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Structural changes and deviations from the PPP within the Euro Area

by

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ABSTRACT

This paper focuses on macroeconomic interdependencies among the euro area member countries over the period 1984-2002. The theoretical framework builds on the generalized purchasing power parity hypothesis, which is empirically tested using vector error correction models with broken deterministic components. The euro area turns out to be an integrated entity, even if national economies still exhibit a certain degree of heterogeneity. The results also suggest that up to now the “euro-effect” in fostering integration within the euro area has been quite weak.

Keywords: Euro area, purchasing power parity, cointegration, structural breaks.

JEL Classification: C32, E31, F36.

NON-TECHNICAL SUMMARY

Since January 1999 European countries joining the third stage of the Economic and Monetary Union (EMU) have shared a common currency and monetary policy. In this respect, the EMU can be considered the best real world approximation of what scholars usually define as an optimal currency area (OCA). The desirability for a given country to join an OCA is generally assessed by a sort of "cost-benefit analysis", allowing to evaluate whether advantages prevail over disadvantages with respect to the country's structural characteristics. Nevertheless, the theory of OCAs does not provide any formal test to evaluate the "optimality" of timing and modalities of implementation of a currency area.

This work focuses on the process of price convergence within the EMU. Such convergence represents a necessary condition in order to stabilize both the nominal (explicit policy target) and the real exchange rate (implicit policy target), allowing to safeguard member countries' intra-regional competitiveness and to avoid the incentive to implement "beggar thy neighbours" policies. Cointegration techniques are used to test the Generalized Purchasing Power Parity hypothesis, after the preliminary assessment of the stationarity of each bilateral real exchange rate. The econometric methodology is based on Vector Error Correction models with broken deterministic components, and provides with robust results because it is expressly designed to reduce the probability to erroneously reject the cointegration hypothesis due to the presence of segmented (instead of unbroken) deterministic components.

The overall picture emerging from the estimates suggests that the EMU is an integrated area, although a certain degree of heterogeneity among national aggregate demand functions still exists. Moreover, the empirical evidence suggests that the "euro-effect" in fostering the integration within the EMU has been quite weak so far. Even though a more precise assessment of the consequences at the national level arising from the integration process in Europe calls for a larger time horizon, these findings suggest that the convergence process within the euro area has not been pushed much further in recent years and that additional steps towards integration need to be done in order to properly make endogenous forces work as predicted by the theory of OCAs.

CAMBIAMENTI STRUTTURALI E DEVIAZIONI DALLA PARITA' DEL POTERE DI ACQUISTO ALL'INTERNO DELL'AREA DELL'EURO

SINTESI

Questo lavoro si concentra sull'analisi delle interdipendenze macroeconomiche esistenti tra i Paesi membri dell'area dell'euro durante il periodo 1984-2002. La struttura teorica si fonda sull'ipotesi della parità del potere di acquisto generalizzata, verificata empiricamente utilizzando modelli vettoriali a correzione del divario in presenza di componenti deterministiche segmentate. L'area dell'euro risulta un'entità integrata, sebbene le singole economie nazionali presentino ancora un certo grado di eterogeneità. I risultati suggeriscono inoltre che l'"effetto-euro" nel promuovere l'integrazione all'interno dell'area integrata europea sia stato piuttosto debole fino ad ora.

Parole chiave: Area Euro, Parità del potere di acquisto, Cointegrazione, Cambiamenti strutturali.

Classificazione JEL: C32, E31, F36.

CONTENTS

1 INTRODUCTION

Since January 1999 European countries joining the third stage of the Economic and Monetary Union (EMU) have shared a common currency and monetary policy. In this respect, the EMU can be considered as the most advanced experiment of monetary integration and represents, perhaps, the only real world approximation of what scholars usually define as an optimal currency area (OCA).

The desirability for a given country to join an OCA is generally assessed by a sort of “cost-benefit analysis”, allowing to evaluate whether advantages prevail over disadvantages with respect to the country’s structural characteristics. Potential gains are mainly related to improvements in economic efficiency, whereas potential losses are mainly related to the impossibility of using a number of instruments of macroeconomic policy in order to face asymmetric shocks (see, among others, Mongelli, 2002). Nevertheless, some recent efforts in this direction notwithstanding (see, for example, Demopoulos and Yannacopoulos, 1999), the theory of OCAs does not provide any formal criterion to evaluate whether timing and modalities of implementation of a currency area can be considered somewhat optimal (Eichengreen, 1990). Moreover, there is no widespread consensus on the effective likelihood to both observe in practice the above-mentioned potential gains and losses and clearly identify their real impact (see, for example, Baldwin, 1991; Buiter, 2000). In addition, there is disagreement on the economic effects of monetary integration with respect to income correlation among member countries and intra-area trade flows. The “specialization hypothesis” (Krugman, 1993; Krugman and Venables, 1996) postulates that as countries become more and more integrated, their industrial structure will develop according to their comparative advantages (Eichengreen and Bayoumi, 1996). In this perspective, the economic systems of each member country of an OCA would become more sectorally concentrated and vulnerable to supply shocks. Opposite implications arise from the “endogeneity hypothesis” (Frankel and Rose, 1997). This paradigm postulates that a positive link between income correlation and trade integration exists, suggesting that countries joining a currency union may satisfy the properties of an OCA *ex-post* even if they do not *ex-ante*.

Even though a large body of research has been done on these issues, the question whether the characteristics of the EMU match those of an OCA has not received a definitive answer yet. The main objective of this paper is to contribute to the ongoing debate, extending the work of Sarno (1997) and

Mouratidis (2001). More specifically, cointegration techniques are employed to compare price dynamics within the EMU taking Germany as a benchmark. The latter is assumed to be the base country in the EMU because of her dominant role during the years of the European Monetary System (EMS) (Giavazzi and Pagano, 1988; Mélitz, 1988). The econometric analysis aims at testing three related hypotheses. First, the purchasing power parity (PPP) condition is tested for each member country of the EMU with respect to the base economy. Second, the generalized PPP (GPPP) hypothesis (Enders and Hurn, 1994) is tested for those economies for which the former arbitrage condition does not hold. According to the GPPP theory, bilateral real exchange rates individually non-stationary may be cointegrated if their long-run macroeconomic determinants (forcing variables) are highly correlated. Thus, the existence of an equilibrium path for a linear combination of real exchange rates allows to rule out the presence of real asymmetries (Bayoumi and Taylor, 1995) and to interpret the empirical validity of the GPPP hypothesis in terms of long-run sustainability of a monetary area in the spirit of Mundell (1961). Third, a measure of the speed of convergence within EMU countries is provided.

It should be noted that the euro area has faced significant structural changes over the last two decades, such as the crisis of the EMS in 1992 and the introduction of the single currency. From the methodological point of view it is advisable to properly take into account at least these major events, which may affect the statistical properties of the variables under consideration. For this purpose, vector error correction (VEC) models with breaks at known times (Johansen et al., 2000) are employed, placing two breaks in 1992M3 and 1998M6, i.e. just before the occurrence of the two above-mentioned episodes.

The paper is structured in five Sections. In the following some stylized facts on the variables involved in the empirical investigation are illustrated. Section 3 presents a formalization of the economic hypotheses to be tested and their econometric specification. Estimation results are shown in Section 4. Concluding remarks follow.

