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Abstract

This article applies the novel concept of global panel cointegration to analyze the role played by trade and foreign direct investment (FDI) activity in driving regional total factor productivity (TFP). Using West German state-level data for the period 1976–2008, the approach allows us to identify the magnitude of direct trade and FDI effects as well as spatial spillovers from these variables. The author finds that the inclusion of spatial lags significantly improves the fit of the empirical model and allows us to strongly reject the null of no cointegration among the variables in the full spatial specification. For the long-run cointegration equation, the empirical results hint at export- and FDI-led growth. Additionally, outward FDI activity shows to have positive spatial spillover effects among German regions, while the spatial patterns of import and inward FDI activity indicate substitution effects of interregional input–output linkages in favor of international ones over the sample period 1976–2008. In the short run, TFP growth is predominantly affected by changes in exports, inward and outward FDI stocks, where the latter variable also provokes positive spillovers. The author uses alternative spatial weighting matrices based on geographical information as well as interregional goods transport flows to check the

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robustness of the obtained spillover effects in the long and short run. Here, the results turn out to be similar for the different empirical specifications employed throughout the analysis. Finally, summing over the four variables to get a direct and indirect net effect of internationalization activity, the author finds that the direct effect is always positive, while the indirect net effect is positive in the short run but slightly negative in the long-run equation.

Keywords

global cointegration, spatial Durbin model, TFP, trade, FDI

Introduction

Intensive research has been done to elaborate the role played by international trade and foreign direct investment (FDI) in fostering economic growth through market-size effects and technology transfer among other channels of transmission. However, while most studies concentrate on the aggregate or sectoral dimension of this subject matter, only recently have economists also become aware of the likely importance of its spatial extent. From a (regional) policy perspective, it is of considerable interest to draw viable conclusions with respect to the evolution of interregional inequalities in the course of enhancing trade integration and foreign capital deepening.

Using panel data for West German states between 1976 and 2008, this article aims to shed more light on the direct and indirect interregional ties between the evolution of technological progress and international business activity. While direct effects measure the region's own comovement of economic development and internationalization activity over time, indirect effects aim at identifying the strength of interregional spillover effects of trade and FDI and thus allows us to measure the spatial extent of internationalization activity at the regional level. Moreover, using a dynamic model context, the analysis looks both at the long-run structural components of the relationship among these variables and the transitory short-run dynamics.

In a recent survey, Brühlhart (2011) has pointed to the need of further research efforts in order to analyze the implications of trade integration on intranational economic geographies. The author reviews the existing theoretical and empirical attempts to describe and explain these within-country adjustments. From a theoretical perspective, the uneven evolution of economic growth due to internationalization activity across space is most prominently discussed by the new economic geography (NEG) literature, where long-run spatial divergence is typically the result of a concentration of economic activity in certain agglomerations. In almost all NEG models, free trade and capital movement play a key role; however, whether agglomeration or dispersion forces dominate depends crucially on the underlying modeling choices: while one bulk of models favors internal economic dispersion in the course of external trade liberalization with unchanged internal transportation

costs (Krugman and Livas Elizondo 1996; Behrens et al. 2007), alternative model specifications instead find increasing agglomeration forces as a result of enhancing trade integration (Monfort and Nicolini 2001; Paluzie 2001). These conflicting predictions basically hold irrespective of whether within-country difference in terms of topography and infrastructure are assumed to be present or not (Brühlhart 2011).

Since the different modeling approaches rely on specific functional forms and thus no *a priori* reasoning can be made to adjudicate between the two, Brühlhart (2011) argues that no clear-cut theoretical conclusion can be drawn and, instead, an empirical identification strategy should be conducted to answer the question related to the spatial extent of internationalization activity. In this article, we therefore measure the likely magnitude of spatial concentration and dispersion forces through the identification of spatial spillovers of trade and FDI activity. Having regional data with a long time dimension at hand, our methodological approach rests on a global cointegration analysis, which aims at identifying comovements among the set of variables within and between regions. The notion of global panel cointegration, which combines elements of aggregate time-series analysis, panel and spatial econometrics, has recently been introduced by Beenstock and Felsenstein (2010).

Besides controlling simultaneously for the well-known problems of spurious regression with trended data as well as cross-sectional correlation among the residuals, global cointegration analysis thus enables a complex interpretation of the importance of space and time in determining the joint evolution of total factor productivity (TFP) and internationalization activity. The basic idea of this approach may be best explained with a simple example: consider the case of two regions, where one region is heavily engaged in international trade or FDI and directly benefits from this activity in terms of productivity growth. The second region is not actively engaged in internationalization activity; however, it may potentially benefit from the first region's openness via forward and backward linkages, which in turn raise output for this region as well. On the other hand, the increased trade integration of region 1 may also lead to an internal reallocation of economic activity, which negatively affects the growth path of region 2. Rather than observing a long-run comovement and its short-run adjustment path for the region's own level of output and internationalization activity, the inclusion of a spatially lagged trade and FDI variable is thus likely to tell a much more complex story of cointegration among regional economic growth, trade, and FDI activity.

The remainder of the article is organized as follows: in the next section, we start with a brief overview of recent empirical contributions regarding the relationship of economic growth, trade, and international capital movement and then discuss the empirical model specification as well as the estimation technique. The section on Data and Stylized Facts presents the data used, while the section on Empirical Results deals with the discussion of the econometric results. Here, a special focus is given to their economic interpretation and analysis of statistical robustness when alternative spatial weighting matrices are used. The final section concludes the article.

Theory and Empirics of Productivity–Trade–FDI Linkages

According to growth theory, a change in the output level over time may either stem from the accumulation of production factors (labor and capital) and/or improvements in the technology level. By fostering the international transfer of technology, trade and FDI are typically expected to affect growth through this transmission channel (Saggi 2010). As a result, the empirical mainstream approach to reveal these linkages typically builds on an aggregate time-series perspective aiming to identify long-run cointegration relationships and analyzing the direction of causality among the variables involved. In a recent meta study for the FDI–growth relationship, the Organization for Economic Cooperation and Development (2002) finds that in eleven of the fourteen studies FDI contributes positively to income growth and factor productivity. A similar meta study by Ozturk (2007) confirms this link. Investigating the simultaneous interference of trade and FDI on growth and vice versa, Ekanayake, Vogel, and Veeramacheneni (2003), Dritsaki, Dritsaki, and Adamopoulos (2004), Wang, Liu, and Wei (2004), Makki and Somwaru (2004), and Hansen and Rand (2006) among others use cointegration analysis to identify the long- and short-run effects among the variables and, by means of Granger causality tests, get general evidence for a bidirectional causal relationship between internationalization activity and economic growth. Using data for North and South American countries between 1960 and 2001 (including Brazil, Canada, Chile, Mexico, and United States), Ekanayake, Vogel, and Veeramacheneni (2003), for instance, report evidence in favor of trade-led growth, while results for (inward) FDI-led growth are mixed. Wang, Liu, and Wei (2004) report for a panel of seventy-nine countries that FDI has a positive impact on growth in high- and middle-income countries, but not in low-income countries. Looking more close at a subsample of developing countries, Hansen and Rand (2006) find that FDI has an impact on gross domestic product (GDP) via knowledge transfers and the adoption of new technology.

