Determinants of agricultural cash rents in Germany: A spatial econometric analysis for farm-level data

Breustedt G. and Habermann H.

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Breustedt G. [1] and Habermann H. [1]

[1] Department of Agricultural Economics, Christian-Albrechts-Universität Kiel, Germany

Abstract- **We empirically analyse the determinants of cash rent levels for agricultural land in Lower Saxony, Germany. We are the first to apply a spatial econometrics approach that accounts for two types of spatial dependence simultaneously to cash rent data at the farm-level. Our empirical results underline the usefulness of such an approach. Farm characteristics which serve as a proxy for the marginal value of rented acreage for the tenant as well as variables which represent local competition on the land market are significant. Among the farm characteristics, operating revenue per hectare, share of high-value crops, soil quality, share of rented acreage, share of arable land relative to rented acreage, and animal density are significant while, ceteris paribus, neither labour nor machinery/buildings per hectare nor farm size are significant. In particular, animal density at the regional level increases the cash rent, underlining the importance of local competition on the land market. The analysis also shows that subsidies which foster competition among farmers for rented land boost landlords' incomes. Thus, evaluation of set-aside programs or evaluation of public support for investment in pig or poultry production or renewable energies has to take such side-effects into account.**

Keywords- **Cash rent, farm-level data, spatial econometrics.**

I. INTRODUCTION

One goal of the Common Agricultural Policy (CAP) of the EU is to support farmers' incomes. However an important question arises: how much of the public money spent on direct payments, intervention buying, investment aid or gas rebate actually transforms into farmers' income, and how much gets passed on to landlords via increased land cash rents. To understand distributional effects of the CAP and certain national policies, empirical analyses of the determinants of cash rents are essential.

While studies analysing determinants of farmland prices are numerous, those that deal with cash rents are less common. Empirical evidence is quite limited. Brümmer and Loy [1], Doll und Klare [2] as well as Drescher and McNamara [3] analyse determinants of regional cash rent levels in Western Germany. Fuchs [4] analyses Eurostat data for Belgium, Denmark, France, Germany and the Netherlands at the level of provinces and federal states, respectively. For the US, land rental market analyses have been conducted by Herriges et al. [5], Bierlen et al. [6], Roberts et al. [7], Lence and Mishra [8] as well as Jannsen and Button [9]. An important shortcoming of most studies is the application of regional data. By aggregating farm-level data into county averages, a considerable amount of variation is eliminated and important information, such as the impact of different factor endowments, cropping patterns or personal abilities among different farmers in a region, is lost. We thus argue that the use of farm-level data is more appropriate for analysing the determinants of land cash rents.

In addition, the spatial nature of the data may result in dependence (spatial autocorrelation) of the observations under study (Anselin [10]). First, cash rents may not only be influenced by covariates in the same location but also by cash rents paid in neighbouring regions (spatial lag dependence). Second, disturbance terms of neighbouring observations may be correlated, because they exhibit the same unobserved characteristics (spatial error dependence). These characteristics may cause standard econometric techniques to become inappropriate. While Lence and Mishra [8] as well as Fuchs [4] explicitly take into account spatial errors for regional data, none of these empirical analyses controls for spatial lag dependence.

Hence the contribution of the study at hand is twofold. First, we apply estimation techniques accounting for both spatial lag and error effects simultaneously. We thereby account for the spatial price transmission of cash rents. This is new to the field of agricultural land markets and the combined spatial estimations are rare within agricultural economics (compare Anselin and Bera [11]; Holloway et al. [12]). Second, we conduct the first empirical analysis of the determinants of cash rents in Europe at the farm-level.

The remainder of the article is organised as follows: section II illustrates the model and sets out the methodology, section III describes the data followed by the presentation of results in section IV. Section V offers conclusions.

