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Land Economics, Volume 95, Number 1, February 2019, pp. 108-123 (Article)

Published by University of Wisconsin Press

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The Spatial and Temporal Diffusion of Agricultural Land Prices

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ABSTRACT *In the last decade, many parts of the world experienced severe increases in agricultural land prices. This price surge, however, did not take place evenly in space and time. To better understand the spatial and temporal behavior of land prices, we employ a price diffusion model that combines features of market integration models and spatial econometric models. An application of this model to farmland prices in Germany shows that prices on a county-level are cointegrated. Apart from convergence toward a long-run equilibrium, we find that price transmission experiences short-term adjustments caused by neighboring regions.* (JEL C23, Q24)

1. Introduction

In the last decade, many parts of the world experienced drastic increases in agricultural land prices. In the European Union, agricultural land prices in Germany surged by almost 150% from an average of 8,909 €/ha in 2006 to 22,310 €/ha in 2016 (Federal Statistical Office of Germany 2017). In France, average farmland prices increased by 33% in the last decade, reaching 5,940 €/ha in 2014, while land prices in the United Kingdom more than doubled during the same time (Eurostat 2016). Likewise, in the United States the average value of cropland increased from 6,252 $/ha to 10,107 $/ha between 2007 and 2017 (U.S. Department of Agriculture 2017). Drivers of this price surge are claimed to be higher land rents due to increased productivity and food prices, the conversion of agricultural land to nonagricultural uses, and speculative activities of financial investors (e.g., Deininger and Byerlee 2011). Farmers and politicians are concerned about this development, since high land prices are an obstacle for the expansion of family-operated farms. In addition, the concentration of farmland in the ownership of large holdings or nonagricultural investors is suspiciously viewed. Indeed, many governments take actions or contemplate measures that target the capping of land prices. For example, in 2014, Belgium laid the foundation for new land market instruments, such as the land observatory, land bank, and updated preemption rights. Belgium also tightened land market regulations, which had previously been liberal. In the same year, new land market regulations aiming to restrict the purchase of agricultural land by foreigners and nonfarmers were released in Slovakia. Likewise, in Germany, the Federal Ministry and the State Ministries of Agriculture are currently discussing bills that target the broad distribution of land ownership, the prevention of dominant land market positions on the supply and demand sides, the capping of land rental and sales prices, and the special treatment of farmers over nonagricultural investors.

It should be noted, however, that the surge of agricultural prices, which triggered the aforementioned policy debate, did not take place evenly in space and time. For example, land prices in western and eastern Germany differ significantly, even 20 years after reunification. Not only do price levels vary, but growth rates of land prices also vary between and within countries. In France, for example, significant double-digit increases took place

Land Economics • February 2019 • 95 (1): 108–123
ISSN 0023-7639; E-ISSN 1543-8325
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from 2011 to 2014 in northern parts (+38%) and western parts (+11%), whereas land prices declined in other regions, notably in the Mediterranean area (–8%) (Eurostat 2016). Italy also witnessed uneven price development: land values almost doubled from 1992 to 2010 in northern Italy, while land values increased by only 15% to 30% in the central and southern regions (Mela, Longhitano, and Povellato 2012).

Yang, Ritter, and Odening (2017) show that even on a regional scale, agricultural land markets may exhibit different dynamics. Potential causes of diverging land prices are different agricultural production systems and disparities in regional growth, in conjunction with the limited mobility of agricultural production. On the other hand, it is widely acknowledged that land prices are sticky across space. This is not only due to the spatial correlation of land price characteristics, such as soil quality, but also an implication of adjustments to demand and supply shocks of land markets. For example, if land prices in the urban fringe increase because agricultural land is converted to commercial land, liquid farmers will likely acquire agricultural land in the neighboring area as a substitute and thus increase land prices. Likewise, if a windfarm is built, this not only creates significant rents in that region, but also generates regional spillover effects since ecological compensation areas have to be established elsewhere. Ritter et al. (2015) provide empirical evidence for this "ripple effect" in Brandenburg, Germany. However, thus far it is not well understood how fast this kind of spatial price transmission works and whether it describes a local or regional phenomenon. From a policy perspective, as well as for the optimal timing of land sales, it is of great interest to know whether regional land price differentials diminish and how price shocks diffuse in space.

At least three types of statistical models can be distinguished that aim to explain the behavior of land prices: spatial econometric models, time series models, and spatiotemporal models. Spatial econometric models, which encompass spatial lag and spatial error models, are more or less the standard now in hedonic models of land prices (e.g., Huang et al. 2006; Patton and McErlean 2003; Hüttel et al. 2013). These models are static in nature and focus on measuring the unbiased impact of land attributes on land prices, while accounting for their spatial relationships. Time-series models are used to estimate trends and structural breaks in land price developments (Gutierrez, Westerlund, and Erickson 2007), test the present value model of prices, and detect price bubbles (Falk 1991). The third modeling approach, spatiotemporal models, seems to be the most suitable approach for our analysis because it captures both dimensions of interest: space and time.

There are only a few applications of spatiotemporal models to farmland prices. Maddison (2009) extends standard hedonic models of farmland values by including spatiotemporal lags of dependent and independent variables, which significantly increase the explanatory power of the regression. Carmona and Rosés (2012) apply panel unit root tests to explore the convergence of farmland prices in Spanish provinces at the beginning of the twentieth century. They find that the Spanish land market is spatially integrated and interpret this finding as an indicator of land market efficiency. More recently, Yang, Ritter, and Odening (2017) apply second-generation panel unit root tests in an iterative procedure to identify "convergence clubs" of regional land markets that share the same price development.[1] Though panel unit root tests give a first impression of the similarity of price trends in different regional land markets, they do not allow for a complete description of price diffusion processes. More specifically, it is not possible to distinguish between convergence, cointegration, and spatial diffusion. In this context, cointegration

[1] The term "convergence clubs" was coined by Baumol (1986) to describe groups of countries showing similar economic development. In the context of land markets, convergence clubs can be understood as regions that share similar price dynamics. A testable hypothesis is that price differences among members of a convergence club vanish in the long run in both absolute and relative terms (e.g., Abbott and de Vita 2013; Montagnoli and Nagayasu 2015 for real estate markets). Members of a convergence club can be characterized, for example, by similar agricultural production structures. The economic forces that cause price convergence include mobility of farmers or investors (i.e., arbitrage processes), diffusion of technologies (e.g., biogas plants), or information spillovers.

