A PANEL DATA ANALYSIS OF THE DETERMINANTS OF FARMLAND PRICE: AN APPLICATION TO THE EFFECTS OF THE 1992 CAP REFORM IN BELGIUM

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Abstract

This study identifies the effects of the 1992 and subsequent CAP reforms on arable farmland price in Belgium. We first propose a brief literature review of studies identifying the determinants of farmland price. Afterwards, we use a panel data set to estimate a capitalization model of farmland price. We first show that the compensatory payments introduce by the 1992 CAP reform exert a positive effect on the arable farmland price in Belgium. We also identify a structural break in the land price equation after the adoption of the new support instruments as well as a regional break between Wallonia and Flanders.

Keywords: farmland price, land market, Common Agricultural Policy, capitalization of agricultural support

Jel classification: Q10, Q15, Q18

1. Introduction

The main objective of most agricultural support policies consists in rising farmers income. One of the major critics opposed to such policy is the weak efficiency of the income transfer. An important part of the income transferred to the agricultural sector leads to larger consumption of variable inputs and creates a rent that capitalizes into asset value. As a fixed factor, land is seen as very sensitive to this capitalization process.

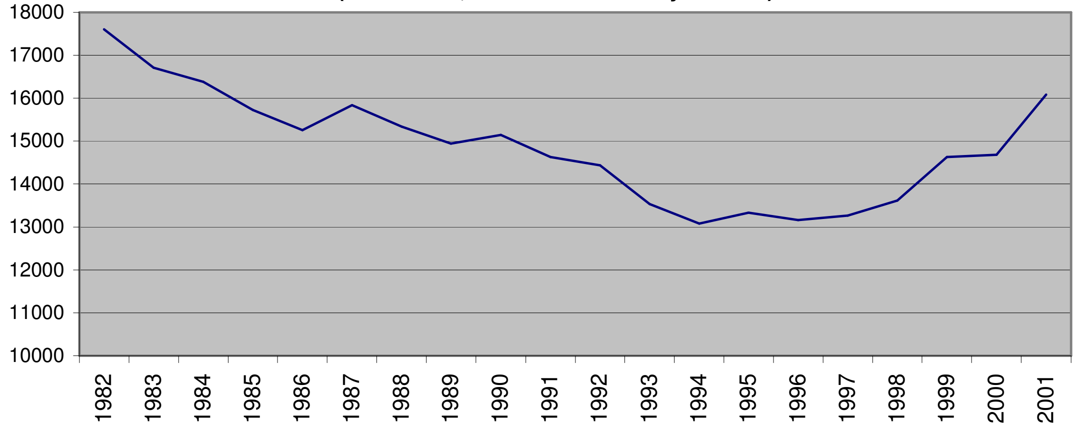
The rise of farmland value can be seen as favorable for an individual farm operator. Indeed, the land value plays a crucial role as collateral to obtain credits and determines the purchasing power of the farmer after retirement. However, that the benefits of the support are capitalized in farmland value rises some equity considerations. Because two thirds of Belgian farmland are cultivated by tenants, not land owners, an important part of the support is captured by non-operators. As inflated land values also imply inflated land rents, this type of policy might even make tenant farmers worse off. Moreover, if a farm operator purchases farmland with the expected future supports already capitalized in its price, he does not fully benefit from the support. Those benefits are used to finance a more expensive investment (Barnard *et al.*, 2001). Finally, the inflated farmland price due to direct payments makes more difficult for a young farmer to start his own business.

In this paper, we are interested in analyzing the effects of the 1992 and subsequent Common Agricultural Policy (CAP) reforms on the price of arable farmland in Belgium. The adoption of the 1992 reform corresponds to a strong inflexion in the CAP support instruments. Relying on price support since its foundation, the CAP progressively switches to direct payments decoupled from production. To understand the reasons of this significant change, we have to remember that the 1992 CAP reform also corresponds to a particular moment of the trade negotiation: the ongoing Uruguay Round. For the first time agriculture was on the global trade agenda. Many specialists pointed out that the elements of the 1992 reform permitted to achieve an agreement at the Uruguay Round. Other announced objectives of the reform were to promote rural development and contain environmental problems.

The adoption of the 1992 and subsequent reforms leads to a substantial cut of the price support but compensated with direct payments. For cereal, oilseed and protein crops, the amount of direct payments depend upon the cultivated area. Depending on the extent of the subsidised area, the direct payments can be conditional to a set aside provision. For beef production, the pre-existing premiums per head of cattle are increased. However, these premiums are conditional to a limit density of cattle per hectare of fodder crops. There is also a ceiling on the number of cattle heads eligible for premium. We see that the reform induces a stronger link between the eligibility to the benefits of the agricultural support and the use of land. Moreover, it is relatively well established that the effect of agricultural support on the price of land is stronger in case of direct payments decoupled from production. As a result, we can suspect that the CAP 1992 reform has caused a substantial change in the determinants of farmland price.

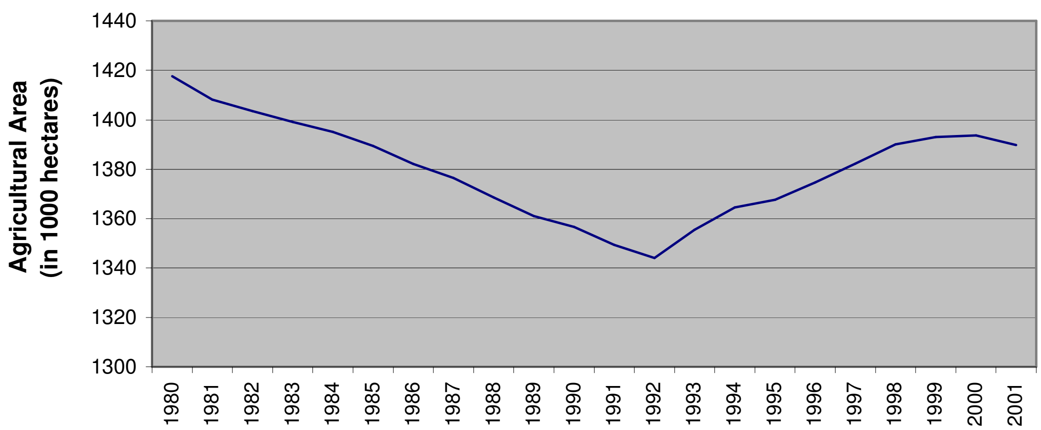
Figure 1 shows the evolution of the arable farmland price in real term during the 1980-2001 period in Belgium. We can see that after a continuous decrease up to 1994, the evolution of farmland price tends to stabilize and then increase in the years following the 1992 CAP reform. During the 1995-2001 period, the real price of arable land increases at a yearly average of about 3.8% in Belgium. Showing the evolution of the Belgian agricultural area, Figure 2 gives us another insight of the effect of the 1992 CAP reform. We can observe a significant change in the evolution of the series during the period succeeding the reform. Although that movement does not correspond to farmland acquisition, it allows us to conclude that the main message of the reform was well understood by farmers and landowners. They anticipate that forthcoming agricultural support will be linked to the farmland use. Consequently, they are more rigorous in reporting their agricultural acreage.

