ENDOGENEITY ANALYSIS OF OUTPUT SYNCHRONIZATION IN THE CURRENT AND PROSPECTIVE EMU

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Abstract

The sustainability of European economic and monetary union (EMU) remains an important issue in light of existing plans for enlargement. This paper conducts an endogeneity analysis of output synchronization, based on panel data estimation from 1994-2013, for different country-groups, including core, periphery, central and eastern European countries, northern European and the prospective candidate countries, which are expected to adopt the euro over the coming years. The quantification of trade-related and direct spillover channels associated with monetary integration provides insight into the relative importance of direct and indirect synchronization gains arising from EMU membership. The use of amplitude and concordance measures of synchronization and a range of estimators enhances robustness. There are important endogeneity implications which emerge from our analysis.

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I. Introduction

Various studies conclude that growing business cycle asymmetries and imbalances arising since the euro's introduction may make in the future more likely the sort of financial crises experienced recently (Schmitz and von Hagen, 2011; Vieira and Vieira, 2012; Chen et al., 2013; Caporale et al., 2015). Given that the recent euro crises involved periphery countries, the prospective membership of the euro area has important implications for future regional stability and beyond.[1] This paper provides a reassessment of its optimality beyond the current crisis in view of the substantial changes to its membership expected over the coming years. The contribution focuses on one prominent criterion, namely the business cycle synchronization.

If one considers the 2004, 2007 and 2013 EU acceding countries of Czech Republic, Hungary, Poland (2004), Bulgaria, Romania (2007) and Croatia (2013), which are all obliged to adopt the euro after satisfying the Maastricht criteria for convergence, this would involve a substantial increase on the current membership of 19 countries. This paper goes beyond previous literature to provide a contemporary analysis of synchronization for the prospective enlarged EMU, which we refer to as EEMU.[2]

This paper considers EEMU both at an aggregate level and analyzes the trends emerging for different country-groups, including core, periphery, central and eastern European countries (CEECs), Northern European and the prospective candidate countries, which are expected to adopt the euro over the coming years. Using a range of estimators and time-varying synchronization measures, including growth rates, cyclical deviations and concordance, this paper sheds light on whether trade integration brings about economic shocks, which fundamentally affect the coincidence of signs or amplitudes of business cycles. Our analysis also enables quantification of

[1] The European economic and monetary union, also known as the euro area, is referred to as EMU. The terms are used interchangeably. As of 2015 this region includes: Austria, Belgium, Germany, France, Italy, Spain, Portugal, Malta, Cyprus, Greece, Netherlands, Ireland, Latvia, Luxembourg, Finland, Slovenia, Slovakia, Estonia and Lithuania.
[2] EEMU includes, in addition to the aforementioned EMU countries, the following candidates: Bulgaria, Croatia, Czech Republic, Hungary, Poland and Romania. Collectively EMU member states are referred to as EMUs and prospective members of EEMU as EEMUs.

both trade-related effects and direct spillovers associated with monetary integration, thus providing insight into the relative importance of direct and indirect synchronization arising from EMU membership channels.

The empirical strategy is based on panel estimation of a sample of 32 countries, which includes the current EMU member states, the prospective candidate countries, as well as a group of additional control countries. Various estimators are used to address endogeneity and sample heterogeneity over 1994-2013, which includes key periods of European integration, such as the euro's introduction and transition of candidate countries to the EU.

The paper is structured as follows. Section II reviews the relevant literature. Section III introduces the empirical methodology and describes the data. Section IV presents and discusses the empirical analysis. Section V summarizes and concludes.

II. Literature Review

Optimal currency area (OCA) theory asserts that countries should consider carefully the costs and benefits of joining a currency area (Mundell, 1961). To the extent that it conveys information about the symmetry and transmission of shocks, business cycle synchronization remains a key indicator of the costliness of a common monetary policy (De Grauwe, 2012).

An important aspect within the OCA literature is the endogeneity of the cost-benefit analysis. According to the *Endogeneity Hypothesis* (Frankel and Rose, 1997, 1998), countries joining a currency area may experience dramatically different business cycles, which may reflect a shift in exchange rate regime and closer international trade with other member states. Positive early evidence adds weight to this claim (Rose, 2008; Flood and Rose, 2010).

The increase in time-series data availability and advances in measurement have increased the scope for empirical research using time-varying measures of synchronicity, including most popularly those based on amplitude (Giannone and Reichlin, 2010; Kalemli-Ozcan et al., 2013a, 2013b;

Caporale et al., 2015). Mink et al. (2012) note, for instance, that perfect correlation of cycles does not mean that the common monetary policy suits all countries equally well if amplitude of cycles differs (see Figure 3, p. 222). However, even if the absolute amplitude differential is small, countries may be in different phases of the business cycle, requiring different policy prescriptions. This distinction between amplitude and concordance seems relevant to both academics and policymakers (De Haan et al., 2007).

Over recent years studies have highlighted the growing divergence of business cycles between core and periphery EMU countries (Schmitz and Von Hagen, 2011; Vieira and Vieira, 2012; Chen et al., 2013; Caporale et al., 2015). The instabilities arising appear to stem partly from a lack of trade gains realized since the euro's introduction, especially for Southern European countries, whose exports have been displaced by more competitive producers, adding to trade and financial imbalances (Chen et al., 2013).

Trade's inability to synchronize core-periphery business cycles has also been highlighted, along with an apparent reinforcement of a 'core' cycle around the euro's introduction (Caporale et al., 2015). Moreover, the trade-synchronization relationship is theoretically ambiguous, since it depends on various factors, including economic structure, the type of trade (e.g. intra versus inter-industry), mixture of shocks and responses of agents to shocks (Krugman, 1993; Fidrmuc, 2004; Kose and Yi, 2006). Since these factors, and others, may differ across EEMU, the impact of trade integration on synchronization may differ as well.[3]

If the *Endogeneity Hypothesis* is favourable only for certain groups of countries and not others, those with inadequate adjustment mechanisms may be pre-exposed to additional risks of rapid deterioration of fiscal and financial balance following large negative macroeconomic and financial shocks (Vieira and Vieira, 2012).

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[3] Luintel et al. (2008) and Arestis et al. (2010) discuss the issue of policy relevance of empirical work in the context of the financial liberalization and economic growth literature. Heterogeneity in the trade-synchronization association raises questions about the policy relevance of existing literature in this area.

In view of the highly uncertain outlook for the prospective candidate countries, this paper goes beyond previous literature to provide a contemporary analysis of synchronization for the prospective enlarged EMU.

