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Paul Feichtinger and Klaus Salhofer

A Spatial Analysis of Agricultural Land Prices in Bavaria

ABSTRACT

This paper empirically analyses a dataset of more than 7,300 agricultural land sales transactions from 2001 and 2007 to identify the factors influencing agricultural land prices in Bavaria. We use a general spatial model, which combines a spatial lag and a spatial error model, and in addition account for endogeneity introduced by the spatially lagged dependent variable as well as other explanatory variables. Our findings confirm the strong influence of agricultural factors such as land productivity, of variables describing the regional land market structure, and of non-agricultural factors such as urban pressure on agricultural land prices. Moreover, the involvement of public authorities as a seller or buyer increases sales prices in Bavaria. We find a significant capitalisation of government support payments into agricultural land, where a decrease of direct payments by 1% would decrease land prices in 2007 and 2001 by 0.27% and 0.06%, respectively. In addition, we confirm strong spatial relationships in our dataset. Neglecting this leads to biased estimates, especially if aggregated data is used. We find that the price of a specific plot increases by 0.24% when sales prices in surrounding areas increase by 1%.

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FACTOR MARKETS Coordination: Centre for European Policy Studies (CEPS), 1 Place du Congrès, 1000 Brussels, Belgium Tel: +32 (0)2 229 3911 • Fax: +32 (0)2 229 4151 • E-mail: info@factormarkets.eu **• web:** www.factormarkets.eu

Contents

List of Tables and Figures

A Spatial Analysis of Agricultural Land Prices in Bavaria

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Factor Markets Working Paper No. 50/June 2013

1. Introduction

The importance of land for farming is beyond doubt. The determinants of land value formation were therefore an important research issue more than 200 years ago (Smith, 1776; Ricardo, 1817; von Thünen, 1842) and have continued to be so since (Lloyd, 1920; Bean, 1938; Scofield, 1957; Klinefelter, 1973; Robison et al., 1985; Shaik et al., 2005) and the matter is not yet fully resolved. Empirical analyses of land prices became more common from the 1960s on. At the same time, the question of the extent to which governmental payments capitalise into agricultural land gained importance. In empirical work, the relevant explanatory variable is realised governmental payments. But as the influence of payments on land price is based on expected values rather than observed values (e.g. Goodwin et al., 2003), one faces an endogeneity problem due to expectation errors. Goodwin et al. (2010) solve this problem by using lagged values of government payments as instruments. In addition to measurement errors introduced through expectation errors, endogeneity in land price studies may be caused by omitted variables or simultaneous determination of the dependent variable and at least one RHS variable.

Until about a decade ago, the spatial dimension of land use and land markets was ignored in land price studies. Land and its specific characteristics such as non-movability lead to a limited spatial extension of farms and to regional land markets. Spatially closer land markets interact with higher intensity than more distant ones. Moreover, in econometric applications, explanatory variables (e.g. precipitation indices) are often fitted to the dimension of the dependent variable land price by spatial interpolation, which can lead to spatially correlated error terms (Anselin, 2002). Hardie et al. (2001) were one of the first to use a spatial error model to account for spatial autocorrelation in the disturbances. Patton and McErlean (2003) also used a spatial error model, while Huang et al. (2006) used a spatial lag model in analysing the determinants of Illinois land prices. [1]

First, we define the land sales price as an outcome of the interaction between supply and demand within an area and the interaction of land markets in neighbouring areas, similarly to Fingleton and Le Gallo (2008). Second, we apply a general spatial model, combining a spatial lag model and a spatial error model to account for spatial dependence and for spatial autocorrelation in the disturbances, to a dataset of more than 7,300 actual arm's length sales transactions of agricultural land in Bavaria. Additionally, we account for endogeneity introduced by explanatory variables other than the spatially lagged dependent variable. According to Fingleton and Le Gallo (2008), this is usually ignored in econometric analysis.

2. Method

Following Fingleton and Le Gallo (2008), we model the observed agricultural land sales price in a specific area as the outcome of the interaction between land supply and demand in this

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* Paul Feichtinger is a PhD student and Klaus Salhofer a professor in the Environmental Economics and Agricultural Policy Group at Technische Universität München (Paul.Feichtinger@tum.de).

[1] Extensive literature reviews on determinants of land prices are provided by Le Mouël (2003), Latruffe and Le Mouël (2009) and Feichtinger and Salhofer (2013).

area and the interaction with land markets in neighbouring areas. Specifically, the quantity of agricultural land demanded in area i (q_i) is modelled as a linear function,

$$q_i = \alpha_0 + \alpha_p\, p_i + \alpha_w \sum_{j \neq i} W_{ij}^D p_j + \sum_{y=1}^{Y} a_y A_{y,i} + \omega_i \tag{1}$$

where p_i (p_j) is the price of agricultural land in area i (j), W_{ij}^D is a row standardised spatial weight matrix with zero elements for non-neighbours and zero elements in the diagonal, $A_{y,i}$ are Y demand shifting variables such as soil quality or distance to the nearest market, ω_i is a stochastic error term with $\omega_i \sim iidN(0, \sigma_\omega^2)$ and all α are coefficients to be estimated. In accordance with standard economic theory, we assume $\alpha_p \leq 0$. High prices for land in area j, which is in close proximity to area i, will reduce demand for land in area j. As a consequence, some demand will be displaced from j to i. Hence, q_i is positively related to the weighted average of land prices in the surrounding areas ($W_{ij}^D p_j$) and $\alpha_w \leq 0$.

Analogously, the supply of agricultural land (q_i) in area i can be modelled as

$$q_i = \beta_0 + \beta_p\, p_i + \beta_w \sum_{j \neq i} W_{ij}^S p_j + \sum_{z=1}^{Z} \beta_z B_{z,i} + \zeta_i \tag{2}$$

where W_{ij}^S is again a row standardised spatial weight matrix, $B_{z,i}$ are Z supply-side shifters such as the share of rented land in a municipality,[2] ζ_i is a stochastic error term with $\zeta_i \sim iidN(0, \sigma_\zeta^2)$ and all β are coefficients to be estimated. In accordance with standard economic theory, we assume $\beta_p \geq 0$. In contrast to the demand side spill-over effect, we assume a negative influence of the weighted average prices in the surrounding areas ($W_{ij}^S p_j$) on the quantity supplied in area i (q_i), because high prices in area j cause a displacement of supply from nearby i to j ($\beta_w \leq 0$).

