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Measuring the World Economy

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Abstract: This paper provides an empirical assessment of whether the world economy has become smaller in terms of economic distance over the last decades. We adopt a cross-sectional spatial econometric approach, relating domestic output volatility to (distance-weighted averages of) other countries' output volatility, using a sample of 135 countries and rolling 10-year time windows over the period 1955 to 2006. Using descriptive measures, test statistics, and spatial econometric estimates, we find that cross-country interdependence was virtually insignificant in the early post-war period but has increased strongly from the mid-1960s to the mid-1980s and remained at a high level since then. Results for the most recent period suggest that common shocks to output volatility have a magnified impact and roughly quadruplicate through international spillover effects, which are transmitted through both trade and financial openness.

Keywords: financial openness, spillovers, trade, output volatility

JEL No: C21, E32, F15, F40

I. Introduction

Over the last decades there has been a sizeable increase in trade and financial openness, triggered by current and capital account liberalization as well as improvements in transport and communication technologies. As a consequence, the importance of spillover effects is likely to have increased at all levels of aggregation: among firms, industries, regions, and even countries. In other words, the relevance of geographic distance should have decreased over time, a view that has been popularized as ‘the death of distance’ by Cairncross (1997).

Yet there is little overall quantitative evidence by how much globalisation has reduced economic distance between countries. A related albeit more specific strand of the literature considers the sensitivity of trade flows to distance in gravity models. In their survey of the empirical trade literature, Leamer and Levinsohn (1995) conclude that the effect of distance on trade patterns has not diminished over time. More recently, Jacks et al. (2009), focussing on the evolution of trade costs over time, find that the role of distance has declined dramatically in the 19th century, but not over the period since 1950.

Another strand of the literature that is closely related to the subject of the present paper considers the synchronisation of business cycles across countries. While there are some well-established stylized facts such as the great moderation of US business cycle volatility since the mid-1980s, the results regarding the occurrence, magnitude and timing of moderations in output volatility vary considerably across countries (Stock and Watson, 2005).

Building on Stock and Watson (2005), Carare and Mody (2010) use factor-structural vector autoregressive models to estimate the contribution of domestic shocks, common international shocks, and international spillovers to output volatility for a sample of 22 countries over the period 1960-2007. They find that the share of volatility due to international spillovers, which

was relatively high in the late 1970s, diminished somewhat in the 1980s and increased again in the 1990s. Accounting for a share of 50-75 percent of total volatility, international spillovers formed the dominant source of instability in the post-1995 period. Carare and Mody (2010) argue that the large degree of global interconnectedness is due to the increased trade content of production as a result of vertical specialization.

Imbs (2004), using data on 25 countries over the 1980s and 1990s, confirms a strong effect of trade on business cycle synchronization, of which a sizable portion is due to intra-industry trade. Moreover, he finds that business cycles of financially integrated regions with similar economic structures show a significantly higher correlation.

Easterly and Stiglitz (2000) use a panel of (up to) 74 countries over the period 1960-1997 to explore the determinants of output volatility. They highlight the strong role of the financial sector and point to a possibly U-shaped relationship with output volatility. As the size of the financial sector grows, the stabilizing effects of consumption and production smoothing possibilities provided by a deep (international) financial system might be offset and ultimately dominated by destabilizing effects due to increases in risk and the propagation of shocks through the international financial system.

Focussing on the recent financial and economic crisis, Eickemeier et al. (2011) emphasize the role of financial shocks originating in the US. They find that size of US financial shocks varies strongly over time, with the 'global financial crisis shock' being very large by historical standards and explaining one third of the variation in GDP growth on average over all countries in 2008-2009.

The growing importance of international spillovers is underlined by Kose et al. (2003), Lumsdaine and Prasad (2003) and Artis et al. (2009), who find strong evidence for a common world component of countries' business cycles. If economic downturns or shocks to countries do not occur mainly locally any more but are becoming more and more synchronized, contagion and spillover effects will be further amplified (Danninger et al., 2009).

Most of the aforementioned studies use time series methods on high frequency data and are thus confined to comparably small sets of countries and time periods. This paper takes an alternative approach in quantitatively assessing the question whether the world economy has become smaller in terms of economic distance, leading to an increased the role and importance of international volatility spillovers over time. Instead of focussing on the transmission of single shocks (which requires high frequency data) or confining the analysis to a particular channel of interdependence (such as trade), it provides a bird-eye's perspective on the world economy, using a large cross-section of 135 countries and considering how cross-country interdependence of output volatility has evolved over the period 1955 to 2006.

From a methodological perspective, we adopt a spatial econometric approach, which is particularly suitable for modelling and estimating spillover effects across countries and has not been employed in this strand of the literature so far. In particular, we relate countries' output volatility to (distance weighted averages of) other countries' output volatility, using descriptive measures, test statistics, and econometric estimates of cross-country interdependence. Moreover, based on bilateral measures of trade and financial linkages between countries, we also provide an assessment of the importance of alternative channels of interdependence, namely that of trade versus financial openness. Thereby we also suggest a new method to approximate unobserved bilateral data (on financial openness) from observed aggregate data.

We find that cross-country interdependence increased significantly till the mid 1980s, decreased slightly afterwards for a decade and returned to its highest levels in the 1990s. In quantitative terms, estimation results for the most recent period suggest that a uniform shock to output volatility in the world economy roughly quadruplicates through spillover effects of incipient shocks and the associated repercussions, which are equally transmitted through both trade and financial linkages.

These findings are remarkably close to other studies that have used entirely different methodological approaches and samples of countries, which underlines the generality of the results. They highlight the increased vulnerability of the world economy to local shocks, which are sizeably amplified due to the high degree of interconnectedness of the world economy. At the same time, they highlight the potential gains (multiplier effects) from international cooperation and policy coordination in the implementation of macroprudential stabilization policies.

The remainder of the paper is organized as follows. Section II outlines the empirical framework. Section III describes the evolution of the world economy's output volatility in the post-war period. Section IV presents estimates of cross-country interdependence of output volatility over time, whereas section V attempts to assess the role of trade and financial openness as transmission channels for cross-country spillover effects. Section VI summarizes the results and concludes.