III. Empirical Methodology

The standard econometric model links business cycle synchronization ($S_{ij,t}$) of country-pair ij, to the natural logarithm of bilateral trade intensity ($T_{ij,t}$), conditional on any control variables (represented by $Z_{ij,t}$), in period t. The association for trade intensity is captured in the slope coefficient φ in equation (1) below:

$$S_{ij,t} = \alpha + \varphi T_{ij,t} + \gamma Z_{ij,t} + \varepsilon_{ij,t} \tag{1}$$

Our panel data econometric approach extends over the basic setup in equation (1) and draws on Artis and Okubo (2011), Mink et al. (2012), Kalemli-Ozcan et al. (2013b) and Caporale et al. (2015), as discussed below.

The first model, which provides an overall view of the empirical association for EEMUs, is dynamically specified with a one period lag of explanatory variables to address possible endogeneity, an issue discussed further below. An EMU membership dummy variable ($EMU_{ij,t}$) is also included, to capture any direct membership effects arising from monetary integration through coefficient β (Flood and Rose, 2010). Cross-sectional and time fixed effects, μ_{ij} and τ_t respectively, are included to account for spatial and dynamic heterogeneity:

$$S_{ij,t} = \alpha + \mu_{ij} + \tau_t + \beta EMU_{ij,t} + \varphi T_{ij,t-1} + \gamma Z_{ij,t-1} + \varepsilon_{ij,t} \tag{2}$$

The first model is extended for a comparative analysis of dynamics for specified country-groups, thus providing a more flexible treatment of trade intensity's impact on synchronization. The second model includes a binary indicator variable, $\lambda_{K,t}$, which takes a value of zero or unity; with K as the group indicator and two interaction variables $\lambda_{K,t}T_{ij,t-1}$ and $(1-\lambda_{K,t})T_{ij,t-1}$:

$$S_{ij,t} = \alpha + \mu_{ij} + \tau_t + \beta EMU_{ij,t} + \varphi_1 \lambda_{K,t} T_{ij,t-1} + \varphi_2 (1 - \lambda_{K,t}) T_{ij,t-1} + \varphi_3 \lambda_{K,t} + \gamma Z_{ij,t-1} + \varepsilon_{ij,t} \tag{3}$$

Coefficient φ_1 is the trade intensity slope coefficient for all country-pairs not identified in the Kth group (i.e. the rest of the world); φ_2 is the trade intensity slope coefficient for the Kth group; φ_3 is the marginal intercept effect for the Kth group.

Endogeneity

It is worth briefly revisiting the endogeneity problem. First, it is possible that synchronization, via its impact on economic policy, affects the set of explanatory variables. Second, there may be unobserved omitted variables, such as institutional and structural characteristics, and common global shocks, that are related both to synchronization and the explanatory variables across the sample. Third, measurement error may affect key variables and could in principle generate endogeneity bias.

To address possible endogeneity, we take several steps. First, following Artis and Okubo (2011) and Kalemli-Ozcan et al. (2013b), specifications are considered as set out in equations (2) and (3). The dynamic specifications, which involve a single lag of exogenous regressors, exploit the low degree of persistence of synchronization measures, which may at least partially address endogeneity in the form of reverse causation.

Second, to account for omitted variables bias arising through unobserved heterogeneity, cross-sectional and time fixed effects are included. We consider both country and country-pair specific effects and include yearly dummy variables to account for dynamic heterogeneity, which helps control for hard-to-measure variables, such as political coordination, cultural factors, informational frictions, and common global shocks (Kalemli-Ozcan et al., 2013b).

Third, a two-step fixed effects efficient Generalized Method of Moments (GMM) estimator is employed, which minimizes the GMM objective function J = Ng'Wg, where N is the sample size; g

are the orthogonality conditions, which specify that exogenous variables are uncorrelated with the error term in the model; W is a weighting matrix (see, e.g., Hayashi, 2000 pp. 200-24). The instrument set includes one lag of endogenous regressors (trade intensity and financial integration), as well as dummy variables for EU and EMU membership, and an index of legal protection (strength of legal rights), taken from the *World Bank's WDI* database.[4,5]

Fourth, to the extent that outliers may reflect omitted variables and measurement errors, a robust regression is also conducted using the iteratively re-weighted least squares M-estimator due to Huber (1964), whereby higher weights are assigned to better behaved observations. This is potentially useful here if sample heterogeneity within and across country-groups results in influential outlier observations. The robust regression is dynamically specified, with control variables and fixed effects.

Sample Selection, Measurement and Definition of Variables

The sample reflects membership of EEMU: Bulgaria, Croatia, Czech Republic, Hungary, Poland, Romania, Austria, Belgium, Cyprus, Estonia, Finland, France, Germany, Greece, Ireland, Italy, Latvia, Netherlands, Portugal, Slovakia, Slovenia, Lithuania and Spain. Luxembourg and Malta are excluded due to data limitations. Also included are the following countries: Canada, Denmark, Japan, Norway, Sweden, Switzerland, Turkey, the United Kingdom and United States. Most data are available from the early 1990s, thus a feasible sample for analysis is 1994-2013. Quality of the available data is another important factor motivating this timeframe (see, e.g. Fidrmuc and Korhonen, 2006).

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[4] Time-invariant variables are perfectly collinear under fixed effects; this limits use of traditional Gravity variables, such as bilateral distance, whether countries are land locked and share a common language.
[5] The p-values from J-tests largely indicate that instruments are valid at conventional significance levels. Instrument weakness testing indicates that endogenous variables are identified through the instrument set.

Synchronization

Consideration of both amplitude and concordance aspects of synchronization adds robustness to our approach. Following Kalemli-Ozcan et al. (2013a, 2013b), the amplitude measure, $S1_{ij,t}$, is based on the negative of the absolute growth rate differential of country-pair ij. The secondary, binary concordance measure, $S2_{ij,t}$, is also considered for panel estimation, which takes a value of 1 if both countries experience an expansion/contraction relative to trend:

$$S1_{ij,t} = -|\Delta \ln y_{i,t} - \Delta \ln y_{j,t}|$$

$$S2_{ij,t} = \{0,1\}$$

Amplitude is formed from output growth rates of country-pair ij, where $\Delta ln\ y_t$ is the first-difference in the natural logarithm of output for period t. For concordance, $\Delta ln\ y_t$ is the HP-filtered cyclical deviation of real GDP from trend, which is preferred to concordance of growth rates because countries are most of the time in an expansionary phase of the business cycle. Both synchronization measures are easily interpretable; a smaller amplitude differential implies a smaller absolute difference in growth rates, whereas larger values for concordance indicate that cyclical phases coincide to a greater extent.