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establishes a long-run equilibrium among (nonstationary) land prices of two or more regional land markets. Convergence of land prices can be considered as a special case of cointegration, in which land prices move toward the same price level (absolute convergence) or toward a constant difference (relative convergence). Spatial price diffusion, on the other hand, can also be driven by short-run effects and contemporaneous spillovers from neighboring regions. Pesaran and Tosetti (2011) suggest a price diffusion model that is able to disentangle these effects, and Holly, Pesaran, and Yamagata (2011) use this model to analyze the spatial and temporal diffusion of house prices in the United Kindgdom. A nice feature of this model is that it enables the testing of whether a specific region is dominant, in other words, that the region is typically the source of price shocks that are then transmitted to neighboring regions with a time delay, and there are no feedback effects. Such a phenomenon is often observed for big cities in the context of house prices (Meen 1999; Lee and Chien 2011).

In this paper, we apply the price diffusion model of Pesaran and Tosetti (2011) to study the behavior of farmland prices in the state of Lower Saxony, Germany. The overall objective of this model class is the inspection of statistical characteristics of land price processes rather than the identification of economic drivers by means of a structural model. Nevertheless, this kind of analysis is relevant from an economic perspective, since price cointegration and convergence are fundamental properties of prices in the context of spatial market integration, which is an important aspect of market efficiency. Within this modeling framework, we are able to answer a set of interesting research questions: Are regional land markets separated or are they integrated such that prices converge in the long run? If low-price regions catch up with high-price regions, how long does this adjustment take? Can we find ripple effects in farmland markets? Is it possible to identify dominant regions in farmland markets, such as in areas with high land rents or in close proximity to urban land markets? Although we target a description of land price dynamics rather than a full economic explanation of these dynamics, it is an important step toward a more comprehensive understanding of farmland markets.

2. Methodology

Beenstock and Felsenstein (2007) develop a spatial vector autoregression model, which is motivated by the ability to explicitly consider the potential impacts of economic events in space. In this model, which consists of temporally lagged terms and spatially lagged terms, the land prices in region i at time t are given by

$$p_{it} = c_i + \sum_{l=1}^{L_{\alpha i}} \alpha_{il} p_{i,t-l} + \sum_{l=0}^{L_{\rho i}} \rho_{il} \overline{p}_{i,t-l} + u_{it}, \qquad [1]$$

where p_{it} denotes the (log) land price in region i at time t, $i = 1,\ldots,N$ and $t = 1,\ldots,T$; c_i is a region-specific fixed effect; $p_{i,t-l}$ is the time-lag of the dependent variable with weights α_{il}; $\overline{p}_{it}$ is the spatially lagged average price with temporal lags $\overline{p}_{i,t-l}$ and weights ρ_{il}. $L_{\alpha i}$ and $L_{\rho i}$ denote the region-specific maximum numbers of temporal lags for the dependent variable and its spatially lagged prices; and u_{it} is an error term, which can consider spatial correlation.

There are several weighting schemes for spatial structures in the spatial econometric literature based on contiguity or distance. Since average land prices per region do not have a distinct spatial core, we employ the queen contiguity scheme, namely, that two regions are considered neighbors if they share a common border. The average neighbor price is then calculated as the weighted average price of neighbors according to $\overline{p}_{it} = \sum_{j=1}^{N} w_{ij} p_{jt}$ with weights w_{ij} defined as follows:

$$w_{ij} = \begin{cases} \dfrac{1}{N_i} & \text{if } i \text{ and } j \text{ share a border, } i \neq j \\ 0 & \text{otherwise} \end{cases}, \qquad [2]$$

where N_i denotes the number of neighbors of region i and it follows that $\sum_{j=1}^{N} w_{ij} = 1$.

Since asset prices are typically nonstationary, it is useful to employ spatial cointegration methods. The relationship between cointe-

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grated variables is captured by vector error correction models. Whereas conventional vector error correction models consider only temporal dynamics, spatial vector error correction models incorporate both spatial and temporal dynamics (Beenstock and Felsenstein 2010). In this framework, the long-run relationship between prices in a region and the average prices in neighboring areas can be modeled through the following spatial autoregressive equation:

$$p_{it} = \delta_i + \beta_i \overline{p}_{it} + \mu_{it}, \tag{3}$$

where $\overline{p}_{it}$ denotes the spatially lagged price as defined above, δ_i is a region-specific fixed effect, and μ_{it} is an error term. If β_i is significant and μ_{it} is stationary, there exists a long-run equilibrium between prices in region i and the average prices in the neighboring area.[2] Temporary deviations from the long-run equilibrium in the previous period, (i.e., $\mu_{i,t-1} = p_{i,t-1} - \delta_i - \beta_i \overline{p}_{i,t-1}$) are corrected toward the equilibrium relation through the adjustment speed ϕ_i:

$$\Delta p_{it} = \gamma_i + \phi_i (p_{i,t-1} - \delta_i - \beta_i \overline{p}_{i,t-1}) + \sum_{l=1}^{L_{ia}} a_{il} \Delta p_{i,t-l} + \sum_{l=0}^{L_{ib}} b_{il} \Delta \overline{p}_{i,t-l} + \epsilon_{it}, \tag{4}$$

where $\Delta p_{it} = p_{it} - p_{i,t-1}$, $i = 1,...,N$, and $t = 1,..., T$; γ_i denotes region-specific fixed effects; $\sum_{l=1}^{L_{ia}} a_{il} \Delta p_{i,t-l}$ describes short-run dependencies of prices in region i; $\sum_{l=1}^{L_{ib}} b_{il} \Delta \overline{p}_{i,t-l}$ describes short-run dependencies of average neighbor prices; $b_{i0} \Delta \overline{p}_{it}$ captures the contemporaneous effect on the average neighbor price; and ϵ_{it} is an error term.[3] If the null hypothesis of no cointegration is rejected by a Johansen test, we further analyze whether the price in region i converges to the average neighbor price. In the case of convergence, prices are cotrending and the cointegrating vector $(1, -\beta_i)$ equals $(1, -1)$. Although this provides evidence of a possible clustering of cointegration outcomes, price convergence is not necessary for spatiotemporal price diffusion.

So far, spatiotemporal price diffusion has been restricted to adjacent regions. However, price changes in one region may also affect regions located further away. The phenomenon of a spillover of shocks from one region to others leading to a global effect on prices in all other regions is referred to as a spatial ripple effect (Meen 1999). This effect can be regarded as a special case of price diffusion since (1) the diffusion area not only includes nearby regions, but also farther areas; and (2) the diffusion direction is one way, which means that a shock starting in one center region spreads to other regions and there are no feedback effects. In empirical applications on house markets (e.g., Holly, Pesaran, and Yamagata 2011; Helgers and Buyst 2016), a large city or major financial center, usually with the highest house prices, is considered the dominant region that drives price development in all other regions. To test whether a region is potentially dominant, the following pairs of equations are estimated for all other $N-1$ regions:

[2] One may question the existence of a stable long-run relationship in the regional land market equation as postulated by equation [3]. We argue that despite of the immobility of land, one can expect that economic responses to spatial price differentials will take place, at least if they are pronounced. For example, after the German reunification, many farmers from West Germany or other Western European countries bought or rented land in East Germany at prices that were considerably lower than in parts of West Germany. Moreover, despite legal barriers, nonagricultural investors participate in agricultural land markets. That is, although land is immobile, the mobility of capital and/or farm managers will likely induce arbitrage processes on land markets toward spatial equilibrium. On the other hand, technological change may affect the long-run equilibrium between spatially lagged prices. However, we presume that changes in transportation technologies (infrastructure), which have an obvious effect on transportation costs and spatial integration in commodity markets, have a smaller impact on land prices because arbitrage mechanisms are different. Nevertheless, we cannot rule out time-varying equilibria. To address this issue, we allow for structural breaks when testing the cointegration relations in our empirical application.

