TRADE AND BUSINESS CYCLE SYNCHRONIZATION IN OECD COUNTRIES A RE-EXAMINATION

ROBERT INKLAAR
RICHARD JONG-A-PIN
JAKOB DE HAAN

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Abstract

This paper re-examines the relationship between trade intensity and business cycle synchronization for 21 OECD countries during 1970-2003. Instead of using instrumental variables, we estimate a multivariate model including variables capturing specialisation, financial integration, and similarity of economic policies. We confirm that trade intensity affects business cycle synchronization, but the effect is much smaller than previously reported. Other factors in our model have a similar impact on business cycle synchronization as trade intensity. Finally, we find that the effect of trade on business cycle synchronisation is not driven by outliers and does not suffer from parameter heterogeneity.

JEL Code: E32, F42.

Keywords: business cycles, trade, synchronization of business cycles.

Robert Inklaar
Faculty of Economics
University of Groningen
PO Box 800
9700 AV Groningen
The Netherlands
R.C.Inklaar@rug.nl

Richard Jong-A-Pin
Faculty of Economics
University of Groningen
PO Box 800
9700 AV Groningen
The Netherlands
r.m.jong-a-pin@eco.rug.nl

Jakob de Haan
Faculty of Economics
University of Groningen
PO Box 800
9700 AV Groningen
The Netherlands
jakob.de.haan@rug.nl

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1. Introduction

In their seminal paper, Frankel and Rose (1998) argue that countries with more intense trade ties have more similar business cycles. This finding has been confirmed in almost all subsequent studies on the determinants of business cycle synchronization regardless of the way in which the trade relationship is modelled. For instance, Baxter and Kouparitsas (2004) find that bilateral trade intensity is robustly related to business cycle synchronization using the Extreme Bounds Analysis (EBA) of Leamer (1983) on a dataset that includes over 100 developed and developing countries.

This paper extends the literature in a number of directions. First, we employ corrected measures of business cycle synchronization. Frankel and Rose (1998) and almost all subsequent studies measure synchronization of business cycles of two countries as the bilateral correlation of some measure of (detrended) real economic activity.[1] Since the dependent variable lies between –1 and 1, the error terms in a regression model of the determinants of business cycle synchronization are unlikely to be normally distributed. We therefore employ transformed correlation coefficients as the dependent variable in our regression models using data for 21 OECD countries for the period 1970-2003.

Second, we examine the issue of endogeneity of trade in a more substantive way than previous studies. The basic problem here is that countries with intense trade relations are more likely to link their currencies, either explicitly or implicitly. This implies that these countries will have similar monetary policies – and possibly other policies – that may synchronize their business cycles. So it is not only trade that causes the business cycles to be correlated but also the similarity of economic policies. Neglecting these other variables in the regression specification renders the trade coefficient biased and inconsistent. Frankel and Rose (1998) and most subsequent studies therefore employ instrumental variables estimation, using gravity variables as instruments. We argue that this is not an adequate solution since the gravity variables are likely to affect other variables that influence business cycle synchronization as well, like participation in a currency union. Instead, we estimate a multivariate model including policy variables as

[1] An exception is the study by Otto *et al.* (2001).

well as structural characteristics and test for the proper estimation method using a Hausman (1978) test.

Third, we examine the effect of specialization on business cycle synchronization. If the degree of specialization between two countries is high, most trade will be inter-industry, and industry-specific shocks will lead to diverging business cycles. However, a dominant role for intra-industry trade can explain the positive association between trade and synchronization that has been found in the literature. Despite these theoretical arguments, this issue has received only scant empirical attention. Gruben *et al.* (2002) include inter-industry and intra-industry trade in their business cycle synchronization model and claim that the effects of both variables are different. We argue that this conclusion is based on unreliable estimates as the correlation between inter- and intra-industry trade is very high. Imbs (2004) accounts for the effect of inter-industry trade by including a measure of industrial specialization. Our approach is similar, but we not only look at industrial structure, but also at the structure of overall exports and the share of (bilateral) intra-industry trade to test the theoretical foundations of the trade relationship more directly.

Finally, we analyse to what extent the relationship between trade intensity and business cycle synchronization is robust across different country pairs. Is the effect of trade on business cycle synchronization the same for country pairs that are already highly synchronized, like Germany and the Netherlands, and countries which are not, like Germany and Japan, say? Or is the effect of trade on business cycle correlations driven by country pairs such as the US and Canada? To examine the importance of sample heterogeneity and outliers we use quantile regressions and least-trimmed squares, respectively.

Our main findings are the following. Trade intensity is found to affect business cycle synchronization, but the effect is much smaller than reported by Frankel and Rose (1998). We also find that apart from the level of trade, specialisation has a strong impact on business cycle synchronization. In addition, similar monetary and similar fiscal policies have a positive impact on business cycle synchronization. The impact of these factors on business cycle synchronization is about as large as the impact of trade intensity. Finally, our results suggest that the effect of trade on business cycle synchronization does not suffer from sample heterogeneity and is robust for outlying observations.[2]The remainder of the paper is organized as follows. Section 2 explains the methodology and section 3 discusses the data sources and methods. Section 4 presents the estimation results and discusses the economic relevance of our findings. Section 5 presents the quantile regressions and least trimmed squares results. The final section offers some concluding comments.

2. Methodology

Theoretically, trade intensity has an ambiguous effect on the co-movement of output. Standard trade theory predicts that openness to trade will lead to increased specialization in production and inter-industry patterns of international trade. If business cycles are dominated by industry-specific shocks, trade-induced specialization leads to decreasing business cycle correlations.[3] However, if trade is dominated by intra-industry trade industry-specific shocks may lead to more symmetric business cycles. Furthermore, in case of intensive trade relations economy-wide shocks in one country will generally have an effect on demand for goods from the other country.

The question how to disentangle the effect of intra-industry and inter-industry trade has been dealt with in different ways in the literature. Imbs (2004) includes an industrial specialization measure to capture the impact of inter-industry trade. Gruben *et al.* (2002) take a more direct approach and split up trade in inter- and intra-industry trade. In a regression in which both intra-industry and inter-industry trade are included, they find that intra-industry trade has a positive effect and that the effect of inter-industry is insignificant. An important problem with this approach is that intra-industry trade is highly correlated with inter-industry trade; in our dataset this correlation is 0.82. This

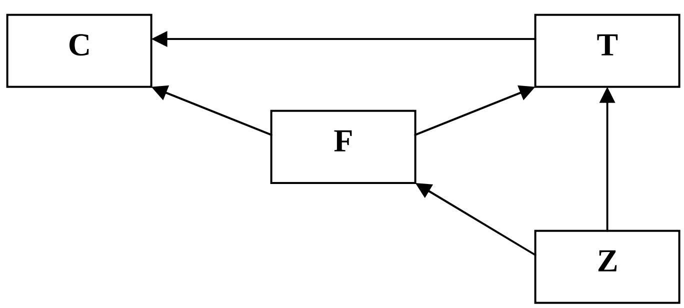
[2] The paper that comes closest to our is Imbs (2004), who also finds that the effect of trade on business cycle synchronization is less than that reported by Frankel and Rose (1998). There are, however, a number of important differences between both studies. Our methodology is quite different as we are primarily interested in the effect of trade intensity on output correlation. Furthermore, we consider a much longer list of potential determinants of business cycle synchronization. Imbs (2004), for instance, does not take the role of monetary and fiscal policy into account, which we find to be important. Imbs also does not examine how sensitive his findings are for sample heterogeneity and outliers.

[3] However, as pointed out by Frankel (2004), a positive shock at one point in the chain of value-added in one country will tend to have positive spill-over effects at the other points along the chain in other countries. Thus trade in inputs and intermediate products gives rise to positive correlations but may be recorded as inter-industry trade.

means that including both variables simultaneously leads to serious multicollinearity problems. Our approach is to take Imb's (2004) solution one step further and consider not only specialization measures based on industrial structure, but also measures based on the structure of exports, and the share of intra-industry trade. These measures will be discussed in more detail in the next section.

Frankel and Rose (1998) acknowledge the possible contrasting effects of inter- and intra-industry trade on business cycle synchronization, but focus on the net effect of total trade on output co-movement. However, even identifying the net effect of trade is not straightforward since trade intensity is endogenous, which makes an OLS regression of business cycle synchronization on trade intensity inappropriate. Frankel and Rose (1998) deal with this problem by using gravity variables (distance, border dummy, common language dummy) as instruments to identify the effect of trade on business cycle correlation. However, as pointed out by Gruben *et al.* (2002), this is not appropriate if the gravity variables (*Z*) not only affect bilateral trade intensity (*T*) but also are also possibly related to some other variables (*F*) that affect business cycle synchronization (*C*), as illustrated in Figure 1. For instance, neighbouring countries are more likely to coordinate their monetary policies, or even to have a common currency, than countries that are further away from each other. In turn, the introduction of a single currency will contribute to reducing trading costs both directly and indirectly, e.g., by removing exchange rate risks (and the cost of hedging) and diminishing information costs (De Grauwe and Mongelli, 2005).