Figure 1 : Arable farmland price evolution
(1982-2001 ; in constant euro by hectare)



Source : Belgian National Institute of Statistics (NIS)

Figure 2 : Evolution of the Belgian Agricultural Area (1980-2001)



Source : Belgian Agricultural Census, Belgian National Institute of Statistics (NIS)

The remainder of this paper is organized as follow. The next section presents a brief literature review of studies identifying the determinants of farmland price. The following section discusses a theoretical model of farmland price formation. The fourth section presents the empirical model and the data used in the empirical analysis. The fifth section presents and interprets the results of the empirical analysis.

We show that the compensatory payments exert a positive effect on arable farmland prices. We also show that the compensatory payments have a distinct effect each year and that the extent to which the agricultural support is capitalized into farmland value is region specific. The final section of the paper offers a summary and concluding remarks.

2. Literature Review

Most research attempting to identify and quantify the determinants of farmland price rely on the capitalization approach. This approach assumes that the price of farmland equals the present value of all future expected cash flow attached to the use of land for productive purposes. In this context, an increasing farmland price should be explained by an increasing land rent. In USA, agricultural economists observed that the evolutions of the real land value and agricultural income go in the same direction from 1910 to 1950. Those trends convince agricultural economics to rely on the capitalization approach. However, a divergent evolution of land value and agricultural income is observed since 1950. Because of the observed decreasing agricultural income and increasing land price the classical model was rejected. Several alternatives hypothesis were proposed to explain the evolution of farmland price.

Feldstein (1980) proposes a first alternative explanation. He points out that the increasing farmland price observed in the seventies took place during a period characterized by a strong inflation. Feldstein (1980) proposes a portfolio choice model with two assets: a classical financial asset and land. He shows that an anticipated inflation could lead to a decrease in the actualization rate applied to land and explain an increase in farmland price. Other authors explain that as a real asset with fixed supply, land tends to hold its real value during inflationary periods. Consequently, there is an inflationary hedging motive to buy land during an inflationary period (Castle *et al.*, 1982). Most studies attempting to evaluate the effect of inflation on farmland prices found an insignificant effect. Alston (1986) shows that the inflation has a significant but negative effect on farmland price. On the basis of this result, Feldstein's hypothesis is rejected.

For Melichar (1979), the diverging trends of agricultural income and farmland price cannot conduct to reject the basic capitalization model. First, we need to consider that the net farm income usually used as a measure of the return to land is an inappropriate indicator. The net farm income corresponds to the return to land but also the return to farm labor and management. The land rent is the net farm income left after subtracting the return to farm labor and management. Second, the share of each type of remuneration in the net farm income cannot be considered as constant through time. During the second half of the twenty century, the capital-labor substitution phenomenon leads to an increase in the share of capital and, consequently, land in the net farm income and a decrease in the share of farm labor. Consequently, it is possible to reconcile both a decrease in the net farm income and an increase in the land rent explaining an increase in land value. Following Melichar's arguments, Alston (1986) shows that between 1962 and 1982 the periods characterized by an increasing land value are also characterized by an increasing land rent. Moreover, the rate of growth of the land rent and the rate of growth of land value are never statistically different.

A number of studies have attempted to estimate the extent to which support policies increase farmland price. Most of those studies investigate the relationship between the benefits of agricultural supports and the value of land using the present value approach. In the land market, the participants accept to pay a higher price as long as they expect to yield a larger stream of cash flow. The subsidies aimed to support the agricultural sector increase the cash flows attached to the property of farmland. Consequently, the agricultural support also increases the land value. In what follow, we present the most significant studies attempting to provide a quantitative estimate of the extent to which government payments are capitalized in farmland value.

Goodwin *et al.* (1992) estimate an empirical model relating land value to expected level of producer support, expected yield and expected producer prices net of government support in six wheat producing regions between 1979 and 1989. The government support is proxy by the producer subsidy equivalent. This study shows that a one percent increase in the producer subsidy equivalent increases land value by 0.38%.

Barnard *et al*. (1997) regress the cropland value on government subsidies and other characteristics like agricultural productivity, non-agricultural influence and state-specific institutional environments.

In their study, government payments are measured by the county-level averages of the annual amount of direct payments received par acre. The authors carry the same analysis for twenty U.S. regions. They show that the elasticity of cropland value to the government subsidies ranges from 0 to 0.69 according to the region. This study also points out that the sensibility of farmland value to government support is spatially variable. Two elements can explain this spatial variability; 1) whether or not the dominant crops in a given region are eligible to the support, and 2) the level of agronomic flexibility of a given region that determines the ability to adjust output in response to changing government policy.

