The capitalization of Area payments into Land Rental Prices: a panel sample selection approach

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Abstract

Previous empirical literature suggests that agricultural subsidies are capitalized into farmland rents and that the introduction of the 2003 decoupling reform of the EU Common Agricultural Policy, attaching the subsidy to land only, might have even extended the phenomenon of capitalization. Employing the FADN dataset for Italy we investigate this issue using methodologies accounting for selectivity, endogeneity and individual heterogeneity simultaneously. The evidence suggests that selectivity bias causes inconsistent estimation of parameters and wrong inference. Results reveal instead that, in Italy, there is no incidence of both coupled and decoupled payments.

1. Introduction

Farmland is by far the most important input in agricultural production. In EU27 land, alongside permanent crops and quotas, accounts for about 65% of total fixed assets of farms in 2009 and the figure rises to 75% when only farms specialized in field-cropping are considered (European Commission - EU FADN 2013). Much attention therefore is paid in the theoretical and empirical literature to the determinants of agricultural land's prices.

Alongside the primary effect of income support, secondary effects can also by produced by agricultural subsidies influencing farmers' decision on amount and composition of factors (i.e., labour, capital and land) to be used in production. Such effects are already documented in literature by, among others, Lence and Mishra (2003) for what concerns land markets, Rude (2008) in relation to farmers' production decisions and Sckokai and Moro (2009) in relation to farmers' investment decisions.

Following the implementation of the 2003 Common Agricultural Policy (CAP) reform, agricultural subsidies have been decoupled from productions and related to land, coining the possibility that payments designed to support farmers' income get capitalized into farmland prices and transferred out of the agricultural sector (Ciaian and Swinnen 2006). Econometric studies of the capitalization effect focus on the US to the largest extent and only a minority approached the topic in the EU (Patton et al. 2008; Breustedt and Habermann 2011; Ciaian and Kancs 2012; Michalek, Ciaian and Kancs 2014), finding rather mixed evidence. For instance in Northern Ireland, Patton et al. (2008) estimate that in the period 1994-2002 as much as 40% of the amount granted to farmers was capitalized into farmland rents; in Germany, the incidence has been estimated by Breustedt and Habermann (2011) of 38% in 2005; Ciaian and Kancs (2012) in their study report estimates of an incidence rate in the range 18-20% in New Member States (NMS); finally an incidence rate between 6% and 10% is found in Michalek et al. (2014) based on the analysis of Old Member States (OMS) in the period 2004-2007. Such diverse results can be associated to differences across studies in terms of country or region analysed, time span of the data, methodological approach and type of agricultural support considered.

In addition to the variation in results related to the heterogeneity in the estimation samples, existing studies differ with respect to the methodological approach applied to estimate the incidence of agricultural payments. Issues like individual heterogeneity, selectivity and endogeneity, in particular, require careful consideration. Concerning heterogeneity, individual characters of the farmer, for instance, are generally not recorded in survey data but likely impact both the decision on the amount of land to be rented and the farmer's willingness to pay for rented land. Concerning selectivity, only a share of observed farms employs rented land in production, implying non-renting farms to be excluded from the estimation of a model in which rent is the dependent variable. Concerning endogeneity, farmers may take their decision on the amount to be paid for rented land based on the expected productivities of land and on the expected level of support; because actual productivities and level of support are instead observed and hence used in the estimation, any deviation of the observed value from the expectation will contribute to the error term. Whereas in some studies the capitalisation effect has been estimated using panel data (in EU: Patton et al. 2008) to account for individual heterogeneity, in some others the issue of selection has been considered (Ciaian and Kancs 2012). In Ciaian and Kancs (2012) cross-section selectivity methods are applied to time-difference data and individual heterogeneity in the decision to rent land (selection equation) is only weakly considered.

This research contributes to the existing empirical literature on the capitalisation effect in Europe providing econometric evidence from different panel-data sample-selection approaches. Econometric estimators have been developed to simultaneously account for selectivity and individual heterogeneity in unbalanced panels (Wooldridge 1995; Kyriazidou 1997; Rochina-Barrachina 1999) but their use in applied econometrics, in facts, has been rather limited (Dustmann and Rochina-Barrachina 2007). More recently, the Wooldridge (1995) estimator has been adapted by Semykina and Wooldridge (2010) to consider endogeneity in some covariates. Furthermore, evidence is provided for Italy, where the incidence of agricultural payment on farmland rents has been investigated to a minor extent only. Finally, the empirical analysis is conducted for both the periods before (1994-2004) and after (2005-2008) the implementation of the decoupling reform in Italy.

The remainder of the work is organized as follows. The next section discusses the issue of subsidy capitalization from both a theoretical and an empirical viewpoint. The third section introduces the econometric strategy adopted to estimate the parameters of interest. Data used in the empirical model are described in the fourth section. Econometric results are presented in the fifth section and a discussion of evidence concludes the work.

2. Farmland rents and agricultural payments

As the agricultural support scheme switched to decoupled payments introduced by the 2003 reform of the CAP, farmers receive support based on the number of entitlements, not on the quantity of output. This support is granted to farmers committed to keep farmland in good environmental and agricultural condition. Implementation of the SPS required Member States (MSs) to adopt either the regional model, consisting in a flat per ha payments for all farms in a region, or the historical model, granting farms entitlements based on the values of payments received during a *reference period*. The choice implied some important consequences on the distribution of payments across the farmers in the MS. It Italy, the historical model has been adopted and the SPS was first implemented in 2005.

Much empirical evidence which is primarily based on US data suggests that coupled subsidies are largely capitalized in land values (Barnard et al. 1997; Ciaian and Kancs 2012; Goodwin, Mishra and Ortalo-Magné 2003; Kirwan 2009; Lence and Mishra 2003; Patton et al. 2008; Roberts, Kirwan and Hopkins 2003; Weersink et al. 1999). Few is known instead on the effect of decoupled subsidies. In particular under the SPS, a single ha of eligible land is required to activate an entitlement. Thus the amount of the support is related to land intrinsically, resulting in higher likelihoods of capitalization.

