The Impact of Interaction Effects among Neighbouring Countries on Financial Reform: A Dynamic Spatial Panel Data Approach

J. Paul Elhorst†, Eelco Zandberg* and Jakob De Haan§

† Faculty of Economics and Business, University of Groningen, PO Box 800, 9700 AV Groningen, The Netherlands
Email: j.p.elhorst@rug.nl

*Faculty of Economics and Business, University of Groningen, PO Box 800, 9700 AV Groningen, The Netherlands and Netspar, Tilburg, The Netherlands
Email: e.d.zandberg@rug.nl

§ De Nederlandsche Bank, Amsterdam, The Netherlands, and Faculty of Economics and Business, University of Groningen, PO Box 800, 9700 AV Groningen, The Netherlands and CESifo, Munich, Germany
Email: Jakob.de.Haan@rug.nl (corresponding author)

Using data from Abiad et al. (2008), we estimate a dynamic spatial panel data model with country and time-period fixed effects reformulated in first-differences to account for non-stable components in the financial liberalization index, to examine financial reform interaction effects, and to test for spatial cointegration. By comparing the performance of different specifications of the spatial weights matrix describing the spatial arrangement of the countries in the sample, we find that the popular regional leader matrix must be rejected in favour of an inverse distance matrix with a cut-off point at 3000 km. Using this matrix, we determine the extent to which a change in a single explanatory variable in a particular country affects financial reform in other countries.

Key words: Financial reform, financial liberalization, spatial dependence, spatial cointegration

JEL Classifications: C21, G3

The views expressed in this paper do not necessarily reflect the views of De Nederlandsche Bank.
1. Introduction

One of the most important developments over the past four decades has been the spread of liberal economic ideas and policies throughout the world. These policies have affected the lives of millions of people, yet our most sophisticated economic models do not adequately capture influences on these policy choices. In this respect, one of the most telling findings of the seminal paper of Abiad and Mody (2005) is that financial reform in the leading country of a region (continent) affects financial reform in other countries, though without considering potential feedback effects. Financial reform is defined as a change towards greater private provision of financial services under fewer operational restrictions, while financial liberalization captures the extent to which the financial system is free of government intervention.

Various theoretical arguments have been put forward to explain the existence of spatial spillover effects among countries. Building on previous studies, Simmons and Elkins (2004) distinguish two broad classes of models. First, economic policy elsewhere can alter the payoffs associated with choosing or maintaining a particular policy. For instance, if financial reform in a particular country attracts more foreign direct investment or trade, other countries may feel competitive pressure to match these policies. Conversely, countries that resist financial reforms may face reputational consequences that cast doubt on their economic policies and potentially the legitimacy of their governance. Second, economic policy elsewhere can change the information set on which governments base their policy decisions. Since countries often lack the crucial information they need to understand the consequences of financial reform, they might learn from the experience of other countries.

To measure the magnitude of spatial spillover effects, Abiad and Mody (2005) use a liberalization index $FL$ that consists of six items each measured on a scale from 0 to 3 (fully repressed to fully liberalized), hence the overall index ranges from 0 to 18. Next, they adopt an ordered logit model to explain the change in the financial liberalization index, $\Delta FL$ (i.e. financial reform). The reason to adopt this particular model is that the reform process in most countries is characterized
by long periods of status quo; 76.2% of the observations do not change if $FL$ is measured on an annual basis. Finally, they add the difference between a country’s degree of financial liberalization and that of the regional leader ($REG_{FL} - FL$), i.e., the country in the region with the highest level of financial liberalization, as one of the explanatory variables of financial reform. They label this variable the regional diffusion effect. The marginal effect of this variable appears to be 0.44 to 0.55 if country-fixed effects are controlled for.\(^1\)

Over the past ten years there has been a growing literature, both empirical and theoretical, dealing with interaction effects among economic units in space. While this increase in interest is relatively recent, spatial econometric modelling has a long history in the regional science and geography literature. In an overview paper on thirty years of spatial econometrics, one of the founding fathers, Luc Anselin (2010), concludes that spatial econometrics has now reached a stage of maturity because of the general acceptance of spatial econometrics as a mainstream methodology and because the number of applied empirical researchers that use econometric techniques in their work sees a near exponential growth. In view of this literature,\(^2\) the setup of Abiad and Mody's study and, therefore, their main results and policy conclusions may be criticized.

The first problem is that it is not clear why a country’s financial reform should only be influenced by the degree of financial liberalization of the leading country in the region. Following Tobler’s (1970) first law of geography, according to which "everything is related to everything else, but near things are more related than distant things", a country’s financial reform may well be affected by the degree of financial liberalization of other countries too. Moreover, not only nearby countries but also countries located outside the region may matter, e.g., because these countries are important trade partners. Policy diffusion will be strongest among countries that are in close contact. In this respect, Simmons and Elkins

\(^1\) Measured as the coefficient of this variable times the probability that $FL$ changes.
\(^2\) See Anselin (2006), LeSage and Pace (2009), and Kelejian and Prucha (2010) for recent overviews of the spatial econometrics literature for cross-section data, and Anselin et al. (2008), Lee and Yu (2010), and Elhorst (2010) for spatial panel data.
(2004) point out that countries may learn from more countries than just the most successful ones.

A related problem is that the regional diffusion effects is treated as an exogenous explanatory variable, which implies that feedback effects are not accounted for. However, the regional leader may liberalize its financial system further, induced by the liberalization process in countries within the region, to maintain its lead over these countries (competition effect). In a star network, a popular structure in the literature on social networks (Bramoullé et al., 2009), individuals are not only connected to the leader (star), but the leader is also connected to all other individuals. Once such feedback effect are accounted for, the regional diffusion effect should be treated as an endogenous rather than an exogenous explanatory variable and the estimation technique should be adjusted accordingly.

A second and, up to now, relatively unexplored problem is that Abiad and Mody (2005) do not test whether the financial liberalization index has a spatial unit root and, related to that, whether the financial liberalization process is spatially cointegrated. Although the financial liberalization index, due to its construction, is bounded from below and above, Abiad and Mody (2005) observe an upward trend in liberalization in all income groups in the last quarter century. Except for time factors, this may be explained by the fact that economies all over the world are integrated with each other. To test for this, Lee and Yu (2010) recommend to estimate a dynamic spatial panel data model and to use the sum of the coefficient estimates of the dependent variables lagged in space, in time, and in both space and time to test whether the dependent variable is stable. If the dependent variable turns out to be non-stable, that is, if the sum of these coefficients is greater than one, they recommend to reformulate the model in spatial first differences. This is because the dynamic spatial panel data model will produce consistent parameter estimates only if the dependent variable is stable or does become stable by taking spatial first differences. Applications of dynamic spatial panel data models are still scarce. Some recent examples are applications to public infrastructure investment (Cohen and Morrison Paul, 2004), housing prices (Brady, 2009), consumption and
habit formation (Korniotis, 2010), and commuting (Parent and LeSage, 2010). However, neither of these empirical studies have considered non-stability and spatial cointegration before.

