Journal of International Money and Finance 112 (2021) 102325

Contents lists available at ScienceDirect

Journal of International Money and Finance

journal homepage: www.elsevier.com/locate/jimf

The dynamics of core and periphery in the European monetary union: A new approach ☆

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ARTICLE INFO

Article history:
Available online 28 November 2020

JEL codes:
C5
E3
N1
F4

Keywords:
Euro
Core-periphery
Competition
Symmetry
Convergence

ABSTRACT

Despite numerous studies about core-periphery in monetary unions, few focus on their dynamics. This paper (i) presents new theory-based, continuous and dynamic measures of the probability of a country being classified as core or periphery; (ii) estimates the determinants of the changes in this probability over time and across countries; and (iii) uses the Phillips-Sul convergence panel framework to investigate the behaviour of core and periphery groups over time. Our main results indicate that the post-EMU decrease of the core-periphery gap that we document was mainly driven by the adoption of the euro and by increasing competition (lower mark-ups).

1. Introduction

The European Single currency (the euro) just commemorated its twentieth anniversary. Despite widespread recognition of the importance of issues related to the existence of core and periphery groups and the distance or imbalances between them (De Grauwe, 2018), the vast majority of the existing research uses a static approach. By static, we mean that existing studies identify which countries are members of the core and the periphery but they do not throw light on the how the probability of being classified as core or periphery evolves or changes over time. This is the main contribution of this paper.

☆ We would like to thank Tamim Bayoumi, Iain Begg, Dionysios Chionis, Paul De Grauwe, Jakob de Haan, Barry Eichengreen, Balazs Egert, Jarko Fidrmuc, Christopher Hartwell, Haekwon Lee, Andrew Martin, Rana Sajedi, Andre Sapir, Jan-Egbert Sturm, Paul Wachtel, Karl Whelan, and seminar participants at London School of Economics, International Monetary Fund, Harvard University, New York University, University College London, University of Orleans, Brunel University London, Bank of England, OECD, Polish National Bank, Czech National Bank, European Economic Association, American Economic Association, Royal Economic Society, Money Macro and Finance Group, ETH Conference on "The Euro Area at 20: Evaluating the Real Side of the Economic and Monetary Union" (Zurich), "Economic Challenges for the Future of Europe" Conference (Paris), Infiniti Conference (Sydney), CASE 25th Anniversary Conference (Warsaw), Monnet Conference on "The Economic Prospects for the EU" (Dusseldorf), and Applied Economics Association Meetings (Valencia) for valuable comments on previous versions of this paper. We are responsible for all remaining errors.
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https://doi.org/10.1016/j.jimonfin.2020.102325
0261-5606/

This paper makes three main contributions: (a) to put forward a new conceptual framework that aims at locating countries along a core-periphery *continuum* (that is, to go beyond the common binary treatment of either core or periphery), (b) to construct a dynamic theory-based measure of these distances so that one can assess the evolution of the core-periphery dynamics, and (c) to identify, based on the theory of endogenous Optimal Currency Areas (hereafter OCA), the main factors that explain the evolution of core and periphery before and after the establishment of the currency union. We study these issues in the context of the European Economic and Monetary Union (EMU).

The seminal paper in this debate is Bayoumi and Eichengreen (1993). This is one of the first to identify a core-periphery pattern in the run-up to the EMU. They forewarn that, if persistent, this core-periphery pattern would be detrimental to the development of the EMU. Using the canonical Aggregate Demand- Aggregate Supply (AD-AS) framework and 1963–1989 data, they estimate the degree of demand and supply shocks synchronization. Bayoumi and Eichengreen show that there is a core where supply-side shocks are highly correlated (Germany, France, Belgium, Netherlands and Denmark) and a periphery where this correlation is lower (Greece, Ireland, Italy, Portugal, Spain and UK). Bayoumi and Eichengreen (as well as the large literature that follows it) define symmetry as a result of high correlation among supply shocks, conditional on a zero long-run restriction on each country's demand shock (Blanchard and Quah, 1989). In this sense, the European Monetary System may have eliminated autonomous monetary policies as a source of idiosyncratic demand shocks, but national fiscal policies remain independent so the cross-country correlation in movements in demand may persist. In their paper, the core-periphery gap is thus identified only based on supply shocks' correlations as demand shocks did not allow to distinguish the core from the periphery.

In their model, Bayoumi and Eichengreen (1993) assume that positive demand shocks across countries raise prices with zero long-run output effects. Based on the AD-AS model, positive supply shocks are expected to lower prices and increase output permanently instead. However, Bayoumi and Eichengreen's (1993) model is exactly-identified because they chose to impose restrictions on demand but not on supply. Santos Silva and Tenreyro (2010) survey of the literature shows that this remains a ubiquitous choice in the literature.

One novelty of our approach is to model these long-run supply considerations as over-identifying restrictions. We impose this over-identifying restriction in a structural VAR framework (SVAR) such that the predicted effects of supply shocks on output are accounted for. We use this to derive new dynamic, continuous and theory-based measures of the probability of a country being classified as core or periphery.

To assess how satisfactory is this new measure, we examine the factors that can potentially explain its evolution over time and across countries. Drawing from the endogenous OCA theory (Frankel and Rose, 1998), our panel estimates suggest that euro membership and competition (more flexible product market regulations or smaller mark-ups) make countries more likely to be classified as 'core'. These results have important academic and policy implications as it remains unclear whether the effects of the euro and competition are closely related or independent (on the latter possibility see Engel and Rogers, 2004). We also find that trade openness and, to a lesser extent, foreign direct investment, help "import competition" and, in so doing, substitute for regulations: imports (but not exports) increase the probability of a country being classified as core. Our findings are robust to numerous changes in specification, samples, measurement, and estimators. Finally, we use our new measure to study convergence. Using the Phillips and Sul (2007) approach we identify a set of core countries (a "hard-core" with Austria, Belgium, Germany, and the Netherlands), an intermediary set (a "soft-core" of France, Italy, Denmark, Spain, and the UK), and an "extended periphery" (that we estimate is composed of Greece, Sweden, Finland, Ireland, Norway, Portugal and Switzerland). Our probability estimates decrease over time for core countries (broadly confirming the endogenous OCA predictions), remains high and constant for the periphery (corroborating Bayoumi and Eichengreen's early warnings), and varies substantially for the intermediate set of countries. Overall, we find the core-periphery gap in Europe decreases after the EMU.

The paper is organised as follows. Section 2 sets up the conceptual framework. Section 3 discusses estimation issues and introduces our new measure of the probability of a country being classified as core or periphery. Section 4 studies the behaviour of our new measure in both static and dynamic settings. Section 5 presents our results regarding its potential determinants over time and across countries, whereas Section 6 uses the Philips and Sul model (2007, 2009) to identify precise country groupings. Section 7 has further detailed discussion and interpretation of our findings, as well as their key policy implications. Section 8 concludes.

2. Core and periphery in currency unions

The main research question driving the OCA scholarship regards the costs and benefits of sharing a currency (Alesina and Barro, 2002). The main cost is the loss of monetary sovereignty, with the ensuing relinquishment of monetary and exchange rate autonomy. Benefits are mostly in terms of reduction of transaction costs and exchange rate uncertainty, and increasing price transparency, trade and competition.[1]

One way of framing the potential trade-offs in an optimum currency area (OCA) is suggested by De Grauwe and Mongelli (2005). They study interactions between symmetry, flexibility and openness. Particularly, they show there exists a minimum combination of these dimensions that countries must reach for a monetary union to generate positive net benefits (see also

[1] The literature is surveyed in De Haan et al. (2008), Santos Silva and Tenreyro, 2010 and Glick and Rose (2016).

Farhi and Werning, 2015). De Grauwe and Mongelli (2005) place the Eurozone (EU) within (to the outside) of the OCA-line suggesting those countries are (not yet) sufficiently integrated to generate efficiency gains that can compensate for the macroeconomic costs of the union. They also note the degree of openness and symmetry may change over time. Before the EMU, there was an intense debate about the extent to which a monetary union affects symmetry (Krugman, 1993). De Grauwe and Mongelli (2005) view is that in the EU specialisation will bring about less symmetry and thus countries would move downwards along the OCA plane.

There are at least two noteworthy recent developments in OCA theory. The original OCA formulation stressed labour mobility, product diversification and trade openness as key criteria and explored the possible endogeneity of currency unions (Frankel and Rose, 1998). Recent work called attention to the additional role of credibility shocks. If there are varying degrees of policy commitment (furthering time-inconsistency problems), countries with dissimilar credibility shocks should find it convenient to join a currency union (Chari et al., 2020).

A second recent and important strand highlights that, although OCA criteria are often thought of as independent, they should instead be considered jointly, e.g., by focusing on the interactions between openness and mobility (Farhi and Werning, 2015).

The optimality of a currency area is thus a function of the relative power (and of the relative "distance") between its members. If these differentials are large, it is common to speak of a core and a periphery. It is expected that core countries would be those more closely meeting the OCA criteria. Given its importance for OCA, it is not surprising there have been various attempts of classifying countries into the core and periphery sets. A basic way of distinguishing these methods is whether or not the core-periphery status is imposed *a priori*.

A statistical technique often used here is cluster analysis. This method helps to determine the overall number and the composition of different grouping (or clusters) given that the elements (countries or regions) which are assigned to a given cluster are more similar to each other than to those in other clusters according to pre-defined criteria. Cluster analysis is a broad statistical methodology that involves various ways of estimating these groups and the distance among them. These can be thought of as different algorithms differing in their criteria and efficiency in determining the composition of each cluster.

