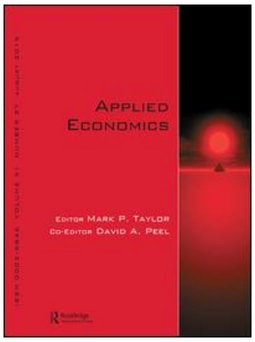

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Land price diffusion across borders – the case of Germany

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ABSTRACT

Land market regulations are often justified by the assumption that activities of foreign and non-agricultural investors drive up prices in domestic land markets. However, empirical knowledge about the dynamics of agricultural land prices across borders is sparse. Using the German reunification as a natural experiment, we study the effect of the former inner German border on the dynamics of agricultural land prices in East and West Germany. We apply a land price diffusion model with an error correction specification to analyse spatial agricultural land markets. A novel feature of our model is its ability to distinguish price diffusion within states and across state borders. We provide evidence for a persistent border effect given that the fraction of spatially integrated counties is larger within states than across the former border. Moreover, we observe non-significant error correction terms for many counties along the former border. From a policy perspective, it is striking to realize that even 25 years after German reunification, pronounced land price differences persist. It is quite likely that price diffusion through existing borders within the EU would take even more time given language barriers, different institutional frameworks, and information asymmetries between domestic and foreign market participants.

KEYWORDS

Agricultural land markets; price diffusion; spatial dependence; border effect

JEL CLASSIFICATION

Q13; Q15

I. Introduction

Recent surges in agricultural land prices and ongoing changes in land use due to urban sprawl, renewable energy production, and growing demands from non-agricultural investors have triggered debates on the effectiveness of existing land market regulations. Although boom and bust cycles are not new to land markets, current changes in the market are considered to result from a new constellation of driving forces. For instance, it is conjectured that the increased demand for land by financial investors has increased land rental and sales prices. These developments have led to demands for stricter regulations of land markets in many countries, including developed countries (cf. Kay, Peuch, and Franco 2015). In 2010, the UK Government Office of Science stressed the need to balance competing pressures on land use and to roll out new land-use policies (Government Office for Science 2010). Four years later, Belgium laid the foundation for new land market instruments, such as a land observatory, land bank, and updated preemption rights. Belgium also tightened land market regulations, which had previously been liberal. Likewise, in Germany, the Federal Ministry and the State Ministries of Agriculture currently aim for a broad distribution of land ownership, the prevention of dominant land market positions on the supply and demand side, the capping of land rental and sales prices, prioritizing agricultural use of farmland, and establishing greater transparency for land markets (Bund-Länder-Arbeitsgruppe "Bodenmarktpolitik" 2015). Although these goals are fairly general, they fall in line with the trend towards stricter land market regulations. The proposed measures envision restricting market access for actors who treat land as an investment asset and do not have farming interests, while simultaneously prioritizing land purchases by farmers and facilitating farm succession and start-ups.

Remarkably, it is mainly the new EU Member States, which carry the legacy of weaker land market institutions from their socialist past, that opt for particularly strong regulations (cf. Swinnen, van Herck, and Vranken 2016). For example, new land market regulations aiming to restrict the purchase of agricultural land by foreigners and non-farmers was released in Slovakia (Lazíková and Bandlerová 2015). In 2016, Poland passed the *Act on the Structuring of the Agricultural System*, which postponed exemptions from EU laws regarding the

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acquisition of land. The bill proposes to stop the sale of state-owned land for the next five years and includes very strict rules on who can sell and buy privately owned land. The objective of the new law is to ensure that farmland remains in the hands of Polish farmers after the transition period. Bulgaria, Hungary, Latvia, and Lithuania followed suit with regulations directly or indirectly restricting the free movement of capital and freedom of establishment.

Most of the aforementioned attempts to regulate land markets have been motivated by the apprehension that in countries with low land price levels, farmers will encounter a drastic price surge unless land markets are protected against demand by foreign and non-agricultural investors. This assumption, however, lacks empirical evidence. Little is known about the spatio-temporal behaviour of agricultural land prices and virtually no empirical study exists that investigates the diffusion of agricultural land prices across borders. In other words, we do not know if and how fast land prices in two neighbouring countries with different price levels would converge if there were no restrictions on the acquisition of land. The main objective of this paper is to address this research gap. Our empirical analysis is conducted for West and East Germany, i.e., we study the effect of the former intra-German border on the dynamics of agricultural land prices. The German reunification constitutes a natural experiment on the establishment and evolution of land markets that allows us to study market integration. It is well known that a gap exists between land prices in West and East Germany, but little is known about how this gap evolves over time and if the same land price dynamics prevail in both parts of Germany. After reunification, regions in Western Germany (especially near the former border between West and East Germany) lost their remoteness since they were suddenly situated in the centre of Europe and thus became more attractive. On the other hand, the supply of cheaper land increased and redirected demand to regions in Eastern Germany, so that the effect of the reunification on land prices remains unclear.

To the best of our knowledge, there are only few studies that test for spatial market integration in the context of agricultural land.[1] Carmona and Roses (2012) investigate the spatial integration of Spanish land markets between 1904 and 1934 from a historical perspective. Their analysis is based on aggregated data and does not take into account heterogeneity of land characteristics and structural breaks in the price series that may bias the test results. More Yang, Ritter, and Odening (2017) explore the spatial pattern of land price development. Based on county-level data for the German state Lower Saxony, they employ stationarity tests and unit root tests to examine whether relative prices between counties converge. Using a sequential testing procedure allows Yang, Ritter, and Odening (2017) to identify several distinct convergence clusters. The closest study to ours investigates the impact of a language border on spatio-temporal price diffusion of house prices in Belgium (Helgers and Buyst 2016). Starting with a pairwise approach to provide insight into the degree of integration among housing prices, the study estimates a bivariate VAR model with error-correcting coefficients. The results indicate that the fraction of pairs for which the regional house price differentials are stationary is higher within a linguistic area than between these areas. Although there are many structural similarities between house markets and land markets, which allow the transfer of methods across these two fields, one should also recognize the differences between these two markets. First, while agricultural land is mainly a production factor, houses have the character of a consumption good. This makes house prices more dependent on buyers' preferences and incomes. Second, potential buyers of houses are usually more mobile than farmers, making it more likely for house prices to converge. Finally, the supply of land follows a different mechanism than the supply of houses. Thus, one cannot readily adopt findings from real estate markets to agricultural land markets.

The remainder of this paper is organized as follows: The following section introduces the spatial price diffusion model and explains the logic of identifying a 'border effect'; Section 3 provides some background information about the study region, the relevant land market environment after reunification, and the derivation of the data; Section 4 presents and discusses the empirical results; and Section 5 concludes.

[1]There is, however, a rich literature on spatial price convergence in real estate markets, particularly in housing markets (cf. Hiebert and Roma (2010) for an overview).

II. A land price diffusion model with a border effect

At the heart of our research lies the question of whether land prices in Germany are integrated through time and space and converge in absence of barriers, such as the former German border. Consequently, the desired empirical application requires a model that allows for the incorporation of time and space. This can be achieved by a price diffusion model as proposed by Holly, Pesaran, and Yamagata (2011) and applied by Gong, Hu, and Boelhouwen (2016). In general, a price diffusion model is based on a Vector Error Correction Model (VECM) since cointegration is a necessity for price convergence in the long-run. The VECM accounts for this cointegration relationship by correcting the short-run responses of prices by deviations from a stable long-run equilibrium.