Theoretical models in Ciaian and Swinnen (2006), Ciaian, Kancs and Swinnen (2008) and Kilian and Salhofer (2008) suggest that the extent to which these payments can be capitalized is rather limited in fact and much depends on the proportion of entitlements to eligible area and, more in general, on the implementation model. A difference between the number of entitlements and eligible ha for farmers is the natural consequence of the decoupling of agricultural support and the introduction of entitlements as the basis for payments. A farmer owning more entitlements than eligible ha has an incentive to either sell the additional entitlements or buy additional eligible land, the price of which will likely capitalize the agricultural support. As a more substantial redistribution of entitlements is associated to the implementation of the regional method, the difference between entitlements and eligible ha is expected larger in this case than in case the historical model is implemented. Because almost all MS choose to implement either the historical model or a hybrid model, one can reasonably expect to find little evidence of capitalization.

Existing empirical studies support the hypothesis of a limited incidence of area payments (Patton et al. 2008; Breustedt and Habermann 2011; Ciaian and Kancs 2012; Michalek et al. 2014), the estimated incidence rate stemming from 6% to 40%. In particular it appears that the incidence estimated by Ciaian and Kancs (2012) in NMS, where a model similar to the regionalized payment has been implemented, is higher than that estimated by Michalek et al. (2014) in OMS, where the historical model, and the hybrid model in some cases, was principally adopted.

Estimation of the incidence rate on farm level data requires that a number of issues are taken into account to retrieve unbiased and consistent estimate of the model parameters. Primarily in view of data limitations, these issues have been considered by previous empirical literature to a limited extent only. In theoretical models for farmland prices (Ciaian and Swinnen 2006) the input demand equation for land is obtained from the FOC of a profit maximization problem in which the assumption that all the land is property of landowners and rented to farmers holds. Let p_k and y_k denote respectively the price and the quantity per ha of output k. Then the equation (1) represents the equilibrium relationship between the costs of a rented ha of land (r) and the average product value across the k productions weighted by the share of land (A) used for the production of the k^{th} output. i and t identify the farm and the year in the sample and e is the standard *iid* vector of disturbances.

$$r_{it} = a_i + b_1 \sum_{k=1}^{K} E(p_k)_k \, y_{k,it} \frac{A_k}{A} + e_{it} \tag{1}$$

Assuming that a per ha amount g_k different for each production is granted to each farmer, the profit function to be maximized is modified accordingly and the equilibrium conditions yield the equation (2).

$$r_{it} = a_i + b_1 \sum_{k=1}^{K} E(p_k) y_{k,it} \frac{A_k}{A} + b_2 \sum_{k=1}^{K} E(g_k) \frac{A_k}{A} + e_{it} \tag{2}$$

Consistent and efficient estimation of the parameters in the equation (2) is subject to many specification issues. Firstly individual heterogeneity (a_i) characterizes the right hand side. Many variables influencing the land rental price are usually not reported in survey data. This is the case of soil quality indicators and more in general of indicators expressing features of the rented land, including goods and rights attached to it. Also the typology of contract is usually not reported although it would be important to know whether the amount paid for rents reported in the survey refers to a rent bargained at the beginning of the season and paid in cash or to the corresponding value in output of a sharecropping agreement received by the landlord at the end of the season. As a result a large part of the variation in rental price across farms may depend on factors which are in fact unobservable to the econometrician. Secondly, some farms may not use rented land in production, causing the dependent variable (r_{it}) to be truncated at zero. If the decision to rent is somehow correlated to the variables in the model right hand side, the issue of selectivity should be taken into account by estimating a probit model of the type $P(s=1) = \Phi(c_i, X, d)$, where s_{it} is the binary indicator taking non-zero value is the farm rents land, X is the matrix of explanatory variables and d is the vector of related coefficients to be estimated. The renting choice is also dependent on factors which are not observable to the econometrician as, for instance, individual characters of the farmer influencing the propensity to rent land, and this requires that fixed effects are also considered in the selection process. Finally, output values and subsidies are subject to endogeneity. In equation (2) the variables p_k are stochastic because the farmer does not know at the beginning of the period the real price at which the output will be sold. Likewise the amounts g_k granted to farmers in Europe are not know certainly as they have been subject to change, especially in years before the 2003 reform, partly because of the progressive reduction of coupled subsidies and partly because of the harmonization of payments for major crops (EC 2000). Farmers take their optimal rental choice based on the related expectations, but expectations are unobservable and the use of current realizations, which are instead observed, causes inconsistent econometric estimations and instrumental variable estimation is required. Allowing Y and G to represent the absolute values of production and payments (not per ha) respectively, the expectation error can be written as $\frac{1}{A}\left(\sum_{k=1}^{K}\left(E(p_k) - p_k\right)Y_{k,it} + \sum_{k=1}^{K}\left(E(G_k) - G_k\right)\right)$ and it is clearly correlated with the realized values of per ha productions and subsidies. Many authors have used past realizations of productions and subsidies as instruments arguing that these are in the information set at the time of decision making and are not correlated with the differential between the expectation and the realization (Lence and Mishra 2003; Kirwan 2009; Patton et al. 2008; Roberts et al. 2003). This implies a considerable loss of observations in the case of strongly unbalanced panels. Alternatively the area shares can be used. Area shares are strongly related with both the agricultural productivity and the amount of payment but are naturally uncorrelated with the expectation error which depends uniquely on the total quantity of land used for each production and not on the relative share.

Decoupled subsidies have been introduced with the 2003 reform, replacing the old payment scheme. In the historical model the amount granted to each farmer is fixed and computed based on payments received by each farmer in the reference period. Under the decoupled scheme the relevance of expectation errors is lower as the farmer knows more certainly the amount that will be granted and hence the use of current realization of payments stops causing endogeneity of the econometric estimates.

