Spatial panel models for the analysis of land prices

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1. Introduction

Farm real estate represents a dominant asset on the farm sector balance sheet in the U.S.A. (it accounted for nearly 84% of total U.S. farm assets in 2009) and is usually the largest investment in the farmers' portfolio. It is therefore considered to be an important indicator of the sector performance and of the producers' welfare (Nickerson et al., 2012). The real values of agricultural land have been increasing dramatically starting from the second half of 2000s, raising many questions about their macroeconomic determinants and whether the boom will turn into a bust (Gloy, 2012), especially after the 2008 global financial crisis. The analysis of land values also raises a number of policy issues, regarding government support, taxation and environmental protection.

For all these reasons, the empirical literature on the determinants of agricultural land values is extensive. The relationship between farmland prices and expected future returns on this asset has been extensively investigated in the past (see, for example, Falk, 1991; Engsted, 1998; Lence and Miller, 1999). However, despite the great amount of research efforts, most economic theories have only met small empirical evidence (Gutierrez et al., 2007).

This work investigates the spatial effects that may characterize the determination of agricultural land values in selected Midwestern U.S.A. States. We adopt the Ricardian Present Value Model (PVM) as the theoretical framework to address farm land values behavior in the long run. We specify and estimate spatiotemporal model that relates land value to its determinants. The spatial econometrics techniques we empoly - designed to account for the spatial effects that may characterize lattice data - represent an important methodological tool that has not yet been extensively applied in this field of analysis. We estimate a model that includes a spatial lag of the dependent variable to account for spatial dependence. We also characterize the temporal dynamics as an autoregressive process of first order. Finally, we present a spatiotemporal lag to account for all possible sources of autocorrelation in the data.

The theoretical framework of analysis and the employed dataset are presented in sections 2 and 3 respectively, while section 4 explores the spatial characteristics of the data. The results of our estimations are given and discussed in section 5. Section 6 presents the necessary checks of the stability conditions for the estimated model and the computation of long-run elasticities of cropland value with respect to the included regressors. Section 7 contains the final concluding remarks and the discussion of possible future developments.

2. The Present Value Model

2.1. The theoretical model

The PVM (Campbell and Shiller, 1988; Campbell et al., 1997) is a financial model that relates the price of a stock to its expected future returns discounted to the present using a constant or time-varying discount rate. When applied to the analysis of land values, we consider the price of the stock to be the price of land (in our case, the value of cropland, CV); the dividends are measured as cash rents (CR) received by the land owners. The value of cropland is therefore related to the capitalized value of the current and future stream of cash rents.

Following Gutierrez et al. (2007), we assume time-varying expected stock returns so that the relationship between prices and returns is non-linear and we define the log of the gross real rate of return on acre of land in State i from period t to $t+1$, (r_{t+1}), as

$$r_{t+1} \equiv log(CV_{t+1} + CR_{t+1}) - log(CV_t) \tag{1}$$

or equivalently

$$r_{t+1} \equiv cv_{t+1} - cv_t + log(1 + exp(s_{t+1})), \tag{2}$$

where $s_{t+1} = cr_{t+1} - cv_{t+1}$ is the natural logarithm of the dividend-price ratio (CR_{t+1}/CV_{t+1}), which is also called spread in financial literature. Lower case letters denote natural logarithms of the correspondent variables.

Equation (2) can be linearized using a first-order Taylor expansion into

$$r_{t+1} \approx k + s_t - \rho s_{t+1} + \Delta cr_{t+1}, \tag{3}$$

where $k = -log(\rho) - (1 - \rho) \cdot log(1/\rho - 1)$ and $\rho = 1/(1 + CR/CV)$. One should notice that equation (3) is a linear difference equation for the log stock price that can be solved forwardly and, under the condition that $\lim_{j\to\infty} \rho^j\, s_{t+j} = 0$, we obtain

$$s_t \approx -k/(1 - \rho) - \textstyle\sum_{j=0}^{\infty} \rho^j\, (\Delta cr_{t+1+j} - r_{t+1+j}). \tag{4}$$

According to equation (4), if the stock price is high today, then there must be some combination of high dividends and low stock returns in the future (Campbell et al., 1997). This relation holds ex-ante as much as ex-post, therefore taking expectations we obtain

$$s_t + k/(1 - \rho) \approx -\mathrm{E}_t\left[\textstyle\sum_{j=0}^{\infty} \rho^j\, (\Delta cr_{t+1+j} - r_{t+1+j})\right]. \tag{5}$$