II. Model Specification and Data

Focussing on a quantitative assessment of cross-country interdependence, we specify a cross-sectional model, which is repeatedly estimated for alternative time periods. Since increasing interdependence will be reflected in an increased propagation of both positive and negative

shocks, it is natural to consider some measure of volatility as dependent variable. In particular, we consider the volatility of GDP per capita growth over (rolling) time periods of 10 years. As it is standard in skedastic regressions, the natural log is used to rule out negative predicted values. For reasons of both data availability and parsimony, the baseline specification includes (a constant) and as single explanatory variable the (log of) initial GDP per capita as an indicator of economic development and institutional quality, which we expect to be negatively related to output volatility. Hence, for a given time period t , our cross-section model reads

$$\sigma_{i,t} = \alpha_t + \beta_t y_{0,i,t} + u_{i,t}, \quad (1)$$

where σ_i is the log of country i 's standard deviation of output growth over a 10-year period, calculated as standard deviation of the residuals from a regression of the log of GDP per capita on a time trend over the period t to $t-10$; y_0 is the log of country i 's initial GDP per capita, referring to the starting year of the period considered ($t-10$). GDP is measured in million 1990 international Geary-Khamis dollars. All data are taken from Maddison (2010), yielding the most comprehensive sample with 135 countries over the period 1955 to 2006. Since we consider 10-year intervals, our first observation refers to $t=1965$ and covers the period 1955 to 1965; we then move forward in time, using a moving 10-year time window, such that the most recent observation refers to $t=2006$, which covers the period 1996 to 2006. Hence, the total number of cross-sections considered is 42, each of them referring to a particular 10-year time period.¹

¹ Actually, data is also available for the first half of the 1950, but its use yielded implausible estimation results throughout the paper (such as a significant positive effect of GDP per capita on output volatility, huge changes in the estimates from moving the time window by one year). Hence, we confine our analysis to data starting in 1955. Apart from data quality, the most likely explanation is

To estimate the strength of cross-sectional interdependence, we adopt a spatial econometric approach. As most general specification we consider a spatial autoregressive model with spatial autoregressive disturbances. In matrix notation, the model reads

$$\boldsymbol{\sigma}_t = \alpha_t + \beta_t \mathbf{y}_{0,t} + \lambda_t \bar{\boldsymbol{\sigma}}_t + \mathbf{u}_t, \text{ with} \quad (2a)$$

$$\mathbf{u}_t = \rho_t \bar{\mathbf{u}}_t + \boldsymbol{\varepsilon}_t, \quad (2b)$$

where $\bar{\boldsymbol{\sigma}}_t \equiv \mathbf{W}\boldsymbol{\sigma}_t$ and $\bar{\mathbf{u}}_t \equiv \mathbf{W}\mathbf{u}_t$ are the spatial lags of the dependent variable and disturbances respectively. The elements of the $N \times N$ spatial weights matrix $\mathbf{W} = (w_{ij})$ reflect the structure of cross-country interdependence and will be specified as decreasing function of geographical distance between countries i and j . Since the weights matrices are time-invariant, we expect a reduction in economic distance (an increase in strength of cross-country interdependence) to be reflected in an increase in the spatial regressive parameters λ and ρ , which measures the strength of spillover effects.

Note that the ‘spatial lag’ of the dependent variable $\bar{\boldsymbol{\sigma}}_t = (\bar{\sigma}_{i,t})$ can be interpreted as weighted averages of other countries’ volatility, since $\bar{\sigma}_{i,t} = \sum_{j=1}^N w_{ij} \sigma_{j,t}$. An analogous interpretation holds for the spatial lag of the disturbances ($\bar{\mathbf{u}}_t$). In economic terms, the spatial regressive structure implies that shocks induce multilateral spillover effects.

As it is standard in the spatial econometrics literature, the $N \times N$ weights matrix \mathbf{W} has zero main-diagonal elements \mathbf{W} (ruling out self-influence) and is row-normalized (such that each row sums to 1) to ensure well-behaved asymptotics. Assuming that spillovers from country j

that many economies were about the return to their pre-war growth path in the early 1950s and that growth (and volatility) in the early 1950s was still driven by adjustments to potential output.

to country i are decreasing in distance between countries i and j ($Dist_{ij}$), the elements of the (unnormalized) weights matrix \mathbf{W}^0 are specified as:

$$w_{ij}^0 = \exp(-\delta Dist_{ij}), \text{ for } i \neq j \text{ and zero otherwise.} \quad (3)$$

Distance is measured in 1000 kilometers and the distance decay parameter is set to $\delta = 2.5$. This implies a half-life distance of shocks of some 280 kilometres; three quarters of a shock have faded away after 550 kilometers, and after 1000 kilometers, spillover effects of local shocks are reduced to 10 percent. Data on bilateral distances is from the CEPII database.

The final weights matrix \mathbf{W} is obtained by row-normalizing the weights matrix defined by (3), such that \mathbf{W} reflects the structure and strength of linkages between country i and j in relative terms, i.e, relative to all linkages of country i . With row-normalized matrices, model (2) implies that – in case of $\lambda = 0$ – a uniform shock to ε leads to an increase in volatility by $1/(1-\rho)$; similarly, for $\lambda \neq 0$, the implied effect amounts to $1/[(1-\rho)(1-\lambda)]$. Notice that model stability requires that $|\lambda| < 1$ and $|\rho| < 1$.

III. Output Volatility of the World Economy

Before turning to the estimation results, we briefly summarize the evolution of output volatility over the period 1960-2006 (Figure 1). We consider averages over all 135 countries of our sample (referred to as ‘world’) and three subsamples: the G-7 countries, 20 OECD countries, and the US. Notice that each observation refers to the period over the last 10 years.

< Figure 1 >

On average over the sample period, global output volatility amounted to 3.4 percent. Starting from a level of 2.5 percent in the mid 1950s, volatility increased steadily and peaked in the

period from the mid-1970s to the mid-1980s at a level of 4.8 percent, following the disruptions in the world economy through the two major oil price shocks in 1973 and 1979, resulting in a period of high inflation. In addition, a high level of discretionary fiscal and monetary policy, a period of turmoil in international financial markets, and the breakdown of the Bretton Woods system contributed to the increased instability in the first half of the 1980s.

Since then output volatility has declined down to some 3 percent in the period from 1996-2006.² As evident from Figure 1, the ‘great moderation’ in US output volatility, which was halved since 1985 was less pronounced at a global scale, where volatility decreased by one third only. The reduction in output volatility of OECD countries is of a comparable magnitude, but materialized with a delay. Whether these reductions in volatility were due to improvements in institutional quality, inventory management, macroeconomic policy or simply good luck (smaller and less synchronized shocks) is still subject to debate, though there is evidence that the reduction in the magnitude of common shocks has been the main reason (see Stock and Watson, 2005).

It is also worth emphasizing that – while for the subgroups of developed countries considered, whose output volatility today is significantly lower than it had been in 1950s – this is not true for the world economy. On average over all countries, output volatility in recent years (slightly above 3 percent) is comparable to the levels experienced in the 1960s.

IV. The Evolution of Cross-Country Interdependence, 1965-2006

In the following we provide some descriptive analysis and specification tests, before turning to more rigorous econometric estimates.

² Obviously, there has been a surge since 2007 in course of the pronounced recent financial and economic crises, causing sizeable reductions in GDP in many countries.