The European Union business cycle was first studies as a phenomenon by Artis in a series of papers in the late-1990s and early 2000s. Artis and Zhang (1997, 1999, 2001, 2002) investigate actual and prospective members of the EMU by applying clustering techniques to a set of variables suggested by the theory of Optimal Currency Areas. The OCA criteria they employ are the extent of synchronisation in business cycles (symmetry in output shocks), volatility in the real exchange rate, synchronisation in the real interest rate cycle, openness to trade, inflation convergence, and labour market flexibility. Their analysis reveals that the member countries may be divided into three groups: those belonging to the core (Germany, France, Austria, Belgium and the Netherlands), those part of a Northern periphery (Denmark, Ireland, the UK, Switzerland, Sweden, Norway and Finland) and those belonging to a Southern periphery (Spain, Italy, Portugal and Greece). Their method delivers a rather intuitive classification of countries and, more usefully, allows researchers to see how these classifications change for each of the OCA criteria and multiple combinations of them. It is important to note that there is relevant related work examining economic growth patterns instead by, among others, Crowley (2008) and Crowley et al. (2013).

There are other, more theory-based, approaches. Bayoumi and Eichengreen (1993) is a seminal piece in this regard. They put forward an approach stressing business cycle synchronisation embedded in a standard Aggregate Demand and Aggregate Supply framework. Using 1960 to 1989 data, they classify countries into core and periphery groups and famously warm that the gap between these groups may put at risk the very planning of the Single Currency. We discuss these results below in greater detail, but it is important to bring to the fore a related paper by these authors that produce an objective index to classify countries into core and non-core. Bayoumi and Eichengreen (1997) construct an "optimum-currency-area index for European countries." The crucial element of their approach is to identify the determinants of nominal exchange rate variability which, they argue, reflects OCA characteristics and support predictions of which countries pertain to each set. This is justified conceptually and empirically. Conceptually, they argue that OCA focuses on criteria that ultimately make exchange rates more stable and monetary unification less costly. In their model, bilateral exchange rate variability is a function of GDP, trade, economic structure dissimilarity, and a measure of output synchronisation. Using 1973 to 1992 data, they find these carry expected signs and are statistically significant so they use them to forecast variability in 1987, 1991 and 1995.

Their econometric analysis allows them to identify three groups: in the first "rapidly converging" group there are Germany, Austria, Belgium, the Netherlands, Ireland and Switzerland. The second group is one that has experienced little convergence and contains the United Kingdom, Denmark, Finland, Norway and France. The third group is a set of countries that are "gradually converging" to the EMU and includes Sweden, Italy, Greece, Portugal and Spain. The two most consequential results from this analysis, with the benefit of hindsight, are that (1) France is positioned in the set that diverges from the Maastricht criteria but (2) that overall "economic integration has thus increased countries' readiness for monetary integration" (Bayoumi and Eichengreen, 1997, p. 769).

A different estimation framework to analyse the issue of asymmetries in the EMU is Basse (2014) which uses cointegration and structural breaks to try to identify EMU core member countries. The OCA dimension he is most interested in is financial, in that he analyses changes in sovereign fixed-income credit risk and the sample includes only selected EMU members (Germany, France, Belgium, Austria, Finland, and the Netherlands.) His main finding raises, yet again, questions about the suitability of France as a core country.

One final noteworthy approach in terms of providing useful insights to understanding core and periphery in the EMU involves studies in which membership is assigned *a priori*. Arestis and Phelps (2016) perform an "endogeneity analysis" of output synchronization, differentiating output amplitude from concordance in cycles, using panel data (covering the period 1994 to 2013), for all members as well as for different country groups, including core, periphery, central and eastern European countries, northern European countries and the euro-candidate countries. The quantification of trade-related and direct spillover channels associated with monetary integration gives insights into the relative importance of direct and indirect synchronisation gains arising from EMU membership through trade. For all members, membership and trade appear to increase both amplitude and coincidence of their business cycles. The individual group analysis shows that core and northern countries have experienced much larger trade spillovers in terms of synchronisation than peripheral and Central and Eastern European countries. They suggest future research on the direct euro membership effects by trying to unpack the role of institutions (see also Campos et al., 2019) and changes in the transmission mechanism since the introduction of the single currency.

3. Implementation

The methodology introduced by Bayoumi and Eichengreen (1993) is an extension of the Blanchard and Quah (1989) procedure for decomposing permanent (supply) and temporary (demand) shocks. Consider a system where the true model is represented by an infinite moving average of a (vector) of variables, X_t, and shocks, ϵ_t. Using the lag operator L, a bivariate vector autoregression (VAR) featuring real GDP and its deflator can be written as an infinite moving average representation of demand and supply disturbances:

$$X_t = A_0\epsilon_t + A_1\epsilon_{t-1} + A_2\epsilon_{t-2} + A_3\epsilon_{t-3} + \cdots = \sum\nolimits_{i=0}^{\infty} L^i A_i \epsilon_t \tag{1}$$

where $X_t = [\Delta y_t, \Delta p_t]$ and the matrices A represent the impulse response functions of the shocks to the elements of X. It follows that

$$\begin{bmatrix} \Delta y_t \\ \Delta p_t \end{bmatrix} = \sum\nolimits_{i=0}^{\infty} L^i \begin{bmatrix} a_{11i} & a_{12i} \\ a_{21i} & a_{22i} \end{bmatrix} \begin{bmatrix} \epsilon_{dt} \\ \epsilon_{st} \end{bmatrix} \tag{2}$$

where y_t and p_t represent the logarithm of the adjusted output and prices and ϵ_t are *i.i.d.* disturbances, which identify supply and demand shocks (Ramey, 2016). For the i-th country, a_{11i} represents element a_{11}, in matrix A_i and so on.

This framework implies that supply shocks have permanent effects on output, while demand shocks have temporary effects. Both have permanent (opposite) effects on prices. The cumulative effect of demand shocks on the change in output must be zero:

$$\sum\nolimits_{i=0}^{\infty} a_{11i} = 0 \tag{3}$$

so it can be estimated using a VAR. Each element can be regressed on lagged values of all the elements of X. Using B to represent these estimated coefficients:

$$\begin{aligned} X_t &= B_1 X_{t-1} + B_2 X_{t-2} + \cdots + B_n X_{t-n} + e_t \\ &= (I - B(L))^{-1} e_t \\ &= (I + B(L) + B(L)^2 + \cdots) e_t \\ &= e_t + D_1 e_{t-1} + D_2 e_{t-2} + D_3 e_{t-3} \end{aligned} \tag{4}$$

where e_t represents the residuals from the VAR equations. To convert (4) into the model in (2) under (3), the residuals from the VAR, e_t, are transformed into demand and supply shocks using the standard relation between the VAR's residuals (e_t) and demand and supply shocks, i.e.$e_t = C\epsilon_t$. For each country, the exact identification of the C matrix requires four restrictions. Two are normalizations, which define the variance of the shocks ϵ_{dt} and ϵ_{st}. The third restriction is from assuming that demand and supply shocks are orthogonal to each other. The fourth that demand shocks have only temporary effects on output (equation (3)).

The canonical AD-AS model implies that demand shocks raise prices in both the short and long run, while supply shocks lower prices and increase demand permanently. To fully reflect the structure of the underlying theoretical model, we impose an additional over-identifying restriction in the VAR such that supply shocks have permanent effects on output.[2]

[2] In addition to trying to implement the AS-AD model more fully, another important reason we adopt the proposed over-identifying restriction is that inflation differentials are often considered a 'normal feature of currency unions' (see e.g., ECB, 2017), hence we impose no restrictions on the reaction of inflation to demand and supply shocks, respectively.

We pay particular attention to modelling the effect of permanent (supply) shocks on output, on top of the usual demand-side one. Since the proposed over-identifying restriction is sufficient to generate structural disturbances in line with AD-AS dynamics, any additional long-run restriction may be redundant in this setting.[3]

Theoretically, testing for the "symmetry" of shocks also reflects the idea that the distinction between permanent and temporary shocks matters when it comes to adjustment within a currency union. According to De Grauwe (2018), when shocks are permanent, the slope of the usual existing trade-off between flexibility and symmetry is likely to depend on the nature of the shocks. Particularly, when permanent shocks dominate this tradeoff is likely to be steeper. Conversely, when temporary shocks dominate, the tradeoff will be flatter. We extend this idea further and suggest that it is not only the nature of the shock that matters but also the direction and extent by which shocks are pushing an economy "out of sync."