The rationale of the PVM is embodied in equation (5) as it expresses the current value of the dividend-price ratio in terms of the present discounted value of expected future values of Δcr_{t+1} and r_{t+1}. The log dividend-price ratio is high only when dividends are expected to grow slowly or the expected stock returns are high and, when the dividend follows a log-linear unit-root process, the log dividend-price ratio is stationary provided that the expected stock return is stationary (Campbell et al., 1997). According to the PVM, if the agents are fully rational, then the asset prices (e.g. farmland values) and the dividends generated from that asset (e.g. cash rents) cannot drift persistently far apart from each other.

Let us also assume that the expected return to our asset $E_t[r_t]$ exceeds the expected return of another asset $E_t[g_t]$ by a constant r that represents the risk premium on investments on our asset; the PVM reduces to

$$s_t + (k - r)/(1 - \rho) \approx \mathrm{E}_t\left[\textstyle\sum_{j=0}^{\infty} \rho^j\, (g_{t+1+j} - \Delta cr_{t+1+j})\right]. \tag{6}$$

By supposing further that the expected rate of return on the alternative asset is stationary and that the logs of dividends and prices are non-stationary but their differences are, then it should be concluded that the RHS of equation (6) is stationary too and the constant expected excess returns version of the PVM holds. According to this finding, the PVM has been tested in the literature by estimating and then testing for cointegration the following equation

$$cv_t = \alpha + \beta cr_t + \varepsilon_t, \tag{7}$$

where $\alpha = -(k - r)/(1 - \rho)$ and ε is a zero-mean disturbance, or equivalently

$$s_t - \alpha = (1 - \beta)cr_t - \varepsilon_t. \quad (8)$$

If $\beta = 1$, intuitively, the log prices move one-to-one with log dividends and their unit-root components cancel out, leaving the spread unaffected. On the contrary, if $\beta \neq 1$, then $(1 - \beta)cr_t$ does not disappear and the spread is non-stationary (Gutierrez et al., 2007).

2.2. Empirical literature on the PVM and farmland prices

Many empirical studies on the determinants of farmland prices refer to the PVM as their theoretical framework. According to it, the value of an income-producing asset such as farmland is the capitalized value of the current and future stream of earnings from owning that asset (often measured, not exclusively, as cash rents). In other words, land values should equal the present value of all future expected cash flows stemming from a productive use of that land and therefore changes in expected returns to farming should explain changes in farmland prices (Du et al., 2007).

The empirical testing of the PVM has consisted in estimating equation (7) for each cross-sectional unit i and then testing the stationarity of the residuals by means of conventional cointegration tests. However, the empirical results do not fully support the PVM as the most appropriate for explaining farmland values. Among the empirical studies on this topic, we recall the analysis on farmland prices in Iowa conducted by Falk (1991), that ended up rejecting the PVM because, although highly correlated, farmland price and rent movements are not consistent with that. Clark et al. (1993) found similar results for Illinois, Tegene and Kuchler (1993) and Engsted (1998) for three U.S. regions (the Lake States, the Corn Belt and the Northern Plains). The failure to find cointegration is addressed by Gutierrez et al. (2007) by allowing structural breaks in the cointegrated relationship that represent a shifting risk premium on farmland investments, thus finding results in favor of the PVM.

Moving from the classical literature on PVM, some other trends have been gaining popularity in the analysis of farmland value. Some researchers concentrated on the influence of urbanization (Hardie et al. 2001; Plantinga et al. 2002; Livanis et al. 2006 among others); others focused on the testing of the PVM in presence of transaction costs (Lence and Miller, 1999; de Fontnouvelle and Lence, 2001). Important contributions tended to make distinctions among the streams of rents, particularly by arguing that farmland rents do not only consist in cash rents and that government payments should be considered as rent sources, but also distinguishing between different types of public subsidies (Clark et al., 1993; Weersink et al., 1999; Goodwin et al., 2003 among the others).

3. The data

All the employed data for the agricultural sector were made available by the United States Department of Agriculture (USDA), National Agricultural Statistics Service[1] (NASS) and Economic Research Service (ERS). The estimates of land values are based on annual survey data and report the market value[2] per acre of cropland only rather than farmland in general (in current dollars), so that problems arising from heterogeneity in land quality and use are limited (pastureland, for example, is not included). Cropland only includes the land used to grow field crops, vegetables or land harvested for hay. This also permits to exclude the value of farm buildings and take the value of land only into consideration.