II. MODEL AND METHODS

A. *Economic background*

Mainly, we follow the income approach also used by Roberts et al. [7] as well as Lence and Mishra [8]. It assumes land rent decisions to be based on profit maximisation. The derivative of the profit function with respect to land then gives the marginal revenue product for land farmed. Now consider on the one hand a farmer operating only his own acreage versus a farmer only using rented acreage. The former might indeed be willing to pay his marginal revenue product for cash rent when he starts to rent additional acreage. The latter is however not able to pay this price for the total rented acreage. He also has to cover fixed costs from his revenues of the rented acreage and hence may just be able to pay lower cash rents. We have to account for this difference when it comes to the choice of variables because on average our sample farms rent more than half of their acreage. But even if the ability to pay for the considered farmer is measured correctly the cash rent may furthermore depend on local competition among farmers. The marginal revenue product of the rented acreage is shared between tenants and landlords with respect to this local competition. This will be taken into account by the use of proxy variables in the empirical analysis.

B. *Spatial dependencies*

Basically spatial dependencies are modelled as extensions of a standard linear regression model: the spatial lag and the spatial error model (Anselin [10]; Anselin and Bera [11]). While the former deals with interactions of agents' decisions by allowing spatial relationships among observations of the dependent variable, the latter addresses spatial patterns in the error terms. In our case the cash rent level at the observed location may be influenced by neighbouring cash rent levels, because farmers may act as tenants in neighbouring communes. Hence cash rents of nearby observations probably influence each other.

Equation (1) illustrates the spatial lag formulation for this case.

$$r = \rho W_1 r + X\beta + u \tag{1}$$

where r is the N by 1 vector of the cash rent per hectare (N = number of observations), W_1 is a N by N spatial weight matrix illustrating the spatial relationship, e.g. distances, among sample farmers, X is the N by K matrix of exogenous explanatory variables, ρ is a spatial autoregressive parameter to be estimated, β the K by 1 vector of regression coefficients to be estimated and u is the N by 1 vector of random, independent and identically distributed (i.i.d.) error terms. The spatial weight matrix W_1 illustrates the assumed spatial relationship between all pairs of observations. Usually W is row-standardized and hence the spatial lag operator $W_1 r$ is a weighted average of cash rents at neighbouring farms. We will explain the weight matrix in detail below. If a spatially lagged dependent variable is falsely ignored a specification error of the omitted variable type occurs and the ordinary least square (*OLS)* estimators will be biased.

Another way to incorporate dependencies over space is through a spatial error specification. The variance-covariance matrix of the disturbance terms, Cov[uu´], exhibits spatial dependence when the off diagonal elements are non-zero following a certain spatial structure. Formally,

$$r = X\beta + u \qquad \text{with} \qquad u = \lambda W_2 u + \varepsilon \tag{2}$$

where the error u of the standard linear regression consists of an error lag $W_2 u$ with the spatial coefficient λ. Here ε is the standard N by 1 vector of i.i.d. error terms. If a form of spatial autocorrelation in the disturbance terms is present and ignored the OLS estimates remain unbiased, but become inefficient.

If a spatial lag model still contains spatial autocorrelation in the disturbance terms both models can be combined (see Case et al. [13]; Anselin and Bera, [11]). Formula (1) can also be expressed as

$$r = (I - \rho W_1)^{-1} X\beta + (I - \rho W_1)^{-1} u \tag{3}$$

and (2) as

$$u = (I - \lambda W_2)^{-1} \varepsilon \tag{4}$$

the combination of (3) and (4) and some reformulation gives

$$r = \rho W_1 r + \lambda W_2 r - \rho\lambda W_1 W_2 r + X\beta - \lambda W_2 X\beta + \varepsilon \quad (5)$$

However in our case we apply the same weight matrix for both spatial components. Hence formula (5) becomes

$$r = (\rho + \lambda)Wr - \rho\lambda W^2 r + X\beta - \lambda WX\beta + \varepsilon \quad (6)$$

Statistical test procedures are available to decide whether the application of such a model combining both spatial components (formula 5 or 6) compared to an *OLS*, lag (formula 1) or error (formula 2) specification is appropriate.

We now turn to the specification of the weight matrix. We define the weights *wij* within the matrix on the base of inverse distances between the communes farms *i* and j are located in (see formula 7). Weights are equal for farms located in the same commune. The weights in *W* satisfy *wij* > *wip* if farms *i* and *j* are located in the same commune C and if *p* is not located in C. Hence weights of farms within the same commune are larger than weights of farms in neighbouring communes. In addition we row standardise *W*. We also introduce a cut-off level at 20 km[1]. We use inverse distances and a cut-off because transportation costs increase with the distance up to a certain level beyond which a distant farmer is not able to compete with local farmers.

