THE EURO-DOLLAR EXCHANGE RATE: IS IT FUNDAMENTAL?

MARIAM CAMARERO
JAVIER ORDÓÑEZ
CECILIO TAMARIT

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Abstract

In this paper we have applied two approaches to the study of the dollar real exchange rate in relation with the Euro-area currencies. First, using dynamic panel techniques, we estimate an error correction model for the dollar real exchange rate versus seven developed countries, four of them Euro-area members. Second, we aggregate the European variables and estimate a model for the Euro-dollar real exchange rate using time series techniques. After identification and model selection, the same specification can be adopted in the two cases, in an eclectic model including real interest rate and productivity differentials, together with relative fiscal policy and net foreign asset positions. This model turns out to be compatible with the very recent results obtained in the context of the New Open Macroeconomics literature.

Keywords: real exchange rate, cointegration, time-series, panel, dollar, Euro-zone.

JEL Classification: C33, F31.

Mariam Camarero
Department of Economics
Jaume I University
Campus de Riu Sec
E-12080 Castellón
Spain
camarero@eco.uji.es

Javier Ordóñez
Department of Economics
Jaume I University
Campus de Riu Sec
E-12080 Castellón
Spain

Cecilio Tamarit
Economics Department II
University of Valencia
P.O.Box 22 006
46071 Valencia
Spain

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1 Introduction.

The evolution of the euro exchange rate vis-à-vis the main international currencies, and particularly, towards the US dollar has given birth to a growing amount of literature. Contrary to the, more or less, general expectations of appreciation, the euro has spent its first three years of existence depreciating against the dollar. Although many arguments have been given in "search of fundamentals" the results are up to now quite discouraging driving to puzzling outcomes (see, for instance, De Grauwe, (2000) or Meredith (2001). Two arguments can be forwarded in order to justify this fact. First, an analysis based on fundamentals cannot be carried out on a short term basis. However, the operators in the money markets seem to be working in a chartist world. On the contrary, from a policy oriented interest analysis, the span of the data set has to be long enough to capture the long run equilibria relationships, being an econometric framework based on cointegration the most appropriate methodology for this purpose. Second, and in connection to the former argument, the absence of historical data for the euro makes necessary the use of aggregate variables (ECB, 2000). This "synthetic" euro and the aggregate euro area variables have an important caveat: they summarize the evolution of the legacy currencies which developed in the framework of rather heterogeneous economic environments[1]. This heterogenous behavior and its importance for the "strength" of the euro was pointed out by De Grauwe (1997). Therefore, in this paper, we propose a complementary methodology in order to overcome these problems. First, we use the Pooled Mean Group (PMG) estimator proposed by Pesaran, Shin and Smith (1999) for non-stationary regressors and estimate a panel for a group of Euro-area currencies. This method constrains the long-run coefficients to be identical, but allows error variances and short-run parameters to differ. This methodology allows us to capture the long run relationships consistently with the medium and long run orientation of the fundamentals exchange rate models and the targets of the European monetary policy. At the same time, it permits us to grasp the different behavior of the euro area countries. Second, we propose the estimation of an aggregate bilateral exchange rate model between the dollar and the euro/ecu using standard Johansen's cointegration analysis methodology in order to find the long-run determinants of the real exchange rate based on the current values of the variables. Under this framework we are also able to test for regime shifts or structural breaks. However, we must bear in mind that these changes can only be detected with a significant delay. Thus, even if the creation of the European Monetary Union has provoked a change in regime, it is still too early to be able to detect it using the available techniques.

The remainder of the paper is organized as follows. Section 2 provides and overview of the recent empirical literature on the issue of exchange rate determination in the euro case. Section 3 describes the theoretical models, whereas the next section presents the econometric results. Finally, in section 5 we report the main results and conclusions.

[1] See ECB (2002)

2 An overview of the recent empirical literature[2].

A traditional starting point for estimating equilibrium exchange rate has been the PPP theory, either in its absolute or relative version. However, due to a different bulk of factors well documented in the literature, the speed of adjustment of the current value of exchange rate to the long run equilibrium is very slow. Therefore, other approaches have been implemented over time. Basically, they can be classified in two strands of literature: first, the so-called "fundamental equilibrium exchange rate" (FEER), and secondly, the "behavioral equilibrium exchange rates" (BEER)[3].A well known caveat of the first approach is its normative nature. This is due to the fact that under the FEER approach the exchange rate has to be consistent with internal and external balance. Thus, we think, according to Clark and MacDonald (1999), that the behavioral approach can be a better empirical approach to exchange rate modelling since its computation is based on current levels of the fundamental factors. Now, the problem is to determine the correct combination of fundamental variables and the answer in mainly empirical. Using different econometric techniques several studies have been implemented for the past two years following the behavioral approach. Alberola et al. (1999) using cointegration techniques for individual currencies as well as for a panel of currencies find only a long run relation with net foreign assets and relative sectoral prices (Balassa-Samuelson effect), Ledo and Taguas (1999) find that the deviations from PPP can be explained largely by productivity differentials and interest rate differentials in an error correction model. Additionally, Closterman and Schnatz (2000) find an equilibrium relationship for the bilateral euro-dollar exchange rate that includes the productivity differential, the interest rate differential, the real oil price and the relative fiscal position. Makrydakis et al (2000) find a relation with the productivity differential and the real interest rate differential as in Alquist and Chinn (2001). Finally, Maeso-Fernández et al. (2001) find that the euro appears to be mainly affected by productivity developments, real interest rate differentials and external shocks due to oil dependence of the euro area. It seems to be that all the models taken together encompass useful information, so that any assessment about the evolution of the real exchange rate should initially build to some extent on such a broad-based multi-approach analysis (ECB, 2002).

3 Theoretical models: an eclectic nested approach.

As mentioned in the previous section in reference to the case euro-dollar, but true in general, the most recent empirical evidence on real exchange rates has

[2]For a complete overview of different empirical approaches, see Williamson (1994) or more recently, MacDonald (2000).

[3]For the sake of simplicity we are omitting the NATREX and the PEER aproaches. We consider that the first one would be clearly connected to the FEER approach and the second to the BEER approach.

not been able to find stable relationships in accordance with the traditional theoretical models. In search of an answer to the problems associated with modelling exchange rates and, in particular, real exchange rates, MacDonald (1998) proposes what he calls an eclectic approach to model real exchange rates.

In a seminal paper, Meese and Rogoff's (1988) studied the link between real exchange rates and real interest rate differentials trying to solve part of the problems related to the monetary models. They define the real exchange rate, q_t, as $q_t \equiv e_t - p_t + p_t^*$, where e_t is the price of a unit of foreign currency in terms of domestic currency and p_t and p_t^* are the logarithms of domestic and foreign prices. Three assumptions are made: first, that when a shock occurs, the real exchange rate returns to its equilibrium value at a constant rate; second, that the long-run real exchange rate, $\hat{q}_t$, is a non-stationary variable; finally, that uncovered real interest rate parity is fulfilled.

Combining the three assumptions above, the real exchange rate can be expressed in the following form:

$$q_t = -\varphi(R_t - R_t^*) + \hat{q}_t \tag{1}$$

where R_t^* and R_t are, respectively, the real foreign and domestic interest rates for an asset of maturity k. This leaves relatively open the question of which are the determinants of $\hat{q}_t$ which is a non-stationary variable.

This model has been very influential in the empirical literature. As Edison and Melick (1995) describe in their paper, the implementation of the empirical tests depends on the treatment of the expected real exchange rate derived from equation (1). The simplest model will assume that the expected real exchange rate is constant, while the models including other variables will specify it using other determinants.

This model was first tested, in its simplest version, in the well-known works of Campbell and Clarida (1987) and Meese and Rogoff (1988). The former paper finds that little of the movement in real exchange rates can be explained by movements in real interest differentials. Also Meese and Rogoff (1988), using cointegration techniques (Engle and Granger single equation tests) cannot find a long-run relationship between the two variables. However, Baxter (1994) found more encouraging results and, in a recent paper, MacDonald and Nagayasu (2000) tested this relationship for 14 industrialized countries using both long and short-run real interest rate differentials and time series as well as panel cointegration methods. After obtaining evidence of statistically significant long-run relationships and plausible point estimates using panel tests, they conclude that the failure of previous researches may be due to the estimation method used rather than to any theoretical deficiency.