Figure 1. The Relationship between Business Cycle Correlation, Trade, Gravity Variables and other Variables



The regression model that corresponds to the figure above is:

$$
\begin{aligned}
C &= \beta_1 T + \beta_2 F + \varepsilon \\
T &= c_1 Z + c_2 F + \mu \\
F &= c_3 Z + \omega
\end{aligned} \tag{1}
$$

The model shows that the business cycle correlation depends on bilateral trade as well as other policy-related and structural variables. Some of these variables may be influenced by the exogenous gravity variables, while, in turn, they may affect trade intensity. Broadly speaking, these variables can be grouped into the following categories: (1) specialisation (see, e.g., Imbs, 2004); (2) monetary integration (see, e.g., Rose and Engel, 2002); (3) financial integration (see, e.g., Imbs, 2004); and (4) similarity of fiscal policies (see, e.g., Clark and van Wincoop, 2001). Apart from these variables many others have been suggested that may be related to business cycle synchronization (see chapter 6 in De Haan *et al.* (2005) for an extensive discussion).

To identify the other variables to be included in our model, we follow Baxter and Kouparitsas (2004) and apply an Extreme Bounds Analysis to examine which variables are robustly related to business cycle synchronization in the OECD area. Using a much longer list of potential explanatory variables than examined by Baxter and Kouparitsas we identify a number of robust variables, including the similarity of monetary policy (proxied by the correlation of short-term interest rates) and the similarity of fiscal policy (proxied by the correlation of cyclically-adjusted budget deficits). In contrast to Baxter and Kouparitsas (2004) we employ the EBA as suggested by Sala-i-Martin (1997) since Leamer's (1983) EBA is extremely restrictive. Table A1 in the Appendix shows the variables that have been used in the analysis and whether they are robust explanatory variables of the business cycle correlation between two OECD countries. When testing for the robustness of these variables, we made sure not to include other proxies for the same "driving force" in the set of control variables. This is especially relevant for financial integration and specialisation, since we have three measures of financial integration and specialization (see section 3 for further details).

Once a suitable set of explanatory variables has been identified, the appropriate method to estimate the model above depends on the correlation between the error terms of the three equations. Given the exogeneity of gravity variables, it is crucial whether μ

and ε are correlated. If so, using OLS for the first equation results in inconsistent estimates and instrumental variables estimation should be preferred. If not, OLS estimates are consistent and at least as efficient. We use the Hausman (1978) test to resolve which estimation method should be chosen.

3. Data sources and methods

In our analysis we use two measures of economic activity, namely (quarterly) GDP and the (monthly) index of industrial production (IIP). The latter is attractive as it is available for a long period of time and (for most countries) at a monthly frequency. However, the coverage of the economy is limited to the manufacturing sector. The main reason for using GDP is that it is the most comprehensive measure of economic activity even though it is available at a quarterly frequency, at most, and time series are generally shorter than for industrial production. These trade-offs argue for using both measures.

Most previous papers on the determinants of business cycle synchronization (including Frankel and Rose, 1998) use the Hodrick-Prescott (HP) filter to detrend the original series. The HP filter can be interpreted as a high-pass filter that removes fluctuations with a frequency of more than 32 quarters and puts those fluctuations in the trend. Baxter and King (1999) argue that the combination of such a high-pass filter and a low-pass filter (which removes high frequencies) is better since the HP filter still leaves much of the high-frequency noise as part of the cycle. If such a so-called band-pass (BP) filter is applied, the resulting cyclical component does not contain any fluctuations with frequencies beyond the predetermined cut-off points. Since most studies find qualitatively similar results for different filtering methods, we restrict ourselves to the Baxter-King filter.[4]

Following most previous studies, our measure of business cycle synchronization is the correlation coefficient of the detrended measures of economic activity (GDP or IIP). Data is available for the period 1970 to 2003 for 21 OECD countries. Most countries report industrial production at a monthly frequency back to at least 1970.[5] Australia, New

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[4] Artis and Zhang (1997) and Calderon *et al.* (2002) conclude that the choice of filtering method is not crucial for their conclusions. Likewise, Massmann and Mitchell (2004), who consider the largest number of business cycle measures, report substantive similarities across alternative measures of the business cycle.
[5] Exceptions are Denmark (1974) and Ireland (1975).

Zealand and Switzerland only report quarterly industrial production, so their correlation vis-à-vis all countries is based on quarterly data.

Figure 2 shows the 8-year moving average of the correlation coefficients. This figure suggests that there is no obvious way to split our sample period in particular sub-periods, so we have split our sample into three periods of equal length (i.e. 11 years: 1970-1981, 1981-1992 and 1992-2003), leaving us with a maximum of 630 observations (0.5*(3*21*20)).[6] For the quantile regression results shown in section 4, we split the sample in eight periods of equal length in order to increase the number of observations.[7]

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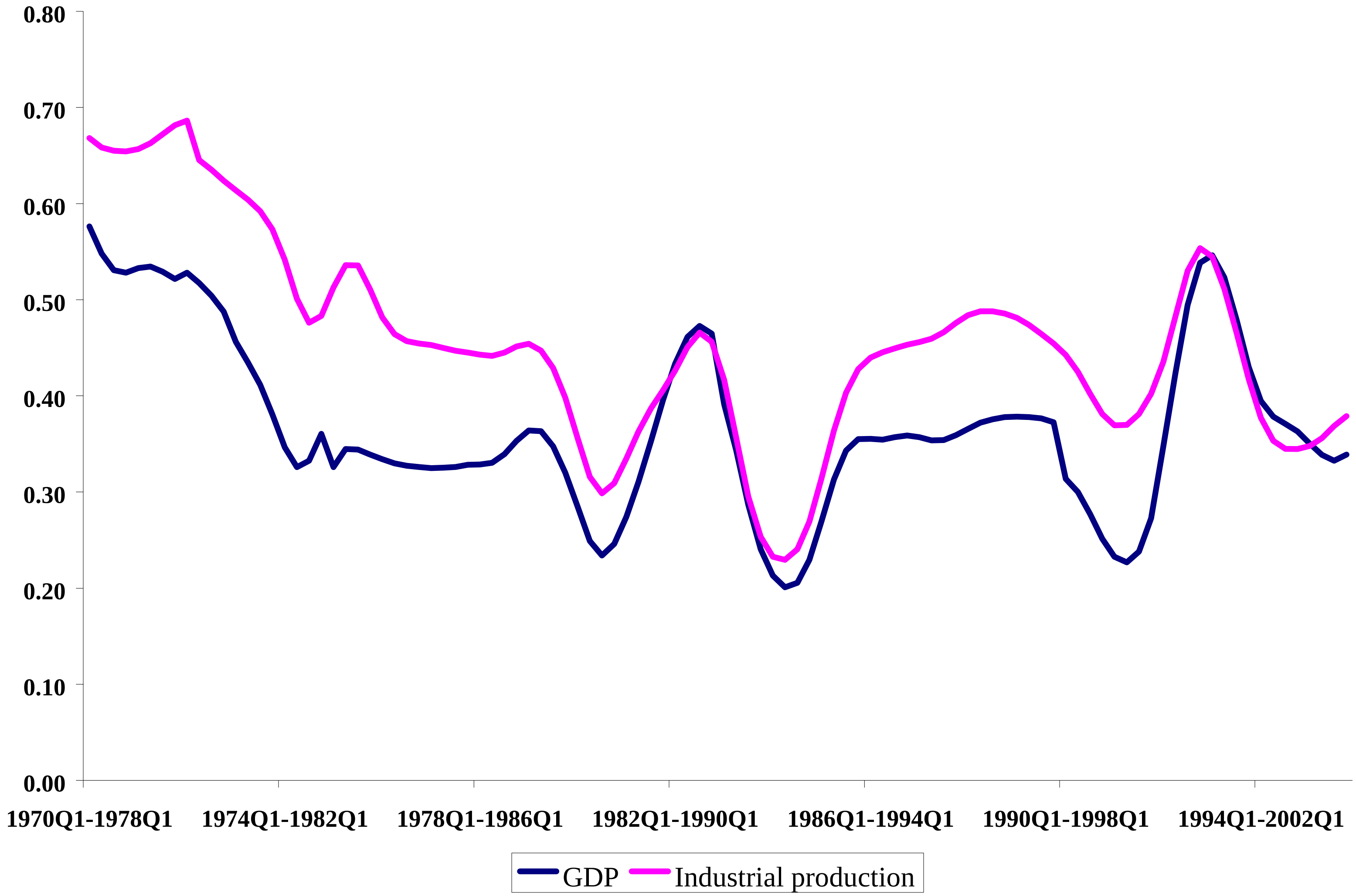
[6] Frankel and Rose (1998) followed a similar approach, using four periods of about 9 years.
[7] The results are generally robust to distinguishing from two up to eight different periods.

Figure 2. Average of bilateral business cycle correlations, 8-year moving windows, 1970-2003



In our regressions we use Fisher's z-transformations of the correlation coefficients as dependent variable. The transformed correlation coefficients are calculated as $C_t = 1/2 \ln((1+C)/(1-C))$, where C is the pair-wise correlation coefficient for each country couple. Since a (Pearson's) correlation coefficient is bounded at –1 and 1, the error terms in a regression model of the determinants of business cycle synchronization are unlikely to be normally distributed if the untransformed correlation coefficients are used. This complicates reliable inference. The transformed correlations do not suffer from this problem, since the transformation ensures that they are normally distributed (see David, 1949). This issue has not been addressed in most previous papers using these types of model, presumably under the assumption that the deviation from normality is sufficiently small. However, Figure 3a – showing kernel density estimates of the untransformed correlation coefficients – suggests that this conjecture is false and hence it is necessary to transform the dependent variable. Figure 3b shows that the transformed correlation coefficients are much closer to being normally distributed.[8]

[8] See also Otto *et al.* (2001).

