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The Eurozone Convergence through Crises and Structural Changes

by

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Abstract

In light of several economic and financial crises and institutional changes experienced by the Eurozone countries, we examine whether the adoption of the euro led to business cycle synchronization or fostered convergence of growth rates. Controlling for reverse causality, we conduct multiple endogenous break tests and find that while output growth was synchronized for some countries, convergence occurred in a nonlinear way for others: (i) convergence was not triggered by adoption of the euro but by international or idiosyncratic shocks; (ii) in several countries convergence started long before the introduction of the euro, accelerated during the 1990s and continued since then, reflecting persistent influence of the core countries; (iii) convergence has been prevalent among the non-Eurozone economies in our sample.

Key words: Convergence, business cycle synchronization, euro, crises, structural breaks
JEL Classification number: E3, F4, F6

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Introduction

The 2008 U.S. financial crisis that caused a global recession also pushed many European countries, in particular those in the periphery, into a stubborn recession from which many are still reeling. But more important, the severity of the crisis is seen as an indictment of the European Monetary Union for failing to generate convergence to its members. Convergence, in real or nominal terms is crucial for the conduct of the single monetary policy where members do not have recourse to fiscal transfer mechanisms nor currency adjustments. In this paper we examine whether the process towards monetary union and the common experience of a single currency led to convergence of output growth for member countries or if convergence was the result of idiosyncratic policies, reforms or global shocks.

The skepticism about the promises of the European Monetary Union goes back to its conception, which occurred among intense controversy about whether the Eurozone was ready to become a currency union since it was not satisfying the original criteria of the Optimal Currency Area[1]. Subsequently, studies of currency unions introduced by Rose (2000, 2004) and Frankel and Rose (1997, 2001) showed a significantly deepened trade integration among member countries of a monetary union. The implication is that the Optimal Currency Area criteria can be satisfied after unification resulting in trade integration and income convergence.

The endogenous business cycle synchronization argument draws on two linkages: currency union increases trade and increased trade leads to synchronization of business cycles. For the second linkage, two opposite views prevail. The European Commission's view was that higher trade integration would trigger synchronization of business cycles. Krugman (1993) cautioned against this view by arguing that higher trade integration leads to specialization, regional concentration of activity and hence to asymmetric

[1] Mundell (1961), McKinnon (1963) and Kenen (1969) argued that if a geographical region exhibits high capital and labor mobility, openness, price and wage flexibility, a risk sharing system such as fiscal transfer mechanism, and financial integration, then it is an Optimal Currency Area where it would be optimal to create a currency union instead of keeping floating exchange rates.

shocks, which reduces synchronization of economies. Empirical results are not uniform though several support the argument of increased synchronization through trade. However, these results are challenged by the view that business cycles in Europe are affected by many factors, which are mostly ignored except the trade channel (Rose, 2008). Exclusion of such factors, such as capital flows or higher sectoral specialization which decrease output comovements, can lead to potentially biased results.

Another inadequacy of most existing studies is that they are panel analyses examining output co-movements while controlling for currency union impact. Among these, few also address the endogeneity issue with trade by using methods with instrumental variables. However, convergence is a process that takes place over an extended period, disrupted by crises and recessions, speeding up during recovery and booms. Panel analyses are neither sufficiently long enough nor suitable to capture such non-linear processes.

Our paper investigates output growth convergence in the European Monetary Union from an angle different than the existing literature. We exploit the time series characteristics of several economies in the European Union, and examine if output growth in the peripheral countries converged towards that of the core countries. Such convergence can be captured directly through the change in the parameter associated with growth in the core countries: the higher the value of the parameter, the greater the synchronization between the core and the periphery.

We use the data of representative peripheral nations from the Eurozone and from non-Eurozone as a control group. In our model each peripheral country's output growth is affected by its past performance, its neighboring countries, and the center economies of the Eurozone and the rest of the world. By directly estimating the impact of "center" output on "peripheral" output and controlling for endogenous breaks, we are able to detect any possible change that may occur in the estimates not only between the center and the peripherals but also within and outside the Eurozone.

The impact of a single currency on the peripheral's economy may be observed either after the formation of the currency union or even before the date of the adoption of the single currency as the economies adjust to brace for the currency union. This impact may show up in one of the following patterns: (i) an endogenous break around the time of the euro adoption and further integration of the center and peripherals' outputs after 1999; (ii) no abrupt break around the time of the euro adoption but instead country-specific breaks throughout the sample, and a gradual and nonlinear integration of the center and peripheral's output growth, reflecting a deliberate economic integration policy initiated by the European countries for over half a century. Our methodology is, therefore, able to establish whether one of these two patterns has characterized convergence in peripheral countries and if so how the model parameters have evolved over time due to structural breaks. We define convergence as an increasing correlation over time between output growth rates of the center and the periphery and as business cycle synchronization a significant and constant correlation between the center and the periphery. Clearly, convergence as we define it can also be labeled as increased synchronization.

Our analysis contributes to the literature in two ways. First, we control for multiple endogenous changes in the convergence parameter and analyze if and how the estimated breakpoints are associated with known shocks, including the creation of the single currency. However, due to bilateral trade links the output of the trade partners may enter into a country's output model as an endogenous regressor, which may lead to biased estimates of the parameters and of the break dates. Therefore, our second contribution to the literature is to control for the endogeneity problem by adopting a methodology that allows instrumenting the output of the trade partners. This is an innovative way of tackling the endogeneity problems in the business cycle synchronization literature. Our approach is consistent with the case of continuous changes in the convergence parameter and is superior in terms of break location accuracy to the alternative time-varying parameter-based approach as we explain below. Our study reaches several interesting findings. First, we identify multiple endogenous break dates that correspond to major shocks in international arena and to idiosyncratic shocks affecting individual countries, suggesting that the convergence process has been nonlinear for a number of countries. Therefore, the evidence we uncover does not support the view that the switch to euro would generate a closer integration of the markets and thus automatically synchronize growth rates between the peripherals and the core countries. Moreover, since the adoption date of the common currency is not detected in the data as an endogenous break in the convergence process, our result indicates that studies imposing the euro's launch date as an exogenous break may lead to biased conclusions. Second, convergence occurred in most of the cases, and it happened as a result of a long process, which accelerated after the 1990s for most countries. Our analysis thus supports the view that the creation of the single currency spurred convergence among several countries in the region well before the introduction of the euro, and continued after that. Third, the non-euro countries' output growth is also significantly impacted by the center. This is not surprising among countries with significant trade ties, as demonstrated in the literature.