In a second group of papers, the assumption that the expected real exchange rate is constant is relaxed and they try to explain it using additional variables. This approach was first introduced by Hooper and Morton (1982) who modelled the expected real exchange rate as a function of cumulated current account. Edison and Pauls (1993) and Edison and Melick (1995) estimate this kind of model using cointegration techniques. While the second paper finds evidence of a cointegrating relationship, Edison and Pauls (1993) fail to find a statistical link between real exchange rates and real interest rates using the Engle-Granger methodology. However, the estimated error correction models are more supportive of such a relation. Wu (1999) has recently obtained also good results (even in terms of forecasting ability) for this type of specification in the cases of Germany and Japan versus the dollar and using the Johansen technique.

MacDonald (1998) also follows this approach, dividing the real exchange rate determinants into two components: the real interest rate differential and a set of fundamentals that explains the behavior of the long-run (equilibrium) real exchange rate, which include productivity differentials, the effect of relative fiscal balances on the equilibrium real exchange rate, the private sector savings and the real price of oil.

We will describe in more detail this eclectic approach, that will be the basis of our analysis.

He assumes that PPP holds for non-traded goods and arrives to the following expression for the long-run equilibrium real exchange rate:

$$\hat{q}_t \equiv q_t^T + \alpha_t(p_t^T - p_t^{NT}) - \alpha_t^*(p_t^{T^*} - p_t^{NT^*}) \tag{2}$$

where q_t^T is the real exchange rate for traded goods; $(p_t^T - p_t^{NT}) - (p_t^{T^*} - p_t^{NT^*})$ is the relative price of traded to non-traded goods between the home and the foreign country and α and α^* are the weights..

Based on (2), MacDonald identifies two potential sources of variation in the equilibrium real exchange rate:

1. Movements in the relative prices of traded to non-traded goods between the home and foreign country (second and third terms in (2)). These differences are likely to be concentrated in the non-traded goods. In particular, according to the traditional Balassa-Samuelson effect, productivity differences in the production of traded goods across countries can cause a bias into the overall real exchange rate, because productivity advances tend to concentrate in the traded goods sectors. Due to the linkages between prices of goods and wages (and wages across sectors), provided that there is internal factor mobility (from the non-traded to the traded goods sectors and conversely), the real exchange rate tends to appreciate for fast growing economies.

2. Non-constancy of the real exchange rate for traded goods (the first term in (2)). Two additional factors may introduce variability in q_t^T :

    - International differences in savings and investment. The real exchange rate for traded goods is also, following MacDonald (1998), a major determinant of the current account, that is in turn driven by the determinants of savings and investment. We can separate two variables that may capture this effect:

- Fiscal policy, whose relation with the real exchange rate depends on the approach. According to the Mundell-Fleming model, an expansionary fiscal policy reduces national savings, increases the domestic real interest rate and generates a permanent appreciation. In contrast, the portfolio balance models consider that a permanent fiscal expansion would cause a decrease in net foreign assets and a depreciation of the currency.
- Private sector net savings may also affect the real exchange rate, influenced in turn by demographic factors. Thus, the cross-country variations of saving rates may affect the relative net foreign asset position.

- Changes in the real price of oil, that tends to appreciate the currencies of the net oil exporters or, in general, the currencies of the less energy dependent countries.

MacDonald's proposal does not rely exclusively on the monetary approach to exchange rate determination, although captures the majority of the fundamental variables mentioned in the literature and makes them compatible with it. Accordingly, the above mentioned factors can be summarized in the following empirical specification:

$$\begin{aligned} q_t &= -\varphi(R_t - R_t^*) + \hat{q}_t = \\ &= f(\underset{(-)}{(R_t - R_t^*)}, \underset{(-)}{(a_t - a_t^*)}, \underset{(-/+)}{(g_t - g_t^*)}, \underset{(-)}{oil_t}, \underset{(-)}{dnfa_t}) \end{aligned} \tag{3}$$

where $(a_t - a_t^*)$ is the difference between the domestic and foreign economies productivity[4], $(g_t - g_t^*)$ is the public expenditure differential, oil_t[5] is the real oil price and $dnfa_t$ is the relative net foreign asset position of the economy.

4 Empirical results.

Two different econometric techniques have been applied to the same data set. First, using dynamic panel techniques[6], the real exchange rate of the dollar versus a group of seven individual countries has been estimated. At the same time, we have studied separately the Euro-countries in the sample from the rest. Second, the dollar-euro real exchange rate is explained using Euro-area aggregated variables using time series techniques.

[4] The breakdown between traded and non-traded goods has not been possible for the sample period, the OECD data available only reaching 1992.

[5] Hamilton (1983) found that the energy price can account for innovations in many US macroeconomic variables. Amano and van Norden (1998) find a stable link between the effective real exchange rate of the dollar and the oil price shocks. They also think that these shocks account for most of the major movements in the terms of trade. According to them, the correlation between the terms of trade and the one-period lagged price of oil is -0.57, -0.78 and -0.92 for the US, Japan and Germany, respectively.

[6] See Appendix A for a detailed description of the Pooled Mean Group Estimates, a technique proposed by Pesaran, Shin and Smith (1999).

4.1 Panel analysis: the dollar in the world.

As already described in the theoretical section of the paper, a wide set of explanatory (fundamental) variables was examined in order to assess the main factors behind the behavior of the dollar real exchange rate. In this first part of the analysis, the countries involved are eight: the US as the domestic country, Japan, Canada and 4 European countries (those with information available for the sample period and variables of interest). Consequently, this first part of the analysis is not strictly a model for the dollar versus the Euro-area. We have chosen to include countries (such as the UK) that do not participate in EMU, as well as Canada and Japan, in order to capture the behavior of the most important world currencies. The methodology used in this part of the analysis will allow for both group and individual approaches. Thus, we consider first, the whole group of countries (where $N = 7$) and, then, we divide the panel into the Euro-area countries ($N = 4$: Germany, Spain, France and Italy) and non Euro-area ($N = 3$: Canada, Japan and the UK). The data are quarterly and the sample goes from 1970:Q1 to 1998:Q4[7].

In the process of selection of the model specification we have tried to follow as close as possible the general to specific methodology. Then, taking as a starting point the models described in the previous section and, in order to make the estimated models comparable, we use a general specification:

$$rerdol_{it} = f(\underset{(-)}{dpro_{it}}, \underset{(-)}{drr_{it}}, \underset{(-)}{oildep}, \underset{(-)}{dnfa}, \underset{(+/-)}{dpex})$$

where $rerdol_{it}$ is the real exchange rate of the dollar versus all the currencies defined as the units of domestic currency necessary to buy a unit of foreign currency in real terms; $dpro_{it}$ is the relative productivity of the US versus each of the other countries: an increase in the value of this variable tends to appreciate the currency.; drr_{it} is the real interest rate differential between the US and the other countries analyzed: an increase in this differential appreciates the currency; finally, $oildep_{it}$ is the real price of oil adjusted by the relative dependency on oil imports in each country as compared to the US: in this case, the dollar will appreciate when the oil dependency of the foreign countries is increasing; $dnfa_{it}$ is the difference in the net foreign asset position over GDP of the US versus the other countries, and the sign should be negative: the currencies of countries increasing its net foreign asset position tend to appreciate; $dpex_{it}$ is the difference in public expenditure over GDP between the US and each of the other countries: in this case, there are two competing theories explaining its relation with the real exchange rate, so that the relation would be positive (depreciation) if the portfolio balance model prevails, whereas a negative sign is associated with the Mundell-Fleming approach.

The models proposed are the following[8]:

[7] A detailed description of the variables can be found in Appendix B. At this stage of the analysis the panel could have been unbalanced, although this has not been the case. We have preferred to exclude two countries (Finland and Sweden) that had only very limited information, due to the distortions they caused.

[8] In addition, other specifications have been estimated in the empirical part of the model.

Model 1: $rerdol_{it} = \alpha_i + \beta_{1i} drr_{it} + \beta_{2i} dpex_{it} + \beta_{3i} dpro_{it} + \beta_{4i} dnfa_{it} + \beta_{5i} oildep_{it}$

Eclectic model.

Model 2: $rerdol_{it} = \alpha_i + \beta_{1i} drr_{it} + \beta_{2i} dpex_{it} + \beta_{3i} dpro_{it} + \beta_{4i} dnfa_{it}$

Restricted eclectic model.

The first model is the general specification described above, whereas the second model is a partial version of model 1, where the oil dependence variable has been excluded. In what follows, these empirical models are tested.