1. Prima-facie Evidence

As simplest measures of cross-country interdependence, we calculate – for 10-year moving time windows over the period 1965 to 2006 – the correlation (r_t) between output volatility σ_t and distance weighted averages of other countries' output volatility $\bar{\sigma}_t$, as well as the partial correlation ($r_{p,t}$), controlling for the initial level of GDP per capita ($y_{0,t}$). We next consider the least squares estimate $\hat{\lambda}_{LS,t}$ from equation (2a), which should be regarded as purely descriptive, however, since the spatial lag $\bar{\sigma}_t$ is endogenous. The three measures used so far relate to equation (2a) only, ignoring the possible cross-sectional dependence in the disturbance term as given equation by (2b).

As more comprehensive and systematic measures, allowing cross-country interdependence to appear through $\bar{\sigma}_t$ in (2a) or \bar{u}_t in (2b) or both, we consider Lagrange multiplier (*LM*) specification tests of model (2) suggested by Anselin et al. (1996). In particular, we calculate i) the LM test of the null hypothesis that $\lambda = 0$ and $\rho = 0$ (LM_t), ii) the LM test that $\lambda = 0$, assuming that $\rho = 0$ ($LM_{\lambda,t}$), and iii) the LM test that $\rho = 0$, assuming that $\lambda = 0$ ($LM_{\rho,t}$).

Figure 2 shows six panels, tracing the evolution of the aforementioned measures over the period 1965 to 2006. To reduce sampling variation, we also show the Hodrik-Prescott filtered series (as dashed line). Note that that LM test statistics are χ^2 -distributed with two degrees of freedom (LM_t) and one degree of freedom ($LM_{\lambda,t}, LM_{\rho,t}$), such that the critical values at the 5 percent level amount to 6 and 3.84, respectively.

< Figure 2 >

For all measures of cross-country interdependence of output volatility, the pattern is fairly similar. Interdependence was very low and typically insignificant in the 1950s and early 1960s, but increased continuously over time to peak in the late 1970s and early 1980s. Afterwards interdependence declined somewhat but remained at relatively high levels over the recent decade and shows a slight upward trend again in the most recent period. Notice that, since the model is estimated for each time period, the average level of output volatility is controlled for by the constant. Overall, results suggest that interdependence has become statistically and economically significant in the early 1980s and remained so since then.

These results hold up under alternative specifications: in fact we obtained qualitatively very similar results, when i) using the standard deviation of output growth instead of the regression-based volatility measures as defined above, and ii) varying the distance decay parameter δ in (3) between 1 and 5. The evidence so far is suggestive but should be regarded as descriptive. In the following we provide a more rigorous econometric analysis.

2. Estimation Results for Spatial Regressive Model

Before turning to the estimation of model (2), we determine whether cross-country interdependence should be modelled through the spatial lag of the dependent variable ($\bar{\sigma}$), the spatial lag of the error term ($\bar{\mu}$), or both. We follow the specification search strategy suggested by Anselin et al. (1996), which is based on a set of five LM tests. In addition to the LM considered above, they also provide ‘robust’ LM tests of the null hypothesis that $\lambda = 0$ (or $\rho = 0$), leaving the ‘other’ spatial regressive parameter ρ (or λ) unrestricted under the null hypothesis. The robust tests are referred to as LM_{λ}^* and LM_{ρ}^* respectively. Results for all five types of LM tests are typically indicative of the proper specification.

In the present paper, results are largely inconclusive. While all three LM tests illustrated in Figure 1 are significant at the five percent level for almost all periods since 1975, the robust LM test for the spatial lag and error model turned out insignificant for essentially the whole time period, each of them thus pointing to the relevance of the ‘other’ source of interdependence, i.e., to the spatial lag or the spatial error model. In light of the results for all five LM tests, we conclude i) that there is significant cross-sectional dependence either through the spatial lag or the spatial error (but not both simultaneously), and ii) that hence both the spatial lag model and the spatial error model are a legitimate choice for the empirical specification. This result is not too surprising, given the parsimonious specification of model (2) with a single explanatory variable.

For both economic and econometric reasons we opt for the spatial error model, i.e., the model given by (2a) and (2b) with λ_t set to zero. From an economic perspective, modelling shocks and output volatility, unrelated to changes in the level of development or institutional quality, through the error process is a natural choice. From an econometric perspective, it avoids estimation problems due to the endogeneity of $\bar{\sigma}_t$.³

Hence, we chose a spatial error model and estimate the spatial regressive parameter ρ_t using the general moments (GM) approach by Kelejian and Prucha (2010), which is robust to heteroskedasticity of arbitrary form in the error term ϵ . Application of this approach proceeds

³ Given the absence of other convincing instruments, the spatial lag model (2a) would require using spatial lags of y_0 as instruments for $\bar{\sigma}_t$. The corresponding two-stages least squares estimates of λ_t turned out implausible, with huge variation in the coefficients over alternative time periods, many of them outside the admissible parameter space $(-1,+1)$. The alternative route – using maximum likelihood (ML) estimation – has the drawback that it produces inconsistent parameters estimates under heteroskedasticity of unknown form (Lee, 2004).

as follows: First, the main equation (2a) (without spatial lag of the dependent variable $\bar{\sigma}_t$) is estimated by least squares, yielding consistent estimates of the disturbances \mathbf{u} . The estimated disturbances are then used in a second step to estimate the spatial regressive parameter (ρ) of the disturbance process (2a), using a GM estimator. In a third step, the main equation (2a) is re-estimated by (feasible) generalized least squares, which can be thought of as least squares estimation of the transformed model $\mathbf{T}\sigma_t = \mathbf{T}\alpha_t + \beta_t \mathbf{T}y_{0,t} + \mathbf{T}u_t$ with transformation matrix $\mathbf{T} = (\mathbf{I} - \rho\mathbf{W})$, where $\mathbf{T}u_t = \varepsilon_t$. For inference, Kelejian and Prucha (2010) derive the joint asymptotic distribution of the estimates of all model parameters, which is robust to heteroskedasticity of unknown form in ε .

< Table 1 >

Table 1 shows the estimates of the spatial error model for alternative time periods, starting from the period 1955-1965 up to the most recent period 1996-2006. In the main equation, the initial level of income shows the expected negative sign but is rarely significant with p-values slightly above 10 percent. The spatial regressive parameter of the disturbance process is both statistically and economically significant as of the 1970s; using its estimate for the most recent period considered, the results suggest that a uniform increase output volatility in all countries roughly doubles through international spillover effects.

