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Full Research Article

Once part-timer always part-timer? Causes for persistence in off farm work state of farmers

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Abstract. Off-farm labour participation is an important way in which farm households adapt their labour resources to farm labour needs, and is often viewed as an income integration and an insurance against risk. Nevertheless, it has also been questioned as a step for exiting agriculture. It is therefore important to assess whether or not it is a permanent status and which are its determinants. Most papers on this issue are based on cross-sectional analyses and thus disregard the problem of persistence in the state. Using a 5-wave panel of Italian family farms we estimate different dynamic nonlinear panel data models of the determinants of off-farm labour participation. We allow for two sources of persistence: unobserved heterogeneity and state dependence, and in addition we control for the initial conditions problem. We find a strong persistence in the state and our findings show that, when taking all these features into account, the present work state is almost totally explained by the previous state and by idiosyncratic characteristics. The variables concerning the farm and the farmer's characteristics, typically found to be relevant in cross-sectional analyses, are not significant in the dynamic setting. The reasons for the inconsistency between our results and those of cross-sectional studies are discussed, and an interpretation of how the determinants influence the off-farm labour participation is presented. The distinction between true state dependence and individual heterogeneity has important policy implications that are discussed.

Keywords. Off-farm work, farm household, state dependence, panel data, CAP reform.

JEL codes. J220, J430, Q120.

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1. Introduction

Agriculture in most Western countries is characterized by an overwhelming share of family farms, which means that farm operation and farm household are tightly linked[1]. In particular, the labour allocation of household members between farm and off-farm labour is an important choice for the household. Indeed, the off-farm labour participation of farmers is a structural feature of both developed and developing economies, and off-farm income is an increasing part of farm household income (Oecd, 2003; Eurostat, 2002). It is part of the secular trend in declining agricultural employment, as productivity in agriculture increases and, due to the finite nature of land, no additional productive and employment basis can be created in the sector (except for intensification). Combining on- and off-farm work, at the individual and/or at the household level (pluriactivity), may indeed be an efficient use of the labour resources of households. Taking into account income opportunities stemming from the farm and from alternative employments allows the growth and stabilization of household income. Nevertheless, part-time farming has also been considered for its influence on the prospects of farms. A discussed related issue is whether it favours the survival of family farms or it is a step towards the exit from farming, with contrasting results. For instance, Weiss (1999) found a negative impact of part-time farming on farm survival, while Breustedt and Glauben (2007) found the opposite. These papers are part of a stream of literature dealing with the changes in agricultural labour. This stream is theoretically based on Todaro's (1969) and Harris and Todaro's (1970) model of labour migration, extended by Barkley (1990), and uses aggregated data (e.g. Breustedt and Glauben, 2007; D'Antoni *et al.*, 2012; Petrick and Zier, 2011, 2012; Olper *et al.*, 2014). Their focus is on the net exits from farming, and they do not deal with the contemporary presence of on-and off-farm work.

Our focus is rather on the process of combining on-and off-farm labour and in our perspective it is important to ascertain whether off-farm labour participation is a permanent status. One important feature in this respect is the flexibility of off-farm work status, i.e., to what extent farmers can and do enter and exit an off-farm work status, and for which reasons. Hence, our theoretical reference is the literature on farm household models. Starting from the seminal paper by Huffman (1980), labour choices in farm households have been widely analysed, initially considering the off-farm labour participation of farm operators (Sumner, 1982) and subsequently 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 Gempesaw, 1992), in addition to including the use of waged labour (Benjamin *et al.*, 1996; Blanc *et al.*, 2006) and the on- and off-farm labour participation of operators, spouses and other household members (Kimhi, 1994, 2004; Benjamin and Kimhi, 2006; Bjørnsen and Biørn, 2010; Corsi and Salvioni, 2012; Biørn and Bjørnsen, 2015). Most of the research has concerned the U.S., but several analyses are also devoted to the determinants of off-farm labour participation in the EU (Corsi, 1994; Benjamin *et al.*, 1996; Weiss, 1997; Woldehanna *et al.*, 2000; Juvančič and Erjavec, 2005; Benjamin and Kimhi, 2006; Salvioni *et al.*, 2008; Bjørnsen and Biørn, 2010; Corsi and Salvioni, 2012; Biørn and Bjørnsen 2015).

[1] The high supervision costs of hired labour, due to the technical features of agricultural production, have been noted to explain this situation: by contrast, family labour does not require supervision because family members are involved in the income that it provides (Pollack, 1985).

A common feature of this stream of research, with a few exceptions, is that the analyses are based on cross-sectional samples[2]. This approach nevertheless prevents analysing the dynamic nature of off-farm labour participation choices and disregards the persistence of the phenomenon. Persistence means that individuals that have experienced the event under study in the past are more likely to experience it in the future than those who have not experienced it in the past. However, the reason for persistence can be true state dependence and/or heterogeneity (Heckman, 1981a). True state dependence exists when the previous state modifies the attitudes, constraints or parameters so that the probability of the present state is affected by the previous state. True state dependence has a truly behavioural origin. For instance, when taking on an off-farm job implies sunk search costs, this creates a difference between the choice of taking on a new off-farm job and the choice of continuing keeping an off-farm job. Or, if an off-farm job is taken on, there might be costs related to the adaptation of the farm operation to the new situation (e.g., higher mechanization or a change in the type of farming or in the farming intensity to cope with lower family labour input), costs that are not incurred if the farmer already has an off-farm job. Obviously, the same holds for the change in opposite direction, i.e., from off-farm occupation to no off-farm occupation. This makes the continuation of the off-farm state more likely than its change.