Figure 3a. Estimated density plot of untransformed business cycle correlations

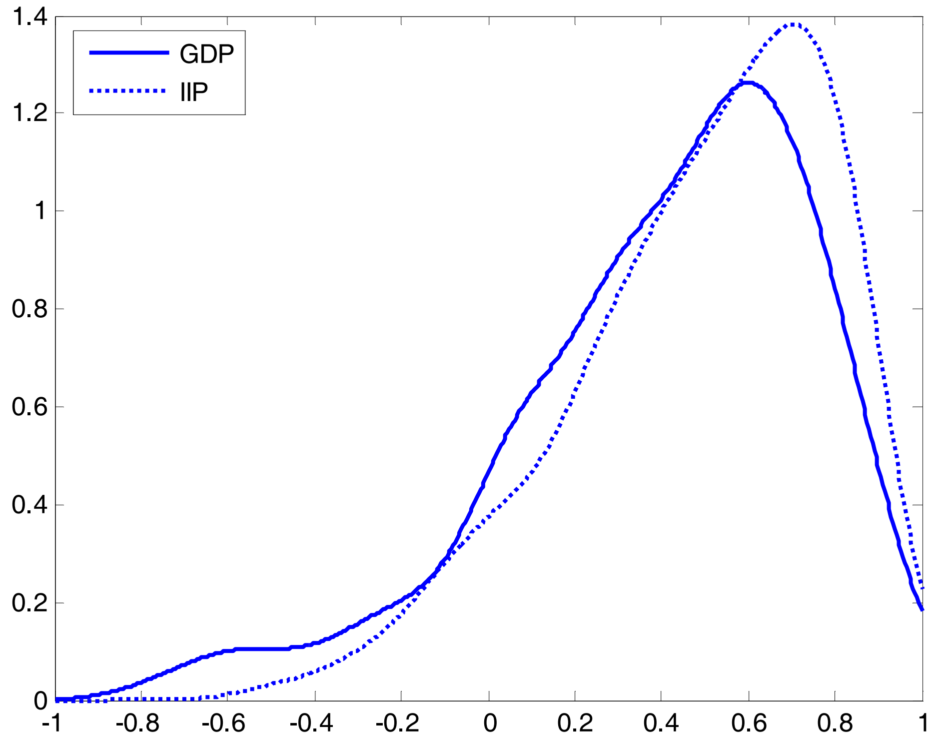
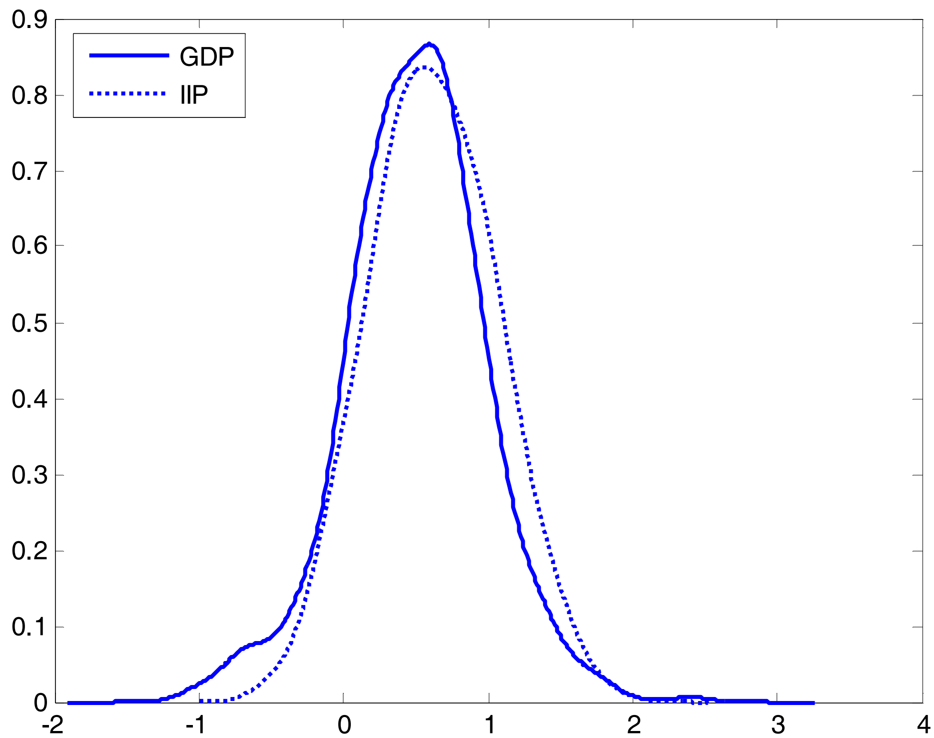


Figure 3b. Estimated density plot of transformed business cycle correlations



In previous studies on the determinants of business cycle synchronization various indicators for trade intensity have been used.[9] For instance, Frankel and Rose (1998) employ total trade (i.e. exports X and imports M) between two countries (i,j) scaled by total GDP (Y) or total trade.[10] Instead of using the sum of trade or GDP of the two countries as scaling factor, some authors prefer scaling by the product of GDP or trade of the two countries concerned (see, for instance, Clark and van Wincoop, 2001) as this indicator is not size-dependent. An alternative indicator is suggested by Otto *et al.* (2001), who take the maximum of:

$$\sum_t \frac{X_{ijt} + M_{ijt}}{Y_{it}}, \sum_t \frac{X_{ijt} + M_{ijt}}{Y_{jt}} \tag{2}$$

arguing that what matters is whether or not at least one country is exposed to the other. In this measure also trade can be used for normalization. We have calculated these six trade intensity measures. Table 1 shows the correlation matrix of these indicators. As these measures are (imperfect) proxies for trade intensity and it is not obvious which one has to be preferred, we combine them into a single measure using principal component analysis. Our trade intensity measure is therefore based on the common variation in the six individual trade intensity measures.[11] This combined measure is based on the largest eigenvalue and accounts for 64 percent of the total variance.[12]

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[9] The source for all our data on trade between countries is the new database by Feenstra *et al.* (2005).

[10] As pointed out by Otto *et al.* (2001), the first measure suffers from obscuring one-way interdependence, the second suffers from not measuring the relative importance of trade in the total economy. Note that when using GDP as a scaling factor, we convert GDP at current national prices to U.S. dollars using purchasing power parities from the OECD (2002) to take price differences between countries into account. All trade data are already converted using current exchange rates.

[11] However, we have also performed all analyses using the different trade intensity measures. Our results are robust for the selection of a particular trade measure (results available on request).

[12] The selection of one principal component is based on both the latent root criterion and the scree plot criterion. Furthermore, a measure based on the largest two eigenvalues has a correlation of 0.99 with the measure we use.

Table 1. Correlation coefficients between trade intensity measures

Correlation	TINT2	TINT3	TINT4	TINT5	TINT6
TINT1	0.52*	0.84*	0.73*	0.27*	0.58*
TINT2		0.58*	0.52*	0.60*	0.48*
TINT3			0.57*	0.29*	0.78*
TINT4				0.64*	0.57*
TINT5					0.51*

Notes: * denotes correlation significantly different from zero at 5% level. TINT1: bilateral trade, normalised by total trade of the two countries. TINT2: normalised by minimum of total trade of the two countries, TINT3: normalised by the product of total trade of the two countries. TINT4-6: same, but with GDP.

As discussed in the previous section, we use three indicators for *specialisation*, namely measures based on industrial specialisation, export similarity and the share of intra-industry trade. Imbs (2004) suggests the following measure for industrial specialisation:

$$\frac{1}{T}\sum_{t}\sum_{n=1}^{N}| s_{in} - s_{jn} | \qquad (3)$$

where $s_{n,i}$ denotes the GDP share of industry n in country i. We have constructed three measures based on industry specialisation. Apart from the index suggested by Imbs, we also use the squared differences – instead of the absolute difference of output shares as in equation (3) – as well as the correlation between the shares. Following Baxter and Kouparitsas (2004), we recast these specialisation measures as similarity measures by subtracting the specialisation measure from one. We have constructed these three similarity indicators using the 60-industry database of the Groningen Growth and Development Centre (GGDC, 2004), which has data on 56 industries covering the entire economy at the 2-digit and sometimes 3-digit level of industry detail (according to the ISIC revision 3 classification).[13] As might be expected, the three measures of output similarity are highly correlated (between 0.87 and 0.96), so following similar reasoning and criteria as for the trade intensity measures, we use the first principal component in the regressions as our first indicator of specialisation.[14]

[13] See www.ggdc.net for a more thorough documentation of this database, as well as the most recent version.
[14] The first principal component accounts for 94 % of the variance.

Furthermore, we follow Baxter and Kouparitsas (2004) and also consider the similarity of exports as our second main indicator for specialisation. As these authors point out, countries with similar baskets of traded goods will be affected similarly in the event of sector-specific shocks hitting their export and/or import sectors. Using the trade data by commodity (at the 4-digit SITC revision level of detail) of Feenstra *et al.* (2005), export shares are calculated for each country. The same three similarity measures as for output shares are calculated for export shares. The correlation between these export similarity measures varies between 0.54 and 0.84, but the first principal component accounts for 78% of the variance and is justified by the selection criteria. Therefore, it will be used as our second specialisation indicator.