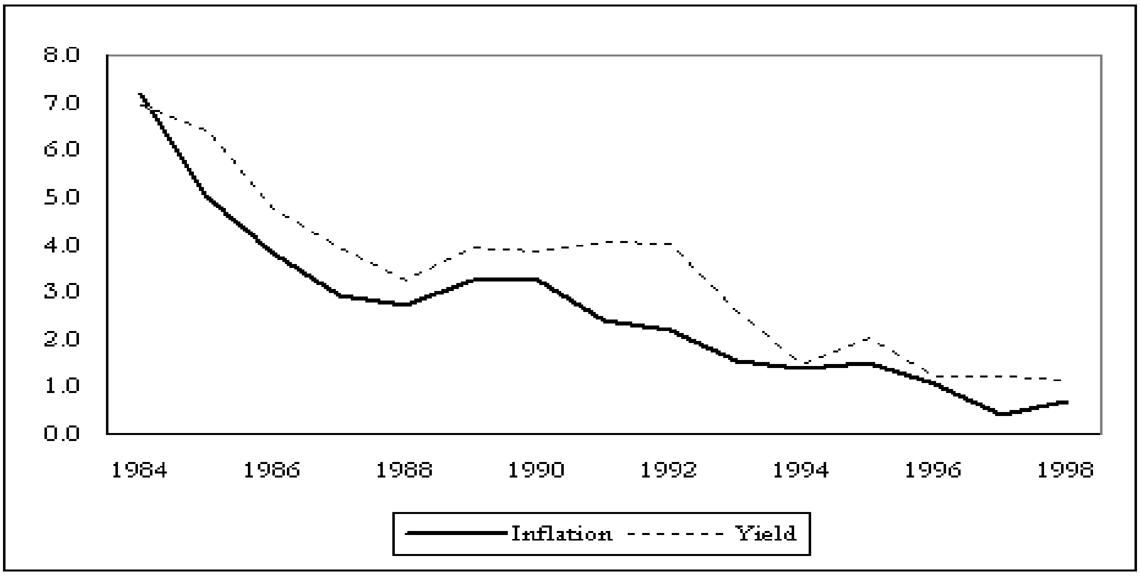
2 SOME STYLIZED FACTS: BOND YIELDS, INFLATION DYNAMICS AND REAL EXCHANGE RATES

During the second stage of the EMU there has been a progressive homogenization of national economic policies and structural features in several European countries, even if the Maastricht criteria have only been partially met. Despite very dissimilar conditions in terms of deficit/GDP and debt/GDP ratios across member countries at their entrance in the third stage of the EMU, an almost complete convergence in terms of interest and inflation rates has occurred.

Figure 1 (dashed line) shows the reduction of the standard deviation of yield differentials between ten-year government bonds issued in the EMU member countries (with the exception of Greece) and the corresponding ten-year government bond issued in Germany, assumed as a risk-free asset (Favero et al., 1997), over the years of the EMS.[1]

Fig. 1 Standard deviation of inflation and long-term bond yield differentials in the EMU countries with respect to Germany: 1984-1998

(percentage values)

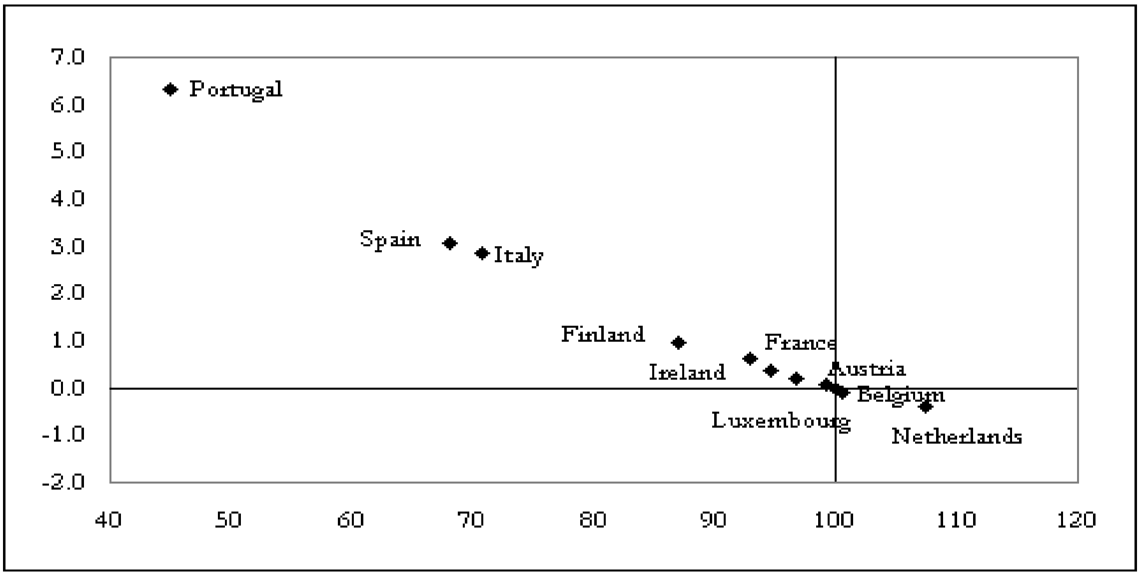


[1] During the Eighties and the Nineties, Germany has been a meta-economic reference point for the other European countries. Even though her central role may be less evident during the most recent years, Germany still weights for roughly one third of the euro area GDP. Moreover, German monetary and fiscal policy strategies have inspired, to some extent, the institutional architecture at the basis of the EMU.

Analogously, the dispersion of inflation differentials with respect to Germany has decreased over time, mainly starting from the first years of the Nineties (Fig. 1, continuous line), suggesting that to some extent a convergence of price levels has occurred in Europe.

This is also shown in Figure 2, illustrating the dispersion of the relative price level at the beginning of the period of analysis (horizontal axis) and of the average inflation differential over the period 1984-1998 (vertical axis) for the EMU member countries with respect to Germany. It should be noted that countries exhibiting higher (lower) price levels as compared to the base country are characterized by lower (higher) inflation rates.

Fig. 2 Relative prices (horizontal axis) in 1984 and 1984-1998 average inflation rate differentials in the EMU countries with respect to Germany
(vertical axis, percentage values)



Nevertheless, in presence of fixed-but-adjustable exchange rates, as in the case of the EMS, the role played by the nominal exchange rate should be taken into account. Such factor turns out to be particularly relevant for those countries which have used this policy tool as a “safety valve” to restore competitiveness of their economic systems during the Eighties and the Nineties. Therefore, for a meaningful comparison the price level of each country should be expressed in terms of a common currency

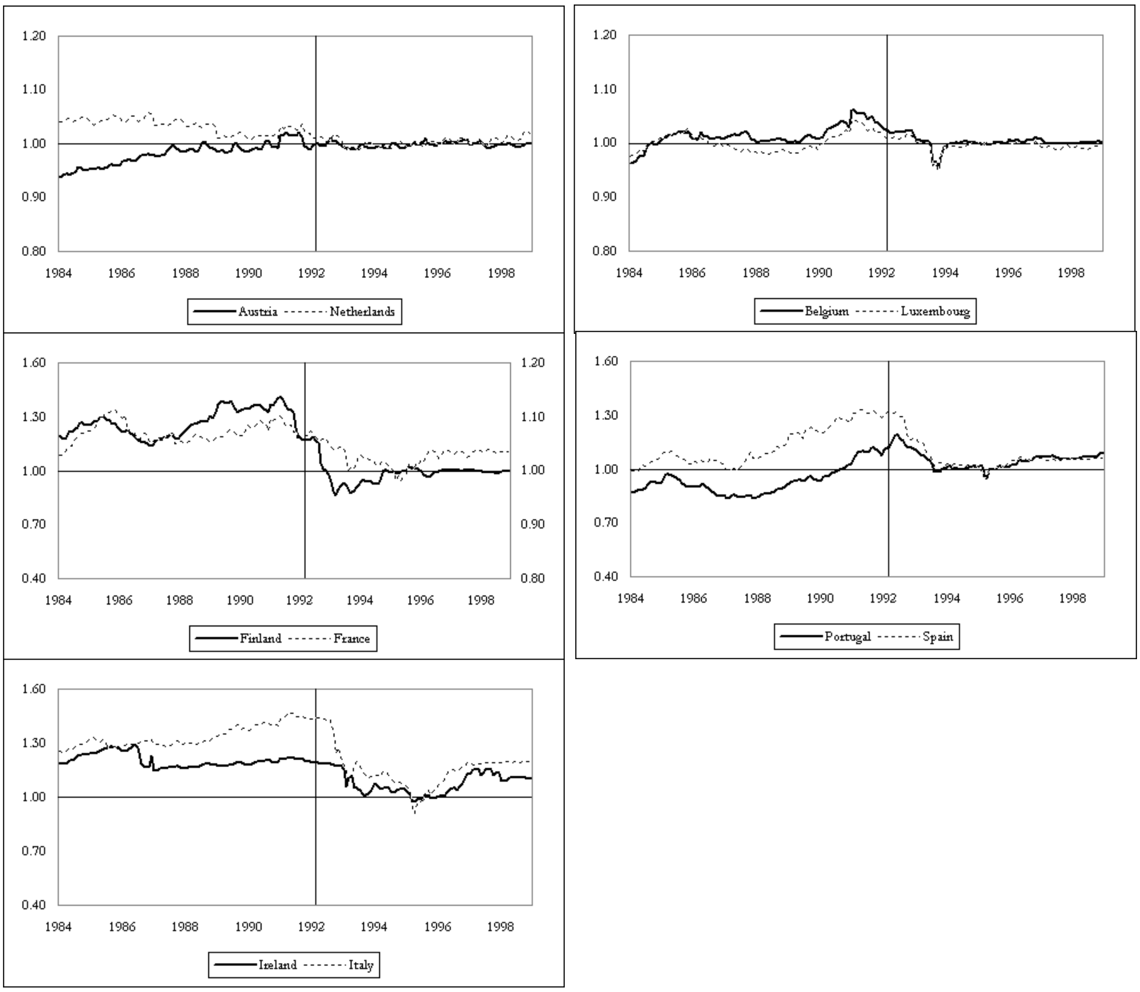
$$RER_{i0,t} \equiv P_{it} / \left(P_{0t} \cdot E_{i0,t} \right) \quad (1)$$