4.1.1 Order of integration of the variables.

Bearing all these considerations in mind, we should start the analysis with the study of the order of integration of the variables. Several panel unit root tests are already available in the literature, from the early works of Levin and Lin (1992)[9], to the Im, Pesaran and Shin (1995) tests. However, due to its higher power, in this section we have applied the LM test for the null of stationarity proposed by Hadri (2000) with heterogeneous and serially correlated errors. These tests can be considered the panel version of the $KPSS$ tests applied in the univariate context. Hadri (2000) proposes two models (with and without a deterministic trend) and their decomposition into the sum of a random walk and a stationary disturbance term. He tests the null hypothesis that all the variables (y_{it}) are stationary (around deterministic levels or around deterministic trends), so that for the N elements of the panel the variance of the errors is such that:

$$H_0 : \sigma^2_{u1} = ... = \sigma^2_{uN} = 0 \tag{4}$$

against the alternative H_1 : that some $\sigma^2_{ui} > 0$. This alternative allows for heterogeneous σ^2_{ui} across the cross-sections and includes the homogeneous alternative ($\sigma^2_{ui} = \sigma^2_u$ for all i) as a special case. It also allows for a subset of cross-sections to be stationary under the alternative. The two statistics are called η_μ for the null of stationarity around an intercept and η_τ when the null is stationarity around a deterministic trend.

The results of the tests applied to the four variables involved are presented in Table 1. The null hypothesis of stationarity can be easily rejected in the two cases (with and without time trend), so that all the panel variables can be considered non stationary.

In particular, the simplest version of the Meese and Rogoff (1988) model ($rerdol_{it} = \alpha_i + \beta_{1i} drr_{it}$), as well as the Rogoff (1992) intertemporal model ($rerdol_{it} = \alpha_i + \beta_{1i} dpex + \beta_{2i} dpro_{it} + \beta_{3i} oildep_{it}$). In the first case, although the Information Criteria were encouraging, the model was not very explanatory (with $\bar{R}^2$ under 0.10 for the individual countries). As for the Rogoff (1992) model, none of the hypotheses concerning the long-run parameters was accepted, and the Information Criteria did not recommend its choice. The results, although not reported in the present paper, are available upon request.

[9] Finally published as Levin, Lin and Chu (2002).

Table 1 Hadri (2000) stationarity tests
($l = 2$)

Variables	η_μ	η_τ
$rerdol_{it}$	23.72**	175.45**
$dpex_{it}$	14.30**	262.49**
$dnfa_{it}$	47.05**	1655.32**
$dpro_{it}$	29.79**	801.21**
drr_{it}	18.23**	167.71**
$oildep_{it}$	18.38**	149.01**

Note: The statistic η_μ does not include a time trend, whereas η_τ does, and are normally distributed. The two asterisks denote rejection of the null hypothesis of stationarity at 5%. The number of lags selected is $l = 2$.

4.1.2 Long-run relationships: Pooled Mean Group estimation results.

Once the order of integration of the variables has been determined, for the analysis of the real exchange rate of the dollar we have followed the methodology proposed by Pesaran, Shin and Smith (1999) and computed the Pooled Mean Group estimators. Due to the specific purpose of this paper, this estimation technique turns out to be specially suited. In our case, we work with seven groups, i.e. four Euro-area countries (Germany, Spain, France and Italy) and three non-Euro countries (Canada, Japan and the UK).

The empirical model is based on the eclectic formulation presented above, starting with Model 1, that includes the main explanatory variables proposed by the literature of real exchange rates. Other theoretical models are restricted versions of Model 1.

Many empirical specifications have been estimated and compared, using Likelihood-based Information Criteria, such as the AIC and the SBC. In addition, in each of these specifications we have tested two particularly important questions: first, the homogeneity restriction using a likelihood ratio test; second, the existence of discrepancies between the Pooled Mean Group Estimates and the Mean Group Estimates, that differ also in the degree of heterogeneity allowed. The Hausman test permits us to decide whether these discrepancies recommend the exclusion of the homogeneity restriction in some of the long-run parameters. Thus, this second test complements the first one, because if homogeneity is rejected using the LR test, the Hausman test for the individual variables helps to identify the variable source of the heterogeneity. Concerning the dynamics of the model, the short-run has been modelled using just two lags (even one for some of the explanatory variables), as derived from the application of the Schwarz Bayesian Criterion for lag selection.

In Table 2 we present the Information Criteria of the two models selected according to this type of information, together with the LR homogeneity tests for the panel, as well as the concrete hypotheses tested. It should be emphasized that we have estimated three different groups of countries: first, the seven countries together; second, the four countries in the Euro-area and, third, three countries that do not belong to the Euro-area, that is, Canada, Japan and the UK. Some slight differences have been detected depending on the group of countries considered, so that we have maintained the three different groups in this part of the analysis.

Table 2

Comparison of the specified models

	$N=7$			*Variables*				
	AIC	*SBC*	*LR test*	drr_t	$dpex_t$	$oild_t$	$dnfa_t$	$dpro_t$
Model 1	1714	1686	$\chi^2(18)=67.81[0.00]$	$\neq$	$=\forall$	$\neq$	$=\forall$	$=\forall$
Model 2	1691	1665	$\chi^2(12)=20.68[0.05]^{**}$	$\neq$	$=\forall$	—	$\neq$	$=\forall$
	$N=4$: Euro-area			*Variables*				
	AIC	*SBC*	*LR test*	drr_t	$dpex_t$	$oild_t$	$dnfa_t$	$dpro_t$
Model 1	1036	1018	$\chi^2(6)=17.95[0.00]$	$\neq$	$\neq$	$\neq$	$=\forall$	$=\forall$
Model 2	998	982	$\chi^2(9)=15.87[0.07]^{**}$	$\neq$	$=\forall$	—	$=\forall$	$=\forall$
			$\chi^2(6)=11.37[0.07]^{**}$	$\neq$	$=\forall$	—	$=\forall$	$\neq$
	$N=3$: Non-Euro			*Variables*				
	AIC	*SBC*	*LR test*	drr_t	$dpex_t$	$oild_t$	$dnfa_t$	$dpro_t$
Model 1	763.74	748.40	$\chi^2(4)=28.51[0.00]$	$\neq$	$\neq$	$\neq$	$=\forall$	$=\forall$
Model 2	763	750	$\chi^2(4)=8.55[0.07]^{**}$	$\neq$	$=\forall$	—	$\neq$	$=\forall$

Note: *AIC* stands for Akaike Information Criterium, *SBC* for Swartz Bayesian Criterium and *LR test* is the Likelihood Ratio Test for equality of either some or all the long-run parameters (probability values appear in parentheses). Two asterisks denote acceptance of the restriction on the long-run parameters at 5% significance level. $\neq$ stands for the assumption of different parameter values for all the N members of the panel, whereas the homogeneity hypothesis is represented by the symbols $=\forall$.

In Model 1, the one including all the variables considered, has higher *AIC* and *SBC* than model 2. Moreover, for none of the groups of countries analyzed the null hypothesis of homogeneity in the long-run parameters can be accepted (see, for example, for $N=7$, $\chi^2(18)=67.81$ with a probability of [0.00]). In addition, the long-run parameter of the variable $oildep_t$ is non-significant. When some heterogeneity was allowed, specifically in the oil dependency variable, the results did not improve[10].

Model 2 is a restricted version of Model 1, where $oildep_t$ has been excluded. The Information Criteria are smaller and, after imposing that not all the long-run parameters are equal for all the countries, the restriction can be accepted for the rest of the variables in the three configurations adopted. If we begin with $N=7$, the homogeneity restriction is accepted for $dpro_t$ and $dpex_t$ ($\chi^2(12)=20.68$ with a probability of [0.05]), whereas there is necessary some heterogeneity in the real interest rate and in the net foreign asset differential. The estimates and the associated t-statistics are presented in the first column of table 3, where all the variables are significant, the only exception being drr_t. It should be noted that the error correction coefficient is highly significant and of a reasonable magnitude (-0.120), so that the adjustment towards

[10] All the results concerning this specification are available upon request.

equili brium takes approximately two years. In Tables 4 and 5, the information concerning the long-run relations between the individual countries and the misspecification tests is provided. Apart from some normality departures in some of the countries, the individual equations pass the misspecification tests. Moreover, the $\bar{R}^2$ of all the equations (with the only exception of Canada) is over 0.80.