Previous studies on the issue of capitalization in Europe have considered these estimation issues separately. The econometric framework proposed in this work attempts to consider them simultaneously. In Wooldridge (1995) (hereinafter W95) a correction procedure is proposed to estimate panel data models in presence of endogenous selectivity and individual heterogeneity in both the selection and the main equations. The methodology relies on a two-step procedure similar to the one proposed by Heckman (1979). The W95 procedure requires T different probit models to be estimated, one for each year. Based on the probit estimates T different Inverse Mill's Ratios (IMRs) are computed and included in the main equation after pooling. Robust standard errors can be estimated using formulas in the appendix of Wooldridge (1995). Alternative estimators have been proposed by Kyriazidou (1997) and Rochina-Barrachina (1999). Both estimators rely on first-differencing of data to eliminate individual heterogeneity and use only observations for which the selection indicator is constant for two consecutive time periods. This generally implies a loss in efficiency with respect to W95, especially in the case of strongly unbalanced panel data. The comparison of the three estimators is discussed in greater detail in Dustmann and Rochina-Barrachina (2007). Semykina and Wooldridge (2010) (hereinafter SW10) extend the W95 approach allowing to correct for endogeneity bias due to non-zero correlation between explanatory variables and idiosyncratic errors. The procedure again works in two steps. In the first step T probit equations are estimated to derive IMRs. In the second step an IV estimator is applied instrumenting endogenous variables with a subset of variables used in first step.

3. Econometric Model

Let r_{it} be the rental price of land for farm i at time t, s_{it} the binary variable indicating whether the farms rents land , Z_{it} a matrix of covariates and X_{it} a subset of Z_{it} which excludes some variables to be used in the second step only (exclusion restrictions), the model in equation can be estimated using the two step procedure in W95

$$\begin{cases} r_{it} = \alpha_i + \beta X_{it} + \varepsilon_{it} \text{ for } s_{it} > 0 \\ P(s_{it} = 1) = \gamma_i + \delta Z_{it} + \mu_{it} \end{cases}. \tag{3}$$

Step 1 – T different probit models $P(s_i = 1) = \gamma + \delta_t Z_i + \mu_i$ are estimated to retrieve estimates of δ and, based on this, the vectors of IMRs. Alternatively a Pooled model $P(s_{it} = 1) = \gamma_i + \delta_1 Z_{it} + \delta_2 Z_i + \mu_{it}$ can be estimated, Z_i being the matrix of individual means of variables. This is a generalization of the Mundlak (1978) and Chamberlain (1982) approach to account for the correlation between the covariates and the individual effects extended to non-linear panel data models (Wooldridge 2010, Ch15).

Step 2 – The estimated IMRs $\left(\hat{\lambda}_{it}\right)$ are pulled into the main equation $r_{it}^{WT} = \alpha + \beta X_{it}^{WT} + \rho \hat{\lambda}_{it}^{WT} + \varepsilon_{it}$ in which the notation ^{WT} applied to a variable indicates that the within transformation has been applied to it. Following W95 an estimate of the parameter ρ different from zero indicates selectivity bias and suggests that a correction is required.

Step 3 – Denoting with Y^j the subset of variable considered endogenous in the X matrix, $j = 1, 2, ..., J$ regressions are run $Y_{it}^j = \eta_i^j + \varphi Z_{it}^j + \omega_{it}^j$ where the Z matrix now includes also appropriate instruments for the endogenous regressors. Estimated vectors $\hat{\omega}_{it}^j$ are then plugged into the equation $r_{it}^{WT} = \alpha + \beta X_{it}^{WT} + \rho \hat{\lambda}_{it}^{WT} + \sum_j \theta_j \hat{\omega}_{it}^{j,WT} + \varepsilon_{it}$. This the equivalent of the Durbin-Wu-Hausman augmented regression test described in Davidson and MacKinnon (1993) but the within transformation is applied to variables to account for fixed effects and the IMR, now estimated excluding endogenous variables from the probit model, is included to control for selectivity. Statistically different from zero estimates of the parameters θ suggest that endogeneity should be also considered in addition to selectivity.

Step 4 – Estimate the model in equation (4) where $X^* \cup Y = X$ and $X^* \cup W = Z^*$, W being a matrix of exclusion restrictions in the probit model and of suitable instruments for the endogenous regressors. In practice, T probits are first estimated using exogenous variables, instruments and exclusion restrictions as explanatory variables and the IMRs are retrieved and stacked by year to form the $\hat{\lambda}_{it}$ vector. Then, for the sample of farms having a positive rent, the rental price is regressed against the endogenous covariates, the exogenous covariates and their individual mean, the IMR and the interaction of this with $T-1$ time dummy variables (T_t) by two-stage lest square using exogenous covariates, instruments, exclusions and all the respective individual means in addition to the IMR and the interaction of this with $T-1$ time dummy variables as explanatory variables in the auxiliary regressions.

$$\begin{cases} r_{it} = \alpha + \sum_j \beta_1^j Y_{it}^j + \beta_2 X_{it}^* + \beta_3 X_i^* + \rho_1 \hat{\lambda}_{it} + \sum_{t=2}^{T} \rho_t T_t \hat{\lambda}_{it} + \varepsilon_{it} \\ Y_{it}^j = \eta^j + \varphi^j Z_{it}^* + Z_i^* + \rho_1 \hat{\lambda}_{it} + \sum_{t=2}^{T} \rho_t T_t \hat{\lambda}_{it} + \omega_{it}^j \\ P(s_i = 1) = \gamma + \delta_t Z_i + \mu_i \end{cases} \tag{4}$$

For the purpose of testing selectivity in the case of endogenous variables it is sufficient to look at the significance of the estimated ρ_1, ether separately or jointly with the estimates of ρ_t, in a simplified fixed effects instrumental variable regression of rents on endogenous and exogenous covariates, the estimated IMR and the interaction of this with time dummy variables. On the opposite, for the purpose of inference on the β parameters, the corrected standard errors for the parameters in the main regression of the model in equation (4) should be computed based on the formula of the variance-covariance matrix provided in the appendix of SW10.

4. Data

The database used for the empirical estimation is the Italian FADN (Farm Accounting Data Network) and the focus is on farms specialized in field-cropping. Considering the years 1994-2008 there are 77913 records in the dataset. Out of these, a value of field-crop production lower or at least equal to zero appears in 938 cases, which have then been excluded from the estimation sample. Considering five main output categories (cereals, maize, oilseeds, proteins and other crops), there are in addition 2327 observations related to farms that report either non-zero production and zero land or zero production and non-zero land in a single output category and these observations are not considered for estimation. Finally it appears that some farms received coupled subsidies for a production of an output without producing that specific output (439 observations). Although it is plausible that certain farms might have received the payment with a delay and hence the amount has been accounted in a year different from that in which the right matured, this is deemed an inconsistency for the purpose of estimation and the observations are dropped from the estimation sample. Overall the incidence of cleaning on sample size is lower than 5% and the final estimation sample consists of 74209 observations, 58582 of which pertain to the period in which payments were coupled to productions and 15627 to the period in which farmers receive decoupled payments. Both panels are strongly unbalanced, the average permanence of a farm in the survey being 2.59 and 2.2 years in the periods before and after the introduction of the reform respectively.