Only very few studies give an explicit account of spatially related variables in the analysis of the trade–FDI–growth nexus. One exception is Özyurt (2008), who estimates a long-run model for labor productivity of Chinese provinces driven by trade and FDI as well as their respective spatial lags. The author finds that FDI and trade volumes have a positive direct effect on regional labor productivity. Her results for the sample period 1979–2006 show that the geographical environment has a subsequent influence on labor productivity in a given region. Besides the spatial lag of the endogenous variable as a “catch-all” proxy for spatial effects, FDI spillovers turn out to be of specific interregional nature. These findings give a first indication that the growth impact of internationalization activity is not restricted to direct effects, but may also be influenced by spillovers from the economic development in neighboring regions. Complementary to the evidence reported by Özyurt (2008), there is a growing literature with respect to third-country effects of FDI activity (see, e.g., Baltagi, Egger and Pfaffermayr 2007; Garretsen and Peeters 2009; Shepotylo 2010). Using regional data, Bode, Nunnenkamp, and Waldkirch (2011) analyze

spatial effects of FDI and other Marshallian externalities on regional productivity growth among US states. The authors find that spatial effects shape these externalities significantly and help avoid drawing false policy conclusions. Similarly, Mullen and Williams (2005) test for the impact of FDI on income convergence among US states. The authors find that FDI has a significantly positive effect on regional income growth.¹

As a common point of departure in the literature summarized above, we start from a stylized regional production function to model the transmission channels from trade and FDI activity to economic growth (Edwards 1998). A spatially extended version is, for instance, presented in Ertur and Koch (2007). The model can be written as

$$Y_{i,t} = A_{i,t} K_{i,t}^{\alpha} L_{i,t}^{\beta}, \quad (1)$$

$$A_{i,t} = A_{0i,t} \left(\mathbf{T}_{i,t}^{\gamma} \right) \left(\mathbf{F}_{i,t}^{\delta} \right) \left(\tilde{\mathbf{T}}_{i,t}^{\phi} \right) \left(\tilde{\mathbf{F}}_{i,t}^{\xi} \right), \quad (2)$$

$$A_{0i,t} = A_{0,i} e^{\mu t} \tilde{A}_{0i,t}^{\rho}, \quad (3)$$

where $Y_{i,t}$ denotes aggregate production of region i at time t . The cross-sectional dimension is specified as $i = 1, \dots, N$ and the time dimension is $t = 1, \dots, T$. K and L are capital and labor inputs, respectively. $A_{i,t}$ is the TFP of region i . The latter in turn is driven by $A_{0i,t}$ as a region- and time-specific productivity as well as international activities including trade and FDI, where \mathbf{T} is a vector of trade variables (export and import volumes) and \mathbf{F} is a vector of FDI activity (inward and outward FDI stocks). Variables denoted by “ \sim ” indicate weighted averages of values for spatial proximate neighbors. In modeling $A_{0i,t}$, we draw on Bode, Nunnenkamp, and Waldkirch (2011) and create it from the following terms: $A_{0,i}$ represents the productivity effects of time-invariant location factors, $e^{\mu t}$ captures the common productivity effects of exogenous technological progress. The last term in equation (3) measures interregional productivity spillovers. α , β , γ , δ , ϕ , ξ , and ρ are the coefficients to be estimated.

The spatial lag term for a variable X is defined as

$$\tilde{X}_{it} = \sum_{j \neq i}^N w_{ij} X_{jt},$$

indicating that the variable is constructed as a weighted average of values in neighboring regions, where w_{ij} are the individual weights taken from a spatial weighting matrix \mathbf{W} . The latter is typically row-standardized with $\sum_j w_{ij} = 1$. Alternative empirical operationalizations for \mathbf{W} are discussed in the section Data and Stylized Facts. Since we are interested in quantifying the role played by international trade and FDI activity on intraregional and interregional technological progress, our main interest is to estimate equations (2) and (3). To do so, we proceed in a two-step

approach. First, we compute a proxy for TFP $A_{i,t}$ as residual from a first-step linear production function model (small letters denote logarithmic transformations) as

$$y_{i,t} = a + \alpha k_{i,t} + \beta l_{i,t} + e_{i,t}, \quad (4)$$

where $y_{i,t} = \ln(Y_{i,t})$ and likewise for k and l , a is a regression constant and the residual $e_{i,t}$ measures the regional TFP level. Then, for estimating equation (2) we need to compute the reduced form of equation (3) and substitute it into equation (2). Using a log-log form, the empirical specification for TFP estimation can be written as

$$\text{tfp}_{i,t} = \psi + \iota_i + \iota_t + \rho \widetilde{\text{tfp}}_{i,t} + \gamma' \mathbf{t}_{i,t} + \delta' \mathbf{f}_{i,t} + \phi' \tilde{\mathbf{t}}_{i,t} + \xi' \tilde{\mathbf{f}}_{i,t} + \mathbf{u}_{i,t}, \quad (5)$$

where $\text{tfp}_{i,t}$ is logged total productivity in region i at time t as a function of its spatial lag ($\widetilde{\text{tfp}}_{i,t}$), logged trade and fdi variables ($\mathbf{t}_{i,t}$, $\mathbf{f}_{i,t}$) as well as their spatial lags ($\tilde{\mathbf{t}}_{i,t}$, $\tilde{\mathbf{f}}_{i,t}$). The empirical specification further contains individual- (ι_i) and time-fixed effects (ι_t); $\mathbf{u}_{i,t}$ is the error term of the model. Since we are dealing with a long time-series dimension and thus potentially nonstationary variables, special attention has to be drawn to a proper model specification: that is, only if both tfp and the vector of endogenous regressors $\mathbf{z} = (\mathbf{t}, \mathbf{f})$ form a stable cointegration relationship, then the error term $\mathbf{u}_{i,t}$ is warranted to be stationary and integrated of order $\mathbf{u}_{i,t} \sim I(0)$. The latter in turn avoids the problem of spurious regression among variables of order $I(d)$, with $d \leq 0$. Moreover, as Beenstock and Felsenstein (2010) point out, in an aspatial specification \mathbf{u}_{it} may be potentially affected by cross-sectional dependence. The presence of spatial lags should capture these effects and account for any bias stemming from omitted variables. Furthermore, since each spatial lag \tilde{X} is a linear combinations of the underlying untransformed series, it has the same order of integration as variable X . This has the important implication for the nonstationary case that the presence of spatial lags enlarges the cointegration space among tfp and the set of regressors \mathbf{z} to ensure stationarity of the residual term \mathbf{u}_{it} .

