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# Do Emu Countries Constitute an Optimum Currency Area? An Empirical Test of the Generalised Purchasing Power Parity Hypothesis

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***Abstract:*** This paper makes use of Generalised Purchasing Power Parity based on the Johansen cointegration technique to examine the issue of Optimum Currency Area (OCA) in five European Monetary System (EMS) countries: France, Greece, Italy, Portugal and Spain. It evaluates the degree of asymmetry in different sub-periods in terms of four (i.e. n-1) equilibria. The existence of n-1 equilibria implies the adoption of a common monetary policy. The empirical results indicate that the only period that these countries constituted an OCA was during 1991-1998. However, this result is the exception rather than the rule since estimation of a more extended period (i.e. 1986-1998) rejects the hypothesis of an OCA.

***JEL Classification:*** E42, F36

***Key words***: Generalized PPP, EU, Real exchange rate

## Introduction

The European economic and monetary union represents the third and final stage of a complete economic and monetary union among the countries participating in the European Union. The legal institutional and monetary aspects of European integration were laid down in Maastricht Treaty (1992). However, the Maastricht criteria contained only demand nominal convergence and it should be cautioned that not all the countries manage to fulfil the fiscal criteria. As is now evident, entry to

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EMU is not restricted to countries which exhibit considerable degree of real convergence. Yet the conditions for a well-functioning monetary union, which already date back to Mundell (1961) are set in real not in nominal terms. Mundell (1961) argues that two regions constitute an optimum currency area (OCA) if they experience similar real disturbances (Bayoumi and Taylor 1995).[1] In a nutshell this theory says that when two regions or countries are subjected to different disturbances the adjustment process will require either the real exchange rate to adjust, or factors of production to move, or fiscal redistribution or a combination of these three. In the absence of real exchange rate flexibility and factor mobility, regional and national concentration of unemployment will be inevitable. The question arises whether the member countries of the EMS constituted an OCA in traditional terms, focusing on the asymmetry of shocks experienced by potential members of the monetary union.

This paper makes use of generalise Purchasing Power Parity, proposed by Enders and Hurn (1994), to identify whether an OCA can be composed by some European Union countries. We use the Johansen maximum likelihood technique to check the number of common stochastic trends shared by the monetary authorities of potential monetary union countries. The number of common stochastic trends is an indicator of the pattern of external shocks affecting the EU countries and whether monetary policy co-ordination was successful in reducing their effects. Monetary policy co-ordination is closely related to the (n-1) problem and consequently to the credibility problem of monetary union.[2] Here, in the context of cointegration, we test the Gaussian properties of the error term. Non-normal residuals can result from expected realignment risk that may not be materialised in the sample period.[3] However, expected realignment has been used by some studies as a credibility indicator of a monetary union (see Rose and Svensson, 1994). The aim of testing for non-normal residuals is to raise the issue of credibility, which might lower the number of cointegrating vectors. Moreover, since it has been seen that the EMS is characterised by periods of stability, interrupted by periods of excess volatility, we test for the existence of OCA in four different sample periods.

The multi-period estimation has two objectives: first to test whether there were periods when EU countries constituted an OCA and second whether EMS countries constituted an OCA in the periods before and after the crises of 1992 and 1993. We investigate mainly countries that most likely do not constitute an OCA. Some empirical evidence by Masson and Taylor (1993) suggests that the EU countries present a reasonable degree of economic convergence if Greece, Portugal and Spain are not taken into account. Therefore if this block of currencies are found to comprise an OCA, then the whole euro area must be optimum right from the inception of a single currency. In this block of currencies we include these three countries along with Italy and France. We use France in this block as a key country. On the other hand, Italy is included in the first block because of its rather disappointing performance on the debt/GDP ratio, which could undermine the constitution of OCA. Thus the countries included in the first block most probably do not comprise an OCA.

In summary, the contribution of this chapter concentrates on three points. First, we estimate the degree of asymmetry in terms of n-1 equilibrium relationships, which implies the adoption of one common policy. Second, we test for the existence of OCA in different periods in order to check the effects of currency crises on the interrelationship between the fundamentals of EU countries. Finally, we use more up to date data than earlier studies.

## Data description and Generalised PPP (GPPP)

This section consists of two parts. The first describes the data, which were used in the empirical estimation. The second part explains the idea of GPPP and its usage as the basis for an analysis of OCA.

### *Data and periods under consideration*

We use monthly seasonally unadjusted data of relative prices and nominal exchange rate, where ten European Union countries are measured against the German mark. All data were taken from the International Monetary Fund's International Financial Statistics (IFS) CD-ROM. The prices are the consumer price index (CPI), IFS line 64. We use the average of the periods nominal exchange rates, quoted in units of national currencies per US dollar which were converted into units of national currencies per German mark.

We test for unit roots and cointegration in the bilateral real exchange rates of the five EMS countries, using as a base country Germany.[4] The sample periods are:

1. January 1979 to March 1998, which includes all the shocks since the inception of the exchange rate mechanism (ERM).

2. July 1986 to December 1990, where no realignment took place. This period was used to test whether the interrelationship between the real exchange rates was stronger in a relatively stable period, implying more cointegrating vectors.

3. January 1991 to March 1998. A period after the relatively stable regime, including the exchange rate crises of September 1992 and July/August 1993. The reason that we use this period is to test whether the number of cointegrating vectors has been reduced by the aforementioned crises.

4. July 1986 to March 1998, combining the last two sub-periods. This allows any small sample bias effects of the last two periods to be corrected and emphasises the

way that the two shocks in the third period affect the number of cointegrating vectors and, consequently, the credibility of the EMS.