Based on equations (1) and (2), and the simplifying assumption that $W^E = W^D = W^S$, we can derive a reduced form pricing equation

$$p_i = \gamma_0 + \gamma_w \sum_{j \neq i} W_{ij}^E p_j + \sum_{k=1}^{K} \gamma_k X_{k,i} + \varepsilon_i \tag{3}$$

where $X_{k,i}$ are $K = Y + Z$ variables of demand and supply shifters. γ are coefficients to be estimated, with $\gamma_0 = \frac{\alpha_0 - \beta_0}{\alpha_p + \beta_p}$, $\gamma_w = \frac{\alpha_w + \beta_w}{\alpha_p + \beta_p}$, $\gamma_k = \frac{\alpha_y}{\alpha_p + \beta_p}$ for demand shifters and $\gamma_k = \frac{-\beta_z}{\alpha_p + \beta_p}$ for supply shifters. Moreover, since $\omega_i \sim iidN(0,\sigma_\omega^2)$ and $\zeta_i \sim iidN(0,\sigma_\zeta^2)$ it holds that $\varepsilon_i = \frac{\beta_P \omega_i - \zeta_i}{\alpha_p + \beta_p}$ $\sim iidN(0,\sigma_\varepsilon^2)$. Equation (3) is recognisable as the well-known spatial lag model (Ord, 1975).

Although equation (3) accounts for spatial dependence, the potential problem of spatial autocorrelation in the disturbances remains. One reason for this might be spatially autocorrelated omitted variables, an inherent problem in land price analysis. In overcoming this problem, spatial error processes are implemented into error terms, with the spatial autoregressive model (SAR) and the spatial moving average model (SMA) being the most common specifications. In the SAR model, an assumed shock in area i is gradually transmitted to all other areas, since the areas are all connected with each other to some degree (global autocorrelation). In contrast, a shock is transmitted only to neighbouring

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[2] Before selling the land, landowners often rent it out for some years. A larger share of rented land may indicate a high number of landowners willing to sell.

areas in the SMA model (local autocorrelation). Hence, the range of the effect is much smaller (Anselin, 2003). As it seems likely in the case of agricultural land markets that a shock in area i is transmitted to more distant units, we choose the SAR model for our error term. Moreover, this seems to be consistent with the (global) autoregressive process of our spatial lag formulation.

Including a SAR model, the error term of equation (3) becomes

$$\varepsilon_i = \gamma_e \sum_{j \neq i} W_{ij}^E \varepsilon_j + \upsilon_i \, , \tag{4}$$

where γ_e is the spatial error coefficient to be estimated and υ_i is an uncorrelated error term $\upsilon_i \sim iidN(0,\sigma_\upsilon^2)$. While a spatial lag coefficient γ_w has a direct interpretation, a SAR model is implemented to obtain unbiased estimates.[3]

3. Data

We utilise a comprehensive dataset of (almost) all arm's length agricultural land sales transactions in Bavaria for the years 2001 (4,055) and 2007 (4,574). The dataset includes transaction-specific information on sales price, soil quality, plot size, municipality affiliation, and whether a public authority was involved as a seller or buyer. Farm take-overs by descendants are not captured in our data; the amount a successive farmer has to pay to other legal heirs as their compulsory portion of inheritance is usually considerably lower than the farm's actual market value (van der Veen et al., 2002).

From this dataset, we exclude plots already legally converted for housing development, land with a special usage such as excavation areas for gravel or sand, and land that also contains buildings. Furthermore, we try to exclude sales not primarily motivated by agricultural usage and therefore do not consider transacted plots smaller than 0.25 hectares. Such plots are more likely to inherit specific rights and easements (e.g. prospective non-agricultural land use) and this may result in price premia that are difficult to capture in our estimations given the information available. To account for other exceptional circumstances (e.g. agricultural land bought by non-farmers in a scenic area at a high premium, or fictitious purchases between closely related persons), we exclude transactions at prices lower (higher) than 2,000 (110,324) €/ha.[4] Additionally, we omit transactions with implausible values such as a soil quality index lower than 7 or higher than 85, or a price/soil quality ratio above 20.[5]

With these restrictions, we are left with 7,369 total observations for the years 2001 (3,539) and 2007 (3,830). On average, sale transactions took place in about 1,200 of the 2,056 Bavarian municipalities in each year. The shape of Bavarian municipalities and the location of municipalities where transactions took place in 2001 are shown in Figures 1a and 1b. At least one transaction took place in either 2001 or 2007 in 1,567 different municipalities. Table 1 shows that a plot of agricultural land sold for 22,642.32 €/ha (21,749.12 €/ha) in 2001 (2007) on average. The dataset does not allow us to distinguish between arable land

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[3] LeSage (1999) and LeSage and Pace (2009) provide, among others, extensive reviews of different spatial models.

[4] Before excluding outliers, we observe an average sales price of 25,289 €/ha with a standard deviation of 28,345 €/ha. To define outliers, we use three standard deviations from the mean as an upper bound. Considering that defining the lower bound analogously would lead us to keep (non-plausible) negative values, one would assume zero to be an appropriate lower bound. To additionally exclude transactions with (exceptionally) low sales prices, we determine 2,000 €/ha instead of zero as a lower bound. After accounting for outliers, our average sales price decreases to 22,198 €/ha with a standard deviation of 14,257 €/ha.

[5] An index system is used in Germany to indicate the soil quality of agricultural land. The point index ranges from zero to 100, with values for Bavaria of between 7 and 85 (Lfl, 2007).