As pointed out in the seminal work of Engle and Granger (1987), cointegration and error correction are mirror images of each other. We may thus move from a specification of the long-run equation in equation (5) to a dynamic specification in first differences, which nevertheless preserves the information of the long-run equation. The resulting (Vector) error correction model ((V)ECM) describes the dynamic process through which cointegrated variables are driven in the adjustment process to their long-run equilibrium. In the following, we build on the concept proposed by Beenstock and Felsenstein (2010) and specify a spatial ECM (SpECM) as a dynamic process, in which time-space cointegrated variables comove over time. We allow for deviations from a stable long-run equilibrium relationship in the short run, however, the “error correction” mechanisms ensures the stability of the system in the long run.

Therefore, the SpECM concept may encompass three important types of cointegration: (i) if cointegration only applies within spatial units but not between them, we refer to “local” cointegration. The latter is the standard concept of cointegration with respect to (panel) time-series analysis; (ii) “spatial” cointegration refers to the case in which nonstationary variables are cointegrated over time between spatial

units but not within them. As Beenstock and Felsenstein (2010) point out, in this case, the long-term trends in spatial units are mutually determined and do not depend upon developments within spatial units; and (iii) finally, if nonstationary spatial panel data are both cointegrated within and between cross-sections over time, we refer to “global” cointegration.

The resulting SpECM associated with equation (5) in its first-order form can be written:

$$\begin{aligned} \Delta \text{tfp}_{i,t} = & \psi + \nu_i + \nu_t + \lambda_1 \Delta \text{tfp}_{i,t-1} + \lambda_2 \widetilde{\Delta \text{tfp}}_{i,t} + \lambda_3 \widetilde{\Delta \text{tfp}}_{i,t-1} + \lambda'_4 \Delta \mathbf{t}_{i,t-1} \\ & + \lambda'_5 \Delta \mathbf{f}_{i,t-1} + \lambda'_6 \widetilde{\Delta \mathbf{t}}_{i,t-1} + \lambda'_7 \widetilde{\Delta \mathbf{f}}_{i,t-1} + \eta_1 u_{i,t-1} + \eta_2 \tilde{u}_{i,t-1} + e_{i,t}, \end{aligned} \quad (6)$$

where Δ is the difference operator such that $\Delta \text{tfp}_{i,t} = \text{tfp}_{i,t} - \text{tfp}_{i,t-1}$; e_{it} is the short-run residual which is assumed to be temporally and spatially uncorrelated. The terms u_{it-1} and \tilde{u}_{it-1} are the (spatially weighted) residuals from the long-term relationships of the system. The coefficients for u and \tilde{u} can be interpreted as error correction coefficients, which drive the system to its long-run equilibrium state. Global error correction arises if the coefficients of these ec-terms, η_1 and η_2 , are nonzero. For the nested case of local error correction, we assume that $\eta_1 < 0$.

It is straightforward to see that if the coefficients for u and \tilde{u} are zero, the long-run information used for estimation drops out and the system in equation (6) reduces to a single equation in a spatial VAR (SpVAR) formulation. Note, that in the short run, \mathbf{z} may affect Δtfp differently from how it affects tfp in the long run. It is also important to note that the coefficient for the time lag of the dependent variable (λ_1) is typically expected to have the same sign as the coefficient for \tilde{u} (η_2), since the dynamics of tfp will be affected by short-run deviations from the cointegration path in neighboring regions. For the case of $\eta_1, \eta_2 \neq 0$, the resulting SpECM specification exhibits “global error correction.”

Since both equations (5) and (6) take the form of a (dynamic) spatial Durbin model (SDM), incorporating spatial lags of the endogenous and exogenous variables (Anselin 1988), an appropriate estimation strategy is to use maximum likelihood (ML) techniques. For panel data settings, Beer and Riedl (2009) have recently proposed an ML estimator for the SDM in a fixed-effects setting, which makes use of a (generalized Helmert) transformation proposed by Lee and Yu (2010). The authors show by means of a Monte Carlo simulation experiment that the SDM-ML estimator has satisfactory small-sample properties. We apply this estimator in the following.

Data and Stylized Facts

For the empirical analysis, we use panel data for the ten West German federal states between 1976 and 2008. Our data set comprises regional GDP, capital and labor inputs, export and import volumes, as well as inward and outward stocks of FDI. All data are used in real terms. For the analysis, all variables are transformed into

Table 1. Data Sources and Summary Statistics of the Variables (in Logs).

Variable	Description	Source	Obs.	In Logarithms			
				M	SD	Min	Max
<i>y</i>	Real GDP (in Euro)	VGR der Länder (VGRdL 2011)	330	11.041	1.193	8.190	13.284
<i>k</i>	Capital stock (in Euro)	VGR der Länder (VGRdL 2011)	330	13.067	1.015	11.076	14.643
<i>l</i>	Total employment	VGR der Länder (VGRdL 2011)	330	7.554	0.992	5.885	9.069
<i>tfp</i>	Total factor productivity	Own calculation	330	1.613	0.587	0.136	2.481
<i>ex</i>	Real exports (in Euro)	Destatis (2011)	330	9.771	1.155	7.193	12.119
<i>im</i>	Real imports (in Euro)	Destatis (2011)	330	9.863	1.051	7.379	12.235
<i>fdi out</i>	Real stock of outward FDI (in Euro)	Deutsche Bundesbank (2011)	330	8.510	2.114	3.002	12.626
<i>fdi in</i>	Real stock of inward FDI (in Euro)	Deutsche Bundesbank (2011)	330	8.337	1.644	5.305	11.912

logarithms.² The sources and summary statistics of the data are given in Table 1. Additionally, Figure 1 plots the time evolution of each variable for the ten West German federal state. All variables show to grow over time throughout the period 1976–2008. The correlation coefficient between GDP and TFP is sufficiently high (0.59), where the TFP variable is derived from a production function model according to equation (4).³

An issue that deserves careful attention when modeling in a spatial regression approach is the use of an appropriate spatial weighting matrix (Stakhovych and Bijmolt 2009; Elhorst 2010). Typically, border- or distance-based matrices are used. As a benchmark case, we use a spatial weighting scheme that contains binary information on whether two states i and j share a common border ($w_{ij} = 1$) or not ($w_{ij} = 0$). The spatial weighting matrix is used in its row-normalized form. In order to check for the stability of our results, we additionally use further weighting matrices, which define pairwise weights based on inverted distances as $w_{ij} = (1/D_{ij})$, where D_{ij} is the distance in road kilometers averaged over pairs of major cities in each state. Finally, as an alternative concept we also use interregional economic linkages based on goods transport flows to define pairwise weights among states. Since a total measure of interregional trade flows among German regions is not available, railway transportation statistics may serve as a proxy for the former. We use data from 1970 to ensure that the observed interregional linkages are exogenous to our estimation system (see the Appendix A for details).