Table 1 describes the variables used for the estimation and provides some descriptive statistics related to the two time periods. The dependent variable, the rental price of land, is measured as the total monetary value of rent paid by the farmer over the total number of rented ha. This value includes also the rent paid for buildings and other structures present on land, if any, and excludes the rent paid for all other rights and quotas which are not attached to land.

Following the theoretical model and a consolidated empirical literature the main explanatory variable of the rental price is the expected average product value per ha. Because farms specialized in field-cropping only are considered in this study, this is the sum of expected production values of cereals, maize, proteins, oilseeds and other crops divided by the total area used by the farm for these cultures. Being the total production value determined by the output prices, which are stochastic and not known by the farmer at the beginning of the year, the observed values are used. Production values are deflated using an appropriate set of deflators retrieved from the Eurostat website[1]., the year 2000 serving as base year. More in detail, a cereal-specific deflator was used for the productions of cereals and maize, a deflator specific for industrial crops was used for oilseeds and proteins and a generic crop deflator for the other crop productions.

In addition to the value of production, the level of payment is included in the rental price equation. This is, in the period 1994-2004, the total value of payments received for cereals,

[1] Database apri_pi00_outa.

maize, proteins, oilseeds and other crops deflated using crop-specific deflators described above and divided by the total number of eligible ha in these output categories. In the period 2005-2008 the variable is computed as the total deflated amount of decoupled payments received over the total amount of eligible ha. Equally to the case of production values, taking the observed value of received payments in place of the expected has important consequences for the model specification. During the period of decoupled subsidies, in particular, the amount of subsidy received by the farmer might have exhibited substantial differences with respect to farmers' expectations. On the contrary, the amount of the grant is more certain if payments are decoupled from production and based on entitlements only. In both periods, a dummy variable for farms receiving the payment is included to test if any statistical difference exists between the rent paid by granted and non-granted farmers.

Table 1- Description of model Variables

Dependent Variable		Mean (SD) 94-04	Mean (SD) 05-08
RENT	Rental rate (€/ha)	189.86 (175.94)	287.44 (274.71)
SEL	Dummy (1 if the farm rents land)	50.98	50.62
Endogenous Covariates			
PROD	Average arable crop productivity (COP, maize and other crops, €/ha)	2097.26 (7387.22)	3924.20 (15592.27)
SUB_C	Average coupled payment rate (COP, maize and other crops, €/ha)	242.21 (146.28)	
Exogenous Covariates			
SUB_D	Average decoupled payment rate (€/ha)		387.12 (415.62)
SUB	Dummy (=1) if the farm receives subsidies (%)	84.14	87.10
ENT	Dummy (=1) if the farm has more entitlements than eligible ha (%)		56.41
SIZE	Farm size (ha)	28.71 (60.13)	45.53 (94.81)
CF	Value of cash flow per ha (€/ha)	1565.20 (5451.80)	2294.12 (8950.23)
LS	Share of livestock and livestock products in total farm output (%)	14.41	10.83
BP	Share of by-product activities in total farm output (%)	1.63	1.09
Additional controls (at the regional level)			
ADENS	Average number of livestock equivalent units per ha (#/ha)	0.13 (0.12)	0.059 (0.047)
WAGE	Average regional wage per hour worked (€/h)	6.55 (2.05)	8.17 (2.04)
LANDTR	Share of farms reporting the acquisition of new land (%)	1.48	0.35
Exclusion Restriction			
LABOUR	Share of family to total worked hours (%)	93.36	85.51
CAPITAL	Assets (buildings, equipment and machinery) value per ha (€/ha)	3846.42 (7641.34)	5374.05 (13536.91)
Instruments			
CEREALS	Share of farmland planted with cereals	48.24	44.70
MAIZE	Share of farmland planted with maize	14.30	15.60
PROTEINS	Share of farmland planted with proteins	1.29	2.31
OILSEEDS	Share of farmland planted with oilseeds	6.61	4.75

Source: own elaboration on EU-FADN - DG AGRI data.

Hence the variables measuring productivity and coupled payments are considered endogenous in the empirical specification of the farmland rent model and the variable measuring decoupled payments is threated as exogenous. A dummy variable is also included in the model taking non-zero value if the farm has received payments to measure if any structural difference exists in the level of paid rent between supported and non-supported farmers which is not associated with the amount of support received. For the 2005-2008 years only, two additional dummy variables are included. One indicates whether the farm has more entitlements than eligible ha and one is obtained as the product between the latter and the dummy for support.

The additional covariates included in the model are selected following the previous empirical literature on the topic. Firstly, the size of the farm is measured as total has and is expected in a negative relationship with the average per ha rent, as larger farms have in fact relatively more power in the bargaining with landlords. Secondly, the cash flow is a proxy for the farmer's capacity of self-financing the agricultural activity and is measured as the income from agricultural activities plus the balance between subsidies and taxes. Due to the barriers a farmer may experience in the access to credit, the capacity of self-financing might be considered a determinant of the willingness to rent land and to pay for the rented land. In this respect, higher values of the cash-flow are expected to correlate positively with the net-rent. Unfortunately the cash-flow is measured in the FADN at the end of the year. This is considered a limitation because rent is accounted among the operating costs, resulting in lower values of cash-flows for farms having paid higher rents. Finally, the incidence of diversification of production on farmland rents is measured with the output share of non-crop activities, namely by-product activities and livestock production. As additional land might be rented by farmers for non-crop production, the inclusion of these variables in the models allows taking into appropriate consideration the variation in farmland rents which, being not related with agricultural productivity and payments in the sector of specialization, result from the specific production choice of the farm.