Data for synchronization are based on real GDP series denoted in US dollars and obtained from *IMF's WEO* database.[6]

Trade Intensity

Following the literature bilateral trade intensity ($T_{ij,t}$) is defined as the natural logarithm of the sum of exports and imports between trade partners ij ($f_{ij,t}$) normalized by the sum of $z_{i,t}$ and $z_{j,t}$ in period t, the nominal GDP of partners i and j respectively:

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[6] IP data are not available for all years in the sample, especially for several non-core countries, which are central to this study.

$$T_{ij,t} = \ln\left(\frac{f_{ij,t}}{z_{i,t} + z_{j,t}}\right)$$

Nominal GDP data are obtained from *International Monetary Fund's (IMF) World Economic Outlook (WEO)* database. Bilateral trade data are obtained from *IMF's Directory of Trade Statistics (DOTS)* database. Both nominal GDP and trade flows are denoted in US dollars.

Financial Integration

To proxy for financial integration, a price-based indicator is adopted as proposed by the European Central Bank (2013). Financial integration is measured using the absolute long-term real interest rate differential on 10-year maturity sovereign bonds, scaled in percentage points:

$$FIN_{ij,t} = |\, r_{i,t} - r_{j,t} \,|$$

Data is sourced from the European Commission's *AMECO* database.

Sectoral Similarity

An economic specialization measure is based on the summation of squared sectoral deviations in gross value added for a given country-pair, following Baxter and Kouparitsas (2005):

$$SPEC_{ij,t} = 1 - \sum_{n=1}^{N}(s_{in,t} - s_{jn,t})^2$$

Sector-n weights in total output in period t for countries i and j respectively, $s_{in,t}$ and $s_{jn,t}$, are based on the proportion of gross value added in total gross value added in industry, agriculture and service sectors. Data is sourced from the *World Bank's World Development Indicators (WDI)* database.

Population Differential

Among the Gravity variables suitable for fixed effects panel estimation is the absolute population

differential of countries *i* and *j*, which may be an important exogenous influence on synchronization

(Baxter and Kouparitsas, 2005):

$$POP_{ij,t} = \mid pop_{i,t} - pop_{j,t} \mid$$

Data on population is taken from the *World Bank's WDI* database

Additional Controls

As well as the dummy indicating joint EMU membership, all models include a dummy variable for

the global crisis years, *CRISIS$_{ij,t}$*, which takes a value of unity for 2008-2011 and zero otherwise.

Data Description and Summary Statistics

Table 1 presents a statistical summary of synchronization variables for all country-pairs in EMU,

EEMU and the full sample, as well as for specific country-groups. Moreover, given their economic

influence on aggregate conditions in the monetary union, this paper pays particular attention to

synchronization of different groups with core countries.[7]

[TABLE 1]

Table 1 indicates that EEMUs exhibit on average comparable synchronization to current EMUs and

are significantly more synchronized than country-pairs in the rest of the world. Considerable

differences also exist across core-non-core country-pair groups and variability of synchronization

across EEMU is somewhat higher than in EMU. Core countries are most synchronized, whereas

CEECs, Northern countries, and candidates are less synchronized than peripherals. CEECs are more

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[7] Classification of countries into core and periphery is by no means straightforward, especially for Italy and Spain (see, e.g. Aguiar-Conraria et al., 2013). Further consideration of this issue is given in the Sensitivity Analysis.

synchronized on average than Northern countries in terms of amplitude, but not concordance, which may reflect in part numerous small Northern countries operating under relatively fixed exchange rate regimes.

[FIGURES 1-2]

Differences in amplitudes for EEMUs have for several years become significantly smaller around the euro's introduction compared with the rest of the world, although this has not been sustained (see Figure 1). Concordance of EEMUs appears to increase somewhat and temporally since the euro's creation and physical introduction in 1999 and 2002 respectively (see Figure 2). However, by 2013 synchronization of EEMUs is no different to most other countries and does not seem much larger than in the earlier years of the sample.

IV. Empirical Analysis

This section sets out and discusses the empirical results.

Model 1: Aggregated Results for EEMU

Table 2 summarizes the trade intensity estimates obtained under the first model using the amplitude measure, which provides an overall indication for EEMUs across the sample period. Various estimators are considered, including the dynamically specified fixed effects estimator (FE), instrumental variables (GMM) and the robust M-estimator due to Huber (1964). For comparison we estimate a pooled ordinary least squares model (POLS). Standard errors are corrected for heteroscedasticty and serial correlation by clustering for individual country-pairs following the approach of Imbs (2010), which is feasible given the large number of country-pairs relative to years in the sample.

[TABLE 2]

The estimated trade intensity slopes appear favourable and in line with Rose's (2008) findings, with positive signs and statistical significance, regardless of the estimation method; coefficients are all significant with a range of 0.26-0.33 for the full EEMU sample estimations. Moreover, there is no indication of an unfavourable specialization channel along the lines of the Krugman paradigm (Krugman, 1993).[8] There is more evidence here for specialization across industries rather than countries; in this case industry-specific shocks may occur alongside common demand shocks, which propagate across vertically integrated and specialized production networks.

There is also evidence of an evolving direct EMU membership effect on synchronization, which is robust to the use of different estimators (see Figures 3-6). Estimates are computed over 5-year rolling windows, estimating for the first full window associated with EMU membership 1999-2003, then for 2000-04, and so on, until the final window of 2009-13. The EMU membership point estimates evolve from near zero in 1999-2003 to approximately 1.4 in the pre-global crisis years. This suggests that members have become increasingly exposed to common monetary and trade-related spillovers since the euro's introduction, extending in an EMU context the finding of Flood and Rose (2010). Therefore, EMU membership has an important direct effect, which goes above and beyond the indirect effect arising through trade.

[FIGURES 3-6]

There are several indicators of endogeneity bias from estimation of the first model. First, fixed effects are always significantly different from zero, according to the outcomes of joint F-tests reported in Table 2. This points to important unobservable factors and common shocks, which if omitted may be correlated with the dependent variable and included regressors, generating endogeneity bias. Second, Granger causality tests provide some evidence of reverse causation between synchronization and trade intensity, although the causal ordering is from trade intensity to

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[8] See also Figures 1-4 in the online Appendix for evidence of temporal evolution of the empirical association. These figures reflect the estimates reported in Table 1 of the online Appendix. There is no evidence here of an unfavourable trade-synchronization relation developing.

synchronization as the null hypothesis is rejected in each case more strongly when trade is the lagged regressor.[9] Third, Hausman's (1978) test suggests that FE and GMM estimates sometimes do differ to the extent that endogeneity may be severe in certain periods.[10] Therefore, inclusion of fixed effects is not always sufficient for overcoming endogeneity bias. GMM estimates may be preferred in this light because the corresponding estimator provides a more general solution to the endogeneity problem, although the full sample results are not very sensitive to the choice of the estimator.