Net cash rents per acre of cropland (in current dollars) are used, rather than gross cash rents, as this reflects the net return to the landowner (Alston, 1986). They measure returns to land from agricultural production, and can be interpreted as a Ricardian land rent. Besides this type of rent, agricultural support programs also represent a land return which may capitalize into land value. Direct government payments per acre of cropland, as estimated by the USDA-Economic Research Service, are therefore used as explanatory variables.

[1] http://www.nass.usda.gov/Quick_Stats/

[2] The land value is the value at which the land used for agricultural production can be sold under current market conditions, if allowed to remain on the market for a reasonable amount of time (USDA-NASS 2012).

All monetary variables were deflated using the GDP implicit price deflator (reference year 2005) from the U.S. Department of Commerce, Bureau of Economic Analysis.

Population density, calculated from the annual estimate of population from the U.S. Department of Commerce, Bureau of Census, is included among the covariates of the model as a proxy for urban pressure, that represents competing demand for land for non-agricultural use (Feichtinger and Salhofer, 2011).

The employed dataset is a panel of annual (1971-2009) observations for 12 Midwestern U.S. States (North Dakota, South Dakota, Nebraska, Minnesota, Iowa, Wisconsin, Illinois, Michigan, Indiana, Ohio, Arkansas, Mississippi), for which more homogeneous data are available, less affected by urban influence (e.g. those for North-eastern States). Moreover, cropland is mostly found in the Midwest, while the Western States, that have lower shares of cropland to total farmland, are less heavily surveyed by NASS for cash rents and the data on cropland per acre are either thinner or not available because sometimes limited only to either irrigated or non-irrigated cropland.

The availability of data on cropland value per acre for the selected variables turned out to be a constraint that led to the exclusion of States such as Louisiana, Missouri and Kansas form the original dataset. The availability of data on cash rents, only limited to 2009 for South Dakota, determined the time-span.

Thanks to the non-commonly considered variables, the employed dataset represents an improvement with respect to earlier studies. Although lower-level data might improve the analysis in terms of better theoretical explanation for spatial dependence (see Breustedt and Habermann, 2008, for a spatial analysis of farm-level cash rents) our focus on State-level data allowed to take a longer time-span into consideration.

4. Exploratory Spatial Data Analysis

Panel data have been frequently used in the field of agricultural economics, but spatial panel data have only recently started to be applied, although it is clear that location plays an important role (Baylis et al., 2001).

When aiming at modeling the spatial dimension of data and take into account the effects of spatial dependence and/or spatial heterogeneity that characterize them (Anselin, 1988), an Exploratory Spatial Data Analysis (ESDA) should be conducted in order to highlight the most appropriate specification of the model. It requires the definition of a spatial weight matrix as a square, non-stochastic and symmetric matrix, whose elements measure the intensity of the spatial connection between spatial units and take on a finite and non-negative value. The elements on the main diagonal are all equal to 0 by definition.

We choose to employ a row-standardized rook spatial weight matrix, $\mathbf{W}$, whose elements, w_{ij}, take on the values of either 0 or 1 depending on whether States i and j share some positive portion of their boundaries or not[3].

In order to determine whether there is overall spatial dependence among the observed cropland values we employed the well-known Moran's I index[4] and scatterplot. The Moran's I index (Table 1) shows significant positive values for all considered years, especially starting from the end of the 1990s, thus leading to reject the null hypothesis of no spatial dependence in favor of positive spatial dependence in the distribution of cropland values. Moran scatterplots[5] confirm that, albeit present in all considered years, spatial dependence appears to be stronger starting from the year 2000. We believe that exploiting the time dimension of the data conveys therefore pieces of information that cross-sectional data would ignore.

[3] We believe it represents a good average picture of the possible connectivity schemes. Nevertheless, the ESDA proved to be robust to the choice of different spatial weight matrices. Results are available upon request.

[4] Moran's I index is calculated as $\mathrm{I} = (\mathrm{n}/\mathrm{S})\mathrm{y}'\mathrm{W}\mathrm{y}(\mathrm{y}'\mathrm{y})^{-1}$, where S is the sum of all the elements of W and y is the vector of the n observations for the considered variable.

