759***	0.707	-2.839***	0.776
Rho	-		-		0.620***	0.048	0.619***	0.038	0.602***	0.049	0.573***	0.053
h (Orme)	-		-		-		-		1.367***		-	
theta (Heckman)	-		-		-		1.255***	0.147	-		-	
Log-likelihood	-3030.73		-1640.47		-1537.42		2269.182		-1545.95		-1509.54	

¹ Only significant four-year averages variables are reported

accounting for the effect of y_{t-1} , i.e. the past off-farm labour state. In all of them, the effect of the past state is strong and highly significant.

Column [2] gives the dynamic pooled probit estimates. This model accounts for state dependence, but does not deal with heterogeneity nor with the issue of the initial conditions. The parameter related to the past state is strongly significant and large. In addition, this model, as compared to the pooled sample model, presents a dramatic increase in the log-likelihood. A likelihood ratio test strongly rejects the restriction implied by the static pooled sample. This means that a dynamic model is statistically superior to the cross sectional one.

Important changes occur even in the covariates since the parameters of all individual characteristics and of several farm characteristics are now no more statistically significant^{3,4}.

The third model adds to the past state the state in the initial year and allows for individual heterogeneity. As compared to the previous one, the parameter of the past state is strongly reduced, since its effect is partly absorbed by the initial year state parameter and by heterogeneity, represented by the correlation term, which are both highly significant.

The following three models control, in different ways, the issue of the initial conditions, while still allowing for both state dependence and heterogeneity. The results are to a large extent similar among them. In all three models the past state is highly significant, and its effect is larger than in model 3. To compare the parameters of random effects models, they must be rescaled by multiplying them by $\sqrt{1 - \rho}$, where ρ is the constant cross-period error correlation (Arumpalam, 1999). After rescaling, the past state parameters for the Heckman's, Wooldridge's, and Orme's models are of comparable magnitude, i.e., 0.874, 0.763 and 0.806 respectively, as compared to 0.704 for model 3. I.e., not controlling for the initial conditions would lead to underestimate the effect of the past state. Also, the values of the correlation coefficient, measuring the effect of unobserved heterogeneity, are of similar magnitude (0.619, 0.604, and 0.573, respectively, for Heckman's, Wooldridge's, and Orme's models). Third, all parameters involved in the correction of the potential bias due to the initial conditions are significant: the *theta* parameter in Heckman's model, the *h* parameters in Orme's, and the initial state, as well as some means of the explanatory variables in Wooldridge's model. From the methodological point of view, this suggests that coping with the initial conditions issue is important to avoid biased estimates. Also, the results suggest that both true state dependence and heterogeneity are important in determining persistence. Heterogeneity represents individual unobservable factors, but the effects of an increase and of a decrease are symmetrical like the other covariates, and the issue of heterogeneity is basically an econometric one, implying biased estimation if not properly controlled. By contrast, true state dependence creates an asymmetry in farmers' response, in the sense that once a change in the work state has taken place, a reversal has not the same effect with an inverted sign. Therefore, true state dependence is an economic issue, stemming from true changes in farm setting. Existence of true state dependence is not simply an econometric issue, it is a behavioural issue.

³ This is not due to collinearity between the past state variable and the covariates that result statistically significant in the pooled model. Inspection of the correlation matrix shows that the correlation coefficient between these covariates and the past state variable is at most 0.06, and below 0.03 for most of them.

⁴ If heterogeneity is added to this model, i.e., if the model including the past state among the explanatory variables is estimated as a random effects model, the estimate of the correlation coefficient is zero. When estimating a random effects model without the lagged state variable, the model does not reach convergence. Since the following models, controlling for state dependence, heterogeneity, and initial conditions do converge, the implication is that controlling for the initial conditions problem is crucial for distinguishing between state dependence and heterogeneity.

To measure the effect of past off-farm labour state on present participation, we calculated its Marginal effect (ME) and the Average Partial Effect (APE). The former is an estimate of the change in probability of the outcome due to a unit change of the relevant dummy variable, evaluated at the mean values of the independent variables (or at their median, in case of dummy variables) and can be interpreted as the effect of the relevant variable for a “representative” farm. The latter is an average of the difference between the probability of the outcome when the past state is set to one and when it is set to zero for each farm at the actual values of the variables. The relevant results of the MEs and APEs are 0.074 and 0.077, respectively, for the Heckman model; 0.058 and 0.048 for the Wooldridge model; 0.054 and 0.008 for the Orme model. They therefore suggest that if the past state was participation, the probability that the present state be participation is 5 to 8 percent higher (only the ME of the Orme model are lower, 0.008 percent). Notice that in the model not accounting for heterogeneity nor for the initial state, the percentages are much higher (17.5 for APE, 68.6 for ME): the past state variable absorbs also the effect of unobservable idiosyncratic characteristics. Hence, the comparison with the models accounting for heterogeneity suggests that an appreciable part of persistence is also due to heterogeneity.