We test for symmetry of permanent shocks in our model by imposing $\sum_{i=0}^{\infty} a_{12i} = \gamma$, where $\gamma > 0$. This assumption implies that demand in each country should respond qualitatively (sign) and quantitatively (size) in the same way to supply shocks. In terms of the structural VAR (SVAR):

$$\sum_{i=1}^{\infty} \begin{bmatrix} d_{11i} & d_{12i} \\ d_{21i} & d_{22i} \end{bmatrix} \begin{bmatrix} c_{11} & c_{12} \\ c_{21} & c_{22} \end{bmatrix} = \begin{bmatrix} 0 & \gamma \\ . & . \end{bmatrix} \tag{5}$$

In order to construct a test for the over-identifying restriction described above, we estimate a SVAR model that is fully consistent with Bayoumi and Eichengreen (1993). Differently from the literature that follows Bayoumi and Eichengreen, we bootstrap the original VAR residuals in a *i.i.d.* fashion and generate K = 10.000 data sets.[4] For each of the k-*th* samples, we proceed with structural analysis and test for the over-identifying restriction based on a LR-test. We record the number of rejections of the over-identifying restriction test at each bootstrap replication and calculate the number of rejections (*NORD*):

$$NORD_i = 100 \times \frac{\sum_{k=1}^{K} \left\{ NORD = 1 | -2(L_r - L_u) > \chi^2_{q-\left(\frac{n^2-n}{2}\right)} \right\}_{i,k}}{K} \tag{6}$$

where L_u and L_r are the maximized values of the (Gaussian) log-likelihood function of the unrestricted and restricted regressions, respectively. Under H_0, the LR statistic has an asymptotic distribution with degrees of freedom equal to the number of long-run restrictions (q) minus $(n^2 - n)/2$, where n is the VAR-dimension (in this case $n = 2$). We calculate $NORD_i$ for different values of γ (Appendix 4). Notice that we do not restrict γ to a fixed value *a priori.* Instead, we vary γ in the interval $[0.1, 2]$. We then chose the value of γ which minimizes the total number of rejections in the sample, and for the average of euro area countries, in particular, so that the number chosen on this basis is $\gamma = 1$ (see Appendix 4). Demand and supply shocks are then retrieved by bootstrap, in particular recalculating the VAR parameters (K = 10.000), identifying the SVAR and considering median values of structural disturbances under $\gamma = 1$.

Three major drawbacks from the existing empirical measures of symmetry in currency unions are that (a) they tend not to be continuous (a country is classified either as core or as periphery), (b) they are often imposed *a priori* and (c) they tend to be time-invariant. One may add that most of these measures are not theory-based. The approach we develop above generates a new type of measure that relaxes most of these main constraints, except the time-invariant aspect. If countries can move from periphery to core, static classifications become very unhelpful.

One way of thinking about how we can construct a dynamic or time-variant measure of the probability of a country being classified as core or as periphery is in terms of parameter constancy. Let T be larger than before (55 years, i.e. 1960–2015) and τ denote the width of a sub-sample (25 years) or window and define the rolling sample metrics:

$$NORD_{\mathrm{ti}}(\tau) = \frac{1}{\tau - 1} \sum_{j=0}^{\tau-1} NORD_{(t-j)_i}(\tau) \tag{7}$$

The windows are rolled through the sample one observation at a time so that the procedure returns $T - \tau + 1$ rolling estimates of each parameter. Using a fixed 25-year window, we obtain a value representing the end-of-period number of rejections for each year, out of the number of the bootstrap replications. This number thus represents the count of the number of times the over-identifying restriction is rejected for each country for the selected time period. The window is then rolled one observation at a time and new estimates are obtained for the whole time-series.

4. Understanding NORD

What is the intuition behind this measure? Our modelling of supply considerations as over-identifying restrictions produces a new theory-based continuous measure. Our measure is the percentage of times the underlying structural macro AD-AS-model is rejected. Not only that: the more frequently this complete set of restrictions is rejected, the less symmetrical (hence, 'more peripheral') a country is said to be compared to the average EU member state in the euro area. The lower is

[3] Among those who argue that demand shocks can and do have long-term consequences for GDP growth are Ball (2014) and Fatás and Summers (2016).

[4] One concern is the possible effects from any regime changes between 1960 and 2015. Using a dummy saturation approach, first proposed by Hendry et al. (2008), we attempt to detect model selection problems and correct the original series for possible regime changes. In the light of the above, we identify three regimes for both GDP and inflation, broadly consistent across countries: 1960:I-1969:I; 1984:I-1992:I; and 2008:I-2015:I.

this frequency, that is, the less often the underlying structural macro-model is rejected for a given country in a given year, the higher – in our interpretation – is the likelihood of the country being classified as ‘core’, again compared to the average EU member states in the euro area.

To better understand our approach, it is important to stress that how different countries respond to macroeconomic shocks may be analysed in a number of ways. To start with, a useful distinction could be drawn between asymmetric and symmetric shocks.

First, our approach tests whether supply shocks would quantitatively and qualitatively affect countries’ GDP growth in similar ways. As Caporale (1993) and Patterson and Amati (1998) observe, there is predominantly symmetric shocks in sectors that manufacture standardised products with few trade barriers. Instead, in sectors where the presence of trade barriers and the degree of shock symmetry seem to have an inverse association asymmetry should decrease over time with the advancing of the EMU.

Secondly, as with shocks, a more helpful distinction can be drawn between short-term and long-term effects (Patterson and Amati, 1998). A monetary policy shock may, for example, have differing effects in different countries because they are at different stages in the economic cycle. However, these may also be due to long-term differences in financial, labour market structure, etc. (see Boeri and Garibaldi, 2006). The relative importance of cyclical versus structural factors in causing asymmetric responses played a large part in the debate on the EMU (see also Patterson and Amati, 1998). By focusing on long-run restrictions, one can argue that, while cyclical misalignment can be seen as a normal feature of a currency union, long-run misalignments reflect more long-term structural differences. Cohen and Wyplosz (1989) for instance, linked asymmetries to transitory shocks and symmetries to permanent shocks, finding that in core countries like France and Germany symmetry prevailed. In this sense, our test about the similarity of shocks across countries does provide an indirect test for core-periphery.

Fig. 1 shows the original Bayoumi and Eichengreen (1993) estimates for the period before the EMU and how do they compare to the ones post-EMU generated using the approach described above. Specifically, the figure compares estimates from pre-Maastricht based on Bayoumi and Eichengreen (1993) sample, covering the period 1963–1988, with our equivalent estimates for the period 1989–2015. Bayoumi and Eichengreen argue that pre-EMU there is a core (Germany, France, Belgium, Netherlands and Denmark) where supply shocks are highly correlated and a periphery (Greece, Ireland, Italy, Portugal, Spain, and the UK) where synchronisation is lower. Importantly, these two groups can be detected visually in terms of supply shocks correlation but not from the perspective of demand correlations (Fig. 1).

Comparing the pre- with the post-EMU static estimates suggests that it may have weakened the original divide. It is also clear this now cannot be fully grasped through visual inspection. A quantitative measure appears to have become necessary.

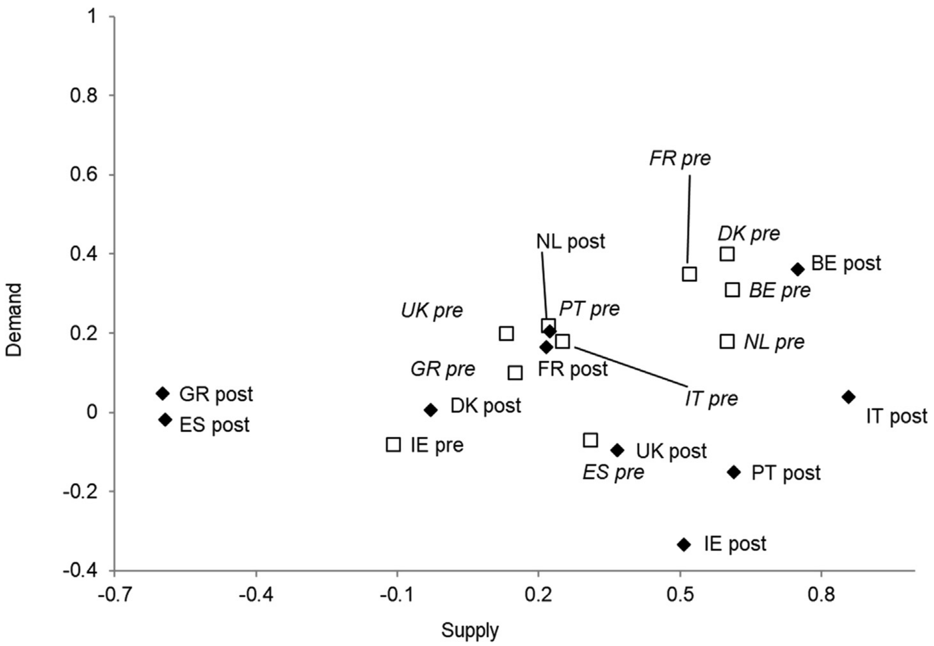

Fig. 1. Pre-EMU core-periphery pattern (1960–1988) from Bayoumi and Eichengreen and our updates Post-EMU (1989–2015) based on the correlations of demand and supply disturbances. Note: This figure reports for the post-EMU estimates median bootstrapped residuals based on 10.000 VAR replications. Structural residuals are retrieved from a SVAR where the over-identifying restriction above is imposed for all countries, with the exception of Ireland, Spain, Greece and Portugal. The sample for this SVAR is 1989–2015, with two lags for all countries and no constant as in Bayoumi and Eichengreen (1993). BE = Belgium; DK = Denmark; ES = Spain; FR = France; IE = Ireland; IT = Italy; NL = The Netherlands; PT = Portugal; UK = United Kingdom; GR = Greece.