[5] Moran scatterplots are available upon request.

Table 1. Results for the Moran's I index for observed cropland value (1971 - 2009).

Year	Moran's I	p-value	Year	Moran's I	p-value	Year	Moran's I	p-value
1971	0.287	0.064	1984	0.414	0.021	1997	0.477	0.012
1972	0.322	0.047	1985	0.414	0.020	1998	0.489	0.011
1973	0.343	0.040	1986	0.460	0.012	1999	0.546	0.006
1974	0.297	0.059	1987	0.464	0.012	2000	0.601	0.003
1975	0.280	0.069	1988	0.356	0.034	2001	0.634	0.002
1976	0.277	0.073	1989	0.273	0.069	2002	0.657	0.002
1977	0.319	0.052	1990	0.291	0.062	2003	0.385	0.027
1978	0.301	0.057	1991	0.267	0.074	2004	0.637	0.002
1979	0.288	0.065	1992	0.279	0.067	2005	0.605	0.003
1980	0.297	0.062	1993	0.327	0.047	2006	0.597	0.003
1981	0.274	0.073	1994	0.286	0.053	2007	0.593	0.003
1982	0.265	0.076	1995	0.336	0.044	2008	0.572	0.004
1983	0.270	0.071	1996	0.327	0.044	2009	0.582	0.004

The results of the ESDA therefore give clear indication in favor of the estimation of a spatial model, capable of taking the spatial dependence among the observations of the dependent variable into account.

5. Estimation and discussion of the results

When dealing with observations that are collected both over space and time, there are numerous reasons to expect both serial dependence between the observations on each spatial unit over time and spatial dependence between the observations on the spatial units at each point in time to be present. This is because economic agents require time in order to collect information and make decisions and because what happens in neighboring locations influences these decisions. Following Elhorst (2010), since we treated space-time data, we conveniently chose to estimate a first-order autoregressive lag model in both space and time: the analysis on the determinants of cropland values in 12 U.S. States over the period 1971-2009 was conducted by estimating a model in which a spatial lag of the dependent variable is included, the temporal dynamics is described as an autoregressive process of first order and a spatiotemporal lag is also introduced so as to make our model a truly time-space dynamic model (Anselin, 2001).

Fixed individual effects were also added to the specification in order to take into account unobserved time-invariant sources of heterogeneity such as climate and land quality (Kirwan, 2009) and different sets of covariates were included, as described in equations (9) and (10):

$$cv_{it} = \lambda \boldsymbol{W} cv_{it} + \gamma cv_{it-1} + \varrho \boldsymbol{W} cv_{it-1} + \beta_1 cr_{it} + \beta_2 pd_{it} + c_i + \varepsilon_{it}; \tag{9}$$

$$cv_{it} = \lambda \boldsymbol{W} cv_{it} + \gamma cv_{it-1} + \varrho \boldsymbol{W} cv_{it-1} + \beta_1 cr_{it} + \beta_2 pd_{it} + \beta_3 gp_{it} + c_i + \varepsilon_{it}, \tag{10}$$

cv is the real cropland value, cr is the real net cash rent for cropland, pd is the population density and gp are real direct government payments. All variables were included in the model after a natural logarithm transformation. Models (9) and (10) were estimated by the Quasi-Maximum Likelihood (QML) estimator by Yu et al. (2008) and the results are shown in Table 2.

Table 2. QML estimates for the coefficients of models (9) and (10).

	Model (9)		Model (10)	
Coeff.	Estimate	t-stat	Estimate	t-stat
λ	0.382	8.899***	0.382	9.074***
γ	0.734	19.824***	0.713	20.359***
ϱ	-0.182	-3.254***	-0.187	-3.529***
β_1 (cr)	0.079	2.720***	-0.012	-0.415
β_2 (pd)	0.328	3.426***	0.548	5.659***
β_3 (gp)			-0.048	-6.906***

Significance level: ***=1% ($|t-\text{stat}| > 2.58$); **=5% ($|t-\text{stat}| > 1.96$) ; *=10% ($|t-\text{stat}| > 1.64$).