Literature survey

Business cycle synchronization and growth convergence were major concerns prior to euro's existence because it imposed a one-size-fits-all monetary policy on Eurozone economies with different sizes and structures. The argument that higher trade volume generated by a single currency would enhance business cycle synchronization is not theoretically well grounded. Standard Ricardian or Hecksher-Ohlin models suggest that industry specialization increases trade but as Krugman (1993) shows, greater industry specialization by country can lead to more asymmetric responses to industry-specific shocks and hence to business cycle divergence. However, if the shocks are common demand shocks or most trades are intra-industry related, business cycles may become *more* similar across countries as trade increases (Frankel and Rose, 1998, Shin and Wang, 2003). Thus, the effect of trade on business cycle synchronization may go in either direction.

Empirical literature is also far from reaching a consensus. On the nexus of trade and currency union, the initial findings of significant relation have been questioned by later research. In an earlier study, Frankel and Rose (1996) indicate that joining a currency union greatly increases trade. The initial estimates of the impact of a common currency on trade suggested a currency union would triple trade between nations sharing the same money (Rose, 2000). Later studies found these estimates to be much smaller (Glick and Rose, 2002, Barr et al., 2003; Micco et al., 2003; Bun and Klaassen, 2007).

Although evidence is more robust on the relation between trade volume and business cycle synchronization, there are notable exceptions. By using thirty years of data for twenty industrialized countries, Frankel and Rose (1998) show empirically that nations with closer trade links tend to have more tightly correlated business cycles. The meta-analysis of Rose (2008) shows that several studies support a significantly positive linkage between trade and business cycles. Yet, a small number find an insignificant one. Among these, Alesina, Barro and Tenreyo (2002) conclude that currency unions do raise trade, but do not generally have a significant impact on output co-movement. Barro and Tenreyro (2007) find that although currency unions enhance trade, they actually *decrease* the co-movement of output.

In addition to these linkages, most analysts agree that a common currency such as the euro may impact the business cycles through other channels. For instance, if currency union formation leads to a sharp increase in foreign borrowing (Eichengreen and Hausmann , 1999), peripheral nations' growth may increase much faster as capital flows from "center" countries to "peripheral" nations with higher returns. This process, which actually occurred in smaller Eurozone countries in the early years of the euro's existence, would tend to lower, rather than raise output co-movements.

Empirical literature on EU convergence and business cycle synchronization also reaches troubling results. Although poorer European countries' growth rates were catching up in the second half of the 1990s, since the financial crisis the failure of the much discussed convergence is frequently documented. Study after study documented not only a decline in convergence but even divergence among some European countries and the new members (Crespo et al 2008, Kaitilia 2014, Lee and Mercurelli, 2014, Benczes and Szent-Ivany, 2015), or showed that convergence among the core is reached at the cost of divergence with the periphery (Ferroni and Klaus, 2015). More recently, Miles and Vijverberg (2016) examine changes in output synchronization since the creation of the euro and find that out of eight eurozone countries investigated, only one country has increased synchronization with the core while three countries have decreased synchronization.

Finally, there are concerns about the techniques adopted in the literature to measure the business cycle synchronization. In the early cross-country literature, some studies use the correlation of either stock market returns or the cyclical components of output levels across countries as a way to evaluate the international integration (Eichengreen, 2001). Later studies estimate the correlations with moving windows over different time periods with instrumental variables (Barro and Tenreyro, 2007; Rose, 2008).

Mink, Jacobs and de Haan (2012) criticize such use of correlations. They argue that a perfectly correlated output gap may still imply a large discrepancy in synchronization because of large differences in cyclical amplitudes. They show that the alternative measures output gap "synchronicity" (similarity of signs) and "similarity" (similarity of output gaps) do not necessarily coincide with the traditional measures of output gap correlation.

However, all the existing studies in the literature, including this study examine the comovements of output gap measures with respect to a reference over a period where convergence was expected to occur. With the exception of Miles et al., none of them examines the dynamic adjustment process towards this state of nature. Unlike the previous studies, we propose to evaluate the evolution of the business cycle from an endogenous multiple structural change perspective. Our study emphasizes the interaction between the breaks in the convergence parameter and the synchronization of the growth rates among the Eurozone economies. The estimated breakpoints in the convergence parameter provide useful information for understanding the internal or external forces that have driven the convergence.

Our approach is in line with Sander and Kleimeier (2004) who test the convergence in the Eurozone retail banking, another strand of the literature examining convergence. The study examines the change of pass-through of interest rates to test if the monetary transmission process in Eurozone is uniform between 1993 and 2002. For this, the authors search for a unique endogenous break; they find that the structural breaks in the transmission mechanism do not coincide with the introduction of euro; rather, the breaks occurred much earlier. Marotta (2009) extends their analysis by allowing for multiple structural breaks; the pattern of the structural break dates indicates that national banking systems adjust slowly to the new monetary regime. He thus cautions the association of structural changes to the introduction of euro.