Table 3

Pooled Mean Group Estimates
Model 2

$$rerdol_{it} = \alpha_i + \beta_{1i} dpro_{it} + \beta_{2i} drr_{it} + \beta_{3i} dnfa_{it} + \beta_{4i} dpex_{it}$$

Variables	*All countries* $(N=7)$	*Euro-area* $(N=4)$		*Non-euro* $(N=3)$
drr_t	-0.005[a]	-0.007[a]	-0.006[a]	-0.008[a]
	(-1.58)	(-1.92)	(-2.38)	(-2.23)
$dpex_t$	0.003	0.003	0.002	0.008
	(2.95)	(2.48)	(2.09)	(2.72)
$dpro_t$	-0.851	-0.870	-0.749[a]	-0.836
	(-27.02)	(-22.34)	(-7.12)	(-15.47)
$dnfa_t$	-0.327[a]	-0.314	-0.288	-0.266[a]
	(-5.57)	(-6.94)	(-6.58)	(-1.59)
ecm_{t-1}	-0.120	-0.126	-0.134	-0.149
	(-3.83)	(-2.99)	(-3.15)	(-4.77)

Note: t-Students in parentheses. [a] indicates that the corresponding variable was not subject to the restriction of equal long-run paramters for all the members of the group. Thus, its estimate is the Mean Group Estimate, instead of the PMGE.

The estimated parameters present the correct signs, as already described in the theoretical section of the paper. First, an increase in the real interest differential tends to appreciate the currency ($\beta_1 < 0$). Next, an expansionary fiscal policy in the US relative to the other countries tends to depreciate the currency ($\beta_2 > 0$), whereas an increase in relative productivity appreciates the currency ($\beta_3 < 0$) due to the Balassa-Samuelson effect. Finally, an increase in the relative net foreign assets position also provokes an appreciation ($\beta_4 < 0$). Notice that in the case of the long-run parameter estimates of drr_t and $dnfa_t$,we do not impose equality of all the cross-section elements. The individual countries' estimates are presented in detail in Table 4.

Although this technique has more advantages the larger N, due to the purpose of this paper, that focuses on the Euro-area, we have also estimated the dynamic panel for the four EMU countries with information available, as well as for the other three countries considered. The results of the long-run parameters estimates, also presented in Table 3, are very similar to those obtained for the larger group.

Table 4
Individual countries estimates

Model 2

Countries	$N=7$ drr_t	$dpro_t$	$dnfa_t$	$dpex_t$	ecm_{t-1}	$N=4$ drr_t	$dpro_t$	$dnfa_t$	$dpex_t$	ecm_{t-1}
Germany	-0.005	−0.851	−0.328	0.003	−0.120	−0.006	−0.74	−0.288	0.002	−0.128
	(-1.58)	(-27.02)	(-5.57)	(2.95)	(-3.83)	(-1.91)	(-8.52)	(-6.57)	(2.09)	(-3.92)
Spain	0.0001	−0.851	−0.372	0.003	−0.117	0.0001	−0.891	−0.288	0.002	−0.123
	(0.08)	(-27.02)	(-2.81)	(2.95)	(-2.99)	(0.09)	(-8.39)	(-6.57)	(2.09)	(-3.10)
France	−0.005	−0.851	−0.330	0.003	−0.215	−0.005	−0.907	−0.288	0.002	−0.246
	(-3.46)	(-27.02)	(-5.27)	(2.95)	(-4.52)	(-4.24)	(-20.85)	(-6.57)	(2.09)	(-4.90)
Italy	-0.004	−0.851	0.127	0.003	−0.096	-0.011	-0.454	−0.288	0.002	-0.039
	(-1.46)	(-27.02)	(1.16)	(2.95)	(-2.72)	(-1.08)	(-1.30)	(-6.57)	(2.09)	(-1.60)
Canada	−0.007	−0.851	−0.350	0.003	−0.144	–	–	–	–	–
	(-3.73)	(-27.02)	(-7.07)	(2.95)	(-4.15)					
Japan	−0.009	−0.851	−0.430	0.003	−0.126	–	–	–	–	–
	(-2.51)	(-27.02)	(-6.95)	(2.95)	(-3.33)					
UK	−0.003	−0.851	0.043	0.003	−0.217	–	–	–	–	–
	(-2.12)	(-27.02)	(2.07)	(2.95)	(-3.91)					

Countries	$N=3$ drr_t	$dpro_t$	$dnfa_t$	$dpex_t$	ecm_{t-1}
Germany	—	—	—	—	—
Spain	—	—	—	—	—
France	—	—	—	—	—
Italy	—	—	—	—	—
Canada	−0.008	−0.836	−0.383	0.007	−0.139
	(-3.79)	(-15.47)	(-15.47)	(2.72)	(-4.18)
Japan	−0.011	−0.836	−0.478	0.007	−0.100
	(-2.26)	(-15.47)	(-5.73)	(2.72)	(-2.95)
UK	−0.003	−0.836	0.063	0.007	−0.207
	(-2.28)	(-15.47)	(2.53)	(2.72)	(-3.84)

For the Euro-area countries, Table 2 shows again the Information Criteria (also smaller than in Model 1), as well as the *LR tests* for homogeneity in the long-run parameters. In this case, after imposing that drr_t is heterogeneous for the members of the group, the homogeneity of the other three explanatory variables can be accepted. However, as an additional test for homogeneity, the Hausman test for the variable $dpro_t$ did not accept the similarity between the coefficient estimated using the *PMG* estimator and the *MG* estimator, where heterogeneity is allowed[11]. Once the two variables are not constrained to be homogeneous, the model passes the Hausman test. However, we present in Table 3 the estimation results for the two cases, that are very similar. All the

[11] The p-values associated with the test for each of the variables are the following: $dpex_t$ [0.40], $dnfa_t$ [0.51] and $dpro_t$ [0.00].

variables are significant and the error correction term is slightly larger in the second case.

Table 5
Individual countries specification tests
Model 2 ($N = 7$)

	$\bar{R}^2$	Correl.	FF	NO	HE
Germany	0.882	0.71	17.43*	34.03*	36.82*
Spain	0.829	0.10	1.67	36.63*	0.03
France	0.890	0.18	1.87	4.39	1.07
Italy	0.850	3.71	1.21	35.98*	0.08
Canada	0.578	1.28	0.52	2.36	0.13
Japan	0.869	0.01	0.54	5.68	0.00
UK	0.844	1.09	0.33	25.79*	0.52

In the case of the other three countries (Canada, Japan and UK), the homogeneity of all the variables is rejected. Only after allowing heterogeneity in drr_t and $dnfa_t$ the homogeneity of the other long-run parameters can be accepted. Model 2 and Model 1 have very similar *AIC* and *SBC,* but only in the case of Model 2 after the restrictions have been imposed the partial homogeneity is accepted, being the test $\chi^2(4) = 8.55$ with a probability of [0.07]. Thus, also for $N = 3$, Model 2 seems adequate. Concerning the long-run estimates the parameters have similar magnitude if compared with the larger model. The only exception is $dpex_t$, whose value is 0.008 in contrast with 0.003. The error correction coefficient takes the value of -0.149 and an associated t-Student of -4.77.

4.2 Aggregate European results: the Euro and the Dollar.

The panel analysis has given us some clues about the behavior of the Dollar versus the main world currencies. As expected, the results do not fit in a simple model (such as the Meese and Rogoff (1988) real interest differential), but rather in an eclectic specification, that includes variables both from the demand and the supply-side of the economy. From the results, the role of productivity differentials supports the fulfillment of the Balassa-Samuelson effect. Moreover, the real interest rate differential is also present, although this is not the exclusive determinant of real exchange rate behavior: the fiscal policy and the net foreign asset position of the countries are also among the explanatory variables. The only variable that did not have a significant contribution from those considered was the real oil price. An additional conclusion that can be extracted from the dynamic panel analysis is that the model estimated for the dollar real exchange rate does not change very much with the different configurations of countries, with some minor exceptions already mentioned.

Once the panel analysis has been completed for the European countries separately, we focus on the "synthetic" Euro-area variables. The two approaches are complementary as the panel allows for an important degree of heterogeneity. In fact, the lack of heterogeneity is one of the main criticisms that are commonly associated with the aggregate analysis. If the results from these two complementary methodologies do not show important discrepancies, we could feel more confident when using the aggregate series for inference and policy analysis.

For this part of the analysis we use the Johansen (1995) methodology for the estimation and identification of cointegrated systems where differentials are no longer calculated for USA relative to each individual country but relative to a representative Euro-area variable.