*A Test for Generalised Purchasing Power Parity (PPP)*

There is a rich literature on the determinants of the real exchange rate. One important contribution of this research is associated with Edwards (1994) who investigated the relationship between a real exchange rate (RER) and a set of exogenous fundamental variables. The reason that we use real exchange rates is that there is a growing agreement that sustained real exchange rate misalignment generates severe macroeconomic disequilibria, and that the correction of external imbalances will require both demand management policies and real exchange devaluation.[5] However, measuring the degree of misalignment is difficult, since it requires an unobserved variable, the equilibrium real exchange rate. In terms of policy debates regarding the equilibrium real exchange rate, much discussion has focused on the concept of a fundamental equilibrium exchange rate (FEER), an explicitly normative approach which offers an appealing way of thinking about the evolution of actual and equilibrium real exchange rates. However, the problem associated with the FEER concept is one of tractability in terms of the need to have a fully specified multilateral structural model and, further, it does not provide an empirical link between the real exchange rate and its determinants. On the other hand, the analysis of a single real exchange rate by using behavioural time series methods is relatively easy and provides the links between the real exchange rate and the underlying fundamentals. Recent studies, based on multivariate cointegration methods, found strong long-run relationships between real exchange rates and a set of macro variables (MacDonald [1997)).

MacDonald (1997) decomposes the real exchange rate into factors introducing systematic trends into the behaviour of the equilibrium real exchange rate. These factors are labelled the fundamentals exclusive of real interest rate (FERID). They include variables such as net foreign asset accumulation, productivity bias and fiscal balances. In particular, MacDonald (op. cit.) used the following general expression for the equilibrium real exchange rate, $q_t^e$:

$$q_t^e = q_t^T + a_t^*(p_t^{T*} - p_t^{NT*}) - a_t(p_t^T - p_t^{NT}) \qquad (1)$$

where $q_t^T$ is the real exchange rate for traded goods, $p_t$ denotes the price level with $T$ and $NT$ denoting traded and non-traded good respectively and an asterisk denotes a foreign magnitude. Lower-case letters denote logarithms of the variables. Underlying equation 1 are three important sources of long-run real exchange rate variability:

non-constancy of the real exchange rate for traded goods; the movement of relative prices of traded to non-traded goods between home and foreign country; differing time-variability of the weights used to construct the overall prices in the home and foreign country. In the next step, MacDonald (op. cit.) presented the real exchange rate for the traded goods as a function of factors that introduce systematic variability into $q_t^T$. These factors can be summarised using the following relationship:

$$q_t^T = f(\underset{+}{FISC}, \underset{+}{PS}, \underset{?}{ROIL}) \tag{2}$$

where FISC denotes the effects of relative fiscal balances on the equilibrium real exchange rate, PS presents private sector savings and ROIL is the real price of oil. The signs below the variables summarise the long-run effects of these variables on the real exchange rate. Finally, Macdonald showed that the second term in equation (1) had the following functional form:

$$(p_t^{T*} - p_t^{NT*}) - (p_t^T - p_t^{NT}) = g(\underset{+}{PROD}, \underset{+}{DEM}) \tag{3}$$

where PROD is a measure of productivity bias and DEM denotes demand side bias. Combining equation (2) and (3) we obtain the general relationship for the equilibrium real exchange rate:

$$q_t^e = h(\underset{+}{PROD}, \underset{+}{DEM}, \underset{+}{FISC}, \underset{+}{PS}, \underset{+/-}{ROIL}) \tag{4}$$

Equation (4) was operationalised by addressing the issue of how the actual exchange rate adjusts to the long-run equilibrium exchange rate. Using real uncovered interest rate parity, MacDonald (op. cit.) defined a static relationship for the current equilibrium exchange rate in terms of FERID variables and the real interest rate differential (RID), namely,

$$q_t = q_t^e - (r_t - r_t^*) \tag{5}$$

The analysis above indicates that there are three main reasons that we use real exchange rates to test whether the EU countries constitute an optimum currency area. The first reason is that the fundamental variables, which determine the equilibrium exchange rate, have been used by policy authorities of individual countries as indicators for optimum monetary and fiscal rules. Second, in a multi-country setting real exchange rates can be interpreted in terms of OCA. In particular, a high interrelationship of the determinants of real exchange rates across the countries of a potential monetary union implies that the four criteria of OCA have been accomplished. The third reason is associated with the econometric method and it was partly explained by MacDonald (1987). In particular the econometric method should permit flexible dynamic adjustment of the real exchange rates toward the equilibrium real exchange rate and it should allow for the influence of short to medium-run macroeconomic and exchange rate policy on the real exchange rate. The last two prerequisite conditions of the econometric method provide the fundamental advantage of the Johansen cointegration technique, enabling the derivation of the simple empirical framework from a much more complicated theoretical model.[6]

For each country $j$, the long-run relationship between its RER against the base country $1$ and the fundamentals is defined as follows:

$$r_{1t} = x_{jt}'\beta_j + \varepsilon_{jt} \tag{6}$$

where $r_{1t}$ denotes the real exchange rate, $x_{jt}$ is the vector of fundamentals, $\beta_j$ is the vector of coefficients, and $\varepsilon_{jt}$ is the stationary error term. There are $n$ independent real exchange rates among $n+1$ countries within the potential currency area under consideration.

Macroeconomic variables in $x_{jt}$ are productivity shocks, the real interest rates and government expenditures. If all of them are stationary then the real exchange rate will be stationary and PPP will hold even in the short run. However, if one of the variables in $x_{jt}$ is not stationary, PPP will no longer hold. In the latter case a long-run relationship between the real exchange rate and the set of fundamentals means that the two sides of (1) are cointegrated. Therefore, if a variable is stationary, it should be omitted without affecting statistically the cointegrating relationship.