A third problem is that Abiad and Mody (2005) do not control for time-period effects. Generally, applied researchers tend to find weaker evidence in favour of spatial interaction effects when time-period fixed effects are accounted for (Elhorst, 2010). Using Monte Carlo simulation experiments, Lee and Yu (2010) report that ignoring time-period fixed effects leads to a large upward bias (up to 0.45) in the coefficient of the diffusion term. The explanation is that most variables tend to increase and decrease together in different spatial units over time (e.g., along the business cycle). If this common effect is not taken into account and thus not separated from the interaction effect among countries, the latter effect might be overestimated.

A final problem is that the level of financial liberalization of the regional leader hardly changes over time. Figure 1 plots financial liberalization for each regional leader identified by Abiad and Mody (2005). The figure demonstrates that there are two regions in which financial liberalization of the regional leader hardly changes over time, namely East Asia and the OECD. As a consequence, changes in the regional diffusion measure, $\text{REG}_\text{FL} - \text{FL}$, are mainly caused by changes in $\text{FL}$ observed in countries following the regional leader. Therefore, the term $\text{REG}_\text{FL} - \text{FL}$ will not only measure regional diffusion but also picks up a convergence effect.

To examine whether the financial liberalization index is stable and financial reform in one country affects reform in other countries, and vice versa, we apply advanced spatial econometric techniques rather than adding a simple regional diffusion term that lacks a firm econometric foundation, as in Abiad and Mody (2005). Recently, Huang (2009) questioned the benefits of the ordered logit approach applied in Abiad and Mody (2005). As the liberalization index is not a
cardinal variable, he argues that it is better to explain the level rather than the change in the liberalization index. In addition, he points out that the ordered logit approach imposes strong distributional assumptions relative to a linear model, and that the parameter estimates may be inconsistent because of an incidental parameter problem. Huang also criticizes Abiad and Mody (2005) for not taking cross-sectional dependence into account. By controlling for common trends and error dependence across countries and over time, following a methodology developed by Pesaran (2006), he finds that the coefficient estimate of the diffusion effect falls from a significant number of 0.06 to a significant number of -0.14. However, Huang does not offer an explanation for this rather striking sign change.

In view of the critique of Huang (2009) and the econometric methodology developed by Lee and Yu (2010), we start by taking the level of financial liberalization as the dependent variable and estimate the parameters of a dynamic spatial panel data model to investigate the properties of this variable in both space and time. For this purpose, we use new data on financial liberalization taken from Abiad et al. (2008) for 62 countries over the period 1976-2005. This is an updated and expanded version of the dataset used by Abiad and Mody (2005). As the financial liberalization index will turn out to be non-stable, we reformulate the model in spatial first differences. Next, the variables of this model are rearranged such that the change of the financial liberalization index of a particular country in a particular year becomes the left-hand side variable, just as in Abiad and Mody (2005). This approach will have the concomitant effect that the regional diffusion effect and convergence effects can be separated from each other. In addition, following LeSage and Pace (2009), we will derive the effects that each explanatory variable in the model has on financial reform in its home country and on financial reform in other countries (i.e., direct and indirect effects).

The remainder of the paper is structured as follows. Section 2 outlines our methodology. Section 3 describes the data used, and section 4 presents estimation results. Section 5 offers our conclusions.
2. A dynamic spatial panel data approach

2.1 The model

The dynamic spatial panel data model takes the form

\[
Y_t = \tau Y_{t-1} + \rho WY_t + \eta WY_{t-1} + X_t \beta + \mu + \alpha_{t0} 1_N + \varepsilon_t,
\]

where \( Y_t \) denotes a \( N \times 1 \) vector consisting of one observation for the financial liberalization index for every country in the sample (\( i=1, \ldots, N \)) at a particular point in time (\( t=1, \ldots, T \)), and \( X_t \) is an \( N \times K \) matrix of exogenous explanatory variables with associated response parameters \( \beta \) contained in a \( K \times 1 \) vector. \( \tau \), the response parameter of the lagged dependent variable \( Y_{t-1} \), is assumed to be restricted to the interval \((-1,1)\). \( \varepsilon_t = (\varepsilon_{t1}, \ldots, \varepsilon_{tN})^T \) is a vector of i.i.d. disturbance terms, whose elements have zero mean and finite variance \( \sigma^2 \). \( \mu = (\mu_1, \ldots, \mu_N)^T \) is a vector with country fixed effects \( \mu_i \). \( \alpha_{t0} \) is the coefficient of a time period fixed effect, and \( 1_N \) is a \( N \times 1 \) vector of ones. Country fixed effects control for all country-specific, time-invariant variables whose omission could bias the estimates, while time-period fixed effects control for all time-specific, country-invariant variables whose omission could bias the estimates in a typical time-series study (Baltagi, 2005).

The variables \( WY_t \) and \( WY_{t-1} \) denote contemporaneous and lagged endogenous interaction effects among the dependent variables. \( \rho \) is called the spatial autoregressive coefficient, while \( \eta \) might be labelled the lagged spatial autoregressive coefficient. \( W \) is a non-negative \( N \times N \) matrix of known constants describing the spatial arrangement of the countries in the sample. Its diagonal elements are set to zero by assumption, since no country can be viewed as its own neighbour. If \( W \) is row-normalized, \( \rho \) and \( \eta \) are defined on the interval \((1/\lambda_{\text{min}},1)\), where \( \lambda_{\text{min}} \) equals the most negative purely real characteristic root of \( W \) (LeSage and Pace, 2009, pp. 88-89).\(^3\) Anselin (2001) considers three simpler types of spatial dynamic models, namely pure space recursive (\( \rho = 0 \)), time-space recursive (\( \rho = 0 \)),

\(^3\) In addition to this, the eigenvalues of the matrix \([W(1-\rho W)^{-1}(\tau I + \eta W)]\) should lie within the unit circle, which further restricts the stationarity region on which the parameters \( \tau, \rho \) and \( \eta \) are defined.
and time-space simultaneous ($\eta=0$) models. However, since we do not wish to preclude any form of dependence in advance, we take model (1) as point of departure.