5.1. *The effects of net cash rents and population density on cropland values*

According to the PVM, we expect net cash rents to have a positive impact on cropland values. The estimation of model (9) (Table 2) indicates a significant, albeit limited, coefficient for the expected net cash rents (0.079), while population density shows a higher positive coefficient (0.328). Indeed, increasing population density may increase the demand for agricultural goods and therefore agricultural land and, at the same time, it may be sign of increasing urban pressure that enhances competing demand for land for non-agricultural use. A stronger effect of changes in population than of returns to farmland on farmland values has already been found for some U.S. regions by applying an entropy-based information approach: Salois et al. (2011) find that, although changes in farmland values are more strongly associated with changes in returns to farmland at the national level, the relationship appears to change over time and region and for some regions (Northeast, Corn Belt, Appalachia, Mountain and Pacific) population has become more informative.

The reasons for such limited effects of the covariates may be numerous. One possible explanation relies in the inclusion of State-specific fixed effects; some results in the literature already support the idea that they may absorb part of the cross-sectional effect of the expected land rent, thus suggesting that structural determinants of the expected rents are more effective in determining cropland value than short-run expected fluctuations (see Duvivier et al., 2005, for a study on a Belgian case). The high and highly significant estimates obtained for the spatial and temporal autoregressive coefficients (λ and γ) suggest that these may also absorb part of the effects of the covariates. The time-space autoregressive coefficient is also significant (ϱ), albeit negative and smaller in absolute value.

5.2. *The inclusion of government payments*

The inclusion of government payments as a covariate into the model does not return straightforward results (Table 2). First, the coefficient associated with direct government payments is significant and negative, indicating a negative impact of public subsidies on cropland value. This result is unexpected and requires deeper analysis and interpretation. Then, when we consider the effects on the other coefficients, it should be noted that the spatial and temporal effects are not significantly affected, whereas the inclusion of government payments enhances the impact of population density (whose coefficient rises from 0.328 to 0.548. Yet the most remarkable consequence is that caused on the estimates of β_1, that turn to be negative and not significant.

The empirical literature has already addressed the issue in various contributions that led to very different conclusions. A central point that should be taken into consideration concerns the fact that agricultural support policy instruments are thought to be highly correlated with land rents and this may cause multicollinearity in the estimates. Indeed, part of the literature concentrates on explaining the relationship between these two variables rather that their effect on land values, trying to assess whether agricultural policy benefits landowners of farmers the most (see, for example, Roberts et al., 2003; Lence and Mishra, 2003; Goodwin et al., 2004; Latruffe and Le Mouël, 2009; Kirwan, 2009).

Moreover, different types of subsidies are expected to have different impacts on cash rents and land values, therefore a distinction between the programs of agricultural support appears to be necessary in order to better interpret these results. Lence and Mishra (2003), for example, find that alternative farm programs have different effects on cash rents in Iowa, with positive effects of market loss assistance and production flexibility contracts, no effects of conservation reserve programs and a negative impact of deficiency payments. Similar conclusions are drawn by Goodwin et al. (2003), who argue that government payments cannot be considered to reflect the long-term expected stream of cash flows, which is the determinant of land values and is a latent variable. The only variables that can be observed and taken into account are the “market and government payment realizations for a sample of farms under a fixed set of policy instruments and market conditions” (p. 745). As Phipps (2003) argues, program payments are extremely variable from year to year and do not appear to have the characteristics of stability that should characterize expectations of returns to land for a given location and policy regime.

Feichtinger and Salhofer (2011) also find different capitalization rates for particular types of payments, with lower elasticity for agro-environmental payments, that often cause land rents to decrease.

The difficulties that arise as a consequence of the inclusion of government payments in the model are therefore numerous and the results obtained through model (10) can only be considered as an indication of the need of further research that takes into account the evolutions of agricultural policy in time and the differences in types of agricultural subsidies.

6. Short run and long run land value elasticity

The estimated β_1 and β_2 coefficients cannot be interpreted exactly as the elasticity of land value to, respectively, cash rents and population density, because of the presence of the variable cv on the RHS of model (9). Another contribution we make is therefore to provide an estimation of the impact and long-run elasticity of cropland values in response to changes in net cash rents and population density.

Before applying long-run value effect analysis, we test the series stationarity, in order to be sure that the process we are analyzing is not an explosive one. In order to do so, from equation (9) we define the N × N matrix

$$\boldsymbol{A} = (\boldsymbol{I} - \lambda \boldsymbol{W})^{-1}(\gamma \boldsymbol{I} + \varrho \boldsymbol{W}) \tag{11}$$

where $\boldsymbol{I}$ is an $N \times N$ identity matrix and $\boldsymbol{W}$ is an exogenous spatial weight matrix of the same dimensions.