The generalised Purchasing Power Parity (GPPP) hypothesis focuses on the interrelationships among the n set of $x_{jt}$. Assuming that each vector contains the same set of $m$ variables that are non-stationary, and stacking $n$ independent representations of $r_{1jt}$ together, we get in a compact form

$$Q_t = \beta X_t + E_t \tag{7}$$

where $Q_t$ is the (n×1) vector of real exchange rates, $\beta$ is the (n×m) matrix of the coefficients of fundamentals variables, $X_t$ is the (m×1) vector of fundamental variables and $E_t$ is the (n×1) vector of the error terms. Although each element in $X_t$ can follow a non-stationary process, the rank of matrix $\beta$ indicates the behaviour of the fundamental macroeconomic variables within the economic region. To be more precise, if the rank is zero, that is every element in $\beta$ is equal to zero, then PPP holds for every bilateral real exchange rate. If $\beta$ has full rank there is no long run relationship among the $n+1$ economies. Moreover, if the rank of β equals unity, it means that all real exchange rates share one single common trend. In general, Enders and Hurn (1994) show that as long as the rank of β is less or equal to $n-1$, there is a linear combination of real exchange rates that is stationary. If the rank is equal to $n-1$ then the cointegrating vector is unique. When real exchange rates are cointegrated, implying a sufficient interrelationship among their underlying economies, GPPP will hold and the set of countries can be considered as satisfying one necessary condition for comprising a potential optimum currency area. The empirical estimation will adopt the suggestion of Serletis et al. (1992), which requires $n-1$ cointegrating vectors between the $n$ countries of a monetary union to constitute an OCA. Therefore, we will pay more attention to the case where the rank of the matrix is equal to one.

When GPPP holds the real exchange rate between country $j$ and the base country can be written as a weighted average of the other real exchange rates in the currency area:

$$r_{12t} = a_0 + a_{13} r_{13t} + a_{14} r_{14t} + \ldots + a_{1n} r_{1nt} \qquad (8)$$

The weights are functions of the parameters in matrix β that represents linkages among the economies. They reflect trade linkages, but also broader linkages such as technology transfers, immigration and financial resource movements.

## Empirical Results

This section describes the results of cointegration analysis for the five EMU countries real exchange rates.

### *Cointegration analysis of the five EU real exchange rates over the whole period: January 1979 to March 1998*

We examine a currency block that includes the following countries: Greece, Portugal, Spain, Italy and France. The reason that we include these countries in this block of currencies is that according to Masson and Taylor (1993) the EMS countries indicated a significant degree of convergence if Greece, Portugal and Spain were left out. In particular, they took real output per capita (as percentage of national average) as the measure of economic performance and presented some evidence on the degree of real convergence during 1960 to 1990 in a monetary union – the US – and the European Union. They show the cross-sectional standard deviation of regional real product per capita for US regions and of real GDP per capita for EC members, including and excluding Greece, Portugal and Spain, relative to the mean. The result for the countries of the EU show that much of the dispersion in real economic performance across the EU is due to the southern countries – i.e. Greece, Portugal and Spain. Excluding these countries, the standard deviation has remained constant at 12 percent since the early 1960s, but including them causes the standard deviation to rise to between 25 percent and 30 percent.

We include these countries in a currency block along with Italy and France. Weber (1991) and Gros and Tygesen (1998) show that the EMS consists of two currency blocks: a 'hard currency block', which included the German mark and the Dutch guilder, and a 'soft currency block' which included other currencies and was centred around the French franc. Thus, we use France in the first block of currencies as the centre country. On the other hand, we include Italy in this block because of its high debt/GDP ratio, which could undermine the structure of an OCA. Apparently, the countries included in the first block most probably do not comprise an OCA.

The cointegration analysis in the block of currencies under consideration commences from an augmented VAR with 8 lags. Using two dummies variables to average out regime shifts, removes the problem of non-normally distributed residuals.[7] The first dummy represents the effects of the eleven realignments that happened during the period January 1979 to July 1986. The second dummy represents the effects of the exchange rate crises in September 1992 and July 1993. Further, dummy variables are used to proxy outliers found in each individual equation.

Diagnostic tests[8] for the block of EU currencies, see Table 1 below, show that there is only a problem of non-normal residuals in the equation for France, Greece and Italy. Similar vector tests also report that there is problem of non-normality. From an economic point of view, the problem of non-normality can be the result of

i) eleven realignments that took place between January of 1981 and June of 1986;

ii) the speculation that erupted in September 1992 and led to the withdrawal of sterling and the Italian lira from the ERM;

iii) the new speculative crisis that took place one year later, involving mainly the French franc but also the peseta, the Belgian franc, and the Danish korone. The last speculative crisis led the EU Ministers of Finance to increase the fluctuation bands to +15 per cent and -15 per cent. Finally, the three devaluations of the Greek Drachma, two took place in the first period between January 1979 and July 1986 and the third in March of 1998, might have affected the residual normality of at least of its own equation. Non-normality of residuals of the VAR may indicate the existence of risk arising from expected realignment before the crisis of September 1992, August 1993 and the stormy period of March 1995. The fact that realignments did not materialise in the sample period implies that the problem of non-normality can be due to three factors. [9]Firstly, the distribution of the residuals may have a heavy tail; secondly the distribution may be heavily skewed. In both of these cases the convergence of the residual distribution to normality may be slow. The third and the most important reason is that, even if the markets are rational and efficient, the residuals are uncorrelated but not independent. In this case the moment generating function of the residual sequence does not exist and consequently the central limit theorem is invalid. It is clear that the problem of non-normality might be the result of the 'peso problem'' (see Krasker 1980).[10]

Table 1: Goodness of Fit Evaluation of the Block of EU Currencies.
Period: January1979 to March 1998.

| Statistic | France | Greece | Portugal | Spain | Italy | VAR |
|---|---|---|---|---|---|---|
| $\sigma$ | 0.0081 | 0.0178 | 0.0181 | 0.0149 | 0.0130 | |
| $F_{ar}$ (7, 79) | 1.18 [0.144] | 2.09 [0.053] | 1.01 [0.426] | 0.38 [0.906] | 0.77 [0.612] | |
| $F_{arch}$(7, 72) | 0.42 [0.884] | 0.26[0.963] | 0.58 [0.768] | 0.16 [0.991] | 1.62 [0.141] | |
| $F_{het}$(82, 3) | 0.10 [1.000] | 0.02 [1.000] | 0.02[1.000] | 0.02 [1.000] | 0.02 [1.000] | |
| $\chi^2_{nd}$ (2) | 11.6[0.003]** | 13.3 [0.001]** | 7.24 [0.026] | 3.08 [0.214] | 12.5 [0.001]** | |
| $F^v_{ar}$(175, 238) | | | | | | 1.18 [0.111] |
| $\chi^{2v}_{nd}$(10) | | | | | | 44.4 [0.000]** |
| $\chi^2$(1230) | | | | | | 1047 [0.999] |

Notes: 1) the value inside the bracket denotes the p values of the null hypothesis to be correct. 2) An asterisk indicates rejection of the null hypothesis at 5 per cent level. Two asterisks indicate rejection of the null hypothesis at 1 per cent level.

