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One Money, One Cycle? The EMU Experience

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Editorial

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September 25, 2013

One Money, One Cycle? The EMU Experience*

Martin Gächter† and Aleksandra Riedl‡

Abstract

We examine whether the introduction of the euro had a significantly positive impact on the synchronization of business cycles among members of Economic and Monetary Union (EMU) which might arise due to the lack of country-specific monetary policy shocks in the euro area. Empirical evidence on this relationship is rare so far and suffers from methodical weaknesses, such as the absence of time variability, which is crucial for addressing this issue. Using a synchronization index that is constructed on a year-by-year basis (1993–2011), we uncover a strong and robust empirical finding: the adoption of the euro has significantly increased the correlation of member countries' business cycles above and beyond the effect of higher trade integration. Thus, our results substantially strengthen the conclusion by Frankel & Rose (1998), i.e. a country is more likely to satisfy the criteria for entry into a currency union ex post rather than ex ante. Remarkably, however, this reasoning is even verified when controlling for the effect of increased trade linkages implied by entering a currency union.

JEL classification: E02, E32, E58, F15, F33;
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1 Introduction

The theory of optimum currency areas (OCA), originally developed by Mundell (1961), McKinnon (1963) and Kenen (1969), proposed a broad set of prerequisites a geographical area should fulfill for considering a currency union (CU).[1] While the topic lost some ground in the literature in the 1970s and 1980s, the Maastricht Treaty in the early 1990s outlined the path towards Economic and Monetary Union (EMU) and brought the discussion back to the center of economic policy relevance. In the years prior to the establishment of the euro area (EA) many studies examined business cycle synchronization among the prospective member states, as this measure can be seen as a 'meta-criterion' for entering a CU.[2] The underlying argument is simple: If the potential members of a monetary union are subject to symmetric economic shocks, the benefits of a common currency are likely to exceed the cost of relinquishing a national autonomous monetary policy (e.g., Bayoumi & Eichengreen 1997, Masson & Taylor 1993).

However, the seminal contributions by Frankel & Rose (1997, 1998) challenged this view, arguing that the suitability of a country for joining a CU is a more complex topic, as the OCA criteria themselves are endogenous. In their paper, the authors demonstrate a clear positive empirical relationship between stronger bilateral trade links and more synchronous cycles. Further, they argue that the expected increase in trade among future EMU member countries would lead to a higher cyclical co-movement in the euro area.[3] Thus, member states of a CU would fulfill the OCA criteria rather ex post than ex ante. Several studies have followed, which confirmed the positive relationship between trade and business cycle correlation (e.g., Artis & Okubo 2011, Inklaar et al. 2008, Baxter & Kouparitsas 2005).

Yet, the described trade link might not be the only channel through which a common currency is connected to more homogeneous business cycles. In fact, economies in a currency union might per se exhibit more synchronized cycles because of common monetary shocks under a single central bank. Furthermore, recent theoretical evidence also suggests a positive impact of EMU membership on business cycle (BC) synchronization. In particular, based on a two-country business cycle model calibrated for the euro area, Enders et al. (2013) show that cycles exhibit a higher co-movement under EMU because of stronger

[1] In particular, the probability that the benefits of establishing a currency union outweigh the costs of a common monetary policy increases if member states exhibit (i) high trade integration, (ii) highly synchronized business cycles, (iii) flexible labor markets, and (iv) an appropriate system of risk sharing (e.g. fiscal transfers, financial integration etc.).

[2] See de Haan et al. (2008) for a literature survey.

[3] It is well established in the literature that a currency union has a significant positive impact on trade integration (Glick & Rose 2002, Baldwin 2006).

spill overs of country-specific shocks to other EMU members.[4] Empirical evidence is rare, however, and the corresponding results are quite mixed. While Rose & Engel (2002) find that business cycles are more synchronized across currency union countries than across countries with sovereign monies, Baxter & Kouparitsas (2005) conclude that this relationship is not robust. In any case, none of those studies can be applied to the EMU experience, as each of them has a different country and time focus.[5]

Against this background, the aim of this paper is to explore whether the adoption of the euro has made EMU member countries' business cycles more homogeneous after their entry into the CU while controlling for the influence of increased bilateral trade. As we take it for granted that monetary policy cannot permanently affect either a country's real income level nor its growth rate, we subsequently focus on the cyclical component of GDP. While our results confirm the findings by Frankel & Rose (1997, 1998) that increased bilateral trade leads to higher business cycle synchronization, we are also able to show that membership in EMU per se leads to further convergence of business cycles above and beyond the effect of higher trade integration most likely due to the common monetary policy conducted by the European Central Bank.

Empirical papers on the euro effect are surprisingly rare and yield quite mixed results not least because of methodological issues related to the measurement and dynamics of business cycle synchronization over time. While the analysis by Christodoulopoulou (2013) suggests a negative euro effect, Gonçalves et al. (2009) find a positive impact of EMU membership on BC synchronization.[6] The observed results are derived using a difference-in-difference approach where the overall sample period is divided into two subperiods, i.e. the pre- and the post-euro period, such that the correlation coefficient of business cycles is observed for two time spans.[7] The difference between those coefficients is then regressed against an EMU dummy and other control variables. Yet, this method is afflicted with several caveats often encountered in the business cycle convergence literature.

[4] The model is calibrated for Germany vis-à-vis the aggregate of six euro area countries. The reason for the observation by Enders et al. (2013) is that domestic shocks depreciate the real exchange rate by less under a common currency (compared to the pre-EMU period) because the nominal exchange rate channel is absent and prices are assumed to be sticky.

[5] The sample of Rose & Engel (2002) is restricted to poor and small countries in the period 1960-1996, and Baxter & Kouparitsas (2005) analyze 100 developed and developing countries up to the year 1995 only.

[6] Note that there are several papers analyzing the euro effect that are not directly related to our work. These include e.g. Giannone et al. (2008) and Lehwald (2012), who focus on output (per capita) fluctuations rather than on the cyclical component of GDP, Weyerstrass et al. (2011) and Furceri & Karras (2008), who base their analysis on visual inspection of various measures of business cycles.

[7] In a robustness analysis, Christodoulopoulou (2013) provides results from panel data, arriving at the same conclusion, i.e. a negative euro effect. However, this result is derived by using correlation coefficients of real GDP growth rates as the dependent variable and is therefore not comparable to our results (see also section 2). Moreover, as the author states in his conclusions, important control variables like trade integration are missing and potential endogeneity issues are not controlled for.

First, and most importantly, the use of cross-correlations is insufficient to answer the research question of interest. More precisely, exploring time variability is crucial for identifying the "euro effect", i.e. whether business cycle synchronization increased across member states after their entrance into the CU ("within variability"), which is in contrast to the question whether countries are more synchronized within currency unions compared to non-members ("between variability"). For this purpose one also needs to evaluate deviations of BC synchronization from individual means over time, which can be accomplished by employing panel data methods. The same line of argument applies to the evaluation of the trade effect. As we will argue below, the direction of correlation between EMU membership and bilateral trade changes when time variability is controlled for. Hence, considering only differences in bilateral trade relations across countries and not over time leads to even opposite conclusions, i.e. higher trade integration is negatively related to business cycle synchronization. This observation might explain why Gonçalves et al. (2009) using difference-in-difference methods find a negative trade effect while Frankel & Rose (1998) who explore variability over time (by evaluating four subperiods) come to the opposite conclusion, i.e. higher trade integration increases BC synchronization, which is also confirmed in our study. Another related caveat of difference-in-difference techniques is that the observed window span to calculate correlation coefficients is set arbitrarily. Therefore, the results are more sensitive with respect to the inclusion (or exclusion) of certain time periods. Finally, in a difference-in-difference analysis, different EMU entrance times cannot be considered, as the concept allows only for one pre- and post "treatment" period. Hence, the analysis is always restricted to a subsample of the group of currently 17 EMU members.[8]

To overcome these caveats, we use a slightly adapted version of an index suggested by Cerqueira & Martins (2009) which allows to observe business cycle synchronization on a year-by-year basis. Based on a newly compiled panel data set including all members of the European Union (EU) in the period 1993-2011, we observe 17 out of 27 EU countries, which sequentially joined EMU during the considered time span.[9] Since our control group comprises all remaining EU countries, which have not implemented the euro, we are able to distinguish the euro effect, i.e. a common monetary policy, from the EU effect, which might have potentially driven up business cycle correlation due to the establishment of a single market (i.e. free movement of people, goods and capital). To estimate the impact of EMU membership on BC synchronization while controlling for bilateral trade and various other potential determinants, we apply a system GMM estimation technique, which allows

[8] Above all, the difference-in-difference approach exhibits another important caveat, as it is based on the "parallel-trend assumption", which cannot be tested empirically.