[3] Since our primary concern is the analysis of the spatial and temporal dispersion of land price shocks and not the explanation of the determinants of agricultural land prices, we do not add other control variables to the model. The dominant region, which will be included in the next step, however, can be regarded as a common factor for the other regions (Holly, Pesaran, and Yamagata 2011; Chudnik and Pesaran 2013).

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$$\Delta p_{0t} = d_{0i} + \phi_{0i}(p_{0,t-1} - \omega_{0i} - \beta_{0i} p_{i,t-1}) + \sum_{l=1}^{L} a_{0il}\Delta p_{i,t-l} + \sum_{l=1}^{L} c_{0il}\Delta p_{0,t-l} + \varepsilon_{0it}, \quad [5]$$

$$\Delta p_{it} = d_{i0} + \phi_{i0}(p_{i,t-1} - \omega_{i0} - \beta_{i0} p_{0,t-1}) + \sum_{l=1}^{L} a_{i0l}\Delta p_{i,t-l} + \sum_{l=1}^{L} c_{i0l}\Delta p_{0,t-l} + \varepsilon_{i0t}, \quad [6]$$

where $\sum_{l=1}^{L} a_{il}\Delta p_{i,t-l}$ and $\sum_{l=1}^{L} c_{il}\Delta p_{0,t-l}$ denote short-run dependencies from price changes in region i and in the dominant region 0, respectively, and ω_{i0} and ω_{0i} are region-specific fixed effects in long-run relations. The adjustment speed ϕ_{0i} in equation [5] describes how fast the price change in a potential dominant region 0, Δp_{0t}, is corrected toward a long-run equilibrium with region i (if existent). In contrast, in equation [6] the adjustment speed ϕ_{i0} depicts how fast the price change in region i, Δp_{it}, is corrected toward a long-run equilibrium with the potential dominant region 0. This estimation is repeated for all candidates for a dominant region. According to the definition of a dominant region, its price should affect prices in the other regions in the long run, that is, ϕ_{i0} should be significant for all i, whereas the price in the dominant region should not be affected by prices in other regions in the long run, that is, ϕ_{0i} should be insignificant for all.

If a candidate for a dominant region 0 passes the aforementioned test, the long-run equilibrium relationship in equation [3] is extended in the following way to account for the special role of the price in the dominant region, p_{0t}:

$$p_{it} = \omega_i + \beta_i \bar{p}_{it} + \beta_{i0} p_{0t} + \mu_{i0t}, \quad [7]$$

where $i = 1,\dots,N-1$ indicates the nondominant regions. Note that for direct neighbors, the dominant region is excluded in the calculation of the average price in the neighboring area $\bar{p}_{it}$.

With the long-run equilibrium in equation [7], the diffusion model from equation [4] can be adapted by adding the prices of the dominant region:[4]

$$\Delta p_{it} = \tau_i + \phi_i(p_{i,t-1} - \omega_i - \beta_i \bar{p}_{i,t-1} - \beta_{i0} p_{0,t-1}) + \sum_{l=1}^{L_{ia}} a_{il}\Delta p_{i,t-l} + \sum_{l=0}^{L_{ib}} b_{il}\Delta \bar{p}_{i,t-l} + \sum_{l=0}^{L_{ic}} c_{il}\Delta p_{0,t-l} + \epsilon_{it}, \quad [8]$$

where $i = 1,\dots,N-1$; the coefficient ϕ_i denotes the adjustment speed of region i to the new long-run equilibrium; and $\sum_{l=0}^{L_{ic}} c_{il}\Delta p_{0,t-l}$ captures the short-run dependencies of the price change in the dominant region, including a contemporaneous effect for $l = 0$. To confirm that region 0 is actually a dominant region, we check the significance of the coefficients ϕ_i and c_{i0}.

With the two abovementioned models shown in equations [4] and [8], the procedure for analyzing the diffusion of prices involves several steps. First, we carry out augmented Dicky-Fuller tests on the individual price series to discern the long-run price development in each region. The next step consists of Johansen tests for the pairwise cointegration between prices of each region and its average neighbor price, as well as the estimation of the long-run equilibrium vectors in the cointegrating equations to confirm that a long-run equilibrium relationship exists. In this case, we can use the error correction term from prices of neighbors to control for price changes. For the model with the average neighbor price and the dominant region, we also test for pairwise cointegration between prices in each region and the long-run equilibrium and estimate their long-run equilibrium vectors. If the null hypothesis of no cointegration is rejected, we can estimate the two diffusion models shown in equations [4] and [8]. Due to the inclusion of contemporaneous effects $\Delta \bar{p}_{it}$ and Δp_{0t}

[4] In general, higher-order terms of spatial dependence could be included in the model. Higher-order dependencies, however, are not ruled out in our model because the average neighbor price itself is influenced by the neighbors' neighbors' price. Since we allow for contemporaneous effects, this indirect influence can happen simultaneously—it just has to go through nearer neighbors. Moreover, more than one dominant region is possible. Given that our later results do not show a clear confirmation that the most suitable candidate for a dominant region is in fact dominant, we refrain from further extending the model.