At first glance, to test the integration of land prices in a study area consisting of N regions would imply to test for $N(N-1)/2$ cointegration relationships. Nevertheless, one price of a cointegration vector can always be expressed by one other price or a combination of cointegrated prices (Holly, Pesaran, and Yamagata 2011). Thus, it is feasible to apply a neighbour approach that reduces the rank of the cointegration vector to unity and the number of equations to be estimated to N. In this parsimonious representation, the cointegration relationship is reduced to the price $p_{i,t}$ of region i and a weighted average price of region i's neighbours j, $\bar{p}_{i,t}^{\text{neighbor}} = \sum_{j=1}^{N} w_{ij}p_{j,t}$, with $\sum_{j=1}^{N} w_{ij} = 1$ if row-standardization is applied. The weights w_{ij} measure connectivity through proximity in geographic, economic, or social terms. Stacking all of the weights in a matrix with the diagonal elements equal to zero gives a spatial weight matrix W, which incorporates the dimension of space into the model. Another benefit of this approach is that no benchmark region has to be selected a priori in the cointegration system (Abbott and de Vita 2013). The regional price diffusion model can be formulated into a VECM:

$$\Delta p_{i,t} = c_i + \phi_i ECT_{i,t-1} + \sum_{k=1}^{K} \gamma_{i,1,k}\Delta p_{i,t-k} + \sum_{l=1}^{L} \gamma_{i,2,l}\Delta \bar{p}_{i,t-l}^{\text{neighbor}} + \lambda_i z_t + \varepsilon_{it}, \quad (1)$$

where $p_{i,t}$ is the land price in region i at time t, $\bar{p}_{i,t}^{\text{neighbor}}$ is a weighted average land price in neighbouring regions, z_t is a vector of exogenous common factors that affect all region prices, ε_{it} is an error term, and Δ is the difference operator. The term c_i is a region-specific constant term to capture unobserved individual effects. The parameter vectors $\gamma_{i,1,k}$ and $\gamma_{i,2,l}$ capture the short-run responses of $\Delta p_{i,t}$ to K own price lags and L weighted neighbour price lags, respectively. λ_i capture contemporaneous responses to the common factors. ϕ_i measures the adjustment speed of corrections given that random deviations $ECT_{i,t-1}$ in the long-run equilibrium relationship between land prices occur. Error correction requires ϕ_i to be negative. A flexible form of the cointegration relationship that includes a constant and a trend is given by

$$ECT_{i,t-1} = p_{i,t-1} - \beta_{0i} - \beta_{1i}\bar{p}_{i,t-1}^{\text{neighbor}} - \beta_{2i}t_i, \quad (2)$$

where βs are parameters defining the cointegration relationship between price pairs. Note that the error correction term $ECT_{i,t-1}$ incorporates the spatial dimension in the long-run relationship through the neighbouring prices. The error correction term incorporates the spatial lag of $p_{i,t}$ and equals the spatial autoregressive cointegration vector of a Spatial Error Correction Model (Beenstock and Felsenstein 2010). While cointegration is sufficient to establish a long-run price relationship, further parameter restrictions have to be fulfilled to assert prices convergence among neighbouring regions (Abbott and de Vita 2013; Yang, Ritter, and Odening 2017). If β_{1i} equals unity and the trend parameter β_{2i} and constant β_{0i} equal zero, prices of neighbouring regions converge to the same level (absolute convergence). If β_{0i} is instead positive, prices converge towards a constant difference (relative convergence) (Waights 2018).[2]

[2]Note that our approach of measuring convergence through a co-integration analysis is closely related to the concept of β-convergence often used in empirical macroeconomic growth theory. In this framework, β-convergence is quantified as partial correlation between growth rate and the initial income level of countries or regions (e.g., Barro and Sala-i-Martin 1992). A negative relationship implies a catch-up of weaker regions and its size relates to the speed of adjustment towards a steady state income level. A similar interpretation holds for ϕ in our diffusion model, though the notion of an equilibrium is less specific. For a link between β-convergence and σ-convergence, we refer to Young, Higgins, and Levy (2008).

To examine whether a predetermined barrier, such as a border, affects the diffusion of prices, we follow Helgers and Buyst (2016) by splitting neighbouring prices into two groups. One group is the weighted price consisting of regions on the same side of the border, $\bar{p}_{i,t}^{\text{same}} = \sum_{j=1}^{N} w_{ij}^{\text{same}} p_{j,t}$, and the other group of regions on the opposite side of the border, $\bar{p}_{i,t}^{\text{opp}} = \sum_{j=1}^{N} w_{ij}^{\text{opp}} p_{j,t}$. Therein weights are based on the individual elements w_{ij} of the original weighting matrix W with the difference that the individual elements of w_{ij}^{same} (w_{ij}^{opp}) are set to zero if region j lies on the opposite (same) side of the border as region i. Again, the elements w_{ij}^{same} and w_{ij}^{opp} are standardized across j columns. In contrast to Helgers and Buyst (2016), we refrain from including a dominant region in the model since a dominant region is less likely to exist in agricultural land markets (Yang, Odening, and Ritter 2019). With this regrouping, the price diffusion model (2) is transformed into:

$$\Delta p_{i,t} = c_i + \phi_{i,1} ECT_{i,1,t-1} + \phi_{i,2} ECT_{i,2,t-1} + \sum_{k=1}^{K} \gamma_{i,1,k} \Delta p_{i,t-k} + \sum_{l=1}^{L} \gamma_{i,2,l} \Delta \bar{p}_{i,t-l}^{\text{same}} + \sum_{q=1}^{Q} \gamma_{i,3,q} \Delta \bar{p}_{i,t-q}^{\text{opp}} + \lambda_i z_t + \varepsilon_{it}. \quad (3)$$

Herein, $ECT_{i,1,t-1}$ captures deviations from the long-run relationship between region i's land price and the within state average neighbours' land price $\bar{p}_{i,t-1}^{\text{same}}$. Accordingly, $ECT_{i,2,t-1}$ corresponds to deviations from the across state neighbours' land price $\bar{p}_{i,t-1}^{\text{opp}}$. Equation (3) allows the empirical investigation of whether a border effect is present in land price diffusion. A border effect can exist under two different circumstances. The first is if deviations from the long-run equilibrium with the weighted average land price of neighbouring regions are not corrected ($\phi_{i,2} \geq 0$). The second is if deviations from the average weighted land price of neighbours within the same state are corrected faster than the average weighted land price of neighbouring regions across the border ($\phi_{i,1} < \phi_{i,2}$). This leads to the following hypotheses:

Hypothesis 1: The former border does not slow down the long-run price diffusion process of region i with neighbours across the border compared to neighbours within the state. Thus, deviations in the cointegration relationship with neighbours across the former border are corrected faster or at the same speed as with neighbours on the same side of the border ($\phi_{i,1} \geq \phi_{i,2}$).

Hypothesis 2: The former border prohibits any correction towards a long-run equilibrium between region i's land price and the land price of neighbouring regions across the border ($\phi_{i,2} \geq 0$).

If Hypothesis 1 is rejected, the former border still affects land price diffusion for region i with its neighbouring land markets across the border. If Hypothesis 2 is rejected, land price changes diffuse across the former border. Thus, we can deduce that if Hypothesis 1 is not rejected and Hypothesis 2 is rejected, land price diffusion to and from region i to its neighbours across the former border is not blocked or slowed down, i.e., there is evidence supporting no border effect. Vice versa, if Hypothesis 1 is rejected or Hypothesis 2 is not rejected, we can conclude that land price diffusion to and from region i to its neighbours across the former border is slowed down and possibly completely blocked, i.e., there is evidence supporting a border effect.

Assuming independence of the error terms, the N regional VECM Equations (3) can be estimated with Ordinary Least Squares (OLS). The seemingly unrelated regression (SUR) allows for the estimation of an unrestricted covariance matrix E_t with possible contemporaneous correlation between the individual region equations, $\mathrm{Cov}(\varepsilon_{it}, \varepsilon_{jt}) \neq 0$ for $i \neq j$. We apply an iterative SUR, which allows updating the covariance matrix in each iteration and converges to maximum likelihood (Greene 2002).

While the system of regional VECM equations is a parsimonious representation of N cointegration relationships and allows one to test whether the former German border still affects long-run land price diffusion, it cannot display the full complexity of the spatio-temporal land price diffusion process and restricts the analysis to regions adjacent to the former border. Regional land markets, however, can be linked over far distances and react to one another, even though no direct cointegration relationship exists due to short-run dynamics and temporal and spatial spillover effects. The price diffusion model in a VECM form is the basis for deriving impulse response function (IRF) specifications. Through impulse response analysis, it is possible to investigate the diffusion of shocks to one region in a regional system over time and space (Holly, Pesaran, and Yamagata 2011). To derive IRFs, the original system of N regional VECM equations with a border effect (3) is stacked and rewritten in matrix notation:

$$\Delta P_t = C + \Pi P_{t-1} + \sum_{l=1}^{L} \Gamma_l \Delta P_{t-l} + \Lambda Z_t + \mathrm{E}_t \quad (4)$$

with

$$C = \begin{bmatrix} c_1 + \phi_{1,1}\beta_{01,1} + \phi_{1,2}\beta_{01,2} \\ c_2 + \phi_{2,1}\beta_{02,1} + \phi_{2,2}\beta_{02,2} \\ \vdots \\ c_{N-1} + \phi_{N-1,1}\beta_{0N-1,1} + \phi_{N-1,2}\beta_{0N-1,2} \\ c_N + \phi_{N,1}\beta_{0N,1} + \phi_{N,2}\beta_{0N,2} \end{bmatrix},$$

$$\Gamma_R = \begin{bmatrix} \gamma_{1,1R} & 0 & \cdots & 0 & 0 \\ 0 & \gamma_{2,1R} & \cdots & 0 & 0 \\ \vdots & \vdots & \ddots & \vdots & \vdots \\ 0 & 0 & \cdots & \gamma_{N-1,1R} & 0 \\ 0 & 0 & \cdots & 0 & \gamma_{N,1R} \end{bmatrix} + \begin{bmatrix} \gamma_{1,2R} w_1^{\mathrm{same}'} + \gamma_{1,3R} w_1^{\mathrm{opp}'} \\ \gamma_{2,2R} w_2^{\mathrm{same}'} + \gamma_{2,3R} w_2^{\mathrm{opp}'} \\ \vdots \\ \gamma_{N-1,2R} w_{N-1}^{\mathrm{same}'} + \gamma_{N-1,3R} w_{N-1}^{\mathrm{opp}\,'} \\ \gamma_{N,2R} w_N^{\mathrm{same}'} + \gamma_{N,3R} w_N^{\mathrm{opp}'} \end{bmatrix}$$