Goodwin *et al.* (2005) use farm level data to estimate the capitalization rate of government payments into farmland value. They first note that the formation of land value is based upon expectations about the long-run stream of returns attached to land. Assuming that the realization of a particular source of return correctly reflects what is expected in the long run could lead to an error in variable problem. Error in variable problem results in inconsistent estimators. To represent the expected payment, the authors use a four-year average value of the realized payments at the county level. They first consider the aggregation of all support programs into a single category. They show that using the actual realized payment of each farm as proxy of the expected rent gives a coefficient of 5.40. With the county average, they obtain a coefficient of 6.09. They conclude that *" the use of the farm observed payments may result in an attenuation bias that forces the implied capitalization rates toward zero"*. Because the four-year county average of the payments is more representative of the long run benefits, a larger coefficient is obtained. They also show that the rate of capitalization of one dollar of payment is program specific. Breaking out the overall measure of government payments into their individual components, they find a specific rate of capitalization for each type of support program. Like Barnard *et al.* (1997), Goodwin *et al.* (2003) show that the extent to which support policies affect land value is spatially variable. They also point out that the implied effect of a support instrument on land value differs from year to year.

Clark *et al.* (1993a) point out a condition for the present value model to hold. The time series of rent and land values must be integer of the same order for the series of the rent to explain the movement in farmland value. Campbell *et al.* (1987) rigorously demonstrate that for the present value model to hold it is necessary that time series of land prices and expected rents are cointegrated. On the basis of this condition, Clark *et al.* (1993b) show that farm incomes alone cannot explain land values. Because the time series of land value is found to contain one unit root while the series of farm income is stationary, the present value model is rejected. However, a time series aggregating farm income and government payments is found to have a single unit root. Moreover, they find some empirical evidence suggesting a cointegration relationship between land value and income including subsidies. They conclude that farm income alone cannot explain land value. However, considering together farm incomes and subsidies, the present value formulation cannot be rejected. This could imply that subsidies are capitalized into farmland value.

Most studies interested in the explanation of farmland prices also take into account the demand of land for alternative uses (residential, recreational, etc.). Generally, the variable intended to capture the influence of non-agricultural demand appears to be significant. Goodwin *et al.* (2005) point out that each additional inhabitant per square mile in a given country adds \$1.85 in value to an acre of farmland. A 1% increase in the population growth rate adds \$64 per acre to the farmland value. Plantinga *et al.* (2002) show that the price of land reflects current as well as potential uses of land. If a potential buyer expects a conversion of farmland for a more profitable purpose in a forthcoming year, the current land value also reflects a conversion value associated with the future potential use of land. The authors show that the importance of that conversion value in farmland value depends upon two main factors: the proximity of large urban area and the availability of farmland. With French data, Cavailhès *et al.* (2003) show a negative relationship between the price of farmland and the distance from a central business district (CBD). The price of farmland is the highest the closest to the CBD. When the distance to CBD increases, the price of land declines reflecting a decrease in the conversion value. At about 40 km from the CBD, the price of farmland becomes independent of the distance. The authors conclude that when farmland is so far from a CBD, the value of that farmland in agricultural production is the sole determinant of farmland price.

3. The Theoretical Model : A Present-Value Model for Farmland Prices

Like for the value of any assets, the classical economic theory suggests that the value of a piece of agricultural land should be equal to the present value of all future expected cash flows. The cash flows associated with the use of land for agricultural production are called the rents (R), i.e., the income left after subtracting them from revenue of all other productive factors (labour, capital and variable inputs). If the price of land is below this fundamental value, buying land allows extra profits. In this case, the competition among potential buyers of land is such that the price increases. We follow Goodwin *et al.* (2002, 2005) and propose an equation explaining the formation of arable land price of the following form:

$$P_t = \int_0^{\infty} R^*_{t+j} \, e^{-r_t j} \, dj \qquad (1)$$

The actual value of land (P_t) is equal to the discounted expected value of all future expected rents from agricultural production (R^*_{t+j}). The term r_t represents the discount rate, e.g., the rate used by land market participants to discount future payments. Interested in estimating the extent to which the CAP compensatory payments are capitalized into farmland value, we break the expected rent from agricultural production into two components: a first component arising from the sale of the agricultural production in the market (MS) and a second component linked to the agricultural support (AS):

$$P_t = \int_0^{\infty} E_t(MS_{t+j}) \, e^{-e_t j} \, dj + \int_0^{\infty} A_t(AS_{t+j}) \, e^{-a_t j} dj \qquad (2)$$

Both functions $E_t(.)$ and $A_t(.)$ represent an expectation formation process for the expected revenue arising from the market and from the agricultural support. For example, the term $A_t(AS_{t+j})$ is the level of agricultural support expected in period t for the period t+j. The terms e_t and a_t are respectively the discount rates apply to the revenue from the sale to the market and to the agricultural support. In any present value model, the weight of any income source depends upon the opportunity cost of capital and a factor reflecting the expected variability of that particular income source. Indeed, if the agents are assumed to be risk averse, the anticipated variability of a future cash flow may affect the determination of the discount rate of that particular cash flow. If a given source of payment is seen as a more variable and/or more temporary, that particular payment takes a smaller weight in the present value. Consequently, we define each discount rate as follows:

$$e_t = o_t + ve_t$$
$$a_t = o_t + va_t$$

Where the term o_t is the opportunity cost of the capital engaged in agriculture in t and the terms ve_t and va_t respectively reflect the expected variability of the revenue arising from the sale to the market and from the agricultural support. In equation (2), the term:

$$A_t(AS_{t+j}) \, e^{-a_t j}$$

represents the present value in t of the agricultural support expected for the period t+j. We use a distinct discount rate for each income source to reflect that the expected variability is payment specific. If a given income stream is seen as more variable, it is discounted at a larger rate. We also constrain the terms $E_t(MR_{t+j})$ and $A_t(AS_{t+j})$ to be constant independently of the period t+j at which the payments is expected :

$$E_t(MR_{t+j}) = E_t(MR)$$
$$A_t(AS_{t+j}) = A_t(AS)$$

However, to remain as general as possible, we allow the participants to the land market to expect increasing (or decreasing) cash flow for each source of payment. For example, if the agricultural support is expected to grow at a constant rate ga_t each year, the present value in period t of a cash flow expected for the period t+j is be given by:

$$A_t\left(AS\right) e^{(-o_t - ve_t)\, j} \; e^{ga_t j} = A_t\left(AS\right) e^{(-o_t - ve_t + ga_t)\, j}$$

A similar expression can be obtained if land market participants anticipate increasing (or decreasing) revenue from the sale in the market at a constant rate ge_t. With the last modifications, equation 2 becomes:

$$P_t = \int_0^{\infty} E_t(MS)\, e^{(-o_t - ve_t + ge_t)\, j}\, dj + \int_0^{\infty} A_t(AS)\, e^{(-o_t - va_t + ga_t)\, j}\, dj \tag{3}$$

where ge_t and ga_t are respectively the rate of growth expected in t for the sources of payment $E_t(MR)$ and $A_t(AS)$.