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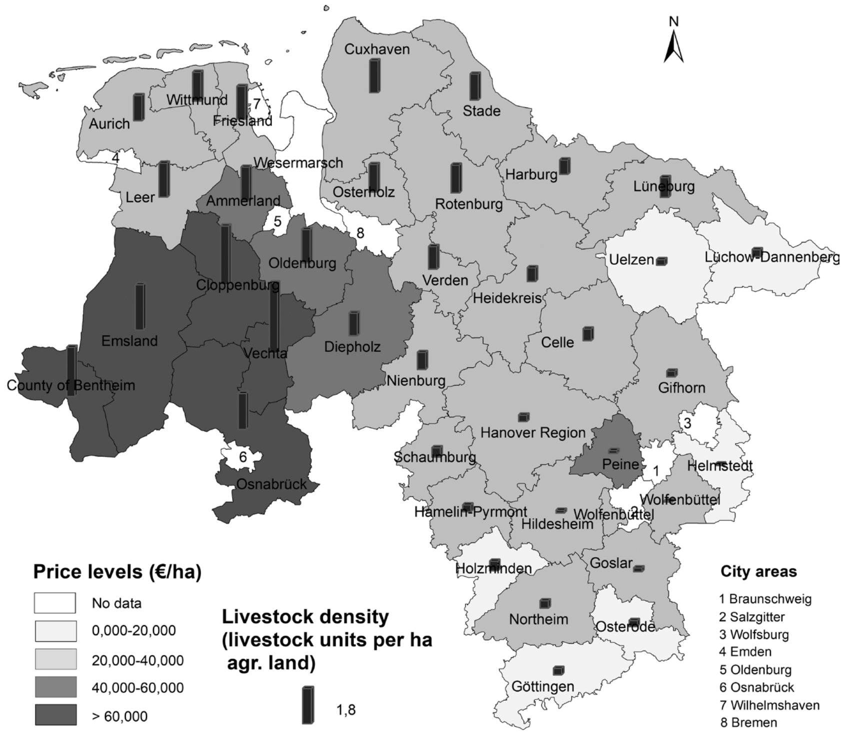
Figure 1
Regional Distribution of Land Prices (€/ha) and Livestock Density in Lower Saxony in 2015



in the two models, an endogeneity problem might appear. Hence, we conduct a Wu-Hausman test. If the Wu-Hausman test rejects exogeneity, we use instrumental variables for the contemporaneous terms. For regions with exogenous contemporaneous terms, we take seemingly unrelated regressions to estimate the system of price change equations to account for correlation in the error terms. Note that exogeneity of prices is a further condition to be fulfilled by a dominant region (cf. Holly, Pesaran, and Yamagata 2011; Helgers and Buyst 2016; Cook and Watson 2016).

3. Study Area and Data

In our empirical analysis, we study the diffusion of land prices in Lower Saxony, Germany. Lower Saxony is located in northwestern Germany and consists of 37 counties. It is the second largest state in Germany, covering an area of 47,600 km^2. About 60% of this area is used for agricultural production. In terms of production value, Lower Saxony is one of the leading states, contributing more than 20% to Germany's revenues from agriculture. However, natural conditions, production structures, and farm size structures largely differ across counties within Lower Saxony. This heterogeneity of agricultural production renders Lower Saxony an interesting study area. Differences in land use intensity translate into differences in both land rental prices and sales prices. Thus, the analysis of price diffusion processes is nontrivial.

Figure 1 depicts the spatial differences of the price levels and livestock density in Lower Saxony in 2015. Table 1 summarizes key variables of agricultural production at the county level. Three different regions can be distinguished. The eastern and southeastern part of Lower Saxony is characterized by fertile soils. In this region, farms are rather large (often more than 100 ha on average) and specialized in cash crops. The livestock density for most

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Table 1
Descriptive Statistics of Agriculture in Lower Saxony

County	Number of Farms	Average Farm Size (ha)	Share of Arable Land (%)	Average Soil Quality (Index Points)[a]	Livestock Density (LSU/ha)[b]	Land Sale Price 2015 (€/ha)	Price Growth Rate, 1990–2015 (%)	Ratio of Rental to Sales Price (%)
Ammerland	841	50.72	48	31	1.74	41,862	228	1.73
Aurich	1,315	62.85	48	42	1.31	28,716	128	1.61
Bentheim	1,140	51.14	86	30	2.55	60,882	214	1.52
Celle	632	82.64	79	35	0.60	20,368	95	1.88
Cloppenburg	1,758	54.33	87	32	3.05	78,441	264	1.56
Cuxhaven	1,857	73.38	45	42	1.65	26,631	134	2.07
Diepholz	1,693	76.51	82	36	1.17	47,312	240	1.80
Emsland	2,812	57.80	90	30	2.35	61,723	304	1.53
Friesland	576	76.19	34	41	1.70	35,670	109	2.16
Gifhorn	817	94.94	83	38	0.30	25,090	209	2.14
Goslar	289	95.19	87	61	0.20	24,348	62	2.40
Göttingen	726	79.15	86	57	0.34	19,707	52	1.91
Hamelin-Pyrmont	482	81.39	89	59	0.35	29,186	41	2.01
Hanover Region	1,481	78.23	84	50	0.34	36,419	69	1.44
Harburg	860	63.86	66	35	0.74	24,984	126	1.40
Heidekreis	900	77.17	69	32	0.71	26,226	148	1.84
Helmstedt	359	115.16	91	51	0.09	18,446	8	2.44
Hildesheim	811	83.73	94	71	0.14	34,539	36	1.54
Holzminden	321	79.65	74	57	0.49	17,829	38	1.92
Leer	1,138	59.05	26	32	1.76	35,941	179	1.97
Lüchow-Dannenberg	587	103.32	80	36	0.37	17,760	118	2.25
Lüneburg	10,480	76.55	65	39	1.02	20,774	176	2.62
Nienburg	1,169	69.98	84	35	0.88	31,244	160	2.23
Northeim	815	69.47	84	66	0.39	22,104	61	1.72
Oldenburg	955	66.87	76	31	1.73	55,414	259	1.34
Osnabrück	2,418	48.44	84	38	1.87	62,253	231	1.46
Osterholz	737	53.75	37	30	1.42	21,528	46	1.59
Osterode	242	64.11	70	55	0.32	14,553	155	2.31
Peine	401	89.19	91	60	0.15	41,094	84	1.85
Rotenburg	1,642	76.76	68	27	1.44	31,650	241	1.93
Schaumburg	440	76.32	86	64	0.48	31,898	42	1.59
Stade	1,276	62.87	52	40	1.34	34,521	211	1.89
Uelzen	693	107.51	90	34	0.29	19,474	68	2.41
Vechta	1,140	56.60	89	39	3.64	90,457	235	1.34
Verden	698	66.09	70	37	1.18	31,318	197	1.49
Wittmund	657	64.30	43	39	1.50	30,323	238	1.84
Wolfenbüttel	403	126.35	96	73	0.05	34,194	82	1.50

Note: Data regarding the number of farms, farm size, share of arable land, average soil quality, livestock density, and land sale price for arable land (1990, 2010, and 2015) are from the Statistical Office of Lower Saxony (2016; see http://www.statistik.niedersachsen.de/themenbereiche/landwirtschaft/themenbereich-land--und-forstwirtschaft-fischerei---statistische-berichte-87592.html). The rental/sale price ratio (2010) is calculated based on rental price data from the Landesbetrieb für Statistik und Kommunikationstechnologie Niedersachsen (2010; see "Landwirtschaftszählung 2010. Heft 10: Eigentums- und Pachtverhältnisse, Pachtentgelte," available at http://www.statistik.niedersachsen.de/download/73728).

[a] Index points for the average soil quality refer to an official index in Germany that ranges from 7 to 104.