The price vector $P_t = \left(p_{1,t}, p_{2,t}, \ \ldots \ , p_{N,t}\right)'$ comprises all N regions' land prices and thus all endogenous time series. Π is the $N \times N$ cointegration matrix to parameterize the long-run spatial relationship in P_t, while the $R \times N$ matrix Γ_R captures the short-run responses to R past changes in P_t.[3] The spatial weight vectors $w_i^{\mathrm{same}\prime} = \left(w_1^{\mathrm{same}'}, w_2^{\mathrm{same}'}, \ \ldots \ , w_{N-1}^{\mathrm{same}'}, w_N^{\mathrm{same}'}\right)$ and $w_i^{\mathrm{opp}\prime} = \left(w_1^{\mathrm{opp}'}, w_2^{\mathrm{opp}'}, \ \ldots \ , w_{N-1}^{\mathrm{opp}\,'}, w_N^{\mathrm{opp}'}\right)$ are the

$$\Pi = \begin{bmatrix} \phi_{1,1} + \phi_{1,2} & 0 & \cdots & 0 & 0 \\ 0 & \phi_{2,1} + \phi_{2,2} & \cdots & 0 & 0 \\ \vdots & \vdots & \ddots & \vdots & \vdots \\ 0 & 0 & \cdots & \phi_{N-1,1} + \phi_{N-1,2} & 0 \\ 0 & 0 & \cdots & 0 & \phi_{N,1} + \phi_{N,2} \end{bmatrix} - \begin{bmatrix} \phi_{1,1}\beta_{11,1} w_1^{\mathrm{same}'} + \phi_{1,2}\beta_{11,2} w_1^{\mathrm{opp}'} \\ \phi_{2,1}\beta_{12,1} w_2^{\mathrm{same}'} + \phi_{2,2}\beta_{12,2} w_2^{\mathrm{opp}'} \\ \vdots \\ \phi_{N-1,1}\beta_{1N-1,1} w_{N-1}^{\mathrm{same}'} + \phi_{N-1,2}\beta_{1N-1,2} w_{N-1}^{\mathrm{opp}\,'} \\ \phi_{N,1}\beta_{1N,1} w_N^{\mathrm{same}'} + \phi_{N,2}\beta_{1N,2} w_N^{\mathrm{opp}'} \end{bmatrix}, \text{and}$$

[3] R is the maximum of the lag numbers K, L, and Q of the lagged own and neighbours' price differences suggested by Schwarz Criterion (BIC).

N rows of the corresponding spatial weight matrices W^{same} and W^{opp}.

The vector autoregression (VAR) representation of (4) is

$$P_t = C + \Phi_1 P_{t-1} + \Phi_2 P_{t-2} + \ \ldots \ + \Phi_R P_{t-R} + \Phi_{R+1} P_{t-(R+1)} + \Lambda Z_t + \mathrm{E}_t,$$

where the parameter matrices $\Phi_1 = I_N + \Pi + \Gamma_1$, $\Phi_R = \Gamma_R - \Gamma_{R-1}$, and $\Phi_{R+1} = -\Gamma_R$ are compounds of the VECM coefficient matrices.

The generalized impulse response function (GIRF) g_i for a one unit (one standard error) shock originating in region i at h time step intervals ahead can be calculated after Pesaran and Shin (1998) by

$$g_i(h) = \frac{\Psi_h \Sigma e_i}{\sqrt{\sigma_{ii}}} \text{for} h = 0, 1, \ \ldots \ , H, \tag{6}$$

where Σ is the covariance matrix, e_i is a $N \times 1$ vector of zeros with exclusion of its i^{th} element set to unity, and σ_{ii} are the diagonal elements of the covariance matrix. The Ψs are calculated recursively with the help of the VAR coefficients by

$$\Psi_h = \Phi_1 \Psi_{h-1} + \Phi_2 \Psi_{h-2} + \ \ldots \ + \Phi_R \Psi_{h-R} + \Phi_{R+1} \Psi_{h-(R+1)}, \tag{7}$$

with $\Psi_0 = I_N$ and $\Psi_h = 0$ for all $h < 0$ (Pesaran and Shin 1998). The GIRF approach is a better representation of dynamic spatial integration since a shock originating in region i will eventually progress to the non-neighbouring region j via spatial linkage through other regions (Abbott and de Vita 2013).

III. Study region and data

The border region of lower Saxony and Saxony-Anhalt

During the division of Germany from 1949 to 1990, the two sides divided by the inner German border were exposed to different political and economic systems. This difference also applied to agricultural land markets. Whereas a free land market was established in West Germany, East Germany was characterized by expropriation and collectivization of land. In 1989, East German agriculture consisted of 464 state-owned farms called *Volkseigene Güter* (VEGs, People-Owned Properties) and 3,844 collective farms called *Landwirtschaftliche Produktionsgenossenschaften* (LPGs, Agricultural Production Cooperatives) (Jochimsen 2010). After reunification in 1990, the property rights in East Germany had to be clarified and former owners were indemnified according to the *Entschädigungs- und Ausgleichsleistungsgesetz* (Indemnification and Compensation Act). The *Landwirtschaftsanpassungsgesetz* (Law on the Adjustment of Agriculture) regulated the decollectivization process and transformation of LPGs towards other legal forms. State-owned land was privatized through the *Treuhandanstalt* (1990–1992) and the *Bodenverwertungs- und -verwaltungs GmbH* (BVVG, since 1992). After 1990, many farmers from West Germany or other Western European countries bought or rented land in former East Germany at prices that were considerably lower than in former West Germany (Koester 2000). This privatization process was recently prolonged to 2030 since the BVVG still holds 136,700 ha of agricultural land in East Germany (BMWi, 2017).

Almost 30 years after the reunification, it could be expected that the open border led to an equalization of conditions on both sides. In this study, we focus on the border region between the state of Lower Saxony (in former West Germany) and the state of Saxony-Anhalt (in former East Germany). After a reform of the counties in Saxony-Anhalt in 2007 (*Kreisreform*), the border region between Saxony-Anhalt and Lower Saxony now consists of four counties on the former east side and six counties on the former west side. With around 415 km, almost one-third of the former inner German border is covered in this analysis.

Table 1 shows similarities and differences between the counties in east and west: The number of farms per county is comparable on both sides of the border (approximately 500 per county), but farms, on average, are more than two times larger in Saxony-Anhalt. This is a result of the history of LPGs: Nowadays, farms in former East Germany are often still organized as cooperatives. In fact, in the former East German border counties, 24% to

Table 1. Descriptive statistics for the border counties (sorted from north to south).

	Border length (km)	Number of farms (2016)	Avg. farm size (ha) (2016)	Share of arable land (2016)	Area hold by juridical person (2016)	Share BVVG (% of transacted agricultural land sold by BVVG) (1991–2016)	Wheat area (ha) (% of arab. land) (2016)	Potato area (ha) (% of arab. land) (2016)	Livestock density (livestock units/ha arab. land) (2016)	Price (€/ha) (2016)	Price growth (2007–16)
Lower Saxony											
Lüchow-Dannenberg	107	587	103	80%	n/a	–	8,045 *(17%)*	5,559 *(11.5%)*	0.37	16,409	127%
Uelzen	15	693	108	90%	n/a	–	14,454 *(21%)*	13,239 *(19.6%)*	0.29	27,761	174%
Gifhorn	71	817	95	83%	n/a	–	9,585 *(15%)*	7,553 *(11.7%)*	0.30	25,519	205%
Helmstedt	122	359	115	91%	n/a	–	16,924 *(45%)*	153 *(0.4%)*	0.09	29,360	144%
Wolfenbüttel	32	403	126	96%	n/a	–	26,603 *(54%)*	60 *(0.1%)*	0.05	29,355	85%
Goslar	69	289	95	87%	n/a	–	12,511 *(53%)*	23 *(0.1%)*	0.20	26,032	67%
Saxony-Anhalt											
Stendal	21	579	269	70%	39%	58%	27,958 *(25%)*	482 *(0.4%)*	0.39	10,755	203%
Altmarkkreis Salzwedel	161	491	256	75%	47%	33%	13,069 *(14%)*	2,008 *(2.1%)*	0.43	9,886	174%
Börde	114	546	277	89%	24%	21%	50,814 *(38%)*	4,569 *(3.4%)*	0.36	18,001	167%
Harz	120	341	303	87%	42%	35%	44,511 *(49%)*	742 *(0.8%)*	0.23	18,494	144%

Data sources: The data for the number and size of farms, the share of arable land, the area held by a juridical person, the wheat and potato growing areas, the livestock density, and the prices for agricultural land in 2007 and 2016 are from the Statistical Office of Lower Saxony and the Statistical Office of Saxony-Anhalt. The area held by a juridical person is not provided by the Statistical Office of Lower Saxony due to the low number of cases and the resulting confidentiality of the information. The border length and share of BVVG in the counties of Saxony-Anhalt are based on own calculations.