As to the other explanatory variables, some results are common to all three models. The two variables related to the household income (pensions and capital income) are strongly significant in all these models. Non-labour income (capital and pension income), is –admittedly poorly-measured by dummy variables. If the household has any capital income, the probability of off-farm labour participation is significantly increased. Since theoretically non-labour income should reduce off-farm work, a possible interpretation of this result is that off-farm employment is induced by off-farm high wages that, in turn, provide more financial assets that yield capital incomes. This probably concerns the wealthiest farms. If a household member has a pension income, the probability that the farm operator has an off-farm job significantly decreases. A possible interpretation is that for poor farm households, off-farm income stemming from pensions of other members decreases the income needs, thus raising the off-farm reservation wage of the operator. In alternative, it is possible that the presence in the household of elders and individuals with a disability increases the time to be devoted to caring activities, in this way discouraging off-farm work.

For other variables, the results are less homogeneous across the different models. For instance, the farm location in mountain areas is significant in Orme’s and Wooldridge’s models, but not in Heckman’s. Farms in those areas are typically less profitable, which might induce more pluriactivity. In addition, the Value Added per inhabitant is significant and positive in Wooldridge’s model. This result is counter-intuitive, since one might argue that a higher income gap between agriculture and the overall economy should induce more off-farm labour participation. An interpretation could be that the most profitable agricultural areas (especially in the North of Italy) are also the ones where more job opportunities are available. All-year-round labour-intensive types of farming are found to significantly decrease the likelihood of working off the farm, though this result is not significant in Wooldridge’s model (which is the one in which less variables are significant). The share of land in property also has a significant positive effect in the same models, probably because this share is generally larger in small farms where participation is usually more frequent. Finally, the amount of working capital is found to be significant and negative, as expected, in Heckman’s and Orme’s models. But in both cases the effect is very small, so that in economic terms it is negligible.

The most striking result –also robust across all models- is that, after introducing state dependence and heterogeneity and controlling for the initial state conditions, in none of these dynamic models the idiosyncratic farmer characteristics have statistically significant effect on the probability that the operator works off the farm. Also, many of the farm characteristics found to be significant in cross-sectional models are no more significant in this dynamic setting. Though

further analyses based on other samples and for other countries are needed to make this conclusion general, this makes a remarkable difference with the current literature based on cross-sectional samples, where typically personal, farm and labour market characteristics are found to be important determinants.

Finally, neither the variable of Coupled Payments (before the CAP reform) nor the one of the Single Farm Payment (after the reform) exhibit significant effects in the three models, apart from SFP in the Heckman's model, in which it is only weakly significant. Again, this makes a difference with the pooled model, for which the SFP variable is strongly significant. Anyway, the sign is negative for semi-coupled payments, and positive for SFP. The former sign would be consistent with theoretical predictions, and would suggest that the substitution effect of coupled payments dominates the wealth effect. By contrast, the sign of the latter would contradict the theoretical expectations, at least if production choices were separable from consumption and family labour allocation choices (if not, there are no a priori predictions). These results on Italian farms are consistent with the results of US studies, that found indeed very small effects of decoupled payments on farm and off-farm labour and on total work hours (Dewbre and Mishra, 2002; El-Hostaet *al.* 2004; Goodwin and Mishra 2004; Ahearn *et al.* 2006; Serra *et al.* 2005). The CAP reform explicitly tried to substitute decoupled for coupled payments without substantially changing total expenditure. Hence, even from a theoretical perspective, its effect is ambiguous (Corsi, 2007 and 2008) and its weak effects on off-farm participation, if any, are not surprising.

5. Summary and Conclusions

In this paper we examine the issue of off-farm labour participation of farm operators taking into consideration the possible dynamic effects of past off-farm labour participation. We estimate different models allowing for both heterogeneity and true state dependence and coping with the problem of the initial conditions, using a panel sample of Italian farms.

The results show that cross-sectional analyses are unable to properly represent farmers' behaviour, since a strong persistence in the state can be observed. The conclusion is that past labour state is a crucial determinant of off-farm labour participation choices of Italian farmers. Persistence is also explained by unobserved heterogeneity. Variables usually found significant determinants of these choices, and also appearing as such in pooled sample estimation, are no more significant when true state dependence and heterogeneity are controlled for. We also find that the effects of the change in CAP, as well as more generally of coupled and decoupled payments, are weak and generally insignificant.

Our results have implications both for research and for policy. Though more research is needed to generalize our results, they are to a certain extent at odds with a whole stream of literature based on cross-sectional samples (including some of our work). The reasons for the discrepancy are nevertheless unclear. It seems that the past state and the unobservable idiosyncratic characteristics explain most of the determinants of off-farm labour participation. A research line that could be tackled to clarify this point might therefore be modelling changes of labour state rather than states *per se*. This is left to further research.

From the policy point of view, our results suggest that, given state dependence, if any policy has to be implemented concerning off-farm labour participation of farmers, an initial effort is needed to change it, but after this initial treatment, incentives are less needed to sustain in time the desired behaviour. Our results also suggest that the unobservable specific characteristics of farms are at stake in determining persistence. This makes identifying the appropriate levers on which to operate more difficult. And, consistently with American studies, the effects of both decoupled and coupled subsidies on off-farm work participation of Italian farmers appear to be weak.

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