[b] Livestock units per hectare of agricultural land.

of the counties in this region is less than 0.5 livestock units per hectare, and the sale prices for agricultural land are rather low (around 20,000 €/ha) and experienced moderate price growth between 1990 and 2015 compared to the rest of Lower Saxony. The northern part of Lower Saxony, which borders the coast, is characterized by a low share of arable land

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(less than 50%). This region is dominated by dairy production and also has a large fruit growing area.

The western part of Lower Saxony is famous for its intensive livestock production. In light of rather poor soil quality (mostly around 30 soil quality points out of a total of 104) and relatively small farm sizes (50–60 ha on average), livestock production has comparative advantages, and its intensity has steadily increased over the last few decades. In fact, 70% of Lower Saxony's hog production and more than 80% of its poultry production are concentrated in the western part. More recently, biogas production has become an important alternative business in this region. The fact that 50% of Lower Saxony's total agricultural revenues are generated in its western part demonstrates the region's important role. Counties within the western part of Lower Saxony had the highest price growth rates between 1990 and 2015 (Emsland +304%, Cloppenburg +264%, and Oldenburg 259%) and the highest absolute land prices in 2015 (Vechta 90,457 €/ha, Cloppenburg 78,441 €/ha, and Emsland 61,723 €/ha). Potential reasons for the large differences in the price growth rates among counties within Lower Saxony are varying livestock intensities and production structures (Yang, Ritter, and Odening 2017), the use of biogas (Habermann and Breustedt 2011; Hennig and Latacz-Lohmann 2017) or wind energy (Ritter et al. 2015), nonagricultural investors (Forstner and Tietz 2013), and urban sprawl (Delbecq, Kuethe, and Borchers 2014; Ritter et al. 2015). The counties Emsland, Cloppenburg, Vechta, and Oldenburg are candidates for the choice of a dominant county in our model since these counties have the highest absolute prices and price growth rates.

The data applied for the estimation of price diffusion models are based on records of individual land sale transactions for arable land from January 1985 to December 2015. The raw data are provided by the committee of evaluation experts in Lower Saxony (Oberer Gutachterauschuss für Grundstückswerte in Niedersachsen),[5] which records all land transactions within Lower Saxony. Besides the price of each sold lot, the data set contains soil quality as a yield index (Ertragsmesszahl) and the size of the lot in square meters. This data set is used to build a balanced panel of quarterly average county prices. There are two reasons why we do not simply average transactions prices. First, the number of transactions per county per year is rather low, so raw observations might not represent the long-term values of plot size and soil quality. Second, differences in soil quality between the counties rule out the absolute convergence of prices, a priori. Hence, we adjust transaction prices to the same overall soil quality and plot size based on a hedonic regression, which is a standard approach when analyzing price convergence for heterogeneous goods (cf. Goldberg and Verboven 2004; Waights 2018). More precisely, we conduct the following steps:

1. We first run a hedonic regression based on the transaction data to quantify the effect of plot size and soil quality according to the following equation:

$$\ln p_k = u_i + \delta soil_k + \rho size_k + \sigma_i trend_k + \varepsilon_k. \qquad [9]$$

   Specifically, we model the logarithm of the transaction price per hectare ($\ln p_k$) as a linear function of the individual soil quality ($soil_k$) and plot size ($size_k$). To reduce the risk of omitted variable bias in the estimation of the coefficients δ and ρ, county-specific fixed effects (u_i) and county-specific linear time trends ($trend_k$) are included. ε_k denotes the error term. The index k corresponds to transactions, and i corresponds to counties.
2. After estimating equation [9] with all observations, we exclude transactions in which the residuals exceed four standard deviations of the empirical distribution of all residuals. These observations are considered outliers, since their prices strongly deviate from the expected price given their soil quality, size, location, and time. Then, we reestimate equation [9] for the cleaned data set. The results of this hedonic regression are presented in Appendix Table A1.
3. The coefficients of soil quality δ and size ρ are found to be statistically significantly

[5] See www.gag.niedersachsen.de/gutachterausschuesse/.

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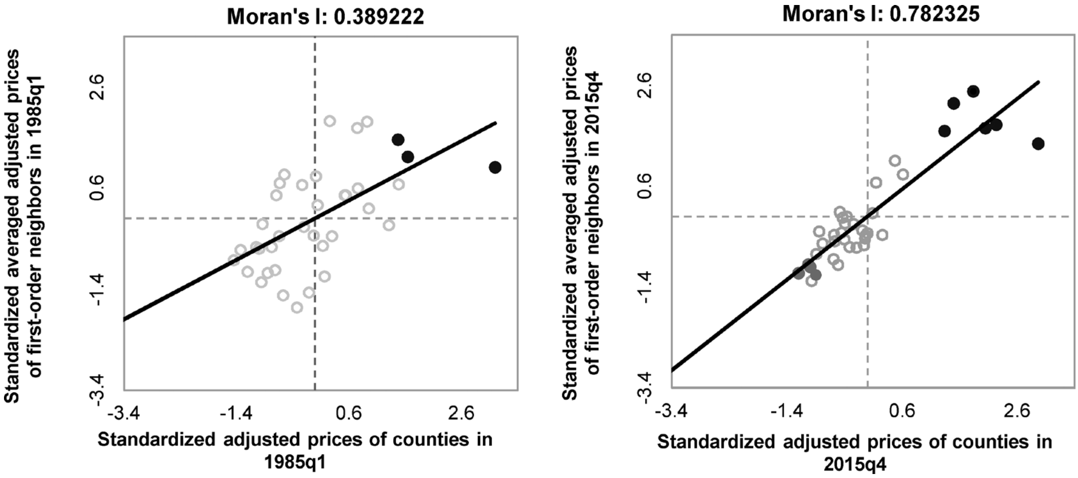
Figure 2

Moran's *I* Indices and Scatter Plots of Adjusted Prices; Black (Gray) Filled Dots Are Counties in a High-High (Low-Low) Cluster according to the Local Moran's *I*



different from zero at the 5% significance level. We use these coefficients to adjust individual prices to the overall averages of soil quality ($\overline{soil}$), and plot size ($\overline{size}$), according to the following equation:

$$(\ln p_k)' = \ln p_k - \hat{\delta}(soil_k - \overline{soil}) - \hat{\rho}(size_k - \overline{size}), \quad [10]$$

where $\hat{\delta}$ and $\hat{\rho}$ denote the estimated coefficients of soil quality and size, respectively, from the reestimated hedonic model (see Appendix Table A1).

