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Time-varying criteria for monetary integration Evidence from the EMU

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Abstract

How can we determine whether an economy will benefit from membership in a monetary union? While economic theory has proposed a number of widely accepted criteria, virtually all empirical studies apply them as if they were time-invariant. The purpose of this paper is to show that unless the dynamic nature of these criteria is explicitly taken into account, the results will be flawed and misleading. We focus on 13 EU countries and two specific criteria: the relative size of output shocks and their synchronization. Using quarterly data from the 1961:1 to 1997:4 period, we show that the parameters relevant to these criteria have exhibited substantial variability over time for essentially all countries in our sample. Our time-varying parameters allow us not just to avoid the flaws of the conventional method, but also to discuss the optimal timing of forming or joining a monetary union.

JEL classification: E42; F36; F42

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1. Introduction

How can we determine whether a certain economy is likely to benefit from membership in a monetary union? Theoretically, a number of economic models have identified various criteria that can be used to determine whether the benefits should be expected to outweigh the

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costs. Some of these are the criteria proposed by the orginal *optimum currency area* literature, such as labor and capital mobility (Mundell, 1961), openness (McKinnon, 1963), and product diversification (Kenen, 1969). Others, emphasized more recently, have to do with the relative magnitude and synchronization of country-specific business-cycle shocks (Eichengreen, 1992, Chap. 3; Emerson, Gros, Italianer, Pisani-Ferry, & Reichenbach, 1992, Chap. 6; Gros & Thygesen, 1992, Chap. 7; Lane, 1999).[1]

Empirically, a vast literature has investigated the extent to which these criteria are satisfied, with particular emphasis on the case of monetary integration in Europe.[2] The main objectives of this line of research are, first, to evaluate the general (overall) prospects of success of a common currency for a certain set of countries, and, second, to identify the countries that are the most promising candidates, in the sense of having the most to gain and/or the least to lose from giving up monetary independence.

It is increasingly recognized, however, that the criteria discussed above are not static, as assumed by the overwhelming majority of the empirical studies, but instead evolve over time. As Frankel (1999, p. 29), the most forceful proponent of this view, put it "[s]uch parameters as openness and income correlations are not fixed for all time, but rather change, in response both to countries' fundamental policy choices and in response to exogenous factors . . ." Virtually all empirical studies, however, have ignored this dynamic element of the optimum currency area criteria, evaluating the desirability of forming or joining a monetary union on the basis of fixed, time-invariant estimates.

The goal of the present paper is to demonstrate that this can be a serious flaw with significantly misleading policy implications. In particular, unless the dynamic nature of these criteria is explicitly taken into account, one can be lead to erroneous conclusions both about the desirability and sustainability of a monetary union, and in terms of comparing and ranking prospective country-candidates.

An additional advantage of estimating and evaluating time-variant criteria is that they enable us to ask, not just whether a common currency for a set of countries is a good idea and who are the most promising countries to join, but also what would be the optimal *timing* for the formation of such a monetary union and *when* should a certain country-candidate join.

The paper uses quarterly data from the 1961 to 1997 period for 13 EU countries, and focuses on two widely used criteria: (1) the relative magnitude of cyclical output shocks, and (2) their degree of synchronization. We selected these two criteria because we expected them to evolve significantly over time, and thus illustrate Frankel's (1999) argument and demonstrate our assertion that imposing time-invariability can limit and bias the empirical results.

The rest of the paper is organized as follows. Section 2 uses a simple model in order to illustrate how the stabilization costs of membership in a monetary union are related to the two criteria. Section 3 describes the empirical methodology and the data. Section 4 contains the

[1] De Grauwe (1997) provides a comprehensive survey of this literature.

[2] For example, Alesina and Wacziarg (1999), Ballabriga, Sebastian, and Valles (1999), Bayoumi and Eichengreen (1992), Bergman (1999), and Karras (1996).

empirical results, first under the assumption that the criteria are time-invariant, and then allowing them to evolve over time. Section 5 discusses and concludes.

2. Theoretical background

This section outlines a simple model in order to illustrate the two criteria used in the paper. The basic approach is that pioneered by Barro and Gordon (1983) and Kydland and Prescott (1977) for monetary policy. Similar versions of this model have been used by Alesina and Grilli (1992, 1994), Alesina and Wacziarg (1999), De Grauwe (1997), Rogoff (1985), and others.