Other factors influencing the variation of farmland rents are measured at the regional (NUTS 2) level. This is the case of the animal density, as measured by the total livestock equivalent number of animals per ha; the average wage, as measured by the compensation per hour of hired labour; and the functioning markets for land on sale, as measures by the proportion of farms which have increased the number of ha of owned land. The motivation for the inclusion of the animal density in the regression is twofold. On the one hand, to comply with the nitrate directive, farms specialized in livestock production usually have to rent additional land for the purpose of manure spreading only, increasing in this way the demand for land with consequences also for the field-cropping sector. Thus a higher density of animals in the region reasonably yields to higher farmland rents. On the other hand, a higher density of animals in a region also indicates the specialization of farms in that region in the production of livestock and related products and lower rents in the region may result from the lack of adequacy of the soil for the purpose of field-crop productions. The relationship with farmland rents may be positive or negative depending on which effect dominates, accordingly. Concerning the regional wage, this is a proxy for the cost of the other variable input in agricultural production, labour. In most theoretical models of agricultural production used for the analysis of farmland prices, land is either considered the only input in agriculture (Ciaian and Swinnen 2006) or is considered alongside other variable inputs but the rental price is derived from a partial equilibrium setting (Lence and Mishra 2003) with the consequence that the land price is not influenced by the price of other inputs. In fact, wage differentials across regions may reflect the varying levels of demand and supply of hired labour which are in turn determined by territorial factors related to the local agricultural system primarily but extending to the socio-economic environment also. Assuming that these factors influence the demand and supply for all the variable inputs, it should be observed that all variable input prices correlate positively. Hence the regional wage is expected to capture the variation in farmland rents determined by territorial characteristics which have an effect on the demand and supply of hired factors. In relation to the functioning of land markets, finally, the demand for rented land responds to the scarcity of land for sale in some regions. Farmers that are willing to expand the production and are not in the position to buy additional land in the neighbourhood are constrained to the alternative of renting land with long-term contracts. The proportion of farms which have increased their owned land is hence a proxy for the general possibility to buy additional land and relatively higher rents can be expected where such a proportion is low.

Two variables are used in the probit model to explain the probability of renting agricultural land. These are the family-to total labour ratio, measured as the ratio in number of hours, and the capital intensity, measured as the monetary value of buildings, machinery and equipment per ha. Both are proxy for the managerial approach to farming and it is expected that rented land is used by managerial farm to a considerably larger extent. The composition of agricultural production, measured by the area shares, is instead used to explain the value of agricultural productivity and coupled payments. The validity of instruments is tested and augmented regression tests are conducted to test the endogeneity of variables using these instruments.

5. Results

A battery of tests is preliminary conducted on data for both the periods attempting to determine to what extent selectivity and endogeneity should both be considered in the specification of the econometric model. It turns that, in both periods, there is strong evidence of selection bias. The value of the coefficient associated to the IMR estimated from a pooled probit in a regression of land rents against endogenous and exogenous covariates, is estimated positive and statistically different from zero. Table 2 reports the estimates of the probit model and of the rental price equation.

The probit model is estimated for the unique purpose of retrieving the IMR $\left(\hat{\lambda}\right)$ which value is then used for testing against the null hypothesis of absence of selectivity and the exclusion restrictions are not considered at this stage. Consequently the probit estimates are not correct but have been reported to demonstrate the association between the covariates and the probability to rent yielding to the selectivity.

To explore the endogeneity issue a modified version of the Durbin-Wu-Hausman augmented regression test described in Davidson and MacKinnon (1993) is implemented and the results are shown in The test for selectivity robust to heterogeneity described in SW10 is finally implemented and the results are shown in

Table 4. A pooled probit is estimated for the probability to rent land, including individual means of the covariates to control for unobserved farm level heterogeneity. The estimated IMR is included in the 2SLS fixed effects regression of rental price. A test against the null of absence of selectivity is conducted based on the estimates of coefficients related to the IMR (ρ_1) and to the interaction of this with time dummy variables (ρ_t). In both the periods the hypothesis of absence of selectivity is strongly rejected, at least when all the interactions are considered jointly.

Table 3. For each endogenous variable a regression is run of that variable against exogenous covariates and instruments, including individual means to account for unobserved heterogeneity. Then a pooled probit regression is run of the selection indication against exogenous covariates and exclusion restriction. Residual from the first step $\left(\hat{e}_j\right)$ and IMR from the second step are finally pulled in the main equation. Based on coefficient estimates from the first step auxiliary regressions and the second step probit estimates there is evidence of endogeneity of considered variables.

Table 2 – Results of the W95 selectivity test

	1994-2004		2005-2008	
	Rent equation	*Probit*	*Rent equation*	*Probit*
PROD	0.268***	0.000	0.986***	0.004*
	(0.035)	(0.000)	(0.095)	(0.003)
SUB_C	10.246***	0.015**	-1.113*	
	(1.006)	(0.007)	(0.574)	
SUB_D				-0.010
				(0.008)
SUB	-34.346***	-0.030	34.734**	0.399*
	(3.943)	(0.023)	(13.934)	(0.221)
ENT			18.416**	-0.334
			(8.426)	(0.317)
*SUB*ENT*			-20.179**	0.273
			(9.679)	(0.327)
SIZE	0.552***	0.008***	-0.496***	0.003
	(0.097)	(0.001)	(0.127)	(0.003)
CF	-0.557***	-0.001	-0.410***	-0.003
	(0.056)	(0.000)	(0.136)	(0.002)
LS	-15.584**	0.096**	-7.097	0.213
	(7.036)	(0.040)	(20.012)	(0.360)
BP	-11.497	-0.174	92.084	-0.190
	(22.285)	(0.117)	(72.627)	(1.116)
WAGE	14.439***	-0.004	-4.877***	0.091***
	(0.712)	(0.006)	(1.483)	(0.021)
ADENS	-70.342***	0.121**	659.195***	-4.198***
	(8.256)	(0.054)	(60.342)	(1.389)
LANDTR	88.543***	0.286***	974.542**	-23.652***
	(17.663)	(0.109)	(377.512)	(5.565)
$\hat{\lambda}$	82.362***		-136.964***	
	(19.325)		(26.261)	
Intercept	8.394***	-0.475***	15.323***	1.224***
	(0.509)	(0.064)	(1.763)	(0.195)

Notes to Table 2: Standard Errors reported in parenthesis. ***, ** and * indicate significance at 1%, 5% and 10%. Probit equations are estimated including individual means of variables (coefficients and SE not reported) to get FE estimates. The rent equation is estimated using within transformed variables.