47% of the agricultural area is operated by legal persons, whereas this percentage is almost zero in former West German border counties. Joint ownership leads to information asymmetries and could prevent Western farmers from buying land on the Eastern side of the former border due to higher transaction costs. At the same time, however, access to information is facilitated for land sold by the BVVG since it uses public auctions. The BVVG is an important player on the East German land market: It has sold between 21% and 58% of the total transacted agricultural land in the Eastern border counties after reunification.

Similar production structures on both sides of the border could also lead to an assimilation of prices. For example, wheat production is quite strong in the south of both border regions where 50% of the available arable land is used for wheat growing. Moreover, there is a cross-border potato cluster in Lüchow-Dannenberg and Uelzen on the western side and in Altmarkkreis Salzwedel and Börde on the eastern side. Livestock densities are, in general, higher on the eastern side and decrease from north to south.

Agricultural land prices in 2016, however, strongly differ with around 25,000 €/ha in Lower Saxony and 15,000 €/ha in Saxony-Anhalt. The percentage increase from 2007 to 2016 is, in general, slightly larger in Saxony-Anhalt. Figure 1 shows that the absolute gap between prices in former East and West Germany rises, so that a tendency of eastern counties to catch up to their western neighbours cannot be observed. The figure also shows that there is only a small overlap of the time series for eastern and western counties and a rather homogenous price development, especially for the eastern counties. These numbers provide a mixed picture. While production structures show similarities across the border, prices seem to evolve differently. In our empirical analysis, we will scrutinize whether the border still influences price development and if there are regional differences between counties in former East and West Germany.

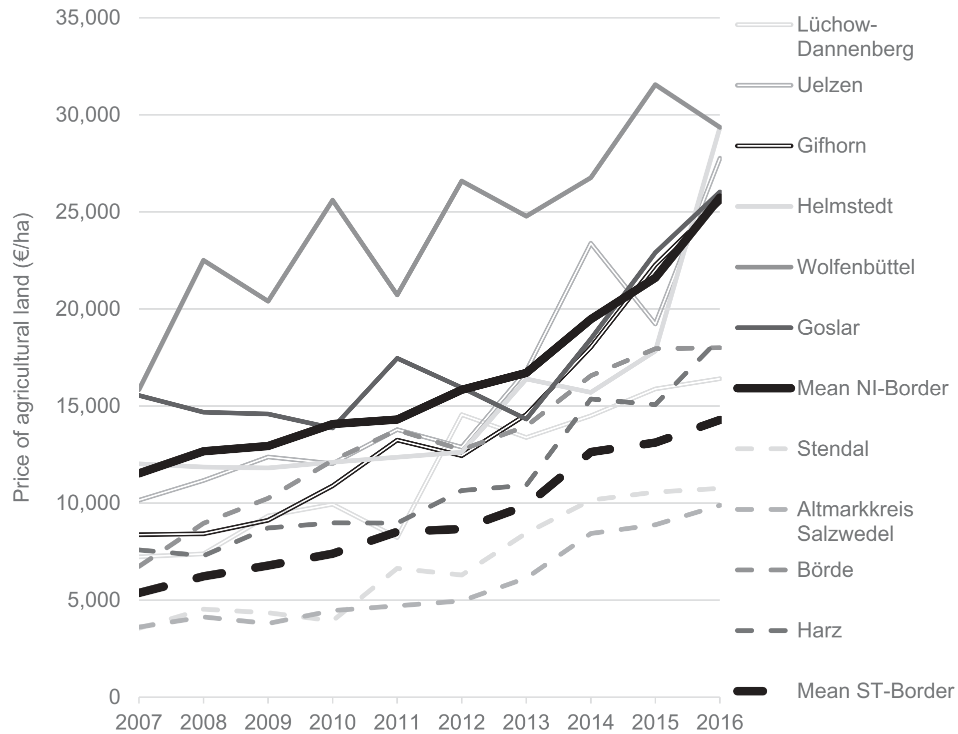

Figure 1. Price development of agricultural land in border counties in Lower Saxony (NI, solid) and Saxony-Anhalt (ST, broken line). Data sources: Statistical Office of Lower Saxony, Statistical Office of Saxony-Anhalt.

Data

The empirical analysis is based on a comprehensive dataset of sale transactions of arable land between 1994 and 2015 in Lower Saxony and Saxony-Anhalt provided by Oberer Gutachterausschuss für Grundstückswerte in Niedersachsen and Gutachterausschuss für Grundstückswerte in Sachsen-Anhalt. It includes information on the price, size, soil quality, and location of sold plots. To conduct the analysis, these data have to be converted into a balanced panel.

Using transaction data has two advantages compared to county averages provided by statistical offices. First, we can derive quarterly instead of yearly average prices and hence obtain a larger panel. Second, the reform of the counties in Saxony-Anhalt in 2007 led to a fusion and reshaping of counties.[4] Through the transaction data, we can create consistent time series for the counties in the pre-reform shape and hence also increase the regional dimension of the panel.

The focus of the study is to evaluate a possible effect of the former German border on land price diffusion. Consequently, to keep the number of regional units at a manageable level, counties in Saxony-Anhalt and Lower Saxony more distant than the 2nd neighbours of border regions are excluded (see Figure 2).[5]

Land price transaction data cannot simply be aggregated to county level cross-section data since land is a heterogeneous factor (Yang, Ritter, and Odening 2017). To homogenize the transaction data, we apply the following hedonic regression to all transactions ($k = 1, \ldots, 82672$):

$$\ln p_k = \delta_{0i} + \delta_{1i} t_i + \delta_2 \text{quality}_k + \delta_3 \text{size}_k + \eta_k, \tag{8}$$

which accounts for soil quality and the size of the transferred plot.[6] The regression also includes

[4]The reform of the counties in 2007 had the following consequences for the border region: Bördekreis and Ohrekreis merged into Börde; Halberstadt, Quedlinburg, Wernigerode, and a small part of Aschersleben-Staßfurt became one county called Harz; and Altmarkkreis Salzwedel and Stendal remained the same.

[5]It could be argued that the empirical application should be confined to border regions. This would, however, prevent the analysis of spillover and spatial effects.

[6]Soil quality is measured by *'Ackerzahl'*, a German evaluation scheme for the quality of agricultural land based on criteria such as soil type, climate, and topography. It has a value that ranges from one ('very poor') to 120 ('very good').