where P_{it} (P_{0t}) indicates the price level of the i-th country (Germany) and $E_{i0,t}$ is the bilateral nominal exchange rate between the i-th country and Germany, obtained by the triangularization $E_{i0,t} = E_{i\$,t} / E_{0\$,t}$, where the numerator (denominator) represents the bilateral nominal exchange rate between the currency of the i-th country (Germany) and the U.S. dollar. Accordingly, expression (1) can be reformulated as

$$RER_{i0,t} \equiv Q_{it} / Q_{0t} \tag{2}$$

where $Q_{it} = P_{it} / E_{i\$,t}$ and $Q_{0,t} = P_{0t} / E_{0\$,t}$. Bilateral real exchange rates (2) between Germany and the other EMU countries are reported in Figure 3.

Fig. 3 **Bilateral rate exchange rate in the EMU countries with respect to Germany: 1984-1998**



The graphs indicate that at the beginning of the third stage of the EMU price levels in Core Europe countries (Austria, Belgium, Luxemburg and the Netherlands) have been aligned with those of Germany. On the other hand, price levels in France and in peripheral countries (Cohesion Funds countries and Italy) have been lower with respect to the base country[2]. In the first group of countries, the volatility of bilateral exchange rates appears rather limited. Such dynamics is even more evident after the EMS crisis in 1992, when the realignment of price levels in Austria and in the Netherlands with respect to Germany is fully accomplished, even though it follows opposite trajectories. In fact, while at the beginning of the sample Austrian (Dutch) price level is slightly lower (higher) with respect to Germany, inflation rate in Austria (Netherlands) during the pre-EMU period results lower (higher), as shown in Figure 2. For France and Finland real exchange rates evolve similarly over time, but in the case of France the oscillations are more limited and comparable with those of the Core Europe countries. Finally, peripheral countries show a more irregular real exchange rate dynamics, with permanent deviations from a constant pattern.

3 METHODOLOGICAL OUTLINE

The empirical analysis derives its main tools from the VEC methodology (Johansen, 1988). This modeling approach takes into account both long-run relationships and short-run dynamic interdependencies among a small set of variables, allowing to associate the economic concept of long-run with the statistical concept of stationarity. The VEC model is particularly suitable to analyze singularly non-stationary time series as it is based on the preliminary identification of stationary linear combinations of such series, known as cointegrating vectors. Ideally, these vectors describe the steady-state configuration which the data tend to revert to in the long-run.

More specifically, since the validity of the PPP implies some sort of mean reversion, the question “do we care if the real exchange rate has a unit root?” (Sarno and Taylor, 2002) deserves a careful answer because the presence of

[2] The dynamics of bilateral exchange rate vis-à-vis Germany closely mimics the picture emerging from the empirical works focused on the relationship between national economic structures and the monetary transmission mechanism (Ehrmann, 2000; Mojon and Peersman, 2003).

non-stationary exchange rates has relevant implication not only from the economic point of view, i.e. the validity of international macroeconomic theories assuming that the PPP holds may be questionable, but also from the statistical point of view, i.e. the econometric tools not expressly designed to deal with units roots may provide spurious estimation results[3].

3.1 PPP condition and beyond

Expressing (2) in log-linear terms

$$rer_{i0,t} \equiv q_{it} - q_{0t} \tag{3}$$

where lowercase variables represent logarithms, the formulation of an empirically testable long-run PPP equilibrium condition involves the specification of a stationary stochastic residual, $\varepsilon_{ppp,t}$, in order to describe the deviations from the steady-state

$$q_{it} - q_{0t} = \varepsilon_{ppp,t} \tag{4}$$

Ahn et al. (2002) stress that even though a PPP-based approach is relatively useful to analyze international competitiveness issues, such point of view is partial and fails to capture the major changes in economic policies and the significant restructuring processes in Europe during the pre-EMU convergence in the Eighties and the Nineties. Conversely, the GPPP theory suggests that the (possible) non-stationarity of real exchange rates may be related to the non-stationarity of their long-run macroeconomic determinants. In practice, the GPPP hypothesis holds if it is possible to identify (at least) one linear combination of bilateral real exchange rates

$$rer_{i0,t} = \sum_{j=1,\, j\neq i}^{k} \beta_j \cdot rer_{j0,t} + \varepsilon_{gppp,t} \tag{5}$$

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[3] In the empirical literature, which uses cointegration techniques to analyze the long-run relationship between exchange rates and relative prices, there is no clear evidence with respect to the *I*(1)-ness or the stationarity of the real exchange rates (Taylor, 1988; Froot and Rogoff, 1995). Discordant results may arise because of several factors, such as time horizon, econometric approach and the choice to use bivariate systems (in which relative prices are treated as a single variable) rather than trivariate systems (in which domestic and foreign prices enter separately). In the present work, the specification of the variables allows to extend the sample span also over the years of the EMU, avoiding the introduction of the nominal exchange rate in the endogenous set.

where $\varepsilon_{gppp,t}$ is stationary. Parameters in (5), β_j's, synthesize the economic interdependencies within the EMU in terms of commercial and financial transactions, technology transfers and migration flows. Their values depend on the functional form of the national aggregate demand functions. Specifically, the more similar the aggregate demand functions the smaller the β_j's (Enders and Hurn, 1994). Expressing the GPPP condition in terms of price levels in the same currency, q_{it}'s, equation (5) becomes

$$q_{it} = \sum_{j=1, j \neq i}^{k} \beta_j \cdot q_{jt} - \beta_0 \cdot q_{0t} + \varepsilon_{gppp,t} \tag{6}$$

where $\beta_0 = \sum_{j=1, j \neq i}^{k} \beta_j + 1$.[4]

3.2 Econometric specification

Provided that all the involved variables have to be at most *I*(1), the k-dimensional VEC model may be written as

$$\begin{bmatrix} \cdots \\ \Delta q_{it} \\ \cdots \\ \Delta q_{0t} \end{bmatrix} = \mathbf{\Pi} \cdot \begin{bmatrix} \cdots \\ q_{it} \\ \cdots \\ q_{0t} \\ t \end{bmatrix} + \sum_{i=1}^{p-1} \mathbf{\Pi}_i \cdot \begin{bmatrix} \cdots \\ \Delta q_{i,t-1} \\ \cdots \\ \Delta q_{0,t-1} \end{bmatrix} + \begin{bmatrix} \cdots \\ u_{q_{it}} \\ \cdots \\ u_{q_{0t}} \end{bmatrix} \tag{7}$$

The identification of the long-term component of the model implies the choice of number and structure of the equilibrium relationships (the cointegrating vectors, in statistical terms). The number of such relationships is equal to the (reduced) rank of the long-run matrix, $\mathbf{\Pi}$, which can be partitioned as the product of two matrices, $\boldsymbol{\alpha} \cdot \boldsymbol{\beta}'$. The adjustment matrix, $\boldsymbol{\alpha}$, contains the feedback coefficients (loadings). The cointegration matrix, $\boldsymbol{\beta}$, contains the $r < k$ theoretical long-run relationships which the data converge to once the effects of transitory shocks have been absorbed (Johansen, 1995).