Table 6
Cointegration Test Statistics

r	Eigenvalues	λ_{max}	λ_{max} (R)	λ_{max} 95%	Trace	Trace (R)	Trace 95%
0	0.3748	40.87*	32.42	39.4	122.7**	97.28*	94.2
1	0.3420	36.42*	28.88	33.5	81.78**	64.86	68.5
2	0.2791	28.48*	22.95	27.1	45.36	35.97	47.2
3	0.1085	9.995	7.927	21.0	16.88	13.39	29.7
4	0.0699	6.312	5.006	14.1	6.883	5.429	15.4
5	0.0065	0.571	0.452	3.8	0.571	0.452	3.8

Note: The critical values are given with 95% critical values based on a response surface fitted to the results of Osterward-Lenum (1992). (R) stands for the small-sample correction of both λ_{max} and *trace* tests statistics proposed by Reimers (1992). $*$ and $**$ denotes rejection of the null hypothesis at 5% and 1% significance level respectively.

In a first stage of the analysis, we studied the order of integration of the variables using a stationarity testing strategy in the context of the VAR system. All the variables turned out to be I(1)[12]. Next, table 6 shows the $\lambda_{\max}$ and *Trace* test statistics for the determination of the number of cointegration relationships[13]. The Reimers' adjusted $\lambda_{\max}$ and *Trace* test statistics are also provided. From this set of test statistics the results are inconclusive: the *Trace* test statistic fails to reject the existence of two cointegration vectors, whereas using the Reimers' adjusted *Trace* test statistic we fail to reject one cointegration vector. The difference is even more noticeable in the case of the $\lambda_{\max}$ test statistic: the non-adjusted statistic fails to reject the existence of three cointegration vectors; in contrast, the adjusted version rejects cointegration. To gain insight on the appropriate choice of the number of cointegration vectors we proceed to complement this evidence by analyzing the roots of the companion matrix: three of them are almost unity and other two are pretty close to unity, implying that five is the number of common stochastic trends. Moreover, when $r = 1$ is set the largest roots are removed, leaving no near unit root in the model, suggesting therefore that this is the appropriate choice for r. In addition, from the time path plot for each of the feasible cointegration vectors only the first one seems to be stationary. The recursive analysis of the

[12] The results are available upon request.

[13] The model has been specified with the constant unrestricted. Previous to this choice, the different possible specifications for the deterministic components were compared using the procedure suggested by Johansen (1996).

system also provides useful information regarding the existence of cointegration: the recursive time path of the non-adjusted Trace statistic suggests that at most there exist two cointegration vectors though one is the most sensible outcome. From all this evidence, the most feasible choice is the existence of one cointegration vector, that is, $p - r = 5$, where p is the number of common stochastic trends.

The cointegration vector is identified imposing the overidentifying restriction that the variable for energy dependence (*oildep*) is excluded from the long-run: the LR statistic is $\chi^2(1) = 3.43$ with a probability value of 0.06. The resulting cointegration vector takes the form (standard errors in parentheses):

$$q_t = \underset{(0.001)}{0.011}\, dpex_t - \underset{(0.001)}{0.007}\, drr_t - \underset{(0.033)}{0.77}\, dpro_t - \underset{(0.032)}{0.36}\, dnfa_t \tag{5}$$

At this stage of the analysis, we can already compare the results obtained using the *PMG* in the dynamic panel with the time series model using aggregate variables. Taking into account the results presented in Table 3 for Model 2, we can observe that the results are very similar. First, the variable relative oil dependency ($oildep_t$) that turned out not to be significant in the panel analysis can be also excluded from the time series cointegration vector. Second, the four variables have the same signs even if we are using quite different estimation techniques. Moreover, the parameters' estimates are not very different in magnitude, the only exception being the case of $dpex_t$, where the time series value is 0.011 and 0.002 for the panel. In other cases, the parameters are almost equal, as for the real interest differential (-0.007 for the aggregate model and -0.006 for the panel) or the productivity differential (-0.77 in the time series model and -0.749 in the panel)[14]. Finally, the net foreign asset position is also in a similar range: -0.36 in the aggregate model and -0.288 in the panel.

Once we have identified the cointegration vector, we formally test for weak exogeneity of the variables in the system. According to our results, all the variables appear to be weakly exogenous with the only exception of the real exchange rate. The joint hypothesis of weak exogeneity and the identifying restrictions on the cointegration space β are accepted: the LR statistic value is $\chi^2(6) = 11.16$ with a probability of 0.08. We present next the error correction model (*ECM* hereafter) for the univariate partial model (t-values in brackets):

$$\begin{aligned}\Delta q_t \;=\;& \underset{[5.010]}{0.291} - \underset{[-7.675]}{0.375}\,\Delta dpro_t - \underset{[-2.999]}{0.185}\,\Delta dpro_{t-1} - \underset{[-2.068]}{0.105}\,\Delta dpro_{t-2} \\ & - \underset{[-2.002]}{0.002}\,\Delta drr_{t-3} - \underset{[-5.007]}{0.184}\, ecm_{t-1} + \varepsilon_t \end{aligned} \tag{6}$$

Misspecification tests:

[14] The magnitude of this parameter also lies in the range commonly found in the empirical literature, as reported by Gregorio and Wolf (1994). According to them, this range is (-0.1,-1.0).

Residual correlation:	$F(5, 76) = 1.0856\,[0.3752]$
ARCH:	$F(4, 73) = 0.8310\,[0.5098]$
Normality:	$\chi^2(2) = 1.1128\,[0.5733]$
Heteroscedaticity (squares):	$F(10, 70) = 1.0960\,[0.3774]$
Heteroscedaticity (squares and cross-products):	$F(20, 60) = 1.1588\,[0.3203]$

where ε_t is a vector of disturbances, ecm_{t-1} is the cointegration vector (5).

The misspecification tests are reported above, and none of them rejects the null hypothesis that the model is correctly specified. In addition, we apply the Hansen and Johansen (1993) approach to test for parameter instability in the cointegration vector. Specifically, we test both whether the cointegration space and each of the parameters in the cointegration vector are stable. We also test for the stability of the loading parameters. If both α and β appear to be stable, we can conclude that our error correction model is well specified for the period analyzed.

Figure 3, panel *(a)*, shows the plot of the test for constancy of the cointegration space. The test statistic has been scaled by the 95% quantile in the χ^2-distribution so that unity corresponds to the 5% significance level. The test statistic for stability is obtained using both the *Z-representation* and the *R-representation* of our model. In the former, stability is analyzed by the recursive estimation of the whole model whereas in the latter the short-run dynamics are fixed and only the long-run parameters are reestimated. Thus, the *R-representation* is the relevant one to assess the stability of the cointegration space, which is clearly accepted.

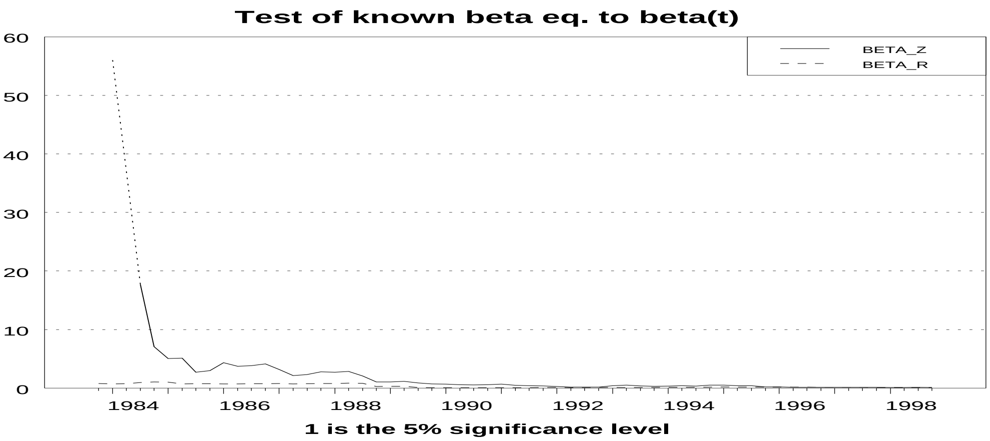
In Figure 3, panels *(b)* and *(c)* show, respectively, the stability tests for each of the beta coefficients and for the loadings to the cointegration vector. In all cases, the recursively estimated coefficients lay within the 95% confidence bounds showing a remarkable stability.