To estimate the parameters of this model, Lee and Yu (2010) consider the log-likelihood function of equation (1), taking into account the Jacobian term that reflects the endogeneity of the $W_{Y_{t}}$ variable, i.e., the fact that one country can affect another country, and vice versa. The estimator that is derived from this log-likelihood function is called the Least Squares Dummy Variables (LSDV) estimator (see Yu et al., 2008 for this terminology). However, by providing a rigorous asymptotic theory, Lee and Yu (2010) show that this LSDV estimator is biased when both the number of spatial units ($N$) and the number of time points ($T$) in the sample go to infinity such that the limit of the ratio of $N$ and $T$ exists and is bounded between 0 and $\infty$ ($0<\lim(N/T)<\infty$). Thereupon, they introduce a bias corrected LSDV (BCLSDV) estimator, which produces consistent parameter estimates, provided that the model is stable, i.e., $\tau+\rho+\eta<1$. When $N/T\rightarrow\infty$, Lee and Yu (2010) find their LSDV and BCLSDV estimators to be T consistent, as is usual. However, when $T/N^{1/3}\rightarrow\infty$ in addition to $N/T\rightarrow\infty$, Yu et al. (2008, theorem 5) demonstrate that their BCLSDV estimator is still an improvement upon the T consistency of the LSDV estimator. This is the case when $N=T^p$ with $1<p<3$. Since $N$ and $T$ in the empirical analysis below are equal to 62 and 30, respectively, we apply Lee and Yu’s BCLSDV estimator in this paper. If the condition $\tau+\rho+\eta<1$ is not satisfied, the BCLSDV needs further adjustment (see section 2.2).

Equation (1) may be rewritten as

$$\Delta Y_t = \rho W \Delta Y_{t_1} + (\tau - 1) Y_{t-1} + (\rho + \eta) W Y_{t-1} + X_t \beta + \mu + \alpha_t + \nu_t + \epsilon_t,$$

where the left-hand side variable is the change of the financial liberalization index, i.e., financial reform. Equation (2) differs from Abiad and Mody’s (2005) model in three respects. First, financial reform of a particular country in a particular year is

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4 The wording “Least Squares Dummy Variables Estimator” suggests that the endogeneity of $W_{Y_{t}}$ is not accounted for, but the opposite is true.

5 For this purpose we used a Matlab routine that has kindly been made available by Jihai Yu.
explained not only by the initial level of the financial liberalization index in that country itself \(Y_{t-1}\), but also by the initial levels of the financial liberalization index in other countries \(WY_{t-1}\). For this model to converge, Elhorst (2001) has found that the sum of the coefficients of the initial level variables in (2) should be smaller than 0, which they will if the model is stable; if \(\tau+\rho+\eta<1\), we also have \((\tau-1)+(\rho+\eta)=\tau+\rho+\eta-1<0\). The coefficient of the initial level in the country itself, \(\tau\), is smaller than zero, indicating that an increase of the financial liberalization index becomes less likely if it reaches the upper limit of the scale. The sign of the coefficient of the initial levels in other countries, \(\rho+\eta\), is unknown but will probably be positive, indicating that an increase of the financial liberalization index becomes more likely if other countries have fewer financial restrictions. The total effect, however, is negative.

Second, financial reform in one country is explained by financial reform in other countries, depending on the specification of the spatial weights matrix \(W\Delta Y_t\). This term represents the regional diffusion effect and allows for different specifications of the spatial weights matrix \(W\). Abiad and Mody (2005) did not consider a spatial weights matrix, but assumed that each country follows the regional leader. Effectively, this corresponds to a spatial weights matrix with just one non-zero element in each row of \(W\) over a particular period of time (see section 4 for more details).

Finally, we control for time-period fixed effects. This precludes that a non-zero and possibly significant coefficient estimate of the regional diffusion effect simply stems from ignoring time effects that are common to all countries.

2.2 Non-stability

The estimation of a dynamic spatial panel data model gets more complicated when the condition \(\tau+\rho+\eta<1\) is not satisfied, i.e., if the model is unstable. To get rid of possible unstable components in \(Y_t\), Lee and Yu (2010) propose to transform the model in spatial first-differences by taking every variable in equation (1) in deviation of its spatially lagged value. Mathematically, this is equivalent with
multiplying equation (1) by the matrix \((I-W)\), where \(I\) denotes the identity matrix of order \(N\),

\[
(I-W)Y_t = \tau(I-W)Y_{t-1} + \rho W(I-W)Y_t + \eta W(I-W)Y_{t-1} + (I-W)X_t \beta + (I-W)\mu + (I-W)e_t,
\]

and where we made use of the property \((I-W)W = W(I-W)\). The resulting equation has some important properties, which require further explanation. First, since \(\alpha_{i0}(I-W)_{iN} = 0\), all time-period fixed effects are eliminated from the model. Note that these fixed effects do remain effective since the estimation of a model formulated in levels produces the same parameter estimates as the estimation of that model reformulated in first-differences without time fixed effects. The reason to renounce the first approach is that we also want to remove the inconsistency caused by the possibly unstable character of the financial liberalization index. Second, just as first differencing in time would reduce the number of observations available for estimation, so does first differencing in space; the former by one for every country and the latter by one for every time period. In contrast to Elhorst et al. (2007), however, who take spatial first differences with respect to one particular reference unit, it is not immediately clear which observation gets lost when the model is transformed by \((I-W)\). Third, since we assumed that \(e_t\) has zero mean and variance \(\sigma^2\), the variance of \((I-W)e_t\) is \(\sigma^2\Sigma\), where \(\Sigma = (I-W)(I-W)'\). This raises the question how to estimate the parameters of equation (3).

Since the eigenvalues of the matrix \((I-W)\) are equal to \(1-\omega_i\) \((i=1,\ldots,N)\), where each \(\omega_i\) denotes a particular eigenvalue of the spatial weights matrix \(W\), and the largest eigenvalue of \(W\) is one \((\omega_{\text{max}}=1)\), provided that \(W\) is row-normalized, at least one eigenvalue of the matrix \((I-W)\) will be zero. This implies that the determinant of \((I-W)\) equals zero and thus that this matrix does not have full rank. If \((I-W)\) does have rank \(N-1\) instead of \(N\), so does \(\Sigma\).\(^6\) Let \(\Lambda_{N-1}\) denote the \((N-1)\times(N-1)\) diagonal matrix of nonzero eigenvalues of \(\Sigma\) and \(F_{N,N-1}\) the

\(^6\) If \(W\) has more than just one eigenvalue that is equal to 1, the number \(N-1\) must be adjusted accordingly.
corresponding orthonormal N×(N-1) matrix of eigenvectors. Then we can transform equation (3) by the matrix $P = \Lambda_{N-1}^{-1/2}F_{N,N-1}'$, to get

$$P(I-W)Y_t = \tau P(I-W)Y_{t-1} + \rho PW(I-W)Y_t + \eta PW(I-W)Y_{t-1} + \Lambda = \eta P(I-W)Y_t + \mu P(I-W)\epsilon_t.$$  \tag{4}

This transformation has three effects. First, since $P$ is a (N-1)×N matrix, the transformation $Y_t^* = P(I-W)Y_t$ reduces the length of $Y_t$ to N-1. The same applies to the length of the transformed matrices or vectors $X_t^*$, $\mu^*$ and $\epsilon_t^*$. This transformation therefore reflects the "observation that gets lost". We use quotation marks here since the information that really gets lost is the perfect linear combination among the observations that occurred due to the multiplication of equation (1) by the matrix (I-W). Second, since the transformation $P$ reverses the transformation (I-W) (except for the observation that gets lost), we have $E(\epsilon_t^*\epsilon_t^*) = \sigma^2 I_{N-1}$. Third, since $W^* = PW(I-W) = \Lambda_{N-1}^{-1/2}F_{N,N-1}'WF_{N,N-1}\Lambda_{N-1}^{1/2}$ (see Lee and Yu, 2010), equation (4) can be rewritten as

$$Y_t^* = \tau Y_{t-1}^* + \rho W^*Y_t + \eta W^*Y_{t-1} + X_t^*\beta + \mu^* + \epsilon_t^*,$$  \tag{5}

whose parameters can be consistently estimated by the same BCLSDV estimator that is used to estimate equation (1). Lee and Yu (2010) show that the transformed model will be stable of $\tau + \omega_{\text{max}-1}(\rho + \eta) < 1$, where $\omega_{\text{max}-1}$ denotes the second largest eigenvalue of the spatial weights matrix $W$. Importantly, the latter restriction is less restrictive than the original restriction $\tau + \rho + \eta < 1$.