Using *NORD*, there is now a larger number of countries in the core, a smaller number in the periphery, and the distance between the two groups is smaller. Based on these static results, the new smaller periphery is comprised solely of Spain, Portugal, Ireland and Greece (Appendix 4). We identify a country as less symmetrical if its *NORD > 50%;* which is, if the number of times the hypothesis of symmetry – based on the over-identifying restriction explained above – is rejected in more than 50% of cases (see Campos and Macchiarelli, forthcoming). The results suggest that more frequently the model under the proposed over-identifying symmetry restriction is rejected, the more peripheral a country is said to be. Conversely, the lower this frequency (i.e. the less often the model is rejected for a given country in a given year), the higher the probability of a country being classified as core. In this sense, based on the chosen parametrization for γ, Germany would have a low probability to be classified as periphery (25%), whereas Greece display a high probability, equal to 92.5% chances of rejection of the same restriction. Importantly, this make the results independent of the choice of anchor country (a role often reserved for Germany) as they only count the number of rejections of the suggested symmetry test.

The static results also show that, after the introduction of the euro, countries such as the United Kingdom have become more synchronised, albeit this value for the UK is close to our 50% threshold. This result has important yet still poorly understood implications in terms of the net benefits of European integration: the European Monetary Union might have fostered integration in the EU as a whole.

A very important extension and update we carry out in the present paper is to use our methodology for the whole EU countries, as opposed to just the EU12 as in Bayoumi and Eichengreen (1993) and Campos and Macchiarelli (forthcoming). This extension shows (Fig. 2) the value of NORD across a much larger set of countries. It suggests the formation of a bigger core with the addition of Sweden, Austria and Slovenia and a bigger periphery, with Spain, Finland, Hungary, Poland, Slovakia and Portugal. More specifically, for the EU as a whole using our new index we find that the periphery is composed of (in decreasing order of the magnitude of NORD) Latvia, Ireland, Lithuania, Estonia, Luxemburg, Czech Republic, Greece, Portugal, Slovakia, Poland, Hungary, Finland, and Spain, while the core contains the UK Sweden, Denmark, Germany, Austria, France, Netherlands, Slovenia, Belgium, Italy (see Appendix 5 for further details). Note that data for Bulgaria, Croatia, Cyprus, Malta and Romania are not included in the OECD Annual Accounts.

These static snapshots cover the post-EMU period and are strictly comparable with the results that Bayoumi and Eichengreen and others have produced for the pre-EMU period. They are also very much in line with the conclusions from the important analysis of business cycle synchronisation of the New Member States of Central and Eastern Europe by Di Giorgio (2016).

5. Evaluating NORD

The main requirement to compute the dynamic *NORD* is the availability of data going back to the early 1960s. In this spirit, one of the main advantages of implementing our approach using a rolling sample is that of obtaining a theory-based classification of symmetry which is non-binary (core-periphery) but rather evolves over time. The requirement of conditioning the estimation onto a longer time-series disqualifies most of the new EU member states, which did not start tran-

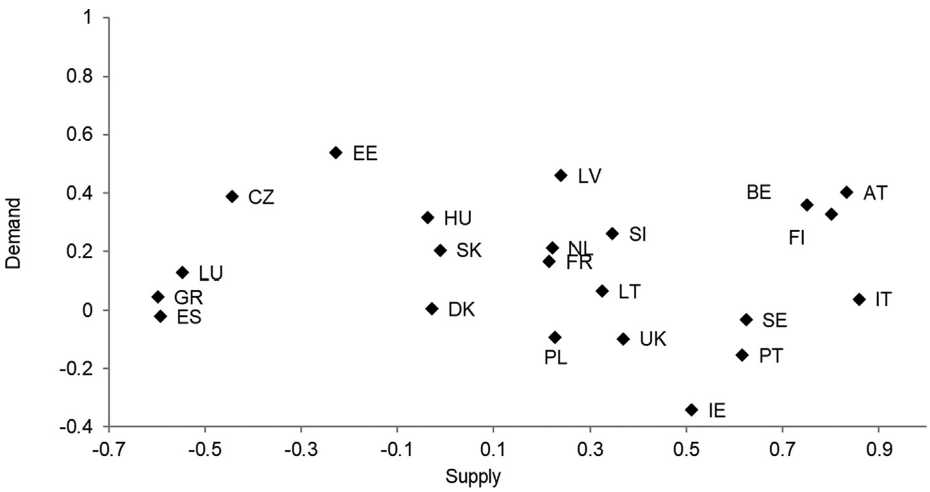

Fig. 2. Core-periphery pattern (1989-2015) based on the correlations of demand and supply disturbances for the EU28. Note: This figure reports for the post-EMU estimates median bootstrapped residuals based on 10.000 VAR replications. Structural residuals are retrieved from a SVAR where the over-identifying restriction above is imposed for all countries, with the exception of countries for which NORD > 50 (see Appendix 5). The sample for this SVAR is 1989–2015, with two lags for all countries and no constant as in Bayoumi and Eichengreen (1993). AT = Austria; BE = Belgium; CZ = Czech Republic; DK = Denmark; EE = Estonia; ES = Spain; FI = Finland; FR = France; HU = Hungary; IE = Ireland; IT = Italy; LT = Lithuania; LU = Luxemburg; LV = Latvia; NL = the Netherlands; PL = Poland; PT = Portugal; SE = Sweden; SI = Slovenia; SK = Slovakia; UK = United Kingdom; GR = Greece. Data for Bulgaria, Croatia, Cyprus, Malta and Romania are not included as they are not part of the OECD Annual Accounts.

sitioning to a market economy up until the early 90 s. In light of the above, our index covers 16 European countries between 1987 and 2015. These are Austria, Belgium, Germany, Spain, Finland, France, Greece, Ireland, Italy, the Netherland, Portugal, plus three EU non-euro area countries (Denmark, Sweden and UK) and two countries that are not EU members (Switzerland and Norway). The objective of this section is to evaluate our new theory-based continuous measure of the probability of a country being classified as periphery (*NORD*).

In order to shed light on the behaviour of *NORD* across countries and over time, we draw from OCA theory (De Grauwe, 2018) and identify a set of candidate explanatory variables. One first group refers to explanatory variables on the fiscal side (Martin and Philippon, 2017). The expectation, in this case, is that countries for which debt to GDP ratios and cyclically adjusted budget balances are larger should be more (less) likely to be classified as peripheral (core).

The second group of variables we use to assess the validity of our index covers external links, in particular, the roles of foreign direct investment (FDI), real effective exchange rate, and trade openness (Rose, 2000; Brouwer et al., 2008). Note that, in this latter case, we examine the potentially different roles of the shares of exports and imports over GDP. Our expectation here is that larger the inflows of FDI, increases in the real exchange rate or in the share of imports plus exports in GDP will – all else equal – make a country less likely to be classified as periphery.

The third group of explanatory variables focuses on financial links, more specifically we use corporate bond spreads, 10-year government bond spreads, 3-month interbank interest rate spreads, average consumer loan interest rate spreads, and returns on equity differential. All spreads are computed *vis-à-vis* Germany. This set of financial variables is consistent with the European Central Bank's definitions of financial integration (ECB, 2017, Spiegel, 2009). We hypothesise that the more financially integrated with the rest of the EU in a given year an individual country is, the more likely it will be classified as belonging to the core.

The fourth and last main group of explanatory variables we draw upon regards structural reforms. We use OECD data on employment protection legislation (EPL), covering both permanent and temporary contracts, and on product markets regulation (PMR). One should expect that the more reforms countries have implemented, the more likely it is that they will not be classified as periphery. In other words, the more extensive is employment legislation and product regulations, the more likely the country can be classified as belonging to the periphery group.

We include a variable for euro area membership as suggested by the endogenous OCA theory (Frankel and Rose, 1998). Eurozone membership takes the values of one for countries joining the EMU starting from accession year.

In summary, we expect that the probability of being classified as core is higher in countries-years in which debt to GDP is lower, trade and foreign investment flows are higher, financial integration is higher, structural reforms are more extensive and the country enjoys full membership in a currency union.

To address possible endogeneity and omitted variables issues, we employ a GMM/IV approach.[5] In the estimation, the number of endogenous variables equals the number of instruments, where we select the instruments to be the lagged dependent variables, the constant and the Eurozone (EZ) membership dummy. This choice tries to address concerns about the instrument proliferation issue in GMM (Roodman, 2009). All regressions account for unobserved heterogeneity.

Table 1 shows our baseline results. We include a dummy variable capturing membership in the Eurozone; this is always statistically significant and carries the expected negative sign. It suggests EZ membership is associated with higher symmetry (it lowers the number of rejections that NORD captures). The first column also shows two important fiscal variables: countries with higher debt to GDP ratios and cyclically adjusted budget balances tend to be less symmetrical, as these are positively associated with *NORD*. Regarding financial integration, we notice that only one of the four dimensions we account for turns out to be statistically significant, namely inter-bank spreads. As for external linkages, the coefficients on both foreign direct investment inflows and the real exchange rate carry the expected negative signs and are statistically significant (we establish below that the FDI result is comparatively more robust). Everything else the same, lower values of *NORD* (that is, higher probability of being classified as core) are associated with countries receiving higher inflows of foreign investment. The results in column (4) for trade openness indicate that the larger the share of the sum of exports and imports over GDP, the more symmetric a country is. Column (5) shows our results for two structural reforms that have played a big role in the debate so far on Eurozone imbalances and they both seem to support *NORD* as a satisfactory measure. They show that the more protective is employment legislation and the more regulated are the product markets, the higher is *NORD* (that is, the higher the probability of the country being classified as periphery).