To summarize, we can conclude that the cointegration space is stable, that is, the long-run parameters as well as the loadings do not show signs of instability.

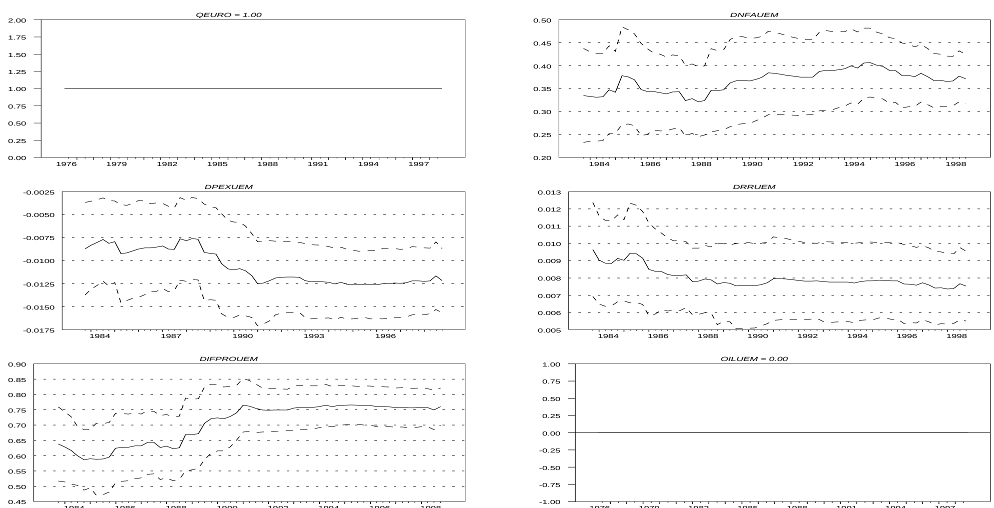
Finally, panel *(b)* in figure 4 presents several recursive tests of parameter stability for the parsimonious conditional model. Accordingly, our model is stable not only concerning the cointegration space but also the model as a whole.

As for the real exchange rate ECM, presented in equation (6), we should note that the error correction parameter presents the correct sign and magnitude (taking into account that the data are quarterly), and passes the Banerjee, Dolado and Mestre (1992) cointegration test. In addition, two are the variables that appear in the dynamics of the real exchange rate. First, with three lags, the real interest rate differential *(drr_t)*, although borderline significative. The negative parameter for this variable, as in the panel analysis, is the one expected from the theory. Second, the productivity differential measure, contemporaneous and lagged from one to two periods, with the same negative sign found in the long-run time series analysis and in the panel section of the paper. It should be also emphasized the important role that the productivity differential has in driving the system towards the equilibrium and the fact that

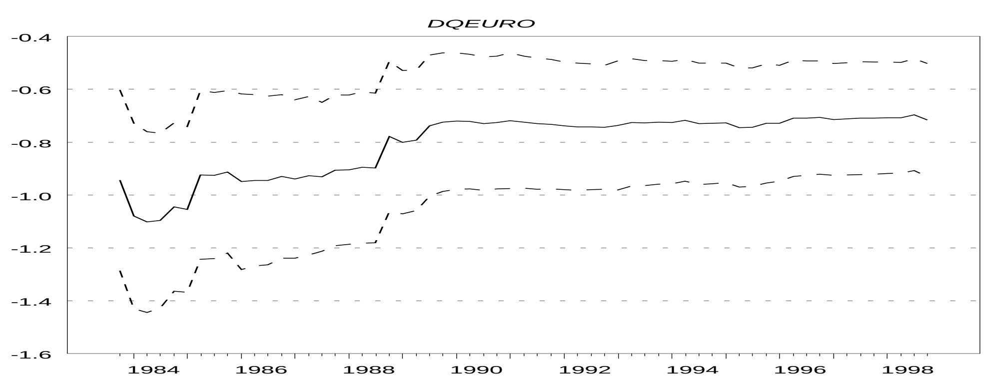
Figure 1: Stability of the cointegration space



(a) Test of constancy of beta

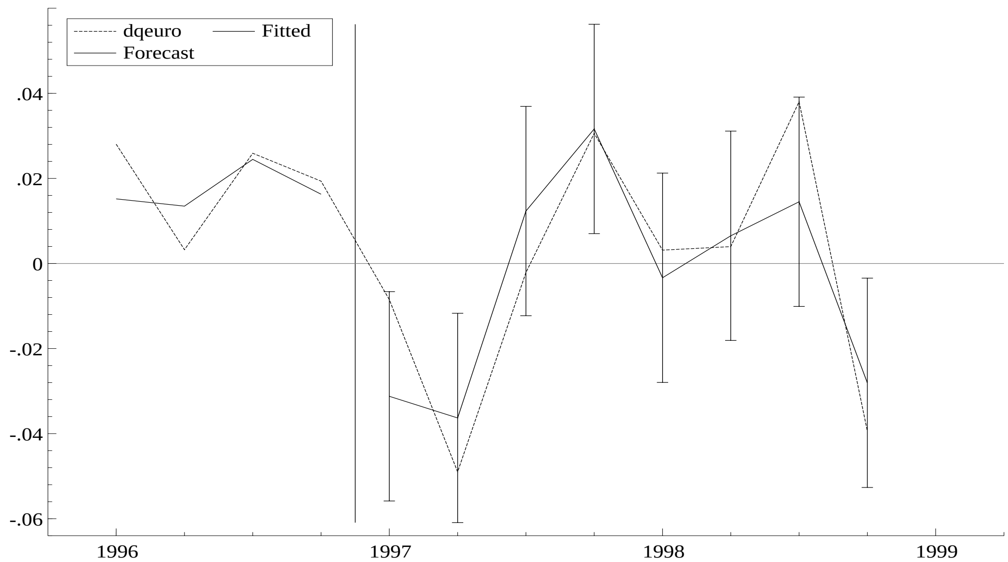


(b) Stability of each of the beta coefficients

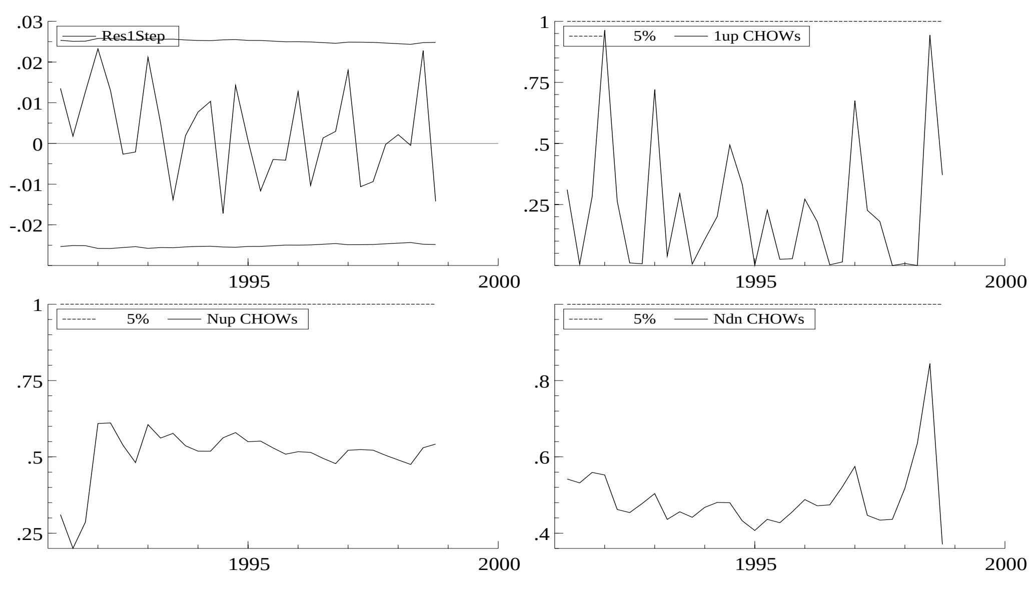


(c) Stability of the adjustment coefficient

Figure 2: Dynamic forecast and recursive estimation



(a) Dynamic forecast



(b) Recursive estimation

the adjustment starts in the same quarter where the shocks have occurred.

We can again compare the error correction model of the aggregate European variables with the results for the panel. As in the time series case, the contemporaneous effects coming from the productivity differential are very important and of the same sign (with a t-statistic of -17.56), but the rest of the variables are not significant. Concerning the error correction coefficient, its magnitude is smaller in the panel (-0.134).