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[4] Obviously, (6) is equivalent to (4) if the condition $\sum_{j=1, j \neq i}^{n} \beta_j = 0$ holds.

As recently stressed in the econometric literature (Clements and Hendry, 2001), structural changes in the deterministic component may affect the integration and cointegration properties of the variables. Moreover, from an operational point of view dummy variables are often requested to properly take into account specific events which may affect the structure of the economic system. Thus, the cointegration analysis is based on the VEC methodology with broken linear trend at known times. Moreover, the class of models used in the analysis, employing segmented (instead of unbroken) deterministic components, allows to reduce the probability to erroneously reject the hypothesis of cointegration, providing more robust results.

Following Johansen et al. (2000) the sample span $(1,...,n)$ is divided into s subsamples $(1,...,n_1)$, $(n_1+1,...,n_2)$, $\ldots$, $(n_{s-1}+1,...,n)$, with $n_0=0$ and $n_s=n$. In each subsample the parameters of the statistical model are assumed to be the same, while the intercept and the linear trend may differ. Under the hypothesis of cointegration in each subsample, expression (7) becomes

$$\Delta\mathbf{y}_t=\boldsymbol{\alpha}\cdot\left[\boldsymbol{\beta}' \,\vdots\, \gamma_1' \,\vdots\, \ldots \,\vdots\, \gamma_s'\right]\cdot\begin{bmatrix}\mathbf{y}_{t-1}\\ t\end{bmatrix}+\boldsymbol{\mu}_j+\sum_{i=1}^{p-1}\boldsymbol{\Pi}_i\cdot\Delta\mathbf{y}_{t-1}+\mathbf{u}_t \qquad (8)$$

where $\mathbf{y}_t=[...\ q_{it}\ ...\ q_{0t}]'$, $\mathbf{u}_t=[...\ u_{q_{it}}\ ...\ u_{q_{0t}}]'$, $j=1,\ldots s$, $n_{j-1}+p<t\le n_j$ and $\boldsymbol{\mu}_j$ is a $(k\times 1)$ vector. The j models (8) can be expressed compactly as

$$\Delta\mathbf{y}_t=\mathbf{A}\cdot\left[\mathbf{B}' \,\vdots\, \gamma'\right]\cdot\begin{bmatrix}\mathbf{y}_{t-1}\\ t\cdot\mathbf{E}_t\end{bmatrix}+\boldsymbol{\mu}\cdot\mathbf{E}_t+\sum_{i=1}^{p-1}\boldsymbol{\Pi}_i\cdot\Delta\mathbf{y}_{t-1}+\sum_{i=1}^{p}\sum_{j=2}^{s}\eta_{ji}\cdot D_{j,t-i}+\mathbf{u}_t \qquad (9)$$

where $\boldsymbol{\alpha}\cdot\left[\boldsymbol{\beta}' \,\vdots\, \gamma_1' \,\vdots\, \ldots \,\vdots\, \gamma_s'\right]=\mathbf{A}\cdot\left[\mathbf{B}' \,\vdots\, \gamma'\right]$,

$$D_{jt}=\begin{cases}1 & \text{if}\quad t=n_{j-1}\\ 0 & \text{otherwise}\end{cases} \qquad\qquad j=1,\ldots,s$$

and $\mathbf{E}_t$ is a $(s\times 1)$ vector in which the j-th element is given by[5]

[5] $D_{j,t-1}$ is a dummy variable referring to the i-th observation in the j-th subsample aiming at excluding the residuals associated to the first p observations in each period. E_{jt} is a dummy variable indicating the observations in the j-th subsample, with the exception of the first p observations.

$$E_{jt} = \sum_{i=p+1}^{n_j - n_{j-1}} D_{j,t-i} = \begin{cases} 1 & \text{if} \quad n_{j-1} + p + 1 \le t \le n_j \\ 0 & \text{otherwise} \end{cases}$$

In the particular case where the intercepts in the cointegrating relationships varying in each subsample are the only deterministic components in the model ($\boldsymbol{\pi}_j = \mathbf{0}$, $1 < j < s$, and vectors $\boldsymbol{\mu}_j$ constrained to belong to the cointegration space), expression (9) can be simplified as

$$\Delta \mathbf{y}_t = \mathbf{A} \cdot \left[\mathbf{B}' \middle| \boldsymbol{\delta}' \right] \cdot \begin{bmatrix} \mathbf{y}_{t-1} \\ \mathbf{E}_t \end{bmatrix} + \sum_{i=1}^{p-1} \boldsymbol{\Pi}_i \cdot \Delta \mathbf{y}_{t-i} + \sum_{i=1}^{p} \sum_{j=2}^{s} \eta_{ji} \cdot D_{j,t-i} + \mathbf{u}_t \qquad (10)$$

where $\boldsymbol{\mu} = \mathbf{A} \cdot \boldsymbol{\delta}'$.

The trace test can be used to verify the existence of r cointegrating relationships (Johansen, 1995). However, in presence of structural breaks the asymptotic distribution of this test is different from the usual one, although it still belongs to the class of multivariate Dickey-Fuller-type tests. Critical values of the trace test for models (9) and (10) can be approximated through a Γ distribution, as shown in Johansen et al. (2000). Once the dimension of the cointegration space is determined, the long-run structure can be identified through the imposition of restrictions on the cointegration matrix. Generally, the statistical validity of such restrictions is tested through a LR test, which is asymptotically distributed as a χ^2_{ω}, where ω is the number of the imposed over-identifying restrictions.

4 ESTIMATION RESULTS

The empirical investigation involves three steps and refers to the eleven countries that joined the EMU in 1999. Monthly seasonally-adjusted U.S. dollar per national currency nominal exchange rates and consumer price indexes (CPI) are taken from the IMF International Financial Statistics database and cover the period 1984M1-2002M12.

a. For the i-th EMU country a bivariate system involving q_{it} and q_{0t} is estimated. If a cointegrating vector with coefficients 1 and -1 turns out to exist ("proportionality and symmetry hypothesis"), then it provides empirical support to the GPPP hypothesis in system (7) for the special case where $k = 2$.

b. On the other hand, the non-stationarity of a bilateral real exchange rate is not taken as an indication of the inadequacy of the i-th economy to form a common monetary area with the base country. Even if a unit root is present in a bilateral real exchange rate, there might exist some real fundamentals (such as terms of trade, tax systems and productivity, among others) that determine the permanent deviation of the real exchange rate from the PPP condition (Fisher and Park, 1991; Kim and Enders, 1991). For these countries, their price levels expressed in the same currency, q_{it}'s, are collected in vector $\mathbf{y}_t$ together with the price level of the base country, q_{0t}, in order to test the GPPP hypothesis.

c. Converting prices in the same currency allows to control for the impact of the nominal exchange rate in the adjustment mechanism through changes in inflation differentials in order to reach the new equilibrium real exchange rate. Therefore, statistically significant feedback coefficients (α_i's) suggest that an EMU country's inflation rate dynamics allows the realignment of relative prices with respect to the base country,[6] assuming that German inflation represents the benchmark ($\alpha_0 = 0$) for the other European countries.

[6] The greater the (absolute) values of adjustment coefficients the quicker the absorption of deviations from equilibrium.