Table 2 presents a test for cointegration in the block of the five EU real exchange rates. This test for cointegration is based on the maximum eigenvalue $\lambda_{max}$ and trace $\lambda_{trace}$ statistics.[11] The $\lambda_{max}$ supports the hypothesis that there are 2 cointegrating vectors while the $\lambda_{trace}$ statistics supports the hypothesis that there is only one cointegrating vector.[12] Moreover, adjusting for degrees of freedom, cointegration tests do not reject the null hypothesis of no-cointegration.[13] Thus, in the first period between January 1979 and March 1998 the five EU countries that belong to the first block do not constitute an optimum currency area.

Table 2: Optimal lags and rank determination of the five EU Real Exchange Rates of the EMU; period: January 1979 to March 1998

Optimal lags **k***

| ENDG | D(FR) | D(IT) | D(GR) | D(PR) | D(SP) | D(CR) | D(RL) |
|---|---|---|---|---|---|---|---|
| 8 | 12 | 12 | 12 | 12 | 12 | 12 | 12 |

| $H_0$:ran=p | -Tlog(1-μ) | using T-m | 95 per cent | -TΣlog(1-μ) | using T-nm | 95 per cent |
|---|---|---|---|---|---|---|
| p<=0 | 40.91* | 33.44 | 37.5 | 91.41* | 74.71 | 87.3 |
| p<=1 | 37.61* | 30.74 | 31.5 | 50.5 | 41.28 | 63.0 |
| p<=2 | 9.105 | 7.442 | 25.5 | 12.9 | 10.54 | 42.4 |
| p<=3 | 2.729 | 2.23 | 19.0 | 3.793 | 3.1 | 25.3 |
| p<=4 | 1.064 | 0.869 | 12.2 | 1.064 | 0.86 | 12.2 |

Eigenvalues: (0.170, 0.157, 0.040, 0.012, 0.004)

Notes: 1) D, CR, and RL stand for Dummy, Crisis and Realignment respectively. 2) The country codes are: FR, France; IT, Italy; GR, Greece; PR, Portugal; SP, Spain.

*Cointegration analysis of the five EU real exchange rates in the first sub-period: July 1986 to December 1990*

In a period where no realignment took place (i.e. July 1986 to December 1990) diagnostic tests for the five real exchange rate of the EU countries indicate that there is no problem of non-Gaussian error terms (see Table 3 below). Under the condition of Gaussian error terms, cointegration tests for the first block of currencies justifies the existence of three cointegrating vectors and leaves open the possibility that there are four cointegrating vectors. In particular, the magnitude of the fourth eigenvalue is quite large and close to being significant at the 5 per cent level. Four cointegrating vectors for the first block of currencies indicate the existence of an optimum currency area. However, adjusting for degrees of freedom, following Reimers (1992) with critical values derived from Osterwald-Lenum (1992), tests for cointegration support the hypothesis of one cointegrating vector (see Table 4 below). The last result indicates, in terms of common stochastic trend, that even in a relatively tranquil period the five EU countries of the first block do not constitute an OCA. The small sample bias can be confirmed as the inclusion of a dummy, proxying German unification in the first period, was rejected. The data of real exchange rates cannot detect the significance of political incidents. It is sensible, before reaching any conclusions about the suitability of some countries to constitute an OCA, to take into account that one of the significant factors (i.e. German unification) that led to the crisis of 1992 cannot be investigated statistically. Consequently, more specialised models about political risk should be applied to test for the significance of particular incidents. Furthermore, it has to be emphasised that the eleven realignments that took place between 1979 and July 1986 do not have any effect on the data generation process (DGP) in the second period. Therefore, expectations of realignment can be detected only backwards (ex-post) and not forward (ex-ante) at least in the context of cointegration.[14]

Table 3: Goodness of fit evaluation of the block of EU currencies.
First Sub-period: July 1986 to December 1990.

| Statistic | France | Greece | Portugal | Spain | Italy | VAR |
|---|---|---|---|---|---|---|
| $\sigma$ | 0.0060 | 0.0101 | 0.0069 | 0.0089 | 0.0049 | |
| $F_{ar}$ (2, 15) | 0.27 [0.763] | 9.35 [0.002]** | 0.16 [0.852] | 3.62 [0.052] | 0.17 [0.844] | |
| $F_{arch}$(2, 13) | 0.34 [0.716] | 0.24 [0.785] | 0.007 [0.992] | 0.45 [0.645] | 0.17 [0.837] | |
| $\chi^2_{nd}$ (2) | 1.60 [0.448] | 1.16 [0.558] | 0.64 [0.725] | 1.37 [0.501] | 0.34 [0.716] | |
| $F^v_{ar}$ (50, 17) | | | | | | 1.26 [0.307] |
| $\chi^{2v}_{nd}$(10) | | | | | | 3.95 [0.944] |

Notes: 1) the value inside the bracket denotes the p values of the null hypothesis to be correct. 2) An asterisk indicates rejection of the null hypothesis at 5 per cent level. Two asterisks indicate rejection of the null hypothesis at 1 per cent level.

Table 4: Cointegration analysis in the block of monthly Real Exchange Rates of the
EU; period: July 1986 to December 1990.