[9] The countries that joined EMU in 1999 are: Austria, Belgium, Finland, France, Germany, Ireland, Italy, Luxembourg, the Netherlands, Portugal and Spain; in 2001: Greece; in 2007: Slovenia; in 2008: Cyprus and Malta; in 2009: Slovakia; in 2011: Estonia.

us to use internal instruments to overcome the problem of endogeneity. Our results indicate that the implementation of the common monetary policy starting in 1999 had a significant positive effect on the synchronization of EMU member countries' business cycles. Thus, we are able to conclude that the assessment whether a country group should establish a CU should not solely rely on an ex ante examination of business cycle synchronization.

The paper is structured as follows. In section 2, we introduce our measure of business cycle synchronization and describe the control variables that enter the regression analysis. Subsequently, we present our econometric model and discuss some estimation issues (section 4). The corresponding results are presented in section 5. At the end of the paper, we provide some conclusions from the analysis (section 6).

2 Measurement of Business Cycle Synchronization

2.1 The Synchronization Index

To measure business cycle synchronization, we use a slightly adapted version of an index developed by Cerqueira & Martins (2009), which in its original form is given by

$$\rho_{ij,t} = 1 - \frac{1}{2}\left(\frac{d_{j,t}-\bar{d}_j}{\sqrt{\frac{1}{T}\sum_{t=1}^{T}(d_{j,t}-\bar{d}_j)^2}} - \frac{d_{i,t}-\bar{d}_i}{\sqrt{\frac{1}{T}\sum_{t=1}^{T}(d_{i,t}-\bar{d}_i)^2}}\right)^2 \tag{1}$$

where $d_{j,t}$ and $d_{i,t}$ are annual GDP growth rates of countries j and i from year $t-1$ to t.[10] This measure captures the correlation of GDP growth rates between individual country pairs at each single point in time, while its mean over time corresponds to the correlation coefficient conventionally applied, i.e. $\frac{1}{T}\sum_t \rho_{ij,t} = \rho_{ij}$ with $\rho_{ij} = \frac{Cov(d_i,d_j)}{\sigma(d_i)\sigma(d_j)}$. Hence, one of the main advantages of this index is that it allows us to exploit the time variability of the data as it distinguishes between specific episodes of asynchronous behavior and periods of highly positive cyclical correlations.[11] This enables us to evaluate deviations of BC synchronization from individual means over time, which is crucial for answering our research question. Another big advantage of this index is that it avoids setting arbitrary time spans to compute correlation coefficients. The bulk of studies on business cycle

[10] Cerqueira & Martins (2009) provide a derivation of this measure in their paper.

[11] Crespo-Cuaresma & Fernández-Amador (2013*a*,*b*) propose the standard deviation of cyclical components as an alternative measure of BC synchronization. While their indicator is also able to assess developments within a country group over time, it is not possible to use the measure in a panel setting where individual country pairs are observed. Moreover, the standard deviation of cyclical components considers the cyclical position (i.e. cyclical amplitude) rather than its co-movement (i.e. cyclical direction) and is therefore not directly comparable to our study.

synchronization apply either one correlation measure for each country pair over the entire time span (resulting in a cross-country analysis, see e.g. Inklaar et al. 2008) or correlation coefficients for different non-overlapping window spans (resulting in a panel data set with a small time dimension, see e.g. Frankel & Rose 1998, Abbott et al. 2008, Calderon et al. 2007). A further alternative would be to use overlapping window spans (see e.g. Massmann & Mitchell 2004). This, however, also implies a loss of observations and – more importantly – leads to a heavily autocorrelated dependent variable, which is difficult to handle in econometric analysis. For these reasons, we employ the index given in equation (1), which we adapt in two important directions.

First, as we are interested in exploring the co-movement of countries' business cycles we use the cyclical component of real GDP rather than real GDP growth rates. This distinction is particularly important when studying the EU-27 countries. Contrary to Cerqueira & Martins (2009), who analyze a sample of 20 OECD countries, our study deals with a more heterogeneous country sample including industrialized as well as transformation economies (former communist countries). These country groups exhibit rather different developments of trend growth rates, as they are at different stages of economic development. Hence, if one country pair from our sample exhibits a poor correlation of GDP growth rates, it does not necessarily imply a poor correlation of their cyclical components or vice versa.

Second, since the measure given in equation (1) is bounded between $3 - 2T$ and 1 (see Cerqueira 2013), we use the transformation proposed by Cerqueira (2013) to yield a symmetric range of the index, i.e. between $-\infty$ and $+\infty$. This is important to consider, as a bounded correlation coefficient is unlikely to result in normally distributed errors in the context of regression analysis (see Inklaar et al. 2008).[12]

Altering the index given in equation (1) by the two modifications described above finally yields our measure of business cycle correlation

$$\text{Correl}_{ij,t}^{Cycle} = \frac{1}{2}\log\left(\frac{1+\frac{\rho_{ij,t}^{Cycle}}{2T-3}}{1-\rho_{ij,t}^{Cycle}}\right) \tag{2}$$

where

$$\rho_{ij,t}^{Cycle} = 1 - \frac{1}{2}\left(\frac{c_{j,t}-\overline{c}_j}{\sqrt{\frac{1}{T}\sum_{t=1}^{T}(c_{j,t}-\overline{c}_j)^2}} - \frac{c_{i,t}-\overline{c}_i}{\sqrt{\frac{1}{T}\sum_{t=1}^{T}(c_{i,t}-\overline{c}_i)^2}}\right)^2 \tag{3}$$

with $c_{j,t}$ and $c_{i,t}$ denoting the cyclical component of real GDP of countries i and j for the years $t = 1, ..., T$. This measure is finally applied to analyze whether EMU member

[12] In Figure 2 in the Appendix, we plot kernel density estimates of the transformed correlation coefficient showing that our dependent variable is very close to being normally distributed.

country pairs exhibit higher business cycle synchronization after accession and amidst other EU members.

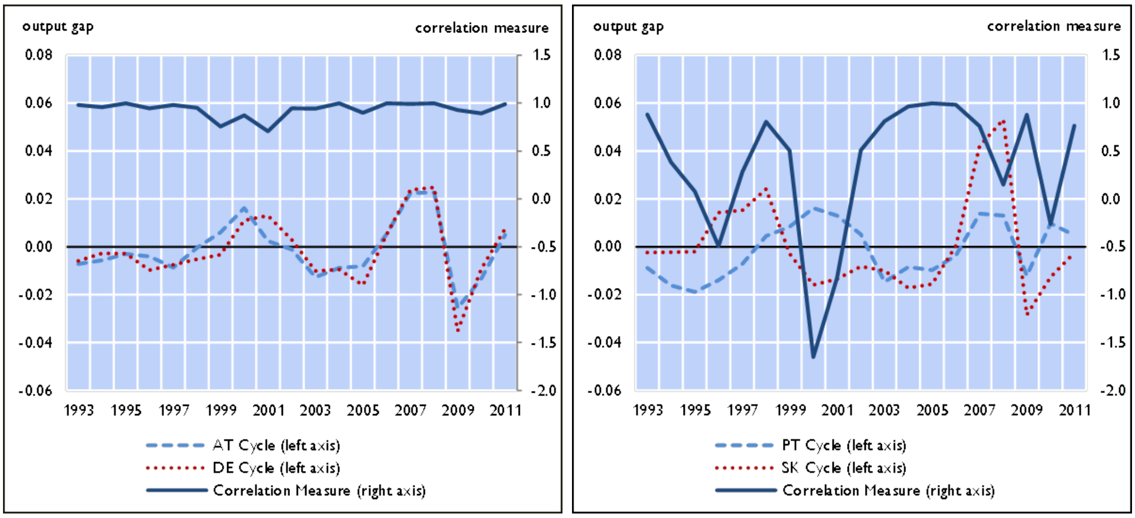
2.2 Filtering Techniques and Data Sample

In order to compute (2), we need to extract the cyclical component from real GDP data first. A handful of business cycle extraction methods are used in the literature. We apply the Hodrick-Prescott (HP) filter (Hodrick & Prescott 1997), which is the most widely applied method in the literature, hence increasing the comparability of our results with those of other studies. The degree of trend smoothing, which has to be determined ex ante, is set to 6.25, as we deal with yearly observations (see e.g. Ravn & Uhlig 2002).[13] Although the choice of the filter method might not influence the degree of synchronization between business cycles as shown by e.g. Massmann & Mitchell (2004),[14] we nevertheless apply an additional filter method – the Baxter-King (BK) band-pass filter (Baxter & King 1999) – in order to check the robustness of our results. As originally suggested by Baxter & King (1999) for annual macroeconomic data, our BK filter admits periodic components between two and eight years, with the lead-lag length of the filter being $K = 3$. Thus, in the case of the BK filter, we lose three years at the beginning and at the end of our sample.