Model and Methodology

We estimate the following model for each peripheral country in our sample (Austria, Belgium, Denmark, Finland, Greece, Ireland, Netherlands, Portugal, Spain and Sweden):

$$\Delta y_t = \alpha + \beta\, \Delta y_t^{FG} + \delta\, \Delta y_t^{N} + \gamma\, \Delta y_t^{US} + \sum_i \omega_i \Delta y_{t-i} + \varepsilon_t \tag{1}$$

where Δy_t , Δy_t^{N}, Δy_t^{FG} and Δy_t^{US} stand for change in the industrial production (IP) of the peripheral country P, change in the weighted average IP of the three largest trade partners of the country P, change in the weighted average IP of the core Germany-France and change in the U.S. IP. Lagged dependent variables are introduced until the first insignificant lag is reached in order to remove possible residual autocorrelation.

Equation (1) shows that growth in a given peripheral country is affected by growth in the core, its neighbors and the United States, and its own past performance. Our goal is to assess β, the convergence parameter, which measures the extent to which a peripheral country's business cycle co-moves with that of the core. We estimate this parameter controlling for the impact of the business cycle in the trading partners as well as the rest of the world, represented by the US economy. Let $(\beta_1, \delta_1, \gamma_1)$ and $(\beta_2, \delta_2, \gamma_2)$ represent values of the parameters (β, δ, γ) before and after a structural break, say the adoption of the Euro, respectively. We hypothesize that there is convergence of the periphery country with the core if $\beta_1 \geq 0$, and $\beta_2 > 0$, and $\beta_2 > \beta_1$ and there is synchronization if $\beta_1 = \beta_2$.

However, this is a very restrictive and unrealistic condition since countries have followed a long process of conversion towards monetary unification and did not switch their economies overnight when they become members of the union. To capture this process, we adopt the less restrictive hypothesis of gradual convergence, which will be reflected in several estimates due to multiple breaks. Reforms or policies in a peripheral country or the ECB's macroeconomic initiatives may affect the convergence parameter anytime during the period under consideration. If such shocks exist, ignoring them by assuming a single break at the introduction of the euro would bias the results and give misleading estimates, especially for β. Since the top trade partners of the periphery often include countries that are part of the EU but not the monetary union (such as UK and Denmark), a significant and positive δ also indicates a larger regional integration that is likely due to the existence of the single market, which increased the volume of trade among all members .

Δy_t^N is potentially an endogenous regressor. This arises when two countries are among the largest three partners of each other. Then the IP growth of a country A depends on growth in one of the major trade partner B through Δy_t^N , while country A appears simultaneously as a major trade partner for country B in another equation similar to (1). Such endogeneities can cause simultaneity biases that are known to yield inconsistent estimates. As we will show below, in our panel of peripheral countries an important subset consists of mutual major partners and therefore subject to the endogeneity problem. The remaining countries are not concerned by this problem.

Since our goal is to check for the instability of the parameters in equation (1) in order to assess whether the core Germany-France has impacted the growth of the peripheral within the monetary union, we implement Bai and Perron's (1998) (BP) methodology of estimation with endogenous structural breakpoints for those countries for which all the regressors are exogenous. If there is a Eurozone influence, we expect to find instability in parameters in the form of an increasing value of β over our sample period. To account for possible endogenous structural breaks, we rewrite model (1) as:

$$\Delta y_t = \sum_{j=1}^{m+1} \theta_j{}' \ z_t 1_{t \in I_j} + v_t \tag{2}$$

where $z_t = (1, \Delta y_t^{FG}, \Delta y_t^{N}, \Delta y_t^{US})'$, $\theta = (\alpha, \beta, \delta, \gamma, \omega_1, \omega_2, ...)'$, *m* is the unknown number of breaks, I_j (*j*=1,…,*m+1*) the segment between break dates t_{j-1} and t_j and $1_{t \in I_j}$ an indicator function such that $1_{t \in I_j}$ = 1 for $t_{j-1} < t \le t_j$ and 0 elsewhere ($t_0 = 1$ and $t_{m+1} = T$). The vector of coefficients θ_j characterizes the effects of the exogenous variables z_t on Δy_t over the *j*th segment (*j*=1,…,*m*+1).[2] A break date $\hat{t}_j$ is estimated as $\arg \min_{1 \le t_j \le T} SSR(t_j)$, where *SSR* is the sum of squared residuals over the sample. Using a sequential F-test procedure, *m* is determined when the null of *m* breaks against *m+1* breaks is not rejected. When no break occurs (*m*=0), the coefficient vector θ is estimated over the full sample and model (2) collapses to model (1).

However, when the model contains endogenous regressors, the OLS-based BP method cannot be used. For these countries we follow Hall et al (2012) (HHB) who extend BP's methodology to linear models with endogenous regressors within an instrumental variables (IV) framework. The idea is to test for changes in the parameters of model (1) using 2SLS method, which provides consistent estimators of parameters in the presence of endogenous regressors. To the best of our knowledge, no previous study has tested endogenous changes in business cycle synchronization by controlling for endogeneity of regressors, although models describing output growth interdependence often involve such feedback effects. [3]

In a recent article, Perron and Yamamoto (2015) (PY) show that, in the presence of endogenous regressors, it is preferable to estimate the break dates of the structural model simply by using the OLS-

[2] The partial structural change model introduced by BP also includes a second vector of independent variables whose parameters are not subject to shifts. We allow all parameters to adjust to the breaks, given that they may prove to be insignificantly different between subsequent segments if they are genuinely constant over the period.

[3] Even though this model has some resemblance to a spatial panel model, this approach would be inadequate to achieve our goal since a spatial panel model with endogenous breaks has not yet been developed. Furthermore, in a spatial panel model, the "spatial correlation" parameter before the "spatial weight" matrix is the same for all countries involved. Thus, given the same "spatial correlation" parameter, the neighboring countries may impact other peripheral countries differently only through the pre-assigned spatial weights. Our model in this paper does not impose this kind of pre-assigned restrictions.

based BP method. For several methodological reasons it is preferable to estimate these break dates by controlling for the endogeneity problem in the regressors by using the HHB method. [4] The latter offers the additional advantage of allowing an explicit breakpoint analysis both for the endogenous regressors and for the structural model, whereas PY's approach does not test for the presence of breaks in the endogenous regressors.