In sum, there are three differences between equations (5) and (1). First, the number of observations in each time period is N-1 instead of N. Second, time-period fixed effects are wiped out, although their effectiveness has not. Third, the matrix $W$ is replaced by $W^*$. It is important to note that the row elements of $W^*$, in contrast to those in $W$, do not necessarily sum up to one. Nevertheless, the N-1 eigenvalues of $W^*$ are identical to those of $W$ that remain after the unit eigenvalue of $W$ is excluded.
2.3 Direct, indirect and spatial spillover effects

Many empirical studies use point estimates of one or more spatial regression models to test the hypothesis as to whether or not spatial spillover effects exist. However, LeSage and Pace (2009, p. 74) have recently pointed out that this may lead to erroneous conclusions, and that a partial derivative interpretation of the impact from changes to the variables of different model specifications represents a more valid basis for testing this hypothesis. They also prove this using a static spatial econometric model (ibid, pp. 34-40). Below we derive the marginal effects of the explanatory variables on the change of the financial liberalization index using (2).

By rewriting this equation as

$$\Delta Y_t = (I - \rho W)^{-1}[(\tau - I + (\rho + \eta)W)Y_{t-1} + (I - \rho W)^{-1}X_t \beta + (I - \rho W)^{-1}(\mu + \alpha_{it} + \nu + \epsilon_t)],$$

the matrix of partial derivatives of $\Delta Y_t$ with respect to the $k^{th}$ explanatory variable of $X_t$ in country 1 up to country $N$ (say $x_{ik}$ for $i=1,\ldots,N$, respectively), both at a particular point in time $t$, is

$$\begin{bmatrix}
\frac{\partial \Delta Y_t}{\partial x_{1k}} & \frac{\partial \Delta Y_t}{\partial x_{Nk}} \\
\frac{\partial \Delta Y_t}{\partial x_{1k}} & \frac{\partial \Delta Y_t}{\partial x_{Nk}} \\
\end{bmatrix}_{t} = (I - \rho W)^{-1} \beta_k.$$

(7)

These partial derivatives have the following properties. First, if a particular explanatory variable in a particular country changes, not only the liberalization index in that country will change but also the liberalization index in other countries. The first is called a direct effect and the second an indirect effect. Note that every diagonal element of the matrix on the right-hand side of (7) represents a direct effect, and that every non-diagonal element represents an indirect effect. Consequently, indirect effects do not occur if $\rho = 0$. Second, both direct and indirect effects are different for different units in the sample. Direct effects are different
because the diagonal elements of the so-called spatial multiplier matrix \((I - \rho W)^{-1}\)
are different for different countries, provided that \(\rho \neq 0\). Similarly, indirect effects
are different because the non-diagonal elements of that matrix are different.

Since the direct and the indirect effects are different for different countries
in the sample, the presentation of both effects is difficult. LeSage and Pace (2009)
therefore propose to report one direct effect measured by the average of the
diagonal elements, and one indirect effect measured by the average of either the
row sums or the column sums of the non-diagonal elements of that matrix.\(^7\) The
total effect is the sum of the direct and indirect effects.

Similarly, we can calculate the effects estimates of convergence. Using (6),
we have

\[
\frac{\partial \Delta Y_t}{\partial \Delta Y_{t-1}} = (I - \rho W)^{-1}[(\tau - 1)I + (\rho + \eta)W].
\] (8)

The average diagonal element of this matrix measures the strength of the
convergence effect of the country itself, and the average row sum of the non-diagonal elements the convergence effect of other countries.

A specific situation occurs if \(\tau + \rho + \eta = 1\). Lee and Yu (2010) label this situation
as spatial cointegration, after conventional cointegration in the time series
literature. The cointegration matrix is \((I-W)\) and the cointegration rank is the
number of eigenvalues of \(W\) that are smaller than 1, which is \(N-1\) (but note
footnote 5). If \(\tau + \rho + \eta = 1\), we have

\[
\frac{\partial \Delta Y_t}{\partial \Delta Y_{t-1}} = (\tau - 1)(I - \rho W)^{-1}(I - W).
\] (8')

If \(W\) is row-normalized, the total effect of this partial derivative equals zero by
construction. In other words, the extent of financial liberalization in different
countries will not converge if this variable turns out to be spatially cointegrated.

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\(^7\) Since the numerical magnitudes of these two calculations of the indirect effect are the same, it does not matter which one is used (LeSage and Pace, 2009, pp. 33-42).
3. Data

The financial liberalization index is taken from Abiad et al. (2008). This index consists of seven components, namely:

1. Credit controls and excessively high reserve requirements
2. Interest rate controls
3. Entry barriers
4. State ownership in the banking sector
5. Capital account restrictions
6. Prudential regulations and supervision of the banking sector
7. Securities market policy.

Each component has been assigned a score between 0 and 3, with 0 representing the highest degree of repression and 3 representing full liberalization. These scores are then summed to obtain a value between 0 and 21 as a measure of overall financial liberalization, and finally normalized to obtain values between 0 and 100. Abiad et al. (2008) acknowledge that the dimension referring to prudential regulation and supervision of banks is different from the other dimensions of financial liberalization. A higher score in this case means better (or more) regulation. In section 4 we therefore also estimate the model using an indicator of financial reform excluding this component.

The database covers 62 countries over the period 1976-2005. Table 1 provides summary statistics for the financial liberalization index and its separate components.

<<< Table 1 here >>

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8 The dataset is an updated version of the data used by Abiad and Mody (2005). There are alternative measures of financial liberalization available, like those of Williamson and Mahar (1998), Kaminsky and Schmukler (2008), Bandiera et al. (2000), and Laeven and Claessens (2003). However, the coverage of these alternatives is more restrictive and therefore we consider the dataset of Abiad et al. (2008) as the most comprehensive source of information on financial liberalization at this moment.