We consider real GDP data (in euro) of 27 EU Member States (EU 28 excluding Croatia). As GDP data for some transition countries are available only from 1993 onward, the subsequent estimations are restricted to the period 1993 to 2011. However, estimates for the output gap are based on the maximum available time span within the range 1988 to 2011. All data are extracted from Eurostat's online database and are thus comparable across countries as well as over time. The number of observations for the synchronization measure amounts to 6669.[15]

To provide a better insight into the applied synchronization measure, Figure 1 plots the development of the index in its untransformed version (given by equation (3)) for two country pairs together with the corresponding business cycles. We choose Austria and Germany as a country pair example for a high co-movement in cycles with $\frac{1}{T}\sum_t \rho_{ij,t}^{Cycle} = 0.93$, while Portugal and Slovakia serve as an example for a relatively low cyclical synchronization over time with $\frac{1}{T}\sum_t \rho_{ij,t}^{Cycle} = 0.34$. As shown in plot (a), Austria and Germany already had highly synchronized cycles before introducing the common currency. Clearly, this is not surprising given their strong bilateral trade links and the currency peg of the Aus-

[13] The HP filter is applied by using the 'hprescott' routine implemented in STATA.
[14] Similar results are obtained by Artis & Zhang (1997) and Calderon et al. (2007).
[15] The number of observations is calculated from $\frac{N\times(N-1)}{2}$ country pairs, with $N = 27$ being the number of countries, which are observed for $T = 19$ years.



(a) Austria (AT) and Germany (DE)

(b) Portugal (PT) and Slovakia (SK)

Figure 1: Graphical Representation of the Correlation Measure

trian schilling to the Deutsche mark since the beginning of the 1980s. In fact, the first years after the euro introduction were characterized by a slight decrease in cyclical correlations, although the correlations increased again thereafter and stabilized at a high level. Panel (b) shows an example of two much less synchronized cycles. In such cases, the time variability of our synchronization measure gains importance, as periods of both convergence and divergence can be detected. The illustration shows highly volatile business cycle synchronization between Portugal and Slovakia before 2000. In the run-up to EU membership, the two cycles clearly converged, before somehow diverging again during the crisis period.

2.3 First Descriptive Results and the Advantages of Panel Data

Returning to the whole country sample, the Fisher-transformed synchronization measure applied in our regression analysis yields a sample average of 1.21 for country pairs that are not in EMU, while the corresponding sample average of EMU country pairs amounts to 1.50 (see Table 1). Obviously, business cycle synchronization tends to be higher for countries which are in a common currency area. Yet, this observation might be driven by the fact that EMU country pairs exhibit higher trade relations leading to a higher cyclical co-movement. As one can see from Table 1 bilateral trade measured in % of GDP is more than twice as high in euro countries than in others.

A closer look at descriptive statistics, however, reveals that the positive relationship be-

Table 1: Descriptive Statistics: Sample Averages for Country Pairs by EMU Membership, 1993-2011

Variable	Non-EMU country pairs	EMU country pairs
Business Cycle Correlation	1.21	1.50
Bilateral Trade (in % of GDP)	0.39	0.85
No. of observations	5,603	1,065

tween EMU membership and business cycle correlation is still significant when bilateral trade is controlled for. This can be seen in the left column of Table 2, which shows the corresponding partial correlation coefficients. Interestingly, repeating this exercise by ignoring time variability yields different results. Constructing the corresponding variables along the lines of the difference-in-difference approach (by taking time averages) shows that bilateral trade is negatively correlated with BC synchronization once EMU membership is controlled for. The different outcomes can be explained by the fact that the increase in bilateral trade relations among EMU members was more sluggish, especially compared to the former communist countries, while the increase in BC synchronization was lower in the latter country group than in the former. The reason for this observation is twofold. First, the decrease in bilateral trade during the financial crises was more pronounced in EMU member countries. Second, the level of bilateral trade between the former communist countries was small due to their lower stage of economic development. Therefore, difference-in-difference techniques would suggest a negative link between BC synchronization and trade relations, as deviations from individual means over time are not considered. Hence, ignoring the dynamics along the time dimension might mask the link between relevant variables and might lead to biased results. The pairwise correlation coefficients given in Table 3 strengthen the argument. Considering only the between variability of the sample changes the sign of the correlation coefficients between trade and EMU membership – an outcome which stands in sharp contrast to the results established so far in the literature.

Table 2: Partial Correlation Coefficients for Business Cycle Correlation

	Panel data approach	Diff-in-diff approach*
EMU	0.06***	0.16**
Bilateral Trade	0.10***	-0.06

* BC correlation is measured as the difference between the correlation coefficient of two country cycles in the periods 1999–2011 and 1993–1998, i.e. $correl(c_i, c_j)^{99-11} - correl(c_i, c_j)^{93-98}$. Correspondingly, bilateral trade is measured as $mean(trade_{ij})^{99-11} - mean(trade_{ij})^{93-98}$. The six countries that joined EMU after 1999 are therefore excluded. For data measurement and sources see section 2.

After having discussed some first descriptive results we will, in section 4, move on to testing

Table 3: Pairwise Correlation Coefficients

	Panel data approach		Diff-in-diff approach	
	BC correlation	EMU	BC correlation	EMU
EMU	0.09		0.23	
Bilateral Trade	0.12	0.22	-0.13	-0.39

whether there is a causal effect of EMU membership on the cyclical co-movement when bilateral trade and various other potential determinants of business cycle synchronization are controlled for.

3 Determinants of Business Cycle Synchronization

The explanatory variable of main interest is a simple dummy variable for EMU membership, denoted as $EMU_{ij,t}$. It takes on the value of 1 if both countries i and j are members of EMU in year t and amounts to 0 otherwise. Following Frankel & Rose (1997, 1998), we would expect increased BC synchronization across EMU countries due to increased bilateral trade relations, as they argue that the optimum currency area (OCA) theory criteria are jointly endogenous. The higher extent of BC synchronization of EMU country pairs has already been shown in Table 1. However, in order to show that the common monetary policy per se has a positive effect on BC synchronization within EMU even beyond the above mentioned trade effect, we have to include several control variables, as explained below.

First of all, many studies find a highly significant positive impact of (bilateral) trade integration on BC synchronization (see, for instance, Frankel & Rose 1998, Baxter & Kouparitsas 2005, Inklaar et al. 2008, Artis & Okubo 2011). Several measures for bilateral trade have been proposed in the literature. Following Frankel & Rose (1998), we consider the two most common measures in the literature. In our baseline model, we use bilateral trade between two countries i and j relative to the sum of their (nominal) GDPs in period t. More precisely, bilateral trade is given by

$$BT^1_{ij,t} = \frac{Exports_{ij,t} + Exports_{ji,t}}{GDP_{i,t} + GDP_{j,t}}. \tag{4}$$

where $Exports_{ij,t}$ refers to all exported goods from country i to country j in time t. For robustness purposes, we will also consider bilateral trade relative to total trade as an independent variable, i.e.

$$BT^2_{ij,t} = \frac{Exports_{ij,t} + Exports_{ji,t}}{Trade_{i,t} + Trade_{j,t}}. \tag{5}$$

where $Trade_{i,t}$ denotes the sum of all imported and exported goods of country i (j) in time t.[16]

Another aspect also highlighted in the literature is the role of industrial specialization patterns for BC synchronization. A higher degree of specialization would likely lead to greater vulnerability to asymmetric shocks. However, the effect of EMU participation on specialization is not unambiguous. On the one hand, as postulated by Krugman (1991), the reduction of trade barriers leads to an increase in inter-industry trade and opportunities for exploiting economies of scale. Therefore, specialization in production would arise whenever countries have a comparative advantage, leading to less diversified production structures and increased vulnerability to asymmetric shocks. On the other hand, as stated by the literature on OCA endogeneity (see, for instance, Frankel & Rose 1998), the participation in EMU leads to greater intra-industry trade integration, more similar economic structures and thus to more synchronized cycles due to the convergence of factor endowments and reduced exchange rate variability. To control for a possible specialization effect, we therefore include the following independent variable.

$$IS_{ij,t} = \sum_{k=1}^{K} |S^k_{i,t} - S^k_{j,t}| \tag{6}$$

where $S^k_{i,t}$ ($S^k_{j,t}$) represents the weight of sector k in the exports of country i (j) at time t. The resulting industrial specialization index thus mirrors absolute differences in export structures across countries, while we distinguish between six sectors ($k = 6$) in our analysis.[17] An increasing index $IS_{ij,t}$ indicates a higher degree of specialization differences between countries i and j. In our sample, the cross-section mean of $IS_{ij,t}$ increased from 0.45 in 1993 to 0.50 in 2011, suggesting a slow strengthening of specialization and diverging export structures over time.