For each country, we select appropriate instruments for the endogenous regressor Δy_t^N such that they are correlated with the latter and orthogonal to ε_t . We consider the following as instruments: the lagged variables Δy_{t-i}^N , i=1,2…, the actual and lagged values of the changes in the IPs of Germany Δy_{t-i}^G , of France Δy_{t-i}^F and of USA Δy_{t-i}^{US} , i=0,1… This IV regression equation is called the Reduced form (RF), whereas model (1) is labeled as the Structural form (SF).

The test procedure suggested by HHB can be summarized by the following two steps. First, test for the presence of structural breaks in the RF using the sequential BP methodology. Second, (a) If there is no break in the RF, test for the presence of structural breaks in the SF using the BP methodology. (b) If the null of no break is rejected in the RF, test for the presence of structural breaks in the SF using the BP methodology over each sub-period delimited by the estimated breaks in the RF. Then, check whether the breaks identified in the RF coincide with additional ones in the SF by performing Wald tests over the appropriate sub-samples evidenced by the estimated breaks in the SF.

[4] The main argument in PY is that in the IV methodology of HHB, the generated regressors, which are obtained as a projection of the original regressors on the space spanned by the instruments, have less quadratic variation than the original regressors involved by OLS, making the break point estimates less accurate. Conditional on OLS-based estimation of breaks, PY propose an IV regression to obtain the estimates of the parameters within each subsample. However, HHB argue that neglected endogeneity in regressors may still continue to contaminate OLS-based estimations of breaks in the structural equation and yield inconsistent OLS estimators of parameters in the sub-periods. Moreover 2SLS is still needed in PY’s approach, making it not different than HHB’s method regarding parameter estimation (Chen, 2015). To check that a change in OLS-based parameters across segments is not due to a change in the simultaneity bias, PY show how an estimation of this change in bias can be computed ex-post and compared with the estimated change in parameters. However, this presents two difficulties: (i) for non-negligible values of this ratio, any conclusion that the estimated change in parameters is due to a change in the true parameters can be dubious; (ii) The need for performing these comparisons systematically for each regressor and each break point can make the whole procedure impractical.

A popular approach in describing the change in the convergence parameter is to assume that this parameter evolves following a dynamic stochastic model. This parameter equation, together with the peripheral country's growth equation, can be represented as a state-space model and estimated using Kalman filtering. However, with this approach, when the change in the parameter is mild and gradual, it may not be easy to locate in the sample the exact beginning date of the change by visual inspection. Yet, identifying such points in time is important to understanding the specific historical events causing the instability in the parameter (e.g., macroeconomic policies, political changes, international shocks). The endogenous nature of convergence towards synchronization necessitates a data-driven identification of the characteristic and timing of the specific events that trigger synchronization. This is in line with Frankel and Rose (1998), who state that cyclical correlation between economies depends on trade integration, which itself is affected by EU policies. Rather than a model with continuously time-varying parameters, we thus need to specify a model where movements in the output growth correlation can be addressed through discrete changes at unknown dates resulting from idiosyncratic or global shocks. This can be achieved by using our methodology, which is superior to alternative methods in terms of breakpoint location accuracy.

It is worth noting that such discrete changes do not mean that convergence occurs with abrupt changes. They, instead, indicate that among the very large number of small breaks that shape the parameter dynamics over a given segment of time, none can be diagnosed as being significant, provided that any significant change initiates a new segment. Thus, the convergence parameter is *statistically* constant over a segment but it changes from one segment to another. Since the (unknown) number of breakpoints is not set a priori but is endogenously determined by using a sequential test procedure, the identification of all the significant breakpoints ensures parameter stability over each sub-period.

Data

Our data consist of monthly industrial production index (IP) from OECD to represent each country's output, spanning from January 1975 to June 2013. Since our focus is on the peripheral countries, our investigation will focus mainly on representative small economies within Eurozone: Austria, Belgium, Finland, Greece, Ireland, Netherlands, Portugal, and Spain. We exclude Italy since it is the third largest economy in the Eurozone, and too large to be considered as a peripheral country. The four periphery countries in our sample are Portugal, Ireland, Greece and Spain, traditionally designated as PIGS (PIIGS with Italy). We added two non-Eurozone small economies, Denmark and Sweden, both part of the EU but not the Eurozone, as a control group to see how the economies that do not adhere to a common currency behaved over the same sample. Denmark decided not to adopt the euro to preserve their economic sovereignty. Sweden rejected the adoption of the euro in a referendum in 2003. Unlike the Swedish krona, which has been floating, Denmark's Krone was shadowing the DM and then the euro from the beginning. Note that Austria, Belgium, Finland, Ireland, Netherlands, Portugal and Spain joined Eurozone in January 1999; Greece was admitted to join in January 2001.