Robustness of Different Synchronization Measures

To explore further for robustness, we conduct three sets of estimations. First, by averaging variables over the sample, we carry out cross-sectional estimation. Second, country fixed effects cross-sectional estimation is conducted for individual years and then aggregated in the spirit of Caporale et al. (2015); aggregated point estimates are computed by averaging yearly point estimates and correcting for serial correlation, following Chakravarti et al. (2004), by multiplying the sample average standard error by $\{[1+\Phi]/[1-\Phi]\}^{1/2}$, where Φ is the autocorrelation of the relevant slope coefficient with its first lag. Third, we conduct a country-pair and time fixed effects GMM panel estimation. Results are reported in Table 3 for the amplitude measure of synchronization (growth rates), an amplitude measure based on the HP-filtered growth rate differential, and for the binary concordance measure, using panel probit and logit estimation, with one lag of explanatory variables and both types of fixed effects to help address endogeneity.

[TABLE 3]

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[9] Granger causality tests operate under the null hypothesis of no conditional correlation between trade and synchronization, based on one lag of explanatory variables.

[10] Table 1 in the online Appendix presents outcomes for Hausman's (1978) test and joint fixed effects tests across 5-year rolling windows. Hausman's (1978) test follows a Chi-square distribution and is based on the ratio of the squared difference of FE and GMM point estimates and their estimated variances. The test operates under the null hypothesis that FE and GMM estimates do not differ significantly.

Trade intensity estimates are all significant under the preferred specifications, which include a full set of controls and robust standard errors, with a range of 0.07-0.3 (amplitude) and 0.11-0.18 (concordance). The HP-filtered measure of amplitude produces estimates which are somewhat lower than under the unfiltered amplitude measure, but estimates remain statistically significant. The direct EMU membership effect is larger with a range of 0.2-0.87 (amplitude) and 0.15-0.27 (concordance); this effect under the HP-filtered amplitude measure is again smaller than under the unfiltered amplitude measure, but it also remains statistically significant.

The EMU membership effect is robust to the inclusion of a dummy variable indicating whether country-pairs operate under a fixed exchange rate regime, using the 1-11 scale of the classification of Ilzetzki et al. (2010), to control for announced, unannounced, direct and indirect exchange rate fixing.[11] In some cases the membership coefficient is smaller and less significant, but not always. Therefore, the favourable spillover effects associated with EMU may reflect common monetary spillovers rather than a fixing of exchange rates to the euro currency.

The economic interpretation of aggregate estimates under the amplitude measure using GMM estimation and the standard set of control variables is as follows. A doubling of trade intensity reduces the (unfiltered) growth rate differential of EEMUs by approximately 0.21 per cent, due to the scaling of amplitude in percentage points, with a 90 per cent confidence interval of 0.17-0.25 per cent. The direct membership effect of adopting the euro reduces the (unfiltered) growth rate differential by 0.6 per cent, with a 90 per cent confidence interval of 0.41-0.8 per cent. Therefore, the total membership effect (direct and indirect) reduces the (unfiltered) growth rate differential by 0.81 per cent, with a 90 per cent confidence interval of 0.58-1.05 per cent. For comparison, using the HP-filtered amplitude measure, the corresponding confidence intervals for indirect, direct and total membership effects are 0.04-0.07, 0.13-0.27 and 0.17-0.34 per cent respectively, which are smaller than those obtained under the unfiltered amplitude measure.

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[11] Upon the inclusion of the fixed exchange rate dummy, the point estimates and significance of the EMU membership dummy are very similar to what is indicated in Figures 3-6.

For the concordance measure, logit estimates are used to analyze the change in the predicted probability that two countries are in the same cyclical phase by altering trade and EMU membership variables.[12] Holding the EMU membership dummy variable constant at unity and doubling trade from its sample mean increases the likelihood of phase alignment by 2.4 per cent. However, changing the EMU membership dummy from zero to unity and holding trade constant at its sample mean increases the likelihood of phase alignment by 5.1 per cent, which is more than double the trade effect. The total membership effect is inferred by changing the EMU membership dummy from zero to unity and doubling trade, which increases the likelihood of phase alignment by 7.5 per cent, with a 90 per cent confidence interval of 3.1-11.7 per cent.

Overall these findings suggest that trade does broadly help promote synchronization, through alignment of growth rates, business cycles and increasing the likelihood of two countries being in the same cyclical phase. However, the trade effects are quite small in absolute terms and in view of existing estimates for the euro's effect on trade of around 5-15 per cent (Baldwin, 2006). The indirect membership effects are also small relative to the direct effects, which seem economically more important.

Model 2: Country-Group Analysis

The second model enables a time-varying decomposition for particular groups, based again on rolling window estimations, computed over 5-year windows, using the GMM fixed effects estimator to address endogeneity.[13] Our preferred measure of synchronization is amplitude, which

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[12] Estimates are computed using the standard set of control variables. We use the average predicted probabilities calculated from the sample values of non-trade and non-EMU predictor variables. Estimated effects are very similar under both probit and logit estimations. All results are provided in Table 2 of the online Appendix for reasons of space.

[13] FE and Huber estimators produce very similar point estimates and levels of significance to the GMM results for all regressors.

exhibits more variation across different periods in the sample.[14] Table 4 presents a summary of the estimates during the pre-crisis years (1995-2007).[15]

[TABLE 4]

The EMU membership estimates (β) are usually significant following the euro's introduction, although estimates are somewhat smaller than in the EEMU sample. Trade intensity estimates for the rest of the world (φ_1) are positive and significant for most years across the sample and the estimates reported for non-trade controls are generally of the expected signs. The findings suggest that more financially integrated countries tend to have more similar growth rates. Sectoral similarity is not always significantly associated with synchronization, but when it is, point estimates are positive; it may be that the relative importance of sectoral versus non-sectoral shocks in influencing growth has changed across the sample. The negatively signed and typically significant estimates for the absolute population differential also provide evidence for a relative population effect; one explanation is that countries with different populations are exposed to rather different economic shocks.

The trade intensity estimates (φ_2) for core-core country-pairs are positive and generally highly significant around the euro's introduction, thus suggesting that the euro has reinforced a 'core' European cycle. Mostly, the slope for core-periphery country-pairs is negative and insignificant. Furthermore, relative to the rest of the world (measured as $\varphi_2 - \varphi_1$), there is an asymmetry between core and periphery estimates (see Figures 7-8). Estimates for core-core countries are often significantly larger than those corresponding to the rest of the world, whereas this is not the case for core-periphery countries. Trade may have reduced synchronization of periphery countries for

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[14] For certain years there is no variation in the binary concordance measure, which makes estimation problematic, particularly around the period corresponding to the global economic crisis.
[15] The rolling-window estimates for all country-groups over the period 1995-2013 are presented in full in Table 3 of the online Appendix for reasons of space.

several years following the euro's introduction, which points to the emergence of a trade-driven specialization channel.

