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The Fischler Reform of the Common Agricultural Policy and Agricultural Land Prices

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Abstract

Based on 7,300 agricultural land sales transactions we estimate the effect of the 2003 reform of the Common Agricultural Policy on land prices. As opposed to the main body of the literature on agricultural land values, we do not start from a demand-oriented net present value approach or hedonic prizing method, but derive our reduced form pricing equation from a spatial land sales market model. Our empirical model accounts for spatial dependence and endogeneity of explanatory variables. A reduction of payments by 50 €/ha would decrease land sales prices by 445 €/ha before and by 984 €/ha after the reform.

Keywords: agricultural land prices, Common Agricultural Policy, spatial econometrics

JEL: Q15, Q18, C21

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I. INTRODUCTION

Since Ricardo's (1817) work, a major argument against agricultural support policies has been that government interventions increase land rental and sales prices. Therefore, part of the economic rents created by policy to support active farmers' incomes are passed through to those who, for example, give up farming and rent out or sell their land. This clearly contradicts the stated objectives of agriculture policy in most developed countries. It might even worsen the situation of active farmers since costs for an important input factor are increasing.[1]

Different government programs will impact agricultural land values to different extents. This was first shown by Floyd (1965) in a simple model with one agricultural output, two production factors (land, labor and capital), and three policies (price support, price support with acreage control, and price support with a quota). Since then, Floyd's theoretical results have been re-examined in alternative ways or extended by relaxing some of his assumptions and/or including alternative policies (e.g. Hertel, 1989; Gardner, 1990; Debrew et al., 2001; Alston and James, 2002; Guyomard et al., 2004; Latruffe and Le Mouël, 2009).

Over the last 20 years, the Common Agricultural Policy (CAP) of the European Union (EU) went through two major changes. Through the MacSharry Reform in 1992 and the AGENDA 2000 Reform, dominant price support policy in the form of intervention prices was gradually replaced by direct payments, mostly coupled to land (e.g. arable area payments) and animal numbers (e.g. suckler cow premiums). In 2003, the subsequent Fischler Reform introduced decoupled payments in form of single farm payments (SFPs). Farmers were now able to receive SFPs by activating entitlements. The number of entitlements each farmer received at the starting point (between 2005 and 2007, depending on the country) was equal to the number of hectares that were farmed at the time of the introduction. Entitlement values were calculated on the basis of direct payments received, on a farm level (historical model), on a regional level (regional model), or on both (hybrid model), in the reference period of

2000 to 2002. To activate a certain number of entitlements, a farmer must at least manage (keep in a cultivatable condition), but not necessarily cultivate, the same number of eligible hectares. Hence, SFPs are regarded as decoupled from the direct production decisions. What a farmer plants, or if he or she plants anything at all, has no influence on the SFPs received. However, since land is necessary to activate entitlements, land values are not necessarily decoupled from SFPs.

Courleux et al. (2008), Ciaian et al. (2008), and Kilian et al. (2012) show, based on different theoretical models, that these decoupled payments still increase land prices. Therefore, part of the payments is capitalized into land values. The degree of capitalization crucially depends on the implemented model (historical, regional, hybrid) and the ratio between the number of entitlements and eligible hectares. Moreover, Kilian et al. (2012) have argued that under some circumstances, the degree of capitalization may have increased with the introduction of SFPs since former animal payments are now closer linked to land as they were before the reform. If this is true, the transfer efficiency, defined as the ratio of benefits to farmers and the costs of all other groups (Gardner, 1983), of the Fischler Reform is ambiguous. On the one hand, decoupling of payments from production decisions clearly decreases market distortions and implied deadweight losses (OECD, 2004). On the other hand, a bigger share of the support may now be captured by untargeted groups. Moreover, a high degree of capitalization clearly contradicts the objective of the CAP and, in particular, the objective of the most recent reform, which is to target "support exclusively to active farmers" (European Commission, 2010, p. 3). Against this background, a major aim of this paper is to compare the degree of capitalization of coupled direct payments before the 2003 Fischler Reform with the decoupled payments after the 2003 reform.

Our paper contributes to the literature on agricultural land sales prices in three ways: First, it is to our knowledge the first study to investigate the impact of the 2013 Reform on land sales prices by explicitly estimating the situation pre and post the reform. In addition to government payments, we investigate the influence of returns from land (Melichar; 1979; Alston, 1986), urban pressure (Capozza and Helsley, 1989; Cavaihles and Wavresky, 2003), and the regional land market structure (Cotteleer et al., 2008; Temesgen and Dupraz, 2014) on land prices.[2] Second, so far no one has applied a spatial autoregressive model with spatial autoregressive disturbances and additional endogenous variables on agricultural land sales prices. Third, almost all empirical contributions to agricultural land prices estimate a reduced form sales price equation. They justify this either by referring to the net present value method or the hedonic pricing approach (Feichtinger and Salhofer, 2013). The net present value approach calculates the maximum willingness to pay for a specific parcel of land as the discounted expected future stream of returns from this land including subsidies (Weersink et al., 1999). Therefore, the net present value approach depicts only the demand side of the market. Likewise, the hedonic pricing approach, anchored in consumer theory (Lancaster, 1966; Rosen, 1974), tries to reveal the willingness to pay for different characteristics (e.g. land quality, location) of a good (e.g. land). Our study provides an alternative justification for a reduced form sales price equation that is based on a spatial land sales market model with demand and supply by following Fingleton and Le Gallo (2008).

II. RELATED LITERATURE

Previous Studies on Capitalization

While several studies on the impact of agricultural policy on land price values exist for the U.S. (Goodwin and Ortalo-Magné, 1992; Barnard et al., 1997; Goodwin et al., 2003; Shaik et al., 2005; Taylor and Brester, 2005; Devadoss and Manchu, 2007) and Canada (Veeman et al., 1993; Weersink, et al., 1999; Carlberg, 2002), empirical evidence for the CAP of the EU and particularly for the impact of the decoupling of payments through the Fischler Reform is scarce, with only two studies published in peer-reviewed journals. Studies investigating the time before the Fischler Reform include Duvivier et al. (2005) and

Pyykkönen (2005). Duvivier et al. (2005) perform a panel data analysis based on average rental prices in 42 Belgian districts from 1980 - 2002. Depending on the year and region, they find elasticities of arable farmland prices to coupled area and animal payments ranging from about 0.1 to 0.5. [3] More in line with our study, Pyykkönen (2005) utilized a sample of more than 6,000 individual sales transactions of arable land in Finland between 1995 and 2002. He estimates capitalization elasticities ranging from 0.2 to 0.6.

More recent contributions evaluating the impacts of decoupled direct payments introduced in the Fischler Reform are Letort and Temesgen (2013), Nielsson and Johannson (2013), and Karlsson and Nielsson (2014). Letort and Temesgen (2013) concentrate on the role of environmental regulations on land prices and use about 4,000 observations of individual land sales transactions in Bretagne from 2007 to 2010. They include SFPs in their land sales equation and report a significant positive coefficient without further commenting on the magnitude of this effect. Based on their estimated coefficients and their descriptive statistics in Table 1, we calculate a capitalization elasticity of approximately 0.2.