As a final indicator for specialisation we use the intra-industry share, *IIT*. The variable *IIT* measures the share of bilateral trade that can be attributed to intra-industry trade. This index is defined as follows:

$$IIT_{ij} = 1 - \frac{\left|\sum_k \left(E_{ij}^k - E_{ji}^k\right)\right|}{\sum_k \left(E_{ij}^k + E_{ji}^k\right)} \qquad (4)$$

The share of intra-industry trade is calculated as one minus the absolute difference between exports of industry k from country i to country j and exports from country j to country i, divided by total bilateral trade (see Grubel and Loyd, 1971). We calculate these indices using the same source as for all our trade data, namely the new database by Feenstra *et al*. (2005). The trade data by commodity are allocated to industries using a detailed concordance.[15]

Financial linkages could result in a higher degree of business cycle synchronization by generating large demand side effects. For instance, a decline in a particular stock market could induce a simultaneous decline in demand in other countries if investors in these countries have invested in this particular stock market. Furthermore, contagion effects that are transmitted through financial linkages could also result in heightened cross-country spill-over effects of macroeconomic fluctuations. However, international

[15] Industries are defined at the 4-digit level of the international standard classification (ISIC rev. 2). See http://www.macalester.edu/research/economics/PAGE/HAVEMAN/Trade.Resources/TradeConcordances.html.

financial linkages could also stimulate specialization of production through the reallocation of capital in a manner consistent with countries' comparative advantages. We consider three indicators for *financial integration*: the correlation of changes in stock market indices, a dummy for capital account restrictions, and the (absolute) difference between the net foreign asset (NFA) positions of a country couple.[16] We collect the stock market data from the IMF's *International Financial Statistics* and calculate the correlation of annual growth rates. The capital account variable is based on information provided by Lane and Milesi-Ferretti (2001) and updated using the IMF publication *Exchange arrangements and exchange* restrictions, which gives an overview of capital and current account restrictions for each country. Our indicator equals one if at least one of the two countries had capital account restrictions during the period considered. For the NFA data, we again rely on Lane and Milesi-Ferretti (2001). They present two estimates, one based on cumulated current account data and one based on cumulated capital accounts. As the capital account-based measure is available for fewer years in most countries, we rely on the cumulated current accounts.

4. Estimation results

The first two rows of Panel A of Table 2 present our replication of the main results of Frankel and Rose (1998), i.e. the OLS and instrumental variables (IV) estimates of the effect of trade on business cycle correlation. In addition to the instruments used by Frankel and Rose (1998), i.e. distance, an adjacency dummy, and a dummy for common language, we also use a variable measuring geographical remoteness and a dummy for common legal origin.[17]

The OLS and IV estimates of the trade coefficient are positive and highly significant and comparable for the two measures of economic activity. Like Frankel and Rose, we find that the coefficients are lower and less significant when bilateral trade

[16] The latter two measures are also employed by Imbs (2004).

[17] All these instruments are highly significant in explaining trade intensity and the F-statistic of the first-stage regression is 157. Legal origin has also been used to directly explain output co-movement (e.g. Otto *et al.*, 2001) but we argue that the main effect of a common legal origin is via trade: the correlation between legal origin and trade intensity is 0.40, while the correlation with the GDP and IP correlations are 0.23 and 0.11, respectively. As the 95% lower bound of the legal origin-trade intensity correlation is 0.27, the link with trade is significantly stronger than the link with output correlations.

intensity is normalized by output. The IV estimates are similar in magnitude as those reported by Frankel and Rose (1998) and considerably higher than the OLS estimates.

Row 3 of panel A of Table 2 shows the results using our preferred indicator of trade intensity (the first principal component of six different measures of trade), while row 4 presents the findings if we transform the dependent variable. The coefficients of our preferred trade indicator are highly significant, suggesting that the qualitative conclusion that trade intensity is positively related to business cycle correlation is not sensitive to the measurement of trade intensity. Transforming the dependent variable yields higher coefficients, but due to the transformation it is not straightforward to compare the coefficients with the estimates of rows 1-3. In order to make a meaningful comparison, Panel B of Table 2 presents the impulse responses of a one standard deviation shock of the trade measure on the business cycle correlation.[18] We not only show the point estimates, but also the 95% upper and lower bound. These results suggest that the use of the transformed dependent variable leads to a somewhat stronger impact of trade on business cycle synchronization.

[18] The impulse response for the model with transformed correlation coefficients is calculated by running the reverse transformation on the estimated impulse response.

Table 2. Replication of the Frankel-Rose model using our data (effect of trade intensity on output correlation)

Panel A	OLS		IV	
	IIP	GDP	IIP	GDP
(1) Bilateral trade, normalised by total trade	0.031	0.025	0.060	0.061
	(6.5)	(4.3)	(7.1)	(5.4)
(2) Bilateral trade, normalised by total GDP	0.009	0.010	0.016	0.016
	(6.7)	(6.3)	(7.7)	(6.2)
(3) Bilateral trade, factor score	0.074	0.086	0.125	0.140
	(7.1)	(6.2)	(8.3)	(6.7)
(4) Bilateral trade, factor score, transformed correlation	0.127	0.125	0.204	0.203
	(7.0)	(6.0)	(8.4)	(6.7)
Hausman test (H0: OLS is consistent; critical 5% value: 6.0)				
Bilateral trade, normalised by total trade			21.0	18.3
Bilateral trade, normalised by total GDP			24.6	11.4
Bilateral trade, factor score			22.2	13.3
Bilateral trade, factor score, transformed correlation			24.5	14.5
Panel B				
Impulse response				
Bilateral trade, normalised by total trade	0.08	0.07	0.07	0.08
Bilateral trade, normalised by total GDP	0.08	0.09	0.08	0.08
Bilateral trade, factor score	0.07	0.08	0.10	0.11
Bilateral trade, factor score, transformed correlation	0.13	0.12	0.16	0.15
[Lower bound response – Upper bound response]				
Bilateral trade, normalised by total trade	[0.06 - 0.11]	[0.04 - 0.10]	[0.05 - 0.09]	[0.05 - 0.11]
Bilateral trade, normalised by total GDP	[0.06 - 0.10]	[0.06 - 0.12]	[0.06 - 0.10]	[0.06 - 0.11]
Bilateral trade, factor score	[0.05 - 0.09]	[0.06 - 0.11]	[0.08 - 0.12]	[0.08 - 0.14]
Bilateral trade, factor score, transformed correlation	[0.09 - 0.16]	[0.08 - 0.16]	[0.12 - 0.20]	[0.11 - 0.19]

Note: t-statistics, consistent for heteroscedasticity, are in parentheses.

Table 3 shows our estimation results for the model outlined in Figure 1. For the variables to be included in *F*, we rely on the results of the Extreme Bounds Analysis (EBA) as described in the Appendix. Our approach is to run a separate analysis for each combination of financial integration and specialization measures. For the financial integration measures we find that only the correlation of stock returns is a robust explanatory variable for synchronization while the capital account restrictions and NFA measures fail to pass the test. We therefore only show regressions with the stock market indicator. In contrast, all three specialisation measures appear robustly related to business cycle synchronization and are therefore each included in a separate regression model.[19]

It follows from Table A1 that apart from the correlation of stock market returns and the specialisation measures also some other variables are considered robust. The correlation of short-term interest rates and the correlation of cyclically-adjusted budget deficits are robustly related to business cycle synchronization no matter whether we focus on GDP correlation or IP correlation. For the GDP-based measure of synchronization, exchange rate variability is also robust.[20]

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[19] The measure of industrial similarity does not pass the test with GDP as the dependent variable, but we include it to facilitate the comparability of results across specifications.
[20] For the IP correlations, measures reflecting differences in capital stocks and arable land are also robust for some combinations of financial integration and specialization measures. Since they frequently fail this test and are also not robustly related to the GDP-based measure of synchronization, we have not included them here.

Table 3. Effect of trade intensity on output correlation using a structural model

Specialisation measure:	*Industrial similarity*		*Export similarity*		*Share of intra-industry trade*	
GDP	OLS	IV	OLS	IV	OLS	IV
Trade	0.043	0.054	0.053	0.121	0.044	0.115
	(2.1)	(2.1)	(2.6)	(3.4)	(2.0)	(2.8)
Specialisation measure	0.032	0.031	0.064	0.050	0.346	0.177
	(1.3)	(1.3)	(3.1)	(2.3)	(2.2)	(1.0)
Correlation of short-term interest rates	0.239	0.236	0.124	0.112	0.129	0.130
	(4.3)	(4.2)	(2.2)	(1.9)	(2.6)	(2.2)
Correlation of cyclically-adjusted budget deficits	0.172	0.171	0.143	0.137	0.136	0.133
	(4.7)	(4.7)	(3.8)	(3.6)	(3.6)	(3.5)
Correlation of stock markets	0.308	0.303	0.214	0.202	0.225	0.216
	(3.9)	(3.8)	(3.3)	(3.1)	(3.5)	(3.4)
Exchange rate variability	-1.600	-1.513	-1.552	-1.089	-1.548	-1.165
	(-3.3)	(-3.0)	(-3.4)	(-2.2)	(-3.4)	(-2.4)
IIP						
Trade	0.080	0.088	0.069	0.113	0.043	0.080
	(3.8)	(3.4)	(3.7)	(4.9)	(2.1)	(3.0)
Specialisation measure	0.070	0.069	0.118	0.105	0.761	0.657
	(4.0)	(3.8)	(7.2)	(6.5)	(6.8)	(5.6)
Correlation of short-term interest rates	0.374	0.372	0.221	0.217	0.211	0.214
	(8.9)	(8.9)	(5.2)	(5.1)	(5.1)	(5.2)
Correlation of cyclically-adjusted budget deficits	0.125	0.106	0.157	0.155	0.143	0.143
	(3.7)	(3.7)	(5.2)	(5.1)	(4.7)	(4.7)
Correlation of stock markets	0.161	0.156	0.064	0.057	0.082	0.077
	(2.7)	(2.6)	(1.2)	(1.1)	(1.6)	(1.5)
Hausman test (H0: OLS is consistent, critical 5% value: 12.6)						
GDP		0.32		6.83		5.00
IIP		0.28		8.51		3.90

Notes: constant included; t-statistics, consistent for heteroscedasticity are in parentheses.