It may be that there has been a shift in the balance of intra and inter-industry trade, with core countries becoming more integrated in the former and periphery countries in the latter around the period corresponding to the euro's introduction. According to Krugman's (1993) specialization paradigm, comparative advantage can lead to a negative association developing between trade and synchronization if specialization occurs across countries. This also seems consistent with findings in the literature, including notably Caporale et al. (2015). Therefore, periphery countries may have become more exposed to rather different industry shocks and isolated from common demand spillovers affecting those involved in European production networks.

[FIGURES 7-10]

Point estimates reported in Table 4 for CEECs, Northern countries and candidates are positive and significant through the early part of the sample and significantly exceed those for all other country-pairs (see Figures 9-10 for Northern and candidate countries). The rapid integration into the European marketplace may have resulted in an increased incidence of trade-related shocks in the build-up to EU accession, which is reflected in changes in the empirical association around the same period. This is suggestive of common spillover and intra-industry trade channels, associated with transition and EU integration (Fidrmuc, 2004).

The empirical association for candidates, many of which are CEECs, has however weakened considerably over time. Point estimates are often smaller than for all other country-groups and no different to the rest of the world (see Figure 10). It may be that other macro-financial/policy-related factors have become more important during the global financial crisis; however, the breakdown for candidates occurs far in advance of the turmoil as it is evident from Table 4. This breakdown is contrary to core and Northern groups, whose empirical associations remain positive and significant

before the global economic crisis and it is suggestive of an emerging specialization channel. Another explanation is that economic and financial institutions in core and Northern countries have become much more harmonized than in periphery, CEEC and candidate countries, leading to differences in the propagation of shocks across EEMU.

Sensitivity Analysis

Several additional checks are carried out. First, we use a measure based on total trade flows in the denominator instead of the sum of GDP. Second, we experiment with additional/alternative control variables. To control for the impact of asymmetric fiscal policies on synchronization, we adopt the measure of Artis et al. (2008), based on the absolute difference of cyclically adjusted fiscal balance, taking data from the *AMECO* database. For financial integration, we employ an alternative, quantity-based measure, constructed using data on total banking assets, liabilities, financial derivatives, equity and portfolio flows from the updated and extended version of the dataset of Lane and Milesi-Ferretti (2007), which yields instead a negative association between financial integration and synchronization (using a slightly smaller sample, which runs until 2011), corroborating previous findings (see, e.g. Kalemli-Ozcan et al., 2013a; Caporale et al., 2015). Estimates and significance for trade and the EMU membership effect are largely unaffected. These checks do not undermine our main conclusions, which are set out below.[16]

As an additional analysis, we re-estimate both models after excluding countries sequentially with replacement; the results suggest that the EEMU trade-synchronization relation is positive and robust at an aggregate level. The direct membership effect is significant and relatively large. The results for the second model, with exclusions of candidate and core countries, are summarized in the online Appendix for reasons of space.[17] The trade intensity slopes for core-candidates are significant and

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[16] All the results referred to in the text are available from the authors upon request.

[17] See Table 4 of the online Appendix.

relatively large up until 2004. However, post-2004 and 2007 EU enlargement rounds, the core-candidate slopes are always insignificant and relatively small.

Economic Implications

At this stage the outlook remains uncertain for the candidate countries, not least due to changes in the empirical association, which may arise from the process of European integration. We can, however, provide some early projections for candidate countries using the full range of point estimates along with standard errors reported in Table 4. In this light we construct 90, 95 and 99 per cent confidence intervals for the direct, indirect and total membership effects on amplitude, where trade gains of up to 20 per cent are considered.[18, 19]

The least optimistic scenario occurs at the lower limit of the 99 per cent confidence interval for membership effects; overall euro adoption increases the growth rate differential by 0.38 per cent under maximum trade gains, although without a direct membership effect the increase is smaller at 0.06 per cent. In the most optimistic scenario, which arises at the upper limit of the 99 per cent confidence interval, the maximum direct effect on amplitude is 1.03 per cent, whereas the range for the maximum indirect effect under 5-20 per cent trade expansions is 0.05-0.19 per cent. This implies a total membership effect, which ranges from 1.03 per cent when trade gains are absent to 1.22 per cent under a 20 per cent trade expansion.

As these results indicate it is possible that further European integration will not lead to an alignment of growth rates or worse. At first glance this may not be especially problematic; for example, on the basis of the EMU experience so far, even for the periphery group, the implications of a negative empirical association developing between trade and synchronization appear rather minor.[20] On this

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[18] The long-term trade gains remain uncertain, but we consider 20 per cent as the most optimistic scenario in view of existing evidence.

[19] The full output is provided in Table 5 of the online Appendix to conserve space.

[20] During 1998-2002 the trade point estimate for periphery countries of -0.145 (see Table 4) implies that a doubling of trade would reduce the growth rate differential by 0.1 per cent.

evidence, the upside from monetary integration seems relatively sizeable. One explanation for the asymmetry in trade effects is that there are numerous positive theoretical channels, which supersede or at least mitigate the negative effects arising from the specialization channel of Krugman (1993). In particular, Keynesian trade spillover channels, emphasized by Frankel and Rose (1998), also provide an outlet for asymmetric shocks, which may spill over to other countries, thus promoting cyclical alignment more broadly across EEMU.

The main finding from this exercise is that the direct membership effect on amplitude goes above and beyond the effect arising through trade; candidates may benefit more from economically important common spillovers associated with monetary integration and consequently may be significantly better positioned ex post. We note that there are other factors, which may be important for synchronization, not least financial integration. However, we leave this for future research as much less is known about the euro's effect on financial integration, while synchronization effects seem very sensitive to how financial integration is measured.

V. Summary and Conclusions

This paper provides a reassessment of EEMU's optimality in view of the substantial changes to its membership expected over the coming years. The outlook from aggregated evidence appears favourable; trade may bring about shocks, which fundamentally affect both the coincidence of business cycle phases and amplitudes. Furthermore, a direct EMU membership effect is identified, which seems larger than the indirect membership effect arising through trade. Member states may be more likely to satisfy the OCA conditions ex post, holding constant all other factors.