4. Finally, adjusted transaction prices are averaged for each quarter t and each county i. Since larger plots rather than smaller plots are more representative of the whole county, the weighted average of adjusted prices by plot size is calculated in the following way:

$$(\ln p_{it})' = \frac{\sum_{k=1}^{Q_{it}} size_k (\ln p_k)'}{\sum_{k=1}^{Q_{it}} size_k}, \quad [11]$$

where $(\ln p_{it})'$ denotes the average adjusted price in county i in quarter t, and Q_{it} denotes the total number of observations in county i in quarter t. In case of missing values for some counties in some quarters, we linearly interpolate and fill longer gaps with annual data from the statistical office of Lower Saxony.[6]

The above procedure results in a balanced panel data set of 4,588 quarterly observations, which forms the basis of our analysis.[7]

To give a first impression of the spatial relationship of land prices, we apply Moran's I on the adjusted log prices in the first and last quarter of the observation period. The p-values significantly reject the null hypothesis that prices are randomly distributed in the study area, indicating spatial autocorrelation of land prices in Lower Saxony in the first and last quarters. As Figure 2 shows, Moran's I is positive for both quarters, indicating that a higher (lower) price in a county is usually linked to higher (lower) prices in the neighboring counties. Since the value for the last quarter is larger than that for the first quarter, this relationship has increased over time. Moreover, the local Moran's I (calculated according to Anselin 1995) depicts that Cloppenburg, Vechta, and Osnabrück form a high-high cluster in 1985q1, as well as in 2015q4, together with Emsland, Oldenburg, and Bentheim. This once again demonstrates

[6] See https://www.statistik.niedersachsen.de/themenbereiche/preise_verdienste/kaufwertestatistiken/.

[7] The data set is available from the authors upon request (matthias.ritter@agrar.hu-berlin.de).

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that these four counties are candidates for the dominant county.

4. Results

Before we turn to the results of the cointegration analysis and the price diffusion models, we inspect the long-run behavior of the adjusted log prices of arable land. Most counties in Lower Saxony show a clear upward trend over the observed period, and augmented Dicky-Fuller tests on the individual quarterly price series cannot reject the nonstationary price development for 33 out of 37 counties at the 5% significance level. Due to the rather low power of the univariate augmented Dicky-Fuller tests, counties that exhibit a stationary price development (Goslar, Leer, Osterholz, and Wolfenbüttel) are not excluded from the subsequent cointegration analysis.

To investigate whether a long-run equilibrium exists between prices in a county and the average prices of its neighbors, which is a precondition for the error correction term in the price diffusion model [4], we test their pairwise cointegration. Table 2 presents the results of the Johansen tests. The trace statistics clearly reject the null hypothesis of no cointegrating relationships for all counties. It is not surprising that there is a long-run connection between land prices in neighboring counties, since neighbors often have similar natural conditions and production structures. Thus, economic factors causing a change of land values in one county, such as new technologies, subsidies, or increased demand for land by financial investors, will likely affect neighboring counties as well.

Cointegration is necessary but not sufficient to establish the convergence of land prices among neighboring counties. To verify that county prices and their average neighbor prices actually converge, we need to further detect whether prices are cotrending and whether the cointegrating vectors $(1,-\beta_i)$ are equal to $(1,-1)$ (Abbott and de Vita 2013). According to Table 2, there are 14 counties in which land prices converge with their average neighbor price. However, this does not imply that prices approach the same level in the long run, which would be in contrast to the rather heterogeneous price paths reported in Table 1. First, one has to recall that we are analyzing adjusted (homogenized) prices, while Table 1 displays actual prices. Second, for 7 of the 14 counties, we find a significant coefficient δ_i in Table 2. This indicates that adjusted land prices among neighbors are equal up to a con-

Table 2

Pairwise Cointegration Tests with Neighbors

County	Trace Statistic	Cointegrating Vector $\hat{\beta}_i$	Constant $\hat{\delta}_i$
Ammerland	34.67***	0.901***+	0.962*
Aurich	54.19***	0.686***	2.802***
Bentheim	43.26***	0.757***	2.603***
Celle	41.78***	0.750***	2.282***
Cloppenburg	31.22***	1.121***	–0.956**
Cuxhaven	41.84***	0.771***	2.042***
Diepholz	45.47***	1.151***	–1.654***
Emsland	52.70***	1.035***+	–0.254
Friesland	50.30***	0.768***	2.071***
Gifhorn	45.25***	1.344***	–3.519***
Goslar	35.64***	0.443***	5.353***
Göttingen	38.73***	0.506***	4.712***
Hamelin-Pyrmont	53.74***	0.616***	3.721***
Hanover Region	38.78***	0.831***+	1.929***
Harburg	45.13***	0.739***	2.558***
Heidekreis	57.45***	1.104***+	–1.162***
Helmstedt	49.67***	0.656***	3.095***
Hildesheim	38.98***	0.795***+	2.076***
Holzminden	65.43***	0.762***+	2.003*
Leer	42.74***	0.896***+	0.881
Lüchow-Dannenberg	35.52***	0.857***+	1.033*
Lüneburg	36.99***	0.792***	1.876***
Nienburg	38.89***	1.174***	–1.767***
Northeim	40.85***	0.857***+	1.284
Oldenburg	35.66***	1.026***+	–0.355
Osnabrück	50.43***	0.830***	1.704***
Osterholz	48.54***	0.634***	3.618***
Osterode	40.60***	0.990***+	–0.413
Peine	39.55***	0.591***	4.129***
Rotenburg	46.13***	1.247***	–2.335***
Schaumburg	32.41***	0.522***	4.581***
Stade	46.55***	1.067***+	–0.547
Uelzen	78.55***	0.878***+	1.267*
Vechta	60.88***	0.866***	1.761***
Verden	38.64***	0.862***	1.258***
Wittmund	43.76***	1.075***+	–0.815
Wolfenbüttel	67.25***	0.658***	3.292***

Note: The trace statistic for testing $H_0{:}r = 0$ vs. $H_1{:}r \geq 1$ was estimated with unrestricted intercepts and restricted trend coefficients; r denotes the number of cointegrating vectors.

*, **, *** Significance at the 90%, 95%, and 99% levels, respectively.

+ $\hat{\beta}_i$ is not significantly different from 1 at the 99% significance level.

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stant in the long run, that is, the relative "law of one price" holds. Since we analyze log prices, this means that absolute prices have a constant ratio, that is, they change with the same rate.