Recognising that each integral corresponds to the present value of a perpetual constant flow, equation (3) simplifies as follows:

$$\begin{aligned} P_t &= \frac{1}{o_t + ve_t - ge_t} E_t(MS) + \frac{1}{o_t + va_t - ga_t} A_t(AS) \\ &= \alpha_t \; E_t(MS) + \beta_t \; A_t(AS) \end{aligned} \tag{4}$$

where α_t and β_t respectively represent the rate of capitalisation of the expected future revenues arising from the sale to the market and the rate of capitalisation of the expected agricultural support. Both terms α_t and β_t represent the reciprocal of the corresponding discount factor.

4. The Econometric Model and Variable Definition

The previous sections suggest a number of factors that may influence the price of arable farmland. This section defines the empirical model and the data used to estimate the determinants of arable farmland price. The empirical analysis proceeds from equation (5) that is estimated using a balanced panel of 42 Belgian districts, each of which has data for the entire period from 1980 to 2001.

$$PAL_{it} = \beta_{rt} EMR_{it} + \beta_{ct} CP_{it} + \beta_d PD_{it-1} + \beta_l \Delta PRL_{t\,t-1} + \beta_p PIG_{i,t-1} + \beta_m MA_{it} + \beta_s MS_{it} + u_{it} \tag{5}$$

The $PAL_{it,}$ variable is the average sale price of arable land (per hectare) observed in district i during year t. The variables EMR and CP are respectively used as indicators of the expected land rent from the sales to the market and the land rent from the agricultural support. Due to limitation in available data, we choose to approach the realized land rent by the sum of the agricultural family income and land costs by hectare. Those data are available in the yearly publications of the Belgian Centre for Agricultural Economics. Because the compensatory payments defined hereafter are also a component of the agricultural family income, we subtract the variable CP of the realized land rent to obtain the land rent arising from sales to the market. Of course, that variable is the current realisation of that particular rent and, consequently, it is a wrong indicator of the expected land rent from sales to the market. To approach the expected market rent, we follow Goodwin *et al.* (2005) and construct a four year average of that land rent realized during the current and past three years.

The variable CP is used to approach the level of support for cereals, oilseed and protein crops. Because it is impossible to obtain the realized payment per hectare, we choose to approach that particular type of support by the following formula:

$$CP_{it} = \sum_j \varpi_{jit} CP_{jit} \ \ with\ \varpi_{jit} = \frac{area_{jit}}{\sum_j area_{jit}}$$

with j = cereals, rapeseed, flax and fallow. The CP_{jit} and $area_{jit}$ variable are respectively the amount of support by hectare of crop j and the area of that crop in district i during year t.

Both variables $PD_{i\ t-1}$ and $\Delta PRL_{i\ t-1}$ are introduced to control for the expected future capital gains inherent in farmland in areas facing non-agricultural pressure. The variable PD is the population density per square kilometre. The variable ΔPRL is the growth rate of the price of residential land. We choose to express the price of residential land in growth rate to break a potential multicollinearity between PD and PRL since the correlation between PD and PRL is close to 0.7. The variables PAL, PD and PRL are directly available in the yearly publications of the Belgian National Institute for Statistics (NIS).

The variable PIG corresponds to the pig density by hectare of agricultural area. The series is calculated from the Belgian Agricultural Census (NIS). The pig density is introduced to control for a possible capitalization of manure spreading rent in the land market (Le Goffe *et al.*, 2005).

With all variables presented above farmland price is mainly explained by demand factors: agricultural profitability, government support, competing uses for land, etc. However, demand factors are the sole determinants of farmland price if and only if farmland supply is perfectly inelastic. We, however, hypothesize that the area of farmland entering into the market in a particular place at a particular date also plays a role in the explanation of farmland value. In other words, we want to control for supply factors explaining farmland price. The variables MS and MA are used to control for this supply. The variable MS represents the size of the arable farmland market. It is approached by the fraction of arable farmland exchanged in a particular district at a particular year. An increase in MS is expected to induce a decrease in the farmland price. The variable MA is the average area of farms in the district. This variable MA reflects the pressure of agricultural activities on farmland market. The variable MA is expected to yield a negative coefficient. The variables PIG, MA and MS are calculated from the Belgian Agricultural Census of NIS.

To correct for the effect of common inflationary pressure, all monetary variable (PAL, EMR, and CP and PRL) are adjusted to 2000 euros using the Gross Domestic Product deflator published in the Belgostat database. For the estimation procedure, all variables are transformed into natural logarithm to obtain a direct interpretation in terms of elasticity.

5. Estimation of the Land Price Equation

5.1. Specification Tests and Model selection

Our model of farmland formation presented in equation (5) can be estimated in several ways. The appropriate estimation method depends upon the structure of the error terms, u_{it}, and the correlation between the observed determinants of farmland prices and the different components of the error term. In what follows, we want to determine the most appropriate estimation technique. We start by assuming a time constant coefficient for both the variable EMR and CP.

The second column of Table 1 gives the OLS estimation of the farmland price equation. This model assumes that all OLS assumptions be satisfied and that the parameters of the equation be constant for all time periods and Belgian districts. However, it might be possible that both district-specific and time-specific unobservable factors affect on farmland price. In case of such time and/or district heterogeneity in the equation explaining the formation of farmland price, OLS estimators are biased. For that reason, the possibility of a cross-sectional or a time specific heterogeneity is considered in the four last columns.