Optimal lags **k***

| ENDG | D(FR) | D(IT) | D(GR) | D(PR) | D(SP) | D(CR) | D(RL) |
|---|---|---|---|---|---|---|---|
| 7 | - | - | - | - | - | - | - |

| *$H_0$:rank=p* | *-Tlog(1-μ)* | *Using T-nm* | *95 per cent* | *-TΣlog(1-μ)* | *using T-nm* | *95 per cent* |
|---|---|---|---|---|---|---|
| *p<=0* | 125.6** | 44.2** | 37.5 | 264.5** | 93.06* | 87.3 |
| *p<=0* | 71.02** | 24.99 | 31.5 | 138.9** | 48.86 | 63.0 |
| *p<=0* | 43.48** | 15.3 | 25.5 | 67.85** | 23.87 | 42.4 |
| *p<=0* | 13.49 | 4.747 | 19.0 | 24.36 | 8.573 | 25.3 |
| *p<=0* | 10.87 | 3.82 | 12.2 | 10.87 | 3.826 | 12.2 |

Eigenvalues: (0.902, 0.731, 0.553, 0.221, 0.182)

Notes: 1) D, CR, and RL stand for Dummy, Crisis and Realignment respectively. 2) The country codes are: FR, France; IT, Italy; GR, Greece; PR, Portugal; SP, Spain. 3) Dash indicates that the dummy or the variable under consideration is not present.

*Cointegration analysis of the five real exchange rates in the second sub-period: January 1991 to March 1998*

In the period between January 1991 and March 1998, where the three currency crises of the EMS took place, diagnostic tests for the block of currencies that we examine show that there is a problem of non-normality both for the individual equations and with the VAR (see Tables 5 below). To be more precise, there is a problem of non-normal residuals in the equations of Greece and Portugal.

Table 5: Goodness of fit evaluation of the block of EU currencies.
Period: January 1991 to March 1998.

| Statistic | France | Greece | Portugal | Spain | Italy | VAR |
|---|---|---|---|---|---|---|
| $\sigma$ | 0.0057 | 0.0149 | 0.0112 | 0.0089 | 0.0135 | |
| $F_{ar}$ (6, 31) | 1.74 [0.144] | 1.75 [0.140] | 2.00 [0.095] | 0.75 [0.607] | 0.49 [0.808] | |
| $F_{arch}$(6, 25) | 0.70 [0.651] | 0.06 [0.998] | 0.23 [0.960] | 0.46 [0.828] | 0.33 [0.909] | |
| $F_{het}$(82, 3) | 0.20 [0.999] | 0.18 [0.999] | 0.18 [0.999] | 0.18 [0.999] | 0.27 [0.999] | |
| $\chi^2_{nd}$ (2) | 0.07 [0.961] | 15.9 [0.003]** | 21.9 [0.000]** | 0.47 [0.788] | 3.27 [0.194] | |
| $F^v_{ar}$(150, 19) | | | | | | 1.85 [0.060] |
| $\chi^{2v}_{nd}$ (10) | | | | | | 49.6 [0.000]** |
| $F^v_{het}$(330, 47) | | | | | | 0.06 [1.000] |

Notes: 1) the value inside the bracket denotes the p values of the null hypothesis to be correct. 2) An asterisk indicates rejection of the null hypothesis at 5 per cent level. Two asterisks indicate rejection of the null hypothesis at 1 per cent level.

Table 6a below reports tests for cointegration in the first block of real exchange rates for the third period and supports the hypothesis of four cointegrating vectors. Adjustment for degrees of freedom does not change the number of cointegrating vectors. Compared to the second period, where the hypothesis of four cointegrating vectors has also been conditionally accepted, the magnitude of the eigenvalues in the third period is significantly lower. Thus, the correlation between the stationary part of the equilibrium correction model (i.e., $Dy_t$) and the linear combination of I(1) series ($ab\check{}y_{t-1}$) is lower in the third period than the corresponding correlation in the tranquil second period. The fact that there are four cointegrating vectors in a five-variable system and that the number of cointegrating vectors has not been reduced compared to those in the second period, implies that these five real exchange rates encompass an optimum currency area in the third period. However, the existence of non-normal residuals undermines the credibility of the inter-relationship between the real exchange rates and a more cautious inquiry is required in a larger sample, where problems of small sample bias can be avoided. Therefore, we need to use additional information to determine the rank of cointegrating space. We look at the five largest eigenvalues of the companion matrix and the graphic analysis of cointegrating vectors and of recursive eigenvalues (see Table 6b below).

Table 6a: Cointegration analysis in the five EU Real Exchange Rates of the EMS. Second Sub-period: January 1991 to March 1998

Optimal lags **k***

| ENDG | D(FR) | D(IT) | D(GR) | D(PR) | D(SP) | D(CR) | D(RL) |
|---|---|---|---|---|---|---|---|
| 2 | 10 | 0 | 6 | 6 | 0 | 10 | - |

| *$H_0$:rank=p* | *-Tlog(1-m)* | *using T-nm* | *95 per cent* | *-Tlg(1-m)* | *using T-nm* | *95 per cent* |
|---|---|---|---|---|---|---|
| *P* | 61.83** | 54.72** | 37.5 | 166.77** | 147.6** | 87.3 |
| *P* | 49.19** | 43.54** | 31.5 | 104.9** | 92.84** | 63.0 |
| *P* | 25.67** | 22.72 | 25.5 | 55.71** | 49.3** | 42.4 |
| *P* | 21.36** | 18.91 | 19.0 | 30.04* | 26.58** | 25.3 |
| *P* | 8.673 | 7.676 | 12.2 | 8.673 | 7.676 | 12.2 |

Eigenvalues: (0.508, 0,431, 0.255, 0.217, 0.094)
Notes: 1) D, CR, and RL stand for Dummy, Crisis and Realignment respectively. 2) The country codes are: FR, France; IT, Italy; GR, Greece; PR, Portugal; SP, Spain. 3) Dash indicates that the dummy or the variable under consideration is not present.