Suppose there are N economies indexed by i ($i = 1, 2, \ldots, N$). Each economy's loss function takes the form

$$L_i = (1/2)E[a_i(y_i - \hat{y}_i)^2 + \pi_i^2], \tag{1}$$

where y denotes output, π inflation, $\hat{y}$ (>0) a target level of output, E the mathematical expectation, and a_i (≥ 0) captures the importance of the output target relative to the inflation target. Aggregate supply is given by an expectations-augmented Phillips curve (with slope normalized to unity and the "natural" rate normalized to zero for simplicity):[3]

$$y_i = (\pi_i - \pi_i^e) + u_i, \tag{2}$$

where π^e is expected inflation, and $u_i \sim (0, \sigma_i^2)$ are economy-specific output shocks. By assumption, the realization of u becomes known after inflationary expectations are set, but before the Central Bank determines π.

Without a monetary union, when each economy's Central Bank can pursue an independent monetary policy, minimizing Eq. (1) subject to Eq. (2) leads to the following dynamically consistent (Nash) equilibrium:

$$\pi_i = a_i\hat{y}_i - \frac{a_i}{1 + a_i}u_i, \tag{3}$$

and

$$y_i = \frac{1}{1 + a_i}u_i. \tag{4}$$

The variability of output is then given by

$$var(y_i) = \frac{1}{(1 + a_i)^2}\sigma_i^2. \tag{5}$$

Note that there is a trade-off between average inflation ($\bar{\pi}_i = a_i\hat{y}_i$) and output variability: if a_i is very low (so that the Central Bank is very "conservative" in the sense of assigning a

[3] Since $\hat{y}$ is assumed to be positive and the natural rate is normalized to zero, it follows that $\hat{y}$ exceeds the natural rate, so that the model has the usual built-in inflationary bias.

higher relative weight to inflation than to output), average inflation will be very low, but output is very unstable.[4]

Next, assume the N economies form a monetary union and monetary authority is delegated to economy 1 ($i=1$). Then, at equilibrium, $\pi_i=\pi_1$ and thus $\pi_i^e=\pi_1^e$ for all i, where π_1 is given as in Eq. (3). So,

$$y_i = (\pi_1 - \pi_1^e) + u_i = u_i - \frac{a_1}{1+a_1}u_1,$$

and thus,

$$\mathrm{var}(y_i) = \sigma_i^2 + \frac{a_1^2}{(1+a_1)^2}\sigma_1^2 - 2\frac{a_1}{1+a_1}\rho_{i1}\sigma_i\sigma_1. \tag{6}$$

It follows, therefore, that from the point of view of country i, joining a monetary union (provided it is dominated by a more "conservative" monetary authority, so that $a_1 \leq a_i$ and $\hat{y}_1 \leq \hat{y}_i$) will reduce the economy's average inflation rate: $\bar{\pi}_i^{\mathrm{UNION}} = a_1\hat{y}_1 < a_i\hat{y}_i = \bar{\pi}_i^{\mathrm{INDEPENDENT}}$. This is one of the main benefits of monetary integration and its main advantage for inflation-prone economies that wish to "import" the price stability of a low-inflation economy.

At the same time, however, comparing Eq. (6) to Eq. (5) shows that membership in the union may very well raise output variability. This is the stabilization cost of membership. From Eq. (6), this cost will depend on the size of σ_i^2 relative to σ_1^2 (Criterion 1). At the same time, the cost will also be smaller, the closer ρ_{i1} is to unity (Criterion 2).

3. Empirical methodology and data

We use two different methods to detrend the output series of each country and estimate the cyclical components.

The first is the Hodrick–Prescott (HP) filter, proposed by Hodrick and Prescott (1980) and extensively used in the business-cycle literature. The HP filter defines the trend component $\bar{x}_t$ of a series x_t as the one that minimizes

$$\sum_{t=1}^{T}(x_t - \bar{x}_t)^2 + \lambda\sum_{t=2}^{T-1}[(\bar{x}_{t+1} - \bar{x}_t) - (\bar{x}_t - \bar{x}_{t-1})]$$

for $\lambda > 0$. The cyclical component is simply $x_t - \bar{x}_t$. Here, we selected $\lambda = 1600$, the value recommended by Kydland and Prescott (1989) for quarterly data.

The second method is the recently very popular Band Pass (BP) method proposed by Baxter and King (1995) and evaluated by Christiano and Fitzgerald (1999) and Stock and Watson (1998), who also compared its properties to those of the HP filter. The low pass (LP)

[4] Rogoff (1985) examines the optimal value for a_i. Fischer and Summers (1989) show that a similar trade-off exists if the source of uncertainty is the Central Bank's inability to determine the inflation rate without error.