Although there is no consensus in the profession on a particular model specification of exchange rate equations inspired by the New Open Macroeconomics literature (Sarno, 2002), the results obtained in this paper are compatible with these models. In particular, according to Lane (2002), net foreign assets positions are an important form of international macroeconomic interdependence. The influence of net foreign asset positions on the values of the real exchange rate has also been studied recently in Cavallo and Ghironi (2002) and Lane and Milesi-Ferretti (2002). In this paper, we have used the net foreign asset dataset constructed in Lane and Milesi-Ferretti (2001), that is the "adjusted cumulative current account", and the results obtained are also compatible with the most recent empirical literature as well as the previous empirical work[15].

To complete our analysis we check the predictive ability of the Euro-area model. Table 7 presents ex-post and ex-ante forecasting results. Eight observations (two years) are left out for forecasting purposes. From the 1-step static forecast analysis, our model appears to deliver sensible and stable forecasts. The estimates for the dynamic forecast are carried out recursively: the estimation period is successively extended quarter by quarter so that the real exchange rate is forecasted for up to eight quarters into the future. Figure 4 in panel *(a)* shows graphically the predictive performance of our model. This graph plots the dynamic forecast for the period 1997(1) to 1998(4) estimated by full information maximum likelihood. The forecasts lie within the 95 per cent confidence interval, shown by the vertical error bars of plus-or-minus twice the forecast's standard error. Moreover, the fit of the model is good and there are no large departures from the actual values.

Finally, the forecast quality of our model is also assessed by comparing its forecast accuracy with a random walk model for the real exchange rate. For this purpose we obtain the ratio between the root mean squared error (RMSE) corresponding to our VECM relative to the random walk. If the VECM presents a better predictive performance, that is, lower RMSE, this ratio will be below 1. In addition, following Diebold (1998) we carried out a formal test to gain insight into whether the random walk model can generate significantly better forecasts from a statistical point of view. Thus, rejection of the null for this test implies that the random walk model does not provide significantly better forecasts than our VECM. Table 4 presents the ratio of the two RMSE for a forecast horizon up to eight quarters as well as the significance level for the Diebold test statistic, which is indicated by asterisks in the third column. According to these results, the VECM outperforms the random walk model even in the shorter horizons. These can be seen from RMSE ratios,

[15]We should note that the real exchange rate is defined in our paper in the opposite way. More precisely, an increase in the real exchange rate corresponds to a real depreciation.

which are well below 1. Moreover, the predictive performance of our model is statistically shown, rejecting for all the forecast horizons the superiority of the random walk model with a probability as low as 1%.

Table 7

Static and dynamic forecasting

1-step (ex-post) forecast analysis: 1997(1) to 1998Q(4).		
Parameter constancy		
ξ_1	$\chi^2(8) = 10.679\,[0.2205]$	$F(8,73) = 1.3349\,[0.2402]$
ξ_2	$\chi^2(8) = 8.9096\,[0.3500]$	$F(8,73) = 1.1137\,[0.3643]$
ξ_3	$\chi^2(8) = 9.5206\,[0.3003]$	$F(8,73) = 1.1901\,[0.3169]$
Forecast tests: $\chi^2(1)$		
	using ξ_1	using ξ_2
1997(1)	3.4134 [0.0647]	2.6766 [0.1018]
1997(2)	1.0618 [0.3028]	0.8527 [0.3558]
1997(3)	1.3785 [0.2404]	1.0663 [0.3018]
1997(4)	0.0069 [0.9337]	0.0062 [0.9369]
1998(1)	0.2791 [0.5973]	0.2503 [0.6168]
1998(2)	0.0428 [0.8361]	0.0380 [0.8454]
1998(3)	3.6488 [0.0561]	3.2989 [0.0693]
1998(4)	0.8479 [0.3571]	0.7203 [0.3960]

Forecast quality: 1997(1) to 1998Q(4).

Forecast Horizon	RMSE (ratio)	Signif.
1997(1)	0.2509	
1997(2)	0.2176	
1997(3)	0.1887	
1997(4)	0.1821	$***$
1998(1)	0.1716	$***$
1998(2)	0.1676	$***$
1998(3)	0.1665	$***$
1998(4)	0.1728	$***$

Note: ξ_1, ξ_2 and ξ_3 are indices of numerical parameter constancy. The former ignores both parameter uncertainty and intercorrelation between forecasts errors at different time periods. ξ_2 is similar to ξ_1 but takes parameter uncertainty into account. ξ_3 takes both parameter uncertainty and intercorrelations between forecasts errors into account. *Forecats test* are the individual test statistics underlying ξ_1 and ξ_2. $***$ stands for 1% error probability.

5 Conclusions.

In this paper we have applied two different but complementary techniques and approaches to the study of the evolution of the dollar real exchange rate in relation with the Euro-area currencies. First, using panel techniques, we study the long-run relationship between the bilateral real exchange rate of the Dollar versus four European countries, Canada and Japan. Second, in a time series framework, we use Euro-area aggregate or "synthetic" variables to study the behavior of the dollar/euro real exchange rate. Our purpose has been to compare the results obtained from the two approaches. Given that the lack of heterogeneity is one of the main criticisms that are commonly associated with the aggregate analysis, with the panel analysis we have allowed for an individual country study. The similarity of the results obtained using the two methods adds robustness to the Euro-area measures. This fact is a distinctive feature of this work compared to previous papers dealing with the real exchange rate of the euro.

We will maintain the above distinction to summarize the most important empirical results. First, concerning the dynamic panel analysis, we use the Pesaran et al. (1999) methodology, that allows for short-run heterogeneity for the individual members of the panel and a formal test of homogeneity in the long-run parameters. We find that both supply and demand-side factors should be accounted for to explain the bilateral real exchange rate of the US dollar. In particular, the estimated error correction models support a specification including relative productivity, the real interest rate differential, the difference in public expenditure and the relative net foreign asset position. Moreover, this type of relation holds not only for the euro-countries but also for the whole group and for the rest of the countries.

The same long-run specification is identified using the Johansen technique in the time series context. Therefore, even if a larger degree of heterogeneity is allowed in the panel and even if we are using different estimation techniques, the results appear to be almost identical. In addition, in the aggregate time series empirical model, the cointegration vector passes all the applied stability tests. Last, the estimated VECM presents a remarkable predictive performance and a better forecasting quality than the random walk both in the short and the medium term.

According to the long-run results, the dollar-euro exchange rate will depreciate when the American fiscal policy is more expansionary than the European one. In contrast, positive productivity and real interest rate differentials, together with the accumulation of net foreign assets will appreciate the currency. Although the work on New Open Macroeconomics has been mainly theoretical and there are no clear empirical exchange rate equations to test, our results are seemingly compatible with it.

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A Pesaran, Shin and Smith (1999) Pooled Mean Group Estimation of Dynamic Heterogeneous Panels.

This approach combines two procedures that are commonly used in panels. First, the Mean Group *(MG)* estimator: separate equations are estimated for each group and then the Mean Group estimator is computed giving consistent estimates of the average of the parameters. However, this estimator does not take account of the fact that some parameters may be the same across groups. Secondly, the traditional pooling estimators (such as the fixed and random effects estimators) allow the intercepts to differ across groups whereas all the other coefficients and the variances are constrained to be the same.

The Pooled Mean Group *(PMG)* estimator involves both pooling and averaging. This estimator allows the intercepts, short-run coefficients and error variances to differ freely across groups, but the long-run coefficients are constrained to be the same. In the context of integration processes, such as in the case of the Euro countries, it seems reasonable to impose equality in the long-run parameters (or in the majority of them) but allowing the short-run slope coefficients and the dynamic specification (i.e. the number of lags included) to differ across groups.

Let us assume that we have data on time periods, $t = 1, 2, ..., T$, and groups, $i = 1, 2, ..., N$.