< Figure 3 >

To provide a more comprehensive picture, Figure 3 shows the estimates for all 10-year time periods in the period from 1965 to 2006. Overall the descriptive results of the previous section are confirmed. We find a substantial increase in cross-country interdependence over time, which has peaked in the 1970s, then decreased somewhat over the 1980s but went up to statistically significant and at high levels since the 1990s again. The same pattern regarding

the evolution of spillover effects is obtained by Carare and Mody (2010), using factor-structural vector autoregressive models for a sample of 22 countries over the period 1960-2007. The similarity of our results for a large sample of 135 countries reinforces their finding that the evolution and rise in the contribution of international spillovers to volatility has played out in a remarkably similar manner across all country groups.

Our results also provide another angle to interpret the results by Stock and Watson (2005). They suggest that the reduced size of common shocks has contributed most to the great moderation of business cycles in the G-7 countries. In light of the present results, it is apparent that such a reduction of shocks implies a magnified reduction in volatility through a stabilization multiplier effect (which of course works into the opposite direction when the size of common shocks were to increase again, as it has been the case during the recent financial and economic crisis). At a more general level, this has an important and highly intuitive policy implication: Stabilization policy, e.g. through imposing fiscal rules to reduce volatility enhancing use of discretionary fiscal policy (Fatas and Mihov, 2004), is way more effective in a group of countries, when it is done in a coordinated way (or even at a supranational level). This should be borne in mind when (re)designing fiscal policy rules in a group of highly integrated countries such as the stability and growth pact in the European Union.

V. The Channels of Cross-Country Interdependence: Trade versus Financial Linkages

So far we have focussed on the evolution of cross-country interdependence over time, using a time-invariant weights matrix based on geographical distance. In the following, we provide an assessment of alternative channels of interdependence (for the most recent time period 1996-2006), using weights matrices that are based on measures of economic rather than geographical distance.

1. Empirical Framework

Assume for now that there are two known weights matrices \mathbf{W}_T and \mathbf{W}_F associated with trade and financial openness, respectively. Their exact definition and construction will be outlined below. In order to assess the relevance of these two channels of interdependence, we specify a second order spatial regressive error model:

$$\boldsymbol{\sigma} = \alpha + \beta \mathbf{y}_0 + \mathbf{u}, \text{ with} \quad (4a)$$

$$\mathbf{u} = \rho_T \mathbf{W}_T \mathbf{u} + \rho_F \mathbf{W}_F \mathbf{u} + \boldsymbol{\varepsilon}, \quad (4b)$$

where the matrices \mathbf{W}_T (\mathbf{W}_F) and the parameters ρ_T (ρ_F) reflect the structure and strength of cross-country interdependence related to trade openness and financial openness respectively. For estimation of model (5), we rely on Badinger and Egger (2010) who extend the GM estimation approach for first order spatial regressive models by Kelejian and Prucha (2010) to the case of higher order spatial regressive models.

2. Construction of Weights Matrices

A straightforward approach would be to include bilateral measures of trade and financial openness as elements of the matrices \mathbf{W}_T and \mathbf{W}_F . Unfortunately, this is not feasible for two reasons: First, weights matrices based on economic rather than geographic distance are likely to be endogenous, invalidating estimation and inference in spatial error models. Second, bilateral data, in particular those relating to financial openness, are available only for a small subset of our sample of 135 countries.

We thus suggest a new method to construct bilateral values from aggregate data using predicted values from bilateral gravity models. Without data on the dependent variable (such as measures of bilateral financial openness) the parameters of the bilateral model cannot be

estimated directly. The basic idea is to combine observed aggregate, country-specific data (such as financial openness) with observed explanatory variables in the bilateral gravity model (such as distance and country size). The parameters of the gravity model can then be ‘estimated’ indirectly such that the correlation between the aggregate predicted values (obtained as sum of the bilateral predicted values) with the observed aggregate data is maximized. By restricting the gravity to include exogenous variables only, this also provides a way to avoid endogeneity concerns, an idea introduced by Frankel and Romer (1999) in a slightly different context as ‘geographical gravity equation’.⁴ The technical details of the approach are outlined in Appendix B.

In particular, we use a simplified version of the geographical gravity model in Frankel and Romer (1999) that takes the following form:

$$\ln w_{ij}^0 = \gamma_0 + \gamma_1 \ln Dist_{ij} + \gamma_2 \ln Pop_i + \gamma_3 \ln Pop_j + \gamma_4 \ln Area_i + \gamma_5 \ln Area_j + \varpi_{ij}, \quad (5)$$

where w_{ij}^0 is the unobserved measure of bilateral (trade or financial) openness. The predicted values from equation (5) will be used to set up the (unnormalized) weights matrices.⁵ This approach motivated by the fact that the gravity model works well not only for trade, but also for FDI and portfolio investment (and thus also financial openness) and that geography plays a significant role in determining the spatial allocation of financial flows (Sariso Guerin, 2006).⁶ Data sources for the explanatory variables are as above.

⁴ Frankel and Romer (1999) consider a cross-country regression of per capita income on trade and suggest using as instrument for trade the country-specific sums of bilateral predicted values from a gravity model that includes geographical variables only.

⁵ Other geographical variables such as common border and landlocked dummies were omitted for the sake of parsimony. Since these variables show relatively little variation they hardly improve the predictive power of the model. As a consequence, their omission does not affect our main results.

⁶ Results by Sariso Guerin (2006) cannot be used directly since her sample, comprising at most 200 bilateral observations, is not representative for our large cross-section of 135 countries; moreover, her

We use the following aggregate, country-specific variables to generate corresponding bilateral (predicted) values: trade openness in terms of imports plus exports, FDI assets and liabilities, portfolio equity assets and liabilities, (portfolio and other investment) debt assets and liabilities, and finally total assets and liabilities, all of them expressed as a share of GDP. Trade openness is taken from the Penn World Tables 6.2. Measures of financial openness are from the dataset by Lane and Milesi-Feretti (2007).

< Table 2 >

Table 2 provides estimation results of the gravity model (5) for the aforementioned openness measures. Since, trade is available at a bilateral level for a reasonably large subsample of our 135 countries, results for model (5) using actual bilateral data on trade openness are reported in column (1a) for comparison.⁷ The coefficients are as expected: trade is decreasing in distance and area and increasing in country j 's size in terms of population. The coefficients in column (1a) are then used as starting values for the indirect estimation procedure for model (5), which is defined in equation (B2) in appendix B and solved numerically by a grid search strategy.⁸ Since the explanatory variables in the parsimonious model (5) are available for most countries, the aggregate predicted values are calculated by summing the bilateral predicted values not only over the 135 sample countries, but over a total of 195 countries, which better matches the definition of the actual data at the aggregate level.

model uses inflows as dependent variable, whereas we are more interested in overall openness measures, including both in- and outflows (relative to county size).

⁷ Data on bilateral trade are from the IFS direction of trade statistics are from Badinger (2008). The sample, on which the results in column (1a) are based, comprises 7574 non-zero bilateral observations and refers to the year 1996.