From a theoretical point of view, fiscal policy has only rarely been mentioned in the OCA literature. In fact, the role of fiscal differentials on BC synchronization is theoretically

[16] Trade data are extracted from Eurostat (EU27 Trade Since 1988 By SITC). Missing data are provided by UNComtrade and the Vienna Institute for International Economic Studies.

[17] The choice of the sectors is based on the Standard International Trade Classification (SITC). The data are available from Eurostat and are provided for the years 1990 to 2011 for the following product groups: (1) SITC0-1 – food, beverage & tobacco (2) SITC2-4 – raw materials (3) SITC3 – mineral fuels, lubricants & related materials (4) SITC5 – chemical products (5) SITC6-8 – other manufactured articles (6) SITC7 – Machinery & transport equipment. Some missing data for the Eastern European countries were complemented by data provided by the Vienna Institute for International Economic Studies (`http://www.wiiw.ac.at`).

ambiguous. On the one hand, more similar budget deficits could indicate less proactive policymaking, leading to fewer fiscal shocks, and thus, to more synchronized business cycles. On the other hand, fiscal policy might also work the opposite way when deficits are used to smooth out cyclical deviations in recession and boom periods, thereby contributing to more similar cycles. Empirically, Darvas et al. (2005) and Artis et al. (2008) show that the former effect seems to outweigh the latter, i.e. countries with divergent fiscal policies tend to have less synchronized business cycles. Darvas et al. (2005) conclude that "irresponsible fiscal policy (a persistently high deficit) coincides with idiosyncratic (fiscal) instability", and thus, as fiscal policies converge, fiscal shocks are also reduced. While the two studies mentioned above are based on cross-sectional data sets, we are also able to examine this relationship in our panel setting. For this purpose, we define fiscal differences ($FD_{ij,t}$) between two countries i and j as follows,

$$FD_{ij,t} = |fb^{ca}_{i,t} - fb^{ca}_{j,t}| \tag{7}$$

where $fb^{ca}_{i,t}$ ($fb^{ca}_{j,t}$) is the cyclically adjusted fiscal balance (net lending / net borrowing in percent of GDP) of country i (j) at time t.[18] According to earlier studies, we expect a negative relationship between fiscal differences and BC synchronization, i.e. that a higher degree of fiscal divergence leads to lower BC synchronization.

Finally, we also include a simple EU dummy in our regression, as the accession to the common market might lead to a further convergence of policies, and thus to increased BC synchronization while controlling for other factors as mentioned above.

4 Estimation

When assessing the impact of the euro on the synchronization of member countries' business cycles two important estimation issues stand out. First, as was already pointed out by Frankel & Rose (1998) and Rose & Engel (2002), EMU accession as well as bilateral trade are endogenous with respect to BC synchronization. The underlying argument is straightforward: Countries with highly synchronized BC have a higher propensity to join a CU and, using the same line of argument, countries with tight trade ties are more likely to adopt a common currency and gain from the absence of transaction costs. Hence, one has to control for the potential endogeneity of these variables when performing the regression analysis. Second, as a test for serial correlation indicates, the applied synchronization measure is autocorrelated. This issue has to be addressed in order to obtain a reliable inference. For these reasons, we employ a dynamic panel data model, where the dependent

[18] The data are extracted from the Ameco Database (`http://ec.europa.eu/economy5_finance/ameco`).

variable enters the list of regressors in its one-period lagged form, i.e.,

$$\text{Correl}_{ij,t}^{Cycle} = \alpha + \beta \text{Correl}_{ij,t-1}^{Cycle} + \gamma EMU_{ij,t} + Z'_{ij,t}\delta + \mu_{ij} + \lambda_t + \nu_{ij,t} \tag{8}$$

where ij denotes the country pair with $ij = 1, ..., 351$ and t denotes time ($t = 1, ..., 19$). $Z_{ij,t}$ is the itth observation on four control variables, which have been discussed in section 3. We also include country-pair specific effects $\mu_{ij} \sim IID(0, \sigma_\mu^2)$ to characterize the unobservable heterogeneity among the individual country pairs as well as time-fixed effects λ_t to control for common global shocks, like commodity price effects. As already discussed in section 3, the variable of main interest is $EMU_{ij,t}$, which is a dummy that takes on the value of 1 if countries i and j both belong to the EMU in year t, and 0 otherwise.

To estimate the model in equation (8), we apply the feasible system GMM estimator developed by Blundell & Bond (1998). This estimator adequately addresses the endogeneity issues by using all available lags in levels as instruments for the differenced variables as well as past differences as instruments for the variables in levels. This has the advantage that time-invariant regressors can be estimated, which would be wiped out by a difference GMM estimator (Arellano & Bond 1991). This is relevant, as we will employ the distance between country pairs (which is, of course, time-invariant) as an additional instrument for bilateral trade. Moreover, applying the feasible estimator allows us to control for arbitrary patterns of heteroscedasticity. The underlying assumption behind the error structure is that the idiosyncratic disturbances are not correlated across country pairs. This assumption is quite reasonable, as the model includes time dummies (Roodman 2009).[19]

[19] As pointed out by Roodman (2009), applying GMM estimation methods involves many specification choices, which should be reported for traceability reasons. In this paper, we use the *xtabond2* routine of the Stata software with the options *two-step robust* to implement the estimator. As the EMU dummy and the trade variable are assumed to be endogenous, they enter (together with the correlation measure) in the *gmmstyle* option in their one-year lagged form. As we will also employ an EU dummy, which might potentially be endogenous as well, it enters the *gmmstyle* option in the same fashion. The remaining variables enter the *ivstyle* option in their current-period form. In addition to the exogenous variables, we include external instruments to instrument bilateral trade. In particular, we use the common determinants found in the literature to significantly impact on bilateral trade (see e.g. Frankel & Rose 1998), i.e. the distance between two country pairs (in logs), a common border dummy and a country's population size (in logs). Those also enter the *ivstyle* option. Finally, to overcome the problem of too many instruments, which might overfit endogenous variables or can weaken the Hansen test (Roodman 2009), we restrict the number of instruments used to up to nine time lags.

5 Results

5.1 Baseline Estimates

The estimation results for the model outlined above (equation (8), section 4) are reported in Table 4. We briefly describe the main features of Table 4 before continuing with the interpretation of our findings. Column (1) starts out by presenting the impact of EMU membership and bilateral trade on business cycle synchronization when no control variables other than time and country fixed-effects are considered. Columns (2) to (4) report the results when control variables are added one at a time. In the remaining columns, we check if our results are robust when we apply an alternative filtering method (5) and use a different synchronization measure (6). At the bottom of Table 4, we report the required test statistics related to the GMM estimation technique. In particular, as the moment conditions are set up under the condition that disturbances are not serially correlated, the differenced residuals should be correlated of order one but not of order two. This is confirmed by the reported Arellano-Bond tests. Moreover, as the number of instruments outweigh the number of regressors, a Hansen test of overidentifying restrictions is reported. According to the test statistics, the set of instruments is valid across all specifications. We also report test statistics related to the joint significance of time dummies, which are highly significant in each specification.