4.1 The convergence process in the pre-EMU years: 1984-1998

The first round of estimates refers to the period 1984M1-1998M6, imposing a structural break in 1992M3. The choice of the first observation allows to leave the early years of the EMS out, when considerable adjustments to the new monetary system have taken place. The last months of 1998 are not included because of the transition towards the third stage of the EMU. The introduced structural break allows to exclude the period immediately preceding the EMS crisis. Figure 3 shows the relevant discrepancy between the trajectories of bilateral real exchange rates with respect to Germany for several European countries (namely Finland, France, Ireland, Italy, Portugal and Spain). In these cases, there is no evidence of a mean reversion to the pre-EMS crisis levels.

- The preliminary analysis encompasses unit root tests of bilateral real exchange rates for the EMU countries with respect to Germany. Table 1 reports ADF and PP test results over the period 1984M1-1998M6, where optimal lags are selected using the AIC criterion, setting the maximum lag equal to six.

  In general, most real exchange rates turn out $I(1)$ variables, although those of the economically and geographically closest partners of Germany (Belgium, Luxembourg and the Netherlands) appear stationary. This is consistent with the GPPP theory, according to which cross-country similarities in economic systems are reflected on real exchange rate dynamics. For variables in levels, the ADF test suggests that the null hypothesis of unit root is rejected only for Austria, Belgium (at the 5% significance level) and Luxemburg (at the 1% significance level). The PP test provides similar results, with the exception of the real exchange rate of the Netherlands, which is stationary at the 10% significance level. For the variables in first differences, both tests strongly reject the null hypothesis.

Tab. 1 **Unit root tests**

	$rer_{i,t}$								$\Delta rer_{i,t}$							
					ADF		PP						ADF		PP	
	d.c.	c.v. 10%	5%	1%	stat.	(lag)	stat.	(lag)	d.c.	c.v. 10%	5%	1%	stat.	(lag)	stat.	(lag)
AUT	*c,t*	-3.14	-3.44	-4.01	**-3.70**	(4)	**-3.87**	(5)	*c*	-2.58	-2.88	-3.47	***-9.36***	(3)	***-15.6***	(4)
BEL	*c*	-2.58	-2.88	-3.47	**-3.09**	(0)	**-3.09**	(0)	-	-1.62	-1.94	-2.58	***-7.78***	(2)	***-12.8***	(2)
FIN	*c*	-2.58	-2.88	-3.47	-1.03	(1)	-1.05	(1)	-	-1.62	-1.94	-2.58	***-8.72***	(0)	***-8.72***	(0)
FRA	*c*	-2.58	-2.88	-3.47	-1.56	(0)	-1.56	(0)	-	-1.62	-1.94	-2.58	***-11.6***	(0)	***-11.6***	(0)
IRE	*c,t*	-3.14	-3.44	-4.01	-2.05	(0)	-2.05	(0)	*c*	-2.58	-2.88	-3.47	***-7.25***	(2)	***-13.5***	(2)
ITA	*c,t*	-3.14	-3.44	-4.01	-1.81	(4)	-1.87	(4)	*c*	-2.58	-2.88	-3.47	***-6.15***	(3)	***-10.0***	(3)
LUX	*c*	-2.58	-2.88	-3.47	-2.55	(0)	-2.53	(0)	-	-1.62	-1.94	-2.58	***-7.91***	(3)	***-17.4***	(3)
NET	*c,t*	-3.14	-3.44	-4.01	-2.41	(6)	*-3.19*	(6)	*c*	-2.58	-2.88	-3.47	***-5.38***	(5)	***-11.1***	(5)
POR	*c,t*	-3.14	-3.44	-4.01	-1.84	(3)	-1.96	(3)	*c*	-2.58	-2.88	-3.47	***-6.35***	(2)	***-9.97***	(2)
SPA	*c,t*	-3.14	-3.44	-4.01	-1.59	(1)	-1.62	(1)	*c*	-2.58	-2.88	-3.47	***-10.2***	(0)	***-10.2***	(0)

Note. For each regression deterministic component (d.c.), critical values (c.v.), test statistics (stat.) and number of lags (lag) are reported. Unit root tests are performed on variables in levels (left part) and in first differences (right part). Statistics in italics, bold, and italics and bold indicate the rejection of the null hypothesis of unit root at the 10%, 5% and 1% significance level, respectively.

- VEC models are specified according to (9) or (10) on the basis of the deterministic component chosen in the ADF and PP test regressions. Table 2 shows the optimal lag, p, for each model using AIC and HQ criteria. As suggested by the econometric literature on cointegrating systems, the second criterion is preferred when discordant results occur. Vector autoregressions include a broken deterministic component in 1992M3.

  Diagnostic tests (not reported) indicate that all models are well specified. Residuals are tested against serial correlation, heteroschedasticity, both for individual equations and for the system as a whole. The distribution of residuals suggests departure from normality mainly due to skewness rather than to kurtosis. However, since cointegration analysis is more sensitive to the latter, such non-normality is not considered as a source of misspecification (Gonzalo, 1994). One-step and break-point Chow tests at the system level are used to check the stability of the models.[7]

[7] Elaborations have been performed using E-views 4.1 (data construction and unit root tests), Malcolm 2.90 (estimation) and Pc-Fiml 10.3 (preliminary analyses and estimation) econometric packages. All test statistics are available upon request.

Tab.2 **Model specification and lag determination**

AUT	BEL	FIN	FRA	IRE	ITA	LUX	NET	POR	SPA
				Model specification					
(9)	(10)	(10)	(10)	(9)	(9)	(10)	(9)	(9)	(9)
				Number of lags					
2	6	5	2	4	5	4	2	2	2

Table 3 reports trace test results which suggest the existence of a cointegrating relationship for six countries: Austria, Belgium, Finland, France, Italy and the Netherlands.

Tab. 3 **Trace test for the PPP hypothesis: 1984-1998**

H_0	H_1	90% c.v.	95% c.v.	AUT	BEL	FIN	FRA	IRE	ITA	LUX	NET	POR	SPA
				Trace test (single structural break)									
$r=0$	$r\geq1$	23.25	25.48		**26.54**	**27.54**	*24.91*			14.24			
$r\leq1$	$r=2$	10.20	11.56		7.33	6.01	7.28			5.30			
$r=0$	$r\geq1$	34.46	37.42	**46.85**				29.07	**56.81**		**40.78**	24.43	33.08
$r\leq1$	$r=2$	16.81	18.95	5.86				6.43	8.74		7.36	12.12	12.07

Note. Upper part: trace test results for models without linear trend in the cointegration space. Lower part: trace test results for models with linear trend in the cointegration space. Statistics in italics (bold) indicate the rejection of the null hypothesis at the 10% (5%) significance level.

The long-run structure of these models is analyzed trying to provide an economically meaningful interpretation. Table 4 reports exclusion tests for each element of $\mathbf{y}_t$ (upper part) and the final specification of the cointegrating relationships (lower part).

In all models, both the elements in $\mathbf{y}_t$ as well as the deterministic component are statistically significant. Nevertheless, the "proportionality and symmetry hypothesis" between domestic and German prices is not rejected only for Belgium, Finland, France and the Netherlands.[8]

[8] The test statistic for Austria (Italy) is equal to 16.09 (6.29), greater than the critical value at the 5% significance level for a $\chi^2(1)$ distribution, equal to 3.84.