The test for selectivity robust to heterogeneity described in SW10 is finally implemented and the results are shown in

Table 4. A pooled probit is estimated for the probability to rent land, including individual means of the covariates to control for unobserved farm level heterogeneity. The estimated IMR is included in the 2SLS fixed effects regression of rental price. A test against the null of absence of selectivity is conducted based on the estimates of coefficients related to the IMR (ρ_1) and to the interaction of this with time dummy variables (ρ_t). In both the periods the hypothesis of absence of selectivity is strongly rejected, at least when all the interactions are considered jointly.

Table 3 – Results of the endogeneity test

	1994-2004				2005-2008		
	Rent equation	*Probit*	*IV 1-st*		*Rent equation*	*Probit*	*IV 1-st*
			PROD	*SUB_C*			*PROD*
PROD	3.777^{***}				1.830^{***}		
	(0.257)				(0.333)		
SUB_C	16.679^{***}						
	(1.733)						
SUB_D					-1.321^{**}	-0.012	0.109
					(0.577)	(0.008)	(0.240)
SUB	-37.779^{***}	0.038	-2.232^{*}	2.694^{***}	39.420^{***}	0.369^{*}	-1.269
	(5.604)	(0.051)	(1.286)	(0.016)	(13.945)	(0.224)	(4.856)
ENT					18.184^{**}	-0.402	-6.667
					(8.432)	(0.317)	(9.213)
*SUB*ENT*					-21.142^{**}	0.348	6.874
					(9.684)	(0.327)	(9.355)
SIZE	0.627^{***}	0.021^{***}	-0.039	0.001^{***}	-0.441^{***}	0.003	-0.013
	(0.088)	(0.003)	(0.036)	(0.000)	(0.128)	(0.003)	(0.048)
CF	-1.647^{***}	-0.006	0.293^{***}	0.000	-0.550^{***}	0.001	0.114^{***}
	(0.096)	(0.001)	(0.007)	(0.000)	(0.138)	(0.002)	(0.017)
LS	8.938	0.401^{*****}	-6.761^{**}	0.001	-2.693	0.216	-15.450^{**}
	(7.219)	(0.168)	(3.141)	(0.040)	(20.075)	(0.353)	(7.506)
BP	52.389^{**}	-0.568	-13.773	0.254^{**}	110.810	-0.342	-12.724
	(22.520)	(0.496)	(9.789)	(0.124)	(73.266)	(1.096)	(28.494)
WAGE	9.850^{***}	0.031^{**}	0.046	0.037^{***}	-4.459^{***}	0.094^{***}	-0.187
	(0.747)	(0.015)	(0.335)	(0.004)	(1.493)	(0.021)	(0.685)
ANIMALD	-55.915^{***}	0.479^{**}	-2.415	-0.728^{***}	665.604^{***}	-4.216^{***}	10.341
	(8.278)	(0.213)	(4.149)	(0.053)	(60.368)	(1.293)	(27.374)
LANDTR	88.446^{***}	0.715	1.707	-0.625^{***}	965.851^{**}	-23.364^{***}	-114.851
	(17.579)	(0.466)	(9.908)	(0.126)	(377.800)	(5.534)	(163.861)
Intercept	8.381^{***}	1.465^{***}	8.206^{***}	-1.366^{***}	15.268^{***}	1.551^{***}	-1.201
	(0.506)	(0.161)	(1.451)	(0.018)	(1.764)	(0.203)	(3.633)
			Correction terms				
$\hat{e}_1$	-3.560^{***}				-0.830^{**}		
	(0.258)				(0.341)		
$\hat{e}_2$	-12.391^{***}						
	(2.112)						
$\hat{\lambda}$	59.211^{***}				-86.410^{***}		
	(11.703)				(25.101)		
			Exclusion Restrictions				
LABOUR		-0.158				0.122	
		(0.239)				(0.296)	
CAPITAL		-0.002^{*}				-0.001^{*}	
		(0.001)				(0.001)	
			Instruments				
CEREALS			-21.465^{***}	-0.451^{***}			-46.442^{***}
			(2.887)	(0.037)			(5.728)
MAIZE			-19.073^{***}	0.840^{***}			-41.648^{***}
			(3.715)	(0.047)			(7.803)
PROTEINS			-20.533^{***}	-0.616^{***}			-44.010^{***}
			(5.391)	(0.068)			(10.198)
OILSEEDS			-21.905^{***}	1.610^{***}			-45.439^{***}
			(3.580)	(0.045)			(8.702)

Notes to The test for selectivity robust to heterogeneity described in SW10 is finally implemented and the results are shown in

Table 4. A pooled probit is estimated for the probability to rent land, including individual means of the covariates to control for unobserved farm level heterogeneity. The estimated

IMR is included in the 2SLS fixed effects regression of rental price. A test against the null of absence of selectivity is conducted based on the estimates of coefficients related to the IMR (ρ_1) and to the interaction of this with time dummy variables (ρ_t). In both the periods the hypothesis of absence of selectivity is strongly rejected, at least when all the interactions are considered jointly.

Table 3Standard Errors reported in parenthesis. ***, ** and * indicate significance at 1%, 5% and 10%.
Probit equations and first stage equations for the endogenous variables are estimated including individual means of variables (coefficients and SE not reported) to get FE estimates. The rent equation is estimated using within transformed variables.