It follows from Table 3 that almost all explanatory variables are significant with the expected sign. So more correlated monetary policy, fiscal policy, more similar industrial and export structures, more intra-industry trade, and less exchange rate variability are related to more similar business cycles.

The main finding in Table 3 is that the trade coefficients are much smaller than those previously found: the coefficient of trade intensity with GDP correlation as dependent variable is only half as large as in Table 2 for both the OLS and IV specification. In addition to the gravity variables that were used as instruments in Table 2, the other explanatory variables are included as instruments too; this specification corresponds to the second line of equation (1). The Hausman tests confirm that the model specification has improved compared to Table 2: the tests no longer reject the null hypothesis that the OLS estimates are consistent. Because Frankel and Rose (1998) did not specify a full model, they overestimated the impact of trade on output correlation.

Figure 4. Impulse response of explanatory variables (intra-industry specification, absolute values)

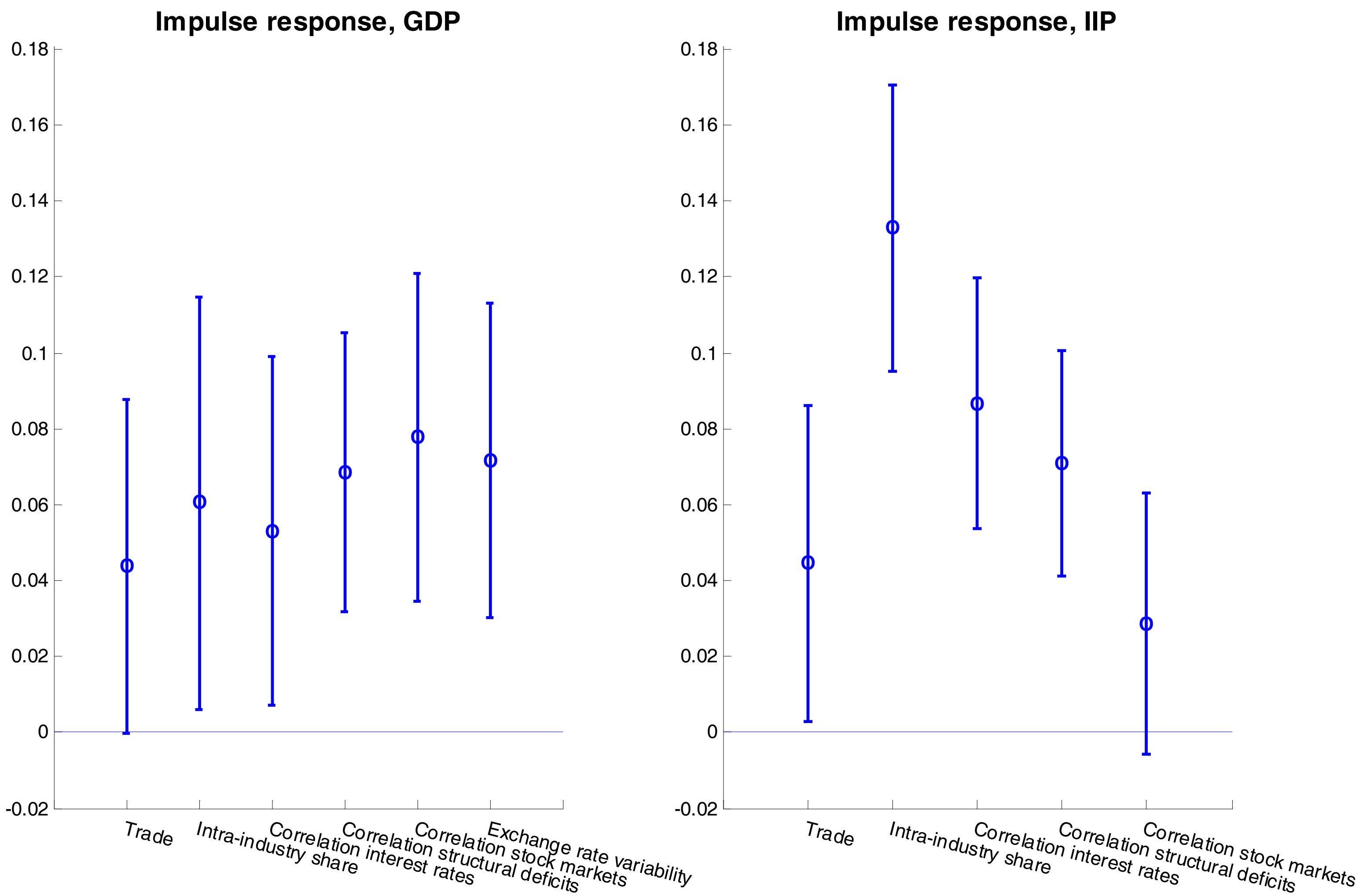
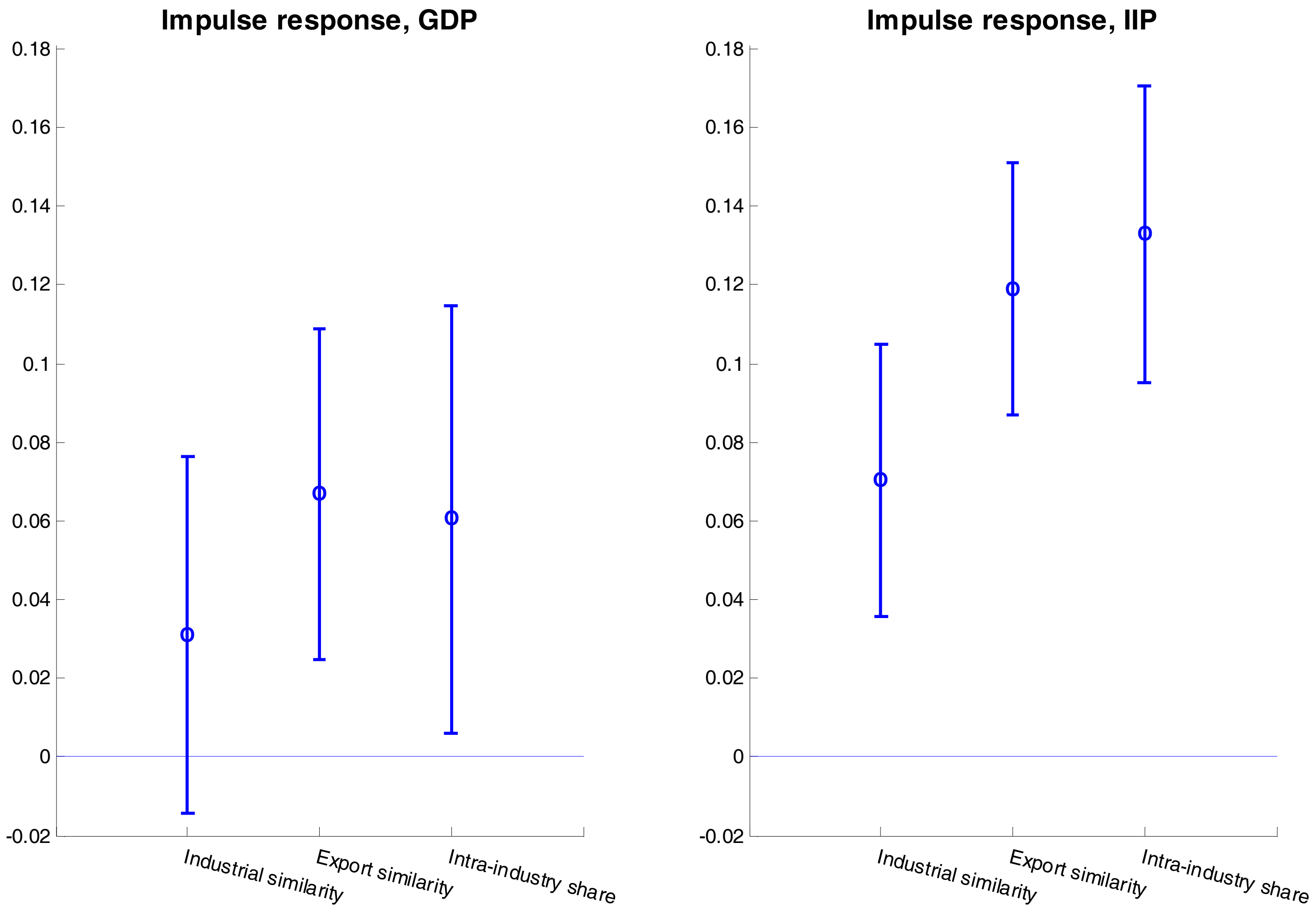


Figure 4 shows the impulse response of an increase of one standard deviation of all the variables included in the model with *IIT* as specialisation measure. The point estimate, as well as the 95% upper and lower bounds are shown. It follows that the point estimate of the impact of almost all variables – like the correlation of short-term interest rates or of cyclically-corrected budget deficits – is larger than the impact of trade intensity. In view of the upper and lower bounds, we cannot conclude that these differences are statistically significant. Still, our evidence suggests that variables that reflect common economic policies and specialisation are at least as important as strong trade ties for synchronization of business cycles.