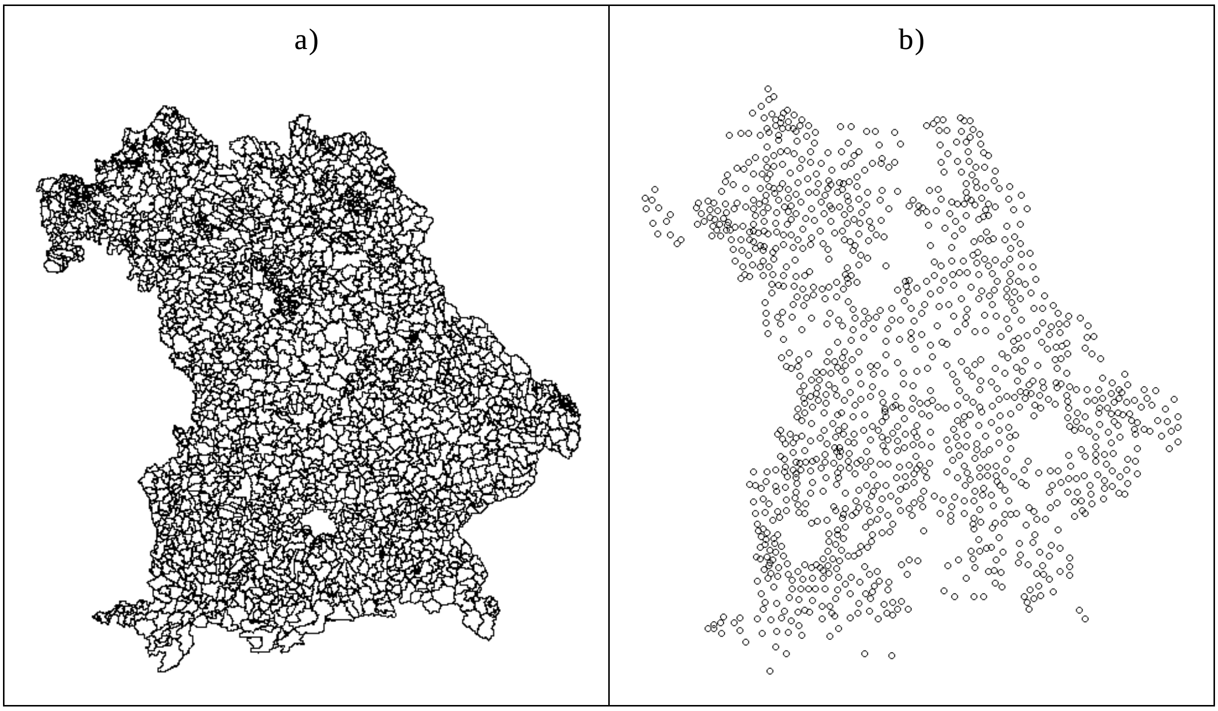
and grassland. In 2001 (2007), public institutions such as municipalities were buyers in 22% (13%) of all transactions. Purchased plots are often dedicated to infrastructure development in the future or are similarly handed over to a landowner as compensation for some of his land being dedicated to developing infrastructure. Public institutions acted as sellers in 2.5% (3%) of the sale transactions. State and municipalities own agricultural land mostly for historical reasons, and this share accounts for transactions of such land as well of sales of plots left over from infrastructure development projects.

Table 1. Descriptive statistics (after excluding outliers)

		No.Obs*	Mean	SD	Min	Max
					2001	
Sales price	€/ha	3,539	22,642.32	14,332.16	2,044.20	102,260.10
Public seller	%	3,539	3.33			
Public buyer	%	3,539	21.87			
Soil quality rating	pt.	3,539	45.19	13.07	7.18	84.00
Size of transacted plot	ha	3,539	1.67	2.26	0.25	73.44
Distance to the next urban centre	km	1,211	29.01	14.14	1.00	80.61
Direct payments	€/ha	1,211	261.28	92.21	7.36	469.03
Share of rented agricultural area	%	82	44.25	10.47	12.75	77.66
Price of building plots	€/m²	82	83.09	66.13	19.21	727.84
Ratio building vs. agricultural land		82	9.43	11.12	2.11	198.24
					2007	
Sales price	€/ha	3,830	21,749.12	14,109.23	2,026.75	102,300.00
Public seller	%	3,830	2.45			
Public buyer	%	3,830	12.74			
Soil quality rating	pt.	3,830	45.50	12.67	7.47	84.00
Size of transacted plot	ha	3,830	1.76	1.94	0.25	31.76
Distance to the next urban centre	km	1,196	29.00	14.62	1.00	72.49
Direct payments	€/ha	1,196	350.31	53.23	122.03	707.74
Share of rented agricultural area	%	86	51.38	9.96	19.26	78.17
Price of building plots	€/m²	86	71.74	50.01	16.07	331.17
Ratio building vs. agricultural land		86	18.15	20.92	2.58	252.84

* A total of 7,369 transactions took place in 1,567 different municipalities and 92 different districts.

Figure 1. Bavarian municipalities (a) and municipalities where transactions took place in 2001 (b)



Source: Authors' presentation.

To broaden the scope of our analysis, we add information on the distance to the next urban centre at the municipality level as well as district averages of the share of rented agricultural land in relation to total agricultural land, the sales prices for building plots, and a ratio of building versus agricultural land. These variables account for regional differences in urban pressure and agricultural land market structure.[6] The average shares of rented agricultural area (in total utilised agricultural area) at the district level vary from 13% to 78%. On average, a transacted plot had a size of about 1.70 ha. The variable helps to explain how economies of scale of larger plots might outweigh potentially greater difficulties in obtaining financing to purchase them. In addition, descriptive statistics reveal that prices of land for construction are, on average, 35 times higher than prices of agricultural land. We also add average direct payments in the respective municipality to account for the fact that agricultural subsidies may capitalise into land values to some extent. The first year of observation (2001) represents the time before the Fischler Reform of the Common Agricultural Policy, and hence includes mainly coupled direct payments for area and animals. The second year (2007) is after the Fischler Reform, with decoupled single farm payments. On average across all municipalities, producers received approximately 261 €/ha/year in 2001 and 350 €/ha/year in 2007 as direct payments (either coupled or decoupled). Low averages in some municipalities, such as the minimum value of 7.36 €/ha (122 €/ha) in 2001 (2007), indicate a comparably high share of milk production, whereas high values (maximum of 469.03 €/ha in 2001 and 707.74 €/ha in 2007) are a sign that arable farming in combination with intensive beef production is predominant.