The latter transport-based weighting matrix is used both in absolute terms and in terms of per capita intensities in order to control for the different absolute size of states. An important motivation for using transport-based weighting schemes is that we are able to give a more straightforward economic interpretation regarding the estimation results. That is, for instance, consider a negative correlation of the neighboring regions' import performance with regional GDP evolution. Opening up for international trade in terms of increased import activity may lead to a substitution effect of intranational forward and backward linkages. Regional supply is thus substituted by its neighbors through international import flows. This, in turn, may slow down economic development in the region under study and can motivate a negative spatial spillover effects. The advantage of transport weighting schemes is that they allow for a sound economic interpretation.

Finally, before proceeding with the estimation approach, we want to investigate the time-series properties of the variables involved. The graphical presentation in Figure 1 already gives an indication that all level variables have a trend pattern and are likely to be nonstationary. To analyze this in more depth, we compute panel unit root tests proposed by Im, Pesaran, and Shin (2003) as well as Pesaran (2007). The latter test has the advantage that it is more robust to cross-sectional correlation brought in by spatial dependence (see, e.g., Baltagi, Bresson and Pirotte 2007.), while the Im, Pesaran, and Shin (2003) test is found

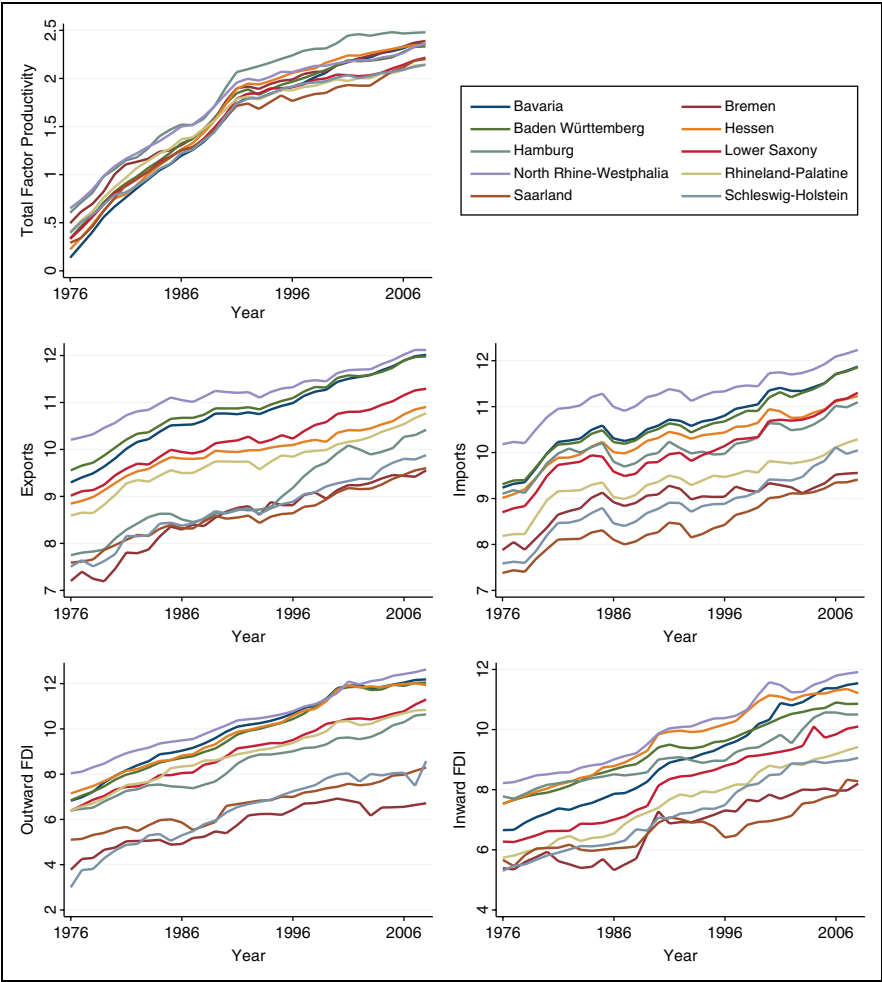


Figure 1. Time evolution of total factor productivity (TFP), trade, and foreign direct investment (FDI) by German states (in logs). Source: See Table 1.

to be oversized, when the spatial autocorrelation coefficient of the residual is large. The results of both panel unit root tests for variables in levels and first differences are reported in Table 2. As the results show, both test statistics give strong evidence that all variables are integrated of order $I(1)$ in their level specification and are tested to stationary after taking first differences. The same holds for the spatial lag transformations underlining the argument made by Beenstock and Felsenstein (2010) that a spatial lag as linear combination of the underlying untransformed series is typically of the same order of integration.

Table 2. Panel Unit Root Tests.

Variable	IPS test for $N = 10, T = 33$			CADF test for $N = 10, T = 33$		
	$W[t\text{-bar}]$	p Value	Av. Lags	$Z[t\text{-bar}]$	p Value	Av. Lags
y	-1.36*	.09	0.70	1.07	.85	
tfp	-1.01	.16	0.70	1.43	.92	
ex	1.36	.91	0.30	-1.21	.11	
im	1.08	.86	0.20	0.87	.81	
fdi out	0.28	.61	0.60	-0.19	.42	
fdi in	1.08	.86	1.00	-0.16	.43	
\tilde{y}	-0.98	.16	0.50	-0.55	.29	
\tilde{tfp}	-0.29	.38	0.60	-0.95	.17	
\tilde{ex}	1.65	.95	0.00	-1.83**	.03	
\tilde{im}	1.55	.93	0.00	2.17	.98	
\tilde{fdi} out	0.28	.61	1.40	1.09	.86	
\tilde{fdi} in	5.04	.99	1.00	-2.04**	.02	
Δy	-4.33***	.00	0.20	-5.17***	.00	
Δtfp	-4.11***	.00	0.40	-5.22***	.00	
Δex	-10.48***	.00	0.70	-8.77***	.00	
Δim	-10.54***	.00	0.50	-6.48***	.00	
Δfdi out	-10.64***	.00	0.30	-5.21***	.00	
Δfdi in	-11.91***	.00	0.30	-7.16***	.00	
$\Delta \tilde{y}$	-2.90***	.00	0.00	-5.62***	.00	
$\Delta \tilde{tfp}$	-2.93***	.00	0.00	-6.12***	.00	
$\Delta \tilde{ex}$	-9.90***	.00	1.00	-8.12***	.00	
$\Delta \tilde{im}$	-10.17***	.00	1.00	-4.81***	.00	
$\Delta \tilde{fdi}$ out	-7.82***	.00	0.60	-4.86***	.00	
$\Delta \tilde{fdi}$ in	-9.80***	.00	0.70	-8.41***	.00	

Note: ***, **, and * denote significance at the 1, 5, and 10 percent levels. For IPS, the optimal lag length is chosen according to the AIC. H_0 for both panel unit root test states that all series contain a unit root. Spatial lags are constructed using the common border-based weighting scheme.