Tab. 4 Exclusion test for the PPP hypothesis and long-run structure: 1984-1998

Exclusion test for each element in $\mathbf{y}_t$						
$\chi^2(1)$ c.v.	5% = 3.84, 1% = 6.64					
	AUT	BEL	FIN	FRA	ITA	NET
$q_{it} = 0$	33.30	11.04	15.20	9.62	34.16	25.98
$q_{0t} = 0$	34.18	10.77	14.10	9.26	27.08	26.05

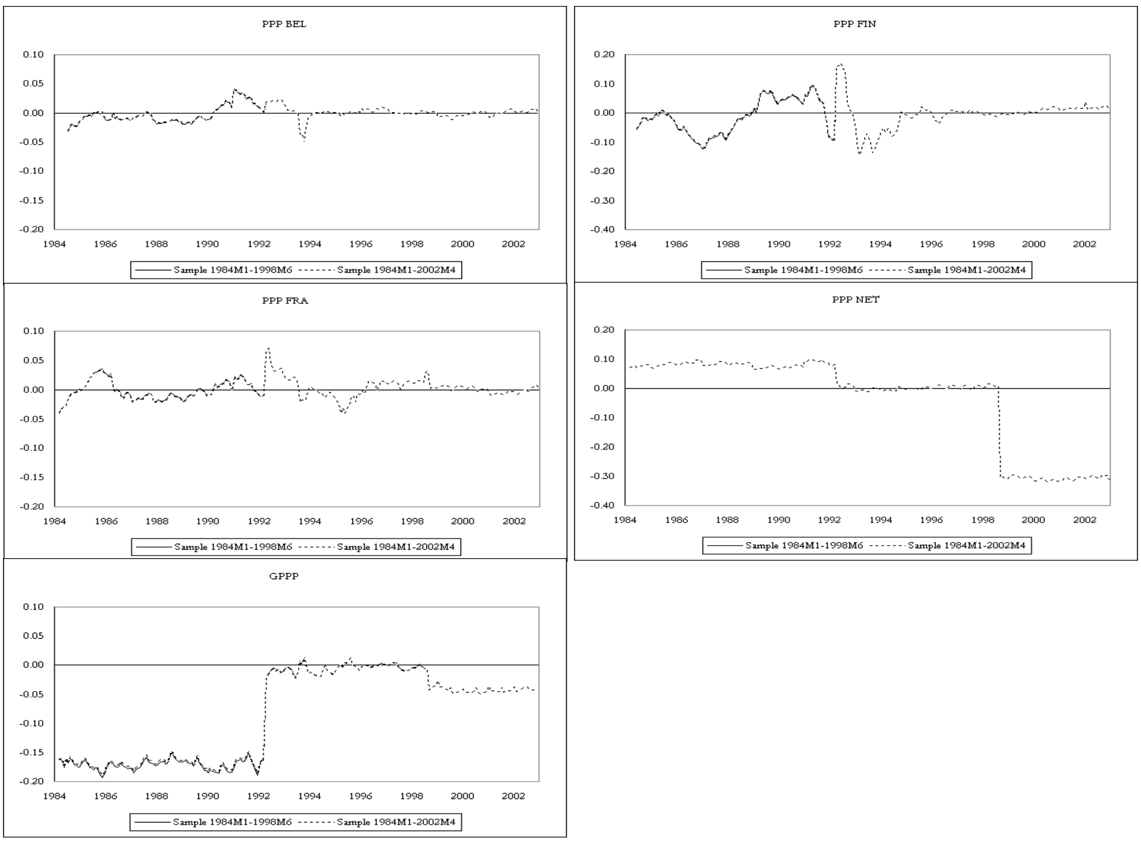
Long-run structure

	q_{it}	q_{0t}	γ_1	γ_2	δ_1	δ_2	(d.f.)	stat.	[p-value]
AUT	1	-1.0512 (0.0003)	-0.0002 (2.9 e-05)	0			$\chi^2(1)$	*0.19*	[0.66]
BEL	1	-1			-0.0202 (0.0043)	0	$\chi^2(2)$	**1.60**	[0.45]
FIN	1	-1			-0.2520 (0.0173)	0	$\chi^2(2)$	**5.46**	[0.07]
FRA	1	-1			-0.0710 (0.0057)	-0.0225 (0.0065)	$\chi^2(1)$	**1.41**	[0.23]
ITA	1	-0.7511 (0.0045)	-0.0034 (0.0003)	-0.0021 (0.0002)				-	
NET	1	-1	0.0004 (2.9 e-05)	0			$\chi^2(2)$	**3.00**	[0.22]

Note. Standard errors in parentheses. Statistics in bold (italics) are referred to hypotheses statistically significant and coherent (incoherent) with economic theory.

For these countries, bilateral real exchange rates result substantially constant over time, as shown in Figure 4 (continuous line). In the Belgian

Fig. 4 Cointegration residuals



and Finnish cases there is a complete price realignment with respect to Germany. In the Dutch case, the linear trend has a negative slope in the first subsample, mainly due to the combination of lower inflation with respect to Germany and substantially fixed bilateral nominal exchange rates. After the realignment, both countries show similar prices dynamics oscillating around a constant value.

- Price levels of countries in which the PPP does not hold are grouped together with German prices in order to test the GPPP hypothesis. System (9) is estimated with two lags, according to the AIC and HQ criteria. The structural break in the linear trend is set at 1992M3. Trace test indicates the existence of a cointegrating relationship, as shown in table 5.

Tab. 5 **Trace test for the GPPP hypothesis: 1984-1998**

	Trace test (single structural break)						
H_0	r=0	r≤1	r≤2	r≤3	r≤4	r≤5	r≤6
H_1	r≥1	r≥2	r≥3	r≥4	r≥5	r≥6	r=7
90% c.v.	177.62	141.33	108.85	80.23	55.5	34.46	16.81
95% c.v.	183.81	146.88	113.75	84.49	59.12	37.42	18.95
stat.	**195.53**	129.17	87.81	57.31	32.88	15.14	5.2

Note. Statistics in bold indicate the rejection of the null hypothesis at the 5% significance level.

The specification of the long-run structure of the model is based on the statistical significance of each element in $\mathbf{y}_t$ (Tab. 6).

Tab. 6 **Exclusion test for the GPPP hypothesis and long-run structure: 1984-1998**

	Exclusion test for each element in $\mathbf{y}_t$						$\chi^2(1)$	
	AUT	IRE	ITA	LUX	POR	SPA	GER	5% c.v.	1% c.v.
q_{it} = 0	18.41	**1.12**	11.03	18.24	8.74	**0.79**	*5.62*	3.84	6.64

Long-run structure											
			q_{it}								
AUT	IRE	ITA	LUX	POR	SPA	GER	γ_1	γ_2	(d.f.)	stat.	[p-value]
1	0	-0.1222	-0.6866	0.2352	0	-0.4264	-0.00082	0	$\chi^2(4)$	**3.99**	[0.41]
		(0.0062)	(0.0669)	(0.0110)		(0.0631)	(3.7 e-05)				

Note. Upper part: statistics in italics (bold) indicate the failure of rejection of the null hypothesis. Lower part: standard errors in parentheses. Statistics in bold are referred to hypotheses statistically significant and coherent with economic theory.

Exclusion tests suggest that price levels of Ireland and Spain can be eliminated from the cointegration space. This finding points out that in the pre-EMU years, Ireland and Spain have been completely disjoined from the other European countries.[9]

The cointegration space is normalized on the prices level of Austria.[10] All variables (including German prices) are statistically significant as well as the deterministic component in the first subsample. As clearly illustrated in Figure 4 (continuous line), prices levels are aligned to those of Germany in the Nineties, consistently with the results of the previous Subsection.

These findings are coherent with those of Mouratidis (2001), who supports the hypothesis that European countries have started to form an OCA during the Nineties, even though this conclusion is rather influenced by the time horizon of the analysis. Conversely, in the present work the employed methodology allows to find a stable cointegrating relationship over the period 1984-1998.

More conventionally, the final specification of the cointegrating vector can be rewritten in terms of real exchange rates

$$rer_{aut0,t} = 0.12 \cdot rer_{ita0,t} + 0.69 \cdot rer_{lux0,t} - 0.24 \cdot rer_{por0,t} + 0.0008 \cdot E_{1t} \cdot t + \varepsilon_{gppp,t}$$

Non-concordant signs and differences in the (absolute) values of long-run coefficients signal a certain degree of heterogeneity of the aggregate demand functions of these economies, highlighting a potential weakness of EMU member countries to real shocks. In any case, according to the "endogeneity hypothesis" (Frankel and Rose, 1997) such vulnerability should progressively disappear as long as the convergence process keeps proceeding.