Heterogeneity represents the unobservable time-invariant idiosyncratic characteristics that affect the choices across time periods. While some time-invariant variables influencing the choice can be observed and hence controlled for, some are not measured or are inherently unobservable, such as, for instance, a farmer's taste or aversion for off-farm work. Such variables create a permanent push towards a specific state, so that the previous state apparently affects the future one. Persistence and its reason are relevant for policy implications. Weiss (1997) notes that "asymmetric adjustment behaviour may cause serious problems for designing an appropriate policy to encourage (or impede) a specific form of agricultural production such as part-time farming", in particular because policies cannot be easily reversed. By contrast, as noted by Corsi and Findeis (2000), if heterogeneity is the principal source of state persistence, then more traditional instruments can be used. Persistence has actually been initially analysed in two waves panel data (Gould and Saupe, 1989; Weiss, 1997; Corsi and Findeis, 2000; Ahituv and Kimhi, 2002; Juvančič and Erjavec, 2005) and, more recently, in long span studies (Bjørnsen and Biørn, 2010 and 2015) accounting for both state dependence and heterogeneity. However, in the case of state dependence it is important to control for both the observed and unobserved determinants of initial participation status. It has been shown (Arulampalam *et al.*, 2000) that even if the model controls for unobserved heterogeneity, in order to disentangle the effect of state dependence from unobserved heterogeneity, the initial conditions need to be modelled instead of being assumed as exogenously given, since they may be correlated with the unobservables. Not accounting for the initial conditions may lead to biased estimates of the effects of true state dependence and individual heterogeneity which, in turn, has important policy implications.

[2] Even the most recent research addressing issues related to off-farm labour participation like, e.g., the interrelationship between off-farm work and agro-tourism (Khanal and Mishra, 2014), or between off-farm labour market decisions and agricultural shocks (Mathenge and Tschirley, 2015) use cross-sectional data.

This paper analyses the off-farm labour participation with a dynamic panel data model[3] allowing for both heterogeneity and true state dependence, and handling the initial conditions problem. Since our estimation is performed on a panel sample of Italian farms for the 2003-2007 period and in this period a major reform of the Common Agricultural Policy (CAP) occurred, this study also accounts for the relevant changes[4].

This paper is structured as follows. Section 2 presents the theoretical framework and the econometric strategy. In Section 3, the data on which the analysis is based are presented. The estimation results are presented and discussed in Section 4, both from the econometric and the factual perspective. The paper is concluded by some final considerations.

2. Theoretical model and econometric strategy

2.1 Cross-sectional modelling

The theoretical model typically employed to analyse farmers' off-farm labour choices is the well-known farm-household model (Nakajima, 1986; Singh *et al*., 1986; Huffmann, 1991), which can be adapted according to Corsi (2007 and 2008) to take into account the CAP reform. The model assumes that the farmer maximizes utility over consumption and leisure, under constraints of income and time. The income constraint comprises both farm income and off-farm wages. The Kuhn-Tucker maximization conditions yield the following first-order conditions:

$$\partial(p+s)Q / \partial F \leq \mu/\lambda \tag{1}$$

$$w \leq \mu/\lambda \tag{2}$$

where Q is the quantity of good produced by the farm; p is its price; s is the coupled payment per unit of output; F is time spent working on the farm; w is the off-farm market wage; and μ and λ are the marginal utilities of leisure and income, respectively.

In a pluriactive farm, equations (1) and (2) hold as equalities: that is, the marginal value product of farm labour (inclusive of the coupled support) is equal to the market wage and to the leisure-income Marginal Rate of Substitution (MRS)[5].

[3] All panel data model are dynamic in that they take into account the time series dimension of the sample. However, functions that specifically model the effect of lagged dependent variables are usually referred to as dynamic panel data models.

[4] The 2003 reform (the so-called Fischler reform) has been the most important step in the process of change in the CAP from trade and market distorting measures to more neutral interventions, leading to a more decoupled, and hence less market distorting, support (Oecd, 2004). Its most important measure was replacing the semi-coupled subsidies (i.e., subsidies per hectare of specific crops and per animal head) with the fully decoupled Single Farm Payment Scheme (SFP), providing income support to farmers regardless of their current production decisions. In Italy the decoupling was immediately implemented in the first possible year, 2005. Furthermore, disregarding the possibility of introducing forms of regional distribution for the direct subsidies, Italy decided to adopt the so-called historical criterion: farmers were granted their SFP according to an entitlement based on the semi-coupled subsidies that they received in the 2000-2002 reference period.

[5] Using this model, the impact of a direct payment has long been established (El-Osta *et al*., 2004; Ahearn *et al*., 2006; Corsi, 2007 and 2008; Hennessy and Rehman, 2008): a decoupled direct payment is tantamount to an increase in non-labour income. Hence, for a farmer participating in off-farm work, a decoupled direct payment

The reduced form typically employed in analyses of off-farm labour participation starts by defining the reservation wage w^*, obtained by setting the off-farm labour supply to zero and solving for w. The reservation wage is compared to the market wage w, which is a function of personal characteristics and the conditions of the local labour market. Therefore, participation (indicated by a dichotomous variable y) occurs if $w>w^*$, which means that prob(y=1) = prob($w>w^*$). From the theoretical model, w^* is a function of the price of agricultural products, variables affecting on-farm labour productivity, and taste shifters. The market wage is a function of personal characteristics that influence human capital and labour market characteristics. Typically, the reduced-form equations are probit or logit equations that estimate the influence of these variables on the probability of (w-w^*) being greater than zero. Four categories of explanatory variables are typically included: individual[6], household, farm and local market characteristics.

2.2 Dynamic modelling

In a dynamic setting, expected utility and expected income streams have to be considered. Moreover, the costs for shifting from one work condition to the other are to be taken into account. To a large extent, these costs are sunk costs, such as, for instance, job search costs and costs for adapting the farm to the new work status and the implied labour needs.