Karlsson and Nielsson (2014) investigate the capitalization of SFPs on farm prices. Their study is based on a sample of approximately 3,400 individual farm sales transactions in Sweden between January 2007 and December 2008. It is important to note that they explicitly concentrate on farm sales rather than farmland sales by including only transactions that contain at least one residential unit. Their dependent variable is the total sales price, rather than price per hectare (ha), ranging from € 6,920 to € 2.9 million.[4] As one of the dependent variables, they use average SFPs per ha at a local sub-district level ranging from € 133 to € 384. Given the absolute nature of the right-hand side variable and the relative nature of the left-hand side, it is not very surprising that they are not able to find any significant influence of per ha payments on total farm value. Based on the same original data pool of individual transactions, but aggregating individual sales to average per hectare prices in 269 municipalities, Nilsson and Johansson (2013) find significant capitalization effects. Their average estimated elasticity of SFPs on sales prices is 0.54. Moreover, based on a quantile regression, they conclude that the capitalization effect is stronger for lower quality land.

Latruffe et al. (2013a) is, to our knowledge, the only paper to include coupled area and animal payments before the Fischler Reform and decoupled payments after the reform. They use simple OLS regression methods on more than 4,000 land transactions in three regions in France between 1994 and 2011. In regard to the capitalization effect of different types of payments, they obtain "rather puzzling estimation results […] when all types of subsidy are considered" (Latruffe et al., 2013a, p. 15). In all of their estimates, the impact of coupled animal and area payments on land prices before the Fischler Reform are either negative or insignificant. In regard to SFPs, they find a "significant positive capitalization impact only for plots located in a [nitrate] surplus zone" (Latruffe et al., 2013a, p. 15), i.e. livestock intensive areas.

Aside from the aforementioned papers on agricultural land sale prices, there is also a literature on the impact of government payments on land rental prices. Though closely related, the theoretical and empirical impact of subsidies on rental prices is different from sales prices. The effect of SPFs, or any other payments linked to land, on land rental prices is much more intuitive and direct. According to the OECD's (2014) Percentage Producer Support Estimate (%PSE), transfers have accounted for approximately 20% of total farm receipts in the EU period 2010 to 2013. If renting land grants this support, this obviously should have an impact on rental prices. However, in the case of land sales under policy uncertainty and an almost perpetual stream of returns from the productivity of land, the sum of discounted expected future payments should account for a much lower share of the total value of the asset. Nevertheless, the influence of SFPs on rental rates is not beyond dispute. While Kilian et al. (2012) and O'Neill and Hanrahan (2013) find clear evidence that a considerable share of the payments is capitalized into land rental prices; Michalek et al.

(2014) find much less evidence, and Moro et al. (2013) reject the hypothesis of a significant degree of capitalization of CAP payments for the time before and after the Fischler Reform.

Empirical Challenges

When estimating a land price model, there are two main empirical challenges: the spatial dimension of land and the potential endogeneity of explanatory variables. The spatial dimension of land leads to a limited spatial extension of farms and to regional land markets. Closer land markets interact with higher intensity than more distant ones, and they cause spatial dependency of the dependent variable. Moreover, unobserved spatial heterogeneity (e.g. in regard to weather or distance to the nearest market) may cause spatial dependency in the error term.

In general, endogeneity in econometric models may arise for three different reasons: omitted variables, measurement error, and simultaneity (Wooldridge, 2002, pp. 50-51). In particular, endogeneity in land price models, beside the possibility of omitted variables, may occur for at least three reasons. First, if a spatial lag model is used to account for the spatial dimension of the problem, endogeneity is automatically introduced since prices in one region are explained by simultaneously determined prices in neighboring regions. Second, other covariates may also not be exogenous given that the multifaceted interactions of demand and supply in land markets are described by a reduced form price equation. Third, land price models may be subject to a measurement error in form of the so called “expectation error” (Goodwin et al., 2003; Kirwan, 2009). Having incomplete foresight, buyers and sellers of agricultural land have to form some expectations about future market returns and government payments. Because farmers’ expectations cannot be observed, actually realized returns and payments are usually used in estimations. If expectations differ from realized values, we get biased estimates.

Neglecting endogeneity and/or spatial relationships can cause biased coefficient estimates. To account for endogeneity, Goodwin et al. (2010) utilized an instrumental variable approach on land sales prices. Kirwan (2009) did the same for rental prices. In solving the problem of spatially correlated error terms, Hardie et al. (2001), Patton and McErlean (2003), and Pyykonen (2005) apply spatial error models in their land sales price analyses. In a different approach to deal with spatial heterogeneity, Karlsson and Nielsson (2014) utilize a spatial multilevel model. To account for spatial dependency in the dependent variable, Huang et al. (2006) use a spatial lag model in their analysis of Illinois land sales prices. As an extension, Kostov (2009) suggests a quantile regression generalization of the (linear) spatial lag model. Maddison (2009) applies a spatio-temporal model where the right hand side variables include spatio-temporally lagged values of the dependent and independent variables. A spatio-temporal model starts from the assumption that farmland sale prices in region *i* are affected by a spatially weighted average of sale prices in neighboring regions in the past rather than in a simultaneous process. Therefore, there is no endogeneity problem introduced by the spatial weight matrix. Given the cross-sectional nature and the lack of information on the exact date of the transaction in our data, this approach is not applicable here.

Recently, Latruffe et al. (2013b) and Letort and Temesgen (2014) estimated a spatial lag model with spatial errors, without accounting for endogeneity of other covariates. Kelejian and Prucha (2010), Arraiz et al. (2010), and Drukker et al. (2013) have developed estimation procedures for spatial autoregressive models with spatial autoregressive disturbances and additional endogenous variables. Breustedt and Habermann (2011) utilized this estimation procedure for agricultural land rental prices in Lower Saxony (Germany). Similarly, we apply this procedure to a rather unique cross-sectional data set of nearly all land sales transactions in Bavaria in 2001 and 2007.

III. Theoretical Framework

Following Fingleton and Le Gallo's (2008) work, 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 area and the interaction with land markets in neighboring areas. Specifically, the quantity of agricultural land demanded in area *i* (Q_i) is modeled as a linear function

$$Q_i = \alpha_0 + \alpha_p P_i + \alpha_w \sum_{j \neq i}^{N} W_{ij}^{D} P_j + \sum_{k=1}^{K} \alpha_k A_{k,i} \qquad (1)$$

where P_i (P_j) is the price of agricultural land in area *i* (*j*) with *N* areas in total, W_{ij}^{D} is a $N \times N$ spatial weight matrix, $A_{k,}$ are *K* demand shifting variables, such as soil quality or distance to the nearest market, and all αs are parameters. 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 that area *j*. As a consequence, some demand will be displaced from area *j* to area *i*. Hence, Q_i is positively related to the weighted average of land prices in the surrounding areas ($W_{ij}^{D} P_j$) and $\alpha_w \geq 0$.

Analogously, the supply of agricultural land (Q_i) in area *i* can be modeled as

$$Q_i = \beta_0 + \beta_p P_i + \beta_w \sum_{j \neq i}^{N} W_{ij}^{S} P_j + \sum_{l=1}^{L} \beta_l B_{l,i} \qquad (2)$$

where W_{ij}^{S} is again a $N \times N$ spatial weight matrix, $B_{l,i}$ are *L* supply side shifters such as the share of rented land in a municipality,[5] and all βs are parameters. In accordance with standard economic theory, we assume $\beta_p \geq 0$. In contrast to the demand side spillover 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 *i* to nearby *j* ($\beta_w \leq 0$). In practice, one can think of a large farm whose owner prefers to sell a plot in a more expensive corner instead of another plot of equal quality in a cheaper corner.