filter $\alpha(L)$, which forms the basis for the BP filter, selects a finite number of moving average weights α_h to minimize

$$Q = \int_{-\pi}^{\pi} |\delta(\omega)|^2 d\omega$$

where $\alpha(L) = \sum_{h=-K}^{K} \alpha_h L^h$ and $\alpha_K(\omega) = \sum_{h=-K}^{K} \alpha_h e^{-i\omega h}$. The LP filter uses $\alpha_K(\omega)$ to approximate the infinite MA filter $\beta(\omega)$. Define $\delta(\omega) \equiv \beta(\omega) - \alpha_K(\omega)$. Minimizing Q minimizes the discrepancy between the ideal LP filter $\beta(\omega)$ and its finite representation $\alpha_K(\omega)$ at frequency ω. The main objective of the BP filter as implemented by Baxter and King (1995) is to remove both the high frequency and low frequency component of a series leaving the business-cycle frequencies. This is formed by subtracting the weights of two LP filters. We define ω_L and ω_H, the lower and upper frequencies of two LP filters as 32 and 6, respectively. We therefore remove all fluctuations shorter than 6 quarters or longer than 8 years. The frequency representation of the BP weights becomes $\alpha_K(\omega_H) - \alpha_K(\omega_L)$, and forms the basis of the Baxter–King filter which provides an alternative estimate of the trend component $\bar{x}_t$. For the implementation, we used the B34S Program documented in Stokes (1997). We tried $K=4$, $K=8$, and $K=12$ in order to investigate how sensitive the results are to the approximation of the theoretical filter. As the relevant results are very similar, we mostly report them for $K=4$ in the sections below.

All data are obtained from the IMF's *International Financial Statistics on CD-ROM* and are quarterly for the period 1961:1–1997:4. The EU economies with data available for the entire period were 13: Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Spain, Sweden, and the UK. Ten of these are European Monetary Union (EMU) members as of January 1, 2000; Greece, Sweden, and the UK are not. Denmark and Portugal are the only two EU countries that are not included in the sample because of data limitations. Output for all countries is measured by Industrial Production, except for Greece where only Manufacturing Production was available. All series are seasonally adjusted at the source, except for Belgium for which only adjusted data were available.

We construct an aggregate series for European output as $y_{E,t} = \sum_{i=1}^{13} w_{i,t} y_{i,t}$, where y denotes output, w the (time-varying) weights used, and i and t index over countries and time, respectively. The weights are computed as $w_{i,t} = \text{GDP}_{i,t}^{\text{PPP}} / \sum_{j=1}^{13} \text{GDP}_{j,t}^{\text{PPP}}$, where $\text{GDP}_{i,t}^{\text{PPP}}$ is Gross Domestic Product expressed in constant (1991) Purchasing Power Parity (PPP) prices in US dollars. Annual data for $\text{GDP}_{i,t}^{\text{PPP}}$ are obtained from the *OECD Statistical Compendium on CD-ROM*. The *average* values of the weights in 1960–1997 are as follows: Austria 2.35%, Belgium 3.15%, Finland 1.40%, France 18.27%, Germany 22.50%, Greece 1.70%, Ireland 0.68%, Italy 16.70%, Luxembourg 0.15%, the Netherlands 4.43%, Spain 8.21%, Sweden 2.81%, and the UK 17.65%.[5]

[5] We are grateful to an anonymous referee for suggesting the use of "Europe" as the benchmark, instead of Germany, which we had used in the original version of this paper. While there are some differences in the results, we think the estimates based on Europe are more interesting and did not report those against Germany in order to preserve space. The German results are available on request.

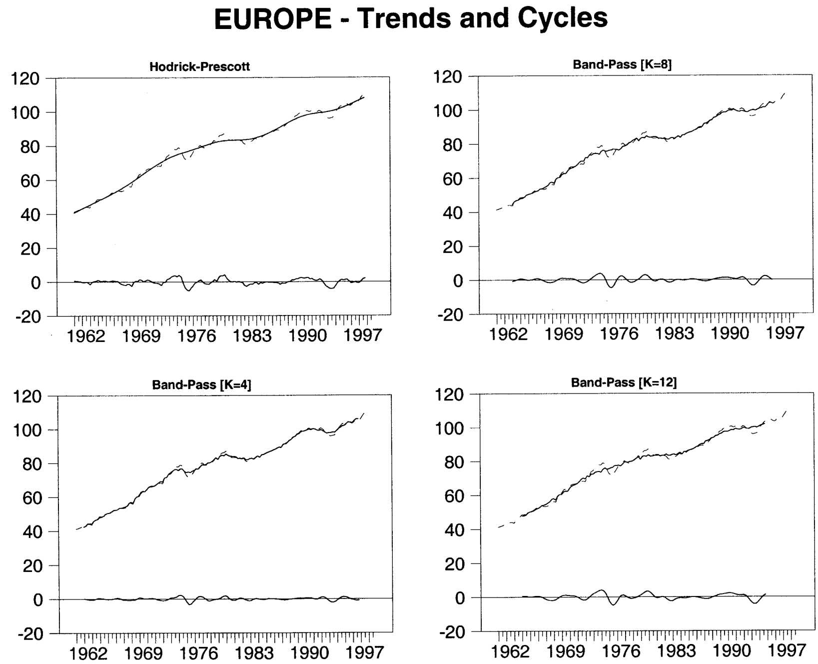

Fig. 1. See text (Section 3) for details on the HP and BP filters.