Analysis at a group level suggests that integration has helped to promote more broadly a European 'core' region. However, unlike for core and Northern countries, trade's effect on synchronization is found to be unstable for CEECs and candidates, and in that way is similar to periphery countries. Therefore, there is some evidence here for Krugman's (1993) specialization paradigm. By contrast

core and Northern countries have experienced favourable trade-related spillovers, in line with the *Endogeneity Hypothesis* and intra-industry trade channels, emphasized by Frankel and Rose (1998). Given the involvement of numerous CEECs and candidates in European production networks, the findings may reflect differences in the transmission of shocks across EEMU. It may be that economic and financial institutions in core and Northern countries have over time become much more harmonized than in periphery, CEEC and candidate countries. A useful task for further research would be to explore further what lies behind the direct membership effect and whether the strength of this channel depends on economic characteristics, such as the structure of production or similarity of institutions.

On this evidence, candidates with low levels of synchronization post EU adoption should be cautious in their euro adoption strategies, particularly those with poor cyclical alignment and weak financial positions. Possible instability in this relation, and the role of real interest rate differentials, might make certain member states more reliant on alternative adjustment mechanisms, namely fiscal policy. These results therefore also point to a reconsideration of the role of fiscal policy in EEMU.

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Table 1: Sample Summary Statistics

Country-Pair Group	Observations	Mean	Standard Deviation	Minimum	Maximum	Difference
Synchronization (Amplitude)						
All	13,249	-3.061	2.769	-26.943	0.000	-
EEMU	4,703	-2.837	2.424	-19.357	-0.001	0.224***
EMU	2,357	-2.715	2.403	-15.874	-0.001	0.346***
Core-Core	420	-1.157	0.806	-5.763	-0.001	-
Core-Periphery	700	-2.188	1.746	-10.002	-0.010	1.031***
Core-CEEC	1,120	-2.704	1.644	-12.257	-0.002	1.547***
Core-Northern	840	-3.023	2.247	-14.953	-0.001	1.866***
Core-Candidate	840	-2.710	2.065	-12.257	-0.002	1.553***
Synchronization (Concordance)						
All	13,249	0.659	0.474	0.000	1.000	-
EEMU	4,703	0.673	0.469	0.000	1.000	0.014***
EMU	2,357	0.681	0.466	0.000	1.000	0.022***
Core-Core	420	0.786	0.411	0.000	1.000	-
Core-Periphery	700	0.696	0.460	0.000	1.000	0.090***
Core-CEEC	1,120	0.653	0.476	0.000	1.000	0.133***
Core-Northern	840	0.664	0.473	0.000	1.000	0.122***
Core-Candidate	840	0.652	0.476	0.000	1.000	0.134***

Source: Author's own calculations.

Notes: Core countries include Austria, Belgium, Finland, France, Germany, Italy, Netherlands; periphery countries include Cyprus, Greece, Ireland, Portugal, Spain; CEEC countries include Bulgaria, Croatia, Czech Republic, Hungary, Poland, Romania, Slovakia, Slovenia; Northern countries consist of Scandinavian and Baltic countries: Denmark, Estonia, Latvia, Lithuania, Norway, Sweden; candidate countries include Bulgaria, Croatia, Czech Republic, Hungary, Poland, Romania. Mean, standard deviation and range are based on yearly cross-sectional averages computed over 1994-2013. Synchronization (Amplitude) is based on measure $S1_{ij,t}$ and scaled in percentage points. Synchronization (Concordance) is based on measure $S2_{ij,t}$. Difference column reports the absolute difference between the reference group (All or Core-Core country-pairs) and the other groups, along with the statistical significance of the difference indicated from a two-tailed t-test for amplitude and a two-group test of proportions for concordance, using *ttest* and *prtest* STATA commands respectively. Significance is indicated by: * (10%), ** (5%), *** (1%).

Table 2: Model 1 – Trade Intensity Estimates (Various Estimators)

Start-End	POLS	FE	Huber	GMM	AR2	F_Effects	JP	HT
All Years								
	0.332***	0.329***	0.260***	0.300***	0.22	YES***	0.44	1.00
	(0.046)	(0.047)	(0.022)	(0.037)				

Granger Causality Tests

$T_{ij,t}$ GC $BC_{ij,t}$	52.55***	48.45***	136.66***	
$BC_{ij,t}$ GC $T_{ij,t}$	24.97***	24.74***	75.70***	Obs. = 4,450

Source: Author's own calculations.

Notes: Panel estimation using POLS, FE, GMM and Huber estimators. Estimates correspond to coefficient φ for trade intensity in the first model using the amplitude measure of synchronization. AR2 is the adjusted R^2 under FE as a measure of model fit. F_Effects reports the outcome and significance of a joint F-test on all fixed effects (time and space), based on FE estimates. JP column reports the p-value for the J-test corresponding to GMM estimation. HT reports Hausman's (1978) Chi-square test statistic. The EMU joint membership dummy, control variables for financial integration, specialization, population differential and the crisis dummy are not reported for brevity. FE, GMM and Huber estimations include time and country-pair fixed effects. POLS estimations exclude fixed effects. All estimations include country-pair specific time trends. Granger causality Chi-square test statistics are reported based on estimations of synchronization on a single lag of trade intensity variable ($T_{ij,t}$ GC $BC_{ij,t}$) and vice versa ($BC_{ij,t}$ GC $T_{ij,t}$). Tests operate under a null hypothesis that the coefficient on the lagged variable does not differ significantly from zero (i.e. no Granger causation). Standard errors (in parentheses) are clustered to correct for heteroscedasticty and autocorrelation. Statistical significance is indicated by: * (10%), ** (5%), *** (1%).

Table 3: Model 1 – Trade Intensity and EMU Membership Estimates (Various Specifications)

	Full Sample Cross-Sectional		Aggregated Yearly Cross-Sectional		Panel			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Standard Controls								
φ	0.192**	0.080**	0.209***	0.072**	0.300***	0.079***	0.110***	0.184***
	(0.084)	(0.030)	(0.060)	(0.027)	(0.037)	(0.015)	(0.017)	(0.028)
β	0.868***	0.558***	0.746**	0.451***	0.606***	0.199**	0.149**	0.267**
	(0.257)	(0.071)	(0.277)	(0.087)	(0.119)	(0.042)	(0.068)	(0.116)
Standard Controls + Fixed Exchange Rate Dummy								
φ	0.217**	0.099***	0.359***	0.119**	0.330***	0.085***	0.108***	0.182***
	(0.078)	(0.027)	(0.085)	(0.054)	(0.039)	(0.015)	(0.017)	(0.029)
β	0.557**	0.318***	0.848*	0.248*	0.400***	0.158***	0.161**	0.286**
	(0.258)	(0.066)	(0.439)	(0.143)	(0.126)	(0.045)	(0.071)	(0.121)
FIX	0.618***	0.476***	0.747	0.524**	0.617***	0.125*	-0.036	-0.058
	(0.123)	(0.068)	(0.513)	(0.256)	(0.143)	(0.070)	(0.067)	(0.112)
Excluding Standard Controls								
φ	0.191**	0.080**	0.188**	0.075*	0.408***	0.144***	0.150**	0.271**
	(0.084)	(0.029)	(0.083)	(0.040)	(0.043)	(0.017)	(0.067)	(0.115)
β	0.827***	0.579***	0.704*	0.210**	0.728***	0.122***	0.102***	0.171***
	(0.256)	(0.078)	(0.403)	(0.084)	(0.145)	(0.044)	(0.016)	(0.027)
Obs.	253	253	4,450	4,450	4,450	4,450	4,450	4,450