Table 6b: Eigenvalues of the Companion Matrix

| *The five largest eigenvalues* | Real | Complex | Modulus |
|---|---|---|---|
| | 0.992 | 0.000 | 0.992 |
| | 0.958 | 0.000 | 0.958 |
| | 0.859 | 0.000 | 0.859 |
| | 0.436 | 0.607 | 0.748 |
| | 0.436 | -0.607 | 0.748 |

The moduli of the two largest eigenvalues are 0.9929 and 0.9582 respectively, indicating that there are three cointegrating vectors. However, graphic analysis of the cointegrating vector and recursive eigenvalues suggests four cointegrating relationships (these graphs are not presented here but can be provided by the author upon request). In particular, graphic cointagration analysis show that there are four stationary cointegrating vectors and four recursive eigenvalues different from zero

In the next step, we deal with the problem of non-normal residuals by testing the real exchange rates of Portugal and Greece for weak exogeneity with respect to the cointegrating vectors. Under this restriction the first block could constitute an OCA if Portugal and Greece are left out. Exogeneity tests in Table 7 below reject the hypothesis of weak exogeneity both for Greece and for Portugal. Moreover, significance tests of these two countries in the cointegrating space indicates that Portugal and Greece must be included in every unique cointegrating vector (see Tables 8 and 9 below). In general, we can accept that in the period January 1991 to March 1998 there are four cointegrating vectors and consequently there is a strong relationship between the fundamental variables across the countries of the first block. The last assessment leads to the conclusion that even in a period where two currency crises took place the countries in the first block comprise an OCA.

Table 7: Weak Exogeneity Test With Respect to the Cointegrating Vectors.

| Exogeneity test | rank=4 $\chi^2(4)$ | LR test | p-value |
|---|---|---|---|
| Greece | | 11.24 | [0.0240]* |
| Portugal | | 35.25 | [0.0000]** |

Table 3.8: Testing Unique General Restrictions on Cointegrating Vectors: Significant Test for Greece.

| TEST | Restricted β | | | | | LR test | p-value |
|---|---|---|---|---|---|---|---|
| | France | Greece | Portugal | Spain | Italy | | |
| $H_{4.1}$ $\beta_1$ | 1.000 | 0.000 | 0.540 | -0.701 | 0.502 | $\chi^2(4)$=32.434 | [0.0000]** |
| $\beta_2$ | 2.624 | 0.000 | 0.268 | 0.961 | -1.201 | | |
| $\beta_3$ | 1.059 | 0.000 | 1.000 | -0.894 | -0.021 | | |
| $\beta_4$ | -3.411 | 0.000 | -0.207 | 1.000 | 0.199 | | |

Table 3.9: Testing Unique General restrictions on cointegrating vectors: significant test for Portugal.

| TEST | Restricted β | | | | | LR test | p-value |
|---|---|---|---|---|---|---|---|
| | France | Greece | Portugal | Spain | Italy | | |
| $H_{4.2}$ $\beta_1$ | 1.000 | -0.277 | 0.000 | -0.380 | 0.139 | $\chi^2(4)$=40.993 | [0.0000]** |
| $\beta_2$ | 3.471 | 1.000 | 0.000 | -0.551 | -0.839 | | |
| $\beta_3$ | 0.908 | -10.59 | 0.000 | 4.055 | 1.279 | | |
| $\beta_4$ | 35.48 | -12.78 | 0.000 | 1.000 | 2.340 | | |

*Cointegration Analysis of Real Exchange Rates Over the Whole Sub-Periods: July 1986 to March 1998*

In the last period we pool the second and the third sub-periods to investigate whether shocks, for example the exchange rate crisis in 1992 and expectation of realignments, affect the number of cointegrating vectors in a larger sample. Equally, we try to test if there is loss of information in the two sub-samples which may have led to a higher number of cointegrating vectors detected than really existed. Table 10 below reports that there is a problem of non-normality in the equations of Greece and in the system. Tests for cointegration, even after adjusting for degrees of freedom, see Table 11a, suggest that there are 2 cointegrating vectors, two less than those found in the sub-periods (i.e., the second and third period). The eigenvalues of the companion matrix in Table 11b below indicates that there are three cointegrating vectors. However, the moduli of the largest root of the companion matrix are outside the unit circle implying that one of the variables is I(2) (see Harris, 1995). Graphs of first difference of the series indicates that non-stationarity must be the result of outliers, which induce a unit root in the variance, rather than of a stochastic trend in the first difference of the series (these graphs are not presented here but are available upon request by the author). Unit root tests for the first difference of the real exchange rates included in the first block of countries reject the hypothesis of non-stationarity (see Table 12 below). Therefore, these variables are not I(2). Moreover, the hypothesis of a trending variance due to parameter instability caused by structural changes can be rejected by tests of parameter stability (Graphic presentation of these tests are not available here but can be provided by the author upon request). Problems of parameter instability can only be justified in the equations of Greece and Italy, a result of the devaluation of the Greek drachma in March 1998 and of the Italian lira in September 1992. In general, we ascribe the explosive eigenvalue of the companion matrix to outliers rather than to I(2) variables and to a trending variance. In the next step, we test whether the Greek real exchange rate is weakly exogenous with respect to the cointegrating vectors. Table 13 below presents tests, which reject the hypothesis of weak exogeneity for Greece. Therefore, no improvement in the stochastic properties of system can be achieved by conditioning on the Greek real exchange rate.

Three cointegrating vectors in a five-variable system imply that the EU countries of this group do not constitute an OCA. Finally, the magnitude of the eigenvalues in the period July 1986 to March 1998 is substantially smaller than those eigenvalues in the second and third period. The last result implies that there is a small sample bias that increases the number of cointegrating vectors in the sub-periods. The fact that the effects of expectations can be detected only backward (ex-post) undermines the credibility of the results based on a small sample.