Finally, Figure 5 compares the impulse response of an increase of one standard deviation of the three specialisation measures that we use. Again, the point estimate as well as the 95% upper and lower bounds are shown. It follows that the point estimate of the impact of industrial similarity is the lowest. In view of the upper and lower bounds of the impulse responses one has to be careful in drawing too strong conclusions, but the evidence suggests that trade-based specialisation measures have a larger impact on business cycle synchronization than industry-structure-based specialisation measures. This is most visible for the impulse responses of the models based on industrial production.

Figure 5. Impulse response of specialization measures

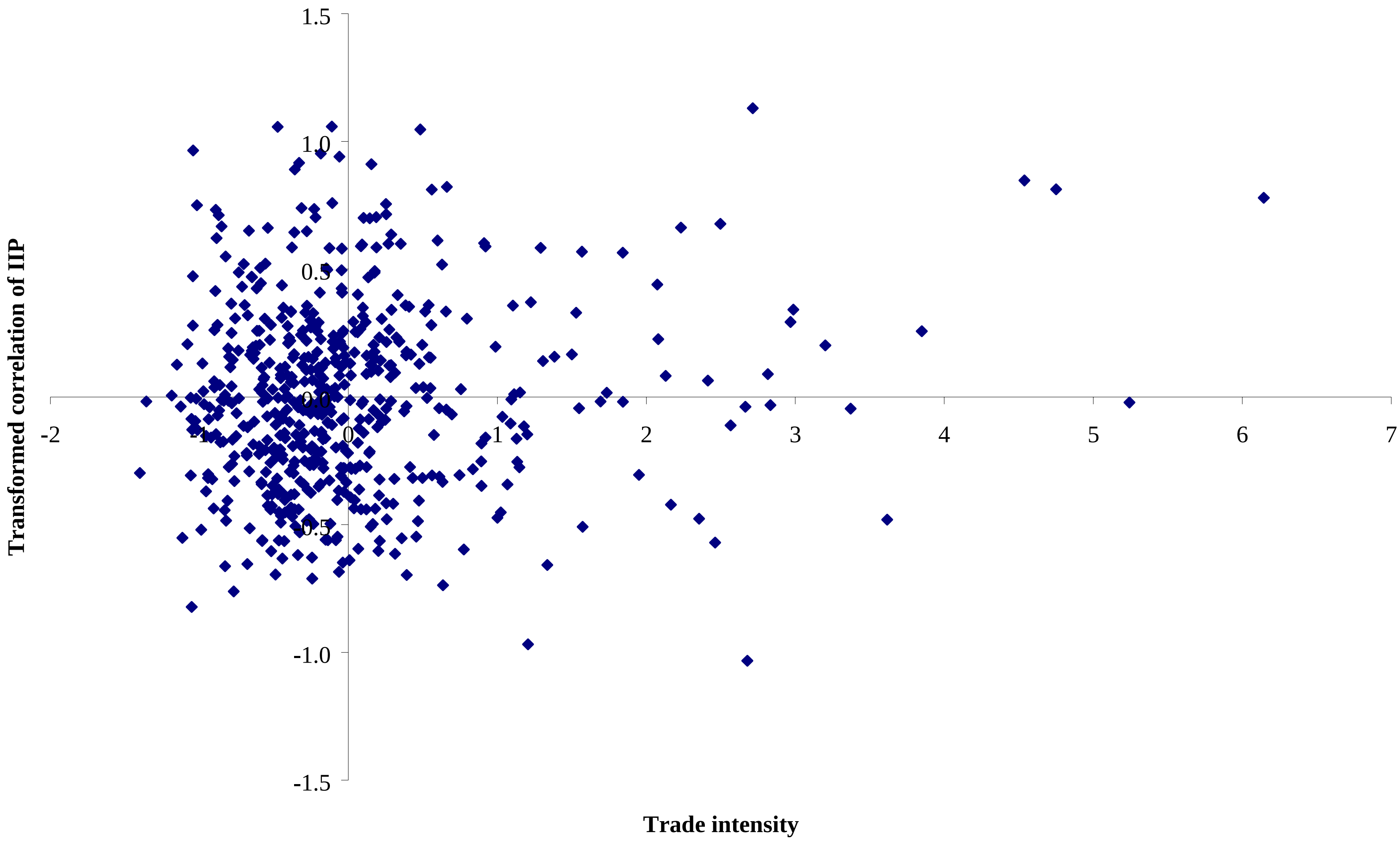


5. Sample heterogeneity and outliers

So far we have focused on the conditional mean of business cycle correlations as a linear function of bilateral trade and other structural and policy related variables. However, it is well known that outliers in the regressand as well as the regressors may seriously influence these OLS estimates. Figure 6, which shows a scatter diagram of industrial production correlations and trade (after conditioning on control variables), suggests that there are various observations that are quite far away from the bulk of the observations and these may drive our results.[21] In this section we therefore report the estimation results using the Least Trimmed Squares (LTS) estimator of Rousseeuw (1984, 1985) to identify outlying observations. Furthermore, we employ quantile regressions to examine sample heterogeneity (see Koenker and Basset, 1978 or Koenker and Hallock, 2001 for a non-technical overview).

[21] Figure 6 shows the residuals of the regression of business cycle correlation for industrial production on the control variables against the residuals of the regression of bilateral trade on these same control variables.

Figure 6. Scatter diagram of industrial production correlations and trade (after conditioning on control variables)



The basic principle of LTS is to fit the majority of the data, after which outliers may be identified as those points that lie far away from the robust fit. LTS typically minimizes the sum of squares over half the observations, the chosen half being the combination, which gives the smallest residual sum of squares. Although this method is particular suited to identify leverage points, it is not suited for inference. As proposed by Rousseeuw (1984), this can be resolved by using re-weighted least squares (RWLS). A simple, but effective, way is to give a weight of zero to all observations identified as outliers and a weight of one to all other observations (Sturm and De Haan, 2005).

Table 4 shows the results of the LTS/RWLS estimates. For comparison purposes, we first repeat the OLS results of Table 3. Overall, there are no large differences between the OLS estimates and the robust estimates. However, there are exceptions. In the models for the GDP-based correlations, the bilateral trade coefficient loses significance in some specifications. This is quite remarkable, as almost all other variables remain significant at the 5% level. Still, in the models for industrial-production-based correlations the significance of the trade variable increases. So we therefore conclude that, in general, the effect of trade on business cycle synchronisation is not driven by outliers.

Table 4. OLS vs LTS/RWLS

Specialisation measure:	*Industrial similarity*		*Export similarity*		*Share of intra-industry trade*	
GDP	OLS	LTS/RWLS	OLS	LTS/RWLS	OLS	LTS/RWLS
Trade	0.043	0.044	0.053	0.033	0.044	0.022
	(2.1)	(2.2)	(2.6)	(1.8)	(2.0)	(1.1)
Specialisation measure	0.032	0.041	0.064	0.059	0.346	0.354
	(1.3)	(2.0)	(3.1)	(3.3)	(2.2)	(2.9)
Correlation of short-term interest rates	0.239	0.274	0.124	0.207	0.129	0.177
	(4.3)	(5.5)	(2.2)	(4.6)	(2.6)	(3.9)
Correlation of cyclically-adjusted budget deficits	0.172	0.191	0.143	0.161	0.136	0.160
	(4.7)	(5.3)	(3.8)	(4.8)	(3.6)	(4.7)
Correlation of stock markets	0.308	0.266	0.214	0.138	0.225	0.158
	(3.9)	(3.9)	(3.3)	(2.7)	(3.5)	(3.1)
Exchange rate variability	-1.600	-1.373	-1.552	-1.920	-1.548	-1.768
	(-3.3)	(-3.3)	(-3.4)	(-4.9)	(-3.4)	(-4.5)
IIP						
Trade	0.080	0.092	0.069	0.074	0.043	0.048
	(3.8)	(5.6)	(3.7)	(4.9)	(2.1)	(3.0)
Specialisation measure	0.070	0.056	0.118	0.136	0.761	0.838
	(4.0)	(3.3)	(7.2)	(8.4)	(6.8)	(8.4)
Correlation of short-term interest rates	0.374	0.392	0.221	0.267	0.211	0.268
	(8.9)	(9.5)	(5.2)	(6.8)	(5.1)	(6.9)
Correlation of cyclically-adjusted budget deficits	0.125	0.117	0.157	0.186	0.143	0.141
	(3.7)	(3.6)	(5.2)	(6.1)	(4.7)	(4.8)
Correlation of stock markets	0.161	0.223	0.064	0.047	0.082	0.101
	(2.7)	(3.9)	(1.2)	(1.1)	(1.6)	(2.3)

Notes: constant included; t-statistics, consistent for heteroscedasticity are in parentheses.

Quantile regression is an appropriate tool to address sample heterogeneity across different quantiles as shown by Koenker and Basset (1978). OLS focuses on the mean of the dependent variable given the explanatory variables. Quantile regressions are used to analyze other parts of the conditional distribution, such as the (conditional) median or specific deciles. In order to increase the degrees of freedom, we divide the sample period 1970-2003 into eight different periods and ran the same regressions as in Table 3.

Figure 7 shows the estimated coefficients of the trade intensity variable for each decile, using the model in which *IIT* is used as specialisation measure.[22] It follows that the relationship between the correlation of business cycles and bilateral trade is fairly robust across deciles. The estimates for each conditional decile are almost always significant at the 5% significance level. Moreover, the figures show that the quantile regression estimates are very similar to the OLS estimates and almost always lie within the 95% confidence band of the OLS estimates. This indicates that the relationship between business cycle correlations and bilateral trade does not differ across the sample.