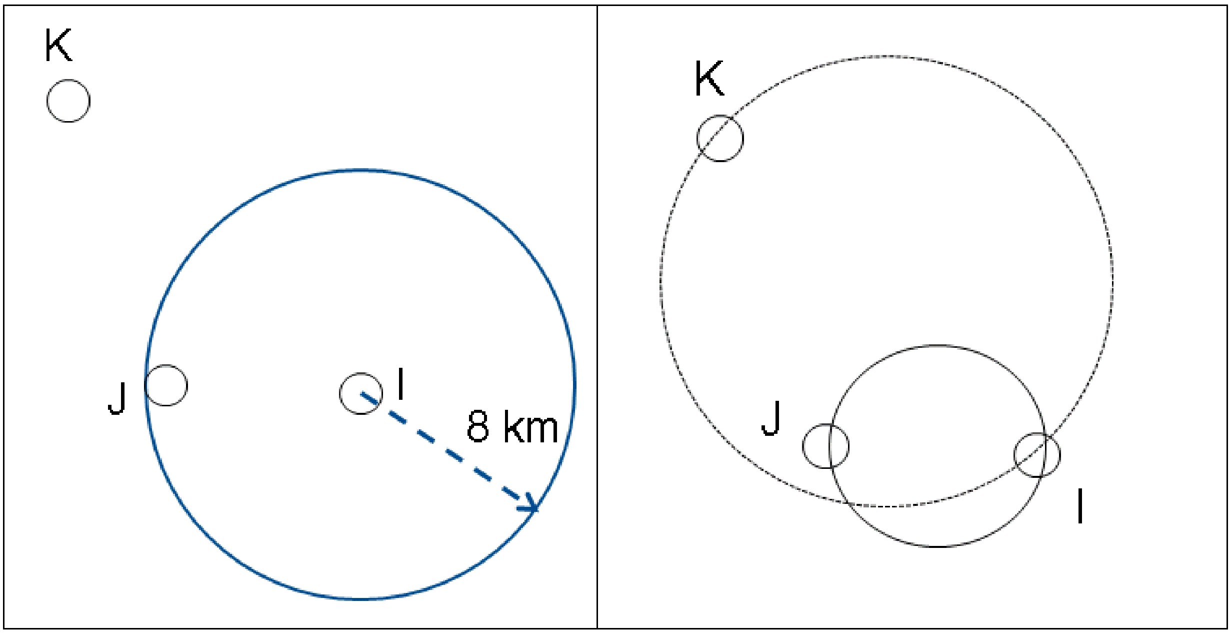
[6] It should be noted that mean and standard deviation of variables based on municipality and district averages are sample weighted because the 7,369 transactions are unequally distributed between municipalities.

4. Empirical Implementation

One critical point in every spatial econometric analysis is the *ex ante* specification of the *n* x *n* spatial weights matrix. What defines neighbours, as well as the weights given to each neighbour, has to be determined exogenously. In our case, *n* is the number of agricultural land transactions in the particular year. Because a land transaction cannot be a neighbouring transaction of itself, diagonal elements of the matrix are zero as are elements for non-neighbours. In contrast, elements for neighbours are non-zero. With regard to choosing weights of neighbours, two commonly used approaches are binary weights assigning a 1 to each neighbour and the inverse distance weight $1/d$, where *d* is the distance between two units. While in the first approach all neighbours are weighted equally, geographically closer transactions would be weighted stronger than more distant transactions in the second approach. We use binary weights as we lack information on the exact location of a transacted plot within a municipality. For the same reason, we assume municipality centroids as the location of any transacted plot in a municipality.

In defining the criteria for whether transactions are neighbours or not, we use two different approaches (Figures 2a and 2b).[7] In the first, a transacted plot (area *j*) is a neighbour of the transacted plot in question (area *i*) if the municipality centroid (*J*) is within a circle of eight kilometres from the centroid of area *i* (*I*). This is depicted in the left panel of Figure 2. In some municipalities, multiple transactions take place in one year and since those transactions are clearly within a circle of eight kilometres, they are also neighbours. Though not necessarily closer in distance to the transaction in question, they are intuitively more closely connected, for example because the flow of information is probably highest within a municipality.

Figure 2. Distance-based and Gabriel-based neighbour definition



Source: Authors' presentation.

In the second approach, illustrated in the right panel of Figure 2, closed discs are drawn between municipality centroids. Areas *i* and *j* are considered neighbours if a closed disc between their centroids (*I* and *J*) contains no other centroids.[8] *K* is not a neighbour of *I* under

[7] Practical advice in defining neighbours and creating weight matrices can be found in Bivand et al. (2008).

[8] The second approach, called a Gabriel graph, was first discussed in Gabriel and Sokal (1969). For an application we refer to Bivand and Brunstad (2006).

either of the two definitions because (a) it is outside an eight kilometre circle, and (b) a closed disc between I and K contains J. When using a distance-based neighbour definition, about 20 transactions per year have to be dropped from our sample due to a lack of neighbours. The reasons for this are generally a low number of sales transactions in the whole region or only one single transaction in a large municipality with the next municipalities' centroids being further away than eight kilometres. An advantage of the second approach is that according to the definition, every area i has at least one neighbouring area j.

Based on those two approaches, we derive two different row standardised weight matrices with every row summing to one, independent of the actual number of neighbours and whether weights are binary or (inverse) distances. This implies decreasing impacts of the single transaction with a rising number of neighbours. Moreover, a row standardised matrix is not symmetric and a transaction in area j may influence a transaction in area i differently from in the reverse case. Most importantly, a row standardised form allows to interpret the coefficient as the weighted average effect of land prices in the surrounding areas ($W_{ij}^{E} p_j$) on land prices in area i (p_i).