As an example, our *W* is

$$W_1 \equiv \begin{bmatrix} \begin{bmatrix} 0 & w_{ji} \\ w_{ij} & 0 \end{bmatrix} & w_{pi} & . & w_{ni} \\ & & & . \\ w_{ip} & 0 & & . \\ . & & 0 & . \\ w_{in} & . & . & . \quad 0 \end{bmatrix} \quad (7)$$

III. DATA

The data we use are taken from profit-and-loss-statements of farms located in the German federal state of Lower Saxony. They are provided by Landdata Ltd.; the market leader of farm accountancy services in Germany. Additionally we include county- or commune-averages from an agricultural census survey conducted by the Federal Statistical Office of Germany. In total we base our estimations on 4564 farm observations including three-year averages of farm-level data and commune- or county-averages for the years 1999-2001 (see Table 1).

1. Different cut-off levels yield minor changes in the results.

Table 1: Variable definition and summary statistics (n = 4564)

Variable	Definition	Mean	Std. Deviation	Minimum	Maximum
rent	Annual cash rent per hectare in € (2001 farm average)	258.66	149.79	13.20	1495.0
revenue	Operating revenue plus wages, rents and interest expenditures; minus crop premiums	509.02	518.45	-7323.76	6967.0
soil	Soil quality (Ertragsmesszahl)	3553.1	1440.5	1000	10000
acreage	Farm size in hectare (1996)	71.91	41.25	7.50	630.68
rentshare	Share of rented acreage to total farmed acreage	0.55	0.26	0.0051	1.39
arablerentshare	Share of rented arable land to total rented area	0.70	0.31	0	1
sbeetshare	Share of sugar beets in cropping pattern	0.051	0.080	0	0.47
potshare	Share of potatoes in cropping pattern	0.044	0.098	0	0.94
vegshare	Share of vegetables in cropping pattern	0.0016	0.0166	0	0.67
andensity_ha	Animal density in 500kg per hectare	1.14	1.10	0	13.45
labour	Employees per hectare	0.0264	0.059	0.000158	3.32
capital	Capital per hectare (€/ha), capital stock minus milk- and sugar beet quota and minus land	3741.2	2361.5	19.58	35848.4
andesity_com	Animal density on commune level (500kg/ha)	0.93	0.52	0.0026	3.40
popden_county	Population density on county level (inhabitant/km^2)	167.51	109.99	42.46	490.81
popchange_county	Population change over eleven years (1995-2005) on county level	0.0470	0.0471	-0.0746	0.139
unempl_county	Unemployment rate on county level (percentage of labour force, 2001)	8.91	2.193	5.5	14.8
income_county	Average income per inhabitant on county level (€)	15557	1174	13222	19056
premiums	Per-hectare premiums for cash crops (in € according to 9 different yield regions)	284.83	36.45	226.50	371.20
Further controls:					
education	Education of farmer: (1) university, (0) other	0.247	0.428	0	1
regularbasis	Operation: Regular (0) versus sideline basis (1)	0.074	0.262	0	1
booking	Gross (0) vs. net book-keeping (1)	0.099	0.298	0	1

The rent variable represents the average rent of the existing cash rent contracts on a single farm. In Germany a considerable share of the rent contracts are signed for time periods around 6, 9 or 12 years and hence our variable also includes contracts signed some years ago. Thus, we take the rent data of 2001 which probably do not suffer from too many decisions about lending before the introduction of premium payments by the European Union in 1993. Additionally a certain share of contracts may have been set up between relatives possibly leading to a downward bias of cash rents compared to contracts at arm's length.

As illustrated above we cannot use the marginal revenue product because our sample farms rent a considerable share of their land. As a natural starting point for the choice of variables we, thus, use a farm's operating revenues from agricultural activities per hectare except premium payments from the EU. To account for marginal revenue products of land which may differ from the average revenue per hectare we incorporate additional farm-level variables. As explained above we expect the share of rented acreage to play a decisive role for a farmer's willingness to pay. Soil quality, farm size, the share of several high-value crops (sugar beets, potatoes and vegetables), labour and capital endowment per hectare as well as animal density are also used. We instrument the farm size by its 1996 value to avoid endogeneity.