We calculate y_t^{FG} as a weighted average of the IP of Germany and IP of France. The weights are obtained by computing each period's total trade first, which is the volume of trade denominated in U.S. dollars. Then, for each time period, we obtain the weight by calculating the ratio of each country's trade to the total trade. We define y_t^{N} as the output of the neighbor of the peripheral country. The notion of "neighbor" here implies economic neighbors rather than geographical neighbors. We calculate this output measure by taking the trade-weighted average of the peripheral neighbors' outputs. We compute the trade weights as follows using the OECD data. For each country, we obtained the US-dollar values of imports and exports with respect to its trading partners from 1990 to 2011 and calculated the average of both imports and exports over these years. The three highest average total trades, excluding Germany, France and non-EU countries, are selected as the dominant trade partners of each country. These three average total trades are summed together to calculate the weight for each of the three countries.[5]

[5] The trading neighbor weights of each country are stated as the following. Austria=0.451*Italy+0.297*Switzerland+0.251*Hungary; Belgium=0.524*Netherlands+ 0.293*UK+ 0.182*Italy; Denmark=0.477*Sweden +0.295*UK +0.228*Netherlands; Finland=0.486*Sweden+ 0.275*UK+0.239*Netherlands; France-Germany= 0.349*Belgium +0.333*Italy +0.318*Netherlands; Greece=0.584*Italy +0.212*UK +0.204*Netherlands; Ireland=0.649*UK+0.244*Belgium+ 0.107*Netherlands;

Empirical Results

Inspection of the trading partners for the peripheral economies in footnote 1 shows that out of ten countries, four do not have an endogeneity problem in the independent variables. For these four countries, Austria, Finland, Greece and Ireland, we apply the BP methodology of endogenous multiple structural breaks (Table 1). For the remaining six countries for which endogeneity was an issue (Belgium, Denmark, Netherlands, Portugal, Spain and Sweden) we use the HHB instrumental variables methodology to test for breaks (Table 3). For sake of simplicity, we will refer to the countries in Table 1 as group 1 and to those in Table 3 as group 2. Tables 2 and 4 present the estimation results for each group, respectively.

In both Table 1 and Table 3, the test is a sequential test of the null of *k* breaks against the alternative of *k+1* breaks, *k*=0,1,2... The null is rejected in favor of the alternative if the minimal value of the overall sum of squared residuals for the *k+1* break model is smaller than the one for the *k* break model. The double maximum statistics $UD\max F_T$ and $WD\max F_T$ test the null of no structural change against an unknown number of breaks up to some upper-bound *M*. They are referred to as double maximum because under the alternative hypothesis the maximization involves the choice for the number of breaks $m=1,...,M$ and for the selection of the highest *F* statistic for a given value of *m*. However, BP note that the *supF* test for *m* breaks may be of low power if *m* is large and suggest to apply appropriate weights to these individual statistics. This gives the weighted double maximum statistic $WD\max F_T$,while $UD\max F_T$ is its unweighted version. We conduct these tests by allowing for regime-specific error distributions; doing so we account for the possible error heterogeneity across sub-periods.

Netherlands=0.457*Belgium +0.361*UK+0.182*Italy; Portugal=0.658*Spain +0.177* UK +0.165*Italy; Spain=0.416*Italy +0.322*UK +0.262*Portugal; Sweden=0.386*Denmark +0.363*UK+0.251*Netherlands; Switzerland=0.509*Italy +0.282*UK +0.209*Netherlands;

Endogenous break dates

According to whether the regressor Δy_t^N is exogenous or endogenous, BP and HHB multiple structural break tests have shown the presence of two breakpoints for Belgium, a single breakpoint for Austria, Finland, Greece, Portugal, Spain and Sweden and no break for Ireland, Denmark and Netherlands (Tables 1 and 3). h double maximum statistics indicate that among group 1 countries there is a parameter shiftat the 5% level for Austria, Finland and Greece while the sequential *F* test rejects the null of 0 break in favor of 1 break but fails to reject the null of 1 break in favor of 2 breaks. The break points for these three countries are March 1991, October 1989, and April 2007, respectively. These three dates are associated with major shocks in the global arena as well as idiosyncratic shocks to individual countries. For Ireland none of the test results supports the presence of a break. it is important to note that any structural change in parameter indicates a major economic event such as a reform, a crisis or an economic policy. , However, all economic events do not necessarily lead to a change in the parameter of the model. If, indeed, a major event affects the dependent and the independent variables simultaneously, the slope coefficients may not change significantly.

The break date of March 1991 in *Austria* is a typical domestic shock-induced breakpoint in that it corresponds to the beginning of the recession that followed a high growth period. In the early 1990s, recession reduced the growth rate from 4.2 percent in 1990 down to 0.4 percent in 1994. The break of 1989 October in *Finland* also characterizes a recession period that started in 1990. Throughout the 1980s, financial deregulation consisting of removal of controls on bank borrowing and foreign borrowing, together with a fixed currency but appreciating real exchange rate culminated in a financial bubble at the end of the decade. The collapse of the Soviet Union in 1991 is widely considered to be the last drop in the cup that triggered the Finnish recession. It reduced the volume of trade by 2/3 and forced the monetary authority to devalue. Explosion of foreign currency denominated debt as a result of devaluation and new emphasis on bank regulation precipitated the burst of the bubble. In *Greece*, the last country of

Group 1, the breakpoint occurs at the beginning of the 2007-2008 financial crisis when the economy entered its protracted recession.

In the second group, *Belgium* exhibits one break in August 1982 that affects the reduced form. In the structural form, we find one break at the left of this break date and another at the right of this break date, which gives the two break dates September 1980 and February 1995. However, the Wald test result indicates that the structural form does not support the break date of the reduced form. For Belgium, therefore, we retain the two break dates in 1980 and 1995. The breakpoint of September 1980 corresponds to the beginning of the first of a series of recessions in the country triggered by the oil crises of the 1970s. The recession of the 1980-1982 was a severe one with large unemployment and deficit, which reached 13% of GDP. The crisis led to a major restructuring of the economy when activity moved from Wallonia to Flanders. Belgium is an advanced industrial economy with a relatively stable average quarterly growth rate of 0.53 percent between 1980 and 2014. The notable exception of an all-time high of 15.80 percent in the first quarter of 1995 was fueled by a sharp rise in gross fixed investment and government spending. The second breakpoint of February 1995 is consistent with this idiosyncratic high growth rate, which followed a severe recession in 1992-93.