Although we have given some theoretical justification for a spatial lag model in Section 2, we also statistically test for spatial autocorrelation using a Moran's I test and for spatial autoregressive processes in the dependent variable as well as the residuals using Lagrange Multiplier (LM) tests. While in Moran's I tests positive (negative) values indicate positive (negative) spatial autocorrelation, values close to zero indicate no autocorrelation. According to Table 2, H_0 of no spatial autocorrelation is rejected in all our cases at the 99% level.[9] To assess the specific form of spatial autocorrelation and to decide whether a spatial error or a spatial lag specification is more appropriate, LM tests are used most frequently. Burridge (1980) proposed a LM test for spatial autoregressive processes in the error term ($H_0: \gamma_e = 0$), while Anselin (1988) proposed a LM test for spatial autoregressive processes in the dependent variable ($H_0: \gamma_w = 0$). The LM test results in Table 2 confirm spatial autoregressive processes in the residuals as well as the dependent variable. In such a case of rejecting both H_0 of no spatial autocorrelation, the robust test versions can be used to determine which autoregressive process is predominant.[10] Robust test version results confirm spatial autoregressive processes in the residuals as well as the dependent variable for all cases. Hence, Moran's I and LM tests confirm (on empirical grounds) the use of a general spatial model as in equation (3) including a decomposed error term as in equation (4).

Table 2. Testing for spatial autocorrelation and for spatial autoregressive processes for alternative spatial weight matrices and spatial dependence test results

	2001		**2007**	
Weight matrix	**Distance based**	**Gabriel**	**Distance based**	**Gabriel**
No. of observations	3539		3830	
No. of municipalities	1211		1196	
Average no. of links	15.32	18.50	16.06	21.07
Breusch-Pagan test for heteroskedasticity	79.40 ***		111.59 ***	
Jarque – Bera test on normality of errors	12.05 ***		85.99 ***	
Moran's I test	0.265 ***	0.246 ***	0.187 ***	0.155 ***
LM error	1,166.80 ***	1,564.58 ***	663.97 ***	729.21 ***
Robust LM error	142.64 ***	300.29 ***	55.43 ***	116.12 ***
LM lag	1,052.77 ***	1,304.67 ***	681.85 ***	699.13 ***

[9] A formula for a Moran's I test is provided by Florax and de Graaff (2004).

[10] Formulae for LM tests can be found in Anselin (2001) and for their robust versions in Florax and de Graaff (2004).

Robust LM lag	28.61 ***	40.38 ***	73.31 ***	86.04 ***

I = Distance based neighbour definition; II = Gabriel neighbour definition.

***p<0,01, **p<0,05, *p<0,10; SE = Standard Error.

Source: Authors' calculations.

Based on non-spatial OLS and instrumental variable estimations, we carry out a series of further tests. As discussed earlier, endogeneity of RHS variables arises from three different sources (Wooldridge, 2010):

1. Measurement error: This is the case if we cannot observe the actual explanatory variable, but only an imperfect measure of it. This is relevant in our analysis in regard to the expectation error with direct payments (Patton et al., 2008). Having incomplete foresight, buyers of agricultural land have to form some expectations about the magnitude and duration of future payments. Another example in this regard is the expectation error about future market returns since, for example, prices for outputs and inputs are not known for certain.
2. Simultaneity: This arises when at least one of the RHS variables is determined simultaneously with the LHS variable. In our analysis, this might be the case for several RHS variables including the "share of rented agricultural area", the "price of building plots", and also "direct payments" since the amount of payments is usually positively correlated with the quality and therefore the price of land. This problem can also evolve from the reduced form nature of our model (see Section 2).
3. Omitted variables: This is the case if a RHS variable is correlated with a variable that influences agricultural land prices but is unavailable in our analysis. Obviously this might be the case for all of our RHS variables.

To detect and to identify appropriate instruments to deal with endogeneity, we apply different tests for it and for instrument weakness. The latter is carried out by looking at the R^2 of the first stage instrumental variable regressions and the Cragg-Donald (1993) statistic as proposed by Stock and Yogo (2001).[11] We can reject instrument weakness for all our RHS variables except for "public authorities as a seller" and "public authorities as a buyer". Since weak instruments can lead to biased inferences on instrument variable estimations, we assume them to be exogenous. We extensively test for endogeneity in many different combinations of all other potential endogenous variables and instruments in two different ways. Using a Durbin-Wu-Hausman test, we can test whether a subset of the endogenous variables are actually exogenous. This is calculated by running a secondary estimation where the test variables are treated as exogenous rather than endogenous, and then comparing the J-statistic between this secondary estimation and the original estimation. In addition, we apply a regression-based test as discussed in Wooldridge (2010, pp. 119). In a first stage, all potentially endogenous explanatory variables of our land price model are regressed on all exogenous variables and instruments based on OLS. In a second stage, we estimate our usual land price model but include the error term of the first stage regression. A variable is endogenous if and only if the added error term has no influence on land prices in the second stage. We can test for this with a standard t-test accounting for heterscedasticity in the usual manner. We apply tests to all RHS variables of our land price model (except "public authority as seller" and "public authority as buyer") in different combinations of instruments and potentially endogenous variables. We obtain a strong indication that the "land quality index" is exogenous. This makes perfect sense given that land quality is defined by natural conditions completely exogenous to our system, although it does not fit perfectly into the expectation error problem with regard to future market returns if the "land quality index" serves as a proxy for expected future market returns. We obtain a mixed indication of endogeneity for the "ratio of building vs. agricultural land". It was clearly endogenous in

[11] The Cragg-Donald statistic is only valid for TSLS and other K-class estimators. However, results of the TSLS and the GMM estimations are very similar in all respects.

2007, but exogenous in 2001. We obtain a clear indication of endogeneity for all other variables. Given this, we decided to keep "public authorities as a seller", "public authorities as a buyer" and the "land quality index" as exogenous and all other variables endogenous. Furthermore, a Breusch-Pagan test clearly indicates heterscedasticity for both years and in a Jarque-Bera test we reject the H_0 of normally distributed error terms.

Given all our tests and considerations above, we estimate a model including endogeneity of the spatial lag and of other RHS variables and with heteroscedastic innovations:

$$\boldsymbol{y} = \boldsymbol{Y}\pi + \boldsymbol{X}\beta + \gamma_w \boldsymbol{W}\boldsymbol{y} + u \tag{5}$$

$$\text{u} = \gamma_e \boldsymbol{W}\boldsymbol{u} + \epsilon \tag{6}$$

where $\boldsymbol{y}$ is an $n \times 1$ vector of observations on the dependent variable; $\boldsymbol{Y}$ is an $n \times p$ matrix of observations on p RHS endogenous variables and π is the corresponding $p \times 1$ parameter vector; $\boldsymbol{X}$ is an $n \times k$ matrix of observations on k RHS exogenous variables and β is the corresponding $p \times 1$ parameter vector; $\boldsymbol{W}$ is a $n \times n$ spatial-weighting matrix and γ_w and γ_e are the corresponding spatial parameters; and ϵ is an $n \times 1$ vector of independently but heteroscedastically distributed innovations (Drukker et al., 2011).