We are now in a position to interpret the magnitude of the variable of interest, namely the EMU dummy variable. The results reported in column (1) support the hypothesis that EMU membership significantly increases the correlation of countries' business cycles above and beyond the effect of increased trade integration. The magnitude of the coefficient suggests that the adoption of the euro raises synchronization among member countries by 0.15 (the sample mean is equal to 1.26). Yet, from an economic point of view, it would be interesting to compare the magnitude of this effect to the one of our trade variable. For this reason, we first calculate the beta (standardized) coefficient for bilateral trade, which is simply the original coefficient estimate multiplied by the ratio of the standard deviations (STD) of the independent and the dependent variable (see e.g. Wooldridge 2002). Hence, based on the coefficient estimate of 0.070, an increase in one STD of bilateral trade raises synchronization by 0.10.[20] Remarkably, this result is very much in line with Frankel & Rose (1998), who arrive at a beta coefficient estimate of 0.13 for bilateral trade. Second, since we will never observe the EMU dummy to increase by one standard deviation (≈ 0.37) in reality, we have to evaluate the impact of an increase of bilateral trade by 2.7 STDs, as this corresponds to setting the EMU dummy to 1. In this case, the impact of bilateral trade on business cycle synchronization is roughly 0.26. Hence, the effect of a common monetary

[20] The respective calculation is $0.070 * \frac{1.657}{1.146} \approx 0.10$.

policy amounts to more than half of the effect of trade integration. While Frankel & Rose (1998) have shown the endogeneity of the OCA criteria with respect to trade integration, our results suggest that the OCA endogeneity theory also applies to the common monetary policy in a CU, which removes any (country-specific) monetary shocks, and thus appears to smooth business cycles. Remarkably, at the same time, we are able to confirm the strong effect of bilateral trade relations shown by Frankel & Rose (1998) even in its exact magnitude.

In columns (2) to (4), we want to observe whether the impact of EMU membership changes when several other control variables are included. In all cases, the significance of the EMU dummy remains unchanged, while the amount of the coefficient estimate even increases slightly. In column (2), we control for industrial specialization, which might have a negative impact on synchronization for the reasons already outlined above. Although this variable slightly increased in the course of the observed time span (e.g. from a cross-section sample mean of 0.45 in 1993 to 0.50 in 2011), it obviously had no significant impact on business cycle synchronization at all. In column (3), we control for the similarity in fiscal policies among countries, as one might argue that the EMU dummy captures the impact of a coordinated fiscal policy among EMU countries due to the existence of the Maastricht criteria. Indeed, we observe that more diverging fiscal policies (i.e., cyclically adjusted public deficits) have led to more asynchronous business cycles, but more importantly, the inclusion of this variable does not dampen the EMU coefficient. Finally, in column (4), we add a dummy variable for country pairs that are in the EU. The inclusion neither has a significant impact on synchronization nor does it alter the EMU dummy. This result is also rather intuitive: While accession to the EU likely leads to stronger trade relations (as the country gains access to the common market), this relationship is already captured by the variable of bilateral trade. On the other hand, EMU accession not only has an impact on trade relations, but also includes a common monetary policy framework, leading to more synchronous business cycles above and beyond the mentioned trade effect.

In column (5), we want to check whether the observed impact of EMU membership is related to our measurement of the dependent variable. Instead of employing the filter suggested by Hodrick & Prescott (1997), we implement the Baxter-King filter. This alteration seems to even increase the impact of EMU membership by 50%, while it turns the coefficient of fiscal policy insignificant. Yet, the changed impact of fiscal policy might be explained by the omission of the crises years (due to the Baxter-King method) where consolidation efforts diverged considerably across countries because of the different impact of the financial crises on their public debt levels.

Table 4: Baseline Results. Maximum Time Span: 1993–2011

Synchronization Measure:	$\text{Correl}_{HP}^{Cycle}$	$\text{Correl}_{HP}^{Cycle}$	$\text{Correl}_{HP}^{Cycle}$	$\text{Correl}_{HP}^{Cycle}$	$\text{Correl}_{BK}^{Cycle}$	$\text{Spread}_{HP}^{Cycle}$
Variable/Model	(1)	(2)	(3)	(4)	(5)	(6)
EMU	0.152***	0.162***	0.164***	0.163***	0.231***	-0.391***
	(0.052)	(0.056)	(0.051)	(0.056)	(0.066)	(0.047)
Bilateral Trade	0.070***	0.062***	0.056***	0.062***	0.078***	-0.158***
	(0.012)	(0.015)	(0.016)	(0.016)	(0.019)	(0.020)
Industrial Specialization		-0.016	0.010	0.049	0.109	0.035
		(0.083)	(0.084)	(0.085)	(0.095)	(0.102)
Fiscal Policy			-0.023***	-0.022***	0.001	0.037***
			(0.006)	(0.006)	(0.008)	(0.006)
EU				0.060	0.011	0.035
				(0.054)	(0.059)	(0.048)
$\text{L1.Correl}_{HP}^{Cycle}$	0.090***	0.093***	0.099***	0.101***		
	(0.014)	(0.015)	(0.014)	(0.014)		
$\text{L1.Correl}_{BK}^{Cycle}$					0.056***	
					(0.019)	
$\text{L1.Spread}_{HP}^{Cycle}$						0.353***
						(0.019)
Constant	1.582***	1.526***	1.583***	1.536***	0.630***	-0.400***
	(0.092)	(0.098)	(0.098)	(0.127)	(0.215)	(0.138)
Time-FE χ^2	267.54***	258.33***	220.93***	227.80***	222.28***	607.71***
Arellano-Bond test AR(1)	-13.718***	-13.540***	-13.687***	-13.671***	-11.500***	-12.210***
Arellano-Bond test AR(2)	-1.418	-1.327	-0.500	-0.472	0.244	-1.423
Hansen test statistic	340.915	340.922	340.788	342.339	342.076	348.830
Hansen p-value	0.926	0.926	0.926	0.997	0.383	0.994
N	6329	6084	5673	5673	4560	5675

Two-tailed significance levels: $*$: 10% $**$: 5% $***$: 1%. Standard errors are reported in parentheses. Out-of-sample instruments included: logdistance, logpopulation, common-border dummy. In-sample instruments: up to 9 lags.

In the last column, we employ a different synchronization measure. In particular, we measure the absolute deviation between two countries' business cycles, i.e. we capture the difference in the amplitude of the cycle instead of the correlation. This alteration is related to the fact that from a policymaker's perspective, it is not only interesting to know whether cycles move in the same direction but also whether output gaps are close to each other in absolute terms. More precisely, a large spread between two countries' cycles, even when the cycles move in the same direction, would call for a dissimilar answer by the central bank, as the country with the higher output gap would prefer a larger interest rate step. If the cycles of two countries have not only co-moved, but their gap has also narrowed due to the introduction of the euro, we would expect to see the EMU dummy impact negatively on this measure, indicating that membership in a currency union decreases the spread between cycles. Indeed, this seems to be the case according to our results. The magnitude of the impact amounts to nearly 0.4, i.e. EMU membership decreases the spread in output gaps among member states by 0.4 percentage points. The coefficients of the remaining variables have the expected sign. Higher bilateral trade reduces the spread between output gaps, while a more asynchronous fiscal policy seems to widen this spread. Repeating this exercise using the Baxter-King filter shows that the negative relationship between CU membership and cyclical spreads is robust.[21]

5.2 Robustness Analysis

In this section, we want to confront our baseline results with several robustness checks of the time and country sample, several independent variables and the estimation technique. The starting point is the model displayed in column (4) of Table 4, which includes all control variables and which is based on the adapted index of Cerqueira & Martins (2009) calculated by using the HP filter. The corresponding results are reported in Tables 5 and 6.

In column (1) of Table 5, we employ a different measure for trade integration, which we introduced in section 3 (see equation (5)). Obviously, normalizing bilateral trade by the sum of total trade does not alter the coefficient estimates considerably. The results point to a slightly lower impact of bilateral trade on business cycle correlation, which translates into a higher coefficient estimate of the EMU dummy.

[21] For brevity reasons, we do not show the results in the table, though they are available upon request.