Based on equations (1) and (2), and the assumption that $W^E = W^D = W^S$,[6] we can derive a reduced form pricing equation which can be written in matrix form as

$$P_i = \gamma + \rho \sum_{j \neq i} W_{ij}^E P_j + \sum_{m=1}^{M} \delta_m X_{m,i} \qquad (3)$$

where $X_{m,i}$ are $M = K + L$ variables of demand and supply shifters and γ, ρ and the M δ's are parameters, with $\gamma = \frac{\alpha_0 - \beta_0}{\alpha_p + \beta_p}$, $\rho = \frac{\alpha_w + \beta_w}{\alpha_p + \beta_p}$, $\delta_m = \frac{\alpha_k}{\alpha_p + \beta_p}$ for K demand shifters and $\delta_m = \frac{-\beta_l}{\alpha_p + \beta_p}$ for L supply shifters.

Rewriting equation (3) in a form that can be estimated by adding an error term, and taking into account that some right hand side (RHS) variables are endogenous, in matrix form we obtain the following:

$$\boldsymbol{P} = \gamma + \boldsymbol{X^e \eta} + \boldsymbol{X^d \mu} + \rho \boldsymbol{WP} + \boldsymbol{\varepsilon} \qquad (4)$$

where $\boldsymbol{P}$ is an $N \times 1$ vector of land sales prices, γ is a constant, $\boldsymbol{X^e}$ is an $N \times Q$ matrix of exogenous variables, $\boldsymbol{\eta}$ is the corresponding $Q \times 1$ vector of coefficients to be estimated, $\boldsymbol{X^d}$ is an $N \times R$ matrix of endogenous variables, $\boldsymbol{\mu}$ is the corresponding $R \times 1$ vector of coefficients to be estimated, $\boldsymbol{W}$ is a $N \times N$ spatial weight matrix, ρ is a spatial lag coefficient to be estimated, and $\boldsymbol{\varepsilon}$ is an error term.

Although equation (4) accounts for spatial dependency in the dependent variable, the potential problem of spatial autocorrelation in the disturbances remains. This may be caused by unobserved spatial heterogeneity, an inherent problem in land price analysis. To overcome this problem, spatial error processes are typically implemented into the 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 because all areas are connected to each other to some degree (global autocorrelation). In contrast, a shock is transmitted only to neighboring areas in the SMA model (local autocorrelation). Hence, the range of the effect is much smaller (Anselin, 2003). In the case of agricultural land markets, a shock in area *i* being transmitted to further distant units seems more plausible. Therefore, we choose the SAR model for our error term. Moreover, this is consistent with the (global) autoregressive process of our spatial lag formulation. The error term of equation (4) becomes

$$\boldsymbol{\varepsilon} = \lambda \boldsymbol{W} \boldsymbol{\varepsilon} + \boldsymbol{\upsilon} \qquad (5)$$

where λ the spatial error coefficient to be estimated. If we allow for heteroskedasticity, $\boldsymbol{\upsilon}$ is a $N \times 1$ vector of independently, but potentially heteroscedastic innovations (Drukker et al. 2011). While a spatial lag coefficient ρ has a direct interpretation, a SAR model is implemented to obtain unbiased estimates.[7] The combined spatial autoregressive model with spatial autoregressive disturbances is often referred to as a SARAR model (Anselin and Florax, 1995).

IV. EMPIRICAL ANALYSIS

Data Sources and Variable Selection

We utilize a comprehensive dataset of almost all arm’s length agricultural land sales transactions in Bavaria for the years 2001 (4,055 transactions) and 2007 (4,574), as collected by the Bavarian State Office for Taxes (Bayerisches Landesamt für Steuern). It 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 takeovers from 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).

We exclude from our dataset plots already legally converted for housing development, land with a special use, 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. Therefore, we do not consider transacted plots smaller than 0.25 ha. Such plots are more likely to inherit specific rights and easements (e.g. prospective non-agricultural land use) and this may result in a price premium 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.[8] 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.[9] Taking those restrictions into account, we are left with 7,369 observations for the years 2001 (3,539) and 2007 (3,830). On average, sales transactions took place in approximately 1,200 out of 2,056 Bavarian municipalities per year. The shape of Bavarian municipalities and the location of municipalities where transactions took place in 2001 are shown in Figure 1a and Figure 1b. Across both years, at least one transaction took place in 1,567 different municipalities and in 92 out of 96 different districts.

Descriptive statistics in Table 1 show that a plot of agricultural land sold on average for 22,642 €/ha (21,749 €/ha) in 2001 (2007). Public institutions, such as municipalities, are buyers in 22% (13%) of all transactions. Plots bought by the public are often dedicated to infrastructure development in the future or are handed over to a landowner as compensation for land dedicated to develop infrastructure. Public institutions act as sellers in 3.3% (2.5%) of the sales transactions. State and municipalities own agricultural land mostly for historical reasons. The share accounts for transactions of such land and for sales of plots left over from infrastructure development projects. The dataset does not allow us to distinguish between arable land and grassland, but we do have the soil quality index for each transacted plot available to account for differences in land quality. The soil quality index has an average value of 45.2 (45.5) and varies between 7.2 (7.5) and 84 (84). The average transacted plot has a relative small size of approximately 1.7 (1.8) ha. This variable helps to test if economies of scale of larger plots outweigh higher potential difficulties in financing to purchase them.

In addition to the information from our main data set on sales transactions, we add information at the municipality and district level. We use average direct payments in the respective municipality from the Integrated Administration and Control System (IACS) of the EU, provided by the Bavarian State Ministry for Food, Agriculture, and Forestry (Bayerisches Staatsministerium für Ernährung, Landwirtschaft und Forsten), to account for the fact that agricultural subsidies may capitalize into land values to some extent. The year 2001 represents the time before the Fischler Reform of the CAP and hence includes mainly coupled area and animal payments. The year 2007 represents the time after the Fischler Reform with decoupled single farm payments. On average, producers received 261 €/ha in 2001 and 350 €/ha in 2007 as direct payments. Low municipality averages, such as the minimum value of 7.36 €/ha in 2001, indicate a comparably high share of milk production on grassland, whereas high values, such as the maximum 707.74 €/ha in 2001, are a sign that arable farming in combination with intensive beef production are predominant.

We add additional covariates, all collected by Bavarian State Agency for Statistics and Data Processing (Bayersisches Landesamt für Statistik und Datenverarbeitung), to account for regional differences in urban pressure and market structure. In particular, we use the distance to the next urban center, the ratio of the sum of building land sold in the respective year and the preceding two years and the farmed agricultural land in the respective year, the sales prices for building plots, and the share of rented agricultural land in the total agricultural area. We expect agricultural land prices to increase in the vicinity of an urban center and in areas where building land is expensive. A high ratio between sold building land and farmed agricultural land indicates progressing urbanization and also a tight agricultural land market. Hence, we expect a positive relationship with agricultural land prices. Since renting land is a substitute to buying it, a higher share of rented agricultural area implies decreases in sales prices.[10]

Estimation Issues

To estimate the model of equations (4) and (5), we utilize a two-step estimation strategy as discussed in Kelejian and Prucha (1999; 2010), Arraiz et al. (2010), and Drukker et al. (2013) and as programmed in in the software package R by Piras (2010; 2013). Each of the two steps consists of alternating Generalized Method of Moments (GMM) and 2SLS estimators.