9 Data before 1976 is incomplete, while data for 1975 is used to cover the (spatially lagged) dependent variable lagged in time.
Our basic set of control variables is roughly similar to that of Abiad and Mody (2005).\textsuperscript{10} We include various financial crisis dummies: BANK, CUR and DEBT are dummy variables that are set to 1 if the country was hit by a banking crisis, currency crisis or debt crisis, respectively, within two years preceding the observation of the liberalization index, and 0 otherwise. These data are taken from Laeven and Valencia (2008). RECESSION and HINFL are dummy variables that are 1 if the annual real GDP growth rate was negative or if the annual inflation rate exceeded 50\%, respectively, and 0 otherwise. The first variable is based on data taken from the Groningen Growth and Development Centre (GGDC), and data for the second variable comes from the International Financial Statistics (IFS) of the IMF. FIRSTYEAR is a dummy variable that is 1 if a newly elected government was in the first year of its term, and 0 otherwise. IMF is a dummy variable that is 1 if the country was under an IMF program, and 0 otherwise (data taken from Sturm et al., 2005). RIGHT or LEFT are dummy variables that are 1 if the main government party was left- or right-wing, according to the Database of Political Institutions (DPI) of the World Bank.

We also examine the impact of two potential determinants of financial reform that have not been considered in Abiad and Mody (2005). According to Huang (2009), democracy is negatively related to financial reform. This finding is remarkable in view of the positive relationship that has been found between democracy and general economic reform (Dethier et al., 1999; De Haan and Sturm, 2003; Pitlik and Wirth, 2003; Lundström, 2005; Pitlik, 2008). Therefore, we consider it useful to examine the nature of the relationship between democracy and financial reform, if any. For this purpose, we use DEMO, the level of democracy as measured by the Polity 2 variable from the Polity IV project. Its value lies between -10 and 10, where a higher value indicates a higher level of democracy. Another political factor that might matter is government fractionalization. Krueger (1993) argues that “strong” governments will react quickly to crises, while “fractionalized” governments will delay stabilizations. The more fractionalized a

\textsuperscript{10} Except for USINT, the U.S. Treasury Bill rate taken from IFS of the IMF, because this variable would be perfectly collinear with the time period fixed effects.
government is, the harder it will be to reach consensus on possible controversial reform measures. Government fractionalization (GOVFRAC) is defined as the probability that two deputies picked at random from among the parties in government will be of different parties. Table 2 provides summary statistics of all the variables.

<< Table 2 here >>

4. Results
As previous studies (Abiad and Mody, 2005 and Huang, 2009) focus on the distance of individual countries to regional leaders, we start by using a spatial weights matrix W that is based on regional leaders. In principle, this spatial weights matrix might change over time when another country becomes the leader of a particular region. In practice, however, this hardly occurs. In those cases, we chose the regional leader to be the country that had the highest financial liberalization index for the longest part of the observation period.11 Just as Abiad and Mody (2005), we also assume that regional leaders do not follow other countries, except if the other country is also a regional leader located in the same region. This occurred in the regions of Africa and the OECD, where there appeared to be two regional leaders for the longest part of the observation period.

It should be stressed that the form of the spatial weights matrix so obtained is rather unusual compared to spatial weights matrices that will be considered later in this paper. First, since most countries are no regional leader, only a limited number of columns contain non-zero elements. Second, since a regional leader does not interact with other countries, unless there happen to be two leaders in the same region, the rows of regional leading countries mostly contain zero elements. Finally, most eigenvalues of W are zero. Actually, whereas the largest eigenvalue is one (which is usual in a row-normalized matrix), the second largest eigenvalue

11 Although the estimators can be modified for a spatial weights matrix that does change over time, their asymptotic properties are unknown, since Lee and Yu (2010) assume that W is constant over time.
happens to be 0. Nevertheless, this spatial weights matrix satisfies all regularity conditions that have been spelled out in Lee and Yu (2010) and related work. Table 3 reports the estimation results based on this spatial weights matrix.

<< Table 3 here >>

To test for the existence of time-period fixed effects, column (1) shows the results of the BCLSDV estimator applied to model (1) without time-period fixed effects and column (2) shows the results with these effects. The results of the corresponding F-test (9.148 with 29 degrees of freedom in the numerator and 1755 in the denominator, p < 0.01) indicate that time-period fixed effects should be included.

To find out whether the model including time-period fixed effects is stable, we calculated $\tau + \rho + \eta$ and carried out a Wald-test to investigate the (null) hypothesis $\tau + \rho + \eta = 1$. The results of this test show that the liberalization index is explosive, which implies that the model should be reformulated in spatial first-differences according to model (3) and be re-estimated according to its transformation in (5) to obtain consistent parameter estimates. The results so obtained are reported in column (3) of Table 3. The transformed model is stable, because the condition that $\tau + \omega_{\text{max}} (\rho + \eta) < 1$ is satisfied. We therefore use the parameter estimates of this model to compute direct and indirect effects of the different explanatory variables, i.e., the average of the diagonal elements and of the row/column sums of the matrix that is obtained when using (7). These direct and indirect effects, as well as the total effects, are reported in columns (4), (5) and (6) of Table 3, respectively.

The coefficients of the dependent variables lagged in time $FL_{t-1}$ and in space $WFL_t$ are both positive and significant in the spatial first-differenced model [column (3)], while the coefficient of the dependent variable lagged in both space and time $WY_{t-1}$ is negative and insignificant. The coefficient of $WFL_t$, which may be
interpreted as the impact of the regional diffusion effect, amounts to 0.126 and therefore is smaller than that found by Abiad and Mody (2005).\footnote{As we pointed before, ignoring time-period fixed effects, when present, may lead to an upward bias up to 0.45 (see Lee and Yu, 2010). The difference between Abiad and Mody’s coefficient estimate of 0.44 to 0.55 and our estimate of 0.126 falls within this range. In his replication study, Huang (2005) also adds time-period fixed effects. In addition, he employs Pesaran’s (2006) common correlated effect pooled (CCEP) approach to allow for the possibility of error dependence across countries. Remarkably, he finds that the coefficient estimate of the diffusion effect falls to a negative and significant number of -0.14. This would mean that countries turn away rather than follow the regional leader, a conclusion that must be rejected in view of our estimation results.}

The coefficients of the initial levels of the liberalization index also represent a significant convergence effect. The direct effect based on the coefficient matrix: 
\[(I-\rho W)^{-1}[(\tau-1)I+(\rho+\eta)W^*]\]
of the variable $Y_{t1}$ is $-1.040$ (t-value $-20.45$), while its indirect effect is $0.936$ (19.78); the higher the indices of financial liberalization in the country itself, or the lower the indices of financial liberalization in other countries, the lower the extent of financial reform, and vice versa. The total effect is small, but still negative and significant ($-0.104$, t-value $-4.86$).