The overall results (last column in Table 1) suggest that a strong role is played by membership in the Eurozone and by the strictness of product market regulation, whereby a high PMR decreases symmetry. On balance, this set of results seem to support *NORD* as a continuous time-variant measure of the probability of a country being classified as core or periphery. There are no unexpected or counter-intuitive results. Membership of the currency union, for the countries in our sample, suggests an important role in making countries less "peripheral", with the size of this reduction being as much as 16 percentage points.[6]

Table 2 extends our baseline results, with the coefficient of the Eurozone membership dummy and that on foreign direct investment remaining significant throughout. In Table 2, we further investigate the other key result from Table 1, regarding

[5] The results are qualitatively the same if we use a Huber/Eicker/White heteroskedasticity-robust variance-covariance matrix or, as we do for the reported results, we set the weighting matrix to equal the 2SLS (these are available upon request from the authors).

[6] The results for the EA membership and PMR remain robust when the EA dummy is set to start in 2002, i.e. the year when euro coins were effectively introduced (see Appendix 6).

Table 1
The Determinants of a Country's Probability of Being Classified as Periphery/Core (NORD, 16 countries annual data for 1991–2015, GMM estimates).

	(1) Fiscal	(2) Financial	(3) External	(4) Trade	(5) Reforms	(6) All
Debt/GDP	0.118***					0.090
	(0.042)					(0.141)
Adj. Budget Balance	0.836**					1.629*
	(0.336)					(0.965)
Corporate bond spread		0.484				0.701
		(0.617)				(0.679)
3-month interbank spread		−4.455***				−0.760
		(1.435)				(2.742)
Avg on consumer loans spread		−0.200				−0.360
		(0.479)				(0.866)
Return on equity diff.		0.634				−0.623
		(0.438)				(0.397)
FDI			−0.479***			−0.419**
			(0.152)			(0.202)
Reer (CPI adj.)			−0.240**			−0.492*
			(0.118)			(0.294)
Trade openness (%GDP)				−0.135**		−0.457**
				(0.065)		(0.191)
EPL temporary					5.300***	4.468
					(1.871)	(3.856)
PMR					6.898*	1.906
					(3.537)	(6.872)
Eurozone membership	−19.284***	−28.330***	−12.038***	−11.612***	−8.911***	−15.363**
dummy (1999)	(2.105)	(3.160)	(1.935)	(2.155)	(3.199)	(6.243)
C	69.516***	82.239***	98.345***	82.566***	49.242***	147.812***
	(3.135)	(2.252)	(11.891)	(4.693)	(7.951)	(45.754)
Adj-R2	0.748	0.679	0.680	0.710	0.790	0.702
Durbin-Watson	0.577	0.970	0.544	0.465	0.717	1.369
J-Stat (p-value)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)

PMR. Specifically, we study whether the greater competitive pressures created by decreasing regulations in product markets can be somehow captured in different ways. We have in mind the notion that these competitive pressures can somehow be "imported" instead.

The first column of Table 2 shows that including exports and imports instead of trade openness creates a problem of multicollinearity, as not one of the two coefficients is statistically significant.[7] When each of these variables is introduced by itself, their coefficients carry the expected signs and are statistically significant. It is also worth noting that the size of the coefficient on imports is substantially larger (almost twice) as that on exports. More importantly, note that the coefficient on product market regulations is significant only when trade openness, exports or imports are excluded from the specification suggesting that these two sets of variables (trade and reforms) may be capturing the same mechanism (i.e., competition).

In Table 2, we also investigate whether the effects of PMR would be identifiable if we account for possible interactions between trade openness and two different aspects of employment protection legislation (covering either permanent or temporary contracts) and, as it can be seen, this does not help to further isolate the effects of PMR (as its coefficient remains statistically insignificant).

As a robustness check on those findings, in Table 3 we use Cette et al. (2019) new indexes of competitiveness.[8] They measure the mark-up at the industry-country level, based on cross-country-industry panel data from the OECD. They calculate firms' (MUR_i^f) and workers' rent rates (MUR_i^l). We collapse those measures at the sector and country level, obtaining annual series for each country which measure competitiveness on the side of the firms (MUR_i^f) and of the workers (MUR_i^l). Cette et al. measures of mark-up are defined as

$$MUR_i^f = \frac{P_iQ_i - (W_iN_i + M_i)}{W_i^rN_i + M_i} \quad \text{and} \quad MUR_i^l = \frac{(W_i - W_i^r)N_i}{W_i^rN_i + M_i}$$

where, in industry i, P_i is the relative production price (i.e. the ratio of production price to GDP price), W_i the average compensation, M_i the intermediate input total cost and Q_i the production at constant prices. In addition, W_i^r and N_i, are the over-

[7] This is because of the high pair-wise correlation between these two variables (as shown in correlation matrix in the Appendix, multicollinearity is not a problem for any of the variables we use, except exports, imports and openness).

[8] Based on Cette et al (2019) paper, and as far as Europe is concerned, the results for the new competitiveness indicators are available for Austria, Belgium, Spain, France, Italy, the Netherlands, plus Denmark, Finland and Sweden.

Table 2
Importing Competition:How Imports/GDP Substitute for PMR when estimating Country's Probability of Being Classified as Periphery/Core (NORD, 16 countries annual data for 1991–2015, GMM estimates).

Debt/GDP	0.057	0.081	0.092	0.090	0.117	0.044
	(0.152)	(0.143)	(0.140)	(0.141)	(0.148)	(0.139)
Adj. Budget Balance	1.609	1.654*	1.580*	1.629*	0.978	1.209
	(1.029)	(0.992)	(0.948)	(0.965)	(0.928)	(0.955)
Corporate bond spread	0.112	0.464	0.873	0.701	0.166	0.929
	(0.928)	(0.724)	(0.656)	(0.679)	(0.661)	(0.672)
Gvt bond spread	−1.022	−0.661	−0.340	−0.617	−3.548*	−0.581
	(1.926)	(1.717)	(1.649)	(1.718)	(1.919)	(1.704)
3-month interbank spread	−0.725	−0.576	−1.053	−0.760	0.717	−2.971
	(2.935)	(2.802)	(2.696)	(2.742)	(2.504)	(2.426)
Avg on consumer loans spread	−0.478	−0.466	−0.228	−0.360	−0.200	0.439
	(0.927)	(0.887)	(0.850)	(0.866)	(0.837)	(0.754)
Return on equity diff.	−0.785	−0.689*	−0.576	−0.623	−0.557	−0.557
	(0.493)	(0.411)	(0.392)	(0.397)	(0.378)	(0.404)
FDI	−0.471**	−0.434**	−0.415**	−0.419**	−0.425**	−0.482**
	(0.228)	(0.207)	(0.200)	(0.202)	(0.206)	(0.211)
Reer (CPI adj.)	−0.480	−0.503*	−0.470	−0.492*	−0.525*	−0.296
	(0.313)	(0.303)	(0.289)	(0.294)	(0.291)	(0.296)
Imports	−2.099	−1.105**				
	(1.547)	(0.442)				
Exports	0.863		−0.718**			
	(1.217)		(0.332)			
Trade openness (%GDP)				−0.457**	1.130**	
				0.191	(0.453)	
EPL temporary	5.509	4.784	4.336	4.468		
	(4.355)	(3.900)	(3.852)	(3.856)		
PMR	−0.58216	0.101	3.969	1.906	5.197	13.228**
	(7.914)	(7.358)	(6.523)	(6.872)	(6.798)	(0.017)
Openness * EPL temporary					−0.032976	
					(0.057)	
Openness * EPL permanent					−0.765***	
					(0.195)	
Eurozone membership dummy	−13.735*	−14.850**	−15.600**	−15.363**	−12.405**	−14.15**
1999	(6.992)	(6.376)	(6.192)	(6.243)	(5.926)	(6.365)
C	158.885***	157.706***	135.533**	147.812***	166.534***	73.414*
	(51.121)	(48.545)	(43.761)	(45.754)	(43.135)	(37.856)
Adj-R2	0.658	0.688	0.708	0.702	0.717	0.691
Durbin-Watson	1.462	1.408	1.341	1.369	1.360	1.324
J-Stat (p-value)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)

all skill average 'reservation wage' and the total number of hours worked at the industry level; where the former is set as equal to or lower than the minimum industry average *observed* workers' compensation across all industries, for a given country and year (Cette et al., 2019).

Using these new indicators, we find that only firms' mark-up is significant (supply) in explaining *NORD*, while workers' mark up (demand) is not. Consistent with our previous findings, imports weaken the effect of the competitiveness indicator, possibly suggesting the role of endogenized competition, and overall suggesting the robustness of our findings also when alternative indicators, sample size and duration are employed.

Table 4 provides an additional important robustness check on the results above. The global financial crisis that started in 2007 and that affected Europe in full in 2010 (the Eurozone sovereign debt crisis) by its very nature generated a similar response across countries. The extent of symmetry increased rapidly so one may be justifiably concerned that this would affect our results. Table 4 shows the simplest test for assessing this idea, namely, of whether splitting the sample in 2010 (or in 2007) substantially affect our baseline results for *NORD*. The results show that, while some coefficients lose statistical significance, our main individual results remain robust with euro membership and trade (or reforms) remaining key factors in explaining *NORD*.

One important remaining concern regards the robustness and magnitude of the effect of euro membership itself. One way of further investigating this issue is to re-estimate the specifications in Table 1 using differences-in-differences. These results are presented in Table 5. The introduction of the euro took place between 1999 and 2002 when twelve European countries started using the single currency.[9] The main coefficient of interest is the interaction between time and the dummy for Euro-

[9] The next country to join was Slovenia in 2017, followed by Slovakia in 2009, Estonia in 2011, Latvia in 2014 and Lithuania in 2015. However, because all of these were communist countries before 1990, data required for our exercise is not available as it does not go back to 1990.