Fig. 1 shows the output series for Europe and its decomposition into trend and cyclical components using the HP filter and three versions of the BP filter ($K=4$, $K=8$, and $K=12$). While the similarities in the decompositions outweigh the differences, the following two observations are worthwhile. First, the HP filter tends to produce a smoother trend component, $\bar{x}_t$, than the BP filters; it follows that the HP filter's cyclical component, $x_t - \bar{x}_t$, , is more choppy than BP's. This is the case for each version of BP we tried. Second, the size of the cyclical component produced by the BP filter seems to increase as the number of MA terms (the values of K) goes up. In addition, the BP decomposition resembles HP more closely for $K=12$ than for $K=4$. As we wanted our results to be as robust to the decomposition method as possible, we selected to report the $K=4$ BP results, together with the HP results, below.

4. Empirical results

4.1. Time-invariant criteria

We begin by imposing the constraint that the criteria under consideration are time-invariant. This means that they have been constant over 1961:1–1997:4 and thus each of them can be captured by a single number for each country for the entire period. We impose this restriction (for now) not because we believe it is valid, but in order to compare the results with those of the rest of the literature, and contrast them with those of Section 4.2, when the restriction will be relaxed.

We start with Criterion 1. Table 1 reports the variances of the cyclical components for each of the 13 economies, and their size relative to Europe's, using both the HP and BP detrending methods. The "Ordering" columns of Table 1 rank the 13 economies in ascending order of economy-specific variance. While one or two of the countries give results that differ between the two filters (such as Sweden and, to a lesser extent, Finland), for the majority of the economies the results are robust to the choice of method. Thus, Austria and the Netherlands appear to have the most stable cyclical components, while Italy, Ireland, and Luxembourg are characterized by the most volatile cycles. Interestingly, Germany's cyclical variability has exactly the same ranking in both methods. Moreover, half of the countries in the sample are found to have cyclical variances higher than Germany's, while the other half have lower.

Criterion 2 depends on how synchronized the cyclical components are across the countries. Table 2 reports the correlation coefficient of each country's cyclical component with Europe's, both for the HP and BP methods. The "Ordering" columns of Table 2 rank the 13 economies in descending order of correlation with Europe. The results between the two filters are even more similar than those of Table 1. Perhaps the most striking fact from Table 2 is that all correlations with Europe are positive. Nevertheless, sizable differences across countries do exist, the coefficients ranging from .35 for Ireland (BP filter), to .89 for France (also BP filter). More specifically, and not at all surprisingly, the "core" countries of France, Germany, Belgium, Austria, and Italy are the most correlated with Europe, while the "periphery" countries of Finland, Greece, Ireland, Sweden, and the UK, the least.

Table 1
Variances of cyclical components

	HP filter			BP Filter		
i	σ_i^2	σ_i^2/σ_E^2	Ordering	σ_i^2	σ_i^2/σ_E^2	Ordering
Austria	3.92	1.31	3	0.76	1.02	1
Belgium	5.56	1.85	6	1.36	1.82	9
Finland	6.57	2.19	8	0.97	1.30	3
France	3.89	1.30	2	1.08	1.44	6
Germany	5.62	1.88	7	1.12	1.49	7
Greece	6.75	2.25	9	1.46	1.95	10
Ireland	7.22	2.41	12	2.05	2.73	12
Italy	7.05	2.35	10	1.93	2.58	11
Luxembourg	12.63	4.21	13	3.19	4.27	13
Netherlands	3.85	1.29	1	0.94	1.26	2
Spain	4.70	1.59	4	1.03	1.38	5
Sweden	7.13	2.38	11	1.01	1.35	4
UK	5.17	1.72	5	1.13	1.51	8
Europe	3.00	1.00	–	0.75	1.00	–

Full period: 1961:1–1997:4.

The HP filter uses $\lambda = 1600$. The BP filter is implemented as in Baxter and King (1985) using $K = 4$ lags. The σ_i^2 is the variance of country i's cyclical component; σ_E^2 is the variance of Europe's cyclical component. "Ordering" ranks the countries in ascending order of country-specific variance.