Spatial Weight Matrix

Specifying the spatial weight matrix $\boldsymbol{W}$ is always subjective to some extent. In particular, the researcher has to determine exogenously what defines neighbors as well as the weights given to each neighbor. In regard to the latter, common approaches are binary weights assigning a 1 to each neighbor and weights based on distance. While in the first approach all neighbors are weighted equally, geographically closer transactions are weighted stronger than more distant transactions in the second approach. We use binary weights since we lack information on the exact location of a transacted plot within a municipality. For the same reason, we assume municipality centroids to be the location of any transacted plot in a municipality.

To determine whether transactions are neighbors, we use two different approaches (Figures 2a and 2b).[11] In the first approach, a transacted plot is a neighbor (area *j*) of a transacted plot in question (area *i*) if the municipality centroid of area *j* (*J*) is within a circle of 8 kilometers from the centroid of area *i* (*I*). This is depicted in Figure 2a.[12] In some municipalities, multiple transactions take place in one year. Because those transactions are clearly within a circle of 8 kilometers, they are also considered neighbors. Though not necessarily closer in distance to the transaction in question, they are intuitively closely connected because the flow of information is most likely highest within a municipality. In the second approach, illustrated in Figure 2b and called a Gabriel graph, closed discs are drawn between municipality centroids. Areas *i* and *j* are considered neighbors if the closed disc between their centroids (*I* and *J*) contains no other centroids.[13] None of the two definitions implies that *K* is a neighbor of *I*. While in the first case this is due to *K* being outside of an 8 kilometer circle, a closed disc between *I* and *K* containing *J* is the reason in the second case. When using a distance-based neighbor definition, approximately 20 transactions per year have to be dropped from our sample due to a lack of neighbors. The reasons for this are generally low numbers of sales transactions in the whole region or only a single transaction in a large municipality, with the next municipalities' centroids being further away than 8 kilometers. Advantageously, no transactions have to be dropped when the second approach is used because every area *i* has at least one neighbor area *j* per definition. In the distance-based approach, the average number of neighbors for each observation was 15.3 in 2001 and 16.1 in 2007. In the Gabriel based approach, it is 18.5 and 21.1, respectively.

Based on these two approaches to define neighbors, we derive two different row-standardized weight matrices with every row summing to one, independent of the actual number of neighbors. This implies a decreasing impact of the single transaction with a rising number of neighbors. Moreover, a row-standardized matrix is not symmetric, and a transaction in area *j* may influence a transaction in area *i* differently than in the reverse case.

Most importantly, a row-standardized form allows us to interpret the coefficient as the weighted average effect of land prices in the surrounding areas on land prices in area *i*.

Instrumental Variables

The main challenges in conducting instrument variable estimates are identifying endogenous variables and finding appropriate instruments. Given the reduced form formulation of our model, most shift variables may suffer a simultaneity problem. For example, a high share of rented agricultural area indicates a relatively large rental market as an alternative to buying land. This will negatively influence the sales price. However, a low sales price will also influence the quantity of land rented out, since buying land, as an alternative to renting it, becomes more attractive. Similar reasoning can be made for most other shift variables. Therefore, we apply different statistical tests for endogeneity. First, we use a Durbin-Wu-Hausman test to determine whether a subset of the endogenous variables is actually exogenous by running a secondary estimation where the test variables are treated as exogenous and by comparing the *J*-statistic of both estimations.[14] Second, we perform a regression based test as discussed in Wooldridge (2002, pp. 119). In the first stage of this test, a potentially endogenous explanatory variable is regressed on all exogenous variables and all instruments. Subsequently, residuals obtained from the first-stage regressions are included in land price regressions in the second stage. If and only if a residual vector added has no influence on land prices in the second stage estimations, the variable of interest is exogenous. This is commonly tested using a standard *t*-test, accounting for heteroscedasticity if necessary.

To test for instrument weakness, we evaluate the R^2 of the OLS estimates of the first stage of Two-Stage-Least-Squares (2SLS) instrumental variable regressions and the Cragg-Donald (1993) statistic, as proposed by Stock and Yogo (2005).[15] Based on all these tests, we can clearly reject endogeneity only for the soil quality index. This makes sense given that soil quality is defined by natural conditions that are completely exogenous to our system. In addition, we are not able to find acceptable instruments for the public seller and public buyer variables. Therefore, we have to assume those variables are exogenous, because using weak instruments can lead to biased inferences in instrumental variable estimations. Hence, our vector of instruments $\boldsymbol{Z}$, which is replacing $\boldsymbol{X^d}$ in estimating equation (4), includes two-year lags of direct payments, the share of rented agricultural area, the ratio of building vs. agricultural land, a one year lag of the price of building plots and municipality averages of the livestock units per hectare, the size of agricultural land parcels, and the standard gross margin per farm. To instrument the spatially lagged dependent variable, we follow Bivand and Piras (2015) and therefore apply the following matrix of instruments: $\boldsymbol{H} = (\boldsymbol{X^e}, \boldsymbol{Z}, \boldsymbol{WX^e}, \boldsymbol{WZ}, \boldsymbol{W^2X^e}, \boldsymbol{W^2Z})$.

Functional Form

We test different functional forms: linear, double-log, semi-log, and mixed-log. The mixed-log is between the double-log and the semi-log with the left hand side (LHS) variable in logs and the RHS variables in logs or absolute values, depending on which variable distribution is closer to a normal distribution. Since the models are not nested in each other, we apply information criteria (Akaike = AIC; Bayesian = BIC) and the Ramsey (1969) regression specification error test (RESET) (Wooldridge, 2003, pp. 292-294). The RESET first estimates an original model (e.g. double-log) and from that derives a fitted value of the LHS variable ($\hat{P}$). In a second stage, the same model, but including polynomials of the fitted values, in our case $\hat{P}^2$ and $\hat{P}^3$, is estimated. If the original model is correctly specified, coefficients of $\hat{P}^2$ and $\hat{P}^3$ should not be significantly different from 0, as tested by a common *F*-Test. To be able to perform these tests, we are restricted to OLS estimates of the spatial lag model. Table 2 represents the results. Based on the information criteria, the semi-log model fits the 2001 data best and the mixed-log model the 2007 data. AIC and BIC values are similar for double-log, semi-log, and mixed-log, but they are very different for the linear model. According to the RESET test, the linear model is clearly rejected for both years. The double log-model cannot be rejected at the 1% level, but it can be rejected at the 10% (5%) only for 2001 (2007). The semi-log and the mixed-log cannot be rejected. Given these results, we chose to continue with the mixed-log model, but final impacts will also be represented for the double-log and semi-log in order to have some indication of how sensitive our results are to different functional forms. We will no longer pursue the linear specification since it is clearly inferior in regard to performance and seems misspecified.

Spatial Model

Although we give some theoretical justification for a spatial lag model in section III, we also statistically test for spatial autocorrelation in general utilizing a Moran's I test and for spatial autoregressive processes in the dependent variable, as well as the residuals utilizing Lagrange Multiplier (LM) tests. In the Moran's I tests, positive (negative) values indicate positive (negative) spatial autocorrelation, and values close to zero indicate no autocorrelation. According to Table 3, the H_0 of no spatial autocorrelation is rejected at the 99% level for all specifications.[16] 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: $\lambda = 0$), while Anselin (1988) proposed a LM test for spatial autoregressive processes in the dependent variable (H_0: $\rho = 0$). LM test results confirm spatial autoregressive processes in the residuals as well as the dependent variable. In such a case, the robust test versions have to be applied (Anselin et al., 1996).[17] Robust test version results again confirm spatial autoregressive processes in the residuals as well as the dependent variable for all specifications. Hence, Moran's I and LM tests confirm (on empirical grounds) the use of a general spatial model of equation (4), including a decomposed error term as in equation (5).