Table 3
Alternative Measures of Competition: Estimating a Country's Probability of Being Classified as Periphery/Core (NORD,16 countries annual data for 1991–2015, GMM estimates).

Debt/GDP	0.729	0.756*	0.742*	1.147	1.008**	0.659**
	(0.455)	(0.417)	(0.393)	(0.945)	(0.394)	(0.282)
Adj. Budget Balance	2.325	1.953	2.404	5.028	1.456	0.327
	(4.976)	(4.282)	(4.458)	(12.860)	(2.925)	(1.578)
Corborate bond spread	1.176	2.509	1.353	0.207	3.428	1.895
	(2.568)	(2.329)	(2.195)	(6.389)	(2.534)	(1.603)
Gvt bond spread	−12.570	−13.059	−12.953	−9.268	−2.969	−8.450
	(13.924)	(12.825)	(11.455)	(29.198)	(12.552)	(5.722)
3-month interbank spread	−4.108	−1.943	−3.905	−5.493	−6.472	−1.891
	(5.179)	(3.651)	(5.398)	(9.993)	(6.622)	(3.139)
Avg on cosumer loans spread	2.539	2.350	2.622	4.995	2.101	1.583
	(2.869)	(2.515)	(2.327)	(8.174)	(1.500)	(1.091)
Return on equity diff.	−0.508	−0.456	−0.509	−1.156	−0.868	−0.313
	(0.686)	(0.588)	(0.682)	(1.663)	(0.857)	(0.435)
FDI	−0.422	−0.251	−0.446	−0.541	−0.320	−0.065
	(1.019)	(0.874)	(0.865)	(1.289)	(0.737)	(0.470)
Reer (CPI adj.)	−0.230	−0.490	−0.230	0.130	−0.422	−0.630**
	(0.787)	(0.564)	(0.784)	(1.807)	(0.432)	(0.283)
Imports	−0.392	1.4251				
	(4.421)	(4.132)				
Exports	2.443		2.187			
	(3.295)		(3.875)			
Trade openness (%GDP)				4.529		
				(10.648)		
EPL temporary	11.488	6.199	11.361		−1.282	2.296
	(16.983)	(11.832)	(17.615)		(7.509)	(4.970)
MURF	6.347*	4.747*	6.295	1.953	0.665	4.305**
	(3.606)	(2.782)	(3.919)	(6.029)	(3.657)	(1.728)
MURL				−5.948	-(6.812)	
				(6.828)	(4.721)	
Openess * EPL temporary				0.096		
				(0.527)		
Openess * EPL permanent				−1.350		
				(2.874)		
Eurozone membership dummy	−79.355	−13.246*	−11.561	−16.696	−25.064**	−13.790**
1999	(313.206)	(6.680)	(8.364)	(25.815)	(12.245)	(5.664)
C	−11.376	15.075	−81.200	69.318	267.280*	88.041**
	(8.106)	(230.533)	(299.873)	(602.515)	(138.229)	(43.564)
Adj-R2	0.545	0.628	0.548	−0.158	0.424	0.725
Durbin-Watson	1.845	1.867	1.860	2.095	2.032	1.728
J-Stat (p-value)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.366)

zone membership. As can be seen, this coefficient is significant throughout and carries the expected negative sign. Adoption of the single currency, everything else equal, makes the probability of being classified as core significantly larger. This coefficient can be interpreted as showing that the actual effect of Eurozone membership is to increase symmetry (or reducing the number of rejections) by about 10 (ten) percentage points. In light of the role played by the various other factors suggested by the OCA theory, this is a non-trivial and, in our view, sensible estimate.

The main results we obtain for *NORD* are also not affected by splitting the sample in before and after the Eurozone crisis of 2010, by using different estimators (logit or OLS), variations on the measurement of the key variables (such as using 2002 for the Eurozone dummy) and the use of standard differences-in-differences.[10]

To conclude, the associations we uncover seem all intuitive: the role of Eurozone membership behaves according to the predictions from endogenous OCA theory (Frankel and Rose, 1998). The results for foreign investment inflows and trade openness (or otherwise product market regulations) are also in line with the aforementioned theoretical predictions (De Grauwe, 2018). The results for fiscal and financial liberalisation variables, while according to theoretical priors, do not emerge as strongly as for other variables.

[10] These are reported in the on-line appendices.

Table 4
The Eurozone Crisis and a Country's Probability of Being Classified as Periphery/Core (NORD, 16 countries annual data for 1991–2015, GMM estimates).

	(1) Fiscal	(2) Financial	(3) External	(4) Trade	(5) Reforms	(6) All
Debt/GDP	0.230***					0.017
	(0.085)					(0.228)
Adj. Budget Balance	0.958**					−0.398
	(0.414)					(1.765)
Corporate bond spread		−0.158				0.794
		(0.623)				(0.846)
Gvt bond spread		2.465				−8.156
		(2.683)				(5.929)
3-month interbank spread		−4.955**				2.741
		(2.106)				(5.042)
Avg on consumer loans spread		−0.026				0.383
		(0.506)				(1.194)
Return on equity diff.		0.519				−0.805
		(0.418)				(0.611)
FDI			−0.626***			−0.370
			(0.181)			(0.248)
Reer (CPI adj.)			−0.308**			−0.820*
			(0.129)			(0.462)
Trade openness (%GDP)				−0.041		−0.436
				(0.096)		(0.385)
EPL temporary					6.008***	2.902
					(2.085)	(4.346)
PMR					8.987**	−1.073
					(4.246)	(12.158)
Eurozone membership dummy	−16.890***	−25.437***	−9.655***	−12.620***	−7.104**	−14.613**
1999	(2.213)	(3.289)	(2.187)	(2.409)	(3.389)	(7.170)
C	62.510***	79.559***	105.507***	75.848***	43.183***	185.747**
	(5.132)	(2.004)	(13.000)	(6.734)	(9.280)	(79.816)
Adj-R2	0.729	0.677	0.659	0.693	0.785	0.605
Durbin-Watson	0.643	1.018	0.684	0.535	0.823	1.509
J-Stat (p-value)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)

6. Dynamics

One important potential concern with our new approach is that our measure may be mostly capturing common (in this case, euro) area factors. In other words, the number of rejections of the symmetry hypothesis may be lower in those countries where common factors are stronger.

To address this concern in full, we use the panel data model proposed by Phillips and Sul (2007). For each time-series of our indicator, the model assumes both common and individual-specific components and is formulated as a nonlinear time-varying factor model, allowing for multiple time paths and individual heterogeneity. This decomposition provides flexibility in idiosyncratic behaviour over time and cross-sections while retaining some commonality across the panel by means of an unknown common component which we identify as a latent "euro area factor." In this case, commonality means that when the heterogeneous time-varying idiosyncratic components converge over time to a constant, a form of panel convergence holds. This is analogous to the concept of conditional sigma convergence. The procedure not only does not impose any particular assumption concerning trend stationarity, thereby being robust to the stationarity property of the time series process, but also has the advantage of singling out a common factor, isolating it from idiosyncratic factors.

Our indicator, *NORD*, is a bounded variable [0,100] and cannot be clearly described by a random walk. However, both from theoretical and empirical points of view, if its adjustment is slow, *NORD* may behave approximately as a random walk over the sample period. In other words, within the chosen sample the process does not show any tendency to be mean-reverting. This is evident by looking at the pattern of the dynamics of our indicator, particularly for what we have identified as 'peripheral' countries. When a process is close to having a unit root, the misspecified model introduced by assuming a unit root can be a good approximation to the true data-generating process. The misspecification introduced by assuming stationarity when in reality there is a unit root leads to more serious errors of inference. The issue may be important in our case if one assumes that structural euro area factors are increasing or reducing the degree of symmetry amongst countries, such that a path dependency of symmetry exists.

Formally, the approach proposed by Phillips and Sul (2007, 2009) allows isolating a common factor from the estimated dynamic indicator, distinguishing it from country-specific factors. The common factor may represent the aggregated common behaviour, but it could also be any common variable of influence on individual behaviour.

The starting point of the model is decomposing the panel data of our variable of interest, *NORD* as:

$$NORD_{it} = g_{it} + a_{it} \tag{8}$$

Table 5
Difference in Differences Evidence on the Effects of Euro Membership and a Country's Probability of Being Classified as Periphery/Core (NORD, 16 countries annual data for 1991–2015).