Using $\boldsymbol{A}$ we can re-write model (9) as

$$cv_{it} = \boldsymbol{A} cv_{t-1} + (\boldsymbol{I} - \lambda \boldsymbol{W})^{-1}(\beta_1 cr_{it} + \beta_2 pd_{it} + c_i + \varepsilon_{it}) \tag{12}$$

The stability conditions of the process described in equation (12) can be now analyzed by computing the eigenvalues of the $\boldsymbol{A}$ matrix.

Depending on the eigenvalues, i.e. the characteristic roots of $\boldsymbol{A}$, we have three possible cases. When all the roots are less than 1 in absolute value, we call it a stable case. When all the roots are equal to 1, we term it a pure unit root case, which generalizes the unit root dynamic panel data model in the time series literature to include spatial elements. When some of the roots (but not all) are equal to 1, we define it as a spatial cointegration case, where the unit roots in the process are generated with mixed time and spatial dimensions.

Using the estimates obtained in section 5 for the autoregressive parameters by using a rook spatial weight matrix[6] ($\hat{\gamma} = 0.734$; $\hat{\lambda} = 0.382$; $\hat{\varrho} = -0.182$), we find the following eigenvalues of matrix $\boldsymbol{A}$ $[0.893, 0.850, 0.773, 0.759, 0.735, 0.710, 0.681, 0.696, 0.693, 0.692, 0.893, 0.663]$. Since all the values are less than 1, we can conclude that the system is stable. Hence the computation of elasticities for cash rents and population density is possible and can be easily done by solving the dynamic equation (12), i.e.

$$cv_{it} = (\boldsymbol{I} - \boldsymbol{A}\mathrm{L})^{-1}(\boldsymbol{I} - \lambda \boldsymbol{W})^{-1}(\beta_1 cr_{it} + \beta_2 pd_{it} + c_i + \varepsilon_{it}). \tag{13}$$

where L is the lag operator, that operates on an element of a time series to produce the previous element, such that, given $X = \{X_1, X_2, X_3, \dots\}$, $X_{it}L = X_{t-1}$, for all $t > 1$.

Using the estimates $\hat{\beta}_1$=0.079 and $\hat{\beta}_2$=0.328 and $t = 0, \dots, 100$, we find that the impact elasticity of cropland value (i.e. the elasticity calculated at $t = 0$) is equal to 0.13 with respect to cash rents and 0.53 with respect to population density. These values represent the expected immediate

[6] The results lead to the same conclusions when the estimates obtained by using the other spatial weight matrices are used in the computations.

percentage changes that a 1% percent change in, respectively, cash rents and population density would cause on cropland values.

Considering long-run impacts instead, the calculated long-run elasticity of cropland value with respect to a 1% increase in cash rents is equal to 1.2, while the long-run elasticity of cropland value with respect to a 1% increase in population density is equal to 4.97 (Figure 1). About 50% of the long-run impact of both cash rents and population density on cropland value is already reached after 6 years and the percentage increases up to 90% after 21 years. Therefore, in the long-run, the effect of population density (hence, according to our assumptions, of urban pressure and competing land uses) is significantly higher than that of cash rents in determining cropland values.