In the district fixed effect (DFE) model, a dummy variable specific for each district accommodates for district heterogeneity. The fifth column of Table 1 presents results from the district random effect (DRE) model where the unobserved district effect is included in the error term. This error term is decomposed into two components: a cross-sectional specific error (v_i) and an individual error (ε_{it}). By introducing the individual effects in the error term, we assume that the intercept is a random outcome variable. The DRE and DFE models differ by one critical assumption: the DRE model assumes no correlation between the cross-sectional effects and the observed regressors. This assumption might be critical for the present study. Some unobserved parameters specific to a district that explain the formation of land price might also have an impact on some independent variables. For example, the agro-climatic conditions of a particular district might influence both farmland price and market rent. In presence of such correlation, the DRE estimator is inconsistent while the DFE estimator remains consistent. In the absence of such correlation both estimators are, however, consistent, but the DRE estimator is more efficient. If we are convinced that the district heterogeneity needs to be taken into account, we need to decide whether the DFE or DRE model is the more appropriate (Sevestre, 2002 ; Baltagi, 2001).

In a similar way, the possibility of a time-specific heterogeneity is considered in fourth and sixth columns of Table 1. The fourth column presents the results from the time fixed effect (TFE) model, which accommodates for time heterogeneity by introducing a dummy variable specific to each time series. The sixth column presents the results from the time random effects (TRE) model in which the error term accommodates for time unobserved heterogeneity.

Table 1. Estimation results of the farmland price equation (5) using ordinary least squares (OLS), time-fixed effect (TFE), district-fixed effect (DFE), time-random effect (TRE) and district-random effect (DRE) estimators (Numbers in parentheses are p-value).

VARIABLE	OLS	DFE	TFE	DRE	TRE
Expected Land Rent from Market Sales$_t$	0.020	-0.203	0.131	-0.189	0.095
(EMR_t)	(0.464)	(0.000)	(0.000)	(0.000)	(0.000)
Compensatory Payment $_t$ (CP_t)	-0.016	-0.025	0.342	-0.025	-0.009
	(0.000)	(0.000)	(0.000)	(0.000)	(0.055)
Population Density $_{t-1}$ (PD_{t-1})	0.228	0.039	0.200	0.260	0.225
	(0.000)	(0.854)	(0.000)	(0.000)	(0.000)
Price of residential Land Growth$_{t-1}$ (ΔPRL_{t-1})	-0.022	-0.002	0.016	-0.004	0.004
	(0.619)	(0.938)	(0.689)	(0.874)	(0.916)
Pig Density$_{t-1}$ (PIG_{t-1})	0.116	-0.021	0.113	0.877	0.115
	(0.000)	(0.527)	(0.000)	(0.000)	(0.000)
Mean Area $_t$ (MA_t)	-0.228	-0.352	-0.177	-0.387	-0.172
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Market Size$_t$ (MS_t)	-0.105	-0.104	-0.028	-0.111	-0.065
	(0.000)	(0.000)	(0.045)	(0.000)	(0.000)

To select the appropriate specification, two questions have to be addressed. First, is there evidence of a need to control for cross-sectional heterogeneity and time-effects? If no, we can rely on the OLS estimator. Second, if such evidence exists, can we reasonably assume that the random district and/or time effect is the Best Linear Unbiased Estimator (BLUE)?

To answer the first question, we first test the joint significance of all district effects in the DFE model. We perform this joint significance test by a traditional F-test. The value of the statistics is 28.65. Compared with a $F(41, 875)$ we reject the null hypothesis at more than one percent significance level. We now test the need to control for time effects by testing the joint significance of all time effects in the TFE model. The value of the statistics is 14.41. Compared with a $F(21, 895)$, we reject the null hypothesis.

The two F-tests indicate the need to accommodate with both unobservable district heterogeneity and unobservable time effects. We now have to decide on whether a fixed or random effect specification is the more reliable one. For the unobservable time effects, we test the assumption that the time specific effects are uncorrelated with the independent variables. Under this assumption, we choose the TRE model which is consistent and more efficient than the TFE model. Under the alternative, we choose the TFE model, which remains consistent. The former assumption is not directly testable. But, Hausman (1978) suggests an indirect test of the null hypothesis. Under the null hypothesis, both estimators are consistent and so converge to the same value. If the difference between the two estimators can be considered as systematic, we reject the null hypothesis and prefer to focus on the TFE model still consistent under the alternative. The results of the estimations reported in Table 1 indicates a large difference in some of the coefficients obtained from the TFE and TRE models, a sign of failure of the null of no correlation. The value of 41.11 of the Hausman statistics confirms this intuition. Under the null hypothesis, this statistics is distributed as a χ^2 with K degrees of freedom equal to the number of regressors. Consequently, the Hausman test rejects the null that the difference in coefficients is not systematic and we rely on the results on the TFE model. For the district heterogeneity, the Hausman statistics takes a value of 51.05. Once again, it is a sign of failure of the no correlation assumption and we prefer to rely on the results of the DFE model.

5.2 Correction for Heteroskedasticity and Serial Correlation

The specification tests indicate the need to control for both district unobservable heterogeneity and unobservable time effects. In both cases, the fixed effect specification seems to be more reliable than the random effect specification. For that reason, we present the results of a two-way fixed effect model. Moreover, we want to stay as general as possible and allow for the compensatory payment to have a distinct effect for each year between 1993 and 2001. Consequently, we estimate the following equation:

$$PAL_{it} = \alpha_i + \lambda_t + \beta_r EMR_{it} + \beta_{ct} CP_{it} + \beta_d PD_{it-1} + \beta_l \Delta PRL_{it-1} + \beta_p PIG_{i,t-1} + \beta_m MA_{it} + \beta_s MS_{it} + u_{it} \tag{6}$$

where the regression disturbances, u_{it}, are assumed to be homoskedastic and uncorrelated. If those assumptions are not satisfied, the two-way fixed effect model is still consistent but no more efficient. Before to present the estimation results, we test for a possible correlation in the residuals. To test the null hypothesis of no first-order serial correlation in the residuals, we use the Baltagi-Wu locally best invariant test statistics. The statistics is distributed as an N(0,1). To test Ho: $\rho = 0$ against Ha: $\rho>0$, we refer to the lower tail of the normal distribution. If the alternative is $\rho<0$, we compare the realisation of the statistics with the upper tail critical value of the distribution (Baltagi *et al.*, 1999). The realization of the statistics is 1.63. Consequently, we reject the null hypothesis of no serial correlation in the residuals. Another test proposed by Bhargava *et al.* (1982) is the generalization of the Durbin-Watson for the fixed effects model. The realisation of the statistics is 1.46. As in the classical Durbin-Watson case, the null hypothesis of no correlation in the residuals is rejected if the realisation of the statistics is below the lower bound critical value, and is not rejected if the statistics is larger than the upper bound critical value. The critical value are reported in Bhargava *et al.* (1982). As the realisation of the statistics is below 1.5, we strongly reject the null of no serial correlation in the residuals.