⁸ For each parameter, we consider intervals of ± 0.5 around the starting values, which corresponds to some 13 to 23-fold of the coefficients standard errors in the directly estimated gravity model.

The indirect gravity estimates in Table 2 are based on data for the year 2001, matching the sample mid-point of the cross-section models considered below. Column (1b) reports the indirect estimates using aggregate data on openness. Most coefficients, in particular that relating to geographical distance, are quite close to that from the ‘direct’ estimation. The correlation of the bilateral predicted values from the models in column (1) and (2) amounts to 0.919, the correlation of the implied aggregate values to 0.424.

Columns (2a) to (2d) show the corresponding indirect estimates for the financial openness measures. Overall results are close to that of the gravity model for trade. This is in line with results by Sarisoy Guerin (2006) in her comparative study on the performance of the gravity model for trade, FDI, and portfolio investment (for a sample of at most 200 bilateral flows). She finds that the estimated coefficients, in particular those for distance, are virtually identical for trade and FDI, whereas distance plays a smaller role for portfolio investment, which matches the results of the indirect estimates in Table 2. On the one hand this is a reassuring result. On the other hand, the similarity of the pattern for the alternative openness measures suggests that it will be difficult to identify their effects separately. In fact, the correlation between the aggregate (bilateral) predicted values is above 0.880 (0.950) across the board. This is not too surprising: Aizenman and Noy (2009), using causality tests and variance decompositions, show that trade and financial openness are intricately intertwined, with two-way feedbacks between both openness dimensions, making it difficult (if not impossible) for a country to increase trade openness while holding financial openness constant.

The estimates of the parameters $\gamma_0, \dots, \gamma_5$ in model (5) are then used to generate predicted values for the bilateral openness measures in levels according to

$$\tilde{w}_{ij}^0 = \exp(\ln \tilde{w}_{ij}^0), \text{ with} \tag{6}$$

$$\ln \tilde{w}_{ij}^0 = \tilde{\gamma}_0 + \tilde{\gamma}_1 \ln Dist_{ij} + \tilde{\gamma}_2 \ln Pop_i + \tilde{\gamma}_3 \ln Pop_j + \tilde{\gamma}_4 \ln Area_i + \tilde{\gamma}_5 \ln Area_j.$$

The elements \tilde{w}_{ij}^0 are then used to set up the (unnormalized) weights matrices. The final weights matrices for trade (\mathbf{W}_T), FDI (\mathbf{W}_{FDI}), portfolio (\mathbf{W}_P), debt (\mathbf{W}_D) and total financial openness (\mathbf{W}_F) are then obtained after row-normalization. Notice that the use of row-normalized matrices makes the choice of the constant $\tilde{\gamma}_0$ irrelevant, which enters the prediction in levels in a multiplicative way such that it cancels out by row-normalization.

3. Estimation Results for Second Order Spatial Regressive Model

In a first step, we check the reliability of our approach to constructing weights matrices from aggregate data. The first two columns in Table 3 compare the estimates of the spatial error model, using the trade based weights matrix, which is constructed from the direct gravity estimates using bilateral trade data (see column (1a) in Table 2) with those using the weights matrix based on the indirect gravity estimates from aggregate data (column (1b) in Table 2). The spatial regressive coefficients of the two models turn out virtually identical, suggesting that weights matrices based on the indirect estimation approach yield results comparable to those obtained from estimates based on actual bilateral data. We next turn to the estimation of the second order spatial regressive model given by (5) for alternative combinations of weights matrices, constructed from the indirect gravity estimates in columns (1b)-(2d) in Table 2.

< Table 3 >

Comparing the point estimates of the spatial regressive parameters across the different specifications from column (2)-(8) suggests that financial openness is a stronger transmission channel than trade openness, and that – within different subcategories of financial openness – debt linkages appear to be more important than (FDI or portfolio related) equity linkages. However, these results should not be overstressed. As expected, the high correlation of the elements of the weights matrices, resulting from the two-way linkages between trade and

financial openness (Aizenman and Noy, 2009), prevents us from identifying their effects in a statistically significant way: In none of the models we can reject the null hypothesis that the two spatial regressive parameters are identical. Overall, we thus conclude that both trade and financial linkages are important transmission channels volatility spillovers, and that there is no statistical significant evidence for a dominance of either of the two channels.

These findings are in line with Imbs (2004) as well as Easterly et al. (2000), who also find the increase in trade and financial openness to be important drivers of business cycle synchronization and output volatility spillovers. It should be added that our sample does not include the 2008-2009 economic crisis, where financial shocks, in particular those originating from the US, appear to have played an increased and dominant role according to Eickmeier et al. (2011).

We proceed with a restricted model, imposing equality of the two spatial regressive parameters associated with the weights matrices for trade openness and total financial openness. In particular, we use $(\mathbf{W}_T + \mathbf{W}_F)/2$ as single weights matrix, where the division by two ensures that the rows sum to one and that the coefficients are directly comparable with the estimates using a single weights matrix. Results for this preferred specification are given in the last column (9) of Table 3. The spatial regressive parameter turns out statistically significant and larger than that in the model using the distance based weights matrix (compare the last column of Table 1). The likely reason for this discrepancy in magnitude is that the weights matrix based on geographic distance only reflects the extent and structure of cross-country linkages less precisely than the weights matrices based on economic distance (in terms of trade and financial openness), causing an attenuation bias that results in a smaller estimate of the spatial regressive parameter. In economic terms, the estimates of our preferred model suggest that common shocks have a magnified impact and quadruplicate as a result of

international spillover effects, which are transmitted to an equal extent through trade and financial linkages.⁹

Notice that Carare and Mody (2010) find that around one half to three-quarters of total volatility is due to international spillovers in the period since the mid-1990s. Although their time series results are not directly comparable with those from the spatial cross-sectional approach used here, the quantitative effects implied by our estimates are very similar. A ‘destabilization multiplier’ of around four in response to a uniform shock across all countries suggests that one quarter of the total increase in volatility is due to the domestic shock, and that the remaining three quarters are triggered by international spillover effects. The similarity of the results with respect to the evolution and magnitude of international volatility spillovers, which are obtained using two alternative econometrics approaches and different samples of countries, is reassuring and underlines the robustness and generality of the findings.

The magnitude of spillovers effects is sizeable and suggest that international spillovers are a dominant source of output volatility. Results for the period of the recent financial and economic crisis by Eickemeier et al. (2011) point to an even stronger role of spillovers, in particular of those related to financial shocks emerging from the US, which explain 30 percent of the variation in GDP growth on average over all countries in 2008-2009, compared to a little less than 10 percent over the 1971-2007 period.