The model is estimated utilising a GMM estimation strategy as discussed in Kelejian and Prucha (e.g. 1999, 2010) and Drukker et al. (2013) consisting of two steps of alternating GM and IV estimators each again consisting of sub-steps.[12]

Following Bivand and Piras (2013), we use as instruments all exogenous variables ($\boldsymbol{X}$; public authority as seller, public authority as buyer, land quality index), their spatial lags ($\boldsymbol{WX}$) and squared spatial lags ($\boldsymbol{W^2X}$) as well as some additional instruments expected to determine $\boldsymbol{Y}$ (two-year lags of direct payments, share of rented agricultural area, the ratio of building vs. agricultural land; one-year lag of the price of building plots; municipality averages of livestock units per hectare; average size of agricultural land parcels; standard gross margin per farm in a municipality).

5. Results

Results for the heteroscedasticity-consistent spatial GMM estimator with an endogenous spatial lag and additional endogenous variables as described in the last section are reported in Table 3 for 2007 and in Table 4 for 2001. These results are based on the distance-based spatial weight matrices. We also report non-spatial White heteroskedasticity-consistent OLS and GMM estimates for comparison. For both years, the highly significant spatial lag (γ_w) and spatial error (γ_e) coefficients indicate the accuracy of our spatial model. Hence, we concentrate on the interpretation of the spatial estimation results. A spatial lag coefficient of 0.24 (0.33) indicates that agricultural land sales prices in area i increase by about 0.24% (0.33%) when sales prices in surrounding areas increase by 1%. In addition, all other model parameters are highly significant except for the "distance to the next urban centre" in the estimation for 2001.

Table 3. Regression results for 2007 using non-spatial OLS and GMM and spatial GMM with a distance based spatial weight matrix

		OLS	**GMM**	**Spatial GMM**			
				coeff.	direct	indirect	total
Constant	*coeff.*	5.1998 ***	5.4920 ***	4.2916 ***			
SE		0.3439	0.4346	0.5810			
Public seller		0.2279 ***	0.2214 ***	0.2072 ***	0.2084 ***	0.0669 **	0.2753 ***

[12] We refer to Piras (2010, 2013) for detailed information on the estimator and the implementation into R.

	0.0501	0.0591	0.0528	0.0530	0.0280	0.0737
Public buyer	0.3414 ***	0.3368 ***	0.3023 ***	0.3050 ***	0.0975 ***	0.4025 ***
	0.0254	0.0280	0.0264	0.0264	0.0317	0.0447
Soil quality rating	0.7996 ***	0.7377 ***	0.6769 ***	0.6819 ***	0.2176 ***	0.8994 ***
	0.0306	0.0341	0.0367	0.0358	0.0674	0.0723
Size of transacted plot	0.0275 ***	0.0926 *	0.1061 **	0.1073 **	0.0341 *	0.1415 **
	0.0091	0.0522	0.0484	0.0481	0.0190	0.0640
Distance to the next urban centre	-0.0610 ***	-0.2410 ***	-0.1607 ***	-0.1611 ***	-0.0490 ***	-0.2102 ***
	0.0123	0.0376	0.0501	0.0503	0.0165	0.0599
Direct payments	0.2941 ***	0.4213 ***	0.2094 **	0.2106 **	0.0630 ***	0.2736 ***
	0.0610	0.0767	0.0864	0.0869	0.0265	0.1063
Share of rented agricultural area	-0.0169 ***	-0.0179 ***	-0.0132 ***	-0.0133 ***	-0.0041 ***	-0.0174 ***
	0.0009	0.0011	0.0017	0.0017	0.0011	0.0017
Price of building plots	0.1006 ***	0.0917 ***	0.0499 **	0.0501 **	0.0153 ***	0.0655 **
	0.0153	0.0191	0.0229	0.0230	0.0078	0.0292
Ratio building vs. agricultural land	0.1702 ***	0.1125 ***	0.1001 ***	0.1008 ***	0.0319 ***	0.1327 ***
	0.0142	0.0242	0.0264	0.0263	0.0125	0.0349
Spatial lag			0.2428 ***			
			0.0606			
Spatial error			0.1943 **			
			0.0805			
Adjusted R-squared	0.4038	0.3547				

***p<0,01, **p<0,05, *p<0,10; SE = Standard Error.

Source: Authors' calculations.

Table 4. Regression results for 2001 using non-spatial OLS and GMM and spatial GMM with a distance based spatial weight matrix

	OLS	GMM	Spatial GMM			
			coeff.	direct	indirect	total
Constant *coeff.*	6.9206 ***	6.2228 ***	3.9200 ***			
SE	0.1618	0.2648	0.6384			
Public seller	0.2605 ***	0.2891 ***	0.2341 ***	0.2374 ***	0.1140 ***	0.3515 ***
	0.0483	0.0529	0.0435	0.0438	0.0411	0.0729
Public buyer	0.2562 ***	0.2822 ***	0.2156 ***	0.2192 ***	0.1050 ***	0.3242 ***
	0.0201	0.0220	0.0200	0.0198	0.0323	0.0394
Soil quality rating	0.6487 ***	0.6121 ***	0.5868 ***	0.5957 ***	0.2865 ***	0.8822 ***
	0.0301	0.0327	0.0338	0.0335	0.0895	0.0980
Size of transacted plot	0.0302 ***	0.2124 ***	0.1404 ***	0.1430 ***	0.0673 **	0.2103 ***
	0.0100	0.0482	0.0525	0.0538	0.0309	0.0787
Distance to the next urban centre	- *** 0.0890	0.0341	-0.0236	-0.0236	-0.0112	- 0.0348
	0.0130	0.0416	0.0459	0.0462	0.0237	0.0691
Direct payments	0.0616 ***	0.0970 ***	0.0441 *	0.0444 *	0.0201 *	0.0646 *
	0.0163	0.0179	0.0233	0.0234	0.0113	0.0331
Share of rented agricultural area	-0.0136 ***	-0.0166 ***	-0.0104 ***	-0.0106 ***	-0.0049 ***	-0.0155 ***
	0.0010	0.0013	0.0018	0.0018	0.0013	0.0023
Price of building plots	0.2055 ***	0.2343 ***	0.1215 ***	0.1234 ***	0.0565 ***	0.1799 ***
	0.0153	0.0218	0.0326	0.0328	0.0156	0.0404
Ratio building vs. agricultural land	0.0502 ***	0.1689 ***	0.0957 ***	0.0973 ***	0.0457 **	0.1430 ***
	0.0153	0.0246	0.0296	0.0298	0.0180	0.0430
Spatial lag			0.3290 ***			