To illustrate the local competition a farmer might face, animal density on the commune level is supposed to proxy for the regional demand for land. This is especially relevant in Germany due to manure- and tax-regulations allowing not more than a certain amount of animal units per hectare. As variables that may represent farmers' benefits from quitting agricultural production we include the unemployment rate and the average income per inhabitant, both on the county level. For the same reason we controlled for population density and population change on the county level. High opportunity costs will probably increase farm exit rates and hence the supply of land. This should induce lower cash rent levels.

We also use a variable for the premium payments that are paid annually on a per hectare basis for cash crops. These premiums are paid according to a historical grain yield reference in 9 different yield regions. Some further control variables are also included.

IV. RESULTS

A Moran's I test (Moran [14]) reveals significant spatial dependence. In our case we yield a test statistic of 16.4, which is highly significant (see Table 2). As a diffuse test (compare Florax and de Graaff [15]) Moran's I is indicative of spatial dependence, but does not point to a specific alternative. However we have to estimate a combined spatial model (according to formula 5/6)[2]. Both spatial components yield significant coefficients.

We also estimate the spatial error and lag model separately. In each of the specifications the spatial components are significant. Compared to the error specification the log-likelihood of our spatial lag formulation is superior suggesting that cash rents may indeed influence each other. A Lagrange multiplier test based on the residuals from this specification model can be used to examine whether the inclusion of the spatial lag term eliminates spatial dependence. In our case the residuals of the model show that significant spatial correlation still exists in the errors and thus the combined model is appropriate (see Anselin and Bera [11], p. 265).

2. We apply the Matlab routines for spatial econometrics from LeSage [16], chapter 3.

Table 2: Results for the general spatial specification (n = 4564)

	OLS		*Spatial error model*		*Spatial lag model*		*Combined spatial model*	
const	104	*	211	**	9.67		-30.0	
revenue	0.0375	***	0.0311	***	0.0332	***	0.0348	***
soil	0.00286		-0.00045		-0.00023		-0.00158	
soil squared	0.00144	**	0.00154	**	0.00139	**	0.00137	**
acreage	-0.0665		0.0879		0.0858		0.1161	
acreage squared	0.000076		-0.00024		-0.000211		-0.000248	
rentshare	37.9	***	25.8	***	27.9	***	28.1	***
arablerentshare	82.3	***	77.8	***	72.6	***	62.7	***
sbeetshare	194	***	225	***	172	***	133	***
potshare	46.6	**	52.7	**	41.3	*	27.3	
vegshare	296	**	316	***	301	***	289	**
andensity_ha	5.68	**	6.50	***	5.68	**	4.62	**
labour	45.4		33.7		35.5		35.3	
capital	0.000341		0.000127		0.000149		0.000072	
andesity_com	92.4	***	65.9	***	59.1	***	44.2	***
popden_county	-0.0421		-0.0241		-0.0096		0.0019	
popchange_county	646	***	546	***	336	***	214	***
unempl_county	-3.12	**	-5.85	**	-0.48		0.67	
income_county	-0.0272	***	-0.0261	***	-0.0144	***	-0.0090	***
premiums	1.23	***	1.02	***	0.53	***	0.28	***
education	3.65		4.73		4.15		3.20	
regularbasis	10.4		16.2	**	14.3	*	12.6	
booking	-13.9	**	-16.8	**	-12.4	*	-8.2	
spatial lag					0.500	***	0.690	***
spatial error			0.564	***			-0.543	***
log-likelihood			-27094		-27075		-24449	
Moran's I	16.4	***						

Note: Significance levels: * 10%, ** 5%, *** 1% or lower.

Additionally to the above given estimation we conducted regressions for arable land cash rents exclusively. The obtained results are quite robust and underline the results of the given specifications.

In the following paragraphs we will discuss the results obtained from the combined spatial model. In line with section 2 we find that the operating revenue per hectare positively affects the paid cash rent. A revenue increase by 1€ lifts the cash rent by 3.5 Cent. Fuchs [4] reports a coefficient of 0.1 for his variable of farm net value added without premiums. Though the comparability between different market return or revenue measures in existing studies is seldom given, most studies yield positive signs for these determinants. Our result is new to the literature since we show the influence on the farm and not on the regional level.