Table 2
Cyclical correlations with Europe

	HP filter		BP filter	
i	ρ_{iE}	Ordering	ρ_{iE}	Ordering
Austria	.78	3	.82	4
Belgium	.78	4	.85	3
Finland	.47	13	.41	11
France	.85	1	.89	1
Germany	.82	2	.87	2
Greece	.52	12	.40	12
Ireland	.53	11	.35	13
Italy	.76	5	.81	5
Luxembourg	.76	6	.81	6
Netherlands	.68	8	.75	7
Spain	.75	7	.75	8
Sweden	.53	10	.60	10
UK	.66	9	.65	9

Full period: 1961:1–1997:4.

See notes to Table 1. ρ_{iE} is the correlation of country i's cyclical component with Europe's cyclical component. "Ordering" ranks the countries in descending order of correlation with Europe.

On the basis of the empirical findings of Tables 1 and 2 and the theoretical implications of Criteria 1 and 2, we would conclude that the "best" (in terms of stabilization costs) candidates for EMU are France, Germany, Austria, Belgium, and the Netherlands, as the cyclical components of these countries are both highly correlated with Europe's and of relatively low variance compared to Europe's. A low σ_i^2 relative to σ_E^2, and a high ρ_{iE} for these countries imply that the cost–benefit calculus of adopting a common European currency is favorable: Europe-wide monetary policy will be stabilizing for these economies because it will be a close substitute for independent monetary policy.

The opposite conclusion holds for Ireland, Greece, Sweden, and the UK, as their cyclical variances are high relative to Europe's, and their correlation with Europe is low. On the basis of the two criteria, these countries should expect monetary union to be accompanied by sizable stabilization costs: Europe-wide monetary policy may end up being destabilizing for these economies, because it is a poor substitute for independent monetary policy. It is worth noting that three of these countries (Greece, Sweden, and the UK) are the only EU members in our sample that are not (yet) also EMU members.[6]

Finally, for completeness, we also examine the properties of the trend components of output.[7] Table 3 reports variance ratios and correlations of each country's trend component relative to Europe's, using both the HP and BP methods. While Criteria 1 and 2 for monetary integration depend on the properties of the cyclical components that determine the direction, efficacy, and desirability of independent monetary policy, the trend results also contain valuable information for the more general process of economic

[6] Although Greece is almost sure to join the *euro* at the beginning of 2001.

[7] We wish to thank an anonymous referee for suggesting we report the trend statistics.

Table 3
Trend variances and correlations with Europe

	HP filter		BP filter	
i	σ_i^2/σ_E^2	ρ_{iE}	σ_i^2/σ_E^2	ρ_{iE}
Austria	1.71	.99	1.70	.99
Belgium	0.83	1.00	0.83	1.00
Finland	2.23	.98	2.23	.97
France	0.97	.99	0.97	.99
Germany	0.81	1.00	0.82	.99
Greece	2.17	.95	2.18	.95
Ireland	5.18	.87	5.12	.87
Italy	1.29	1.00	1.30	1.00
Luxembourg	0.85	.90	0.87	.90
Netherlands	1.39	.99	1.39	.99
Spain	1.96	.99	1.96	.99
Sweden	0.80	.97	0.83	.95
UK	0.60	.98	0.62	.98

Full period: 1961:1–1997:4.
σ_i^2 is the variance of country i's trend component; σ_E^2 is the variance of Europe's trend component. ρ_{iE} is the correlation of country i's trend component.

integration. Note, for example, that variance relative to Europe ranges from 0.6 for the UK to 5.18 for Ireland. Also, as expected, virtually all correlation coefficients are between .90 and 1.00.[8]

4.2. Time-variant criteria

In Section 4.1, each country's parameters relevant to Criteria 1 and 2 (σ_i^2, σ_E^2, and ρ_{iE}) were estimated for the entire period 1961:1–1997:4. This is equivalent to imposing the restriction that these parameters have been constant during this period for each of these countries. The purpose of this section is to relax this assumption and demonstrate that it is not innocuous.

To that end, we estimate time-varying parameters, $\sigma_{i,t}^2$, $\sigma_{E,t}^2$, and $\rho_{iE,t}$, computing them for consecutive "rolling" windows of fixed length, k. Specifically, the estimated $\sigma_{i,t}^2$, $\sigma_{E,t}^2$, and $\rho_{iE,t}$ are based on the HP- and BP-filtered data over the time period $t-k$ to t, where $1961{:}1+k<t<1997{:}4$, and k is the number of observations in (i.e., the "length" of) the window. Note that, even though the HP algorithm does not sacrifice data points at the beginning or the end of the window, we make sure to define both HP and BP over the same window in order for them to be comparable. This exercise allows an assessment of how the parameters evolved over time. For the results reported below, we select $k=48$, i.e., a window of 48 quarters, or 12 years. By construction, then all estimated time-varying parameters will

[8] Note, however, that the trend components of output are nonstationary series, so (1) individual (although not necessarily relative) variances may not be finite, and (2) the computed correlations are subject to the well known Granger and Newbold (1974) "spurious" correlation critique.

start in 1973:1 (=1961:1 + k) which roughly coincides with the beginning of the floating exchange rate period.