Table 4 – Results of the SW10 selectivity test robust to endogeneity

	1994-2004			2005-2008		
	Rent equation	*Probit*	*Individual means*	*Rent equation*	*Probit*	*Individual means*
PROD	0.182***			1.019***		
	(0.037)			(0.110)		
SUB_C	2.028*		0.003			
	(1.196)		(0.040)			
SUB_D				-1.442**	-0.003	0.024***
				(0.685)	(0.005)	(0.006)
SUB	-18.250***	-0.001		48.284*	0.108	-0.061
	(4.548)	(0.032)		(19.682)	(0.100)	(0.114)
ENT				83.817*	0.273	-0.294***
				(43.733)	(0.187)	(0.078)
*SUB*ENT*				-81.942*	-0.286	0.656***
				(44.542)	(0.190)	(0.078)
SIZE	0.081	0.007***	-0.006***	-0.931***	0.001	-0.001
	(0.119)	(0.001)	(0.001)	(0.212)	(0.001)	(0.001)
CF	-0.352***	0.000	0.001***	-0.367**	0.000	0.000
	(0.059)	(0.000)	(0.000)	(0.161)	(0.001)	(0.001)
LS	-9.315	0.122	0.183**	-13.546	0.031	0.091
	(7.663)	(0.078)	(0.083)	(23.757)	(0.154)	(0.166)
BP	7.012	-0.151	0.203	142.730*	-0.124	1.123
	(23.930)	(0.241)	(0.301)	(84.275)	(0.591)	(0.732)
WAGE	3.922***	-0.015*	0.043***	-4.117*	0.024*	0.101***
	(0.880)	(0.008)	(0.009)	(2.385)	(0.014)	(0.016)
ANIMALD	-64.594***	0.089	1.007***	457.713***	-0.551	3.480***
	(9.157)	(0.107)	(0.121)	(92.331)	(0.558)	(0.617)
LANDTR	77.458***	0.300	6.784***	426.511	-5.696*	9.827**
	(18.697)	(0.281)	(0.383)	(595.538)	(3.382)	(4.388)
CEREALS		0.069	-0.278***		0.028	-0.064
		(0.071)	(0.077)		(0.117)	(0.127)
MAIZE		0.020	0.368***		-0.127	0.291*
		(0.092)	(0.099)		(0.159)	(0.170)
PROTEINS		0.135	-0.250		0.048	-0.153
		(0.133)	(0.169)		(0.210)	(0.259)
OILSEEDS		0.187**	0.015		-0.063	-0.380*
		(0.089)	(0.102)		(0.178)	(0.212)
LABOUR		-0.091	0.300**		0.028	-0.049
		(0.110)	(0.117)		(0.146)	(0.156)
CAPITAL		-0.001***	-0.001***		0.000	0.000
		(0.000)	(0.000)		(0.000)	(0.000)
Intercept	121.007***	-0.705***		324.366***	-1.353***	
	(20.856)	(0.050)		(120.341)	(0.084)	
			Tests for selectivity bias			
$H_0: \rho_1 = 0$	1.51			0.12		
	[0.2184]			[0.7254]		
$H_0: \rho_1 = \rho_2 = \rho_t = 0$	439.45			52.41		
	[0.0000]			[0.000]		

Notes to

Table 4: Standard Errors reported in parenthesis. ***, ** and * indicate significance at 1%, 5% and 10%. Probit equations and first stage equations for the endogenous variables are estimated including individual means of variables to get FE estimates. The rent equation is estimated by 2SLS using the within transformed variables. The test statistic for the hypothesis of no selection bias is χ^2 distributed with degrees of freedom equal to 1 in the case the only coefficient on the IMR is considered and 11 and 4 in the case the interactions with time dummy are also considered, respectively in the 1994-2004 and 20058-2008 periods.

The final results are presented in Table 5. Coefficients and standard errors estimates are presented for the SW10 method and for the standard 2SLS fixed effect to show the relevance of the applied correction for selectivity. For the estimation with the SW10 method the IMR employed in the regression is computed based on the pooled probit estimates reported in

Table 4. Individual heterogeneity at the farm level is instead accounted for through the inclusion of individual means of variables. For the purpose of inference, only the coefficients on normal variables (not individual means) should be considered, however.

Turning to the estimation results, the coefficient for land productivity is always positive and statistically larger than zero. The size of the coefficient changes across the two periods, increasing in the most recent one, and across estimators, being considerably larger when the SW10 method is employed. Evidence about the incidence of agricultural subsidies suggests that, in Italy, coupled subsidies have not been capitalized into farmland rents. Also decoupled subsidies are not directly capitalized into farmland rents but there is weak evidence of granted farmers paying higher rents, on average. On the contrary it appears that no additional premium is paid by those farmers renting land and owning more entitlements than eligible ha. When considering the incidence of agricultural payments, large differences in the slope direction and size are exhibited between the values of coefficients estimated with the two methods.

The contribution of farm size to farmland rent is ambiguous. Evidence indicates that larger farms have paid higher rents in the period 1994-2005 and the opposite is true in the period 2005-2008. The result can be in part related to the role played by large farms in the Italian agriculture, which, according to the last census of agriculture (ISTAT 2013) has changed in last decades. In particular, the dynamic of the Italian agriculture has been shaped by a dramatic reduction in the number of farms and by the increase as well of the average farm size, likely the effect of small and possibly uncompetitive farms leaving the market. This change toward a larger farm size is noticeable also in the sample used for the estimation in this work. In addition to the evidence on the average size suggested by the summary statistics in the Table 1, it is worth noting that the proportion of enterprises in the sample farming more than 50 ha has increased to the 23.83% in 2008 from the 7.75% in 1994 and that of enterprises farming more than 100 ha to 10.07 from 1.57. We speculate that this changing farm structure might have impacted the land market in some way, possibly rebalancing the bargaining power between tenants and landlords in favour of the formers and causing larger farms to pay lower rents.

The coefficient related to the cash flow is negative and weakly significant in both periods. The unexpected sign associated to the low level of significance suggest that the variable can be used to proxy the farmers access to credit to a limited extent only. As explained in the section four, it is possible that, on the contrary, the weak negative association results from a spurious correlation generated by the method used to compute the cash flow in the FADN databank.