Table 5: Robustness Tests

Robustness checks: Variable / Model	Bilateral Trade (1)	Excluding Crisis (2)	EMU Countries (3)	ERM (4)	GFA (5)	Bond Spreads (6)	RE-IV (7)
EMU	0.183***	0.171***	0.321***		0.114**	0.174***	0.107**
	(0.052)	(0.060)	(0.102)		(0.056)	(0.051)	(0.052)
Bilateral Trade	0.050***	0.066***	0.047*	0.078***	0.085***	0.071***	0.067***
	(0.014)	(0.017)	(0.026)	(0.016)	(0.016)	(0.016)	(0.018)
Industrial Specialization	0.025	0.053	-0.157	0.076	-0.009	0.075	0.117
	(0.084)	(0.092)	(0.126)	(0.089)	(0.086)	(0.096)	(0.087)
Fiscal Policy	-0.021***	-0.014**	-0.027***	-0.021***	-0.019***	-0.020***	-0.029***
	(0.006)	(0.007)	(0.010)	(0.006)	(0.006)	(0.007)	(0.005)
EU	0.054	0.048	-0.052	-0.040	-0.022	-0.082	0.138**
	(0.053)	(0.055)	(0.114)	(0.065)	(0.058)	(0.060)	(0.057)
ERM				0.179***			
				(0.059)			
Financial integration[1,2]					0.148***	-0.033***	
					(0.040)	(0.011)	
L1.Correl$_{HP}^{Cycle}$	0.098***	0.108***	0.064**	0.088***	0.083***	0.072***	
	(0.015)	(0.015)	(0.028)	(0.015)	(0.015)	(0.017)	
Constant	1.454***	0.826***	1.493***	1.681***	1.555***	1.801***	1.658***
	(0.120)	(0.189)	(0.256)	(0.136)	(0.133)	(0.138)	(0.148)
Time-FE χ^2_{17}	233.85***	165.24***	99.70***	250.17***	216.21***	155.48***	
Arellano-Bond test AR(1)	-13.633	-12.863	-8.201	-13.557	-13.523	-12.362	
Arellano-Bond test AR(2)	-0.519	-1.623	-0.779	-0.763	-0.882	-0.742	
Hansen test statistic	343.880	341.658	121.339	336.617	343.431	335.058	
Hansen p-value	0.668	0.389	0.959	0.172	0.973	0.550	
N	5673	4621	2216	5673	5673	4565	5674

Two-tailed significance levels: $*$: 10% $**$: 5% $***$: 1%. Standard Errors are reported in parenthesis. Out-of-sample instruments included: logdistance, logpopulation, common-border dummy. In-sample instruments: up to 9 lags. *Models:* (1) Alternative variable for "Bilateral Trade" (relative to total trade), (2) excluding the financial crisis (time sample restricted to 1993–2008), (3) including EMU countries only (i.e. EMU dummy captures time / within dimension only), (4) ERM dummy variable instead of EMU (including both ERM I and ERM II) (5 and 6) including two different measures for 'Financial Integration' as further explanatory variables (7) Random effects estimator, where bilateral trade, EMU and EU are instrumented by logdistance, logpopulation, common-border dummy, first time lag of EU and EMU.

In column (2), we exclude the period of the financial and debt crisis in Europe, i.e. we restrict our sample to the years 1993 to 2008. This alteration has nearly no effect on coefficient estimates except the one for fiscal policy, which becomes slightly lower. This is in line with our observation made in the previous section, namely that the relevance of fiscal policy for the synchronization of business cycles has increased since the onset of the financial and debt crises due to diverging consolidation efforts in Europe.

In column (3), we exclude all countries that were not EMU members in any of the observed years. In this case, the coefficient estimate of the EMU dummy is predominantly based on the variability over time ("within" variation) and hence allows us to identify whether CU entry increases business cycle correlation across member states over time. Restricting the sample to EMU countries even increases the EMU dummy, while leaving the other coefficient estimates unchanged. From this result we are able to conclude that the significantly positive effect of a common monetary policy observed in our baseline model is not solely driven by the cross-section variability, i.e. ot only reflects the difference in synchronization across members and nonmembers of a currency union.

In column (4), we widen the definition of the EMU dummy by additionally allowing all country pairs that were at the same time members of the Exchange Rate Mechanism I or II(ERM) to take on values of 1. Although the currency of these countries was allowed to float by a certain margin, their monetary policies were restricted, as their currencies were tied to the euro.[22] Interestingly, the size of the coefficient estimate of the ERM dummy is nearly equal to the one of the EMU dummy. This result is a strong indication that the EMU variable adequately measures the impact of common monetary policy shocks.

Still, one might argue that the effect captured by the EMU variable arises due to the increased degree of financial integration. This development of course might be related to EMU, as the establishment of a common currency area might have enhanced cross-border capital flows. Therefore, in columns (5) and (6), we try to capture the effect of financial integration on business cycle synchronization. Basically, there a two channels through which increased financial integration might play a role. First, higher capital mobility is associated with faster cross-country spillovers and therefore might lead to a higher degree

[22] The ERM I encompassed fixed currency exchange rate margins, while exchange rates were variable within those margins. Before the introduction of the euro, exchange rates were based on the European Currency Unit (ECU), whose value was determined as a weighted average of participating currencies. The magnitude of allowed fluctuations around the central rate was altered several times to accommodate speculation against some currencies, but never exceeded ±15%. ERM II replaced ERM I in 1999, and allowed a similar maximum float within a range of ±15%. In practice, however, many countries voluntarily decided to keep the range much smaller within one or two percent (e.g. Slovenia, Denmark). Moreover, participation in ERM II for at least two years is a necessary precondition to joining the euro area. Therefore, the ERM dummy captures common monetary policy across countries in a much broader definition than the EMU dummy, which only captures country pairs where both countries have already adopted the euro.

of business cycle synchronization. On the other hand, more developed financial markets can provide a significant source of insurance against asymmetric shocks. When countries are insured against idiosyncratic shocks, they can afford to specialize more strongly (see Kalemli-Ozcan et al. 2005), which in turn should have a negative impact on business cycle synchronization. Hence, the relationship is theoretically ambiguous. Yet, the task of measuring financial integration is afflicted with several difficulties mainly arising due to data limitations. In particular, there is no information on cross-border capital flows on a bilateral basis for the time period considered. Therefore, we resort to an indicator developed by Lane & Milesi-Ferretti (2007), who have collected data of external assets and liabilities for individual countries. Based on these data, we construct a bilateral measure of financial integration, which is calculated as the sum of two countries' external (foreign) assets and liabilities as a share of the sum of their GDPs.[23] A high value of this quantity-based measure indicates that both countries' financial markets are likely to be well integrated. To account for possible endogeneity, financial integration is being instrumented. Estimation results show that country pairs with highly integrated financial markets experience a higher co-movement of business cycles. The relationship is highly significant, and the inclusion of this additional variable has nearly no impact on the other regression coefficients. Only, the impact of the EMU variable is slightly alleviated but still remains positive and significant.

In column (6), we employ a further measure of financial integration as proposed by the European Central Bank (2013), which belongs to the group of priced-based indicators. In particular, to reflect the degree of financial segmentation in Europe, we calculate absolute differences of two countries' sovereign bond yields for ten-year maturities.[24] Especially since the onset of the financial crises, the dispersion between sovereign bond yields have risen markedly across euro area countries, reflecting both diverging country fundamentals but also flight-to-liquidity effects. The latter arise due to safe-haven flows (most notably to Germany), which have depressed bond yields of nondistressed countries and increased the ones of distressed countries. As a result, a spread arises between equally rated bonds that exhibit differences in liquidity, which reflects the process of financial disintegration especially since 2008. In column (6), we report the estimation results, which underline the outcome obtained by using the previous indicator. Obviously, increased financial disintegration seems to dampen business cycle synchronization across country-pairs. Again,

[23] In particular, the measure is defined as $FI^1 = log(\frac{A_{i,t}+L_{i,t}+A_{j,t}+L_{j,t}}{GDP_{i,t}+GDP_{j,t}})$, where $A_{i,t}$ and $L_{i,t}$ represent a country i's total external assets and liabilities in year t. Assets and liabilities include portfolio equity, foreign direct investment, debt and financial derivatives. The database can be downloaded from `http://www.philiplane.org/EWN.html`.

[24] Data are obtained from the database of the European Commission (`http://ec.europa.eu/economy_finance/ameco/user/serie/SelectSerie.cfm`).

the inclusion of this variable has no effect on other coefficient estimates.[25] Based on our robustness tests, we are therefore able to conclude that the EMU effect on business cycle synchronization does not arise due to the increased degree of financial integration, but is likely to capture the impact of a common monetary policy.

In the last column of Table 5 we check whether a different estimation technique alters our basic findings. We consider the static version of our baseline model and decide to use the random effects IV estimator, as suggested by the Hausman test. Trade is instrumented by the distance between country pairs (in logs), the population of a country (in logs) to capture the size of the economy and finally a common border dummy. The variables EMU and EU are instrumented by their one-period lags. The first stage results confirm the high explanatory power of the instruments.[26] The results outlined in column (6) show that the coefficient estimates remain stable and that the significantly positive impact of EMU membership on business cycle synchronization cannot be rejected.

We also checked whether the inclusion of distance between two countries alters the results of the EMU dummy. This point is linked to Baxter & Kouparitsas (2005), who find that a common currency is not a robust determinant of business cycle synchronization, as the inclusion of gravity variables renders the impact of CUs insignificant. Yet, according to our results,[27] the impact of common currencies is unaffected by the inclusion of distance. The only difference to our baseline model is that bilateral trade becomes insignificant, most likely due to multicollinearity issues related to the high correlation of trade intensity and distance.