Empirical Results

First, we estimate the long-run cointegration equation for TFP, trade, and FDI. The results for the SDM-ML model with different spatial weighting matrices are shown in Table 3. All model specifications report a positive and statistically significant coefficient for the region i 's own export and outward FDI performance on TFP. While the region's own import and inward FDI activity are found to generally add little explanatory power, the models show significant correlations between TFP and the spatial lag variables. That is, while all specifications find a significant negative coefficient for inward FDI, with the exception of the transport per population weighting scheme all models report significant positive correlation between the spatial lag of outward FDI activity and TFP. Similarly, the coefficient of the spatial lag of export volumes is found to be statistically

Table 3. SDM-ML Estimates for TFP, Trade and FDI with Alternative Spatial Weighting Schemes.

Type of W	Border	Distance	Transport	Transport/Population
$ex_{i,t}$	0.521*** (0.080)	0.166*** (0.028)	0.168*** (0.027)	0.409*** (0.034)
$im_{i,t}$	-0.105 (0.091)	-0.089*** (0.031)	-0.022 (0.031)	-0.062 (0.045)
$fdi\ out_{i,t}$	0.227*** (0.050)	0.065*** (0.014)	0.061*** (0.013)	0.184*** (0.019)
$fdi\ in_{i,t}$	-0.055 (0.043)	0.024* (0.013)	0.031** (0.013)	0.020 (0.020)
$\widetilde{ex}_{i,t}$	0.032 (0.043)	-0.163*** (0.047)	-0.131** (0.053)	-0.210*** (0.040)
$\widetilde{im}_{i,t}$	-0.129*** (0.037)	0.066 (0.041)	-0.049 (0.044)	-0.001 (0.021)
$\widetilde{fdi\ out}_{i,t}$	0.221*** (0.026)	0.076*** (0.028)	0.081** (0.034)	0.033 (0.027)
$\widetilde{fdi\ in}_{i,t}$	-0.099*** (0.024)	-0.079*** (0.024)	-0.088*** (0.027)	-0.067*** (0.021)
$\widetilde{tfp}_{i,t}$	-0.236*** (0.020)	0.790*** (0.026)	0.816*** (0.021)	0.356*** (0.021)
<i>N</i>	330	330	330	330
<i>R</i> ²	0.96	0.96	0.94	0.94
Log likelihood	589.01	505.83	508.11	377.96

Note: ***, **, and * denote significance at the 1, 5, and 10 percent levels. Standard errors in parentheses. *R*² defined as squared correlation between *tfp* and *tfp*.

significant negative in three out of four specifications. The same holds for the spatial lag of the endogenous variable, which shows a positive correlation among regions in all cases except for the border-based weighting scheme.⁴

However, one has to note that the regression parameters for the set of explanatory variables from equation (5) cannot be interpreted directly as marginal effects. As LeSage and Pace (2009) have recently shown, for model specifications including a spatial lag of the endogenous variable, economic interpretation is more complex. Table 4 therefore additionally computes summary measures for the SDM-ML based on a decomposition of the average total effect into its direct and indirect elements. The direct effect thereby measures the impact of changes in region *i*'s outcome variable for a change in the value of regressor *x* for region *i*, the indirect effect arises from spatial spillovers to region *i*'s outcome variable due to the averaged change in the observations of all neighboring regions *j* (for details, see LeSage and Pace 2009). Statistical inference is done based on simulation runs used to create distributions for the effect measures.⁵

Since all variables are specified as logarithms, the reported marginal effects can be interpreted as elasticities.⁶ The results obtained confirm the correlations from

Table 4. Direct, Indirect, and Total Long-Run Effect of Variables in TFP Model.

Type of W	Border	Distance	Transport	Transport/ Population
Direct				
$ex_{i,t}$	0.581*** (0.104)	0.153*** (0.034)	0.173*** (0.038)	0.378*** (0.039)
$im_{i,t}$	-0.026 (1.182)	-0.091*** (0.034)	-0.056 (0.036)	-0.072 (0.044)
$fdi\ out_{i,t}$	0.109 (0.067)	0.120*** (0.017)	0.123*** (0.020)	0.231*** (0.023)
$fdi\ in_{i,t}$	0.004 (0.051)	-0.001 (0.012)	0.001 (0.167)	-0.008 (0.023)
Indirect				
$ex_{i,t}$	-0.135* (0.079)	-0.134 (0.187)	0.038 (0.257)	-0.071 (0.044)
$im_{i,t}$	-0.135*** (0.064)	-0.025 (0.143)	-0.348** (0.170)	-0.024 (0.032)
$fdi\ out_{i,t}$	0.252*** (0.050)	0.560*** (0.099)	0.661*** (0.159)	0.108*** (0.028)
$fdi\ in_{i,t}$	-0.128*** (0.042)	-0.259** (0.101)	-0.319** (0.134)	-0.065*** (0.023)
Total				
$ex_{i,t}$	0.445*** (0.063)	0.018 (0.207)	0.211 (0.287)	0.308*** (0.072)
$im_{i,t}$	-0.186*** (0.066)	-0.116 (0.159)	-0.405** (0.188)	-0.096 (0.059)
$fdi\ out_{i,t}$	0.362*** (0.038)	0.681*** (0.111)	0.785*** (0.174)	0.339*** (0.043)
$fdi\ in_{i,t}$	-0.123*** (0.037)	-0.261** (0.114)	-0.318** (0.149)	-0.074* (0.040)

Note: ***, **, and * denote significance at the 1, 5, and 10 percent levels based on partial derivatives and parameter simulations as described in LeSage and Pace (2009). Simulated standard errors are given in parentheses.

Table 3 and report a significant direct effect of export flows on the regional TFP level in all model specifications. This finding can be directly associated with the commonly known trade-led growth hypothesis. The estimated output elasticity for TFP varies between 0.15 percent and 0.6 percent, given a 1 percent change in regional export volumes. Except for the border-based specification, all models also show a significant direct effect of outward FDI activity and thus confirm the FDI-led growth hypothesis. The estimated elasticity is somewhat smaller than in the export case and ranges from 0.12 percent to 0.23 percent. Although all models report a negative direct effect of import activity on TFP, this effect turns out to be statistically significant only in the distance-based specification.

Looking at the significance and size of spillovers from neighboring regions, all models find a strong and positive effect for outward FDI, while inward FDI is associated with negative spillovers. The same holds for import activity. Both effects hint at reallocation activity stemming from increased inward-oriented internalization activity in Germany, which leads to a substitution of regional input–output linkages by international ones. On the other hand, there is a strong correlation between the regions' external internationalization activity, which leads to a positive technology transfer both in the host region and in its neighborhood. Summing up both effects to a measure of the total marginal impact, this shows that the TFP level is predominantly driven by FDI activity, positive in the case of outward FDI stock, negative for inward FDI. The results for trade activity hint at the same direction, although the reported statistical significance is mixed. Nevertheless, all specifications show a positive coefficient for the trade effect, while the output elasticity of import activity is generally negative.