The diversification of the agricultural activities of the farm, as proxy by the output share of non-crop activities such as by-product and livestock, has no direct effect on farmland rents as the estimated coefficients are not statistically different from zero in both periods

Table 5 – Estimation Results: IV and SW10 compared

	1994-2004			2005-2008		
	IVFE	*SW10*	*Individual means*	*IVFE*	*SW10*	*Individual means*
PROD	0.236***	2.002***		1.060***	2.199***	
	(0.038)	(0.689)		(0.110)	(0.764)	
SUB_C	8.514***	1.183				
	(1.073)	(2.963)				
SUB_D				-1.361**	-1.283	15.283***
				(0.658)	(0.904)	(2.584)
SUB	-33.217***	-10.675	-16.362	43.909***	45.246*	-129.931**
	(4.234)	(9.644)	(13.400)	(16.987)	(25.110)	(53.424)
ENT				75.072**	72.838	-33.429
				(36.889)	(77.871)	(39.377)
*SUB*ENT*				-76.139**	-76.908	-10.778
				(37.197)	(77.756)	(36.004)
SIZE	-0.147***	0.265*	-0.207	-0.863***	-0.407**	0.444**
	(0.079)	(0.143)	(0.146)	(0.192)	(0.195)	(0.196)
CF	-0.367***	-1.331*	-0.941	-0.471***	-1.574*	-0.693
	(0.060)	(0.718)	(0.858)	(0.160)	(0.820)	(0.731)
LS	-24.263***	-3.565	-92.215***	-10.472	12.590	-163.183***
	(7.567)	(10.056)	(12.311)	(23.765)	(22.756)	(34.535)
BP	12.476	26.437	-244.164***	133.213	102.624	-633.138***
	(24.072)	(26.918)	(56.106)	(84.377)	(77.421)	(155.729)
WAGE	11.094***	6.878***	11.039***	-4.686***	-0.364	22.364***
	(0.704)	(1.525)	(2.086)	(1.734)	(2.456)	(5.448)
ADENS	-67.079***	-79.054***	163.881***	733.491***	252.904***	-81.472
	(8.878)	(16.575)	(28.850)	(69.148)	(141.963)	(276.755)
LANDTR	77.012***	90.868	-198.165	1035.714**	-718.819	-6779.654***
	(18.675)	(65.998)	(136.981)	(436.670)	(546.669)	(1175.532)
CEREALS			-131.466***			-170.489***
			(22.274)			(32.863)
MAIZE			-46.042***			-60.054*
			(16.324)			(34.449)
PROTEINS			-246.089***			-280.661***
			(31.096)			(86.152)
OILSEEDS			-161.204***			-214.660***
			(22.904)			(48.124)
LABOUR			-97.413***			-106.608***
			(24.437)			(39.745)
CAPITAL			-0.090			-0.293
			(0.102)			(0.245)
Intercept	138.836***	200.773***		262.389***	405.367***	
	(6.143)	(40.326)		(24.731)	(108.115)	
$\hat{\lambda}$		38.736**			-29.792	
		(16.715)			(57.125)	
$\hat{\lambda}\cdot T_2$		-6.339			-68.016**	
		(13.751)			(32.143)	
$\hat{\lambda}\cdot T_3$		-25.058			-42.580	
		(17.981)			(45.581)	
$\hat{\lambda}\cdot T_4$		-15.509			21.364	
		(17.318)			(45.434)	
$\hat{\lambda}\cdot T_5$		-25.482				
		(22.064)				
$\hat{\lambda}\cdot T_6$		44.077				
		(41.890)				
$\hat{\lambda}\cdot T_7$		56.957**				
		(25.032)				
$\hat{\lambda}\cdot T_8$		3.488				
		(20.350)				
$\hat{\lambda}\cdot T_9$		32.132				
		(32.229)				
$\hat{\lambda}\cdot T_{10}$		41.495				
		(34.482)				
$\hat{\lambda}\cdot T_{11}$		30.983				
		(37.576)				

Notes to Table 5: Standard Errors reported in parenthesis
***, ** *and* * *indicate significance at 1%, 5% and 10%.*

Regional variables control for any source of variation related to the socio-economic and productive environment in which farms operate. The positive association between farmland rents and the cost of labour is confirmed by the empirical results, at least for the period before the introduction of decoupled payments. There is no evidence of this association in the period 2005-2008, on the opposite. The density of animals is correlated with farmland rents negatively in the period 1994-2004 and positively in the period 2005-2008. The lower land price in regions where crop production is relatively less important may result from a spurious correlation caused by the unobserved quality of soil. The effect of the nitrate directive, on the contrary, may have become important in the second period only, when the regulation of nitrates turned more stringent from farmers. According to the Commission report (EC 2010) during years 2004-2007, in fact, in almost all EU15 member states the portion of territory subject to the implementation of action programmes has grown and in countries like Italy the portion of vulnerable zones has also grown. Finally the results indicate that no effect on prices can be attributed to the availability of land for sell in the region.

6. Discussion and Conclusion

Much academic and policy discussion accompanied the introduction of agricultural payments decoupled from production, debating the extent to which these payments, being attached to land only, get capitalized into agricultural land prices. A methodology is proposed in this paper to estimate the incidence of agricultural payments on farmland rents accounting for a number of specification issues which are typical of the rental price equation and which have been barely considered in previous micro-econometric literature at the EU level. The method allows consistent estimation of the parameters of interest in presence of selectivity and endogeneity of some explanatory variables in the framework of strongly unbalanced panels. Bothe are issue of substantial relevance in the specification of a rental price equation, as also suggested by the large number of theoretical contributions and empirical works which have so far considered them separately. Firstly, not all the farms rent land, requiring a panel probit model to be estimated to account for the relationship between the covariates and the probability to rent land in addition to the rent equation. Secondly, the main variables used to explain variation in the rental prices, land productivity and government payments, are likely correlated with the errors because the farmer’s optimal decision is taken based on expectations, which may actually differ from the current realizations used in the econometric model. Furthermore, the methods allows considering unobservable heterogeneity at the individual level in both the selection and the main equations.

The methodology is applied to a sample of Italian farms taken from the FADN databank. The focus is on farms specialized in field-cropping only and the time span considered is long enough to break the whole sample in to two sub-samples based on the implementation of coupled (1994-2004) or decoupled (2005-2008) payments.

Preliminary specification tests suggest that both the selectivity and endogeneity bias characterize the model specification and corrections are implemented accordingly. The application of the methods returns substantially different estimates compared to the standard instrumental variable estimation which does not consider the selectivity.

Results suggest that neither coupled nor decoupled payments have been capitalized into farmland rents. This is only apparently in contrast with previous empirical evidence on the EU. The extent to which the incidence of decoupled support has been estimated relevant for farmland rents depends in fact on the implementation methods adopted by each MS to introduce decoupled payments. Having Italy adopted a hybrid model, somehow more balanced toward historical payments, the incidence of decoupled payments was expected lower if not altogether negligible.

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