Finally, we want to check whether our results are driven by an individual country. For this purpose, we re-estimate our baseline model excluding one country at a time, which corresponds to excluding 26 (of 351) country pairs from the sample. In Table 6, we report coefficient estimates of the EMU dummy for all 27 alterations, showing that the exclusion of any country changes neither the sign nor the significance of the EMU coefficient.

In view of our baseline results and the various robustness checks, we therefore conclude that the implementation of a common monetary policy has a positive and significant effect on business cycle synchronization, with the common monetary policy effect likely to increase cyclical correlations in the range of 0.11 to 0.32.

[25] At this point however, we want to stress that there is one important caveat in using bond spreads to measure financial integration. Price-based indicators can only reflect financial integration properly once risk premia are controlled for. As increasing bond spreads were also related to increasing risk premia the employed measure is likely to overestimate market segmentation.

[26] Estimations are performed using the STATA routine *xtivreg*. The results of the fixed effects model, first stage results and the corresponding test statistics are available from the authors upon request.

[27] For brevity reasons we do not display the estimation output, the results however are available upon request.

Table 6: Sensitivity of EMU Coefficient to Country Exclusion

Excluded Country	EMU	Excluded Country	EMU
Austria	0.140***	Ireland	0.211***
	(0.061)		(0.058)
Belgium	0.146***	Italy	0.160***
	(0.059)		(0.059)
Bulgaria	0.171***	Lithuania	0.190***
	(0.055)		(0.054)
Cyprus	0.182***	Latvia	0.184***
	(0.057)		(0.057)
Czech Republic	0.197***	Luxembourg	0.115**
	(0.057)		(0.059)
Germany	0.149***	Malta	0.121**
	(0.059)		(0.056)
Denmark	0.181***	Netherlands	0.178***
	(0.053)		(0.056)
Estonia	0.154***	Poland	0.155***
	(0.054)		(0.058)
Spain	0.153***	Portugal	0.192***
	(0.056)		(0.059)
Finland	0.138***	Romania	0.157***
	(0.055)		(0.056)
France	0.145***	Sweden	0.166***
	(0.062)		(0.054)
Great Britain	0.158***	Slovakia	0.143***
	(0.056)		(0.057)
Greece	0.240***	Slovenia	0.194***
	(0.055)		(0.055)
Hungary	0.153***		
	(0.057)		

Notes: One-tailed significance levels: $*$: 10% $**$: 5% $***$: 1%. Standard errors are reported in parentheses. Regression based on baseline specification (column 4 of table 4).

6 Discussion and Conclusion

As outlined by Frankel & Rose (1998, p. 1010), *"Countries that enter a currency union are likely to experience dramatically different business cycles than before. In part this will necessarily reflect the adoption of a common monetary policy; but it will also be a result of closer international trade with the other members of the union."* While their seminal paper shows that closer trade links lead to more synchronized business cycles, we are able to show empirically that the adoption of a common monetary policy per se leads to higher

BC synchronization. Hence, along these lines, this paper adds the second piece of the puzzle on the endogeneity of the OCA criteria.

By applying novel synchronization measures in a panel setting, we are able to overcome the short time-series data issue and analyze the impact of a common monetary policy on BC synchronization. While our results confirm the strong trade effect on BC synchronization found by Frankel & Rose (1998), even in its exact magnitude, we find an additional effect stemming from common monetary policy of roughly half the size of trade integration. Thus, the estimated coefficient is not only statistically, but also economically significant. Our findings are robust to a large set of different model specifications controlling for financial integration, differences in national fiscal policies and industrial specialization patterns. Remarkably, our results further indicate that EMU membership has not only led to a higher co-movement in business cycles but has also decreased the spread in output gaps among member states, which is also a novel and important insight for policymakers in central banks.

As a consequence, we are able to conclude that increased bilateral trade is not the only channel through which a common currency is connected to more homogeneous business cycles. In fact, economies in EMU seem to exhibit more synchronized cycles because of common monetary shocks in the euro area above and beyond the commonly cited trade integration effect. The lack of country-specific monetary policy shocks seem to smooth business cycles. Thus, the frequently mentioned costs of entering a CU, i.e. losing an important policy instrument to smooth business cycles by relinquishing autonomous monetary policy, are probably much lower than previously thought. On this view, our results substantially strengthen the conclusion by Frankel & Rose (1998), i.e. a country is more likely to satisfy the criteria for entry into a currency union ex post rather than ex ante. Remarkably, however, this reasoning is even verified when controlling for the effect of increased trade linkages implied by entering a currency union.