The coefficient estimates of the independent variables show that banking and debt crises significantly set back financial liberalization. Conversely, periods of high inflation and the imposition of an IMF program lead to financial reform. Left- and right-wing governments also appear to reform their financial system significantly more than centrist governments. Recessions and newly elected governments have a positive and a negative effect, respectively, but the coefficients of both variables are insignificant. Also the potential determinants of financial reform that have been put forward by Huang (2009) and Krueger (1993), respectively, namely the level of democracy ($DEMO$) and government fractionalization ($GOVFRAC$), are insignificant.

The direct effects in the spatial first-differenced model [column (4)] are different from the estimates of the response parameters [column (3)]. This is caused by the feedback effects that arise as a result of impacts passing through other countries and back to the country itself. These feedback effects, however, turn out to be very small. For example, since the direct effect and the coefficient estimate of banking crises ($BANK$) are $-0.898$ and $-0.917$, respectively, its feedback effect is
only -0.019. By contrast, the indirect effect of a change in one of the explanatory variables in a particular country on other countries appears to be approximately 13% of the direct effect within that country. Furthermore, based on the t-statistics calculated from a set of 1,000 simulated parameter values13, one of these indirect effects appears to be significantly different from zero at a significance level of 5% (inflation) and another three at 10% (debt crises, IMF program and left-wing government). In other words, if one of these explanatory variables in a particular country changes, not only the liberalization index of that country itself but also that of other countries will change. The proportion of the change in other countries and the change in the country itself is approximately 1 to 7.7.

The results that have been presented and discussed so far strongly depend on the regional leader matrix. An interesting question is whether the performance of the model will improve and the conclusions will change when other spatial weights matrices are used. Table 4 reports the performance of the spatial first-differenced model based on three alternative specifications of W. The first is a bilateral trade flow matrix (imports plus exports) based on data extracted from the World Trade Organization (WTO) for the year 1998 (row 2 of Table 4). The second is an inverse distance matrix based on the physical distance between the capitals of every pair of countries according to the distance dataset of the CEPII (row 3 of Table 4). The third is an inverse distance matrix with a cut-off point after which the interaction between two countries is assumed to be zero. Since the value of this cut-off point is an empirical question, we consider a range of different values (rows 4-8 of Table 4).14 All matrices have been row-normalized, so that the entries of each row add up to 1. The results obtained for the regional leader matrix are also reported (row 1 of Table 4).

13 In order to draw inferences regarding the statistical significance of these effects, we used the variation of 1,000 simulated parameters combinations drawn from the multivariate normal distribution implied by the maximum likelihood estimates (see LeSage and Pace, 2009 for mathematical details).


15 For values greater than or equal to 3000, every country in the sample appeared to have at least one neighbour. To avoid "islands" for values that are smaller than 3000, we also assumed that every country at least interacts with its nearest neighbour.
To determine the spatial weights matrix that best describes the data, one may compare log-likelihood function values and rely on the model that exhibits the highest value. Since the number of eigenvalues equal to one turned out to be different for different Ws (see Table 4), and the estimation of the dynamic spatial panel data model reformulated in spatial first-differences is based on different numbers of observations as a result (see the explanation in footnote 5), we consider log-likelihood function values relative to the number of observations. Another approach is to compare Bayesian posterior model probabilities and to rely on the model with the highest probability (see LeSage and Pace, 2009, pp. 167-168). Finally, one may select the model that exhibits the lowest parameter estimate of the residual variance ($\sigma^2$).

The results reported in Table 4 show that the regional leader matrix describes the interaction effects of financial reform among countries rather poor. It has the lowest log likelihood function value, the lowest Bayesian posterior model probability and the highest parameter estimate of the residual variance. A much better description is obtained when adopting the inverse distance matrix with a cut-off point at 3000 km, which has the highest log likelihood function value, the highest Bayesian posterior model probability and the lowest parameter estimate of the residual variance. Table 5 reports the estimation results when employing this matrix.\(^{17}\)

---

\(^{16}\) To be able to calculate the BCLSDV estimator, one needs the LSDV estimator (see section 2.1). LeSage (1999) furnishes Matlab routines to determine the cross-sectional version of this estimator using maximum likelihood (ML) and the Bayesian Markov Chain Monte Carlo (MCMC) approach (www.spatial-econometrics.com), while Elhorst (2003, 2010) extends the ML routine for spatial panels (www.regroningen.nl/elhorst). Unfortunately, the Bayesian MCMC routine for spatial panels, required to be able to compute Bayesian posterior model probabilities, does not exist yet. As an alternative, one may replace all cross-sectional arguments of this routine by their spatial panel counterparts, e.g., a block-diagonal NT×NT matrix, diag(W,…,W) as argument for W. Although less efficient from a computational viewpoint, this works well if N×T is not too large.

\(^{17}\) Estimation results for the dynamic spatial panel data model formulated in levels (with or without time dummies) are left aside here, because this model again turned out to be non-stable.
When comparing the estimation results in Table 5 with those in Table 3, we see noteworthy changes. First, whereas $\tau + \rho + \eta$ appeared to be significantly greater than one when using the regional leader matrix, $\tau + \rho + \eta$ is smaller than one when using the inverse distance matrix. In addition, the hypothesis $\tau + \rho + \eta = 1$ and thus that the financial liberalization process is spatially cointegrated cannot be rejected. For this reason, the estimation results when imposing this restriction are reported too. As pointed out, the side effect of this so-called spatially cointegrated model (see section 2.2) is that the convergence effect of the initial levels of the liberalization index is zero. Second, whereas the interaction effect of financial reform among countries according to the regional leader matrix amounts to 0.126, it rises to 0.191 according to the inverse distance matrix. Related to that, whereas a country's financial reform in the first case is influenced by the leading country in the region only and feedback effects are not accounted for, it is influenced back and forth by all countries within a radius of 3000 km in the second case. Third, according to the regional leader matrix, debt crises set back financial liberalization but not according to the inverse distance matrix. In the latter case, both the coefficient estimate and the direct effect of this variable are no longer significant. Also the results for currency crises change. Whereas the willingness to liberalize is not significantly affected by a currency crisis according to the regional leader matrix, the effect of currency crises becomes significant according to the inverse distance matrix (significance level of 10% with respect to the coefficient estimate and 5% with respect to the direct effect). Finally, whereas the indirect effects as a percentage of the direct effects appeared to be approximately 13% when using the regional leader matrix, this ratio increases to approximately 29% when using the inverse distance matrix. Both spatial weights matrices have in common that the indirect effects of inflation, participation in an IMF program and left-wing governments are significant (at 5% or 10%). The difference is that the inverse distance matrix also points to a significant indirect effect of banking crises rather than that of debt crises.
Finally, we investigated whether the results are robust to changing the financial liberalization index by excluding the component prudential regulations and supervision of the banking sector (see section 3). Although the coefficient estimates and the t-values change slightly, our main conclusions remain the same (results available on request).