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 that the farmer could earn by working off the farm in time t; f_{1t} and f_{2t} the farm income that 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 is (for simplicity, we consider utility as only stemming from income; non-pecuniary benefits from different jobs can be easily incorporated):

$$prob\left(y_t=1 \mid y_{t-1}=0\right)=prob\left\{E\left[\sum_{t=1}^{T}\left(w_t+f_{1t}-f_{2t}\right)\frac{1}{(1+r)^t}-C_1\right]>0\right\} \tag{3}$$

and the probability of continuing not participating is:

$$prob\left(y_t=0 \mid y_{t-1}=0\right)=prob\left\{E\left[\sum_{t=1}^{T}\left(w_t+f_{1t}-f_{2t}\right)\frac{1}{(1+r)^t}-C_1\right]\leq 0\right\} \tag{4}$$

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will decrease off-farm work. However, if production and labour allocation decisions are not separate, then it is unclear which type of work – on-farm or off-farm – will be reduced. By contrast, a decrease in coupled support has both a wealth and a substitution effect. The decrease in the marginal value product of family farm labour induces a reduction of on-farm work. Simultaneously, the decrease in income decreases the MRS, which means that the farmer consumes less leisure. Hence, the overall result is an increase in off-farm work. Given that the CAP reform is a combination of an income payment and a decrease in the average revenue, the two effects operate in opposite directions.

[6] Observable personal characteristics are taken as proxies for individual idiosyncratic preference shifters; some attempt has been undertaken to capture non-pecuniary benefits of farm work through personal statements (Howley *et al.*, 2012).

The probability of not participating in time t for a farmer who participates in time t-1 is:

$$prob\left(y_t = 0 \mid y_{t-1} = 1\right) = prob\left\{E\left[\sum_{t=1}^{T}\left(f_{2t} - w_t - f_{1t}\right)\frac{1}{\left(1+r\right)^t} - C_2\right] > 0\right\} \tag{5}$$

or:

$$prob\left(y_t = 0 \mid y_{t-1} = 1\right) = prob\left\{E\left[\sum_{t=1}^{T}\left(w_t + f_{1t} - f_{2t}\right)\frac{1}{\left(1+r\right)^t} + C_2\right] < 0\right\} \tag{6}$$

Hence, the probability of continuing to participate for a farmer already participating is:

$$prob\left(y_t = 1 \mid y_{t-1} = 1\right) = prob\left\{E\left[\sum_{t=1}^{T}\left(w_t + f_{1t} - f_{2t}\right)\frac{1}{\left(1+r\right)^t} + C_2\right] \geq 0\right\} \tag{7}$$

From a comparison of (3) to (7), it can be concluded that, for every strictly positive C_1 or C_2, $prob(y_t=1|y_{t-1}=1) > prob(y_t=1|y_{t-1}=0)$. Hence, costs for shifting from one condition to another create a state dependence, and not including them among the explanatory variables would lead to an omitted variable bias. Since potential shifting costs are not observable, the issue can be addressed by including the 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.

A further reason for observed persistence in the labour state can be heterogeneity, i.e., farmers' or farms' unobservable idiosyncratic characteristics that make participation more or less likely; thus, when observing panel data, heterogeneity translates into an apparent dependence of the present state on the previous state (spurious state dependence).

2.3 Econometric model

Unlike most previous studies, we exploit the panel nature of our data to analyse the determinants of off-farm labour participation, controlling for both true state dependence, created by shifting costs, and heterogeneity, and handling the initial conditions problem. The probability that a farm operator will work off the farm is estimated by applying a dynamic non-linear (probit) random effects model that accounts for both unobserved heterogeneity and true state dependence.

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

$$Prob(y_{it}=1) = \Phi(\beta_0 x_{it} + \gamma y_{it-1} + u_i + \varepsilon_{it}) \tag{8}$$

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,…,T), and 0 otherwise, so that y_{it-1} is used to test the effect of true state dependence; Φ is the normal cumulative density function; x is a vector of observable time-variant and time-invariant explanatory variables; β_0 and γ are parameters to be estimated; u_i is an individual idiosyncratic term that represents time-constant unobservable individual characteristics, i.e., heterogeneity; and ε_{it} is a random component, uncorrelated across individuals and years.

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, and hence of the magnitude of state dependence, is the initial conditions problem – whether y_{i1} is independent of u_i. The problem is caused by the presence in the equation of both the past value of the dependent variable and of an unobserved heterogeneity term and the correlation between them. The treatment of initial conditions is crucial because misspecification will result in an inflated parameter of the lagged dependent variable that measures the magnitude of the cost of shifting employment.

Different estimators that cope with these issues for the nonlinear dynamic panel data model have been proposed: Heckman (1981), Wooldridge (2005) and Orme (2001). 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 most cases, all three estimators display a satisfactory performance, except when the number of time periods is very small. We estimated all three models to test the robustness of our results, but for the sake of brevity, we present here only Wooldridge's (2001) model. The results of the other models are largely consistent with the results that we present and are reported in the Appendix.

In Wooldridge's approach, the individual term is assumed to be a function of the initial condition and the means of the explanatory variables[7]:

$$u_i = \alpha_0 + \alpha_1 y_{i1} + \alpha_2 \mu_i + \nu_i \quad (9)$$

where y_{i1} is the state in the first observed year; μ_i is a vector of the means of the explanatory variables over the entire observed period; and ν_i is a random individual component, uncorrelated across individuals.

The overall model is therefore:

$$Prob(y_{it}=1) = \Phi(\alpha_0 + \beta_0 x_{it} + \gamma y_{it-1} + \alpha_1 y_{i1} + \alpha_2 \mu_i + \nu_i + \varepsilon_{it}) \quad (10)$$

The model is estimated as a random effects probit by a maximum likelihood method.

3. Data

This study relies on data collected by the Italian Farm Accountancy Data Network (FADN) survey. The survey started to be conducted on a statistically representative basis

[7] To reduce the number of parameters to be estimated, we adopt the popular version of Wooldridge's (2005) solution to the initial conditions problem, and we use the within means of time-varying explanatory variables calculated over the 4 years following the initial one. Skrondal and Rabe-Hesketh (2014) showed that for periods larger than 3 this model does not produce significant bias.

in 2003. The sample is stratified according to criteria of geographical region, economic size (European Size Units -ESU) and type of farm (TF)[8]. The field of observation is the total of commercial farms, that is, farms with an economic size greater than 4 ESU (4,800 euro)[9].

In this study, we employ a 5-wave balanced panel of 3,294 farms for which data were collected in all years from 2003 to 2007. We only kept family farms[10]. The models are estimated over the 2004-2007 period because the 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[11].