Results

Estimation results for the mixed-log model with distance-based spatial weight matrices are reported in Table 4 for 2007 and in Table 5 for 2001. Results for the Gabriel weight matrices are in the appendix in Table A1 and Table A2. Here we concentrate on the interpretation of the heteroscedasticity-consistent spatial 2SLS/GMM estimator, although we also report non-spatial White heteroscedasticity-consistent OLS and GMM estimates for comparison. A spatial lag coefficient of 0.21 (0.31) in 2007 (2001) indicates that agricultural land sales prices in area *i* increase by approximately 0.21% (0.31%) when sales prices in surrounding areas increase by 1%. The significant spatial autocorrelation coefficient of 0.26 (0.32) confirms our SARAR model. In addition, all other model coefficient estimates are highly significant and have the expected signs, except for the distance to the next urban center in 2001.

It is important to note that coefficient estimates in a spatial lag model cannot be interpreted analogously to those obtained from models without a spatial lag. For example, a coefficient of 0.1108 for the variable log(size of a transacted plot) in 2007 only covers the initial effect of a change in the plot size. However, an increase in the plot size and a subsequent increase in agricultural land prices in area *i* will, in turn, spillover to all neighboring areas *j* through the spatial lag parameter and affect agricultural land prices in *j*.[18] Increased prices in area *j* cause a feedback effect, though smaller in size, 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, a direct effect gives the average impact over all regions (including feedbacks) of changing a particular explanatory variable in one area. While this might be the appropriate measure to reveal the effect of the soil quality index or the size of the transacted plot on land prices, it is so it is probably not to capture the impact of government support payments on land prices, because an altered support regime causes changes of direct payments in many (or most likely all) regions at the same time. Hence, we add the effect of changing direct payments in all neighboring areas *j* on area *i*, which is called the indirect effect. Total effects, obtained by summing direct and indirect effects, essentially report the total average effect of changing direct payments in all regions simultaneously on agricultural land prices.

Comparing the estimates of our spatial model to those obtained from non-spatial OLS and GMM regressions shows that signs and significance levels are not markedly different, while coefficient values differ to some extent. Results for the semi-log and the double-log model are in the same ranges. Comparing the results for a distance-based weight matrix (Table 4 and Table 5) with those based on a Gabriel weight matrix (Appendix Tables A1 and A2) reveals slightly stronger spatial effects with other coefficients being relatively comparable.

Table 6 reports the effects of changes of our determinants on land sales prices for all estimated models and a distance-based weight matrix. We discuss the results for the mixed-log model and provide the results of the double-log and semi-log model as a sensitivity analysis. Very interestingly, involvement of a public authority, either as a buyer or a seller of a plot, increases sales prices quite substantially. The impact at the median sales price of € 18,525 in 2007 (€ 19,476 in 2001) is estimated to be 5,705 (4,292) €/ha if a public buyer is involved and 3,873 (4,731) €/ha if a public seller is involved.[19] Plots with public authorities involved in the transaction are probably more likely located in more densely populated areas. Moreover, public authorities often buy agricultural land for infrastructure development. Another possible explanation for this phenomenon could be a downward bias of officially stated land prices when only private parties are involved, in order to avoid taxes.

With regard to the influence of government support on land prices, we find that for land with a median sales price and median direct payments of 354 €/ha in 2007 and 282 €/ha in 2001, a decrease of direct payments by e.g. 50 €/ha will cause the sales price to drop by 984 €/ha and 444 €/ha, respectively. These numbers clearly indicate an increased degree of capitalization of government support payments into agricultural land prices between 2001 and 2007.

Furthermore, 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 index has a positive impact on land sales prices because it is a relatively direct measure of productivity. The difference in sales prices between a median plot and one with a soil quality index10 points higher, all other characteristics equal, is 3,045 (2,782) €/ha in 2007 (2001). Analogously, a plot that is 1 hectare larger than the median plot costs is 2,063 (2,782) €/ha more. A positive influence of plot size makes sense due to lower transaction costs in the transfer and lower operating costs thereafter.

Agricultural land sales prices clearly increase with increased urban pressure. This is confirmed by the coefficients of all three variables: distance to the next urban center, ratio between sold building land and farmed agricultural land, and the price of building plots. First, an increase in the distance to the next urban center from a median distance of 28.2 km by 10 km to 38.2 km decreases the price by 1,338 €/ha in 2007. The impact in 2001 is slightly positive (148 €/ha) based on an insignificant coefficient estimate. Second, doubling the ratio between sold building land and farmed agricultural land from a median value of 7.9 (14) increases the sales price of land by 3,920 (2,180) €/ha. This positive relation can be justified in the following way: a high numerator indicates a high demand for building land, putting pressure on agricultural land prices. Moreover, a high number of sold building parcels usually increases farm income and increases farmers' willingness to pay for agricultural land as reinvestment and to save on income tax. A low denominator indicates a potentially thin agricultural land market, implying a higher price per hectare. Third, agricultural land use competes with other potential usages, in particular housing. Therefore, an increase of the sales price for building land from a median price of 56 (63) €/m^2 to 106 (113) €/m^2 increases the sales price of agricultural land by 2,541 (2,160) €/ha.

Finally, an increase in the share of rented land from a median value of 50% (43%) by 10% points decreases the sales price by 2,541 (2,081) €/ha. A large rental share indicates a busy rental market and increases farmers' potential to acquire land through the rental market as a substitute to buying land.

V. CONCLUSIONS AND DISCUSSION

The study at hand is the first to directly compare the effects of coupled government payments before the 2003 Fischler Reform with the effects from decoupled single farm payments (SFP) after the reform. We find significant differences in the degree of capitalization of payments into land prices. While the effect of a decrease in payments by 50 €/ha is estimated to decrease land prices between 227 €/ha and 445 € ha in 2001, the same reduction in payments would cause land price reduction by a range of 723 €/ha to 1,397 €/ha in 2007. To put it differently, the capitalization elasticity, defined as the percentage change in sales prices given a 1% change in government payments, increased from somewhere between 0.07 and 0.09 in 2001 to somewhere between 0.20% and 0.28% in 2007.

This finding is very much in line with theoretical considerations by Courleux et al. (2008), Ciaian et al. (2008), and Kilian et al. (2012), who argue that single farm payments, though decoupled from production decisions, are by no means decoupled from land values because land is the crucial and limited factor to receive SFP. For land rental markets, Kilian et al. (2012) and Feichtinger et al. (2014) empirically confirm that the Fischler Reform increases the capitalization effect.

The degree of capitalization increasing from the 2003 reform is problematic for two reasons. First, it contradicts the objectives of the EU Common Agricultural Policy (CAP), particularly the objective of the most recent reforms, to target "support exclusively to active farmers" (European Commission, 2010, p. 3). Second, whether the reform increased the transfer efficiency, defined as the ratio between benefits of the targeted group and costs to all other groups (Josling, 1974; Gardner, 1983; Bullock and Salhofer, 2003), of the CAP remains ambiguous. On the one hand, decoupled payments are clearly less distortionary than coupled payments (OECD 2004). On the other hand, the capitalization effect causes some leakage of transfers to unintended groups (OECD, 1995; Salhofer and Schmid, 2004). Hence, whether overall transfer efficiency has improved remains questionable.