Estimator	OLS	OLS	FE	FE	GMM	GMM	PROBIT	LOGIT
Dynamic	-	-	YES	YES	NO	NO	YES	YES
Country FE	-	-	YES	YES	-	-	-	-
Country-Pair FE	-	-	-	-	YES	YES	YES	YES
Time FE	-	-	-	-	YES	YES	YES	YES
Country-Pair Trends	-	-	-	-	YES	YES	YES	YES

Source: Author's own calculations.

Notes: Cross-sectional estimation, columns (1)-(2); cross-sectional estimation aggregated over individual years, columns (3)-(4); and panel estimation, columns (5)-(8). Coefficients for trade and EMU membership dummy reported based on EEMU full sample estimation. The standard control set includes measures of financial integration, specialization and a dummy variable for the 2008-2011 crisis years. Coefficients for control variables are not reported for brevity. Coefficients in the row FIX correspond to the fixed exchange rate regime dummy, using the classification of Ilzetzki et al. (2010). The dependent variable in columns (1), (3) and (5) is the amplitude measure based on (unfiltered) growth rates. The dependent variable in columns (2), (4) and (6) is the amplitude measure based the differential of HP-filtered cyclical deviations from trend. The dependent variable in columns (7)-(8) is the binary concordance measure. Estimator (OLS, FE, GMM, PROBIT and LOGIT) indicates estimator. Dynamic indicates whether the explanatory variable set is lagged by one year. Country, Country-Pair, and Time FE indicates whether country, country-pair and time fixed effects are included in the estimation. Country-Pair Trends indicates whether country-pair time trends are included. Robust standard errors (in parentheses) are clustered to correct for heteroscedasticty and autocorrelation. Statistical significance is indicated by: * (10%), ** (5%), *** (1%).

Table 4: Model 2 – Country-Group Estimations (Pre-Crisis)

Start-End	ROW	Core	Periphery	CEEC	Northern	Candidate	β	γ_1	γ_2	γ_3	JP
1995-1999	-0.007	0.298**	-0.093	0.386**	0.173**	0.458**	0.129	-0.503***	-0.001	-0.001**	0.644
	(0.041)	(0.145)	(0.111)	(0.179)	(0.075)	(0.202)	(0.312)	(0.047)	(0.012)	(0.000)	
1996-2000	0.030	0.370***	-0.077	0.408***	0.287***	0.474***	-0.103	-0.422***	0.000	-0.001**	0.885
	(0.038)	(0.133)	(0.107)	(0.157)	(0.080)	(0.175)	(0.268)	(0.044)	(0.011)	(0.000)	
1997-2001	0.064*	0.372***	-0.047	0.456***	0.365***	0.545***	0.079	-0.413***	0.000	-0.001**	0.661
	(0.035)	(0.123)	(0.098)	(0.139)	(0.083)	(0.155)	(0.201)	(0.043)	(0.010)	(0.000)	
1998-2002	0.113***	0.413***	-0.145	0.483***	0.625***	0.618***	0.189	-0.365***	-0.001	-0.000	0.811
	(0.034)	(0.132)	(0.097)	(0.137)	(0.091)	(0.153)	(0.186)	(0.045)	(0.010)	(0.000)	
1999-2003	0.181***	0.407***	-0.111	0.457***	0.873***	0.636***	0.213	-0.350***	0.000	-0.000	0.601
	(0.035)	(0.140)	(0.097)	(0.136)	(0.108)	(0.152)	(0.182)	(0.056)	(0.010)	(0.000)	
2000-2004	0.143***	0.311**	-0.139	0.236**	0.888***	0.356***	0.416**	-0.315***	-0.002	-0.001**	0.579
	(0.034)	(0.121)	(0.107)	(0.114)	(0.118)	(0.126)	(0.172)	(0.063)	(0.010)	(0.000)	
2001-2005	0.161***	0.363***	-0.120	0.170*	1.023***	0.269**	0.390**	-0.342***	-0.001	-0.001***	0.607
	(0.035)	(0.112)	(0.108)	(0.102)	(0.119)	(0.110)	(0.165)	(0.062)	(0.010)	(0.000)	
2002-2006	0.209***	0.443***	-0.118	0.147	1.250***	0.258**	0.442***	-0.268***	0.004	-0.001***	0.821
	(0.036)	(0.134)	(0.106)	(0.097)	(0.112)	(0.104)	(0.150)	(0.043)	(0.009)	(0.000)	
2003-2007	0.212***	0.438***	-0.093	0.117	1.406***	0.257**	0.643***	-0.172***	0.004	-0.002***	0.980
	(0.038)	(0.115)	(0.107)	(0.099)	(0.112)	(0.106)	(0.145)	(0.033)	(0.009)	(0.000)	
										Obs. = 8,256	

Notes: 5-year rolling window GMM panel estimation with country and time fixed effects, reported for the second model using the amplitude measure. Estimates reported for All other country-pairs (ROW), Core-Core (Core), Core-Periphery (Periphery), Core-CEEC (CEEC), Core-Northern (Northern) and Core-Candidate (Candidate) country-groups, EMU

membership dummy (β), financial integration (γ_1), specialization (γ_2), population differential (γ_3). Trade intensity estimates reported in columns correspond to the second model as follows: Trade intensity for country-group K (φ_2), rest of the world (φ_1) JP column reports the p-value for the J-test. Coefficients for the dummy variable to control for the marginal country-group effect on the intercept and the crisis dummy are not reported for brevity. Estimates for β, φ_1, γ_1, γ_2, γ_3 and JP statistic based on estimation of the second model for core-candidates. Standard errors (in parentheses) are clustered to correct for heteroscedasticty and autocorrelation. Statistical significance is indicated by: * (10%), ** (5%), *** (1%).