The dependent variable is a dummy variable that indicates whether the farm operator works off the farm. The share of farms operators who have an off-farm job is 6.6 percent over the entire 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. These percentages are much lower than the participation in off-farm work revealed by the USDA's Agricultural Resource Management Survey data (fifty-two percent of farm operators) in the US (Fernandez-Cornejo, 2007) and by the Farm Structure Survey (thirty-six percent of EU-27 family farm managers) in the EU (European Commission, 2008). The lower participation rate recorded in the Italian FADN is partly because the FADN refers to professional, larger farms that typically participate less in off-farm employment and partly because pluriactivity is less widespread in the southern countries of the EU (European Commission, 2008).

Following previous research, we use four categories of explanatory variables to specify the model for the off-farm participation decision: individual, household, farm and local market characteristics.

Individual attributes include age, age squared and gender of the operator; it was not possible to control for the effect of education because this information is not collected by the survey. Previous research with cross-sectional data typically has found a positive effect of age on participation but with a curvilinear pattern (shown by the age squared variable), reaching a peak and then declining.

The household non-labour incomes, i.e., capital income and pensions, are used to explore the existence of a wealth effect. Non-labour income (capital and pension income), is –admittedly, poorly – measured by dummy variables. Larger values are theoretically expected to have a negative effect on off-farm participation. Unfortunately, we can only measure the presence of these types of income with dummy variables, and we have no information on the number and characteristics of other household members.

Farm characteristics include farm size (in hectares); farm location (mountain, hills and plains as base category); total debts in thousand Euro; degree of mechanization (horse

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[8] FADN defines a farm as specialized in a TF if the Standard Gross Margin (SGM) for the particular production covers more than 2/3 of the total SGM. SGMs are obtained on the basis of the farm area (number of heads) and a crop (livestock) and area specific standardised gross margin.

[9] Small farms are more interested in pluriactivity. Though 4 ESU (4,800 Euro) is a small size, the exclusion of farms below this threshold leads to underestimate the importance of the phenomenon. The results should therefore be interpreted as concerning the observation field of FADN.

[10] Farms that were not characterized as sole ownership or private partnership were not considered in this analysis.

[11] Ours is an artificial panel data set obtained from a series of cross-sectional surveys. Hence, the weights provided by the FADN, which are based on Standard Gross Margin, Standard Output, and costs, are no more valid for inference to the general population. Therefore, we decided to apply no weight to the observations in our sample. The generalization of our results to the general population should therefore be performed with some caution.

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

	Mean	Std.Dev.
Share of off-farm labour participants 2003-2007	0.07	0.07
Share of off-farm labour participants 2003	0.06	0.01
Share of off-farm labour participants 2004	0.06	0.01
Share of off-farm labour participants 2005	0.07	0.01
Share of off-farm labour participants 2006	0.07	0.01
Share of off-farm labour participants 2007	0.07	0.02
Personal characteristics		
Operator's age	54.19	13.60
Operator's age squared	3121.84	1513.44
Operator's gender (1=M, 0=F)	0.83	0.38
Farm characteristics		
UAA (ha)	33.70	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	0.50
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
Household characteristics		
Pension income (0/1)	0.23	0.42
Capital income (0/1)	0.01	0.11
CAP reform		
Coupled Payments (1000 Euro)	3.48	33.79
Single Farm Payment (1000 Euro)	7.84	33.21
Economic environment		
Agricultural to total employment (%)	6.35	4.03
VA per inhabitant (1000 Euro)	21.72	5.23

power per hectare of used land); working capital in thousand Euro; specialization in the production of labour intensive, seasonal and all-year-round Types of Farming; the presence of direct selling; and dummies for organic farming and agro-tourism. Farm size is typically expected to decrease off-farm participation because larger farms are usually more profitable, which makes off-farm labour comparatively less attractive. Mountain and hill farms are typically characterized by low returns. Accordingly, the farm location in mountains and hills should be expected to positively affect the off-farm participation of farm household members attempting to increase total household income with alternative off-farm incomes. Nevertheless, these areas typically also provide less employment opportunities, which would have the opposite effect. The higher the debt is, the higher the interest that farmers have to pay to lenders. This situation might provide an incentive to farm households to work off-farm to find new sources of income to pay back the debt. There are no clear theoretical expectations regarding the sign of the coefficient of mechanization. The use of machines could reduce the labour hours required on farms and therefore increase the probability of off-farm labour participation. By contrast, a high investment in machinery could be a sign of the deep commitment of the household in farm activities and, more importantly, could raise farm income, thus discouraging off-farm labour participation. High levels of working capital are expected to increase productivity and to lower off-farm labour attractiveness. Off-farm work is also expected to be discouraged in those cases in which farms are specialized in all-year-round labour intensive and, to a lesser extent, in seasonal labour intensive types of faming. The use of organic farming is anticipated to reduce the likelihood of off-farm work, given the higher labour requirements of these farming systems as compared to conventional farming. Finally, direct selling, in addition to offering agro-tourism services, is labour intensive and should reduce off-farm activities[12].

The variables that describe the local labour market and the external economic environment are at the provincial level (Provinces are administrative bodies that correspond to the NUTS-3 level of Eurostat). All data are drawn from national accounting data published by Istat, i.e., the national official statistical agency. The agricultural to total employment ratio is introduced to account for employment opportunities outside agriculture[13]. Value Added per inhabitant is an overall measure of the degree of economic development of the area where the farm is located and is also a proxy for the opportunity cost of agricultural labour. These two variables are expected to positively influence the probability of off-farm participation because the higher these ratios, the more job opportunities are available.

The variable called Single Farm Payments (SFP) is equal to zero before the CAP reform implementation in 2005, and is equal to the relevant payment in the following years. Hence, it is an indicator of both the policy structural change and of the intensity of the intervention. Given that SFP is allocated based on the historical production mix, it can be considered as exogenous (indeed, decoupling aims at making public support exogenous to production choices). The amount of the SFP is farm-specific, and its impact is a pure

[12] Some of these variables (in particular, the share of rented land, the mechanization variable, and the semi-coupled payments, presented below) may be suspected of being endogenous because, in principle, the off-farm labour choice can influence them. We tested the use of lagged values of these explanatory variables to overcome the possible issue, but the results were almost identical to those presented here. The results are available from the authors upon request.