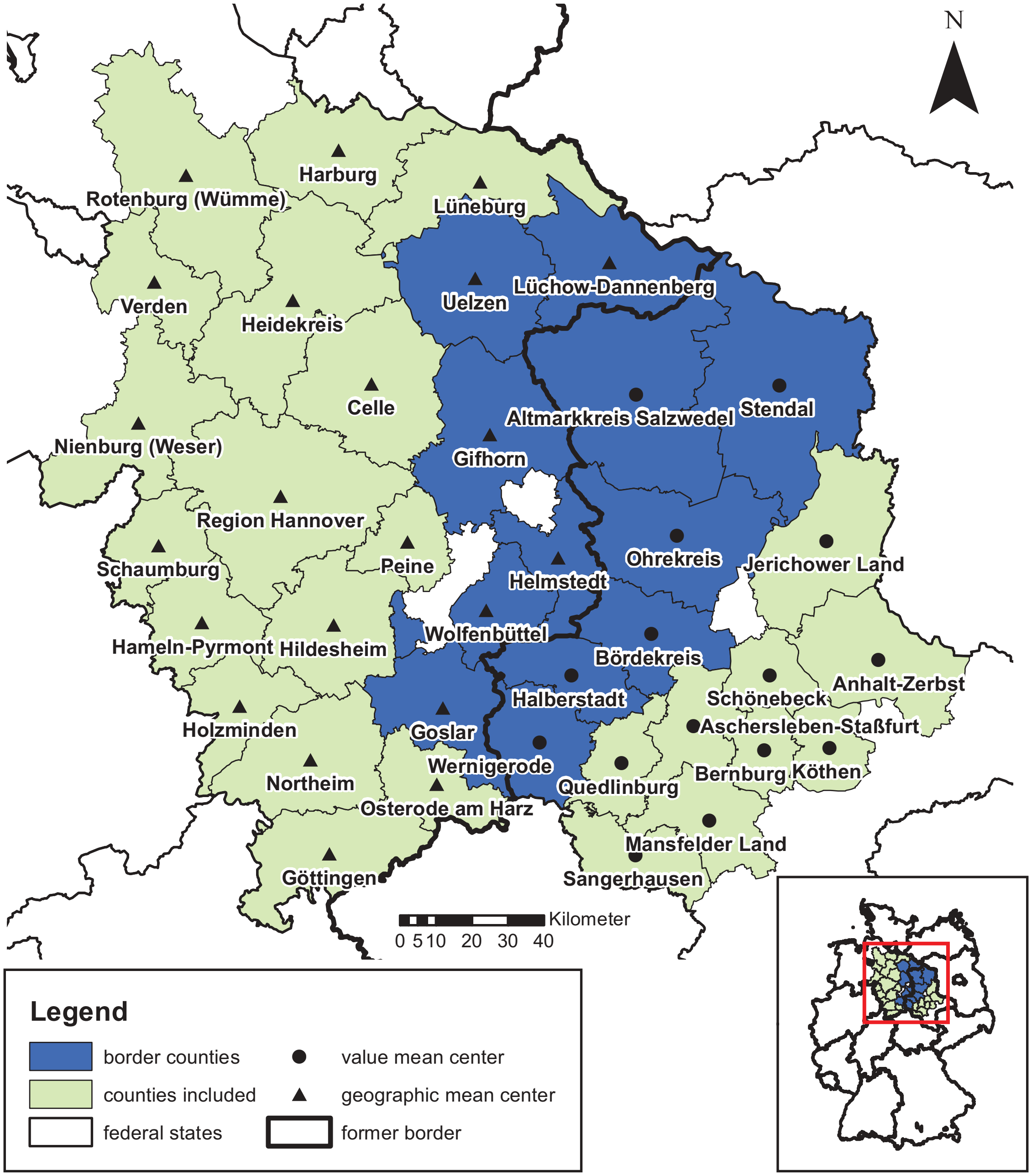

Figure 2. Counties included as well as the geographic location of value and geographic mean centres; the shape of the counties corresponds to the situation before 2007.

a county-specific constant δ_{0i} and time trend t_i to account for county-individual effects that otherwise could bias the estimated effects of size and quality. The hedonic regression is estimated via OLS. Then, the 5% observations with the largest and smallest residuals $\hat{\eta}_k$ are removed and (8) is re-estimated. As expected, soil quality and the size of the transferred land have a positive effect on the price of arable land ($\hat{\delta}_2 = 0.012$, $\hat{\delta}_3 = 0.003$). With these coefficients at hand, log land prices are adjusted to average soil quality and average size:

$$\ln p_k^* = \ln p_k - \hat{\delta}_2\left(\text{quality}_k - \overline{\text{quality}}\right) - \hat{\delta}_3\left(\text{size}_k - \overline{\text{size}}\right) \quad (9)$$

where $\overline{quality}$ and $\overline{size}$ denote the sample means of soil quality and plot size, respectively. The adjusted transaction prices p_k^* are then averaged to quarterly county-level data. The resulting time series are smoothed to eliminate outliers, which can occur due to infrequent transactions for some counties and time periods. A standard exponential moving average of up to four time periods before t are applied in form of $p_{it} = \frac{p_{it}+\sum_{s=1}^{S}(1-\alpha)^s p_{it-s}}{1+\sum_{s=1}^{S}(1-\alpha)^s}$ with $\alpha = \frac{1}{p+1}$. The resulting panel dataset with average prices in 37 counties from the first quarter of 1994 to the fourth quarter of 2015 ($37 \times 88 = 3,256$ observations) is used to estimate the price diffusion model. Equation (3) allows the incorporation of common factors that might influence the development of land prices across the study region. We follow Helgers and Buyst (2016) and add the change in real GDP growth for the same time period as a possible explanatory variable for the price development at county-level within the entire study region. Since we do not expect the price change in a county to influence the real GDP growth, we consider this variable exogenous.

IV. Empirical application

Model specification

To model the spatial relationship between the counties and to estimate the VAR of price pairs, a spatial weighting matrix representing spatial dependencies has to be chosen *a priori*. Although its specification is arbitrary, it is influential for the results of the price diffusion model (Meen 1996). Hence, we apply and test several widely used weighting matrices in our empirical application: two inverse distance matrices $W^{id}, w_{ij}^{id} = \left(\frac{1}{\text{distance between } i \text{ and } j}\right)^v$, with a decay factor v of 1 and 2, respectively, a simple binary neighbourhood matrix W^b, with each i^{th} row element j set to one ($w_{ij}^b = 1$) for a (direct) neighbour county of region j and zero otherwise[7], and a so-called 'binary/distance[2]' matrix W^{bid}, which is the product of the binary and inverse distance matrix ($w_{ij}^{bid} = w_{ij}^b w_{ij}^{id}$) with a decay factor v of two.[8] The latter matrix extends on the simple binary relationship, but limits the influence to first neighbours. Distance is measured between value mean centres of the land transactions for Saxony-Anhalt where geographic coordinates for all transactions are available. For Lower Saxony, where no coordinates are available, the geographic mean of a county is used as its centre. The Moran's I value of mixed binary-distance is the highest, but only slightly surpasses the binary's value. Thus, we proceed with these two spatial weight matrices for the model selection procedure.

The spatio-temporal price diffusion model (3) is based on the assumption of cointegration between the prices of region i and its neighbours. To test this prerequisite, we apply a Johansen trace test (Johansen 1991) with a cointegration constant and unrestricted β_{1i}. More specifically, we test the following two equations individually for all counties:

$$\begin{aligned} ECT_{i,1,t-1} &= p_{i,t-1} - \beta_{0i}^{\text{same}} - \beta_{1i}^{\text{same}} \bar{p}_{i,t-1}^{\text{same}} \\ ECT_{i,1,t-1} &= p_{i,t-1} - \beta_{0i}^{\text{opp}} - \beta_{1i}^{\text{opp}} \bar{p}_{i,t-1}^{\text{opp}} \end{aligned} \tag{10}$$

Table 2 shows that prices in all counties are cointegrated with their neighbours' prices in the same state with the exception of Peine. Moreover, land prices in Uelzen, one of the border neighbouring counties, are not cointegrated with prices across the former border. Note that only significant cointegration relationships (at the 10% significance level) enter the price diffusion model via $ECT_{i,1,t-1}$ or $ECT_{i,2,t-1}$. Furthermore, we test whether land prices converge in the long-run by testing the hypothesis $H_0 : \beta_{1i} = -1$. The hypothesis of (relative) convergence is rejected in 48 out of 49 cointegration relationships at the 10% significance level. Thus, price convergence is rare, a finding which is also reported in other studies (Yang, Odening, and Ritter 2019).

Three specifications of the full system of price time series are estimated representing different assumptions about the convergence process: absolute convergence ($\beta_{1i} = -1$ and $\beta_{0i} = 0$), relative convergence ($\beta_{1i} = -1$ and $\beta_{0i} \neq 0$), and non-convergence ($\beta_{1i} \neq -1$). Moreover, we compare

[7]The diagonal of any weighting matrix W, i.e., county i's element in the i^{th} row, is set to zero ($w_{ii} = 0$).
[8]w_{ij}^{bid} equals either w_{ij}^{id}, if the county j is a direct neighbour of county i, or zero, if they are not neighbouring counties.

Table 2. Johansen cointegration test with constant (trace statistic) and test for restrictions on cointegration vector.