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7 Appendix

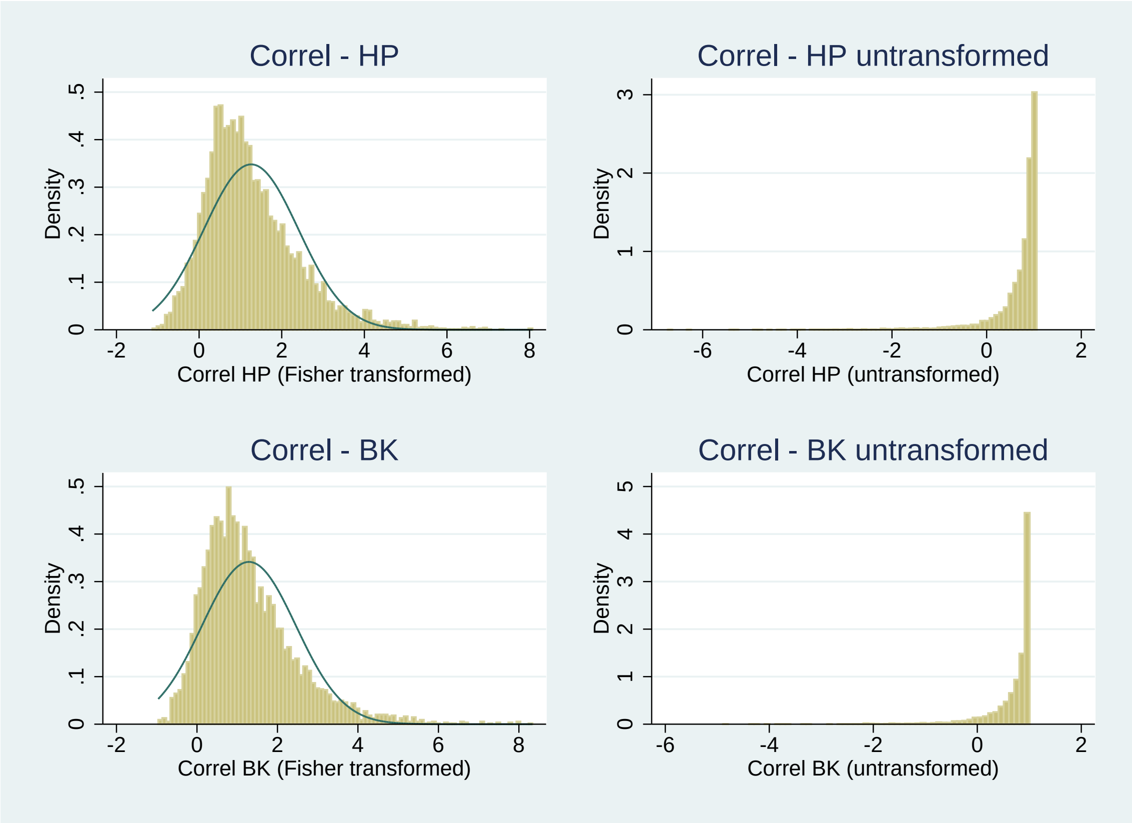

Figure 2: Kernel Density Plots of Dependent Variables

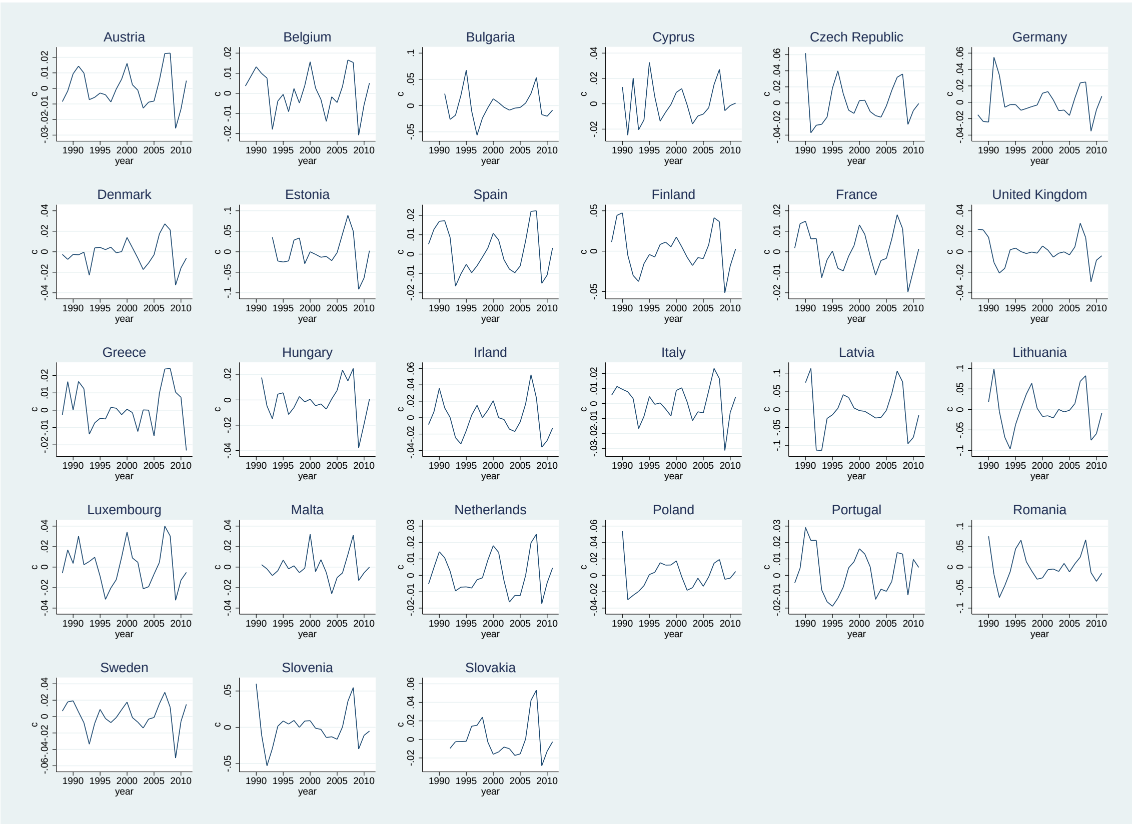

Figure 3: Cyclical Components: HP-Filter (Hodrick & Prescott 1997)

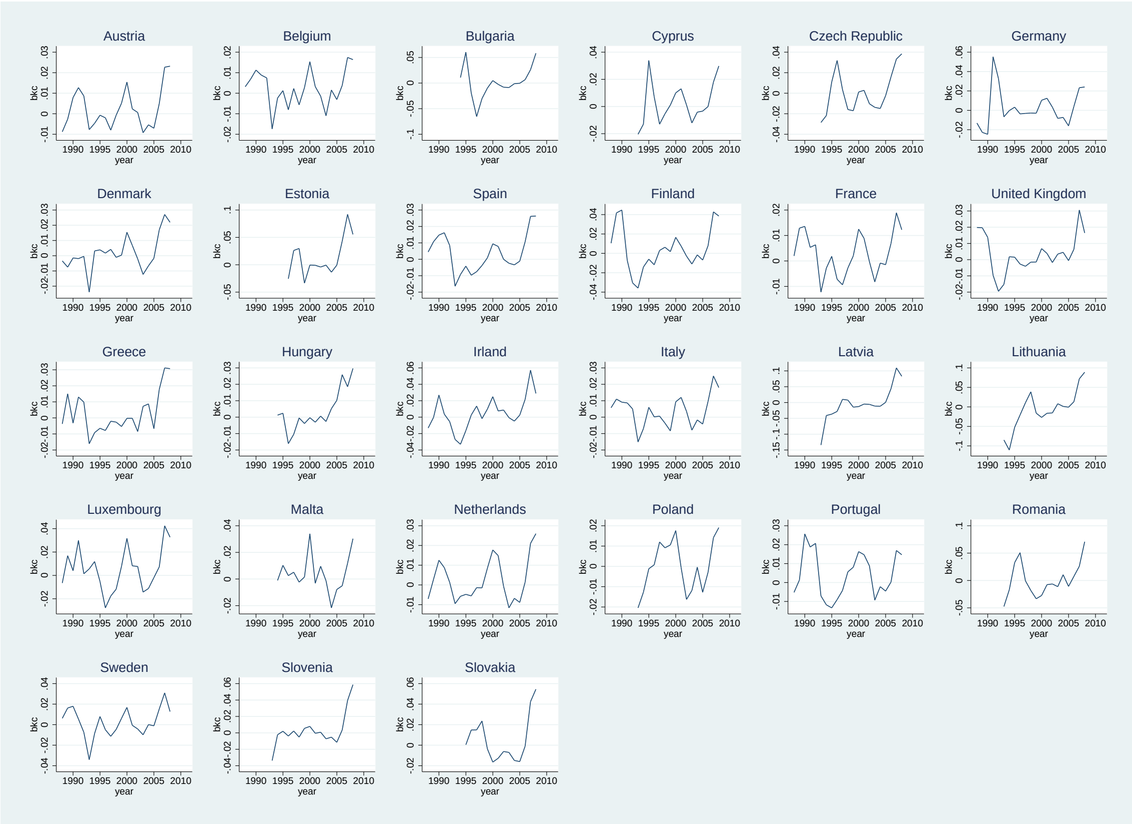

Figure 4: Cyclical Components: BK-Filter (Baxter & King 1999)

Table 7: Descriptive Statistics

Variable		Mean	Std. Dev.	Min	Max	Observations
$\text{Correl}_{HP}^{Cycle}$	overall	1.260	1.146	-1.130	7.985	N = 6668
	between		0.335	0.344	2.338	n = 351
	within		1.097	-1.522	7.255	$\bar{T}$ = 18.997
$\text{Correl}_{BK}^{Cycle}$	overall	1.292	1.168	-0.962	8.332	N = 5417
	between		0.340	0.475	2.442	n = 351
	within		1.118	-1.269	7.904	$\bar{T}$ = 15.433
$\text{Spread}_{HP}^{Cycle}$	overall	1.541	1.745	0.000	14.867	N = 6669
	between		0.906	0.288	3.973	n = 351
	within		1.492	-2.351	14.672	T = 19
EMU	overall	0.160	0.366	0.000	1.000	N = 6669
	between		0.256	0.000	0.684	n = 351
	within		0.262	-0.525	1.107	T = 19
ERM	overall	0.322	0.467	0.000	1.000	N = 6669
	between		0.357	0.000	1.000	n = 351
	within		0.302	-0.468	1.006	T = 19
Bilateral Trade[1,*]	overall	-6.562	1.657	-12.612	-2.199	N = 6350
	between		1.586	-10.120	-2.498	n = 351
	within		0.488	-10.247	-3.257	$\bar{T}$ = 18.091
Bilateral Trade[2,**]	overall	-6.127	1.669	-12.698	-1.995	N = 6350
	between		1.628	-10.116	-2.707	n = 351
	within		0.414	-9.821	-2.977	$\bar{T}$ = 18.091
Industrial Specialization	overall	0.492	0.217	0.042	1.314	N = 6106
	between		0.191	0.123	1.040	n = 351
	within		0.106	-0.093	0.892	$\bar{T}$ = 17.396
Fiscal Policy	overall	3.693	3.149	0.002	32.067	N = 6019
	between		1.577	1.043	9.459	n = 351
	within		2.723	-3.659	30.890	$\bar{T}$ = 17.148
EU	overall	0.560	0.496	0.000	1.000	N = 6669
	between		0.269	0.263	1.000	n = 351
	within		0.417	-0.335	1.296	T = 19
Financial Integration[1]	overall	1.249	0.983	-0.827	5.350	N = 6669
	between		0.888	-0.204	5.003	n = 351
	within		0.425	-0.175	2.255	T = 19
Financial Integration[2]	overall	1.610	2.127	0.000	16.910	N = 4713
	between		1.113	0.101	4.426	n = 351
	within		1.810	-2.750	14.196	$\bar{T}$ = 13.427
Distance, in logs	overall	7.093	0.649	4.007	8.236	N = 6669
	between		0.650	4.007	8.236	n = 351
	within		0.000	7.093	7.093	T = 19
Common border	overall	0.194	0.503	0.000	2.000	N = 6669
	between		0.504	0.000	2.000	n = 351
	within		0.000	0.194	0.194	T = 19
Population, in logs	overall	16.009	1.417	12.825	18.229	N = 6669
	between		1.419	12.888	18.223	n = 351
	within		0.038	15.871	16.129	T = 19

* measured in % of GDP, in logs; ** measured in % of total trade, in logs.

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