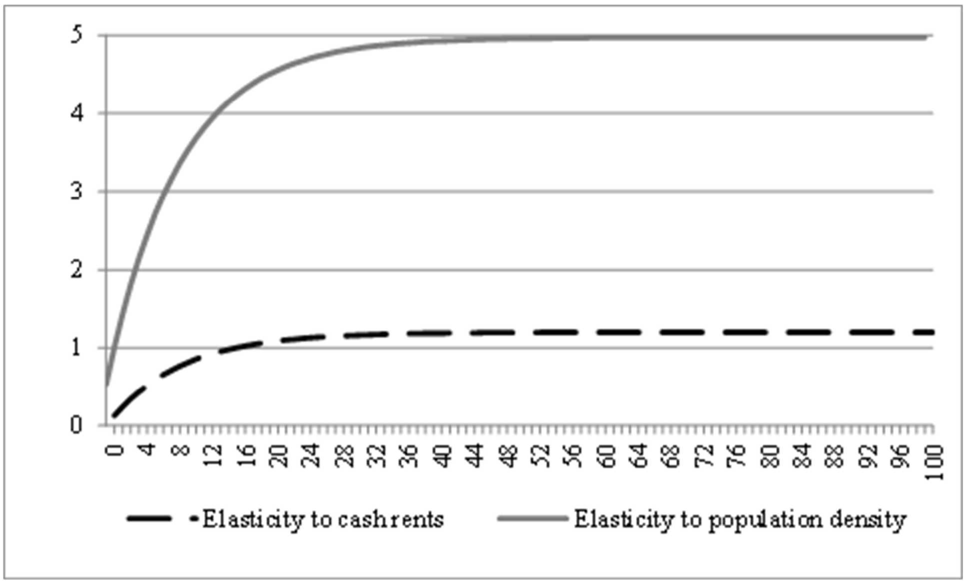

Figure 1. Long-run elasticity of cropland value with respect to net cash rents and population density.

Such a close-to-unity estimated long-run elasticity of cropland values to cash rents is close to what one would expect according to the PVM and that is usually not verified in empirical analyses. Gutierrez et al. (2007) find similar results by allowing for structural breaks in the cointegration relationship between the two time series, for a large panel of 31 U.S. States for the period 1960-2000. Previous empirical contributions, mainly based on time-series analysis, lead to different conclusions and, as previously said, end up rejecting the PVM and generally finding evidence of divergence between the present value of future cash flows and the market price of farmland (Falk, 1991; Clark et al., 1993a; Engsted, 1998).

7. Concluding remarks

The analysis of the determinants of land value in the U.S.A. is a relevant field of study given the importance of farm real estate on the farm balance sheet and because of the great number of policy issues that it raises. We adopted the PVM framework, according to which the value of land is the capitalized value of the current and future stream of earnings from owing that asset. In order to consider a more homogeneous dataset, only 12 States of Midwestern U.S.A., for which more reliable agricultural data are available, were included in the analysis and only cropland was taken into consideration when collecting data on land value and cash rents. Our model also introduced population density among the regressors as a proxy for urban pressure, in order to take into account the effects that competing alternative land uses might exert.

Although a fairly large body of literature has been devoted to this topic, spatial econometrics has only found limited application in this empirical field so far. We believe, as the ESDA confirmed, that data on land values are characterized by effects of spatial dependence that should be taken into account in estimating an econometric model that aims at explaining the factors that contribute to land value formation. In order to do so, we chose to estimate a model in which a spatial lag of the dependent variable is included. The temporal dynamics is described as an autoregressive process of first order and a spatiotemporal lag was also introduced so as to make our model a truly time-space dynamic model.

The results that we obtained confirm the existence of significant spatial and temporal dependence and therefore the need to take them into consideration. Our estimate of the long-run elasticity of cropland value with respect to net cash rents, which is close to unity, is an element favorable to the validity of the PVM assumptions. This is a result that has found only limited support in the literature on land values, which generally ends up rejecting the PVM. Gutierrez et al. (2007) find similar evidence in favor of the theoretical model when allowing for structural breaks in the time series. However, further checks on the estimated elasticity of 1.2 are required before drawing a conclusion on this. The effect of cash rents in determining land values is smaller than that of population density, which also has a positive significant effect on cropland values. Both variables appear to exert the biggest part of their influence on land values in about 20 years, as the computation of long-run elasticities revealed, even if about half of that impact is already reached after about 6 years.

The inclusion of government payments among the regressors was motivated by the fact that they can also be considered as an expected future stream of earnings from owning land, with relevant policy implications. However, the obtained results so far do not allow to draw final conclusions on the impact of agricultural support programs on cropland values. As suggested by the vast literature on this topic, a deeper reasoning and more disaggregated data are needed in order to provide a better model specification, capable of taking into account the evolution of U.S. agricultural policy in time and the differences between various instruments of government intervention.

Future developments of this analysis should therefore follow two main paths. On the methodological point of view, the econometric model that was estimated is one that has not been widely employed in empirical analyses, because of the complexity of its estimation and the lack of already available routines in econometric software. No standard and widely known testing procedures are available yet. Nevertheless we consider running precise specification testing as a priority in order to complete the present analysis. Moreover, following Gutierrez et al. (2007), the model should also be tested for structural breaks that may occur in the time series. This is not only a methodological extension of the study because detecting and allowing for structural breaks may also serve as a means for adding to the analysis of government support intervention. A deeper reasoning on the role of government payments and the best way to treat available data on policy intervention is also a path that should be followed.

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