County	Trace statistic $\bar{p}_{i,t}^{same} H_0$: $r = 0$	p-value H_0: $\beta_1^{same} = -1$	Trace statistic $\bar{p}_{i,t}^{opp} H_0$: $r = 0$	p-value H_0: $\beta_1^{opp} = -1$
Lower Saxony				
Celle	29.10	0.01		
Gifhorn	24.95	< 0.01	22.25	0.01
Göttingen	24.47	0.04		
Goslar	24.10	< 0.01	22.14	< 0.01
Hameln-Pyrmont	20.99	< 0.01		
Harburg	31.32	< 0.01		
Heidekreis	37.56	< 0.01		
Helmstedt	21.46	< 0.01	31.24	0.02
Hildesheim	19.35	0.01		
Holzminden	24.22	< 0.01		
Lüchow-Dannenberg	35.56	0.03	22.52	0.01
Lüneburg	44.83	0.01		
Nienburg/Weser	33.23	< 0.01		
Northeim	27.87	< 0.01		
Osterode am Harz	21.55	< 0.01		
Peine	13.84	0.06		
Region Hannover	30.30	0.02		
Rotenburg (Wümme)	36.78	< 0.01		
Schaumburg	33.32	< 0.01		
Uelzen	37.87	0.03	16.57	0.05
Verden	45.43	< 0.01		
Wolfenbüttel	21.66	< 0.01	29.46	< 0.01
Saxony-Anhalt				
Altmarkkreis Salzwedel	31.92	0.02	25.98	0.05
Anhalt-Zerbst	49.75	< 0.01		
Aschersleben-Straßfurt	55.05	< 0.01		
Bernburg	43.47	< 0.01		
Bördekreis	30.73	0.02	25.73	0.05
Halberstadt	28.13	< 0.01	21.95	0.01
Jerichower Land	50.14	< 0.01		
Köthen	35.95	0.03		
Mansfelder Land	33.82	< 0.01		
Ohrekreis	32.15	< 0.01	22.92	0.04
Quedlinburg	35.99	< 0.01		
Sangerhausen	25.85	0.26		
Schönebeck	45.62	< 0.01		
Stendal	27.14	0.08	25.92	0.02
Wernigerode	31.50	< 0.01	18.02	< 0.01

The critical values for 10%, 5%, 1% level of significance are 17.85, 19.96, and 24.6, respectively.

Table 3. Share of significant parameters for the N price diffusion equations (p-value smaller or equal to 0.05).

	$\phi_{i,1}$	$\phi_{i,2}$	$\gamma_{i,1}$	$\gamma_{i,2}$	$\gamma_{i,3}$	λ_i
Presented Model (No. 1 by ΔAIC)	75.0%	72.7%	65.1%	59.2%	20.0%	40.5%
Alternative Model with W^b	80.8%	25.0%	58.1%	44.2%	13.3%	45.9%
Presented Model OLS-estimation	63.9%	36.4%	34.9%	8.2%	6.7%	29.7%

models with and without inclusion of common factors. Finally, we estimate model variants with a border effect (3) and without a border effect (2) to address our main research question. Out of the overall 24 model specifications, the model with the lowest AIC value, and therefore the specification that best fits the data generating process, uses the 'binary/distance[2]' spatial weight matrix, does not restrict county prices to converge, incorporates a common factor, and separates neighbouring prices into two groups, thus representing the border effect.

Estimation results

Table 3 reports the shares of significant parameters for the N price diffusion equations for the iterative SUR estimation results.[9] In line with the previous Johansen trace tests, we observe a large share of significant adjustment coefficients (75.0% and 72.7% respectively) pointing at a long-run equilibrium of land prices with their neighbour counties'

[9]In other applications, a potential endogeneity problem of the price diffusion model is addressed. Holly, Pesaran, and Yamagata (2011) conduct a Wu-Hausman test and use an IV estimator if required. Fortunately, the endogeneity issue is less severe here. In contrast to Holly, Pesaran, and Yamagata (2011) and Helgers and Buyst (2016), we do not consider contemporaneous effects of a dominant region. Unlike Yang, Odening, and Ritter (2019), we neither included contemporaneous effects of other regions. The only contemporaneous effect may result from the common factor z_t, which is the real GDP growth in our case. However, it is reasonable to assume that land price changes in a county do not influence the real GDP growth, i.e., we consider z_t as an exogenous variable.

Table 4. Estimation results and hypotheses test statistics of border counties' price diffusion Equations (3) with a binary/distance^2 spatial matrix (standard errors in brackets) and no convergence.

	$\phi_{i,1}$	p-value	$\phi_{i,2}$	p-value	$\gamma_{i,1}$	p-value	$\gamma_{i,2}$	p-value	$\gamma_{i,3}$	p-value	λ	p-value	*K/L/Q*	$\phi_{i,1} \geq \phi_{i,2}$	$\phi_{i,2} \geq 0$
Saxony Anhalt															
Bördekreis	−0.15	< 0.01	−0.05	0.01	−0.30	< 0.01	0.12	0.24	0.03	0.51	0.31	< 0.01	1/1/1	−2.80	−2.65
	(−0.03)		(−0.02)		(−0.07)		(−0.10)		(−0.04)		(−0.06)				
Halberstadt	−0.25	< 0.01	0.05	0.11	−0.15	0.10	−0.27	0.01	0.01	0.94	0.12	0.09	1/1/1	−4.77	1.59
	(−0.06)		(−0.03)		(−0.09)		(−0.11)		(−0.10)		(−0.07)				
Ohrekreis	−0.08	0.11	−0.08	0.11	−0.05	0.62	−0.10	0.55	−0.24	0.02	0.03	0.75	1/4/1	0.11	−1.59
	(−0.05)		(−0.06)		(−0.09)		(−0.17)		(−0.10)		(−0.10)				
Stendal	0.02	0.19	−0.09	< 0.01	−0.04	0.59	0.23	< 0.01	0.05	0.06	0.04	0.37	1/1/1	4.10	−4.09
	(−0.02)		(−0.02)		(−0.07)		(−0.08)		(−0.03)		(−0.05)				
Wernigerode	−0.18	0.01	0.02	0.01	−0.28	< 0.01	0.33	< 0.01	−0.10	0.01	0.05	0.52	2/1/1	−2.83	2.50
	(−0.07)		(−0.01)		(−0.08)		(−0.10)		(−0.04)		(−0.08)				
Altmarkreis-Salzwedel	−0.07	0.01	−0.08	< 0.01	0.15	0.03	−0.51	< 0.01	0.00	0.97	−0.44	< 0.01	1/1/1	0.26	−5.20
	(−0.03)		(−0.02)		(−0.07)		(−0.09)		(−0.07)		(−0.13)				
Rest of Saxony-Anhalt	−0.25				−0.10		0.22				0.26				
Lower Saxony															
Gifhorn	0.03	0.66	−0.35	< 0.01	−0.10	0.21	0.42	< 0.01	−0.07	0.61	0.54	< 0.01	1/1/1	3.86	−4.96
	(−0.07)		(−0.07)		(−0.08)		(−0.15)		(−0.14)		(−0.11)				
Goslar	−0.13	0.78	−0.20	0.67	−0.21	0.02	0.82	0.02	−0.42	0.07	0.05	0.80	1/1/1	0.11	−0.44
	(−0.44)		(−0.44)		(−0.09)		(−0.34)		(−0.23)		(−0.17)				
Helmstedt	−0.03	0.78	−0.45	< 0.01	−0.20	0.03	0.10	0.46	−0.06	0.71	−0.22	0.01	1/1/4	2.58	−3.49
	(−0.10)		(−0.13)		(−0.09)		(−0.13)		(−0.16)		(−0.09)				
Wolfenbüttel	0.09	0.15	−0.33	< 0.01	−0.06	0.32	−0.20	< 0.01	−0.03	0.55	−0.01	0.85	1/1/1	4.26	−4.43
	(−0.06)		(−0.07)		(−0.06)		(−0.06)		(−0.05)		(−0.07)				
Lüchow-Dannenberg	−0.46	< 0.01	−0.14	0.04	−0.57	< 0.01	0.25	< 0.01	−0.17	0.32	0.51	< 0.01	1/1/1	−3.44	−2.03
	(−0.06)		(−0.07)		(−0.07)		(−0.09)		(−0.17)		(−0.13)				
Uelzen	−0.43	< 0.01			−0.32	< 0.01	0.15	0.19	0.48	< 0.01	0.04	0.77	1/1/1		
	(−0.09)				(−0.08)		(−0.11)		(−0.12)		(−0.13)				
Rest of Lower Saxony	−0.37				−0.31		0.21				0.07				

The last two columns provide the tests statistics for the testing of the two hypotheses by one-sided t-tests ($H_0 : \phi_{i,1} \geq \phi_{i,2}$ and $H_0 : \phi_{i,2} \geq 0$), the critical value for $a = 0.05$ is -1.64.

prices. In only six cases neither $\phi_{i,1}$ nor $\phi_{i,2}$ are statistically different from zero at the 5% level. The average speed of adjustment for all is rather slow (-0.23). This value is comparable to other studies in the real estate market (e.g., Helgers and Buyst 2016; Holly, Pesaran, and Yamagata 2011).