We now proceed to the results of the price diffusion model [4]. Since contemporaneous terms $\Delta\overline{p}_{it}$ are included in the model, we test whether this term is weakly exogenous. The Wu-Hausman test statistic is the t-value for testing H_0:$\lambda_i = 0$ in the augmented regression:

$$\Delta p_{it} = \gamma_i + \sum_{l=1}^{L_{ia}} a_{il}\Delta p_{i,t-l} + \sum_{l=0}^{L_{ib}} b_{il}\Delta\overline{p}_{i,t-l} + \phi_i(p_{i,t-1} - \delta_i - \beta_i\overline{p}_{i,t-1}) + \lambda_i\hat{\varepsilon}_{0t} + \varepsilon_{it}, \qquad [12]$$

where $\hat{\varepsilon}_{0t}$ denotes the residuals of the average neighbor price $\Delta\overline{p}_{it}$ regressed by the instrumental variables ($\Delta\overline{p}_{i,t-L_{ib}-h}$, $\Delta\overline{p}^{s}_{i,t-d}$). $\Delta\overline{p}_{i,t-L_{ib}-h}$ with $h \in \mathbb{N}$ are the temporally lagged average prices that were not already included in the model. $\Delta\overline{p}^{s}_{i,t-d}$ with $d \in \mathbb{N}_0$ are the (temporally lagged) average prices of the second-order neighbors of county i. The error correction coefficients in equation [12] are restricted as described above. The lag numbers h and d are chosen so that the instruments pass the Sargan-Hansen test of overidentifying restrictions for validity and an F-test for the strength of the instruments at the 5% significance level.[8] If the Wu-Hausman test rejects H_0 at the 95% significance level, the variables $\Delta\overline{p}_{i,t-L_{ib}-h}$ and $\Delta\overline{p}^{s}_{i,t-d}$ are used as instrumental variables for $\Delta\overline{p}_{it}$. According to Table 3, this is the case only for Goslar. Moreover, the Breusch-Pagan test rejects error independence at the 95% significance level for all counties with exogenous contemporaneous terms, so we use seemingly unrelated regressions to estimate the system of price equations.

The estimation results of the price diffusion model [4] are shown in Table 3. The adjustment speed coefficients $\hat{\phi}_i$ are all significant and negative, which indicates that land prices move toward the long-run equilibrium with the average neighbor price. The adjustment coefficients amount to 67% per quarter on average, which is rather slow compared to agricultural commodity markets that have an adjustment speed that is usually greater than 90% per quarter (e.g., Wang and Tomek 2007). This finding is not surprising since land is immobile and economic equilibria cannot simply be attained by trading and transport. Adjustment processes in the land market are more complex and are sometimes related to the diffusion of new technologies. For example, Hennig and Latacz-Lohmann (2017) show that the boom in biogas plants has led to an increase in land rental prices in Germany. Moreover, land markets are less liquid compared to commodity markets. Moreover, information on price changes is processed more slowly in land markets. Note that there is regional variation in the adjustment speeds. Smaller absolute values imply a lower impact from the average neighbor price. We find that the five counties with the smallest adjustment speeds (Lüchow-Dannenberg, 24%; Aurich, 31%; Cuxhaven, 43%; Cloppenburg, 43%; and Göttingen, 44%) are located adjacent to the state border, with the exception of Cloppenburg. Whereas border counties could also be affected by price development beyond the border, which is not considered in our analysis, the slow adjustment of prices in Cloppenburg to its average neighbor price could indicate that Cloppenburg is a dominant county.

Regarding the short-term development of land prices, we find that most of the values for the own lagged effects and some of the neighbors' lagged effects are significant. Most of these effects are negative, which means that short-term deviations are compensated in later periods. About 80% of the counties have a significant and positive coefficient b_{i0} for the neighbors' contemporaneous effects, such that land price changes in one county will immediately spill over to adjacent counties. Economic drivers of these price changes include subsidies or regulations that affect land prices in neighboring counties simultaneously.

To summarize, the evidence for static spatial autocorrelation of land prices within Lower Saxony, which we found from Moran's I, is confirmed in a dynamic context by our price diffusion model.

To put more structure on the price diffusion process, we now examine whether land

[8] In some cases, several lags had to be combined to obtain strong instruments at the 5% level. In four cases, we could not obtain strong instruments at the 5% level.

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Table 3

Estimation Results for the Price Diffusion Equations with Neighboring Counties

County	Adjustment Speed	Own Lagged Effects	Neighbors' Lagged Effects	Neighbors' Contemporaneous Effect	Wu-Hausman Test
Ammerland	–0.714***	–0.144*	–0.219*	0.423***	0.00
Aurich	–0.310***	–0.426***	–0.025	0.192***	1.68
Bentheim	–0.557***	–0.365***	–0.057	0.388**	1.11
Celle	–0.704***	–0.369***	0.101	0.754***	0.57
Cloppenburg	–0.426***	–0.181**	0.004	0.271**	1.49
Cuxhaven	–0.425***	–0.513***	–0.203**	0.177**	0.59
Diepholz	–0.760***	–0.236**	–0.127	0.325***	0.84
Emsland	–0.549***	–0.166*	–0.231***	0.232***	1.38
Friesland	–0.702***	–0.135*	–0.175	0.410***	0.58
Gifhorn	–0.472***	–0.233**	0.023	0.510***	1.56
Goslar	–0.770***	–0.102	0.588*	1.264*	5.69**
Göttingen	–0.433***	–0.368***	–0.015	0.383***	0.57
Hamelin-Pyrmont	–0.722***	–0.183**	—	0.311***	2.03
Hanover Region	–0.764***	–0.151**	—	0.731***	0.44
Harburg	–0.718***	–0.122*	0.305*	0.562***	0.02
Heidekreis	–0.957***	0.154*	–0.405***	0.262**	0.02
Helmstedt	–0.811***	–0.062	–0.114	0.304***	0.72
Hildesheim	–0.573***	–0.068	—	0.157	0.25
Holzminden	–0.912***	–0.180*	—	0.732***	0.04
Leer	–0.710***	—	–0.153	0.842***	0.19
Lüchow-Dannenberg	–0.241***	–0.245***	—	0.088*	0.96
Lüneburg	–0.718***	–0.394***	–0.245*	0.128	1.31
Nienburg	–0.576***	–0.227***	–0.115	0.341***	0.22
Northeim	–0.556***	–0.289***	—	0.291***	0.05
Oldenburg	–0.674***	–0.300***	–0.069	0.387***	0.00
Osnabrück	–0.622***	–0.340***	0.158	0.449***	0.03
Osterholz	–0.871***	–0.178**	–0.513**	–0.172	0.91
Osterode	–0.576***	–0.277**	–0.134	0.313***	0.61
Peine	–0.744***	–0.138	–0.073	0.171	0.91
Rotenburg	–0.443***	–0.351***	–0.155	0.251**	0.31
Schaumburg	–0.774***	–0.151*	–0.048	0.208*	0.33
Stade	–0.685***	–0.170*	0.293**	0.575***	3.19*
Uelzen	–1.247***	–0.158	–0.498**	0.519***	0.14
Vechta	–0.887***	–0.114	–0.270	0.435**	2.17
Verden	–0.519***	–0.275***	–0.089	0.146	2.33
Wittmund	–0.657***	–0.167**	—	0.603***	0.44
Wolfenbüttel	–0.914***	—	–0.446***	–0.127	0.16

Note: The lag order for each county is selected separately using the Bayesian information criterion using a maximum lag order of four. The reported coefficient for the lagged effects is the value with the lowest *p*-value. "—" denotes that the lag order equals zero. All regressions include an intercept term.