[22] For brevity, only the estimates across deciles for bilateral trade are shown. Full results are available upon request.

Figure 7a. Quantile Regression Plot, GDP model

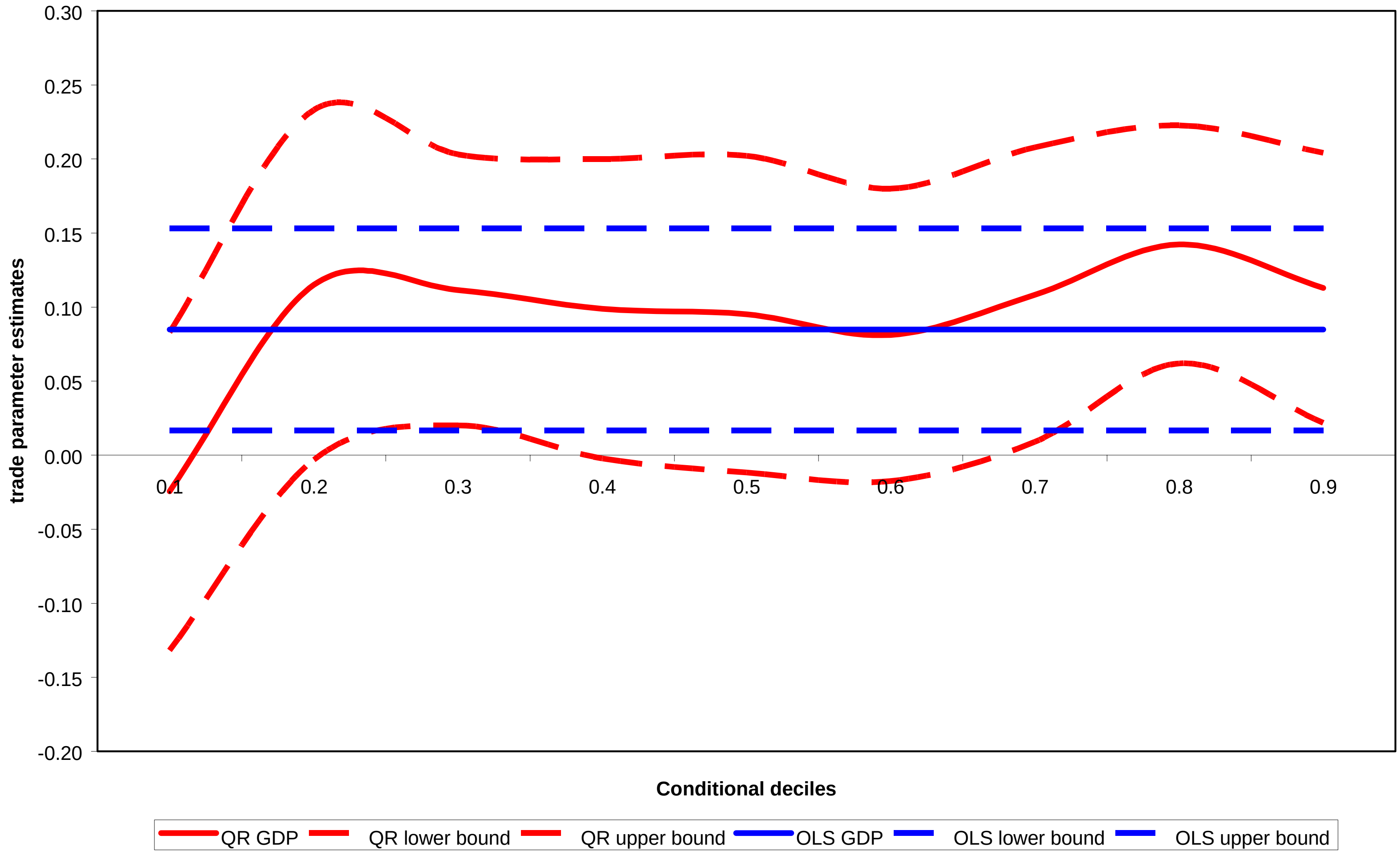
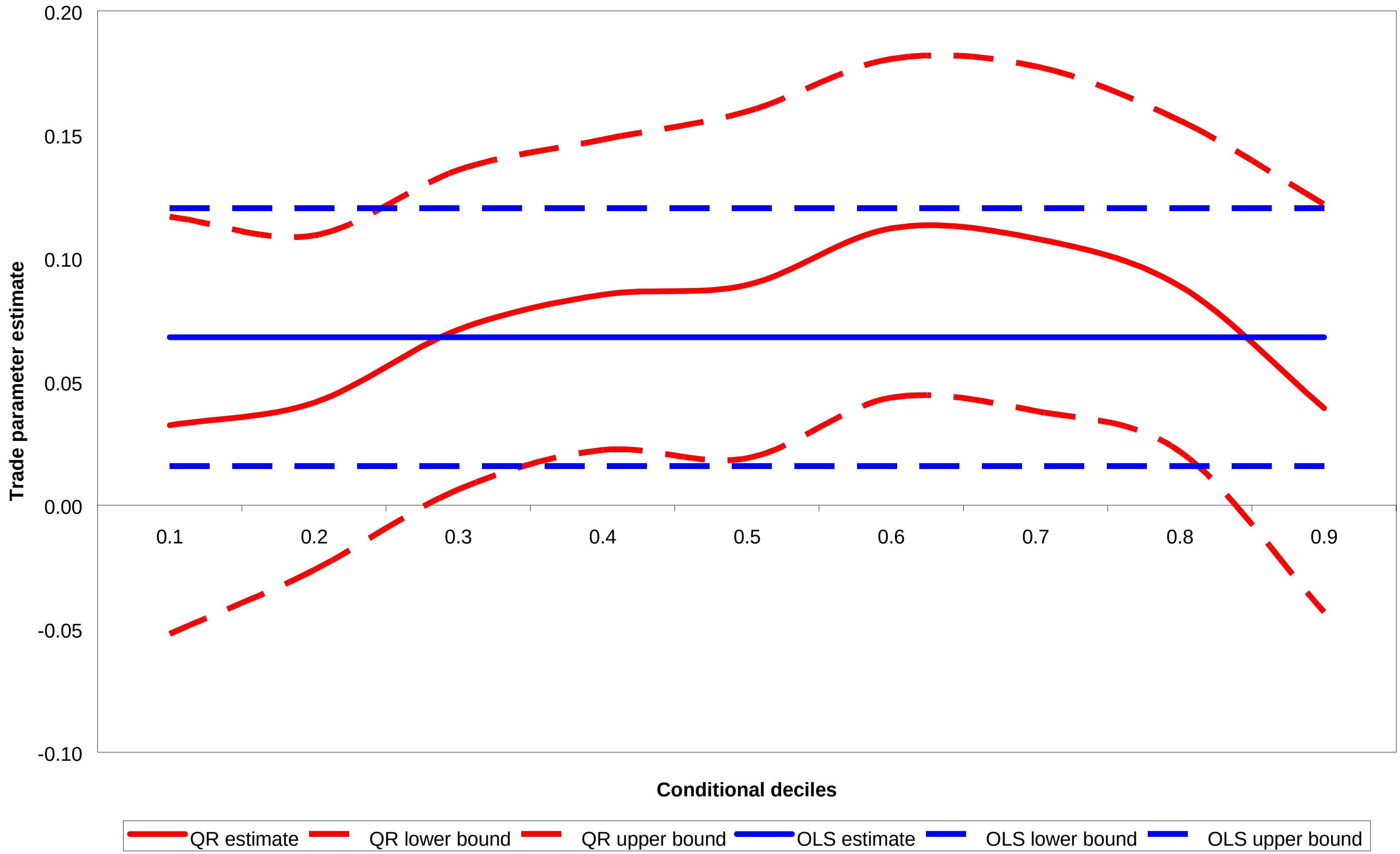


Figure 7b. Quantile Regression Plot, IIP model



6. Concluding comments

We have re-examined the relationship between trade intensity and business cycle synchronization for a sample of 21 OECD countries over the period 1970-2003, using the bilateral correlation of detrended real economic activity (GDP and industrial production) as dependent variable. Since a correlation coefficient lies between –1 and 1, the error terms in a regression model of the determinants of business cycle synchronization are unlikely to be normally distributed. We therefore employ transformed correlation coefficients as the dependent variable in our regression models. Including variables capturing similarity of monetary and fiscal policies, financial integration, and specialisation in a multivariate model, instead of using instrumental variables estimation, we confirm the finding that trade intensity affects business cycle synchronization, but the effect is much smaller than previously reported. Furthermore, the other factors included in the model have at least as strong an effect on business cycle synchronization as trade intensity. Finally, our results suggest that the effect of trade on business cycle synchronization does not suffer from sample heterogeneity and is robust for outlying observations.

Our findings are good news for supporters of the Economic and Monetary Union (EMU) in Europe. A common monetary policy will be easier to implement if the member countries' business cycles are aligned. If various countries in the monetary union are not at the same points in the business cycle, decision-making on the appropriate monetary policy stance becomes a difficult task.[23] However, our results suggest that the well-known critique on EMU that a common monetary policy may not be equally good for all countries in the union ("one size does not fit all"), has lost force due to the economic and monetary integration process. Not only more trade and especially more intra-industry trade – which has increased substantially over time in the EMU countries – leads to business cycles that are more in sync, also similar economic policies lead to more business cycle synchronization. These findings lend support to Trichet's claim that "we can be reasonably confident in the increasing integration of European countries, and in the fact that economic developments are becoming more and more correlated in the area. This has been highlighted, in the academic field, by several empirical investigations …… [that] found

[23] However, as pointed out by Kalemli-Ozcan *et al.* (2001), insurance possibilities against idiosyncratic shocks could increase aggregate utility and the more so with asynchronous business cycles.

evidence that business cycles are becoming more synchronous across Europe” (Trichet, 2001, pp. 5-6).