As a robustness check, Table 3 also reports the shares of significant parameters for another model specification using a binary spatial matrix as well as the results of an OLS estimation, which are in both cases considerably lower.

Details of the iterative SUR-estimation results of the price diffusion Equations (3) for 12 border counties as well as the average results for the remaining counties' equations of each state are presented in Table 4.[10] Inspection of Table 4 reveals regional differences in land price diffusion. In most cases (8 out of 11), $\phi_{i,1}$ is larger (in absolute terms) than its counterpart $\phi_{i,2}$, i.e., border counties' prices adjust faster to a long-run equilibrium with neighbours in the same state compared to neighbours across the former border. Moreover, non-border counties show a higher level of land market integration than border counties in terms of their adjustment speed parameters. A possible explanation is that former border counties were located in the periphery of West and East Germany. This remoteness led to a decoupling from the economic development of the rest of the country (Redding and Sturm 2008).

Short-run dynamics from the regional price diffusion equations are captured by lagged variables based on differences in the own price and neighbours' average price within the state and across the former border. The coefficient of the changes in the lagged own price is significant in almost all border county Equations (10 out of 12) and has a negative sign. The spatial pattern of short-run dynamics is more heterogeneous compared to long-run estimates, but differences between the two groups of neighbours exist. The parameter estimates for changes in the lagged land price for within state neighbours $\gamma_{i,2}$ are of a greater magnitude (0.12), on average, for the former border counties compared to the parameters $\gamma_{i,3}$ for across border lagged land price differences (0.04).

The overall effect of these rather heterogeneous parameters is captured by the GIRFs, which are presented below. Shocks in (real) GDP growth are included in the model as a common factor for all regional price equations. The corresponding parameters λ are statistically significant for more than 40% of the estimated equations. However, sign and size of λ estimates vary considerably among counties. A similar finding is reported by Helgers and Buyst (2016).

To examine the presence of a border effect more explicitly, we empirically test the two hypotheses from Section 2 using one-sided t-tests. This test procedure leads to the following classification: counties with a border effect (reject Hypothesis 1 or fail to reject Hypothesis 2), counties with no border effect (fail to reject Hypothesis 1 and reject Hypothesis 2), and counties that cannot be assigned to one of the former groups (fail to reject both hypotheses). The spatial distribution of these categories is displayed in Figure 3.

Land prices in counties that belong to the 'border effect' group do not have a long-run relationship with land prices from neighbouring counties across the former border or they react more slowly to deviations for the cointegration relationship compared to within state neighbours. This group comprises Bördekreis, Halberstadt, Lüchow-Dannenberg, Uelzen, and Wernigerode. Most of these counties are located in Saxony-Anhalt and form a regional cluster at the southern intersection of both states. This finding may be traced back to the relative low share of BVVG administrated transactions in these counties, which, in turn, may lead to larger information asymmetries and higher transaction costs for West German buyers (see Table 1). Furthermore, in the East German counties of this group, a high share of land transactions took place more than 10 kilometres away from the former border (see Figure 4). It is unlikely that these land plots were attractive to farmers from Lower Saxony since it is unprofitable to operate them due to high transportation costs. Another peculiarity of this group is that Lüchow-Dannenberg and Uelzen are characterized by high levels of potato production, while counties across the former border focus on wheat production.

[10]In case of higher-order lags, Table 4 reports the parameter of the first lagged variable only.

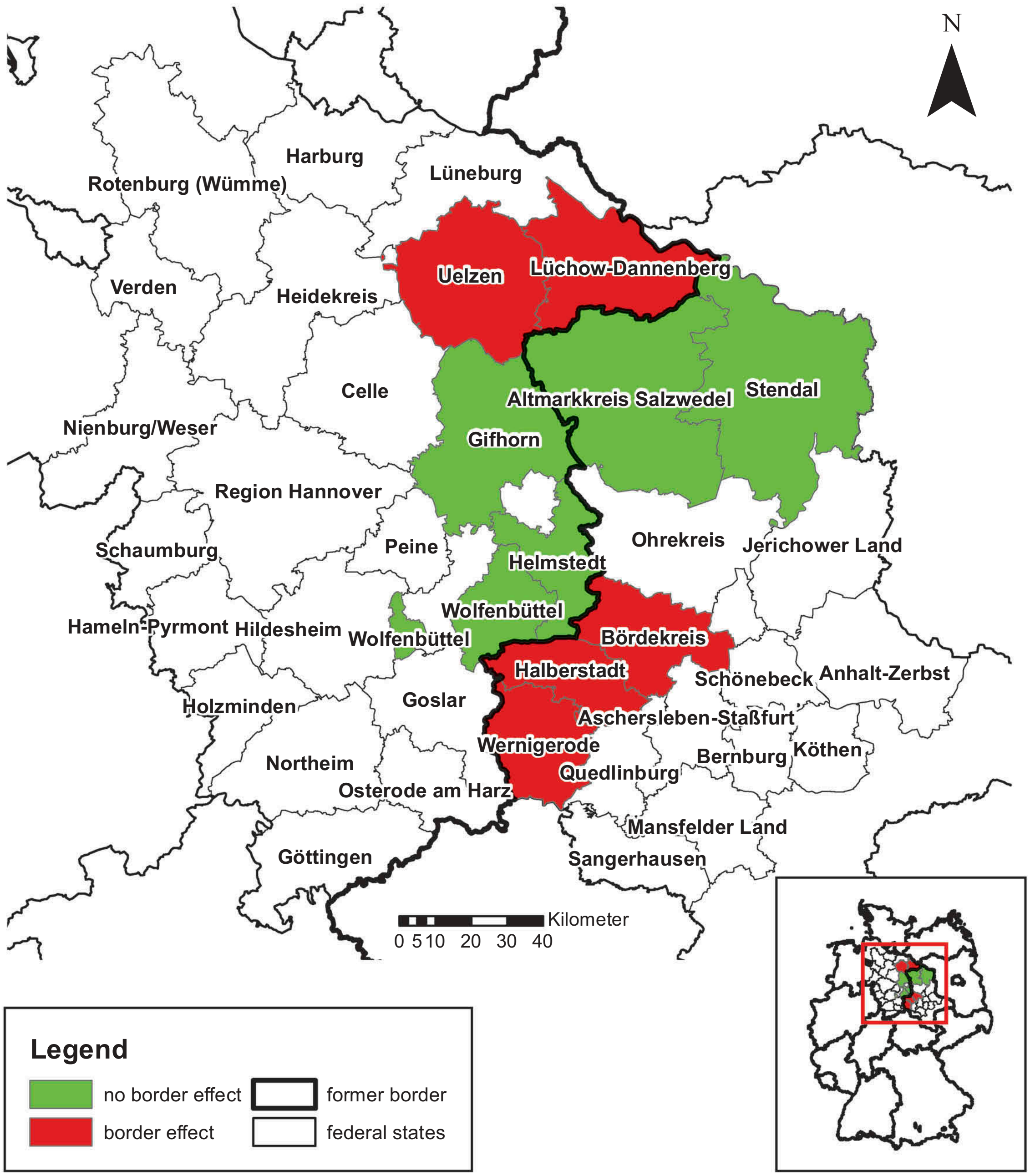

Figure 3. County groups based on iterative SUR estimation results and hypotheses testing.

The hypothesis of a border effect is rejected for five counties located in the northern part of Saxony-Anhalt (Altmarkkreis-Salzwedel, Stendal) and in Lower Saxony (Gifhorn, Helmstedt, and Wolfenbüttel). Counties in this group show a relatively high share of BVVG administrated transactions and transactions are located close to the border, lowering transaction and economic costs for farmers operating across the border (see Table 1 and Figure 4). Only for two counties, Goslar and Ohrekreis, neither hypothesis can be rejected.