A general result obtained from the long-run estimation is that the inclusion of spatial lags helps to significantly improve the fit of the different models and shows more clear-cut results regarding a stable cointegration relationship among the variables. For the full model including all variables and their spatial lags, the null hypothesis of no cointegration can be strongly rejected by the Pedroni (1999) test (Panel ρ statistic). Here, the test statistic is 3.24 (p value = .00) versus 1.77 (p value = .11) in the aspatial case, showing only weak support for a stable cointegration vector with stationary residuals when we discard the spatial lags.⁷ Our results can thus be interpreted along the lines of Beenstock and Felsenstein (2010), who find that the inclusion of spatial lags may be of vital importance to ensure a stable cointegration relationship in regional economic models.

We then move on and use the information obtained from the long-run cointegration relationship in a SpECM framework for regional TFP growth to analyze the short-run correlation between technology growth and internationalization activity. The estimation results of the SpECM are shown in Table 5. Here, the inclusion of time and spatial lags of the endogenous variable results in a more complex specification. We use a “time-space-simultaneous” model (see, e.g., Anselin, Le Gallo, and Jayet 2008), which includes a contemporaneous spatial lag of the endogenous variable and a one period time lag.⁸ The results generally show that the short-run dynamics of TFP growth is mainly driven locally and are therefore different from long-run impacts. Looking at the regression coefficients describing the short-run correlation among the variables, we see that regional TFP growth is mainly driven by short-run movements in outward and inward FDI. For both variables, the estimation results show a positive and statistically significant regression coefficient. We also obtain some evidence for a positive short-run correlation between changes in the regional export volumes and TFP growth, as well as a negative relationship between short-run changes in the spatial lag of import flows and TFP growth.

For the included error correction term ($u_{i,t-1}$), we get a significant and negative regression parameter supporting the findings of the Pedroni (1999) cointegration

Table 5. Spatially Augmented Short-Run Estimates of TFP Growth Driven by Trade and FDI.

Type of W	Space-Time Simultaneous			
	Border	Distance	Transport	Transport/ Population
$\Delta tfp_{i,t-1}$	0.383*** (0.093)	0.134** (0.056)	0.111** (0.055)	0.301*** (0.053)
$\Delta ex_{i,t}$	0.044 (0.044)	0.051** (0.021)	0.051** (0.020)	0.045* (0.024)
$\Delta im_{i,t}$	0.054 (0.040)	-0.022 (0.025)	-0.010 (0.025)	0.010 (0.022)
$\Delta fdi\ out_{i,t}$	0.037** (0.018)	0.019** (0.008)	0.013 (0.008)	0.033*** (0.010)
$\Delta fdi\ in_{i,t}$	0.055*** (0.017)	0.017** (0.008)	0.017** (0.008)	0.024*** (0.009)
$u_{i,t-1}$	-0.140*** (0.053)	-0.077*** (0.026)	-0.063*** (0.022)	-0.049*** (0.017)
$\widetilde{\Delta tfp}_{i,t}$	-0.237*** (0.021)	0.694*** (0.039)	0.693*** (0.035)	0.338*** (0.024)
$\widetilde{\Delta ex}_{i,t}$	0.021 (0.021)	-0.013 (0.034)	-0.022 (0.035)	-0.001 (0.250)
$\widetilde{\Delta im}_{i,t}$	-0.008 (0.015)	0.019 (0.030)	0.014 (0.030)	-0.005 (0.014)
$\widetilde{\Delta fdi\ out}_{i,t}$	0.017* (0.010)	-0.003 (0.016)	0.007 (0.017)	-0.006 (0.011)
$\widetilde{\Delta fdi\ in}_{i,t}$	0.054*** (0.010)	0.034** (0.015)	0.038** (0.015)	0.013 (0.009)
$\tilde{u}_{i,t-1}$	-0.075* (0.039)	0.044 (0.107)	0.151** (0.076)	0.047 (0.043)
N	310	310	310	310
R^2	0.45	0.52	0.56	0.50
Z_{STMI}	-1.236	-0.569	-0.940	-1.033
$p\ value$.216	.568	.347	.301

Note: ***, **, and * denote significance at the 1, 5, and 10 percent level. Standard errors in parentheses. R^2 defined as squared correlation between Δtfp and $\widetilde{\Delta tfp}$. Z_{STMI} is the z statistic of the spatiotemporal version of Moran's I .

test, that is, local error correction is in order, although with a rather small adjustment speed. The regression results for the coefficient for the spatialized ec term ($\tilde{u}_{i,t-1}$) are highly mixed and statistical significance is weak. Only for the transport-based weighting scheme we get a significant coefficient. As argued by Beenstock and Felsenstein (2010), it also found to be in line with the coefficient sign of the spatial lag of the endogenous variable. The positive coefficient indicates that temporary deviations from the stable long-run relation between TFP and internationalization activity in neighboring regions foster region i 's TFP growth in the short run. However, the

regression results in the other three specifications hint at the absence of such interregional spillovers stemming from the short-run adjustments to long-run equilibrium.

Moreover, we find strong evidence for the role of autoregressive adjustment processes measured by $\Delta \text{tfp}_{i,t-1}$ as well as significant positive coefficient for the contemporaneous spatial lag of the endogenous variables. The latter result mirrors earlier findings for German regional growth, which shows a positive intra-national growth nexus among spatial neighbors (see, e.g., Niebuhr 2000 as well as Eckey, Kosfeld, and Tuerck 2007). Testing for remaining cross-section correlation in the error term of the SpECM by means of the spatiotemporal extension of Moran's I (STMI, see Lopez et al. 2011), Table 5 shows that the time-space-simultaneous specification is well equipped to capture all underlying spatial patterns in the data.⁹

Finally, Table 6 computes the direct, indirect, and total effects of the SpECM in order to interpret the regression results regarding the significance of growth determinants and spatial spillovers. To do so, we first derive the reduced form of the model by multiplying both sides of equation (6) with $(I_N - \lambda_2 \mathbf{W})^{-1}$, where I_N is an identity matrix of order N , so that the model is only written as a function of predetermined endogenous and exogenous variables. From this transformation, we can compute the direct, indirect, and total time-contemporaneous impact for each exogenous variable in analogy to the long-run cointegration equation. We can then compute the final form of equation (6) by multiplying it with $(1 - \lambda_1)^{-1}$. One has to note that this approach is based on the assumption of time-space separability, that is, we assume that spatial processes are fully processed through the system for each period of time observation such that there are no time-space covariances. This in fact is a simplification of a more complex time-space structure as, for instance, discussed in Debarsy, Ertur, and LeSage (2011). Since most of the spatial lag terms were estimated to be statistically insignificant, this simplification should still give a reasonably good approximation of the underlying dynamic process.