Spatial error			0.0710 0.2835 *** 0.0740
Adjusted R-squared	0.3191	0.2303	

***p<0,01, **p<0,05, *p<0,10; SE = Standard Error.

Source: Authors' calculations.

As discussed earlier, regression coefficients in a spatial lag model cannot be interpreted analogously to coefficients obtained from models without a spatial lag. The coefficient of, for example, 0.1061 for the "size of the transacted plot" in 2007 only covers the initial effect of a change in the plot size. However, an increase in the plot size and the subsequent increase in agricultural land prices in area *i* will also affect agricultural land prices in all neighbouring areas *j* through the spatial lag, and subsequently feeds back to the land price in area *i*. This feedback effect is included in what is usually defined as the "direct effect" in a spatial model (LeSage and Pace, 2009). Hence, this direct effect gives the average (over all regions) of the impact (including feedbacks) of changing one particular explanatory variable in one region. While this might be the appropriate measure to reveal the effect of soil quality rating or the size of the transacted plot, it is not appropriate for discussing the impact of a change in direct payments, since a policy change will usually occur in all regions. Hence, we have to add the effect of a change of direct payments in all neighbouring regions on area *i*. Those are covered by the indirect effect. Therefore, the total effect reports the total average effect of changing, for example, direct payments in all regions simultaneously on agricultural land prices in all regions.

All coefficient estimates in Tables 3 and 4 have the expected sign. Interestingly, the involvement of public authorities as buyer or seller increases the sales price quite substantially: the impact on the average land with a sales price of €21,749 in 2007 (€22,642 in 2001) is estimated to be 4,532 (5,376) €/ha if a public authority is a seller, and 6,633 (4,963) €/ha if it is a buyer. Plots with public authority involvement in the transaction are more likely to be located in more densely populated areas and land is eventually considered for prospective infrastructure development. Another possible explanation for this phenomenon could be a downward bias of official land prices when only private parties are involved as buyer and seller in order to reduce taxes.

Our analysis confirms the influence of agricultural factors such as land productivity, of variables describing the regional land market structure, and of non-agricultural factors such as urban pressure on agricultural land prices. As expected, the soil quality rating has a high positive impact on land sales prices, since it is directly connected to productivity. An increase in the soil quality rating by 1% will increase the sales price by 0.68% in 2007 (0.60% in 2001).[13] In other words, the difference in sales price between two otherwise identical plots with average sales prices, one having an average quality rating of approximately 45.5 in 2007 (45.2 in 2001) and the other ten points higher, is 3,259 (2,985) €/ha. Similarly, the difference between an average plot of 1.76 ha and a 3 ha plot, all other characteristics being equal and average, is 1,638 (2,563) €/ha. This positive influence of plot size can be explained by lower transaction costs in the transfer and lower operating costs. A 10 km increase in the distance to the next urban centre from an average of 29 (29) km to 39 (39) km decreases the price by 1,208 (184) €/ha. A 10 percentage point increase in the share of rented land from an average of 51% (44%) decreases the sales price by 2,884 (2,389) €/ha. The negative impact of a higher share of rented land on land sales prices is explained by the substitutive relation between renting and buying land. Moreover, land competes with other potential usages, in particular housing. Therefore, an increase of the sales price for land for construction from an average of 72 (83) €/m^2 to € 82 (93) €/m^2 increase the sales price of agricultural land by 152

[13] In accordance with our discussion above, we use the direct effects to discuss the impact for all determinants except direct payments.

(336) €/ha. A 1% increase in the ratio of building land to agricultural land increases the sales price of land by 0.1% (0.1%).

With regard to the influence of government support on land prices, a decrease in direct payments by 1% will decrease land prices by 0.27% in 2007, but only by 0.06% in 2001. To put this into numbers, for land at an average sales value and with average direct payments of 350 €/ha in 2007 and 261 €/ha in 2001, a decrease of direct payments by 50 €/ha will cause the sales price to drop by 849 €/ha and 280 €/ha, respectively. These numbers clearly indicate that the capitalisation of government payments into agricultural land prices increased between 2001 and 2007.

Signs, significance levels and, to a great extent, values are not markedly different for non-spatial estimates. Estimation results for the spatial model with a Gabriel neighbour weight matrix are reported in Tables A1 and A2 in the appendix for 2007 and 2001, respectively. Results do not differ much, though spatial effects are slightly stronger.

6. Conclusions

The purpose of this study is to investigate the main relationships determining agricultural land prices in Bavaria. We empirically analyse a dataset of more than 7,300 arm's length agricultural land sales transactions in 2001 and 2007.

Most of the preceding studies have ignored the spatial dimension of land use and land markets. Neglecting spatial relationships can lead to biased coefficient estimates for non-spatial explanatory variables. Combining a spatial lag model and a spatial error model (general spatial model) allows us to account for spatial dependence as well as spatial autocorrelation in the disturbances. Additionally, we consider the problem of endogeneity introduced by explanatory variables other than the spatially lagged dependent variable, a problem usually ignored in spatial econometrics (Fingleton and Le Gallo, 2008).