The rejection of the null hypothesis of no serial correlation indicates the need to correct the standard errors for serial correlation. Table 2 reports the results of the estimation of the two-way fixed effect model using robust standard error. We assume that the disturbance follows an AR(1) process with a specific coefficient for each cross-section. To correct for possible heteroskedasticity, we assume that each panel has its own variance.

The estimation results of equation (6) are presented in Table 2. Estimates of the district and time dummy variables (generally significant) are not presented here due to space limitation. Table 2 shows that the compensatory payments are consistently significant with the correct sign from 1996 to 2001. Both variables aimed to represent the supply side are also found to be significant with the expected sign. However, the variable EMR plays only a very weak effect on the price of arable farmland. An increase in EMR of 1% only induces a increase in farmland price of 0.078 %. Moreover, EMR is only significant at 6% significant level. Both variables representing the non-agricultural demand (PD_{t-1} and ΔPRL_{t-1}) are found to be insignificant.

That PD_{t-1} and ΔPRL_{t-1} are found insignificant and that the coefficient of EMR is low and weakly significant do not mean that those variables are not explicative of land price variation. To understand the source of these counterintuitive results, the third column of Table 2 presents the estimation results of the following equation:

$$PAL_{it} = \lambda_t + \beta_r EMR_{it} + \beta_{ct} CP_{it} + \beta_d PD_{it-1} + \beta_l \Delta PRL_{it-1} + \beta_p PIG_{i\,t-1} + \beta_m MA_{it} + \beta_s MS_{it} + u_{it} \tag{7}$$

Table 2. Estimation results of the farmland price equations (6), (7) and (8). Prais-Winston regression with panel corrected standard errors and AR[1] disturbances (Numbers in parentheses are p-value).

Variable	Equation 6	Equation 7	Equation 8 with τ=1993	Equation 8 with τ=1994
Expected Land Rent from Market	0.078	0.182		
$Sales_t$ (EMR_t)	(0.064)	(0.000)	-	-
Expected Land Rent from Market			0.215	0.244
$Sales_t$ (EMR_t) ($t<\tau$)	-	-	(0.000)	(0.000)
Expected Land Rent from Market			0.186	0.182
$Sales_t$ (EMR_t) ($t\geq\tau$)	-	-	(0.000)	(0.000)
Compensatory $Payment_{1993}$ (CP_{1993})	0.067	0.065	0.065	0.070
	(0.280)	(0.309)	(0.313)	(0.270)
Compensatory $Payment_{1994}$ (CP_{1994})	0.068	0.070	0.072	0.069
	(0.214)	(0.264)	(0.251)	(0.268)
Compensatory $Payment_{1995}$ (CP_{1995})	0.099	0.147	0.150	0.147
	(0.123)	(0.052)	(0.043)	(0.047)
Compensatory $Payment_{1996}$ (CP_{1996})	0.170	0.240	0.243	0.240
	(0.011)	(0.003)	(0.002)	(0.002)
Compensatory $Payment_{1997}$ (CP_{1997})	0.197	0.260	0.263	0.261
	(0.010)	(0.004)	(0.003)	(0.003)
Compensatory $Payment_{1998}$ (CP_{1998})	0.247	0.366	0.368	0.364
	(0.002)	(0.000)	(0.000)	(0.000)
Compensatory $Payment_{1999,}$ (CP_{1999})	0.174	0.287	0.286	0.282
	(0.014)	(0.001)	(0.000)	(0.000)
Compensatory $Payment_{2000}$ (CP_{2000})	0.340	0.467	0.470	0.466
	(0.000)	(0.000)	(0.000)	(0.000)
Compensatory $Payment_{2001}$ (CP_{2001})	0.246	0.377	0.380	0.378
	(0.000)	(0.000)	(0.000)	(0.000)
Population Density $_{t-1}$ (PD_{t-1})	0.079	0.189	0.187	0.185
	(0.697)	(0.000)	(0.000)	(0.000)
Price of residential Land $Growth_{t-1}$	0.004	0.005	0.007	0.006
(ΔPRL_{t-1})	(0.849)	(0.781)	(0.710)	(0.739)
Pig $Density_{t-1}$ (PIG_{t-1})	-0.006	0.099	0.098	0.098
	(0.840)	(0.000)	(0.000)	(0.000)
Mean Area $_t$ (MA_t)	-0.265	-0.197	-0.183	-0.175
	(0.001)	(0.000)	(0.000)	(0.000)
Market Size $_t$ (MS_t)	-0.063	-0.068	-0.067	-0.067
	(0.000)	(0.000)	(0.000)	(0.000)

Equation (7) only differs from equation (6) by removing the district dummies. Estimation results of equation (7) show a larger elasticity for the expected land rent from market sales and indicate that the population density has an expected and significant effect. That the elasticity of EMR is of smaller size in the estimation results of equation (6) indicates that the cross sectional difference across districts plays a very important role in explaining farmland value. In equation (6), the district dummies absorbed most of the cross-sectional effect of the expected land rent from market sales.