Fig. 2 looks at the evolution of Criterion 1 during 1973:1–1997:4 by plotting each country's variance relative to Europe's, $\sigma_{i,t}^2/\sigma_{E,t}^2$, for each of the 13 countries. With the interesting exception of France, the relative variances of all other countries have exhibited significant variability over time. Note that both the HP and BP (reported here for $K=4$) decompositions give quite similar results. For most of the countries (such as Belgium, Germany, Greece, Italy, the Netherlands, Sweden, the UK, and others), the variance relative to Europe peaked in the early 1990s, and has been declining since. Whether this peak was related to the 1992 crisis in the European Monetary System (EMS) is unclear, but it strongly suggests that, in terms of relative cyclical variability, the late 1990s was a much better timing for instituting the *euro*, than almost any other time since the beginning of the floating-rate period.

Fig. 3 plots the estimated $\rho_{iE,t}$ for each country over 1973:1–1997:4. Once more, there is no appreciable difference between the BP and HP estimates. One of the most important results of Fig. 3 is that for all countries, and for the entire period under consideration, correlation with Europe has been positive.[9] Equally important, however, these correlation coefficients have varied widely both across countries and over time, sometimes coming very close to zero. As expected, France, Germany, and (since the late 1970s) Italy have the highest and, overall, most stable correlation with Europe. However, a decline in $\rho_{iE,t}$ during the late 1980s and early 1990s is conspicuous in every country's plot. The prime suspects for this pan-European drop in business-cycle correlation are the economic consequences of German unification and (perhaps related) the EMS crisis of 1992. While all 13 correlations have recovered since then, we note that both the size of the decline and the speed of recovery have been very different across countries. Thus, France and Italy experienced the smallest loss of synchronization with Europe, while Ireland, Greece, Spain, and the Netherlands the largest. On the other hand, Spain and France have fully recovered (and they may be even more cyclically synchronous with Europe by now than they were in the early 1980s), while Ireland, the Netherlands, and the UK are still below their average levels.

A very interesting question is how (if at all) the evolution of $\sigma_{i,t}^2/\sigma_{E,t}^2$ and $\rho_{iE,t}$ has affected, or been affected by, structural changes within the EU. While a rigorous investigation of this question is beyond the scope of the present paper, we attempt to shed some light on it in a simple way. In both Figs. 2 and 3, we use two vertical lines in every country's plot to indicate the timing of two important developments in the progression towards the *euro*: (1) the formation of the EMS and the establishment of the Exchange Rate Mechanism (ERM) and the European Currency Unit (ECU) in 1979, and (2) the signing of the Maastricht Treaty by the EU heads of state and government in 1992. No obvious patterns can be detected in either figure, but the importance of the issue makes it a very promising subject for future research.

[9] Strictly speaking, however, this is true only for the $K=4$ case. The correlation patterns for $K=8$ and $K=12$ are remarkably similar to those for $K=4$ (and thus not reported), but do give some negative values. All results are available on request.