- Table 7 reports feedback coefficients for the models employed to test the PPP hypothesis for Belgium, Finland, France and the Netherlands (upper part) as well as for the model employed to test the GPPP hypothesis (lower part), once the long-run restrictions in Tables 4 and 6, respectively, have been imposed.

[9] The result for Ireland can be explained taking into account her stronger economic linkages with the U.S. and the U.K. with respect to continental European countries. The result for Spain, as suggested by Carlucci and Girardi (2004), may be due to the rigidities in her economic structure.

[10] The specification of the equilibrium condition is not influenced by the chosen country for the normalization. The unrestricted long-run relationship is very similar to the one reported in the text.

Tab. 7 **Feedback coefficients: 1984-1998**

	BEL		FIN		FRA		NET	
α_i	-0.1237		-0.0805		-0.0960		-0.2634	
	(0.0302)		(0.0207)		(0.0252)		(0.0466)	
α_0	0		0		0		0	
$\chi^2(4)$	1.61	[0.81]	5.88	[0.21]	3.44	[0.49]	3.11	[0.54]

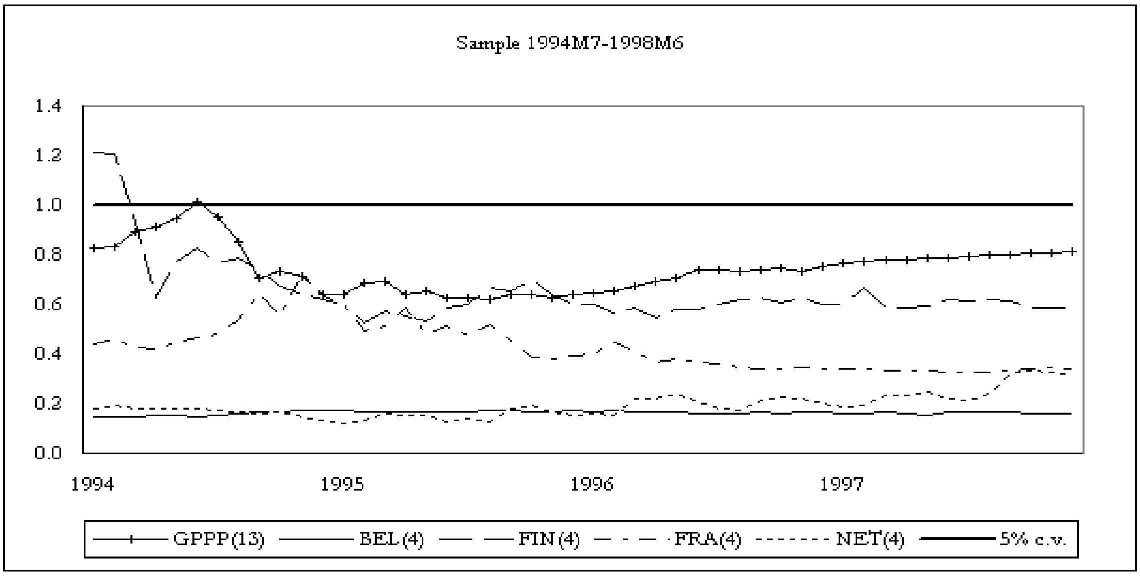
	AUT	IRE	ITA	LUX	POR	SPA	GER
α_i	-0.2657	0	0	0.1411	0	0	0
	(0.0424)			(0.0536)			
$\chi^2(13)$	18.23 [0.15]						

Note. Upper part: results for PPP-based models. Lower part: results for GPPP-based model. Standard errors in parentheses.

In countries with rather constant bilateral real exchange rates (Figure 3), the speed of adjustment towards equilibrium is very high. For Austria and the Netherlands the adjustment coefficient is substantially similar (around twice the coefficient of Belgium and Luxemburg). In Finland and France the adjustment mechanism is slower, reflecting their later entrance in the Core Europe. On the other hand, for peripheral countries there is evidence of a step-by-step alignment to German prices (through nominal exchange rate devaluations) rather than in the "continuous time".

Finally, the over-identified structure of the long-run matrix, $\mathbf{\Pi}$, appears markedly stable in each model (Fig. 5). The only exceptions concern the

Fig. 5 **Long-run matrix stability test: 1994M7-1998M6**



Note. Under the null hypothesis the parameters of the model are stable. The horizontal line, normalized to unity, indicates the 5% significance level. Number of over-identifying restrictions in parentheses.

months 1994M7 and 1994M8 for Finland, in which the null hypothesis is not rejected at the 3% significance level.

4.2 The convergence process during the EMU years: 1984-2002

This Subsection replicates the previous procedures extending the sample over the period 1984M1-2002M12, i.e. including the introduction of the single currency, in order to test possible endogenous effects within the new European economic and institutional architecture (Frankel and Rose, 1997). To this aim, a second structural break, excluding from the estimates the transition months before the beginning of EMU is introduced in 1998M6.

- Trace test results for the same models illustrated in Table 2 are shown in Table 8.

Tab. 8 **Trace test for the PPP hypothesis: 1984-2002**

				AUT	BEL	FIN	FRA	IRE	ITA	LUX	NET	POR	SPA
H_0	H_1	90% c.v.	95% c.v.	Trace test (double structural break)									
r=0	r≥1	29.40	32.11		**36.78**	**34.43**	*30.97*			20.45			
r≤1	r=2	14.02	15.92		10.66	7.98	8.25			7.55			
r=0	r≥1	44.78	48.15	**60.86**				40.31	**71.49**		**60.04**	29.78	41.61
r≤1	r=2	22.41	24.93	8.62				10.23	12.04		9.93	14.07	14.69

Note. See Tab. 3.

The test provides similar indications to those previously reported, confirming the existence of one cointegrating relationship for Austria, Belgium, Finland, France, Italy and the Netherlands. The long-run structure of each model is specified according to the outcome of the stationarity test for the elements of $\mathbf{y}_t$ (Table 9, upper part) and turns out to be very close to the results for the pre-EMU years (Table 9, lower part). Cointegration residuals are plotted in Figure 4 (dashed line).

With the exception of Austria and Italy, there is clear evidence that bilateral real exchange rates are stationary around a segmented deterministic component.[11] The latter is highly statistically significant in the third subsample (with the exception of Finland), providing additional support for the introduction of the second structural break. In this subsample, for

[11] The test statistic for Austria (Italy) is 18.44 (5.43). Therefore, it is greater than the critical value at the 5% significance level for a $\chi^2(1)$ distribution equal to 3.84.

Belgium and France there is an upward shift of the intercept from the 1993-1998 “convergence” levels. For the Netherlands, the first years of the EMU correspond to the widening of the inflation differential with respect to Germany. In this case, the slope of the linear trend is around four times that of the first subsample, even though it has an opposite sign.

Tab.9 Exclusion test for the PPP hypothesis and long-run structure: 1984-2002

Exclusion test for each element in $\mathbf{y}_t$						
$\chi^2(1)$ c.v.	5% = 3.84, 1% = 6.64					
	AUT	BEL	FIN	FRA	ITA	NET
$q_{it} = 0$	41.8	15.06	18.26	13.52	41.62	39.94
$q_{0t} = 0$	42.66	14.81	17.12	13.02	34.46	40.17

Long-run structure											
	q_{it}	q_{0t}	γ_1	γ_2	γ_3	δ_1	δ_2	δ_3	(d.f.)	stat.	[p-value]
AUT	1	-1.0462	-0.0002	0	-0.0004				$\chi^2(1)$	*0.25*	[0.62]
		(0.0003)	(2.6e-05)		(1.0e-05)						
BEL	1	-1				-0.0191	0	-0.0152	$\chi^2(2)$	**2.56**	[0.28]
						(0.0037)		(0.0052)			
IN	1	-1				-0.2487	0	0	$\chi^2(3)$	**6.77**	[0.08]
						(0.0158)					
FRA	1	-1				-0.0720	-0.0225	-0.0313	$\chi^2(1)$	**1.40**	[0.24]
						(0.0050)	(0.0058)	(0.0069)			
ITA	1	-0.7807	-0.0032	-0.0023	0				$\chi^2(1)$	*1.55*	[0.21]
		(0.0022)	(0.0002)	(1.0e-05)							
NET	1	-1	0.0004	0	-0.0013				$\chi^2(2)$	**4.69**	[0.10]
			(2.6e-05)		(1.0e-05)						

Note. See Tab. 4.