5.3 Tests of the Null Hypothesis of Non-Stationarity

The classical panel-data estimation method presented upon relies on the assumption that the series are stationary. The presence of non-stationarity series could invalidate the results presented above. When both dependent and independent variables are trend-dominated, we are very likely to find significant coefficient and high R^2, even in the absence of causality between those trending variables. This phenomenon known as the spurious correlation problem might explain the very attractive results presented above. We, therefore, examine the order of integration of our series using the panel unit root test proposed by Levin *et al.* (2002) and another similar test proposed by Im *et al.* (2003).

In both tests the null hypothesis assumes a unit root in the data generating process. In the Levin *et al.* test, it is assumed that the coefficient of the autoregressive process is common to each cross-section. In contrast, the Im *et al.* test allows a specific coefficient for each cross-section. More detailed information on the construction of those two tests can be founded in Baltagi (2001, chapter 12). For each test, we incorporate an intercept. It is left to decide whether or not it is appropriate to introduce a time trend. The inspection of the data indicates to introduce such a trend only for the EMR series. Of course, the absence of a significant trend in the series representing land prices mainly reflects that we have withdrawn the effect of inflation by working on real term data.

Table 3 presents the results of the Levin *et al.* and the Im *et al.* tests of the null hypothesis of unit root process for each variable. Both tests reject the null of non stationarity for the series representing the price of arable land, the expected land rent from market sales and the price of compensatory payments. That those series are found to be integer of the same order is to be linked with the results of the study of Clark *et al.* (1993). We cannot reject the assumption that the movements in the expected land rent from market sales and the compensatory payments explain the movements in land price. For the series of PD, PIG and MA, the Im *et al.* test does not allow to safely reject the null of non stationarity.

Table 3. Panel unit root tests

Variable	Levin, Lin and Chu		Im, Pesaran and Shin	
	Test statistic	p-value	Test statistic	p-value
Price of Arable Land (PAL)	-10.27	0.000	-9.73	0.000
Expected land Rent from the market sales (EMR)	-9.58	0.000	-3.86	0.000
Compensatory Payment (CP)	-14.90	0.000	-3.27	0.000
Population Density (PD)	-2.62	0.004	-0.47	0.31
Price of residential Land Growth (ΔPRL)	-7.59	0.000	-11.43	0.000
Pig Density (PIG)	-3.84	0.001	-1.13	0.12
Mean Area (MA)	11.09	1	18.11	1
Market Size (MS)	-5.57	0	-7.23	0

That three series are non-stationary could invalidate the results presented upon. However, by forming the linear combination of the PD, PIG and MA series using the estimated coefficient reported in Table 2 (third column), we find that the linear combination is stationary. The Levin *et al.* statistics takes a value of –15.18. We can reject the null of non stationarity of the linear combination at more than 1% significant level. This means that the left and right hand sides of equation (7) are balanced. Hence, the results of equation (7) presented in Table 2 cannot be invalidated by spurious correlation. Unfortunately, it is not possible to show that both parts of equation (6) are balanced by a similar argument. That both sides of equation (6) are unbalanced makes any inference drawn from the results of equation (6) questionable.

5.4 Structural Break Test

In what follows, we want to test the assumption that the 1992 CAP reform has caused a structural break in the land price equation. We expect to find that the elasticity of the variable EMR is smaller after the 1992 reform. The assumption of a decrease in the rate of capitalization of the expected land rent from market sales relies on two arguments. First, because of the 1992 reform and the concomitant price cut, the CAP instruments are less able to stabilise agricultural incomes. Consequently, the expected land rent from market sales could be expected to be more volatile and, hence, have a larger capitalization factor attached to it. Second, the adoption of the Agreement on Agriculture at the Uruguay Round together with the discussions concerning the future of the CAP instruments indicate subsequent price cuts and, hence, subsequent decrease in farm incomes. On the basis of equation (4), we can conclude that both a larger anticipated variability of the expected land rent from market sales (increase in ve_t) and an expected decreasing land rent from market sales ($ge_t<0$ for $t>1992$) should induce a decrease in the rate of capitalization of the expected land rent from market sales.

To test the existence of a structural break in a given year τ, we estimate an extended version of equation (7):

$$PAL_{it} = \lambda_t + \beta_{rt} EMR_{it} + \beta_{ct} CP_{it} + \beta_d PD_{it-1} + \beta_l \Delta PRL_{it-1} + \beta_p PIG_{i,t-1} + \beta_m MA_{it} + \beta_s MS_{it} + u_{it}$$
$$where\ \ \beta_{rt} = \beta_r\ if\ t < \tau\ and\ \beta_{rt} = \beta_r + \theta\ if\ t \geq \tau. \qquad (8)$$

Table 2 (fourth and fifth columns) presents the results of the estimation of two versions of equation (8). The fifth column of table 2 shows the estimation results of equation (8) introducing the possibility of a structural break at the first year of the introduction of the reform, i.e.,1993. The elasticity of EMR is quantitatively smaller in the period succeeding the break. We test the null of equality of the EMR elasticity before and after 1993. The statistics of the test takes a value of 1.71. Compared with a χ^2 with one degree of freedom, we cannot reject the null hypothesis of elasticity equality. The last column of Table 2 shows the estimation results of equation (8) considering the possibility of a structural break in 1994. Testing again the assumption of elasticity equality, we find a statistics of 5.66. Consequently, we can reject the hypothesis of elasticity equality.

5.5 Regional Effect of the CAP Compensatory Payments

Because the share of arable land planted in crops eligible to direct support (cereals, rapeseeds, flax and fallow) is greater in the Belgian region of Wallonia (50%) than in the region of Flanders (29%), we check whether or not the elasticity of the CAP compensatory payments have a regional-specific effect. We allow for all variables of equation (8) to have a separate effect in each region. Estimation results reported in Table 4 show that the elasticity of farmland price to compensatory payments is systematically stronger in Wallonia than in Flanders. On the basis of that result, we cannot reject that the regional specialisation of agriculture has an impact on the capitalisation rate of direct agricultural support. Results in Table 4 also show an increase in the rate of capitalisation of the compensatory payments in both regions. Those trends can reflect the expectation of the land market participants. An increase in the rate of capitalisation of the compensatory payments could be explained by the belief that this type of direct payments will continue in the forthcoming years.