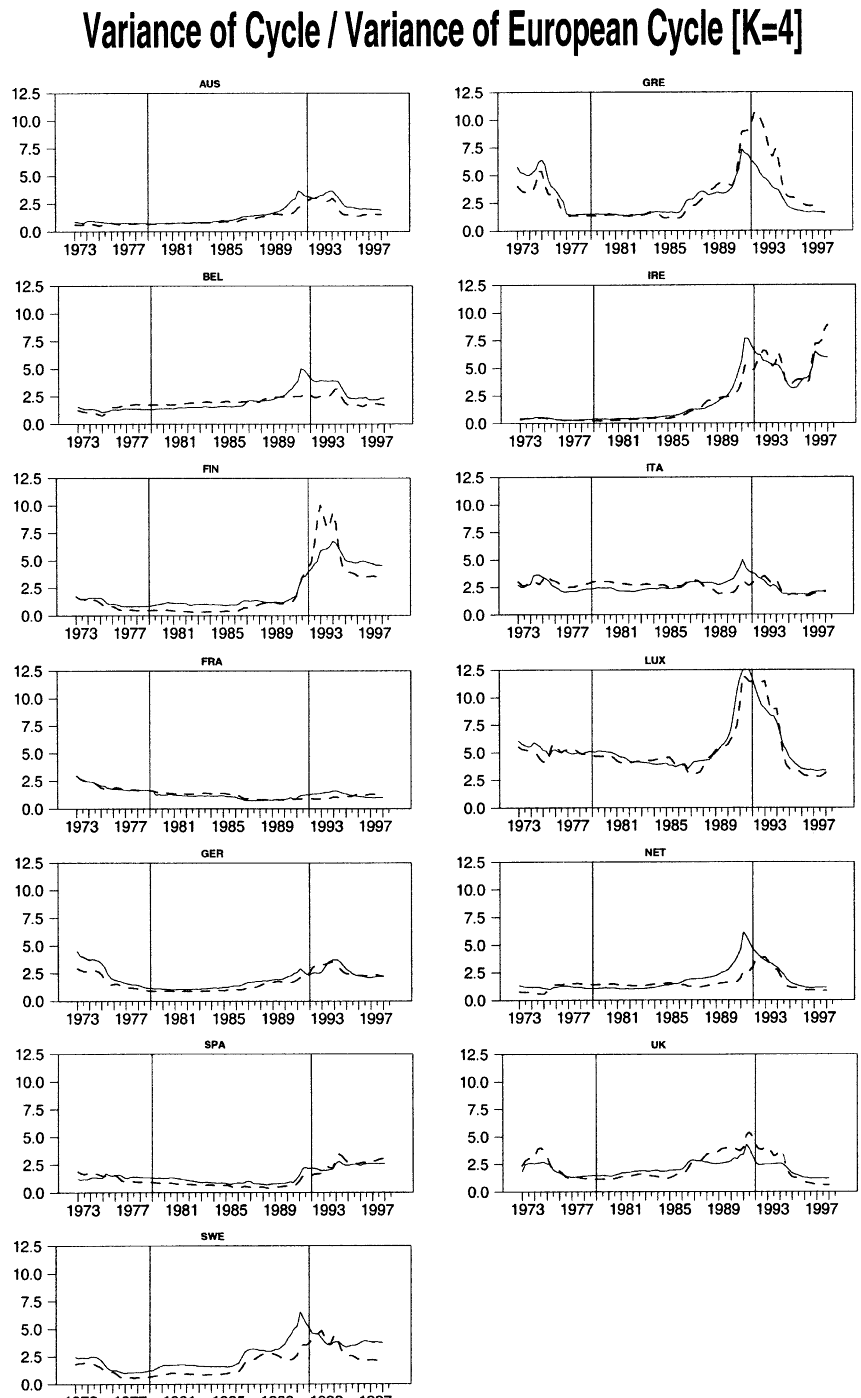

Fig. 2. The graphs plot estimates based on the HP (solid line) and BP (dashed lines) filters. The vertical lines indicate the years 1979 (EMS) and 1992 (Maastricht Treaty).