The group of farm characteristics intended to account for acreage with marginal revenue products that differ from average revenue also contributes to the explanation of cash rents. The squared soil quality, the share of rented acreage, the share of rented arable acreage as well as the livestock density show a positive influence on the endogenous variable. The signs of these factors are in line with the studies of Bierlen et al. [6], Drescher and McNamara [3] and Fuchs [4] except the sign of the share of rented acreage. While Fuchs [4] yields a negative sign ours is positive in contrast to our expectation. The variable may hence illustrate a tenant's willingness to pass a higher share of the marginal revenue product of acreage to the landlord. One can probably not identify this relationship by means of regional data.

Also cash rents increase with higher shares of sugar beets or vegetables within the cropping pattern. Thus the expected positive influence of high value crops is confirmed. Bierlen et al. [6] also show significant cash rent increases if the considered acre is planted with soybeans for example.

Furthermore the significance of county averages like livestock density, the population change or the average income per inhabitant indicates the importance of the local competition a farmer is embedded in. An increase of the livestock density in the commune by one unit (500kg) per hectare leads to a cash rent increase of 44.2€. Existing studies such as Drescher and McNamara [3] or Fuchs [4] come to equal results at least with respect to the sign. While the former use a different measure of animal density Fuchs [4] applies the same measure of animal units per hectare and obtains a coefficient six times higher than ours. This impact of animal density is important for policy makers deciding on investment aid for pig fattening.

The negative sign of the variable for average income suggests that in regions with good opportunities (e.g. in urban fringes) farmers may quit farming more easily and offer their land to tenants. An income increase by one € reduces the cash rent by approximately one Cent. However if regions have shown strong population growth this has a contrary impact. The growth may result in a lower availability of land for agricultural utilisation.

The coefficient of premium payments calls for further analysis because it reacts quite sensitive on the specification. Nearly 30% of the premium payments are capitalised into land rents. Existing studies like Brümmer and Loy [1] or Janssen and Button [9] yield somewhat lower incidence levels of government payments, while Lence and Mishra [8] yield coefficients of nearly one for market loss assistance (MLA) and production flexibility contracts (PFC) in the US. An interesting new study from Patton et al. [17] reveals that decoupled payments which are linked to land fully capitalise into land rents.

Interestingly, farm size does not seem to impact the level of cash rents. By the inclusion of measures like capital- and labour-intensity as well as animal density we account for farm size in the sense of the European Size Units (ESU).

Comparable to Bierlen et al. [6] the education of the farmer does not play a significant role within our estimation.

Lastly both spatial components are significant. The positive sign of the spatial lag estimator ρ with a coefficient of 0.69 indicates that an increase of the average neighbouring cash rent by one € rises the cash rent paid by the considered farm around 70 Cents.

V. CONCLUSIONS

The investigation of agricultural cash rents has yielded limited insights into the determinants of rent rates in the empirical literature so far. However thorough analyses may be of interest not only for agricultural economists but also for policy makers. We show that the use of both farm-level data and estimation techniques of spatial econometrics may contribute to the limited knowledge that exists.

We find that the average operating revenue per hectare determines cash rents to a considerable degree. Furthermore, farm characteristics such as soil quality and the share of certain high-value crops increase rents, while farm size or endowment with labour and capital appear not to affect land rents.

Our findings imply that government support increasing agricultural income will be partly passed on to landlords via inflated rent payments. This holds for the per hectare premiums of the CAP on the one hand. On the other hand variables representing the degree of competition on local land markets are also an important determinant of the payment shares that end up in the landlords' pockets. The strongly positive impact of (regional) livestock densities on cash rents should be of interest to policy makers deciding on investment aid schemes for livestock production. Such subsidies step up competition among farmers for rented land, thus boosting landlords' incomes. Evaluation of set-aside programs, investment aid schemes or renewable energies schemes should take such distributional side-effects into account.

Future research should focus on the incidence of premium payments: what percentage of these payments gets passed on to landlords? An in-depth analysis of this issue should be based on newly contracted cash rent decisions – rental agreements concluded after the entering into force of the Luxembourg Agreement. Another useful line of enquiry might focus on rental agreements between relatives and those concluded at arm's length.

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