Figures 1-2*:* Synchronization of EEMUs and ROW under Amplitude and Concordance Measures
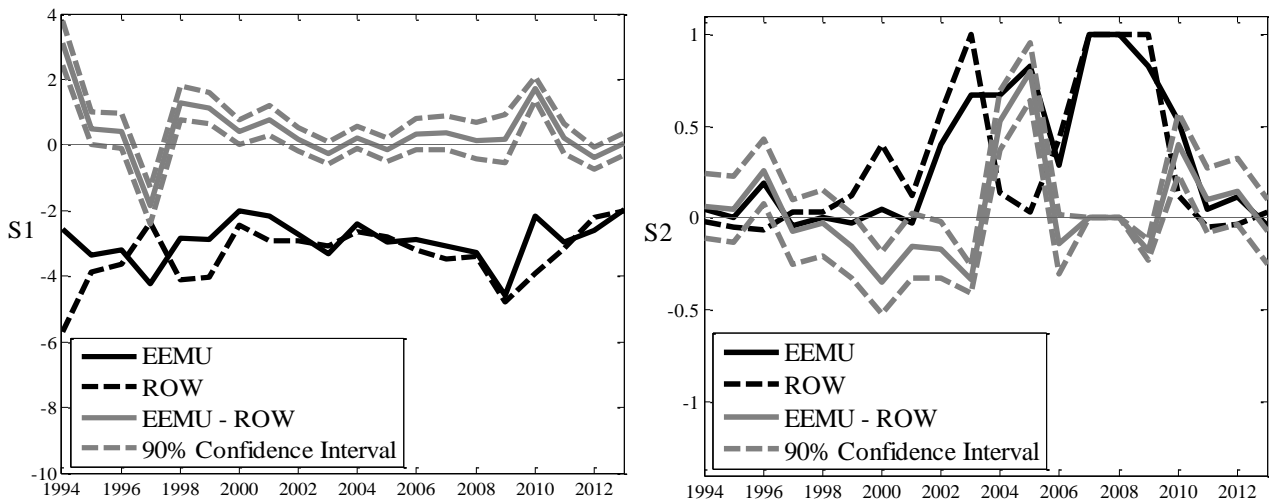
Source: Author's own calculations.

Notes: S1 and S2 (vertical-axis) correspond to amplitude and concordance-based synchronization of EEMUs (black solid line), synchronization in ROW (black dotted line), the difference (grey solid line), and the 90 per cent confidence interval (90% CI) for the difference (dotted grey line). S1 is scaled in percentage points.

Figures 3-6: 5-Year Rolling Window Estimation of Direct EMU Membership Effect
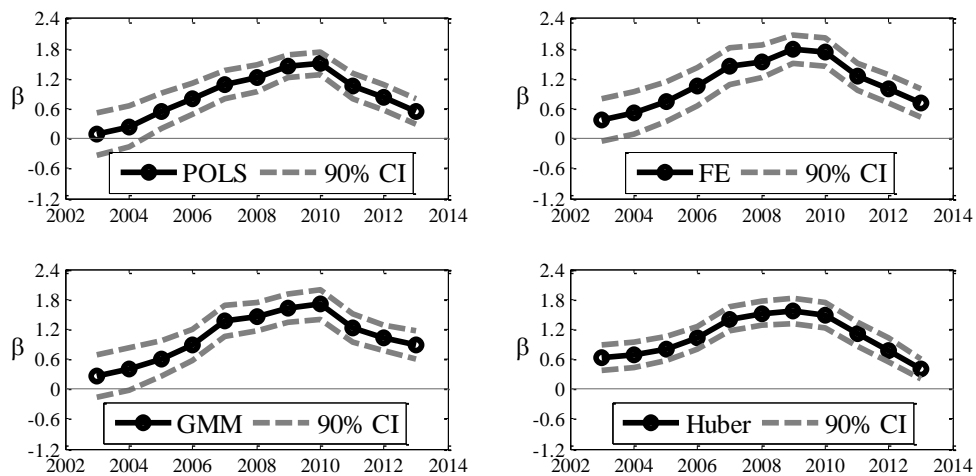
Source: Author's own calculations.

Notes: β (vertical-axis) corresponds to the EMU membership coefficient as set out in the first model, based on POLS, FE, GMM and Huber estimators. The 90 per cent confidence interval (90% CI) is based on heteroscedasticity and autocorrelation robust standard errors.

Figures 7-10: Evolution of the Relative Slope for Core-core, Core-periphery, Core-Northern and Core-candidates.
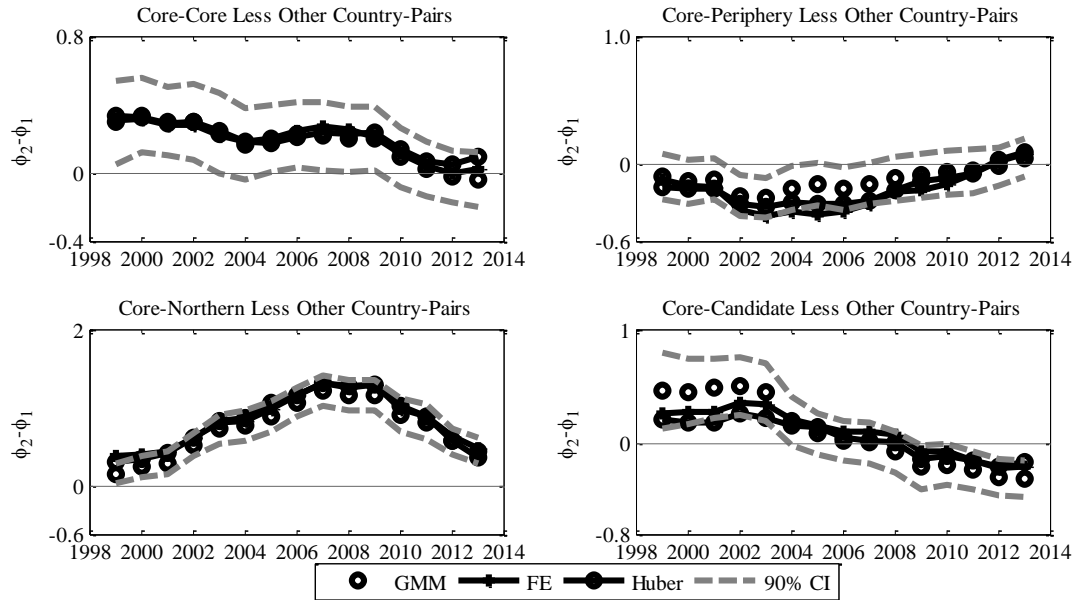
Source: Author's own calculations.

Notes: $\varphi_2 - \varphi_1$ (vertical-axis) corresponds to the difference in the trade-synchronization slope for core-non-core country-pairs over all other country-pairs estimated over 5-year rolling windows, based on GMM, FE and Huber estimators using amplitude synchronization measure. The 90 per cent confidence interval (90% CI) is based on heteroscedasticity and autocorrelation robust standard errors, computed under the GMM estimator.