Figure 5 displays the reaction of a shock (one standard deviation) in county i across the study region over ten years. The origin of the price shock in county i shows a value of unity at $t = 1$. The other counties are ordered by distance from the origin of the shock county. Counties to the left (with

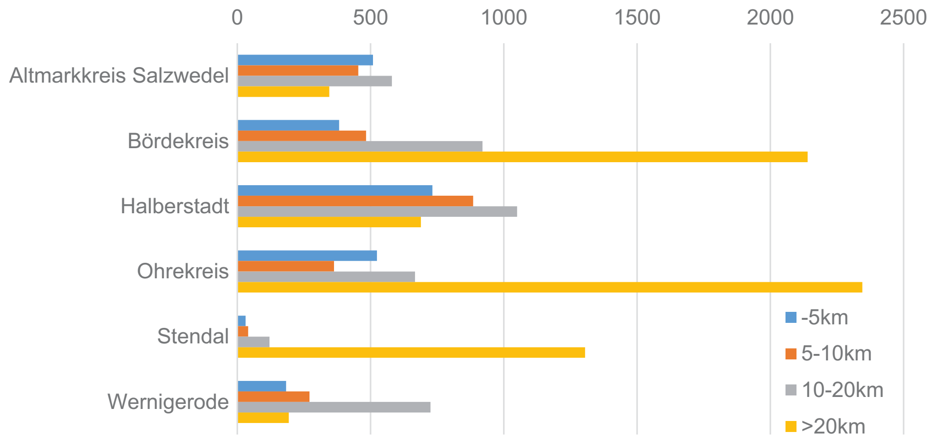

Figure 4. Number of transactions in border counties in Saxony-Anhalt from 1994 to 2016 sorted by distance from the former border.

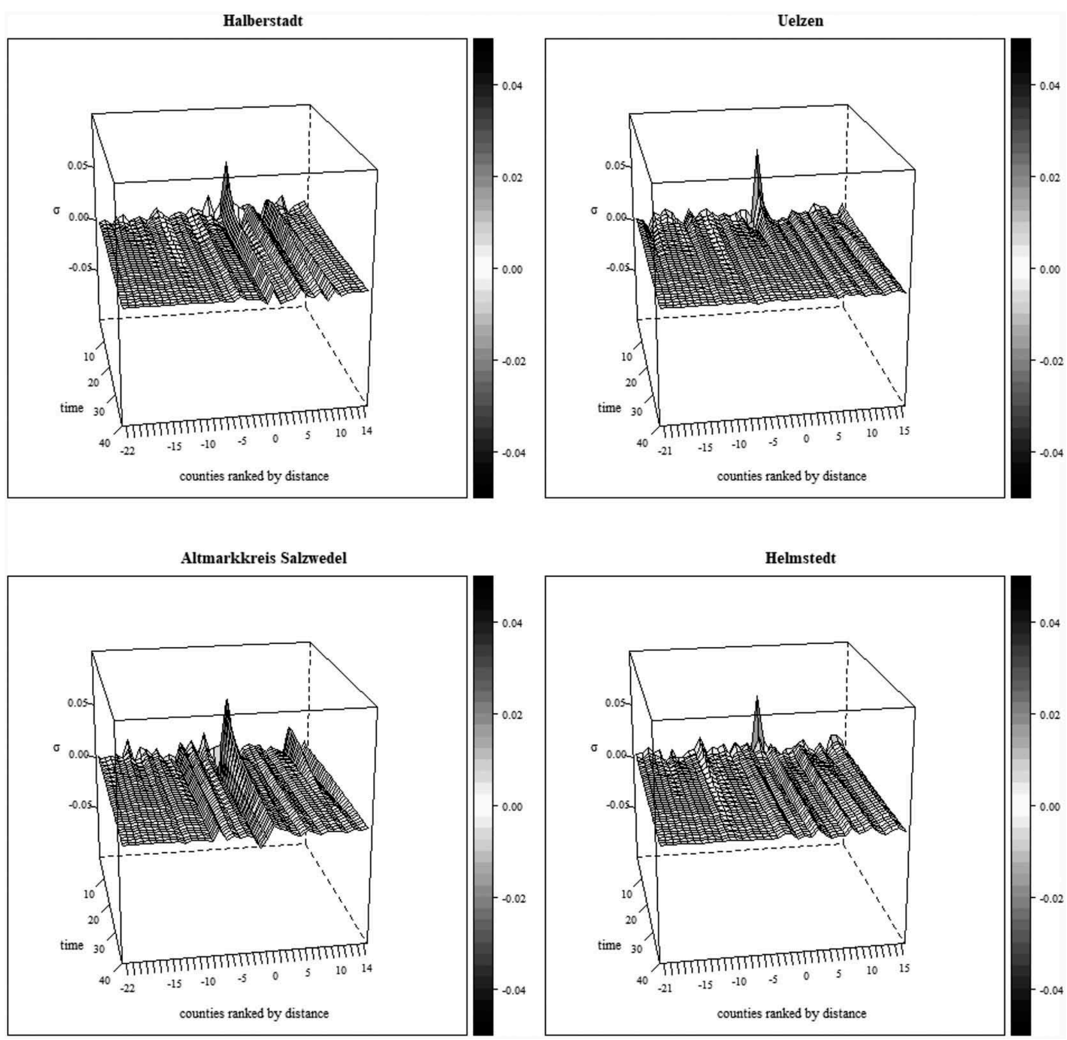

Figure 5. GIRFs for selected border counties with shocks of one standard deviation σ.

a negative value) are located in Lower Saxony (the Western state) while counties to the right (with a positive value) are located in Saxony-Anhalt (the Eastern state). The graphs at the top of Figure 5 show shocks originating in counties in which land prices are affected by the former border, Halberstadt (left) and Uelzen (right). The price diffusion model revealed that Halberstadt's land price interacts only with neighbours' prices in the same state and not with neighbours across border. Hence, we see that a shock to Halberstadt causes a permanent positive increase in its own price and prices for most of its neighbouring counties in Saxony-Anhalt, while there is only little response in Lower Saxony caused by the short-run dynamics of the VECM. A shock to the land price in Uelzen causes only a temporary deviation from the pre-shock level in Lower Saxony and is completely absorbed within the considered time horizon. In contrast, prices across the border do not return to their pre-shock level. The two graphs at the bottom of Figure 5 showcases of the 'no border group' on both sides of the former border (Altmarkkreis-Salzwedel and Helmstedt). In both cases, the prices do not return to their pre-shock level. These permanent increases are transmitted through the long-run spatial market integration to neighbours' prices on both sides of the border.

The GIRFs confirm the results of the price diffusion model with a border effect (3). The former German border still affects land price diffusion in Germany, not just in counties' land markets located directly at the former border. While most counties in the southern part of the study region are only integrated with within state land markets, a number of counties, particularly those in the northern study region with similar production structures, have a cointegration relationship with across neighbours' prices and are sometimes separated from their own states' overall land price diffusion process.

V. Conclusions

This article examines whether there is a diffusion of agricultural land prices across the former border separating East and West Germany. The research question was motivated by concerns among policymakers in several EU countries that in unregulated land markets, the activities of foreign investors may cause high prices to spill over into neighbouring countries that initially have lower land price levels. On the other hand, the European Commission recently appealed to EU Member States to adjust their land market regulations according to European Law, which requires the free movement of capital within the EU (European Commission 2016). To shed some light on this controversial discussion, we consider the German reunification as a case study on land price development in two different states after the border was removed. We apply a land price diffusion model with an error correction specification that estimates to what extent agricultural land markets are spatially integrated. A novel feature of our model is its ability to distinguish price diffusion within states and across state borders. We find that local agricultural land markets in Germany are spatially integrated, i.e., prices in one county are linked with prices in neighbouring counties by a long-run equilibrium relationship. Spatial market integration, however, does not hold among all counties in our study area. In line with earlier studies, there is evidence for convergence clubs, which differ in their land price dynamics.

With regard to our main research question, we find evidence for a persistent border effect given that the fraction of spatially integrated counties is larger within states than across the former border. Moreover, for many counties along the former border, we observe non-significant error correction terms for prices of neighbouring counties across the former border. It is noteworthy that the former border does not act as a strict barrier for price alignment between former East and West Germany. In fact, it is permeable at several locations. In some cases, it even happens that counties share similar land price dynamics with neighbours across the border, but not with neighbours within the same state. By virtue of its reduced form character, our model cannot provide a clear answer on what is behind this border effect and why it appears to be local. We conjecture that differences in farm size structures, long-lasting rental contracts, local market power, non-transparency of sellable land plots and information asymmetries regarding the property status of land may explain non-integration and stickiness of land prices in parts of East and West Germany. Providing empirical evidence for the role of these economic factors is a promising direction for further research.

Our results are not only interesting from a historical perspective, but they are also relevant for a better understanding of the functioning of agricultural markets. From a policy perspective, it is striking to realize that even 25 years after German reunification, pronounced land price differences persist. Even if regional land markets across the former border are integrated, land prices react rather slowly and only in relative terms, i.e., land prices do not reach the same level. It is quite likely that price diffusion through existing borders within the EU would take even more time given language barriers, different administrative procedures for land acquisitions, different tax systems, and more pronounced information asymmetries between domestic and foreign market participants. Therefore, proposals for stricter land market regulations aiming at the protection of local farmers and capping of land prices through the discrimination of foreign buyers appear questionable.

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