As pointed out by Debarsy, Ertur, and LeSage (2011), it is also possible to track the temporal distribution of the temporary and cumulated impact (both aspatial and spatial) over time as known from standard multiplier analysis in the time-series literature (Lütkepohl 2005). For each period s , we thus calculate the interim multiplier D_s as $D_s = (\lambda_1)^s \times \mathbf{B}^{-1}$, where $\mathbf{B} = (I_N - \lambda_2 \mathbf{W})$. For $s = 0$, we use the coefficient values from the reduced form specification according to Table 6. The total cumulated effect up to period S can then be calculated as $\sum_{s=0}^S D_s$.¹⁰

Table 6 thus shows the contemporaneous direct, indirect, and total effects of the SpECM's reduced form specification. The results confirm the observation that TFP growth is mainly driven by outward and inward FDI, where the latter variable has both positive direct and indirect spatial spillover effects. Although inward FDI is found to be negatively correlated with TFP in the long run, short-run changes are found to have a positive temporary effect on the evolution of technical progress. Export growth is

Table 6. Contemporaneous Direct, Indirect, and Total Short-Run Effects in TFP Model.

Type of W	Border	Distance	Transport	Transport/ Population
Direct				
$ex_{i,t}$	0.035 (0.059)	0.057** (0.022)	0.054** (0.023)	0.052** (0.025)
$im_{i,t}$	0.069 (0.054)	-0.022 (0.024)	-0.007 (0.021)	0.009 (0.022)
$fdi\ out_{i,t}$	0.032 (0.024)	0.021** (0.010)	0.017* (0.010)	0.035*** (0.012)
$fdi\ in_{i,t}$	0.026*** (0.022)	0.029*** (0.009)	0.031*** (0.009)	0.034*** (0.010)
Indirect				
$ex_{i,t}$	0.016 (0.042)	0.069 (0.098)	0.036 (0.013)	0.018 (0.024)
$im_{i,t}$	-0.032 (0.034)	0.008 (0.058)	0.022 (0.065)	-0.003 (0.015)
$fdi\ out_{i,t}$	0.012 (0.018)	0.029 (0.052)	0.051 (0.059)	0.004 (0.012)
$fdi\ in_{i,t}$	0.061*** (0.017)	0.143*** (0.040)	0.154*** (0.040)	0.025** (0.010)
Total				
$ex_{i,t}$	0.052 (0.035)	0.126 (0.109)	0.091 (0.111)	0.070* (0.039)
$im_{i,t}$	0.037 (0.027)	-0.013 (0.064)	0.014 (0.068)	0.006 (0.026)
$fdi\ out_{i,t}$	0.044*** (0.014)	0.051 (0.059)	0.069 (0.065)	0.039* (0.021)
$fdi\ in_{i,t}$	0.088*** (0.014)	0.172*** (0.044)	0.186*** (0.045)	0.059*** (0.014)

Note: ***, **, and * denote significance at the 1, 5, and 10 percent levels based on partial derivatives and parameter simulations as described in LeSage and Pace (2009). Simulated standard errors are given in parentheses.

shown to have a positive direct effect in most specifications as well. To illustrate the short-run effect of changes in the exogenous variables on TFP growth, Figure 2 gives a graphical presentation of their interims multipliers.¹¹ As the Figure shows, export and outward FDI are mainly driven by its direct effect. For inward FDI, the indirect spatial spillover effect is almost of equal size as the direct effect, while in the case of import flows the positive direct and negative indirect effect offset each other. In general, all effects fade out after five years at the most, indicating a rapid absorption of short-run shocks.

Finally, to answer the politically relevant question whether internationalization activity is altogether beneficial or damaging for regional TFP, we compute net direct and indirect trade effects in the long- and short-run by aggregating the individual

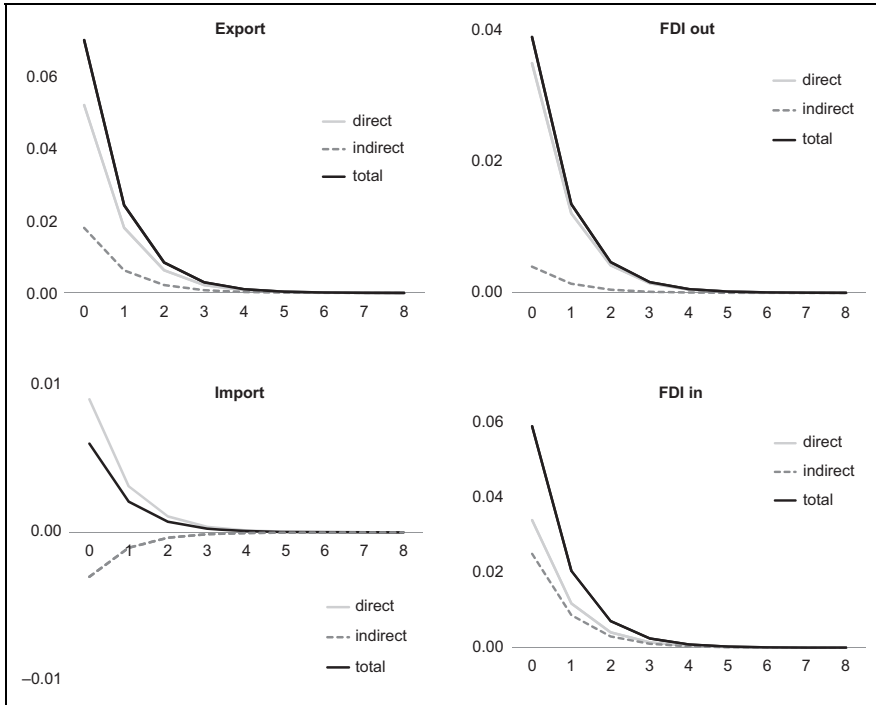


Figure 2. Interims multiplier for dynamic short-run effects of internationalization activity on total factor productivity (TFP) growth. *Note:* Multipliers and the different effects are defined in the text. Spatial weights based on transport intensities.

effects over all four variables. We therefore take the sample mean of each of the explanatory variables (both in levels and first differences) and calculate the variable's individual absolute contributions by multiplying the sample mean with the obtained coefficients from Tables 4 and 6. Then, by summing up the obtained absolute contributions to the total net effect and normalizing it to 100 percent, we can compute the relative importance of the direct and indirect elements as a share of the total effect. The results for the case of long- and short-run specifications based on the transport intensities as weighting scheme are shown in Figure 3. The figure shows that in the short run, both effects are positive. Here, the net direct effect measured over all four variables amounts to roughly 75 percent, whereas one-fourth of the total effect stems from spatial spillovers. For the long-run equation, we find that the overall positive net effect is driven by the sum of the direct effects, while—collectively—the sum of the individual direct effects is slightly negative. The latter relationships shows that—measured in absolute terms—the positive effect from outward FDI is outweighed by the negative spillovers induced from inward FDI and



Figure 3. Short- and long-run “net” effects of internationalization activity on regional total factor productivity (TFP; $I = 100$ percent). *Note:* Effects calculated for the regression results using transport intensities in Tables 4 and 6.

trade activity. As a result of these two opposing effects, the relative contribution of the spillover part to the total effect is rather small.