References

Artis, M. (2003), Is there a European Business Cycle?, CESifo Working Paper, No. 1053.

Artis, M. J. and Zhang, W. (1997), International Business Cycles and the ERM, International Journal of Finance and Economics, 2 (1), 1-16.

Baxter, M. and King, R.G. (1999), Measuring Business Cycles: Approximate Bandpass Filters for Economic Time Series, Review of Economics and Statistics, 81(4), 575-593.

Baxter, M. and Kouparitsas, M. (2004), Determinants of Business Cycle Comovement: A Robust Analysis, NBER Working Paper, No. 10725.

Beck, T., Demirgüç-Kunt, A. and Levine, R. (1999), A New Database on Financial Development and Structure, World Bank Working Paper, No. 2146.

Bergman, U.M. (2004), How similar are European business cycles? Lund University, mimeo.

Calderón, C., Chong, A. and Stein, E. (2002), Trade intensity and business cycle synchronization: Are developing countries any different? Central Bank of Chile Working Paper, No. 195.

Camacho, M., Perez-Quiros, G. and Saiz, L. (2005), Are European Business Cycles Close Enough to Be Just One? CEPR Discussion Paper, No. 4824.

Clark, T.E. and van Wincoop, E. (2001), Borders and Business Cycles, Journal of International Economics, 55, 59-85.

David, F. N. (1949), The Moments of the z and F Distributions, Biometrika 36, 394-403.

De Grauwe, P. and F. P. Mongelli (2005), Endogeneities of optimum currency areas: what brings countries sharing a single currency closer together?, ECB Working Paper, No. 468.

De Haan, J., Eijffinger, S.C.W., Waller, S., (2005), The European Central Bank Credibility, Transparency, and Centralization, MIT Press, Cambridge (MA).

Frankel, J.A. (2004), Real convergence and euro adoption in Central and Eastern Europe: Trade and business cycle correlations as endogenous criteria for joining EMU, John F. Kennedy School of Government Working Paper, No. RWP04-039.

Frankel, J.A. and Rose, A.K. (1998), The Endogeneity of the Optimum Currency Area Criteria, Economic Journal, 108, 1009-1025.

Feenstra, R.C., Lipsey, R.E., Deng, H., Ma, A. C. and Mo, H. (2005), World Trade Flows: 1962-2000, NBER Working Paper, No. 11040.

Grubel, H. G. and Lloyd, P. J. (1971), The Empirical Measurement of Intra-Industry Trade, Economic Record, 47 (120), 494-517.

Gruben, W.C., Koo, J. and Millis, E. (2002), How Much Does International Trade Affect Business Cycle Synchronization?, Federal Reserve Bank of Dallas, Working Paper, No. 0203.

Hausman, J.A. (1978), Specification tests in Econometrics, Econometrica, 46, 1251-71

Imbs, J. (2004), Trade, Finance, Specialization and Synchronization, Review of Economics and Statistics, 86, 723-734.

Kalemli-Ozcan, S., B.E. Sørensen and O. Yosha (2001), Economic integration, industrial specialization, and the asymmetry of macroeconomic fluctuations, Journal of International Economics, 55, 107-137.

Koenker, R. and Bassett, G. (1978), Regression quantiles, Econometrica, 46, 33–50.

Koenker, R. and Hallock, K.F. (2001), Quantile Regression, Journal of Economic Perspectives, 15(4), 143-56.

Lane, P. R. and Milesi-Ferretti, G. M. (2001), The External Wealth of Nations: Measures of Foreign Assets and Liabilities in Industrial and Developing Countries Journal of International Economics 55(2), 263-294.

Leamer, E.E. (1983), Let's take the Con Out of Econometrics, American Economic Review, 73(3), 31-43.

Levine, R. and Renelt, D. (1992), A sensitivity analysis of cross-country growth regressions, American Economic Review, 82, 942-963.

Massmann, M. and Mitchell, J. (2004), Reconsidering the evidence: are Eurozone business cycles converging? Journal of Business Cycle Measurement and Analysis, 1(3), 275-308.

OECD (2002) Purchasing Power Parities and Real Expenditures, 1999 Benchmark year, OECD: Paris.

Otto, G., Voss, G. and Willard, L. (2001), Understanding OECD output correlations, Reserve Bank of Australia Research Discussion Paper, No. 2001-5.

Rose, A. and C. Engel (2002), Currency unions and international integration, Journal of Money, Credit, and Banking, 34, 1067-1089.

Rousseeuw, P. J. (1984), Least Median of Squares Regression, Journal of the American Statistical Association, 79, 871-880.

Rousseeuw, P. J. (1985), Multivariate Estimation with High Breakdown Point, in Grossmann, W., G. Pflug, I. Vincze and W. Wertz (eds), Mathematical Statistics and Applications, Dordrecht: D. Reidel Publishing Co.

Sala-i-Martin, X., (1997), I just ran two millions regressions, American Economic Review 87(2), 178-183.

Sturm, J.E. and De Haan, J. (2005), Determinants of long-term growth: New results applying robust estimation and Extreme Bounds Analysis, Empirical Economics, 30, 1-22.

Trichet, J-C. (2001), The Euro after Two Years, Journal of Common Market Studies, 39 (1), 1-13.

Appendix. The EBA used to select the variables used in the structural model

Variable:	Source:	Suggested by:	Robust in model for:	
			GDP correlation	IP correlation
Short-term interest rate correlation	IMF, International Financial Statistics (IFS)	Otto et al. (2001)	Yes	Yes
Cyclically-adjusted budget deficits correlation	OECD Economic Outlook (vol. 76)	Camacho et al. (2005)	Yes	Yes
Correlation of changes in the stock market	IFS	Otto et al. (2001)	Yes	Yes
Capital account restrictions	Milesi-Feretti and IMF	Imbs (2004)	No	No
Difference (absolute) in Net foreign asset positions	Milesi-Feretti and IMF	Imbs (2004)	No	No
Share of intra-industry trade (IIT)	Feenstra *et al.* (2005)		Yes	Yes
Industrial similarity	GGDC 60-industry database	Imbs (2004)	No	Yes
Export similarity	Feenstra *et al.* (2005)	Baxter and Kouparitsas (2004)	Yes	Yes
Exchange rate variability	IFS	Otto et al. (2001)	Yes	No
Average openness	IFS & GGDC Total Economy Database	Baxter and Kouparitsas (2004)	No	No
Import similarity	Feenstra *et al.* (2005)	Baxter and Kouparitsas (2004)	No	No
Human capital difference	OECD Labour Force Statistics	Baxter and Kouparitsas (2004)	No	No
Physical capital difference	GGDC Total Economy Growth Accounting Database	Baxter and Kouparitsas (2004)	No	No
EMS-dummy		Frankel and Rose (1998)	No	No
Average oil import share	World Bank, World Development Indicators (WDI)	Artis (2003)	No	No
Correlation of inflation rates	IFS	Camacho et al. (2005)	No	No
Variability in inflation rate difference	IFS	Camacho et al. (2005)	No	No
Current account restrictions	Milesi-Feretti and IMF	Imbs (2004)	No	No
Human capital (tertiairy education)	OECD Labour Force Statistics	Baxter and Kouparitsas (2004)	No	No
Arable land difference	WDI	Baxter and Kouparitsas (2004)	No	No
Relative labour productivity level	GGDC Total Economy Database	Baxter and Kouparitsas (2004)	No	No

Relative financial structure (credit/stock)	Beck *et al.* (1999)	Artis (2003)	No	No
Difference in national savings ratio	OECD National Accounts	Camacho et al. (2005)	No	No

Notes: A more detailed description of the variables and sources, as well as the data is available at www.rug.nl/economics/inklaarrc

The Extreme Bounds Analysis (EBA) as suggested by Leamer (1983) and Levine and Renelt (1992) is used to determine the list of variables to be included in the structural model outlined in the main text. The EBA has been widely used in the economic growth literature (see Sturm and De Haan (2005) for a further discussion). Baxter and Kouparitsas (2004) also use this methodology (using a different set of countries and a more limited number of possible explanatory variables than in the present paper) to examine which variables are robustly related to business cycle synchronization. The EBA can be exemplified as follows. Equations of the following general form are estimated:

$$Y = \alpha M + \beta F + \gamma Z + u \qquad (A1)$$

where Y is the dependent variable (output correlation); M is a vector of 'standard' explanatory variables; F is the variable of interest; Z is a vector of up to three (here we follow Levine and Renelt (1992)) possible additional explanatory variables, which according to the literature may be related to the dependent variable; and u is an error term. In our analysis only trade intensity is included in the M vector. As explained in the main text, the various proxies for financial integration and specialisation are not considered simultaneously. Following Sala-i-Martin (1997), we use the unweighted cumulative distribution function (CDF(0)), i.e. the fraction of the cumulative distribution function lying on one side of zero, and the percentage of the regressions in which the coefficient of the variable of interest differs significantly from zero. Following Sturm and De Haan (2005), a variable is considered to be robust if the CDF(0) test statistic > 0.95 and if the variable has a significant coefficient (on the 5% significance level) in 90% of all regressions ran.

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