In *Denmark* and *Netherlands* we detect one break in the reduced form (Table 3, top panel), respectively. However, evidence does not support any break in the structural form conditional to the breaks in the reduced form (Table 3, bottom panel). Therefore, for these two countries no break point is retained. By contrast, for *Portugal and Spain*, no break appears in the reduced form (top panel). We, therefore, test for multiple breaks in the structural form over the whole sample. Test results reject the null and find breaks in January 2004 in Portugal and June 1992 in Spain (bottom panel). The Portuguese economy grew at a healthy average rate of 3 percent in the run up to the adoption of the euro. However, during the 2000s it went through a recession and anemic growth, the lowest in Europe, while its competitors were expanding. Portugal's breakpoint of January 2004 corresponds to the end of the recession but the beginning of a dismal growth period. This is attributed to a decline in productivity caused by misallocation of capital flows to unproductive subsectors of nontradables (Reis, 2013). Spain's break coincides with many domestic and international factors. These include the European financial crisis in the summer of 1992, three times devaluation of its currency between 1992 and 1993 and an economic downturn in the early 1990s that lasted till mid-1990s, after which it returned to a vigorous growth spurred by strong foreign investment before entering the European Monetary Union in 1999.

In our sample, *Sweden* is the only country where a break of February 1982 found in the reduced form is significantly supported by the Wald test. No additional break is detected in this equation on either side of the break date obtained from the reduced form. This break date falls in the middle of a major restructuring effort that Sweden took up during the 1980s to escape the 1970s' slow growth rates. During this period, which was also called Sweden's "lost decade", output growth decreased from one of the highest in Europe to one of the lowest. During the early 1980s policy makers launched a series of deregulations in many sectors including government and financial markets, devaluing the Swedish krona by 24 percent and cutting back the deficit. By 1984 these measures started slowly to invigorate economic activity through domestic investment, although growth did not reach its healthy levels until the early 1990s.

Our results lend further support to the argument that exogenously imposed break dates can be quite misleading and lead to biased estimates. It is striking that adoption of a single currency in 1999 did not create any instability in the regression results. In all economies that exhibit instability in the parameters of equation (1), the break points correspond to dates when the countries experienced an economic recession, slowdown or a demand shock. By inspecting the changes in the parameter estimates, we can examine whether there was any integration and if so, when it started.

Parameter estimates

From the estimated parameters across regimes reported in Tables 2 and 4, it can be seen that our parameter of interest, β, is positive for each peripheral country (PIGS) and becomes highly significant after the break. This suggests a higher synchronization with the center and a clear significant impact of

Germany and France continuously or over certain periods in time. Our results do not support the notion that the switch to euro in 1999 triggered a closer integration between the peripherals and the core in euro zone. Convergence did occur in several cases but it was a result of a long process, which accelerated after the 1990s for most countries.

Among the first group of countries with a break the estimates of the convergence parameter β are positive but not significant in the pre-break period (Table 3, top panel, first 3 columns). But for Austria and Greece, the estimates of β become larger and highly significant after their respective breaks, suggesting strong convergence to the core. It is interesting to note that at no point in the data Finland exhibits convergence to the core. In Ireland the convergence estimate is positive and significant at the 10 percent confidence level throughout the sample period.

The estimates of the parameters δ and γ show the measure of integration with the European trade partners' economies and the US economy, respectively. Inspection of δ shows a higher level of business cycle synchronization after the break points in Austria and Finland with their European partners, but a similar synchronization happens before the breakpoint in Greece. The United States acquires significance as a trade partner only in Finland after the break point. The Irish economy is also affected not only by its European neighbors but especially by the US.

Among the second group of countries, the estimate of the convergence parameter is similarly insignificant before the respective breaks of Portugal and Spain (Table 4). It becomes highly significant in Portugal but surprisingly remains insignificant in Spain. Both countries' economic activities are highly correlated with those of their neighbors (these are UK, Italy, Portugal for Spain and UK, Italy Spain for Portugal). In Belgium, the only economy with two breaks, the estimate of the convergence parameters weakly significant in the first subsample, during the period of accelerating European integration. The convergence seems to be adversely affected by the series of recessions that the country went through during the 1980s, but it picks up speed after the second break, which also corresponds to the start of the Schengen agreements in 1995. Except during the first subsample, the Belgian economic activity is not synchronized with its neighbors (Netherlands, UK, Italy). The last country in this group, the Netherlands, has no break and a strong and positive convergence parameter indicating a well synchronized economic activity with the center, and a strong and significant synchronization with the US but none with its neighbors (Belgium, UK, Italy).

When we consider the results of the periphery, PIGS, two points need to be stressed. First, the strong convergence of Greece after 2007 is simply due to the country's recession coinciding with the slowdown in the core and therefore, not a reliable indication that the economy strongly converged to the core.. Second, Ireland's convergence parameter estimate is weakly significant, while Spain never shows any significant convergence. These facts suggest that results do not support unequivocally the hypothesis of gradual synchronization for three out of four periphery economies.

It is interesting to note that economic activity in the two non-euro countries, Sweden and Denmark, does not display a particularly different behavior compared to that of the euro countries. Denmark, which does not have a break, exhibits a strong synchronization with the center but an insignificant one with its neighbors (Sweden, UK, and Netherlands) and none with the US. After the break, Sweden's output correlation with the center becomes significant and positive and its economic ties with the US strengthen considerably. This result suggests that even if they do not belong to the Eurozone, the Netherlands from the beginning and Sweden after its break became candidates to the European currency union.