[13] It would have been desirable to introduce other important variables concerning local labour markets, the activity rate and the employment and unemployment rates. Unfortunately, in 2004 Istat changed the methodology of the labour force surveys; thus, the series from 2004 onward is not comparable to that of previous years. For this reason, we could not add these variables. Regardless, if these variables are introduced in the models for the 2004-2007 period, they are never significant.

wealth effect. 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 overall effect of these latter payments is ambiguous because they may have both an income and a substitution effect.

It is important to note that all the above expected effects of the explanatory variables refer to the static farm household model. Hence, they are what the theory predicts when changing state implies no cost and no change in farm setting.

4. Results

4.1 Econometric results

Table 2 shows the results of the estimates. In addition to the results of Wooldridge's model, we report the estimation of a static pooled probit model that assumes no heterogeneity and no state dependence. This estimation is tantamount to having cross-sectional estimates, which means that, in a sense, it is a benchmark for "traditional" cross-sectional analyses of off-farm labour participation. Both models are estimated for the 2004-2007 period to allow comparisons. We also report the estimates of the marginal effects of the variables on the probability of off-farm labour participation. As usual, they are estimated at the mean values of the continuous variables and at the median value for the dichotomous variables.

Column [1] gives the standard pooled probit estimates. Actually, the results are largely similar to those typically found in cross-sectional analyses. The probability that the farm operator participates in off-farm work is found to be affected significantly by the idiosyncratic characteristics of the farmer (age, sex) and of her household (the presence of pensions and of capital income), whereas no statistically significant influence is found for the socio-economic conditions at the provincial level. Regarding 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-farm. By contrast, the use of organic farming techniques, higher percentages of working capital and being specialized in the production of all-year-round labor-intensive agricultural products significantly decrease the probability of the operator working off-farm. All these results are commonly found in cross-sectional studies.

Column [2] gives the estimates of Wooldridge's model, which accounts for the effect of y_{t-1}, i.e., the past off-farm labour state (state dependence), and controls for the initial conditions and heterogeneity[14]. The past state parameter is highly significant[15]. The cor-

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[14] We also estimated two further models. The first model (dynamic pooled probit) accounted for state dependence but did not address heterogeneity or the issue of the initial conditions. The parameter related to the past state was strongly significant and large. In addition, this model, compared to the pooled sample model, presented a dramatic increase in the log-likelihood, and a likelihood ratio test strongly rejected the restriction implied by the static pooled sample. Important changes also occurred even in the covariates, because the parameters of all individual characteristics and of several farm characteristics were no more statistically significant.. A further model (dynamic random effects) added the state in the initial year and allowed individual heterogeneity. These results are presented in the Appendix.

[15] As a robustness check, the past state parameter has been compared to the estimates of the other models. To compare the parameters of random effects models, they must be rescaled by multiplying them by $\sqrt{1-\rho}$ where

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

	[1] Pooled			[2] Wooldridge		
	Coeff.	Std.Err.	Marg. Eff.	Coeff.	Std.Err.	Marg. Eff.
y(t-1)	-			1.234***	0.127	0.048
y(0)	-			2.724***	0.293	0.107
Personal characteristics						
Operator's age	0.011	0.010		0.029	0.040	0.001
Operator's age squared	-0.000***	0.000		0.000	0.000	0.000
Operator's gender (1= male)	0.258***	0.050		0.045	0.321	0.002
Farm characteristics						
UAA (ha)	-0.001	0.000	-0.000	-0.004	0.006	0.000
Share of land in property (%)	0.370***	0.051	0.040	0.140	0.350	0.006
Working capital (1000 Euro)	-0.001***	0.000	-0.000	0.000	0.001	0.000
Total Debts (1000 Euro)	0.000	0.000	0.000	0.000	0.002	0.000
Hp/UAA	0.001**	0.000	0.000	0.001	0.005	0.000
TF labor int. all year round (0,1)	-0.380***	0.048	-0.038	0.129	0.239	0.005
TF labor int. seasonally (0,1)	-0.091**	0.043	-0.010	0.258	0.223	0.010
Mountain (0,1)	0.107**	0.050	0.012	0.276**	0.137	0.011
Hills (0,1)	0.027	0.042	0.003	0.140	0.121	0.006
Direct sales (0,1)	0.175***	0.043	0.021	-0.019	0.133	-0.001
Organic farming (0,1)	-0.296***	0.101	-0.026	-0.135	0.401	-0.005
Agritourism (0,1)	-0.094	0.116	-0.009	-0.906	0.623	-0.036
Household characteristics						
Pension income (0,1)	-0.271***	0.049	-0.026	-1.384***	0.127	-0.054
Capital income (0,1)	0.641***	0.129	0.111	0.695**	0.336	0.027
CAP reform						
Coupled Payments (1000 Euro)	-0.002	0.002	0.000	-0.001	0.005	0.000
Single Farm Payment (1000 Euro)	0.002***	0.001	0.000	0.002	0.002	0.000
Economic environment						
Ag. to total employment (%)	0.001	0.006	0.000	-0.005	0.071	-0.000
VA per inhabitant (1000 Euro)	0.000	0.005	0.000	0.085**	0.039	0.003
2004-2007 averages[1]						
Pension income (0,1)	-			1.205***	0.197	0.047
TF labor int. all year round (0,1)	-			-0.607**	0.273	0.014
Constant	-1.660***	0.290		-2.839***	0.776	
Rho	-			0.573***	0.053	
log likelihood	-3030.73			-1509.54		
Chi-squared (df)	438.25*** (21)	3480.62*** (41)				

[1] Only significant four-year averages variables are reported.
*, **, *** = significant at the 10%, 5%, 1%, respectively.

relation coefficient, which measures the effect of unobserved heterogeneity, is 0.573 and is also highly significant. Hence, the results suggest that both true state dependence and heterogeneity are important in determining persistence. Heterogeneity represents individual unobservable factors, but its effect is symmetrical, as is the case with the other covariates, and the issue of heterogeneity is basically an econometric issue, implying biased estimation if not properly controlled. By contrast, true state dependence creates an asymmetry in farmers' behaviour and is an economic issue that stems from true changes in the farm setting. The existence of true state dependence is not simply an econometric issue; rather, it is a behavioural issue. Furthermore, the means of the explanatory variables, i.e., the parameters involved in the correction of the potential bias due to the initial conditions, are also significant. This finding suggests that coping with the initial conditions issue is important to avoid biased estimates[16].