5. Conclusions

Abiad and Mody’s (2008) financial liberalization index turns out to be unstable in a dynamic spatial panel data model. As shown by Lee and Yu (2010), under such circumstances the empirical model should be reformulated in spatial first-differences and be re-estimated to obtain consistent parameter estimates. By rewriting the dynamic spatial panel data model, we show that the change in the liberalization index of a particular country, i.e., financial reform, depends on the change in the liberalization index of other countries, on the initial levels of the liberalization index in the country itself and that of other countries and on a set of exogenous explanatory variables. Using this approach, we find that financial reform in one country affects financial reform in other countries. Our results also point out that the regional leader matrix used in Abiad and Mody (2005) and Huang (2009) must be rejected in favour of an inverse distance matrix with a cut-off point at 3000 km, provided that time period fixed effects are controlled for. The degree of interaction among countries amounts to 0.191, which is substantially smaller than that in Abiad and Mody (2005).

Explanatory variables that appear to have a significant effect on financial reform in the country itself are banking and currency crises, periods of high inflation, the imposition of an IMF program, and the dummy for left-wing governments. Most of these variables also have a significant effect on neighbouring countries, although the significance level tends to be lower (10% rather than 5%). The magnitude of spillover effects compared to the effect on countries themselves is approximately 1 to 3.4. The level of democracy and government fractionalization put forward by Huang (2009) and Krueger (1993), respectively, as potential determinant of financial reform turn out to be insignificant.
References


Yu, Jihai, Jong, Robert de and Lung-fei Lee (2008), ‘Quasi-maximum likelihood estimators for spatial dynamic panel data with fixed effects when both n and T are large’, *Journal of Econometrics* 146: 118-134.
Table 1. Financial liberalization: Summary statistics of the index and its components

<table>
<thead>
<tr>
<th>Variable</th>
<th>Mean</th>
<th>Standard deviation</th>
</tr>
</thead>
<tbody>
<tr>
<td>Credit Controls</td>
<td>1.699</td>
<td>1.150</td>
</tr>
<tr>
<td>Interest Rate Controls</td>
<td>1.972</td>
<td>1.262</td>
</tr>
<tr>
<td>Entry Barriers</td>
<td>1.800</td>
<td>1.166</td>
</tr>
<tr>
<td>Bank Regulation and Supervision</td>
<td>0.837</td>
<td>0.981</td>
</tr>
<tr>
<td>Privatization</td>
<td>1.364</td>
<td>1.165</td>
</tr>
<tr>
<td>Capital Account</td>
<td>1.795</td>
<td>1.085</td>
</tr>
<tr>
<td>Securities Market</td>
<td>1.637</td>
<td>1.115</td>
</tr>
<tr>
<td>Financial Reform Index</td>
<td>11.104</td>
<td>6.262</td>
</tr>
<tr>
<td>Index normalized</td>
<td>52.9</td>
<td>29.8</td>
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</table>

Table 2. Summary statistics of control variables

<table>
<thead>
<tr>
<th>Variable</th>
<th>Mean</th>
<th>Standard Deviation</th>
<th>Minimum</th>
<th>Maximum</th>
</tr>
</thead>
<tbody>
<tr>
<td>BANK</td>
<td>0.093</td>
<td>0.291</td>
<td>0</td>
<td>1</td>
</tr>
<tr>
<td>CUR</td>
<td>0.139</td>
<td>0.346</td>
<td>0</td>
<td>1</td>
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<tr>
<td>DEBT</td>
<td>0.048</td>
<td>0.214</td>
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<td>1</td>
</tr>
<tr>
<td>RECESSION</td>
<td>0.147</td>
<td>0.354</td>
<td>0</td>
<td>1</td>
</tr>
<tr>
<td>HINFL</td>
<td>0.075</td>
<td>0.263</td>
<td>0</td>
<td>1</td>
</tr>
<tr>
<td>FIRSTYEAR</td>
<td>0.199</td>
<td>0.400</td>
<td>0</td>
<td>1</td>
</tr>
<tr>
<td>IMF</td>
<td>0.335</td>
<td>0.472</td>
<td>0</td>
<td>1</td>
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<td>0.318</td>
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</tr>
<tr>
<td>RIGHT</td>
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<td>0.465</td>
<td>0</td>
<td>1</td>
</tr>
<tr>
<td>DEMO</td>
<td>4.209</td>
<td>6.812</td>
<td>-10</td>
<td>10</td>
</tr>
<tr>
<td>GOVFRAC</td>
<td>0.208</td>
<td>0.274</td>
<td>0</td>
<td>1</td>
</tr>
</tbody>
</table>
Table 3. Explaining the financial liberalization index (FL) using different model specifications; spatial weights matrix based on regional leaders

<table>
<thead>
<tr>
<th>Determinants</th>
<th>No time dummies</th>
<th>Time dummies</th>
<th>Spatial first-differences and effects estimates</th>
<th>Coeff. (1)</th>
<th>Direct (3)</th>
<th>Indirect (5)</th>
<th>Total (6)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)(^a)</td>
<td>(2)(^)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>FL(_{t+1})</td>
<td>0.979</td>
<td>0.940</td>
<td>0.959</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(106.62)</td>
<td>(76.66)</td>
<td>(102.63)</td>
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<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>W FL(_t)</td>
<td>-0.238</td>
<td>0.075</td>
<td>0.126</td>
<td>-1.040(^d)</td>
<td>0.936(^d)</td>
<td>-0.104</td>
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<tr>
<td></td>
<td>(-8.17)</td>
<td>(2.48)</td>
<td>(3.24)</td>
<td>(-20.45)</td>
<td>(19.78)</td>
<td>(-4.86)</td>
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<tr>
<td>W FL(_{t+1})</td>
<td>0.253</td>
<td>-0.147(^e)</td>
<td>-0.052</td>
<td></td>
<td></td>
<td></td>
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<tr>
<td></td>
<td>(10.31)</td>
<td>(-0.29(^e))</td>
<td>(-0.07)</td>
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<td>-0.972</td>
<td>0.456</td>
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<td>-1.013</td>
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<tr>
<td></td>
<td>(-1.17)</td>
<td>(-2.52)</td>
<td>(-1.99)</td>
<td>(-2.13)</td>
<td>(-1.55)</td>
<td>(-2.12)</td>
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<tr>
<td>CUR</td>
<td>0.912</td>
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<td>0.085</td>
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<tr>
<td></td>
<td>(1.54)</td>
<td>(0.79)</td>
<td>(1.63)</td>
<td>(0.68)</td>
<td>(0.76)</td>
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<tr>
<td>DEBT</td>
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<td>(-1.65)</td>
<td>(-1.65)</td>
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<tr>
<td>RECESSION</td>
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<td>0.236</td>
<td>0.287</td>
<td>0.037</td>
<td>0.070</td>
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<tr>
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<td>(0.61)</td>
<td>(0.48)</td>
<td>(0.81)</td>
<td>(0.68)</td>
<td>(0.76)</td>
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<tr>
<td>HINFL</td>
<td>2.206</td>
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<td>0.272</td>
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<td>(2.92)</td>
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<td>(-2.07)</td>
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<td></td>
<td>(3.49)</td>
<td>(2.22)</td>
<td>(2.94)</td>
<td>(1.83)</td>
<td>(3.04)</td>
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<tr>
<td>LEFT</td>
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<td>1.111</td>
<td>1.324</td>
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<td>0.171</td>
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<td>(3.64)</td>
<td>(2.89)</td>
<td>(3.33)</td>
<td>(1.87)</td>
<td>(3.16)</td>
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<td>RIGHT</td>
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<td>0.080</td>
<td>0.693</td>
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<td></td>
<td>(2.26)</td>
<td>(1.82)</td>
<td>(1.98)</td>
<td>(1.21)</td>
<td>(1.50)</td>
<td>(1.50)</td>
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<tr>
<td>DEMO</td>
<td>0.029</td>
<td>-0.015</td>
<td>0.025</td>
<td>0.004</td>
<td>0.004</td>
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<tr>
<td></td>
<td>(1.18)</td>
<td>(-0.61)</td>
<td>(0.53)</td>
<td>(0.56)</td>
<td>(0.56)</td>
<td>(0.56)</td>
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</tr>
<tr>
<td>GOVFRAC</td>
<td>0.224</td>
<td>0.223</td>
<td>0.118</td>
<td>0.017</td>
<td>0.017</td>
<td>0.132</td>
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<tr>
<td></td>
<td>(0.37)</td>
<td>(0.05)</td>
<td>(0.28)</td>
<td>(0.19)</td>
<td>(0.19)</td>
<td>(0.19)</td>
<td></td>
</tr>
</tbody>
</table>