To sum up, although the estimates of the convergence parameters are in general positive throughout the sample period, they mostly become significant after the country-specific break dates. Among the economies in our sample, five of the eight Eurozone countries (Austria, Belgium, Netherlands, Portugal and Ireland) show significant convergence to the core. Among six Eurozone countries that experienced breaks, four (Austria, Belgium, Greece and Portugal) experience stronger convergence in the last part of their sample period. Two Eurozone countries that do not have breaks, Netherlands and Ireland, also show convergence to with the US growth rates as well as the core. The two non-Eurozone countries, Sweden and Denmark show both strong convergence to the core and to the US in the case of the latter. Thus, Austria, Belgium, and Portugal indicate a rising integration with the core after their respective endogenous breaks; Netherlands and Ireland had strong integration with the core from the beginning, while, the two non-Eurozone economies, Denmark was all along integrated with the core, while Sweden became integrated after the break.

Conclusion

In this study we examine the economic growth convergence characteristics of several peripheral countries that are currently economically struggling in the European Union. We analyze whether the process towards monetary union and the common experience of a single currency led to convergence of output growth for these member countries. To test for convergence, we use a methodology that accounts for endogenous structural breaks, nonlinearities and endogeneity problem in the regressors when reverse causality arises between a country and its trading partners. To our knowledge, this approach has never been applied before in this literature.

Our results indicate the following. First, none of the break dates detected endogenously by the data correspond to the date of the euro adoption. In all economies that exhibit instability in the parameters, breaks occurred when the countries experienced an economic recession, slowdown or a demand shock. Our results lend further support to the argument that exogenously imposed break dates can be quite misleading and lead to biased estimates.

Second, for peripheral countries such as Portugal, Ireland and Greece, the parameter of convergence is either positive at all times, or becomes increasingly significant after their respective breaks. The same pattern is also observed in the rest of the Eurozone countries in our sample. An exception to this result is the case of Finland and Spain, which become more integrated with their European trade partners or the US instead of the core.

Third, we also observe a strong integration with the core of the control group consisting of two countries outside the Eurozone. Our results thus suggest that the euro adoption date of 1999 is not the pivotal force in enhancing the integration between the peripherals and the core in Eurozone. Rather, the impact of the single currency has been gradual, and it started before the adoption of the euro in 1999, i.e., some countries were in a long convergence process in anticipation of their membership in the monetary union, which strengthened subsequently after they joined the union. But adoption of a single currency, likely also increased economic integration in the whole region, and led to similar patterns of economic growth in the nonmember countries of our sample.

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Table 1. Multiple break tests with exogenous regressors (Bai-Perron)

c	Austria	Finland	Greece	Ireland
$UD\max F_T$	23.51 (1)**	28.56 (1)**	32.04 (1)**	17.13 (3)
$WD\max F_T$	23.51 (1)**	28.56 (1)**	32.04 (1)**	22.22 (5)
$SupF_T(k+1/k)$	23.51 (1/0)** 5.74 (2/1)	28.56 (1/0)** 19.71 (2/1)	32.04 (1/0)** 22.33 (2/1)	11.48 (1/0)
$\hat{t}_{SF}$	1991.03	1989.10	2007.04	-

Notes: $\hat{t}_{SF}$ denotes the estimated break in the structural form. $UD\max F_T$, $WD\max F_T$ and $SupF_T(k+1/k)$ are the unweighted and weighted double maximum statistics reached at the indicated number of breaks and the sequential *F*-statistic for the null of *k* break(s) versus *k+1* break(s), respectively. ** represents 5% level of significance. Critical values at the 5% level for ($UD\max F_T$, $WD\max F_T$, $SupF_T(1/0)$, $SupF_T(2/1)$) are (20.30, 21.86, 20.8, 22.11) for Austria, (22.04, 23.81, 21.87, 24.17) for Finland, (23.87, 25.63, 23.70, 25.75) for Greece and (23.87, 25.63, 23.70) for Ireland. A trimming percentage of 15% is employed and error distributions are allowed to differ across breaks.

Table 2. Estimation of models with exogenous regressors (Bai-Perron)

	Austria	Finland	Greece	Ireland
	1979.02 - 1991.02 [145]	1975.02 -1989.09 [176]	1975.02 - 2007.03 [386]	1975.12 - 2013.06 [451]
α_1	0.005*** (3.41)	0.007*** (3.08)	0.003** (2.38)	0.01*** (5.04)
β_1	0.02 (0.14)	0.15 (1.07)	0.04 (0.40)	0.26* (1.72)
δ_1	0.10 (1.45)	0.20 (1.57)	0.15* (1.65)	0.28* (1.67)
γ_1	0.20 (1.02)	-0.30 (-1.13)	0.14 (0.74)	0.62** (2.33)
	1991.03 - 2013.06 [268]	*1989.10* - 2013.06 [285]	*2007.04* - 2013.06 [75]	
α_2	0.004*** (3.46)	0.001 (0.70)	-0.01*** (-3.39)	
β_2	0.40*** (4.34)	0.18 (1.54)	0.72*** (3.30)	
δ_2	0.15** (2.19)	0.33*** (2.84)	-0.18 (-0.68)	
γ_2	0.004 (0.02)	0.66*** (3.25)	-0.18 (-0.54)	
$\bar{R}^2$	0.30	0.20	0.31	0.29

Notes. Numbers in parentheses are t-values. Subscript i=1,2 attached to parameters denote the *i*'th sub-period. ***,** and * represent 1, 5 and 10% levels of significance, respectively. Dates in italics are the significant breakpoints at the 5% level as reported in Table 1.