To measure the effect of the past off-farm labour state on present participation, in addition to the marginal effect (ME) shown in Table 2, we calculated 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, which is evaluated at the mean values of the independent variables (or at their median, in the 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 results of the MEs and APEs are 0.048 and 0.058. That is, the probability that the present state is participation is approximately 5 percent higher if the past state was participation. This finding provides a measure of the part of persistence that is due to heterogeneity. In the dynamic model (not shown here), not accounting for heterogeneity or for the initial state, the percentages are much higher (0.175 for APE, 0.686 for ME). The reason is that, in the dynamic model, the past state variable also absorbs the effect of the unobservable idiosyncratic characteristics[17]. Hence, the comparison with the model accounting for heterogeneity suggests that an appreciable part of persistence is due to heterogeneity.

Although our results provide strong evidence that there is indeed persistence in off-farm labour participation and that both heterogeneity and state dependence determine it, it is difficult to appreciate how much of the persistence is due to either. However, the above-mentioned considerations would suggest that heterogeneity could be of greater importance than true state dependence. This notion contrasts with the results of Biørn and Bjørnsen (2015), who actually find very low levels of covariance among the state-specific random effects that account for heterogeneity. This can be due to the fact that they do not control for the initial conditions problem, which may induce an overestimate of true state dependence (Heckmann, 1981b). From their study, it is nevertheless unclear

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ρ is the constant cross-period error correlation (Arulampalam, 2009). After rescaling, the past state parameters for the Heckman, Orme, and Wooldridge models are of remarkably similar magnitude (0.874, 0.763 and 0.806, respectively).

[16] The past state parameter estimated in the dynamic random effects model, including the past state and the initial state but not controlling for the initial state bias, is 0.704 compared to 0.773 in Wooldridge's model (rescaled values), thus suggesting that not controlling that bias would cause the effect of the past state to be underestimated.

[17] Vice versa, excluding the past state variable would cause the effect of the other variables to be overestimated; see Ahituv and Kimhi (2002).

what the relative weight of true state dependence is, although they also find evidence of persistence.

In summary, from the econometric perspective, the main conclusions are that cross-sectional pooled estimation is rejected vs. models that account for heterogeneity and state dependency, and that correcting for the initial state is also crucial.

4.2 Effects of the variables

From the factual perspective, after introducing state dependence and heterogeneity and controlling for the initial state conditions, none of the idiosyncratic farmer observable characteristics, usually found to be significant in cross-sectional models, is found to have statistically significant effects on the probability that the operator works off the farm (see section 4.3 for a discussion of these differences in results).

The two variables related to household income (pensions and capital income) are strongly significant. If the household has any capital income, the probability of off-farm labour participation is significantly increased, albeit in a limited measure. At the mean values of the variables, the probability is increased by 2.8 percent. This finding is not consistent with our expectations and with the results of previous studies (Mishra and Goodwin, 1997; Mishra and Goodwin, 1998; Mishra and Holthausen, 2002), according to which farm household wealth acts as a substitute for off-farm work. This effect is interpreted as a lower need for extra work income because wealth helps these farms to smooth consumption when income falls. A possible interpretation of the positive result is that off-farm employment may be induced by high off-farm wages that, in turn, provide more financial assets that yield capital incomes[18]. If a household member has a retirement or invalidity pension income, then the probability that the farm operator has an off-farm job, as expected, significantly decreases. In this case, the marginal effect is -5.4 percent. A possible interpretation is that, for poor farm households, off-farm income that stems from pensions of other members decreases the income needs, thus raising the off-farm reservation wage of the operator. Alternatively, the old age of the operator and of the other members of the family make them less interested in off farm work. Finally, the presence of elders and individuals with a disability in the household increases the time that is devoted to caring activities, thus discouraging off-farm work (Salvioni *et al.*, 2008).

A farm's being located in mountain areas has a significant and, as expected, positive effect. Farms in those areas are typically less profitable, which might induce more pluriactivity. The Value Added per inhabitant is significant and, as expected, positive. This result confirms our expectations about larger probabilities of off-farm work in more economically developed areas i.e., where more job opportunities are available and the alternative income provided by off farm jobs is higher. All-year-round labour-intensive types of

[18] Remark nevertheless that if this were the case, a problem of endogeneity would result (we are grateful to a referee for drawing our attention to this point). Though, applying to non-linear models, such as the one we estimate, the 2SLS method usually employed in linear models leads to "forbidden regression" (Angrist and Pischke, 2009). Forbidden regressions produce consistent estimates only under very restrictive assumptions that rarely hold in practice (see for instance Wooldridge, (2010), pp. 265-268). The risk of producing inconsistent estimates further increases when the endogeneity problem is referred to a binary explanatory variable such as capital income in our model.

farming, the share of land in property and the amount of working capital are not significant in Wooldridge's model[19].

Neither the variable of Coupled Payments (before the CAP reform) nor the variable of the Single Farm Payment (after the reform) exhibits significant effects. Again, this finding makes a difference in respect to the pooled model, for which the SFP variable is strongly significant. Regardless, 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 positive sign of the SFP 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). The non-significance of decoupled payments is not consistent with the results of many previous parametric US studies (Dewbre and Mishra, 2002; El-Hosta *et al.*, 2004; Goodwin and Mishra, 2004; Ahearn *et al.*, 2006; Serra *et al.*, 2005), which have found – indeed, very small[20] – effects of decoupled payments on farm and off-farm labour and on total work hours, but they support more recent results (Pandit *et al.*, 2013), which find that neither direct nor indirect government payments have any impact on the off-farm labour supply of farm operators. Also Olper *et al.* (2014) find for the EU that the evidence of a negative effect of decoupled payments on agricultural out-migration is weak in a dynamic setting, and in no model by Petrick and Zier (2012) do direct payments have any effect on agricultural employment. The CAP reform explicitly attempted to substitute decoupled for coupled payments without substantially changing the 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.