Conclusion

This article has examined the role of direct and indirect spillover effects for the German regional TFP–trade–FDI nexus. We have used a combination of time-series analysis as well as spatial econometric tools to tackle this issue. While the empirical analysis of long-run comovements among nonstationary variables is by now a common standard in time-series econometrics, less attention has been paid to the importance of spatial lags in the long-run formulation of such a regression model. Applying the novel concept of global cointegration recently proposed by Beenstock and Felsenstein (2010) enables us to estimate SpECM for West German data between 1976 and 2008.

Our results show that both direct and indirect spatial links between the variables matter when tracking their long-run comovement. In particular, export flows show a significant and positive long-run impact on the level of TFP, supporting the export-led growth hypothesis from regional and international economics. Furthermore, we also find evidence for long-run FDI-led growth and identify negative spillovers from inward FDI and import activity. The latter results may hint at the potential role played by substitution effects, when intranational input–output linkages are scaled down in the process of enhanced trade integration and inward-directed FDI activity. Similar to Beenstock and Felsenstein (2010), we can conclude that the inclusion of

spatial lag of the endogenous variable and set of regressors has important implications for the stability of a cointegration relationship among variables for a regional economic system. Residual-based cointegration tests strongly reject the null of no cointegration among the set of variables including their spatial lags.

For the short-run determinants of TFP growth, we observe that this process is mainly driven locally rather than globally. We find a negative and statistically coefficient for the error correction term of the SpECM, which gives a further indication that the system is dynamically driven to its stable long-run value. In this case, besides an autoregressive component and the spatial lag of the endogenous variable, outward and inward FDI stocks in particular are found to drive technical progress in the short run. Inward FDI is also found to have significant positive spillover effects.

Regarding the choice of the spatial weighting scheme, we employ geographical information (common border and road distance–based specifications) as well as individual weights stemming from interregional goods transport flows. Our results are generally robust regarding the choice of the spatial weighting matrix. The advantage of (exogenous) transportation data is that the estimated effects can be better interpreted in the context of economic theory. That is, for instance, the negative long-run spillovers stemming from increasing regional import flows can be seen as a substitution effect for intranational input–output linkages affecting the transmission channels of diffusion in technical progress. Summing up the four variables to get direct and indirect net effects of internationalization activity, we find that the direct effect is always positive, while the indirect effect is positive in the short run but slightly negative in the long-run equation.

Appendix A

Road Distances and Transport Flows among West German States

Table A1. Road Distances between West German States (in Kilometers).

From./To:	SH	HH	NIE	BRE	NRW	HES	RHP	BW	BAY	SAAR
SH	0	76	272	192	447	510	629	732	745	754
HH	76	0	177	110	363	473	553	667	666	690
NIE	272	177	0	130	265	310	427	529	541	544
BRE	192	110	130	0	273	424	483	633	650	590
NRW	447	363	265	273	0	234	251	410	501	349
HES	510	473	310	424	234	0	163	231	308	255
RHP	629	553	427	483	251	163	0	207	339	146
BW	732	667	529	633	410	231	207	0	262	226
BAY	745	666	541	650	501	308	339	262	0	378
SAAR	754	690	544	590	349	255	146	226	378	0

Source. Average distances in road kilometers among major cities per state calculated with www.map24.de.

Table A2. Interregional Railway Transportation Flows in 1970 (in 1,000 Tons).

From:/To:	SH	HH	NIE	BRE	NRW	HES	RHP	BW	BAY	SAAR	Total
SH	966	176	679	94	340	102	63	206	289	9	2,924
HH	321	896	2,297	374	933	342	118	361	747	27	6,416
NIE	1,303	1,033	20,434	1,593	7,288	1,465	391	890	2,140	3,726	40,263
BRE	42	61	3,182	3,158	1,449	386	193	420	600	111	9,602
NRW	2,064	2,191	11,056	4,705	1,02,530	5,114	3,271	4,821	7,737	2,064	1,45,553
HES	195	491	958	517	1,823	4,512	782	942	1,583	158	11,961
RHP	181	177	618	231	2,013	895	2,337	2,916	1,722	1,097	12,187
BW	68	308	305	254	961	852	696	10,853	2,711	517	17,525
BAY	143	468	578	364	1,644	813	451	2,225	19,349	145	26,180
SAAR	21	108	181	276	659	407	774	1,471	898	7,761	12,556
Total	5,304	5,909	40,288	11,566	1,19,640	14,888	9,076	25,105	37,776	15,615	2,85,167

Source: Statistisches Bundesamt, Fachserie H, R 4, Eisenbahnverkehr, 1970.

Note: For the design of spatial weighting matrices, the diagonal elements have been set to zero.

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Notes

1. An alternative strand of the empirical literature, regressing trade openness on regional concentration measures is discussed in Brühlhart (2011). Focusing mainly on developing countries, these contributions generally reveal a positive correlation between trade openness and spatial concentration of economic activity.
2. It would be desirable to have a higher degree of regional disaggregation rather than $N = 10$ with $T = 33$, however, no such data on trade and FDI activity are available. The panel structure of the data is nevertheless still comparable to Beenstock and Felsenstein (2010) with $N = 9$ and $T = 18$, so that it should be feasible to apply their proposed method to our regional data.

3. The regression coefficients come close to standard parameter values reported in the literature with $\alpha = 0.38$ and $\beta = 0.60$. We use both an unrestricted specification and a restricted Cobb-Douglas form with $\beta = (1 - \alpha)$. Detailed results can be obtained from the author upon request.
4. The opposite result for the binary border-based matrix may stem from its (too) strict neighborhood identification.
5. Computations are performed in Matlab using the code by Beer and Riedl (2009) based on the spatial econometrics toolbox by LeSage (2010).
6. As pointed out by one referee, causal inference in a single equation framework has to be done very carefully given the reversed causality problem. In fact, one has to assume exogeneity of right-hand side regressors to do so, whereas a full account of the causal impact of each variable can only be done in a system approach. However, the latter approach goes beyond the scope of this article and dealing with this issue is left for future research.
7. For the computation of the Panel p statistic, I include individual intercepts for each cross-sectional unit but no individual time trends. The optimal lag length is chosen by means of the Schwartz information criterion. Similar results also hold for the Fisher-type Johansen cointegration test as reported in chapter 7 of Mitze (2012).
8. We thus assume that λ_3 in equation (6) is zero, which greatly simplifies the interpretation of the model results. We also tested for the statistical significance of λ_3 in a general “space–time dynamic” model framework, however, the coefficient turned out to be statistically insignificant in most specifications. Detailed results can be obtained upon request.
9. We used both asymptotic and bootstrap-based versions of the STMI. A description of the latter is given in chapter 7 of Mitze (2012).
10. For the general discussion of the concept of time–space multiplier analysis, see Debarsy, Ertur, and LeSage (2011).
11. Here, we only focus on the distribution of the short-run effect, disregarding from the long-run cointegration equations. Computations are done for the spatial weighting scheme based on transport intensities.

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