Table 3. Multiple break tests with endogenous regressors (Hall-Han-Boldea)

	Belgium	Denmark	Netherlands	Portugal	Spain	Sweden
Break test in the IV- reduced form of Δy_t^{NP}						
$UD\max F_T$	25.23 (1)**	31.52 (1)**	20.07 (2)**	15.08 (3)	12.95 (1)	27.95 (1)**
$WD\max F_T$	25.23 (1)**	31.52 (1)**	23.60 (2)**	20.89 (5)	18.58 (5)	27.95 (1)**
$SupF_T(k+1/k)$	25.23 (1/0)** 12.11 (2/1)	31.52 (1/0)** 12.71 (2/1)	18.56 (1/0)** 11.45 (2/1)	9.20 (1/0) -	12.95 (1/0) -	27.95 (1/0)** 7.92 (2/1)
$\hat{t}_{RF}$	1982.08	1978.01	2005.11	-	-	*1982.02*
Break test in the structural form over the sub-periods conditional on $\hat{t}_{RF}$						
$UD\max F_T$ $WD\max F_T$	27.60 (1)** 27.60 (1)** ——— 44.64 (1)** 44.64 (1)**	(a) ——— 16.86 (1) 16.86 (1)	19.44 (1) 20.16 (3) ——— 20.94 (2) 29.00 (5)**	31.28 (1)** 31.28 (1)**	24.92 (1)** 24.92 (1)**	25.49 (2)** 30.53 (3)** ——— 22.08 (2)** 25.52 (2)**
$SupF_T$ $(k+1/k)$	27.60(1/0)** 9.52(2/1) ——— 44.64(1/0)** 7.42(2/1)	(a) ——— 16.86 (1/0) -	19.44 (1/0) - ——— 11.90 (1/0) -	31.28 (1/0)** 11.10 (2/1)	24.92 (1/0)** 12.96 (2/1)	13.98 (1/0) - ——— 19.35 (1/0) -
$\hat{t}_{SF}$	*1980.09* ——— *1995.02*	- ——— -	- ——— -	*2004.01*	*1992.06*	- ——— -
Wald test for the presence of the break $\hat{t}_{RF}$ *in the structural form*						
Wald p-value Sub-sample	0.07 1980.09-1995.01	0.75 1975.02-2013.06	0.06 1975.02-2013.06	-	-	0.0001 1975.02-2013.06

Notes. $\hat{t}_{RF}$ and $\hat{t}_{SF}$ denote the estimated breaks in the reduced and structural forms, respectively. $UD\max F_T$, $WD\max F_T$ and $SupF_T(k+1/k)$ are the unweighted and weighted double maximum statistics reached at the indicated number of breaks and the sequential *F*-statistic for the null of *k* break(s) versus *k+1* break(s), respectively. Numbers above (below) the bars in the second panel are the values of the mentioned test statistics over the sub-period at the left (at the right) of $\hat{t}_{RF}$. No bar means that the tests are run over the whole sample. ** represents 5% level of significance. Critical values at the 5% level for ($UD\max F_T, WD\max F_T, SupF_T(1/0)$, $SupF_T(2/1)$) in the case of the reduced form and of the structural form are (16.37, 17.83, 16.19, 18.11) and (20.30, 21.86, 20.08, 22.11) for Belgium, (23.87, 25.63, 23.70, 25.75) and (18.42, 19.96, 19.23) for Denmark, (16.37, 17.83, 16.19, 18.11) and (22.04, 23.81, 21.87) for Netherlands, (23.87, 25.63, 23.70) and (22.04, 23.81, 21.87, 24.17) for Portugal, (16.37, 17.83, 16.19) and (18.42, 19.96, 18.23, 19.91) for Spain and (20.30, 21.86, 20.08, 22.11) and (20.30, 21.86, 20.08) for Sweden, respectively. A trimming percentage of 15% is employed and error distributions are allowed to differ across breaks. (a) indicates that the breakpoint test cannot be performed because of a too short sub-sample (35 observations). Dates in italics are the significant breakpoints at the 5% level.

Table 4. Estimation of models with endogenous regressors (Hall-Han-Boldea)

	Belgium	Denmark	Netherlands	Portugal	Spain	Sweden
1st sub-period	1975.02 - 1980.08 [67]	1975.02 - 2013.06 [461]	1975.02 - 2013.06 [461]	1975.02 - 2003.12 [347]	1975.02 - 1992.05 [208]	1975.02 - 1982.01 [84]
α_1	-0.001 (-0.18)	-	-	0.01*** (4.12)	-	-
β_1	0.40* (1.72)	0.42*** (3.60)	0.35*** (3.60)	0.12 (1.07)	0.20 (0.92)	0.38 (1.10)
δ_1	-0.41 (-0.79)	0.14 (1.27)	-0.06 -(0.26)	0.38** (2.31)	-0.02 (-0.02)	-0.76 (-1.28)
γ_1	0.81** (2.13)	0.24 (1.16)	0.42*** (2.65)	-0.18 (-0.88)	0.15 (0.48)	0.79* (1.74)
2d sub-period	*1980.09* - 1995.01 [173]			*2004.01*- 2013.06 [114]	*1992.06* - 2013.06 [253]	*1982.02* - 2013.06 [377]
α_2	0.002 (1.10)			-0.005** (-2.44)	-	-
β_2	0.12 (1.00)			0.67*** (4.22)	-0.24 (-1.59)	0.21** (2.26)
δ_2	0.03 (0.32)			0.57** (1.99)	2.60*** (4.49)	0.19 (1.23)
γ_2	-0.09 (-0.38)			-0.19 (-0.74)	-0.11 (-0.68)	0.63*** (4.51)
3d sub-period	*1995.02* - 2013.06 [221]					
α_3	0.002** (2.10)					
β_3	0.54*** (5.30)					
δ_3	-0.17 (-0.96)					
γ_3	0.06 (0.40)					
$\overline{R}^2$	0.29	0.10	0.29	0.30	0.30	0.24

Notes. Numbers in parentheses are t-values. Subscript i=1,2,3 attached to parameters denote the i'th sub-period. ***,** and * represent 1, 5 and 10% levels of significance, respectively. When intercepts where insignificant they have been removed from the model before reestimation. Dates in italics are the significant breakpoints at the 5% level as reported in Table 3.