⁹ This argument is also supported by results of a time series regression of the estimated interdependence parameter on measures of trade and financial openness. As in the cross-section, the time series of the two openness measures are highly correlated (0.88) such that their effect cannot be disentangled. Regressing the estimate of ρ in Figure 3 on a joint openness measure, calculated as sum of trade and financial openness, over the period 1965 to 2006 suggests that an increase in openness by 10 percentage points increase the spatial regressive parameter by 0.007 (with a t-value of 2.384).

The finding that both trade and financial linkages turn out as important transmission channels also provides an explanation for the evolution of cross-country interdependence over time as illustrated in Figures 2 and 3, namely an increase up to late 1970s, a decrease in the 1980s, and an increase in the early 1990s again. One could argue that the increase in trade openness in the early post-war period has increased countries' vulnerability to foreign shocks, but that economic agents have become more alert to this phenomenon during the period of high volatility in the 1970s (see Figure 1) and have developed and adopted strategies to insure better against trade-related spillovers from foreign shocks (e.g., through increased diversification of firms' exports strategies and international risk sharing), leading to a decrease in the strength of cross-country interdependence. However, the potential of stabilizing effects from international diversification is limited and has been reduced by the increased synchronization of business cycles and shocks documented by Kose et al. (2003), which may be one explanation for the further increase of spillovers as of the 1990s.

Another explanation for the decrease and resurgence of spillover effects since the mid 1980s is provided by the results in Easterly et al. (2000), who find empirical support for a U-shaped effect of financial openness on volatility. Starting from moderate levels of financial openness in the 1970s, the increases in capital mobility during the 1980s seems to have first dampened volatility due to the possibility of production and consumption smoothing. Following its sharp increase in the mid-1980s and early 1990s, however, financial openness appears to have exceeded the threshold level associated with the turning point, as of which further increases in financial openness have destabilizing effects due to an increase in risk and the propagation of local shocks through the international financial system (e.g., when international investors pull short-term money out of a country, thereby causing credit constraints that enforce economic downturns).

Capital mobility appears to have emerged as a new transmission channel of international volatility spillovers, which is more complex, less transparent and less well understood than the trade channel, resurrecting countries' vulnerability to foreign shocks and increasing cross-country interdependence to the highest level in the post-war period.

VI. Conclusions

This paper provides evidence that the world has become significantly smaller in terms of economic distance over the last decades. The impact of shocks was rather local in the early post-war period, but has increased substantially up to the mid 1980s as a result of increases in trade and financial openness, and remained at a high level since then.

As the recent financial and economic crisis has shown, common shocks rapidly spread over the globe and have a magnified impact. According to our results, a global shock quadruplicates by propagating through the world economic system via both trade and financial linkages. The large magnitude of destabilizing multiplier effects underlines the importance of establishing a new international financial architecture. At the same time our results suggest that the potential gains from international policy coordination in terms of stabilizing multiplier effects are sizeable.

The results of the present paper also suggest some extensions for future research. In light of the recent financial and economic crisis it would be of particular interest to consider explicitly destabilizing spillover effects through macro-financial linkages and interactions between real instability and financial instability within and across countries.

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Appendices

Appendix A. Data Description

Our final sample comprises 135 countries, for which the main variables are available over the whole period 1955 to 2006. Table A1 provides a list of the countries used. Data on real GDP (in Geary Khamis PPPs) and population are taken from Maddison (2010, www.ggd.net/Maddison). Data on area and bilateral distances are from the CEPII database (www-cepii.fr). Measures of financial openness are calculated from the updated and extended version of the External Wealth of Nations Mark II database developed by Lane and Milesi-Ferretti (2007).

< Table A1 >

Appendix B. Indirect Estimation of Gravity Models from Aggregate Data

In the following we suggest a method to construct weights matrices involving bilateral relationships from aggregate data. This approach should be useful in a variety of applications considering multilateral interdependence, when there is no bilateral data on particular channels of interdependence, but when aggregate data is available and there is a theoretically motivated bilateral model with a set of observed explanatory variables.

Consider the bilateral model

$$l_{ij} = f(\mathbf{x}_{ij}, \boldsymbol{\beta}) + \varpi_{ij}, \quad (\text{B1})$$

where l_{ij} is the unobserved value of the dependent variable reflecting some bilateral linkage (e.g., financial openness) between countries i and j , and \mathbf{x}_{ij} is a $1 \times K$ vector of (bilateral or

country-specific) observed explanatory variables. Since actual data for l_{ij} are not available, the parameter vector β in model (B1) cannot be estimated directly by standard methods, e.g. by least squares ($\hat{\beta}$).

Now suppose that aggregate data on the variable l is available for each country i , given by

$l_i = \sum_{j=1}^{N_i} l_{ij}$ (e.g., country i 's total financial openness), where N_i is the number of country i 's

partner countries. Then, a reasonable approach to approximate bilateral (predicted) values on the variable l is to choose the 'estimate' of β such that the correlation between the aggregate predicted value (obtained as sum from the bilateral predicted values) and the actual value of the aggregate variable is maximized, i.e.,

$$\tilde{\beta} = \arg \max_{\beta} \text{Corr}[\mathbf{l}, \mathbf{l}(\beta)], \quad (\text{B2})$$

where $\mathbf{l} = (l_i)$ is a vector with observations of the aggregate, country-specific values of the variable l , and $\mathbf{l}(\beta) = [l_i(\beta)]$ is vector with the predicted counterpart (for the parameter values

β) with elements $l_i(\beta) = \sum_{j=1}^{N_i} l_{ij}(\beta)$.

In general, the (least squares) estimate $\hat{\beta}$ from (B1) and the indirect estimate $\tilde{\beta}$ from (B2) are not equivalent. However, we argue that it is reasonable to assume that the predicted values implied by the estimator in (B2) are a reasonable approximation of those implied by the direct estimate from model (B1). This will indeed be confirmed in the present paper in section V, subsections 2 and 3, for trade, where a reasonable subsample of bilateral observations is available.

Notice that the correlation between the aggregate predicted and actual values is invariant to overall level shifts, such the choice of the constant in is arbitrary in the first place, i.e., only the slope parameters are ‘identified’ by (B2). In the present paper, estimation of the constant in not required, since it is eliminated through the row-normalization of the weights matrices.