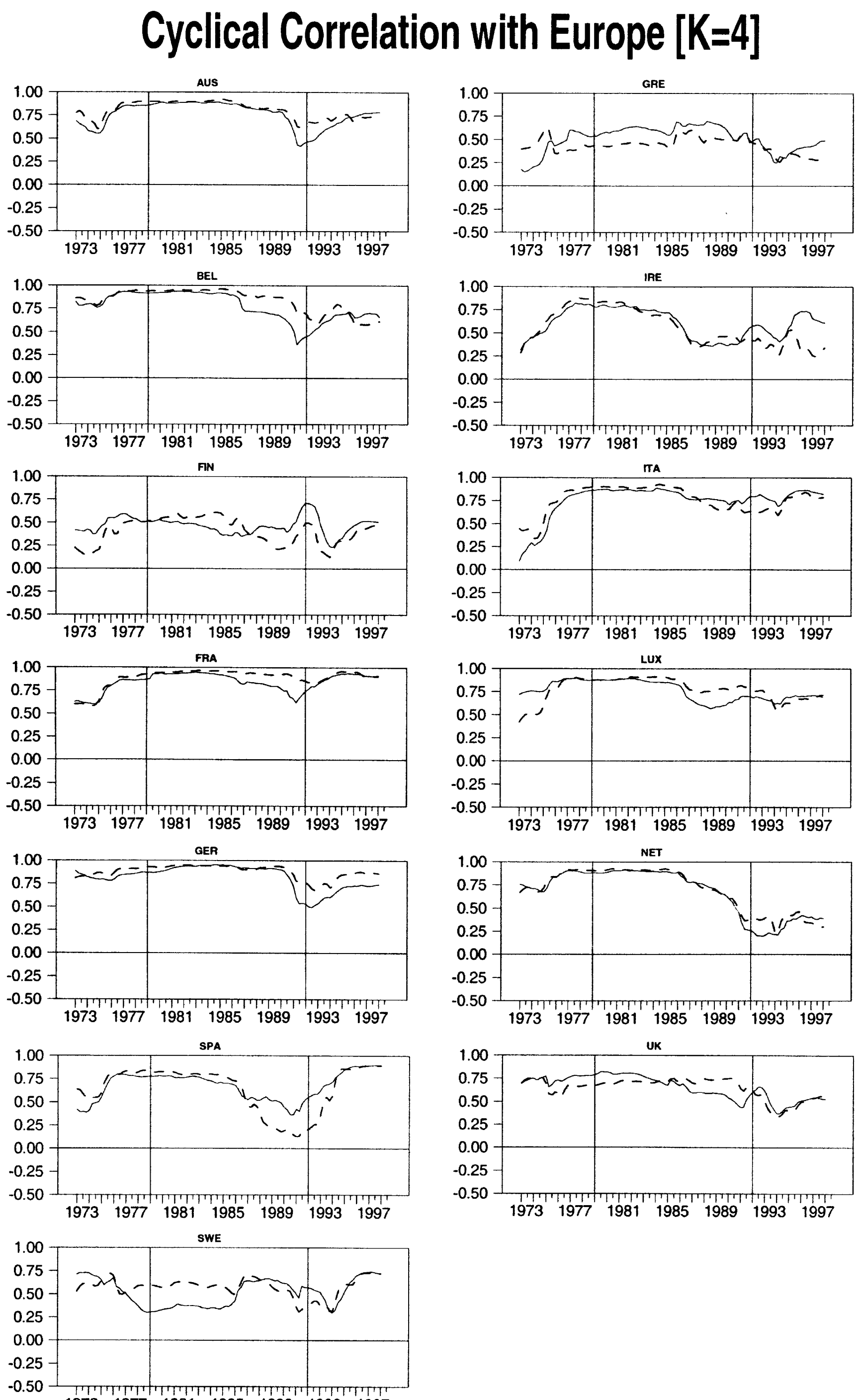

Fig. 3. The graphs plot estimates based on the HP (solid line) and BP (dashed lines) filters. The vertical lines indicate the years 1979 (EMS) and 1992 (Maastricht Treaty).

5. Discussion and conclusions

Most of the empirical literature on currency regimes (fixed vs. floating rates, optimum currency areas, monetary unions, dollarization, etc.) evaluates the validity of various theoretical criteria as if they were constant over time. The purpose of the present paper was to demonstrate empirically that, as Frankel (1999) warned, ignoring how these criteria evolve over time can lead to seriously flawed results with significantly misleading policy implications.

To show that this is the case, we used quarterly data from the 1961:1 to 1997:4 period for 13 EU countries, and focused on two widely used criteria: (1) the relative magnitude of cyclical output shocks, and (2) their degree of synchronization. Cyclical output was obtained by detrending industrial production using the HP and BP filters. We started by imposing the time-invariance assumption, and then we relaxed it to allow for dynamically evolving parameters.

Our findings strongly show that the estimated parameters, such as relative variances and correlations, are not constant over time. On the contrary, they generally exhibit substantial variation over time. It follows that, unless the time dimension is explicitly modeled, any evaluation of monetary-union criteria will be biased. This can lead to misleading conclusions about both the desirability of forming a monetary union and the benefits and costs associated with membership for a certain economy.

An additional advantage of allowing the parameters to evolve over time, is that we can ask *when* (rather than just *if*) a monetary union should be established, and *when* (rather then just *if*) a certain country should join. This is made clear if one compares Tables 1 and 2, which give the static results, with Figs. 2 and 3, which present the time-variant estimates. Tables 1 and 2 mask the time dimension and can only report estimates averaged over a period of time. The usefulness of these numbers, while considerable, is obviously limited. Figs. 2 and 3, on the other hand, by showing how the criteria are changing over time, enable us not just to evaluate for each country the expected stabilization costs and benefits associated with membership in a monetary union, but also to determine when this cost–benefit calculation will be the most favorable.

In practical terms, one of the potential uses of the time-variant technique proposed in this paper can be to determine the optimal timing of EU enlargement. For example, only an investigation of time-variant criteria for the next wave of countries (such as Cyprus, the Czech Republic, Estonia, Hungary, Poland, and Slovenia, all of which began detailed negotiations with EU in 1998) can determine when *euro* membership would maximize the benefits and minimize the stabilization costs of giving up monetary independence for each of these economies.

Finally, it has to be acknowledged that EU and *euro* membership are political processes, involving more than strictly economic decisions. This is almost always the case with similar international arrangements, other examples of which are NAFTA, the accession of China to the WTO, and the prospect of dollarization for various Latin American countries. The fact that political issues are highly important, however, does not change the economic part of the equation. If political criteria are more prominent than economic ones, an economy may adopt the *euro* when it is not yet optimal to do so, or may be prevented from adopting it when the situation is optimal. In this case, fulfilling the economic criteria may not be a good predictor of actual membership. However, the economic effects will always depend on these criteria. Thus, whether *euro* membership will benefit or harm a country's economy depends on the economic criteria only.

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