*, **, *** Significance at the 90%, 95%, and 99% levels, respectively.

prices in Lower Saxony are driven not only by prices in neighboring counties, but also by a dominant county. In this case, land price changes would be unidirectional and ripple out from a dominant county to other counties. In contrast to studies in the housing market, in which large metropolitan areas are a natural candidate for a dominant county, it is not obvious where such a county—if it exists at all—is located in the agricultural land market of Lower Saxony. To select a potentially dominant county, we proceed as follows: First, we focus on counties showing the highest land price level and the most pronounced price increase during the observation period. According to Table 1, these are Vechta, Cloppenburg, Emsland, and Oldenburg—counties that are characterized by intensive livestock production. Next, we estimate the pairwise error correction models (equations [5] and [6]) for

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these four counties with all other counties in Lower Saxony. We expect a dominant county to have significant adjustment speeds ϕ_{i0} and to not have a reverse effect, that is, the values of ϕ_{0i} are not significant. Appendix Table A2 reveals that Cloppenburg and Oldenburg pass this test (with one exception in each case). Recalling the previous finding that Cloppenburg has a slow adjustment speed in the price diffusion model [4], we finally select this county as the most suitable candidate for a dominant county.

Appendix Table A3 depicts the results of the cointegration test for the extended model [8], which allows for joint effects of neighbors and a dominant county. We observe that the coefficient β_{i0} in the cointegrating vector is significant in most but not all cases, meaning that Cloppenburg contributes significantly to the joint long-run equilibrium. Counties that are influenced by their neighbors but not by Cloppenburg are either remote from Cloppenburg (Goslar and Uelzen), adjacent to the Netherlands (Leer and Bentheim), or show a very similar production structure (Vechta), so that it remains unclear which county leads or follows in the price diffusion process.

Estimation of the diffusion model [8] follows the same procedure as before. Since contemporaneous terms Δp_{0t} are included in the model, we test for endogeneity using the Wu-Hausman test with instrumental variables $(\Delta p_{0,t-L_{ic}-h}, \Delta\bar{p}_{0,t-d})$. The lags $h \in \mathbb{N}$ and $d \in \mathbb{N}$ are again chosen so that the instruments are valid and strong at the 5% level, which could be achieved in all cases. The results for the Wu-Hausman test and the estimation of model [8] are provided in Appendix Table A4. Again, the coefficients $\hat{\phi}_i$ are all significant and negative, which implies a correction toward a long-term equilibrium with neighbors and the dominant county.

It is, however, difficult to disentangle this effect and to separate the contribution of Cloppenburg. Comparing the results with the previous model [4] shows that the inclusion of Cloppenburg has increased the absolute value of the coefficients $\hat{\phi}_i$ on average, in other words, the observed adjustment is faster in the model with Cloppenburg as a dominant county. Regarding the short-run effects, we find that Cloppenburg has a significant impact on land prices in only a few counties. In addition, spillover effects can be measured for some neighboring counties such as Emsland and Oldenburg. Overall, the contemporaneous effects of neighbors seem to be more relevant. We conclude that Cloppenburg cannot be clearly characterized as a strong dominant county, and ripple effects are less pronounced in land markets compared with real estate markets. This finding can be explained by differences in the underlying economic mechanisms, which drive the price diffusion process. In housing markets, migration plays a central role in the emergence of ripple effects, whereas in the case of farmland, ripple effects rely on the mobility of farmers, which is restricted by transport costs, as well as natural and legal conditions.

5. Conclusions

Politicians and other stakeholders in agriculture are concerned about the recent surge in farmland prices observed in many parts of the world. While this price increase is rather unambiguous at the aggregate level, the development of land prices is more subtle and differentiated at the regional level. Our case study from Germany documents that land prices may grow at different rates even within a country or state. Notably, despite extant empirical work on explaining the determinants of farmland price levels, detailed analyses on the spatial development of land prices at a regional level are rare. We contribute to this research by employing a price diffusion model that combines features of market integration models and spatial econometric models. This approach identifies long-run equilibrium relationships among local land markets and separates short- and long-run price transmission. An application of this model to farmland prices in the state of Lower Saxony shows that prices at a county level are in fact cointegrated, as they are linked by long-run equilibria. However, this does not imply that land prices in all counties necessarily converge to the same level or a constant difference even after adjusting for land quality differences. This

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result confirms earlier findings by Yang, Ritter, and Odening (2017) that local land markets may exhibit distinct convergence clubs. Not surprisingly, the adjustment rates that we measure are smaller compared to commodity markets and are similar to those of other real estate markets. In some cases, apart from convergence toward a long-run equilibrium, we find that price transmission also takes place through short-term adjustments caused by prices changes in neighboring counties. A modification of the price diffusion model allowed us to examine whether some counties dominate others in the sense that price diffusion is unidirectional, namely, that price shocks spill over from a dominant county to neighboring counties, but not vice versa. We found that Cloppenburg, a center of intensive livestock production in Germany, actually mimics some of these behaviors.

Our findings are relevant from a policy perspective for at least two reasons. First, our results assert that land markets are spatially integrated, a feature that is generally considered as being indicative of market efficiency. Market inefficiency, on the other hand, is often used as an argument to justify market regulation. In this regard, our results do not provide evidence for land market regulations. Second, in view of the recent land price surge, many E.U. countries have implemented price monitoring systems to increase transparency of price formation in farmland markets. Our results support this task, since knowledge of this diffusion process can be useful to predict how price changes in local land markets will affect neighboring regions. More specifically, the estimated price diffusion model can be used to derive impulse response functions that allow an assessment of the spatial land price impact that is caused by a price shock in a specific area (e.g., Helgers and Buyst 2016). This is particularly useful for understanding whether the price impact of a regional policy or land market intervention is bounded.

So far, our analysis targets the identification of patterns in farmland price diffusion. It is, however, rather silent about the economic forces that cause these patterns. Thus, a natural step toward a more comprehensive understanding of the spatial dynamics of land prices would be the inclusion of common factors into the price diffusion model, such as interest rates, land rental prices, or income variables. Furthermore, the consideration of structural variables that characterize local economic activities, such as farm exit rates, farm size, or production intensity, could enhance our empirical analysis.

Acknowledgments

Financial support from the China Scholarship Council (CSC No. 201406990006) and the Deutsche Forschungsgemeinschaft (DFG) through Research Unit 2569 “Agricultural Land Markets—Efficiency and Regulation” is gratefully acknowledged. The authors also thank Oberer Gutachterausschuss für Grundstückswerte in Niedersachsen (Peter Ache) for providing the data used in the analysis.

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