Table 10: Goodness of fit evaluation of the block of EU currencies.
Third Sub-period: July 1986 to March 1998.

| Statistic | France | Greece | Portugal | Spain | Italy | VAR |
|---|---|---|---|---|---|---|
| $\sigma$ | 0.0069 | 0.0134 | 0.0110 | 0.0131 | 0.0147 | |
| $F_{ar}$ (7, 52) | 0.54 [0.796] | 1.16 [0.336] | 0.93 [0.486] | 2.76 [0.016] | 0.49 [0.831] | |
| $F_{arch}$ (7, 45) | 0.46 [0.857] | 0.18 [0.987] | 0.37 [0.914] | 0.52 [0.809] | 0.61 [0.741] | |
| $F_{het}$(82, 3) | 0.08 [1.000] | 0.03 [1.000] | 0.14 [1.000] | 0.05 [1.000] | 0.07 [1.00] | |
| $\chi^2_{nd}$ (2) | 2.57 [0.276] | 25.4 [0.000]** | 0.287 [0.866] | 0.44 [0.798] | 4.52 [0.104] | |
| $F^v_{ar}$ (150, 19) | | | | | | 1.01 [0.4736] |
| $\chi^{2v}_{nd}$ (10) | | | | | | 57.9 [0.000]** |
| $\chi^2$(780) | | | | | | 733.7 [0.880] |

Notes: 1) the value inside the bracket denotes the p values of the null hypothesis to be correct. 2) An asterisk indicates rejection of the null hypothesis at 5 per cent level. Two asterisks indicate rejection of the null hypothesis at 1 per cent level.

Table 11a:Optimal lags and rank determination of the five EU real exchange rates.
Period: July 1986 to December 1998

Optimal lags **k***

| ENDG | D(FR) | D(IT) | D(GR) | D(PR) | D(SP) | D(CR) | D(RL) |
|---|---|---|---|---|---|---|---|
| 5 | 6 | 6 | 6 | 6 | 6 | 12 | 6 |

| *$H_0$:rank=p* | *-Tlog(1-μ)* | *using T-nm* | *95 per cent* | *-TΣlog(1-μ)* | *using T-nm* | *95 per cent* |
|---|---|---|---|---|---|---|
| *p<=* | 75.03** | 61.72** | 37.5 | 155.2** | 127.7* | 87.3 |
| *p<=* | 48.48* | 39.89** | 31.5 | 80.2* | 65.98* | 63.0 |
| *p<=* | 18.9 | 15.55 | 25.5 | 31.7 | 26.08 | 42.4 |
| *p<=* | 10.52 | 8.657 | 19.0 | 12.81 | 10.54 | 25.3 |
| *p<=* | 2.284 | 1.879 | 12.2 | 2.284 | 1.879 | 12.2 |

Eigenvalues: (0.412, 0.291, 0.125, 0.071, 0.01)

Table 11b: Eigenvalues of the Companion Matrix

| The five largest eigenvalues | Real | Complex | Modulus |
|---|---|---|---|
| | 1.020 | 0.000 | 1.020 |
| | 0.985 | 0.000 | 0.985 |
| | -0.411 | 0.764 | 0.868 |
| | -0.411 | -0.764 | 0.868 |
| | 0.851 | -0.148 | 0.864 |

Notes: 1) D, CR, and RL stand for Dummy, Crisis and Realignment respectively. 2) The country codes are: FR, France; IT, Italy; GR, Greece; PR, Portugal; SP, Spain.

Table 12: Test for unit root in the first difference of real exchange rates

| Country | ADF t-stat | Lags |
|---|---|---|
| France | -9.594** | 1 |
| Italy | -7.898** | 1 |
| Greece | -9.342** | 1 |
| Portugal | -8.05** | 1 |
| Spain | -7.679 | 1 |

Note: Critical values for the individual countries are from Fuller 1976

Table 13: Weak Exogeneity Test With Respect to the Cointegrating Vectors

| Exogeneity test | rank=3 $c^2(3)$ | LR test | p-value |
|---|---|---|---|
| Greece | | 15.338 | [0.0015]** |

## Conclusions

This paper has investigated the issue of OCA in a block of currencies of EMU countries. In agreement with the empirical evidence previously established, we conclude that the countries in question do not constitute an OCA. In particular, Masson and Taylor (1993) argue that EMU countries could be an OCA if Greece, Portugal and Spain were left out. Therefore, if a block of countries in EMU including the three countries that comprise an OCA, then EMU is optimum from the inception of a single currency. We use a GPPP based on Johansen cointegration method to estimate the issue of OCA. We focus on the number of common stochastic trends in terms of n-1 equilibria. In case that there are more than one common trend (i.e. less than n-1 equilibria) then an individual country emphasises more domestic issues like growth and unemployment than issues pursued by a common monetary policy. Moreover, since the EMS characterised by periods of stability interrupted by periods of excess volatility, we examine the question of OCA in four different periods.

Empirical evidence from cointegration analysis indicate that the hypothesis of an OCA, in terms of n-1 cointegrating vectors, has been rejected for all periods except for the third period. In particular, during the period January 1991 to March 1998, the results from cointegration analysis indicate that the hypothesis of OCA cannot be rejected. However, this conclusion might be the result of small sample bias effects since the estimation of the last period, which include the second and the third period, shows that the countries under consideration do not constitute an OCA.

## NOTES

[1] For a more detailed analysis see De Grauwe (1997).

[2] The credibility problem in terms of expected realignment does not belong to the scope of this paper and for a relevant study see Gomez-Puig and Montalvo (1997).

[3] Frankel (1997) wrote that if there is a drastic event which has a small probability, $\delta$, of occurrence in a given month as $T \rightarrow \infty$ this event will occur $100\delta$ percent of time. However, in a moderate sized sample, it is possible that this event will not occur at all, thereby making it seen as though expectations were biased. This is actually description how Type I error (i.e. the probability of rejecting a hypothesis when is true) and is not an argument that the test was inappropriate. Krasker (1980) explains why the sampling distributions might not be normal. The third and most serious reason was that even if the market is rational and efficient the error (i.e. the risk premium) are usually not independent, as required by the central limit theorem, but only uncorrelated.