Earlier studies (e.g. Barnard et al., 1997; Duvivier et al., 2005) empirically confirmed the influence of government support payments on agricultural land prices. Goodwin et al. (2003) show that different forms of government support capitalise into agricultural land prices differently. According to our findings, a 1% reduction in EU direct payments would lead to a decrease in agricultural land prices of 0.27% in 2007 and 0.06% in 2001. Evaluated at mean levels, a 50 €/ha reduction in direct payments leads to reductions of 849 €/ha and 280 €/ha, respectively. We therefore find a significantly higher capitalisation of government support into land after the decoupling of direct payments under the Fischler Reform in 2004. Whether the reform or other occurrences are responsible for a larger capitalisation needs further inquiry. Apart from direct payments, we find a substantial influence of land productivity, urban pressure and regional land market structure on land prices. Our spatial results show that land prices increase by about 0.24% when land prices in surrounding areas increase by 1%.

In Bavaria, the share of agricultural land sold every year is relatively low. This might entail an unbalanced market structure with a small number of sellers on the one hand, and probably multiple potential buyers on the other. Accounting for this potentially imperfect competition and its implications for the determinants of agricultural land prices merits future investigation.

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Appendix

Table A1. Regression results for 2007 using spatial GMM with a Gabriel neighbour spatial weight matrix

		Spatial GMM			
		coeff.	direct	indirect	total
Constant	*coeff.*	3.9044 ***			
	SE	0.5803			
Public seller		0.2162 ***	0.2170 ***	0.0897 **	0.3067 ***
		0.0522	0.0529	0.0354	0.0804
Public buyer		0.2997 ***	0.3021 ***	0.1244 ***	0.4266 ***
		0.0265	0.0268	0.0376	0.0512
Soil quality rating		0.6720 ***	0.6767 ***	0.2775 ***	0.9541 ***
		0.0365	0.0365	0.0774	0.0799
Size of transacted plot		0.0989 **	0.0996 **	0.0408 *	0.1404 **
		0.0480	0.0478	0.0234	0.0684
Distance to the next urban centre		-0.1697 ***	-0.1699 ***	-0.0676 ***	-0.2375 ***
		0.0452	0.0455	0.0209	0.0581
Direct payments		0.2149 **	0.2154 **	0.0839 **	0.2992 ***
		0.0848	0.0851	0.0334	0.1108
Share of rented agricultural area		-0.0123 ***	-0.0124 ***	-0.0050 ***	-0.0174 ***
		0.0017	0.0017	0.0011	0.0017
Price of building plots		0.0355	0.0355	0.0136	0.0491
		0.0224	0.0228	0.0092	0.0310
Ratio building vs. agricultural land		0.0827 ***	0.0837 ***	0.0340 **	0.1177 ***
		0.0251	0.0251	0.0136	0.0356
Spatial lag		0.2904 ***			
		0.0610			
Spatial error		0.1683 **			
		0.0792			

***p<0,01, **p<0,05, *p<0,10; SE = Standard Error.
Source: Authors' calculations.

Table A2. Regression results for 2001 using spatial GMM with a Gabriel neighbour spatial weight matrix

		Spatial GMM			
		coeff.	direct	indirect	total
Constant	*coeff.*	3.4648 ***			
	SE	0.6487			
Public seller		0.2308 ***	0.2342 ***	0.1536 ***	0.3877 ***
		0.0430	0.0434	0.0541	0.0846
Public buyer		0.2188 ***	0.2220 ***	0.1452 ***	0.3673 ***
		0.0197	0.0199	0.0434	0.0502
Soil quality rating		0.5815 ***	0.5900 ***	0.3870 ***	0.9770 ***
		0.0329	0.0329	0.1161	0.1235
Size of transacted plot		0.1199 **	0.1221 **	0.0782 **	0.2002 **
		0.0490	0.0497	0.0372	0.0814
Distance to the next urban centre		-0.0581	-0.0595	-0.0392	-0.0986
		0.0449	0.0458	0.0337	0.0775
Direct payments		0.0420 *	0.0426 *	0.0264 *	0.0689 *
		0.0231	0.0236	0.0154	0.0374
Share of rented agricultural area		-0.0088 ***	-0.0090 ***	-0.0057 ***	-0.0147 ***
		0.0018	0.0018	0.0015	0.0026
Price of building plots		0.0963 ***	0.0977 ***	0.0602 ***	0.1579 ***
		0.0325	0.0327	0.0184	0.0456
Ratio building vs. agricultural land		0.0755 ***	0.0767 ***	0.0492 **	0.1259 ***
		0.0282	0.0288	0.0222	0.0473
Spatial lag		0.3975 ***			
		0.0718			
Spatial error		0.3045 ***			
		0.0768			

***p<0,01, **p<0,05, *p<0,10; SE = Standard Error.

Source: Authors' calculations.

Comparative Analysis of Factor Markets for Agriculture across the Member States

245123-FP7-KBBE-2009-3

The Factor Markets project in a nutshell

Title	Comparative Analysis of Factor Markets for Agriculture across the Member States
Funding scheme	Collaborative Project (CP) / Small or medium scale focused research project
Coordinator	CEPS, Prof. Johan F.M. Swinnen
Duration	01/09/2010 – 31/08/2013 (36 months)
Short description	Well functioning factor markets are a crucial condition for the competitiveness and growth of agriculture and for rural development. At the same time, the functioning of the factor markets themselves are influenced by changes in agriculture and the rural economy, and in EU policies. Member state regulations and institutions affecting land, labour, and capital markets may cause important heterogeneity in the factor markets, which may have important effects on the functioning of the factor markets and on the interactions between factor markets and EU policies. The general objective of the FACTOR MARKETS project is to analyse the functioning of factor markets for agriculture in the EU-27, including the Candidate Countries. The FACTOR MARKETS project will compare the different markets, their institutional framework and their impact on agricultural development and structural change, as well as their impact on rural economies, for the Member States, Candidate Countries and the EU as a whole. The FACTOR MARKETS project will focus on capital, labour and land markets. The results of this study will contribute to a better understanding of the fundamental economic factors affecting EU agriculture, thus allowing better targeting of policies to improve the competitiveness of the sector.
Contact e-mail	info@factormarkets.eu
Website	www.factormarkets.eu
Partners	17 (13 countries)
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