	(1) Fiscal	(2) Financial	(3) External	(4) Trade	(5) Reforms	(6) All
Time	−0.231	−0.324***	0.075	0.187	−0.276	−0.041
	(0.159)	(0.168)	(0.127)	(0.151)	(0.287)	(0.371)
Time*EZ membership	−1.093***	−0.673*	−0.744***	−0.744***	−0.585*	−0.851**
	(0.282)	(0.260)	(0.245)	(0.245)	(0.310)	(0.364)
Debt/GDP	0.260***					0.135**
	(0.044)					(0.066)
Adj. Budget Balance	0.729***					0.342
	(0.260)					(0.315)
Corporate bond spread		0.053				0.307
		(0.360)				(0.441)
Gvt bond spread		0.787				0.294
		(0.435)				(0.566)
3-month interbank spread		−2.758				−1.496
		(0.764)				(1.224)
Avg on consumer loans spread		−0.487				−0.549
		(0.271)				(0.461)
Return on equity diff.		0.006				−0.034
		(0.004)				(0.056)
FDI			−0.038			0.041
			(0.072)			(0.071)
Reer (CPI adj.)			−0.147			−0.255*
			(0.091)			(0.151)
Trade openness (%GDP)				−0.094		−0.269**
				(0.075)		(0.108)
EPL temporary					1.800	0.314
					(1.641)	(1.930)
PMR					−3.960	−2.295
					(4.590)	(4.964)
Eurozone membership dummy	4.574	−11.604	−0.973	−0.196	−3.748	−11.643*
1999	(5.116)	(4.877)	(4.665)	(4.518)	(5.317)	(6.684)
C	62.573***	84.011	87.293	78.055***	81.737***	122.577***
	(3.638)	(2.511)	(9.355)	(4.738)	(12.946)	(23.234)
Adj-R2	0.758	0.793	0.709	0.709	0.784	0.819
Durbin-Watson	0.564	0.807	0.471	0.459	0.664	0.948
J-Stat (p-value)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)

where g_{it} represents systematic components (such as a permanent common component), and a_{it} reflects country-specific transitory components. In order to separate common from idiosyncratic components, equation (8) is further transformed as:

$$NORD_{it} = \left(\frac{g_{it} + a_{it}}{ea_t}\right) ea_t = \delta_{it} ea_t \tag{9}$$

where δ_{it} is a time-varying idiosyncratic element and $ea_t$is a single common ('euro area') component. Equation (9) is thus a dynamic factor model.

In the general case, it is impossible to estimate the model directly without imposing some structure on δ_{it} and a_t. Phillips and Sul (2007) propose removing the common factor by scaling so as to obtain the relative loading coefficient:

$$h_{it} = \frac{\delta_{it}}{\frac{1}{N}\sum_{i=1}^{N}\delta_{it}} \tag{10}$$

which they call the "relative transition parameter," measuring the loading coefficient relative to the panel average at time t. By definition, h_{it} traces out a transition path of individual i in relation to the panel average (Phillips and Sul, 2007).

We use the Philipps and Sul model to evaluate the convergence properties of the estimated time-varying idiosyncratic components. Their suggested procedure for clustering panel data into clubs with similar convergence characteristics is as follows:

$$lim_{t\to\infty}\frac{NORD_{it}}{NORD_{jt}} = 1 \ \ for \ \ \forall i,j \tag{11}$$

Phillips and Sul (2007) define this condition as "relative convergence." This condition is equivalent to the convergence of the time-varying factor-loading coefficient:

$$lim_{t\to\infty}\delta_{it} = \delta \ \ for \ \ \forall i \tag{12}$$

Showing that the cross-sectional mean of h_{it} is unity and the cross-sectional variance of h_{it} satisfies the condition $H_{it} = \frac{1}{N}\sum_{i=1}^{N}(h_{it}-1)^2 \rightarrow 0 if lim_{t\rightarrow\infty}\delta_{it} = \delta for \forall i$, Phillips and Sul (2007) develop a regression test of the null hypothesis of convergence. Specifically, the hypothesis is implemented by means of the following 'log t' regression model:

$$log\left(\frac{H1}{H_t}\right) - 2\log(\log(t)) = a + b\log(t) + \epsilon_t \tag{13}$$

$$for\ \ t = [rT], [rT]+1, ..., T\ \ with\ \ r > 0$$

where r is a selection of the initial sample's fraction. Phillips and Sul's Monte Carlo simulations indicate that $r \in [0.2, 0.3]$ achieves a satisfactory performance. In line with their results, we set $r = 0.3$because of the short time series in our panel ($T = 29$).

When we run this *log t* regression for the convergence test under the null hypothesis of convergence for the whole panel, the value of the t statistic is -5.587 which is less than -1.650, hence the null hypothesis of overall convergence is rejected at the 5% level. However, the rejection of the hypothesis does not exclude the possibility of the existence of convergence clubs. When testing for convergence among sub-groups, the clustering of convergence sub-groups suggests the following partition:

Group 1: Ireland, Norway, Switzerland;
Group 2: Finland, Greece, Portugal, Sweden;
Group 3: Denmark, Spain, UK;
Group 4: France, Italy;
Group 5: Austria, Belgium, Germany, the Netherlands.

The results suggest all the clubs fulfil the convergence hypothesis. However, the interpretation of the t-statistics requires some care: while the Phillips and Sul's routine may have a natural interpretation for the extent of incremental economic growth and convergence across countries, the intuition regarding the convergence pattern in the degree of symmetry reflected by *NORD* may be less straightforward. In fact, in our setting countries showing the strongest degree of divergence from the core are often the ones for which the *NORD* has the strongest root.

To further investigate this issue, we perform the tests for possible club stemming from Phillips and Sul (2007) approach (see Appendix 10). These show statistical support to the possibility that the original groups 1 and 2 and the original groups 3 and 4 can be merged to form larger convergence clubs.

This implies that three main clubs of countries may have emerged since the euro introduction, namely an *extended periphery* (Group 1 + 2) composed by Finland, Ireland, Norway, Portugal, Switzerland, Sweden, Greece; an *intermediate group* (Group 3 + 4) formed by Denmark, Spain, UK, France, and Italy; and a *hard-core* (Group 5) where one finds Austria, Belgium, Germany, and the Netherlands.

Fig. 2 shows a first group, the core, that have in common a sustained decline in *NORD*. Using the convenient 50% mark for these classification purposes, one can see that the three countries to move below it are Germany and Austria, both by 1999. This is, of course, the year in which the euro was introduced. Belgium joins this core group in 2000, while the Netherlands joins around 2007. If one equates core membership with a strong indication of being part of an optimum currency area (OCA), our results depict the gradual formation of an OCA. These results provide evidence in favour of endogenous OCA theory (Frankel and Rose, 1998) as countries enter or join the core after the EMU. Having *NORD* below 50% is, however, a necessary but insufficient condition for being more symmetrical, or part of the core, as the examples of France and Italy demonstrate. Based on the classification proposed beforehand, we recognize the latter two are fundamentally different from the 'core' based on the Philip and Sul test: they rather belong to an intermediate group (Fig. 3). Within this group, one also distinguishes Denmark, which does not seem to show any clear trends between 1987 and 2015, and the UK, which while still trendless, displays the highest in-and-out volatility. To this intermediate group also belong Spain, whose *NORD* is flat until 1999 and starts to decline afterwards.

The group of countries we call an extended periphery is shown in Fig. 4. Not only these countries all have very high levels of *NORD*, but they also display (on average) a high number of rejections throughout. In other words, their scores are consistently above the 50% mark, showing, worrisomely, a trendless pattern. To this group belong also Greece and Sweden distinctively moving away, in other words, becoming less symmetric over time. It is interesting to note, however, that between the introduction of the euro in 1999 and the onset of the financial crisis in 2007, Greece was moving towards the core (see Fig. 5).

These results provide confirmation of Bayoumi and Eichengreen (1993) early warning about the EMU, specifically about the possibility of an entrenched core and periphery pattern being driven by the latter being deep-seated. Among the periphery, it is noteworthy that neither Norway, Sweden nor Switzerland are eurozone countries.

In summary, the results above reveal that the core and periphery pattern has changed considerably since the late 1980s, both in terms of distance between main groups of countries but also in terms of the trajectories of individual countries. For instance, it is interesting to see how different countries join the core at different points in time. Overall, our new theory-based index, *NORD*, can be seen to decrease over time for the set of countries that the Phillips and Sul (2007) procedure single out and that we call "core" (a result which broadly confirms endogenous OCA predictions), it remains high for most of the

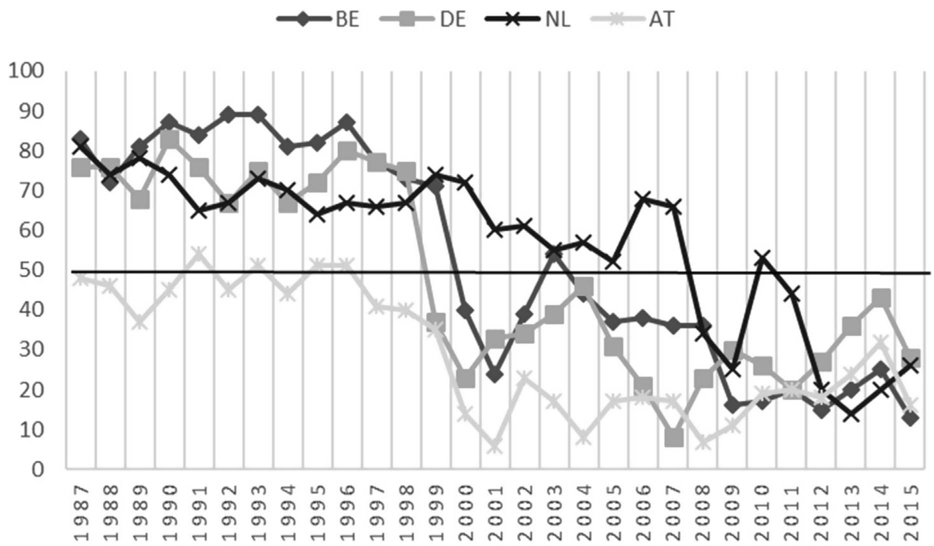

Fig. 3. *NORD* for core countries, 1987–2015 (Core-periphery dynamics). Note: The figure reports the value of NORD, yearly, for BE = Belgium, DE = Germany, NL = The Netherlands, AT = Austria based on the clustering approach of Phillips and Sul (2007). The estimation period for the SVAR model starts in 1962 over a fixed 25-year window.