[4] Test for unit root in real exchange rates are not presented in this paper but can be provided by the author upon request.

[5] Diaz-Alejandro (1984), drawing on the experience of Latin America, argued that real exchange misalignment and especially overvaluation can be detrimental to an export-oriented development strategy. Caballero and Corbo (1989) emphasise the importance of real exchange rate stability for export stability and export promotion, while Servent and Solimano (1991) found real exchange rate stability to have a significant positive effect on private investment.

[6] For a more formal description of the advantages of cointegration technique to estimate Equilibrium Real Exchange Rate (ERER) models see Meese and Rogoff (1988), Edison and Dianne (1993) and Elbadawi (1994).

[7] When regime shifts are known (like the exchange rate realignments and the two currency crises in the EMS) then they can be treated by simple dummy variables. However, in the case where the regime shifts are unknown then we have to make assumptions about the process that an unobserved state variable follows. In such a case many authors (see Kim and Nelson 1999) have used a Markov-switching model.

[8] Diagnostic test involves F-tests for the hypothesis of serial autocorrelation ($F_{ar}$ against seventh-order autocorrelation), Autoregressive Conditional Heteroscedasticity ($F_{arch}$ against seventh-order), Heteroscedasticity ($F_{het}$), and lastly a $x^2$-test for normality ($x^2_{nd}$) Similar vectors test are also reported.

[9] The sample under consideration (first period) includes all the realignments that have taken placed since the inception of the EMS (1979). In this case the problem of non-normality can be the result of a regime-shifts that occurs in each realignment. The regime here represents the high volatility stage before the realignment and the low volatility stage after the realignment. The regime itself can be the outcome of an unobserved Markov chain. In such a case the distribution of the error term of real exchange rates is known as an i.i.d. mixture distributions and the assumption of normality is wrong. More precisely, we can index the process that the regime follows at date $t$ by an unobserved random variable $s_t$ where there are $M$ possible regimes. In the case of expected realignment of nominal exchange rates in the EMS we assume that there are two states, the high volatility and low volatility state. Hence, when the process is in regime *1*, the observed variable of the real exchange rate is presumed to have been drawn from an $N(\mu_1, \sigma_1^2)$ distribution. If the process is in regime *2*, and then the real exchange rate is drawn form an $N(\mu_1, \sigma_1^2)$ distribution, and so on. Therefore, the density of the real exchange rates is a mixture of densities and conditional on the value of high or low volatility state that has been taken on by the random variable $s_t$. The density function of real exchange rate can be written as $f(y_t|s_t=j;q)=1/(2p)^{1/2}s_j\,exp\{-(y_t-m_j)^2/2s_j^2\}$ where $y_t$ denotes the real exchange rate, $s_t$ denotes the state variable and q is a vector of population parameters the includes $\mu_1\ldots\mu_M$ and $\sigma_1,\ldots,\sigma_M$.

[10] The 'peso problem' suggests that a perceived small probability of a large discrete change in the exchange rate (such as expected realignment), which does not materialise in-sample will induce serial dependence into the forecast errors.

[11] The trace statistic can be written as: $\lambda_{trace} = -T\sum_{i=r+1}^{n}\ln(1-\hat{\mu}_i)$ where $\hat{\mu}_i$ is the estimate eigenvalues of the long-run equilibrium matrix and r=0, 1, 2...n-1. The trace statistic tests the null hypothesis that the number of dinstict cointegrating vectors is less than $r$ against a general alternative. The maximum eigenvalue statistic can be written as: :$\lambda_{max} = \ln(1-\hat{\mu}_{r+1})$. This statistic tests the null hypothesis that the number of cointegrating vectors are $r$ against the alternative of $r+1$.

[12] Critical values for the trace and maximum eigenvalues have been tabulated by *inter alia* Johansen (1988), Johansen and Juselius (1990) and Osterwald-Lenum (1992). However, if dummy variables enter the deterministic part of the multivariate model, then these critical values are only indicative.

[13] The problem of small samples has also been mentioned, and Reimers (1992) suggests that in such a situation the Johansen procedure over-reject when the null is true. Reimers suggests taking into account the number of parameters to be estimated in the model and making an adjustment for degrees of freedom by replacing $T$ in λmax and λtrace by $T$-$n\kappa$, where $T$ is the number of observations in the sample, $n$ is the number of variables in the model and κ is the number of lag-length set when estimating the model. However, Doornik and Hendry (1994) suggest that it is unclear yet whether this is the preferred correction. Furthermore, Cheung and Lai (1993) argue that the finite sample bias of Johansen tests is a positive function of $T/(T-n\kappa)$ and the finite-sample bias towards over-rejection of no cointegration hypothesis enlarges with increasing values of $n$ and κ. Cheung and Lai noted that an equivalent way to make finite-sample corrections is to adjust the critical values and not the test statistics. The scaling factor that they have used to adjust the critical values is $T/(T-n\kappa)$.

[14] Modern econometric techniques can calculate expectations for realignment ex-ante based on the inferred probabilities derived from the estimation of a Markov-switching regime model. Gomez-Puig

and Montalvo (1997), using euro-interest rate differentials between the interest rate of EMS countries and German interest rate, show that the probability of high volatility state (i.e. the probability of expected realignment) was higher more than one month before the crisis of 1992 for the Spanish peseta, the Portuguese escudo and the Italian lira. Recently, using short-term borrowing between financial institution interest rates, Mouratidis (2001) show that the probability of a high volatility state before the crisis of 1992 was high only for the Italian lira. For the other currencies, both in our estimation and in the estimation of Gomez-Puig and Montalvo (op. cit), the probability of realignment was contemporaneous to the crisis. Thus, beside the experience of previous realignments a more profound analysis about the nature of currency crisis is necessary to make any conclusions about the factors, which led to the crises of 1992 and 1993

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