Table 1. *Estimates of spatial error model for alternative time periods*

	1965	1970	1975	1980	1985	1990	1995	2000	2006
constant	-3.709*** (0.520)	-3.106*** (0.837)	-3.144*** (0.806)	-2.69*** (0.706)	-2.97*** (0.572)	-3.31*** (0.478)	-2.853*** (0.469)	-2.218*** (0.47)	-2.908*** (0.538)
y_0	-0.001 (0.065)	-0.076 (0.106)	-0.044 (0.1)	-0.086 (0.087)	-0.046 (0.07)	-0.02 (0.058)	-0.074 (0.057)	-0.156*** (0.057)	-0.101 (0.064)
mean ¹⁾	-3.713	-3.666	-3.479	-3.347	-3.366	-3.482	-3.445	-3.466	-3.731
sd ²⁾	0.823	0.879	0.841	0.778	0.979	0.913	0.665	0.741	0.717
Disturbance process									
ρ	-0.160 (0.122)	0.233** (0.109)	0.306*** (0.102)	0.388*** (0.079)	0.353** (0.140)	0.165*** (0.052)	0.288*** (0.102)	0.210* (0.126)	0.447*** (0.124)
LM_ρ	0.822	0.747	4.782**	9.032***	15.759***	10.134***	8.147***	1.800	11.3***
LM_ρ^*	0.113	0.666	0.884	0.421	0.369	1.223	1.933	4.882**	0.051
LM_λ	0.825	0.899	5.164**	9.797***	16.193***	10.417***	7.693***	1.043	11.332***
LM_λ^*	0.116	0.818	1.266	1.185	0.802	1.506	1.479	4.125**	0.083
LM	0.938	1.565	6.048**	10.218***	16.561***	11.64***	9.626***	5.925*	11.384***
σ_u	0.823	0.873	0.838	0.767	0.975	0.912	0.663	0.727	0.704
σ_ε	0.818	0.869	0.815	0.727	0.934	0.891	0.639	0.718	0.662

Notes: Dependent variable is the log of σ_y . Main equation is estimated by feasible generalized least squares, using the transformation matrix

$(\mathbf{I} - \hat{\rho}\mathbf{W})$. Spatial regressive parameter (ρ) is estimated by GM (Kelejian and Prucha, 2010), based on least squares residuals. Standard errors in

parentheses are robust against heteroskedasticity in ε . ¹⁾ Mean of dependent variable. ²⁾ Standard deviation of dependent variable.

Table 2. Gravity 'estimates' for alternative openness measures

	(1a)	(1b)	(2a)	(2b)	(2c)	(2d)
	Trade ¹⁾	Trade	FDI	Portfolio	Debt	Total
$\ln DIST_{ij}$	-1.222	-1.342	-1.342	-1.162	-1.442	-1.342
$\ln Area_i$	-0.114	-0.726	-0.734	-0.602	-0.734	-0.734
$\ln Area_j$	-0.386	-0.134	-0.586	-0.198	-0.062	-0.122
$\ln Pop_i$	0.0170	0.249	0.417	0.537	-0.007	0.121
$\ln Pop_j$	1.268	1.064	1.300	0.812	0.648	0.648
Correlation ²⁾	0.324	0.650	0.610	0.356	0.629	0.657

Notes: ¹⁾ Column (1a) reports estimates of model (5) using bilateral trade data. Columns (1b)-(2d) report indirect estimates based on aggregate data, using the estimator defined in (B2) (see Appendix B). ²⁾ Correlation between actual and predicted values of the respective aggregate (country-specific) openness measure.

Table 3. *Estimates of first and second order spatial error model using weights matrices based on economic distance, 1996-2006*

	(1a)	(1b)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
W_1	Trade ¹⁾	Trade	Trade	Trade	Trade	Trade	FDI	FDI	Portfolio	Trade
W_2	-	-	FDI	Portfolio	Debt	Total	Portfolio	Debt	Debt	Total
constant	-2.938*** (0.525)	-2.984*** (0.527)	-2.867*** (0.519)	-2.895*** (0.601)	-2.999*** (0.613)	-2.966*** (0.603)	-2.891*** (0.474)	-2.903*** (0.564)	-2.902*** (0.573)	-2.934*** (0.571)
y_0	-0.092 (0.057)	-0.097* (0.058)	-0.095 (0.058)	-0.097 (0.059)	-0.099* (0.06)	-0.098 (0.06)	-0.101* (0.055)	-0.098 (0.061)	-0.098 (0.062)	-0.098 (0.060)
Disturbance process										
ρ_1	0.704* (0.372)	0.679** (0.321)	0.135 (0.455)	0.167 (0.445)	0.471 (0.346)	0.416 (0.288)	0.010 (0.370)	0.010 (0.265)	0.010 (0.711)	0.751** (0.322)
ρ_2	-	-	0.516*** (0.16)	0.639*** (0.182)	0.338 (0.416)	0.384 (0.351)	0.456 (0.674)	0.696* (0.409)	0.714 (0.632)	$\rho_1 = \rho_2$
$H_0: \rho_1 = \rho_2$ ²⁾			(0.523)	(0.41)	(0.854)	(0.954)	(0.667)	(0.294)	(0.595)	
σ_u	0.704	0.704	0.704	0.704	0.704	0.704	0.704	0.704	0.704	0.704
σ_ε	0.692	0.688	0.688	0.687	0.683	0.684	0.692	0.677	0.677	0.683

Notes: See Table 2. ¹⁾ Based on weights matrix implied by gravity estimates of in column (1a) of Table 2. ²⁾ p-value of Wald test that $\rho_1 = \rho_2$. Spatial regressive parameters in second order models estimated using by GM (Badinger and Egger, 2010).

Table A1. *List of countries*

Albania	Guatemala	Panama
Algeria	Guinea	Paraguay
Angola	Guinea Bissau	Peru
Argentina	Haïti	Philippines
Australia	Honduras	Poland
Austria	Hong Kong	Portugal
Bahrain	Hungary	Puerto Rico
Bangladesh	India	Qatar
Belgium	Indonesia	Romania
Benin	Iran	Russian Federation
Bolivia	Iraq	Rwanda
Botswana	Ireland	São Tomé and Príncipe
Brazil	Israel	Saudi Arabia
Bulgaria	Italy	Senegal
Burkina Faso	Jamaica	Seychelles
Burma	Japan	Sierra Leone
Burundi	Jordan	Singapore
Cambodia	Kenya	Somalia
Cameroon	Kuwait	South Africa
Canada	Laos	South Korea
Cape Verde	Lebanon	Spain
Central African Republic	Lesotho	Sri Lanka
Chad	Liberia	Sudan
Chile	Libya	Swaziland
China	Madagascar	Sweden
Colombia	Malawi	Switzerland
Comoro Islands	Malaysia	Syria
Congo 'Brazzaville'	Mali	Taiwan
Costa Rica	Mauritania	Tanzania
Côte d'Ivoire	Mauritius	Thailand
Cuba	Mexico	Togo
Denmark	Mongolia	Trinidad and Tobago
Djibouti	Morocco	Tunisia
Dominican Republic	Mozambique	Turkey
Ecuador	Namibia	Uganda
Egypt	Nepal	United Arab Emirates
El Salvador	Netherlands	United Kingdom
Equatorial Guinea	New Zealand	United States
Finland	Nicaragua	Uruguay
France	Niger	Venezuela
Gabon	Nigeria	Vietnam
Gambia	North Korea	Yemen
Germany	Norway	Zaire
Ghana	Oman	Zambia
Greece	Pakistan	Zimbabwe