Pesaran et al. (1999) present the following Autoregressive Distributed Lag Model, $ARDL(p, q, q, ..., q)$:

$$y_{it} = \sum_{j=1}^{p} \lambda_{ij} y_{i,t-j} + \sum_{j=0}^{q} \delta_{ij}' \mathbf{x}_{i,t-j} + \mu_i + \varepsilon_{it}, \tag{7}$$

We work with the following error correction re-parametrization of the system:

$$\Delta y_{it} = \phi_i y_{it} + \beta_i' \mathbf{x}_{it} + \sum_{j=1}^{p-1} \lambda_{ij}^* \Delta y_{it-j} + \sum_{j=0}^{q-1} \delta_{ij}^{*'} \Delta x_{it-j} + \mu_i + \varepsilon_{it} \tag{8}$$

where y_{it} is the dependent variable, $\phi_i = -(1 - \sum_{j=1}^{p} \lambda_{ij})$, and $\beta_i = \sum_{j=0}^{q} \delta_{ij}$. In addition, $\lambda_{ij}^* = -\sum_{m=j+1}^{p} \lambda_{ij},\ j = 1, 2, ..., p-1$ and $\delta_{ij}^* = -\sum_{m=j+1}^{q} \delta_{im}$, $j = 1, 2, ..., q-1$. $\mathbf{x}_{it}(k \times 1)$ is the vector of explanatory variables for group i, μ_i represent the fixed effects, the coefficients of the lagged dependent variable, λ, are scalars, and those of the explanatory variables, δ, are $(k \times 1)$ coefficient vectors. It should be noted that, although for convenience the lags p and q were set equal across groups, this is not necessary. In addition, T can also differ. A very interesting feature of this methodology is that some of the long-run parameters can be also unconstrained, so that they may be different for each group. This possibility can be tested using LR-type tests.

The disturbances ε_i, are independetly distributed across i and t, with zero means, variances $\sigma_i^2 > 0$, and finite fourth order moments. They are also distributed independently of the regressors, $\mathbf{x}_{it.}$ The assumption that they are independent also across time is not very restrictive, and can be satisfied commonly by increasing the distributed lag orders on y_{it} and $\mathbf{x}_{it}$. The independence of the disturbances and the regressors is needed for the consistent estimation of the short-run coefficients, but Pesaran (1997) shows that it is relatively straightforward to allow for the possible dependence of $\mathbf{x}_{it}$ and ε_{it} when estimating the long-run coefficients, so long as $\mathbf{x}_{it}$ have finite-order autoregressive representations. Thus, for the $ARDL$ model in (7) to be stable the roots of $\sum_{j=1}^{p} \lambda_{ij} z^j = 1$, $i = 1, 2, ..., N$, should lie outside the unit circle.

This assumption ensures that $\phi_i < 0$, so that there exists a long-run relationship between y_{it} and $\mathbf{x}_{it}$ such as

$$y_{it} = -(\beta_i'/\phi_i)\mathbf{x}_{it} + \eta_{it} \tag{9}$$

for each $i = 1, 2, ..., N$, where η_{it} is a stationary process.

Pesaran et al. (1999) also assume the long-run coefficients on $\mathbf{X}_i$[16], defined by $\theta_i = -\beta_i/\phi_i$, to be the same across the groups, so that $\theta_i = \theta$, $i = 1, 2, ..., N$. They write the error correction model representation more compactly as:

$$\Delta y_i = \phi_i \xi_i(\theta) + \mathbf{W}_i \kappa_i + \varepsilon_i,\ \ i = 1, 2, ..., N \tag{10}$$

where

$$\xi_i(\theta) = \mathbf{y}_{i,-1} - \mathbf{X}_i\theta,\ i = 1, 2, ..., N \tag{11}$$

is the error correction component, $\mathbf{W}_i = (\Delta\mathbf{y}_{i,-1}, ..., \Delta\mathbf{y}_{i,-p+1}, \Delta\mathbf{X}_i, \Delta\mathbf{X}_{i,-1}, ..., \Delta\mathbf{X}_{i,-q+1}, \iota)$, and $\kappa_i = (\lambda_{i1}^*, ..., \lambda_{i,p-1}^*, \delta_{i0}^{*'}, \delta_{i1}^{*'}, ..., \delta_{i,q-1}^{*'}, \mu_i)'$. It should be noted that the group specific equations in the panel (10) are non-linear in ϕ_i and θ, and since θ is common across groups the panel is subject to cross-equation parameter restrictions. They also allow the error variances $Var(\varepsilon_{it}) = \sigma_i^2$ to differ across groups.

They then estimate the panel adopting a likelihood approach. These ML estimators are called "pooled mean group" *(PMG)* estimators due to both the pooling implied by the homogeneity restrictions on the long-run coefficients and the averaging across groups used to obtain means of the estimated error-correction coefficients and the other short-run parameters of the model.

[16] When stacking the time-series observations for each group.

B Data sources.

The data is quarterly and covers the period 1970:Q1 to 1998:Q4. We consider the following European countries: France, Germany, Italy, Spain and the United Kingdom. In addition, the United States are the home country and Canada and Japan are also analyzed. The data has been obtained from the magnetic tapes of the International Monetary Fund International Financial Statistics (IFS) with the exception of employment and oil balances data which have been obtained from the International Sectoral Database (OECD). The net foreign assets data has been taken from Lane and Milesi-Ferretti (2001), L-M hereafter. The nominal exchange rate for the Euro relative to the USD has been taken from the database for European variables of the Banco Bilbao Vizcaya Argentaria (BBVA).

The panel data has been constructed as follows:

rerdol$_{it}$: bilateral real exchange rate of the USD relative to the other currencies considered. The nominal exchange rate, s_t, has been defined as currency units of USD to purchase a unit of currency j.

$$rerdol_t = \log\left(\frac{p_t^j}{s_t \times p_t^{usa}}\right)$$

where p_t^{USA} and p_t^j are respectively the CPI for the Unites States and the foreign country.

Source: IFS.

drr$_{it}$: real interest rate differential. The nominal interest rates are *call money rates* as defined by the IMF. In order to obtain the real variables, the expected inflation rate is the smoothed variable based on CPI indices using the Hodrick and Prescott filter.

$$\pi_t = \frac{p_t - p_{t-1}}{p_{t-1}} \times 100$$

$$\pi_t^e = \pi_t - \pi_t^t$$

$$rr_t = r_t - \pi_t^e$$

$$drre_t = rr_t^{USA} - rr_t^j$$

where π_t^e is expected inflation filtered using the HP filter; π_t^t is the transitory component of inflation; rr_t^{usa} is the American real interest rate and rr_t^j the foreign rate.

Source: IFS.

dpro$_{it}$: apparent productivity differential in labor:

$$dpro_t = pro_t^{USA} - pro_t^j$$

where pro_t^{USA} and pro_t^j are respectively the American and the foreign apparent labor productivity and are calculated as:

$$pro_t^j = \log\left(\frac{gdp_t^j}{employment_t^j}\right) \times \frac{1}{s_t}$$

and:

$$pro_t^{USA} = \log\left(\frac{gdp_t^{USA}}{employment_t^{USA}}\right)$$

being the exchange rate

Source: IFS and OCDE.

dpex$_{it}$: public expenditure differential calculated as

$$dpex_t = pex_t^{USA} - pex_t^j$$

where pex_t^{USA} and pex_t^j are respectively the American and the foreign government spending. The government spending is calculated relative to GDP:

$$pex_t = \frac{pexn_t}{gdpn_t} \times 100$$

where $pexpn_t$ is nominal public expenditure.

Source: IFS.

dnfa$_{it}$: net foreign assets differential:

$$dnfa_t = rnfa_t^{USA} - rnfa_t^j$$

where $rnfa_t^{USA}$ and $rnfa_t^j$ stands respectively for the American and the foreign's net foreign asset position relative to the GDP in USD:

$$rnfa_t^j = \frac{nfa_t^j}{gdp_t^j \times \frac{1}{s_t}}$$

and

$$rnfa_t^{USA} = \frac{nfa_t^{USA}}{gdp_t^{USA}}$$

Source: L-M.

oildep$_{it}$: relative dependence of petroleum:

$$oildep_t = \frac{bal_t^j}{bal_t^{USA}} \times \frac{brent\ price}{cpi_t^{USA}} \times 100$$

where bal_t^{USA} and bal_t^j are measures of energetic dependence for USA and the foreign country respectively and have been obtained as:

$$bal_t = \frac{Net\ oil\ imports}{gdpn_t}$$

Source: IFS and OCDE.

For the time series analysis, differentials are no longer calculated for USA relative to each individual country but relative to a representative European variable. The latter is obtained as the weighted average of the corresponding national values already used in the panel analysis. The weights are the share of national GDP relative to the GDP for the Euro-area. The bilateral real exchange rate (q_t) of the USD relative to the Euro has been obtained as in the panel where now s_t is defined as units of Euro require to purchase a unit of USD. The source for s_t are BBVA (from 1970:Q1 to 1997:Q4) and IFS for the rest of the sample.

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