4.3 Discussion

In this paragraph, we discuss the meaning and the implication of persistence in terms of farmers' behaviour and of policy implications, by comparing cross-sectional and dynamic panel analyses. Most likely the most striking result – also robust across all estimated models – is that, after introducing state dependence and heterogeneity and controlling for the initial state conditions, in none of these dynamic models do the idiosyncratic farmer observable characteristics have statistically significant effects on the probability that the operator works off the farm. Additionally, many of the farm characteristics found to be significant in cross-sectional models are no more significant in this dynamic setting.

Although more research is needed to generalize our results, our findings are at odds with an entire stream of literature based on cross-sectional samples (including some of our previous studies), where personal, farm and labour market characteristics are typically found to be important determinants. We conclude that past experience has a genuine behavioural effect in the sense that otherwise identical individuals who did not experience participation would behave differently from those who did experience it. In other

[19] Their coefficients are nevertheless significant in both the Heckman and Orme specifications.

[20] For instance, the coefficients estimated by Dewbre and Mishra (2002), although statistically significant, indicate that a thousand-dollar payment decreases the total work by approximately or less than one hour per year. Regarding off-farm work, the largest estimate is that by Ahearn et al. (2006), who estimate that a thousand-dollar payment decreases the probability of off-farm labour participation by 0.4-0.5 percent.

words, we find structural state dependence. We also find that unobservable idiosyncratic characteristics do have an influence on off-farm labour choice and most likely have a stronger effect, but we find scarce evidence of an effect of observable farm and personal characteristics. Importantly, this difference emerges when estimating different models from the same data and comparing the dynamic model accounting for heterogeneity and state dependence vs. the pooled model equivalent to a cross-sectional setting. Thus, the problem is how to explain the relevance of the personal and farm characteristics found in cross-sectional studies. This problem calls into question the interpretation given to the results of cross-sectional studies, that are typically interpreted in causal terms, implying, e.g., that an older age or a smaller farm size makes farmers more likely to take on an off-farm job. Tentatively, we conjecture that the relationship between personal and farm characteristics, such as age and farm size, found in models in which past experience is not taken into account could be interpreted in the following manner: i) some unobservable characteristics, of both the farm and the farmer, exert a permanent push in the direction of a particular state. For example, low soil fertility (hence, unobservable lower returns from the farm) or a farmer's preference for non-agricultural work, may create a tendency to look for an off-farm job. ii) Some observable characteristics of the farm and the farmer (such as small farm size or education) also create a tendency to take on an off-farm job. iii) These two groups of variables influence the choice whether to take an off-farm job the first time it is made, which is nevertheless contingent on the availability of sufficiently remunerative job opportunities and on a sufficiently low cost of changing status. iv) The influence of the age variable found in cross-sectional studies is therefore actually because an older age means that, for those who look for an off-farm job, more years have passed, which implies that more opportunities of finding such a job may have occurred. v) Once the choice is made, the off-farm work status tends to remain the same because some farm setting is changed after this choice, because the unobservable and observable characteristics remain at work, and because of the costs implied by changing status again. Hence, once a status is chosen, the subsequent situation is dominated by the previous situation; thus, the influence of the observable variables apparently vanishes. In this sense, one might think of an "off-farm job trap" and of a "full-time on-farm job trap". This interpretation of the discrepancy between the cross-sectional and the dynamic models in terms of "status traps" should nevertheless be made with some caution, given that our panel is not very long, even if it has been proven that 4 years are enough to produce valid and reliable results (Arulampalam and Stewart, 2009). Anyway, we feel that our results deserve attention on a possible misinterpretation of the results of cross-section models. The effect of personal and farm characteristics are at risk of being severely exaggerated if the econometrician fails to recognise the presence of persistence.

Hence, our results and their interpretation also have policy implications. If pluriactivity is considered by policy makers to be desirable, e.g., because it helps maintain the agricultural activities in a rural area and thus prevents depopulation, then appropriate policies have to be chosen to induce farmers to engage in off-farm activities in the area and to choose not to quit farming. Our results suggest that the unobservable specific characteristics of farms and farmers are prevailing in determining persistence. Hence, the levers on which to operate are those that permanently change the individual propensity to work off the farm, e.g., focusing on education and development of personal skills. Our results show that also true state dependence determines persistence. Hence, along with policies affecting individual characteristics, also policies facilitating farmers' access to the off-farm labour market could be appropriate. Given true state dependence, an initial effort is necessary to induce the off-farm labour participation of farmers, but after this initial treatment, incentives are less necessary to sustain the desired behaviour over time. Finally, our results suggest that both decoupled and coupled payments are not significant variables in explaining the off-farm labour participation of Italian farmers. This is consistent with recent US studies (Pandit *et al.*, 2013), while the related literature on the effects of farm subsidies on labour outflow from agriculture gives mixed results. For instance, a negative effect of Pillar I subsidies on net labour outflow from agriculture had been found by Olper *et al.* (2014) using panel data at the aggregate level, and similar effects are found for subsidies by Breustedt and Glauben (2007), though other papers do not find significant effects (Glauben *et a*l., 2006; Pietrick and Zier, 2011 and 2012). Remark nevertheless that these results apply to quitting agriculture, not to combining on- and off-farm labour. The implication of our results is that policymakers should not increase government spending in the form of agricultural subsidies to reduce unemployment.

5. Summary and conclusions

In this paper, we examine the issue of the off-farm labour participation of farm operators. We estimate different models that allow both heterogeneity and true state dependence, and that cope with the initial conditions problem, using a panel sample of Italian farms.