- The trace test for the model specified in Table 5, i.e. including those countries for which the PPP condition does not hold, confirms the existence of one cointegrating relationship, as shown in Table 10.

Tab. 10 Trace test for the GPPP hypothesis: 1984-2002

Trace test (double structural break)							
H_0	r=0	r≤1	r≤2	r≤3	r≤4	r≤5	r≤6
H_1	r≥1	r≥2	r≥3	r≥4	r≥5	r≥6	r=7
90% c.v.	211.1	170.2	133.02	99.71	70.36	44.78	22.41
95% c.v.	217.88	176.33	138.49	104.51	74.47	48.15	24.93
stat.	**241.51**	163.31	109.43	71.17	41.7	19.22	7.13

Note. See Tab. 5.

Table 11 reports exclusion test results for each element of $\mathbf{y}_t$ (upper part) and the final specification of the cointegrating vector normalized on the prices of Austria[12] (lower part).

Tab. 11 Exclusion test for the GPPP hypothesis and long-run structure: 1984-2002

Exclusion test for each element in $\mathbf{y}_t$									
	AUT	IRE	ITA	LUX	POR	SPA	GER	$\chi^2(1)$ 5% c.v.	$\chi^2(1)$ 1% c.v.
$q_{it} = 0$	19.50	**0.30**	8.56	17.12	10.73	**0.64**	8.10	3.84	6.64

Long-run structure												
q_{it} AUT	IRE	ITA	LUX	POR	SPA	GER	γ_1	γ_2	γ_3	(d.f.)	stat.	p-value
1	0	-0.1219	-0.6495	0.2246	0	-0.4532	-0.00080	0	-0.00015	$\chi^2(4)$	**5.83**	0.21
		(0.0059)	(0.0588)	(0.0104)		(0.055)	(3.5e-05)		(1.0e-05)			

Note. See Tab. 6.

From an econometric point of view the estimated equilibrium condition confirms that: i) variables referred to Ireland and Spain do not belong to the cointegration space; ii) real exchanges rates for Austria and Luxemburg, as well as those of Portugal and Italy, show strong similarity with each other and constitute (all four together) a stationary linear combination. From an economic point of view such relationship also suggest that: i) endogenous forces arising from the formation of the EMU seem too weak to narrow the existing heterogeneities within European countries; ii) the slope of the linear trend in the third subsample suggests that the (temporary) convergence process in the late Nineties is followed by a (temporary) divergence with respect to German prices in the first years of the EMU, as shown in Figure 5 (dashed line). Unlike the previous years, this result seems mainly attributable to prices too low in Germany rather than prices too high in other European countries.

- Table 12 reports the feedback coefficients for the models estimated over the period 1984M1-2002M12, i.e. taking into account the introduction of the single currency.

[12] The unrestricted long-run relationship is very similar to the one presented in the text.

Tab. 12 **Feedback coefficients: 1984-2002**

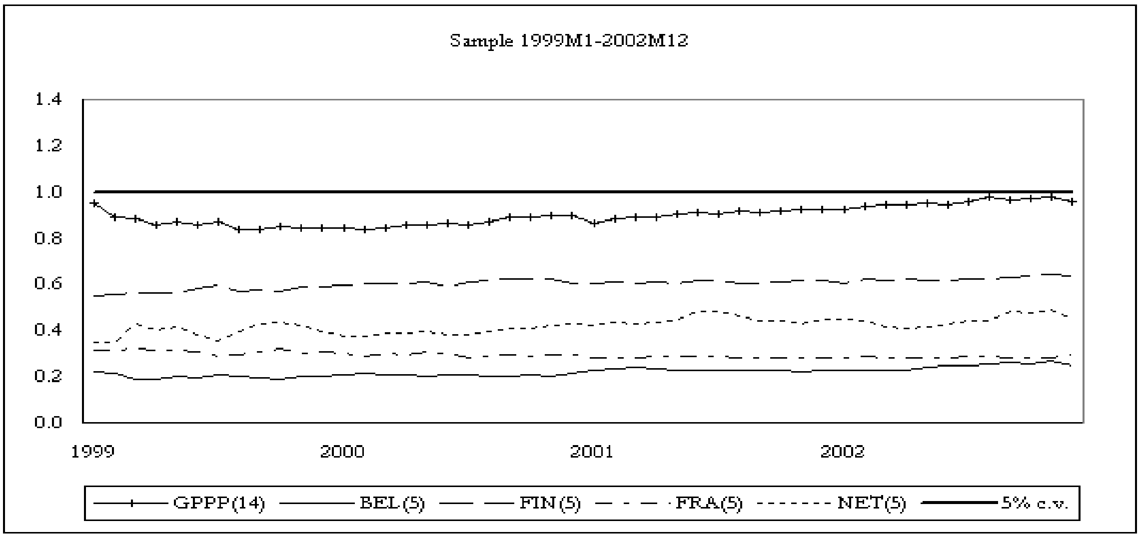
	BEL		FIN		FRA		NET	
α_i	-0.1272		-0.0813		-0.1007		-0.2935	
	(0.0270)		(0.0188)		(0.0266)		(0.0424)	
α_0	0		0		0		0	
$\chi^2(5)$	2.77	[0.74]	7.07	[0.22]	3.27	[0.66]	4.97	[0.42]

	AUT	IRE	ITA	LUX	POR	SPA	GER
α_i	-0.2658	0	0	0.1101	0	0	0
	(0.0383)			(0.0590)			
$\chi^2(14)$	18.23 [0.07]						

Note. See Tab. 7.

The over-identified structure of the long-run matrix, $\mathbf{\Pi}$, turns out to be stable in each model, as shown in Figure 6.

Fig. 6 **Long-run matrix stability test: 1999M1-2002M12**



Note. See Fig. 5.

Short-run adjustments confirm the results for the pre-EMU years. Peripheral countries differ from those belonging to the Core Europe because the former lack an equilibrium restoration mechanism. This finding assumes a more “alarming” connotation with respect to the pre-EMU period, because intra-area nominal exchange rate fluctuations no longer can substitute for such mechanism.

5 CONCLUSIONS

This paper aims at contributing to the debate on whether the euro area can be defined as an OCA. The empirical work focuses on a specific topic of applied economics and investigates the process of price convergence within the EMU. Such convergence represents a necessary condition in order to stabilize both the nominal (explicit policy target) and the real exchange rate (implicit policy target), allowing to safeguard member countries' intra-regional competitiveness and to avoid the incentive to implement "beggar thy neighbours" policies.

Cointegration techniques are used to test the GPPP hypothesis, after the preliminary assessment of the stationarity of each bilateral real exchange rate. The econometric methodology is based on VEC models with broken deterministic components, and provides with robust results because it is expressly designed to reduce the probability to erroneously reject the cointegration hypothesis due to the presence of segmented (instead of unbroken) deterministic components.

The overall picture emerging from the estimates suggests that the EMU is an integrated area with the exception of Spain and Ireland. However, a certain degree of heterogeneity among national aggregate demand functions still exists. Moreover, the empirical evidence suggests that the "euro-effect" in fostering the integration within the EMU has been quite weak so far, in accordance with the most recent literature on the euro area (see, for example, de Nardis and Vicarelli, 2003). Even though a more precise assessment of the consequences at the national level arising from the integration process in Europe calls for a larger time horizon, these findings suggest that the convergence process within the euro area has not been pushed much further in recent years and that additional steps towards integration need to be done in order to properly make endogenous forces work as predicted by the theory of OCAs.

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