At first sight, the CAP Reform 2014-2020 includes some major changes. Decoupled (former single farm) payments have been divided into basic payments and some additional payments, including green direct payments, redistributive payments, payments for areas with natural or other specific constraints, and payments for young farmers. To receive basic payments farmers will still need entitlements and the same number of eligible hectares. Green payments account for 30% of all direct payments and are paid on the condition that farmers undertake practices that are beneficial to the climate and to the environment. Other additional payments are either linked to farm and/or farmer characteristics. However, for all these additional payments, receiving basic payments is a precondition. Hence, also under the new scheme, payments are linked as closely to land as before the reform and will be capitalized to a similar amount as SFPs.

Courleux et al. (2008), Ciaian et al. (2008), and Kilian et al. (2012) all argue that the ratio between entitlements and eligible hectares is one of the crucial factors in determining the degree of capitalization. If the number of allocated entitlements is considerably smaller than the number of eligible hectares in a country, competition for land, necessary for activating entitlements, would decrease. While the exact ratio between allocated entitlements and eligible area is unknown, Ciaian et al. (2014) show that at least for half of the old member states including Germany, the ratio between activated entitlements and utilized agricultural area (UAA) is close to 1. Though there might be differences between allocated and activated entitlements and between eligible area and UAA, this is an indication of strong competition for land, which is necessary for activating entitlements. Consequently, decreasing the number of entitlements could decrease the capitalization effect. One example in this regard might be Ireland. Given the short-term nature of rental contracts in Ireland, as part of the Fischler Reform, farmers were allowed to consolidate entitlements where rental contracts have expired to other rented or owned land. Hence, the value of up to two entitlements can now be activated with one hectare of eligible area. This considerably changes the ratio between entitlements and eligible area and might explain why O'Neil and Hanrahan (20013), in their rental price study, found the degree of capitalization to decrease with the Fischler Reform.

Apart from that, we find a substantial influence of land productivity, the regional land market structure, and urban pressure on land prices. In contrast to previous studies of land sales prices, we account for the spatial dimension of land markets and for the endogeneity of explanatory variables. Each of these issues can potentially lead to biased estimates. In regard to spatial dependency, we show that land prices within a region are significantly influenced by prices in neighboring regions. Hence, not taking this into account may cause biased estimates for the coefficients of all determinants.

Based on LfStat (2008; 2013) we find that approximately 0.20% of total Bavarian agricultural land was sold in 2007. This number does not change considerably over the years. Hence, in general, the share of agricultural land sold each year is relatively low. This might entail an unbalanced market structure with a small number of sellers and most likely multiple potential buyers. Accounting for this potential imperfect competition, and its implications on the determinants of agricultural land prices, would be worth further investigation in the future.

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Footnotes

[1] For example, farm expenditures for land rentals in Germany add up to € 2.434 billion in 2013. This corresponds to approximately 45 % of all direct payments that German farmers receive from the EU under pillar 1 of the CAP, or 39 % of the agricultural sector's total net added value, defined as the production value (not including subsidies) minus input costs (not including rents) minus depreciation (BMELV, 2014).

[2] Feichtinger and Salhofer (2013) provide a review of the variables used in previous agricultural land price studies. Other related literature reviews on agricultural land prices are Oltmer et al. (2001), Le Mouël (2003) and Latruffe, and Le Mouël (2009).

[3] A capitalization elasticity of 0.2 means that a 1% increase in government payments increases land prices by 0.2%.

[4] Values in Euros are calculated using their results in Swedish kronor and the exchange rate given in their footnote 2.

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

[6] This assumption implies that agricultural land demand and supply in region i is influenced by the exact same neighboring regions j. Areas which are too far away to compete in demand are also too far away to compete in supply.

[7] LeSage (1999) and LeSage and Pace (2009) provide extensive reviews of different spatial models.

[8] Before excluding outliers, the average sales price was 25,289 €/ha, with a standard deviation of 28,345 €/ha including both years of observation. After accounting for outliers, our average sales price drops to 22,178 €/ha, with a standard deviation of 14,223 €/ha.

[9] In Germany, an index system is used to indicate the soil quality of agricultural land. This index ranges from zero to 100, with values for Bavaria between 7 and 85 (LfL, 2007).

[10] Mean and standard deviations of variables based on municipality and district averages are sample weighted because the 7,369 transactions are unequally distributed across municipalities.

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

[12] Choosing a radius of 8 kilometers is to some extent random. It is driven by considerations about farmers' knowledge about and interests in the land market in their vicinity. From a technical point of view, if the chosen radius is too short, many observations have no neighbor at all and have to be excluded from the analysis. If the chosen radius is too long, each observation receives a large number of neighbors.

[13] For an application of the Gabriel graph (first discussed in Gabriel and Sokal, 1969), we refer to Bivand and Brunstad (2006).

[14] All endogeneity tests are conducted with Eviews.

[15] The Cragg-Donald statistic is only valid for 2SLS and other K-class estimators. However, our results of the 2SLS and the GMM estimations are very similar in all respects.

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

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

[18] Please note, because we assume a spatial autoregressive model, the shock spreads further.

[19] In accordance with our discussion above, we use the direct effects to simulate the impact for all determinants except for direct payments, where we use the total effect.

TABLE 1

Descriptive Statistics

		No. obs	Mean/share	Median	Std. Dev.	Min	Max
2001							
Sales price	€/ha	3,539	22,642.32	19,476.49	14,332.16	2,044.20	102,260.10
Public buyer	%	3,539	21.87				
Public seller	%	3,539	3.33				
Soil quality index	Pt.	3,539	45.19	44.01	13.07	7.18	84.00
Size of a transacted plot	Ha	3,539	1.67	1.07	2.26	0.25	73.44
Direct payments	€/ha	1,211	261.28	282.03	92.21	7.36	469.03
Distance to the next urban center	Km	1,211	29.01	28.19	14.14	1.00	80.61
Ratio building vs. agricultural land		82	9.43	7.85	11.12	2.11	198.24
Price of building plots	€/m^2	82	83.09	63.15	66.13	19.21	727.84
Share of rented agricultural area	%	82	44.25	42.56	10.47	12.75	77.66
2007							
Sales price	€/ha	3,830	21,749.12	18,524.79	14,109.23	2,026.75	102,300.00
Public buyer	%	3,830	12.74				
Public seller	%	3,830	2.45				
Soil quality index	Pt.	3,830	45.50	44.91	12.67	7.47	84.00
Size of a transacted plot	Ha	3,830	1.76	1.13	1.94	0.25	31.76
Direct payments	€/ha	1,196	350.31	354.41	53.23	122.03	707.74
Distance to the next urban center	Km	1,196	29.00	28.14	14.62	1.00	72.49
Ratio building vs. agricultural land		86	18.15	14.02	20.92	2.58	252.84
Price of building plots	€/m^2	86	71.74	55.99	50.01	16.07	331.17
Share of rented agricultural area	%	86	51.38	49.62	9.96	19.26	78.17