Results in Table 4 also show that the rate of capitalization of the expected land rent from market sales remains steady in Flanders. In contrast, we observe a significant change in the rate of capitalization of that expected land rent in Wallonia. Before 1994, a 1% increase in the expected land rent from market sales leads to a 0.15 % increase in farmland price. Since 1994, the movements in the expected land rent from market sales badly explain farmland price movements. Once again this result must be linked to the agricultural specialisation in each part of Belgium. Due to its greater agricultural specialisation towards farm activities supported by interventionist prices, an important part of the expected land rent from market sales arises from the price support in Wallonia. Consequently, the 1992 and forthcoming announced intervention price cuts have a stronger impact on the rate of capitalisation of the expected land rent from market sales in Wallonia than in Flanders.

Table 4. Estimation of equation (8) with τ=1994 and allowing a regional specific elasticity for each variable. Prais-Winston regression with panel corrected standard errors and AR[1] disturbances (Numbers in parentheses are p-value).

Variable	Wallonia	Flanders
Expected Rent from Market Sales$_t$ (EMR$_t$)	0.150	0.189
(t<1993)	(0.012)	(0.003)
Expected Rent from Marke Sales$_t$ (EMR$_t$)	0.080	0.173
(t≥1994)	(0.158)	(0.003)
Compensatory Payment$_{1993}$ (CP$_{1993}$)	0.100	0.106
	(0.135)	(0.123)
Compensatory Payment$_{1994}$ (CP$_{1994}$)	0.121	0.056
	(0.094)	(0.480)
Compensatory Payment$_{1995}$ (CP$_{1995}$)	0.164	0.097
	(0.048)	(0.259)
Compensatory Payment$_{1996}$ (CP$_{1996}$)	0.254	0.191
	(0.004)	(0.026)
Compensatory Payment$_{1997}$ (CP$_{1997}$)	0.286	0.223
	(0.003)	(0.021)
Compensatory Payment$_{1998}$ (CP$_{1998}$)	0.386	0.341
	(0.000)	(0.001)
Compensatory Payment$_{1999,}$ (CP$_{1999}$)	0.284	0.226
	(0.002)	(0.016)
Compensatory Payment$_{2000}$ (CP$_{2000}$)	0.469	0.415
	(0.000)	(0.000)
Compensatory Payment$_{2001}$ (CP$_{2001}$)	0.384	0.335
	(0.000)	(0.000)
Population Density $_{t-1}$ (PD$_{t-1}$)	0.208	0.141
	(0.000)	(0.000)
Price of residential Land Growth$_{t-1}$ (ΔPRL$_{t-1}$)	-0.0003	0.015
	(0.992)	(0.493)
Pig Density$_{t-1}$ (PIG$_{t-1}$)	0.077	0.097
	(0.000)	(0.000)
Mean Area$_t$ (MA$_t$)	-0.201	-0.226
	(0.004)	(0.000)
Market Size$_t$ (MS$_t$)	-0.037	-0.080
	(0.063)	(0.019)

6. Summary and conclusion

Few existing studies provide information on the extent to which CAP support instruments are capitalized into farmland sale price. In this study, we estimate the effect of the CAP 1992 reform on the price of arable land in Belgium. Indeed, after a long period of declining prices, the real price of arable farmland increased at a yearly average of 3.8% during the 1995-2001 period. The CAP 1992 and subsequent reform also correspond to an important inflexion of the CAP support instruments. Starting from an intervention system largely based on the price support, the CAP progressively shifts to direct payments partially decoupled from production. However, economic theory suggests that farmland prices also reflect direct agricultural support.

Using a panel of 42 Belgian districts from 1980 to 2001, we first observed that the exchange price of arable farmland is affected by the compensatory payments. Depending on the year and region considered, the elasticity of arable farmland price to compensatory payments ranges from 0.12 to 0.47. Those parameters are in the same range of those found by other studies like, for example, Goodwin *et al.* (1992) and Barnard *et al.* (1997). This result also indicates that, by creating a rent that capitalizes into land value, the new CAP instruments also benefit to landowners. Because about two thirds of the Belgian agricultural land are rented by farmers, non-operators capture an important share of the agricultural direct support. Our results also indicate that a temporal variability exists in the elasticity of arable farmland value to compensatory payments. We show that the sensitivity of arable farmland values to compensatory payments increases during the 1993-2001 period.

All results obtained from the time fixed effect model show that, besides the time dummies and other control variables, the expected land rent from market sales exert an important effect on arable farmland price. Moreover, because results reported in Table 2 prevent to reject the presence of a structural break in the period subsequent to the CAP 1992 reform, we can conclude that the intervention price cut induced by the 1992 CAP reform conducts to a decrease in the rate of capitalisation of the expected land rent from market sales. Before the reform, an 1% increase in the expected land rent from market sales induces a farmland price increase of 0.24%. In the years following the reform, the farmland price elasticity declines to a value of 0.18%. Results from Table 4 also show that the sensitivity of arable farmland price to both expected land rent from market sales and compensatory payments is region specific. The elasticity of farmland price to compensatory payments is much stronger in Wallonia than in Flanders. Second, the decrease in the capitalization rate of the expected land rent from market sales is mostly attributable to a decrease in the capitalisation rate in Wallonia.

The comparison of the results of the two-way and one-way time effect models drawn from Table 2 provides us another important information concerning the determinants of farmland prices. The one-way time model gives a stronger and more significant elasticity of farmland prices to expected land rents from market sales than the two-way fixed effect model. To understand this, we need to consider that our measured expected land rent from market sales has two components. The first component is a strictly structural and time-invariant component. The second one is a transitory and time varying component. In the two-way fixed effect model, the district dummies absorb the structural component of the expected land rent from market sales. In contrast, in the one-way fixed effect model, the elasticity of farmland price to the expected land rent from market sales totally reflects the structural component. As a result, we can conclude that the structural determinants of the expected land rent from market sales are the sole determinants of farmland prices. Farmland price is insensitive to short-run expected rent fluctuations. This result is coherent with the theory stating that long-run expectations of rents are the true determinants of current land prices. Since our sample covers only 20 years, the long run expectations are absorbed in the district fixed effects.

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