We find a strong persistence in the state, which is mainly explained by unobserved heterogeneity. The past labour state is also an important determinant of the off-farm labour participation choices of Italian farmers. These results imply that policies aimed at fostering off farm employment should focus on interventions that permanently change the individual propensity to work off the farm, e.g., focusing on education and development of personal skills.

Variables typically found to be significant determinants of these choices, and also appearing as such in pooled sample estimations, are no more significant when true state dependence and heterogeneity are controlled for. We also find that the effects of the change in CAP, and more generally of coupled and decoupled payments, are weak and generally insignificant.

To some extent, our results are at odds with the usual interpretation given to the results of cross-sectional studies concerning the effect of farm and operator characteristics. In general, our findings point to the conclusion that persistence effects are at work, because the choice of a work state modifies the preferences or the constraints that the farm operator faces (e.g., because changing status may imply sunk costs), or because an operator's unobservable preference toward off-farm work can be met contingent on local job opportunities, and the cumulative probability of finding such a job increases along with time.

The discrepancy between our results and those of the traditional cross-sectional studies calls for more research to clarify these issues. Promising research lines could include modelling the changes of labour state rather than the states *per se*, in the line undertaken by Biørn and Bjørnsen (2015), but coping with the issue of the initial values, or using job search models and modelling the permanence in the states. This task is left for further research.

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Appendix

We present here the different models that we estimated for dealing with the problems of true state dependence heterogeneity, and the initial conditions issue. In the following y_{it} is a dummy variable equal to 1 if the farm operator *i* participates in off-farm work in year *t* (*t*=1,…,T), else 0; 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 benchmark is the pooled model, equivalent to a cross-sectional model, since it assumes that past states have no effect and that all time-invariant variables are included:

$Prob(y_{it}=1) = \Phi(\beta_0 x_{it} + \varepsilon_{it})$ with $Corr(\varepsilon_{it}, \varepsilon_{is}) = 0 \ \forall \ s \neq t$ (Pooled model)

A second model accounts for state dependence, but does not deal with heterogeneity nor with the issue of the initial conditions.

$Prob(y_{it}=1) = \Phi(\beta_0 x_{it} + \gamma y_{it-1} + \varepsilon_{it})$ (Dynamic pooled model)

The third one adds to the past state the state in the initial year (y_1) among the explanatory variables and allows for individual heterogeneity (u_i), modelled as a random effect.

$Prob(y_{it}=1) = \Phi(\beta_0 x_{it} + \gamma y_{it-1} + \iota y_1 + u_i + \varepsilon_{it})$ (Dynamic random effects model)

The *Heckman* approach to the initial conditions problem involves assuming the following specification of the initial value of the dependent variable:

$y_{i1} = z_{i1}\pi + \eta_i$

where z_{i1} is a vector of exogenous variables including x_{i1} and other instrumental variables (for example pre-sample variables) and η_i is assumed to be correlated with u_i, but uncorrelated with ε_{it} for $t \geq 2$. Then η_i can be written as $\eta_i = \theta u_i + \varepsilon_{i1}$ ($\theta > 0$), with u_i and ε_{i1} independent of one another. It is also assumed that ε_{it} and u_i are normally distributed with variance 1 and σ_u, respectively. The function for the initial time period is therefore specified as

$y_{i1} = z_{i1}\pi + \theta u_i + \varepsilon_{i1}$

and this equation is estimated jointly with the one of the remaining time periods. The output is therefore an estimate of the initial year equation, and of the following years equation, including the correlation term.

Since the initial conditions problem is due to the correlation between the y_{it-1} regressor and the individual heterogeneity term, *Orme*'s approach is to insert a correction term in the equation. He substitutes u_i with another unobservable term uncorrelated with the initial observation, under the assumption that u_i and ε_{it} follow a bivariate normal distribution. This term:

$h_i = E(u_i|y_{i1}) = (2y_{i2}-1)\sigma_\upsilon\phi(\zeta' z_i /\sigma_u)/\Phi[(2y_{i2}-1)\ \zeta' z_i /\sigma_u]$

is the generalized error term from a probit estimated for the initial year, where z are the covariates, ζ the relevant parameters, ϕ and Φ the normal density and distribution functions. The generalized error terms are inserted as regressors in a random effect probit model for the remaining years, so that the overall model is:

$Prob(y_{it}=1) = \Phi(\alpha_0 + \beta_0 x_{it} + \gamma y_{it-1} + \tau h_i + \nu_i + \varepsilon_{it})$ (Orme's model)

The following table presents the results of all the above models, also including for comparison Wooldridge's model.

Table A1. Estimates of the models of off-farm labor participation.

	[1] Pooled		[2] Dynamic pooled		[3] Dynamic RE		[4] Heckman		[5] Orme		[6] Wooldridge	
	Coeff.	Std.Err.	Coeff.	Std.Err.	Coeff.	Std.Err.	Coeff.	Std.Err.	Coeff.	Std.Err.	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(1)	-		-		2.953***	0.302	-		-	0.142	2.724***	0.293
Personal characteristics												
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
Farm characteristics												
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 labor 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 labor 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 A1. (continued).

	[1] Pooled		[2] Dynamic pooled		[3] Dynamic RE		[4] Heckman		[5] Orme		[6] Wooldridge	
	Coeff.	Std.Err.	Coeff.	Std.Err.	Coeff.	Std.Err.	Coeff.	Std.Err.	Coeff.	Std.Err.	Coeff.	Std.Err.
Household characteristics												
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
CAP reform												
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
Economic environment												
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
2004-2007 averages[1]												
Pension income (0/1)	-		-		-		-		-		1.205***	0.197
TF labor 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		2269.18		-1545.95		-1509.54	
Chi-squared (df)	438.25*** (21)	3218.763*** (22)	3334.87*** (24)	400.17*** (22)	1664.38*** (25)	3480.62*** (41)						

Note: Significance levels are denoted by one asterisk (*) at the 10% level, two asterisks (**) at the 5% level, three asterisks (***) at the 1% level.
[1] Only significant four-year averages variables are reported.