TABLE 2

RESET and Information Criteria for Different Functional Forms

		Linear	Double-log	Semi-log	Mixed-log
2001					
RESET	*F-statistik*	18.416	2.349	0.448	1.656
	P-value	0.000	0.096	0.639	0.191
AIC		75,722.9	4,465.7	4,452.9	4,464.2
BIC		75,797.0	4,539.7	4,526.9	4,538.3
2007					
RESET	*F-statistik*	19.459	3.450	2.298	0.859
	P-value	0.000	0.032	0.101	0.424
AIC		81,872.0	4,940.3	4,929.7	4,917.8
BIC		81,947.0	5,015.3	5,004.7	4,992.8

TABLE 3

Spatial Autocorrelation Tests

	2001				2007			
Weight matrix	Distance based		Gabriel		Distance based		Gabriel	
Average no. of neighbors	15.32		18.50		16.06		21.07	
Moran's I test	0.271	***	0.252	***	0.186	***	0.156	***
LM error	1,220.31	***	1,638.58	***	662.30	***	735.45	***
Robust LM error	155.53	***	319.16	***	71.48	***	144.35	***
LM lag	1,086.92	***	1,351.18	***	644.26	***	650.01	***
Robust LM lag	22.15	***	31.75	***	53.43	***	58.91	***

***p<0.01, **p<0.05, *p<0.10.

TABLE 4

Regression Results for 2007 Using OLS, GMM and Spatial 2SLS/GMM for the Mixed-log Model with a Distance-based Weight Matrix

		OLS	**GMM**	**Spatial 2SLS/GMM**			
		coeff.	coeff.	coeff.	direct	indirect	total
Constant	coeff.	8.5603 ***	8.7326 ***	6.9444 ***			
	SE	0.1031	0.1255	0.5879			
Public buyer		0.3406 ***	0.3435 ***	0.3066 ***	0.3078 ***	0.0796 ***	0.3874 ***
		0.0253	0.0271	0.0263	0.0266	0.0296	0.0426
Public seller		0.2232 ***	0.2258 ***	0.2081 ***	0.2095 ***	0.0543 **	0.2637 ***
		0.0479	0.0533	0.0500	0.0500	0.0243	0.0662
Direct payments		0.0011 ***	0.0015 ***	0.0008 ***	0.0008 ***	0.0002 ***	0.0011 ***
		0.0002	0.0002	0.0003	0.0003	0.0001	0.0003
Soil quality index		0.0187 ***	0.0176 ***	0.0164 ***	0.0164 ***	0.0042 ***	0.0207 ***
		0.0007	0.0007	0.0008	0.0008	0.0015	0.0016
Log(size of a transacted plot)		0.0270 ***	0.0945 ***	0.1108 **	0.1114 **	0.0287 *	0.1402 **
		0.0090	0.0502	0.0506	0.0509	0.0174	0.0648
Distance to the next urban center		-0.0031 ***	-0.0094 *	-0.0072 ***	-0.0072 ***	-0.0018 **	-0.0090 ***
		0.0006	0.0016	0.0023	0.0023	0.0007	0.0027
Log(ratio building vs. agricultural land)		0.1886 ***	0.1934 ***	0.1627 ***	0.1638 ***	0.0412 ***	0.2050 ***
		0.0142	0.0221	0.0291	0.0291	0.0138	0.0326
Log(price of building plots)		0.0981 ***	0.0837 ***	0.0496 ***	0.0503 **	0.0125 *	0.0628 **
		0.0153	0.0195	0.0252	0.0253	0.0075	0.0311
Share of rented agricultural area		-0.0171 ***	-0.0175 ***	-0.0137 ***	-0.0138 ***	-0.0035 ***	-0.0172 ***
		0.0009	0.0011	0.0017	0.0017	0.0011	0.0017
Spatial lag				0.2063 ***			
				0.0614			
Spatial error				0.2628 ***			
				0.0738			
Adjusted R-squared		0.4114	0.3832				

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

TABLE 5

Regression Results for 2001 Using OLS, GMM and Spatial 2SLS/GMM for the Mixed-log Model with a Distance-based Weight Matrix

		OLS	GMM	Spatial 2SLS/GMM			
		coeff.	coeff.	coeff.	direct	indirect	total
Constant	coeff.	8.5346 ***	8.2163 ***	5.6070 ***			
	SE	0.0943	0.1557	0.6458			
Public buyer		0.2651 ***	0.2770 ***	0.2177 ***	0.2208 ***	0.0993 ***	0.3201 ***
		0.0208	0.0216	0.0197	0.0199	0.0337	0.0408
Public seller		0.2761 ***	0.2946 ***	0.2399 ***	0.2434 ***	0.1097 ***	0.3531 ***
		0.0477	0.0512	0.0429	0.0432	0.0419	0.0717
Direct payments		0.0005 ***	0.0008 ***	0.0003 *	0.0003 *	0.0001 *	0.0005 **
		0.0001	0.0001	0.0002	0.0002	0.0001	0.0002
Soil quality index		0.0153 ***	0.0144 ***	0.0141 ***	0.0143 ***	0.0065 ***	0.0208 ***
		0.0007	0.0007	0.0008	0.0007	0.0022	0.0023
Log(size of a		0.0279 ***	0.1683 ***	0.1314 **	0.1330 **	0.0597 *	0.1927 **
transacted plot)		0.0101	0.0488	0.0537	0.0539	0.0321	0.0804
Distance to the next		-0.0019 ***	0.0033 *	0.0008	0.0008	0.0003	0.0011
urban center		0.0006	0.0019	0.0023	0.0023	0.0011	0.0034
Log(ratio building		0.0813 ***	0.1700 ***	0.1106 ***	0.1126 ***	0.0494 ***	0.1619 ***
vs. agricultural land)		0.0147	0.0206	0.0270	0.0271	0.0172	0.0368
Log(price of		0.2222 ***	0.2386 ***	0.1383 ***	0.1405 ***	0.0599 ***	0.2004 ***
building plots)		0.0154	0.0210	0.0349	0.0351	0.0157	0.0400
Share of rented		-0.0140 ***	-0.0169 ***	-0.0106 ***	-0.0107 ***	-0.0047 ***	-0.0154 ***
agricultural area		0.0010	0.0013	0.0020	0.0019	0.0013	0.0023
Spatial lag				0.3141 ***			
				0.0749			
Spatial error				0.3192 ***			
				0.0761			
Adjusted R-squared		0.3172	0.2580				

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

TABLE 6

Effects of Changes of Determinants on Land Sales Prices for All Estimated Models and a Distance-based Weight Matrix

	Mixed-log	Semi-log	Double-log
2007			
Public buyer (yes)	5,705.37	5,627.31	5,638.09
Public seller (yes)	3,873.23	3,597.54	3,863.07
Direct payments (+50 €/ha)	984.22	1,396.95	722.84
Soil quality index (+10 points)	3,044.78	3,009.91	2,810.78
Size of transacted plot (doubled median)	2,062.64	993.36	1,747.57
Distance to next urban center (+10 km)	-1,338.02	-1,749.63	-1,064.71
Ratio building vs. agricultural land (doubled median)	3,920.17	504.61	2,416.36
Price of building plots (+50€/m2)	824.38	1,118.88	831.40
Share of rented agricultural area (+10 percentage points)	-2,540.88	-2,372.24	-2,462.05
2001			
Public buyer (yes)	4,292.42	4,101.82	4,257.92
Public seller (yes)	4,730.76	4,620.78	4,623.13
Direct payments (+50 €/ha)	444.95	418.46	226.84
Soil quality index (+10 points)	2,782.13	2,942.52	2,633.61
Size of transacted plot (+ 1 ha)	2,591.75	256.11	2,580.42
Distance to next urban center (+10 km)	154.13	199.73	-165.49
Ratio building vs. agricultural land (doubled median)	2,180.44	810.85	1,889.73
Price of building plots (+50€/m2)	2,159.67	817.97	1,899.67
Share of rented agricultural area (+10 percentage points)	-2,081.72	-2,011.17	-2,049.81