# observations 1860 1860 1830
\(\sigma^2\) 25.286 21.965 23.106
\(\tau+\rho+\eta\) 0.994 1.015 Not appl.
Wald-test \(\tau+\rho+\eta=1\) 0.452 0.489 7.641
p-value Wald-test 0.501 0.484 0.006
\(\tau+\theta_{max-1}(\rho+\eta)\) Not appl. Not appl. 0.959

* BCLSDV estimator Equation (1)
* BCLSDV estimator Equation (5)
* Direct/Indirect effects using (7)
* Direct/Indirect effects initial levels financial liberalization using (8)
Table 4. Spatial weights model comparison

<table>
<thead>
<tr>
<th>W</th>
<th>Eigenvalues equal to one</th>
<th>Log likelihood function value per observation</th>
<th>Bayesian posterior model probability</th>
<th>$\hat{\sigma}^2$</th>
</tr>
</thead>
<tbody>
<tr>
<td>(1) Regional leader</td>
<td>1</td>
<td>-2.89</td>
<td>1e-25</td>
<td>23.11</td>
</tr>
<tr>
<td>(2) Trade flows</td>
<td>1</td>
<td>-2.87</td>
<td>0.001</td>
<td>22.15</td>
</tr>
<tr>
<td>(3) Inverse distance</td>
<td>1</td>
<td>-2.87</td>
<td>0.019</td>
<td>21.96</td>
</tr>
<tr>
<td>(4) cut-off 2000 km</td>
<td>9</td>
<td>-2.85</td>
<td>0.263</td>
<td>22.05</td>
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<tr>
<td>(5) cut-off 3000 km</td>
<td>4</td>
<td>-2.84</td>
<td>0.314</td>
<td>21.32</td>
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<td>(6) cut-off 4000 km</td>
<td>3</td>
<td>-2.85</td>
<td>0.181</td>
<td>21.53</td>
</tr>
<tr>
<td>(7) cut-off 5000 km</td>
<td>2</td>
<td>-2.86</td>
<td>0.180</td>
<td>21.73</td>
</tr>
<tr>
<td>(8) cut-off 6000 km</td>
<td>1</td>
<td>-2.87</td>
<td>0.042</td>
<td>21.97</td>
</tr>
</tbody>
</table>
Table 5. Explaining the financial liberalization index (FL) using different model specifications; spatial weights matrix based on inverse distances with cut-off point at 3,000 km

<table>
<thead>
<tr>
<th>Determinants</th>
<th>Spatial first-differences</th>
<th>Spatially cointegrated model $\tau+\rho+\eta=1$ and effects estimates</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)$^a$</td>
<td>Coeff. (2)$^b$</td>
</tr>
<tr>
<td>FL$_t$1</td>
<td>0.938 (70.75)</td>
<td>0.942 (78.86)</td>
</tr>
<tr>
<td>W FL$_t$</td>
<td>0.185 (3.05)</td>
<td>0.191 (3.32)</td>
</tr>
<tr>
<td>W FL$_t$1</td>
<td>-0.150 (-2.11)</td>
<td>-0.133 (-2.29)</td>
</tr>
<tr>
<td>BANK</td>
<td>-1.107 (-2.71)</td>
<td>-1.104 (-2.57)</td>
</tr>
<tr>
<td>CUR</td>
<td>0.711 (1.14)</td>
<td>0.732 (1.91)</td>
</tr>
<tr>
<td>DEBT</td>
<td>-0.701 (-1.55)</td>
<td>-0.657 (-1.08)</td>
</tr>
<tr>
<td>RECESSION</td>
<td>0.321 (0.67)</td>
<td>0.327 (0.93)</td>
</tr>
<tr>
<td>HINFL</td>
<td>1.439 (1.95)</td>
<td>1.416 (2.56)</td>
</tr>
<tr>
<td>FIRSTYEAR</td>
<td>-0.580 (-1.94)</td>
<td>-0.578 (-1.94)</td>
</tr>
<tr>
<td>IMF</td>
<td>0.777 (2.60)</td>
<td>0.789 (2.48)</td>
</tr>
<tr>
<td>LEFT</td>
<td>1.019 (2.71)</td>
<td>1.050 (2.50)</td>
</tr>
<tr>
<td>RIGHT</td>
<td>0.432 (2.03)</td>
<td>0.426 (1.02)</td>
</tr>
<tr>
<td>DEMO</td>
<td>0.004 (0.11)</td>
<td>0.001 (0.03)</td>
</tr>
<tr>
<td>GOVFRAC</td>
<td>0.011 (0.40)</td>
<td>-0.02 (-0.03)</td>
</tr>
</tbody>
</table>

# observations 1740 1740
$\sigma^2$ 21.320 21.542
$\tau+\rho+\eta$ 0.973 1
Wald-test $\tau+\rho+\eta=1$ 0.927 Not appl.
p-value Wald-test 0.336 Not appl.
$\tau+\omega_{\text{max}}(\rho+\eta)$ 0.967 Not appl.

$^a$ BCLSDV estimator of Equation (5)
$^b$ BCLSDV estimator of Equation (5), provided that $\tau+\rho+\eta=1$.
$^c$ Direct/Indirect effects using (7)
$^d$ Direct/Indirect effects initial levels financial liberalization using (8)
Figure 1. Financial liberalization of regional leaders over time