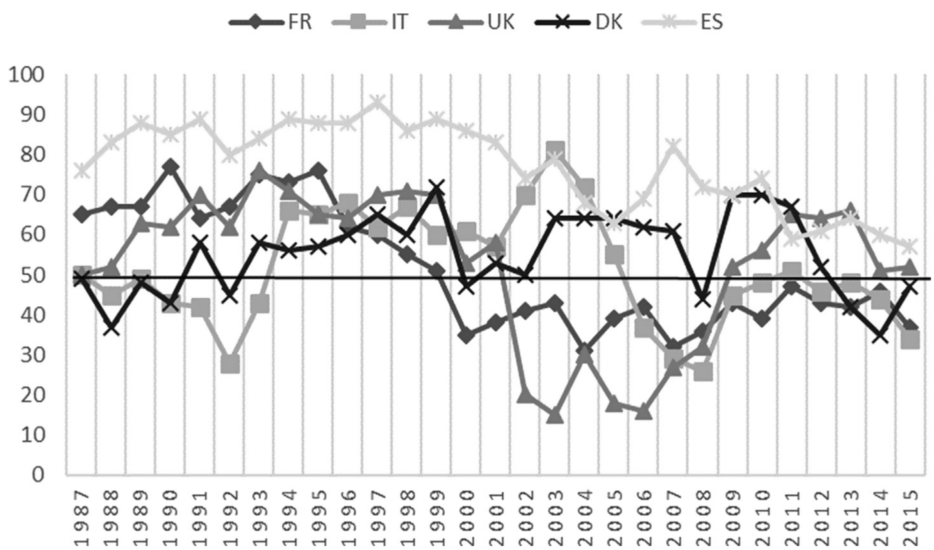

Fig. 4. *NORD* for intermediate countries, 1987–2015 (Core-periphery dynamics). Note: The figure reports the value of NORD, yearly, for FR = France, IT = Italy, UK = United Kingdom, DK = Denmark, ES = Spain based on the clustering approach of Phillips and Sul (2007). The estimation period for the SVAR model starts in 1962 over a fixed 25-year window.

peripheric group (worrisomely confirming Bayoumi and Eichengreen's (1993) warning about a deep-seated periphery), and it varies substantially for an intermediate set of countries between 1987 and 2015.

7. Discussion and policy implications

The debate which took place following Bayoumi and Eichengreen (1993) is centred on the idea of core-periphery and the strengthening of the EMU. According to these authors, what defines core and periphery is the correlation of supply-side shocks. Arguing along the lines of Bayoumi and Eichengreen (1993), Viñals (1996) note that some of the observed heterogeneities across countries resulted from the imperfect coordination of monetary policies, exchange rate movements and currency substitution well before the EMU. It is therefore unsurprising that historical studies had found smaller asymmetries among the so-called core countries, especially as these countries had over the years maintained closer monetary and exchange rate ties (Patterson and Amati, 1998). With the EMU, however, the asymmetries affecting participating non-

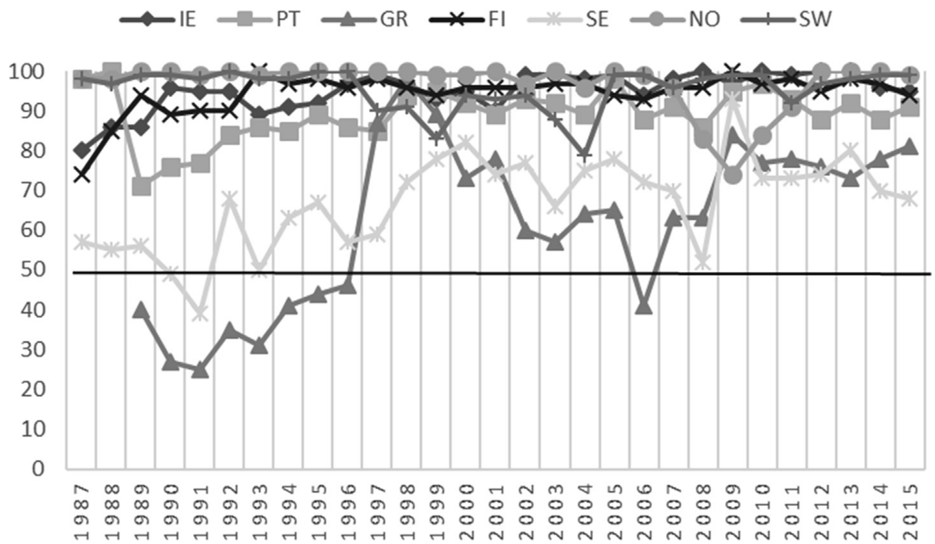

Fig. 5. *NORD* for periphery countries, 1987–2015 (Core-periphery dynamics). Note: The figure reports the value of NORD, yearly, for IE = Ireland, PT = Portugal, GR = Greece, FI = Finland, SE = Sweden, NO = Norway, SW = Switzerland based on the clustering approach of Phillips and Sul (2007). The estimation period for the SVAR model starts in 1962 over a fixed 25-year window.

core countries was expected to fall into line with those of the former so-called core as such imperfect coordination of monetary and exchange rate policies would have "disappear[ed] instantaneously" (Patterson and Amati, 1998).

Our approach is close to some early approaches in analysing the feasibility of the EMU from a business-cycle perspective (see also, among others, Eichengreen, 1991; De Grauwe and Vanhaverbeke, 1993; Patterson and Amati, 1998). Christodoulakis et al. (1995), for instance, asked whether integration implies that countries would have a similar and synchronous response to shocks, or – equally – whether countries would display similar cyclical properties in terms of intensity, timing and duration (see also Macchiarelli, 2013). They found that the different economies of the EU in most respects responded similarly, regardless of the nature of the shock.

The 1992 Maastricht Treaty was effectively a balance between two ideals prevalent in the 1980s of nominal convergence, and an independent monetary policy with a supranational independent central bank (see also Patterson and Amati, 1998). Despite the objectives of the EMU being clearly identified and formulated, this balance left much of the battlefield of real convergence open for further debate. The extent to which asymmetries have declined within the euro area can therefore be seen as another formulation of an old question (Patterson and Amati, 1998): have monetary union led to an economic union? Our answer is yes, but that depends on main drivers behind membership to the single currency itself, with competition being a case in point.

Synchronicity (i.e. the strengthening of the core) since 1999 is likely to have generated substantial benefits. However, two caveats are in order.

One is that these benefits are far from entrenched or irreversible: policy inconsistencies, delays and mistakes can diminish them (and one can argue that this has indeed happened since 2010). In this sense, irreversibility cannot be taken for granted.

The second caveat is that this process of rooting of an endogenous OCA, that our results illustrate, suggests the possibility of substantial economic costs in the case of exit, above and beyond the (maybe symmetric) loss of benefits.[11]

There are various other important features that also deserve close scrutiny, empirically, such as the interactions among trade openness, labour mobility and business cycle synchronisation. Equally, the role of finance (financial cycle) in explaining real convergence, despite not receiving much support from our econometric results, is clearly an area that needs further research. Yet these should be carried out acknowledging that the EMU has changed and will continue to do so. The construction of a Genuine EMU is on-going and a crucial aspect of this debate (Begg, 2016).

8. Conclusions

The objective of this paper is to attempt to further our understanding of symmetry and its dynamics in currency unions, specifically in the context of the European Economic and Monetary Union. One main concern was to try to go beyond the mainly static and binary ways of framing this issue. We put forward a new simple theory-based measure that we think can successfully locate countries in a core-periphery *continuum*. We construct such measure for a set of 16 European coun-

[11] See Campos and Macchiarelli, 2020.

tries yearly from 1987 to 2015 and provided an assessment, based on the endogenous OCA theory, of the main potential underlying explanatory factors of the dynamics of this measure over time and across countries.

Our main conclusion is that this new theory-based measure allows us to clearly identify three sets of countries on the basis not only of the level of our new measure but also in terms of its dynamic behaviour. Using the Phillips-Sul (2007) procedure, we show the emergence of an *extended periphery* composed by Finland, Ireland, Norway, Portugal, Switzerland, Sweden, Greece; an *intermediate group* formed by Denmark, Spain, UK, France, and Italy; and a *hard-core* where one finds Austria, Belgium, Germany, and the Netherlands.

There are valuable lessons from the dynamics of this measure. Our findings suggest that the irreversibility of being in the core should not be taken for granted. Our index increases for core countries (which confirms endogenous OCA predictions), remains worrisomely constant for most of the periphery and varies substantially for the intermediate set of countries. Among the periphery, Sweden and Greece become consistently less symmetrical over time. By contrast, among the intermediate group, Denmark's remains constant and the UK moves in and out of over time. Our panel estimates on a specification suggested by endogenous OCA theory imply that euro membership and product market competition (or trade openness) make the reactions to shocks more symmetric hence affecting the likelihood to have them classified as 'core'. These results have important academic and policy implications as it remains unclear whether the effects of the euro and of competition are indeed closely related or completely independent. Future research will do well to try to throw further light on these important effects.

Appendix A. Supplementary material

Supplementary data to this article can be found online at https://doi.org/10.1016/j.jimonfin.2020.102325.

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