FIGURE 1

Bavaria with its Municipalities (a) and Municipality Centroids where Transactions Took Place in 2001 (b)

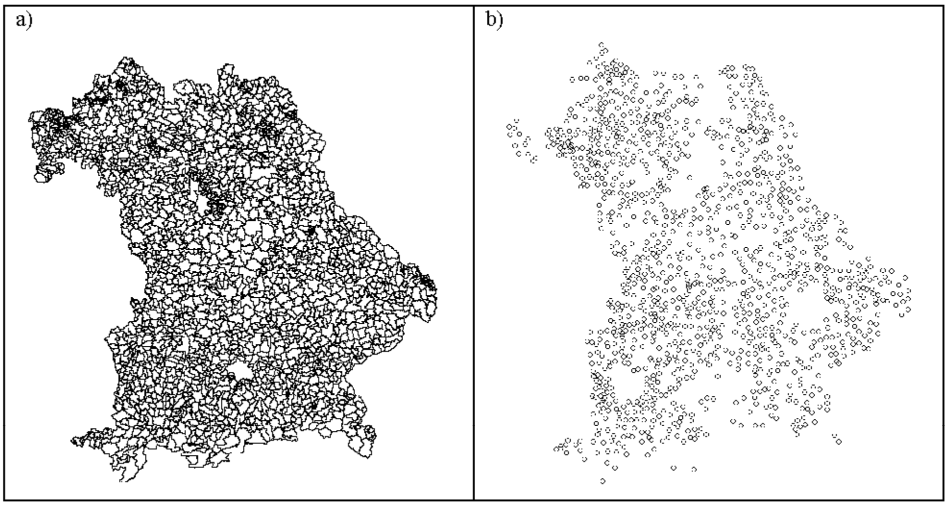
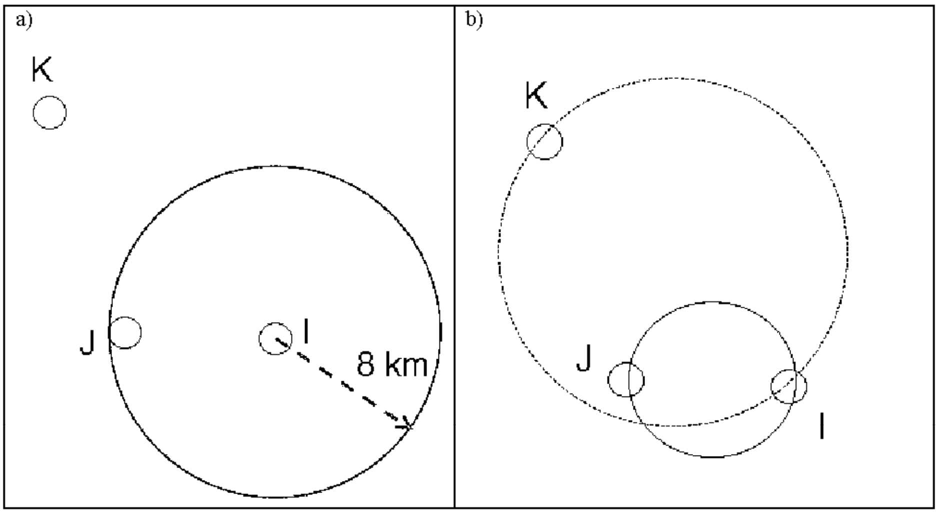


FIGURE 2

Distance-based (a) and Gabriel (b) Neighbor Definition



APPENDIX

TABLE A1

Regression Results for 2007 Using Spatial 2SLS/GMM for the Mixed-log Model with a Gabriel Weight Matrix

		Spatial 2SLS/GMM			
		coeff.	direct	indirect	total
Constant	coeff.	6.7292 ***			
	SE	0.6362			
Public buyer		0.3054 ***	0.3064 ***	0.0953 ***	0.4017 ***
		0.0266	0.0265	0.0350	0.0478
Public seller		0.2183 ***	0.2193 ***	0.0684 **	0.2877 ***
		0.0493	0.0496	0.0298	0.0702
Direct payments		0.0009 ***	0.0009 ***	0.0003 ***	0.0012 ***
		0.0003	0.0003	0.0001	0.0003
Soil quality index		0.0163 ***	0.0164 ***	0.0051 ***	0.0214 ***
		0.0008	0.0008	0.0018	0.0018
Log(size of a transacted plot)		0.1122 **	0.1128 **	0.0351 *	0.1479 **
		0.0513	0.0516	0.0211	0.0689
Distance to the next urban center		-0.0082 ***	-0.0082 ***	-0.0024 ***	-0.0106 ***
		0.0023	0.0023	0.0009	0.0027
Log(ratio building vs. agricultural land)		0.1520 ***	0.1521 ***	0.0461 ***	0.1982 ***
		0.0279	0.0280	0.0153	0.0333
Log(price of building plots)		0.0372	0.0380	0.0111	0.0490
		0.0260	0.0260	0.0086	0.0334
Share of rented agricultural area		-0.0130 ***	-0.0130 ***	-0.0039 ***	-0.0169 ***
		0.0018	0.0018	0.0011	0.0018
Spatial lag		0.2344 ***			
		0.0658			
Spatial error		0.2922 ***			
		0.0775			

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

TABLE A2

Regression Results for 2001 Using Spatial 2SLS/GMM for the Mixed-log Model with a Gabriel Weight Matrix

		Spatial 2SLS/GMM			
		coeff.	direct	indirect	total
Constant	coeff.	4.9160 ***			
	SE	0.6546			
Public buyer		0.2229 ***	0.2261 ***	0.1498 ***	0.3759 ***
		0.0195	0.0194	0.0467	0.0535
Public seller		0.2386 ***	0.2423 ***	0.1605 ***	0.4028 ***
		0.0424	0.0429	0.0567	0.0854
Direct payments		0.0003	0.0003	0.0002	0.0004
		0.0002	0.0002	0.0001	0.0003
Soil quality index		0.0140 ***	0.0142 ***	0.0094 ***	0.0236 ***
		0.0007	0.0007	0.0029	0.0030
Log(size of a transacted plot)		0.1122 **	0.1138 **	0.0748 **	0.1886 **
		0.0491	0.0500	0.0404	0.0854
Distance to the next urban center		-0.0008	-0.0008	-0.0007	-0.0015
		0.0023	0.0023	0.0016	0.0039
Log(ratio building vs. agricultural land)		0.0958 ***	0.0974 ***	0.0631 ***	0.1605 ***
		0.0248	0.0252	0.0220	0.0408
Log(price of building plots)		0.1102 ***	0.1118 ***	0.0696 ***	0.1815 ***
		0.0341	0.0338	0.0185	0.0449
Share of rented agricultural area		-0.0088 ***	-0.0089 ***	-0.0057 ***	-0.0146 ***
		0.0019	0.0019	0.0015	0.0027
Spatial lag		0.3981 ***			
		0.0752			
Spatial error		0.3289 ***			
		0.0789			

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

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