Figure 1. Output volatility of the world economy and selected subgroups, 1955-2006 (in percent)

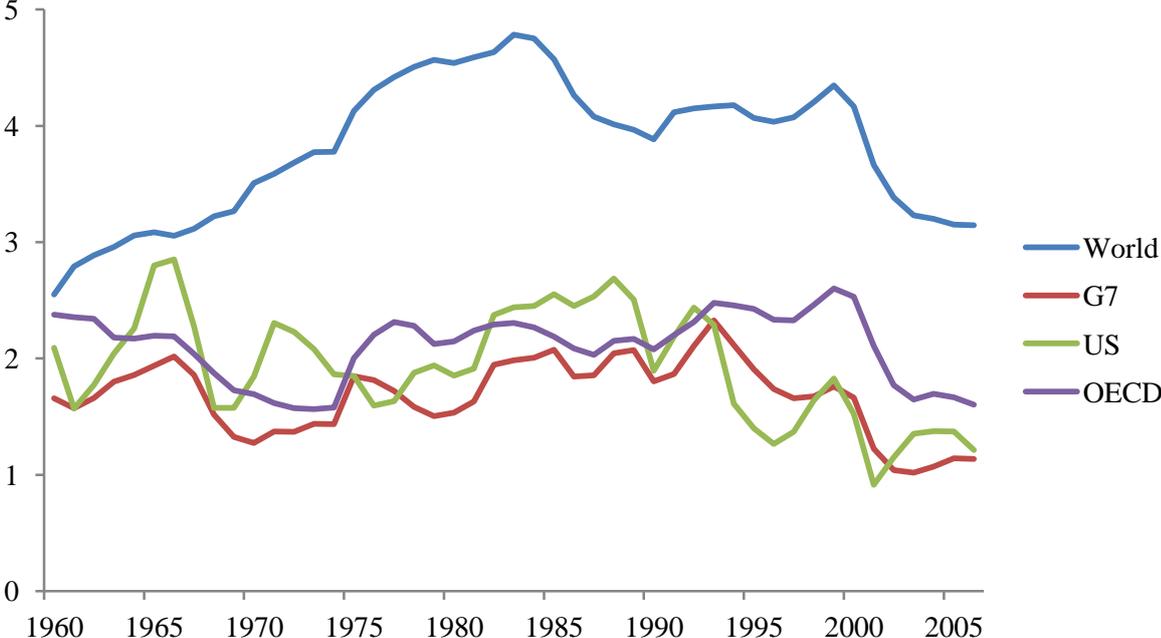


Figure 2. Descriptive measures of cross-country interdependence, 1955-2006

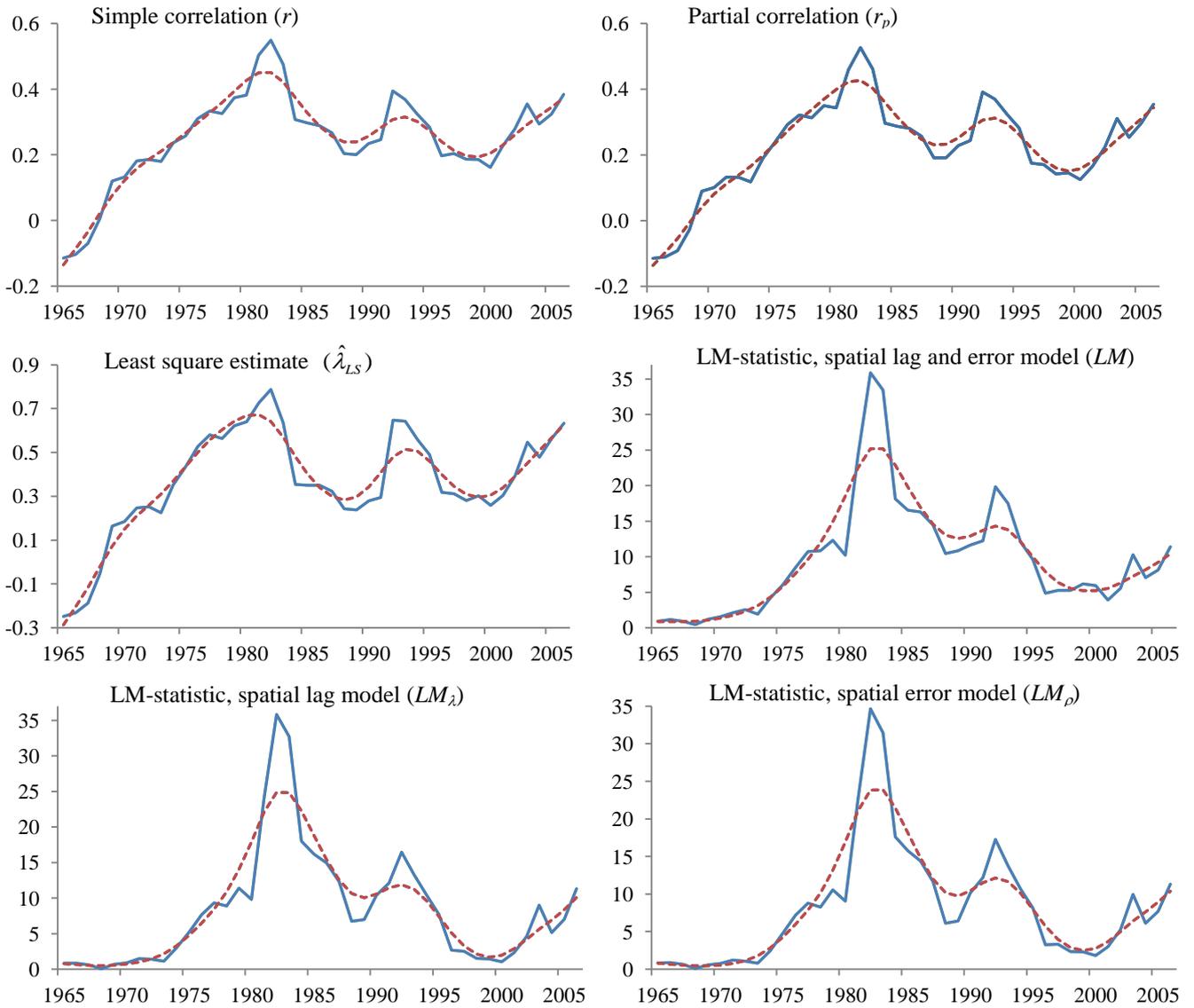


Figure 3. GM estimates of spatial regressive parameter ρ , 1965-2006

