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The Euro-Dollar exchange rate: 
Is it fundamental?*

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Abstract
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 the panel cointegration methodology for both homogeneous and heterogeneous panels, we study the long-run relationship between the bilateral real exchange rate of the dollar versus 10 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. The selected specification obtained using the panel techniques is an eclectic one, that supports the Meese and Rogoff (1988) real interest rate differential model augmented with two supply-side variables: the real oil dependence and the relative productivity in the non-tradables. The Euro-area variables support this type of results, although an additional determinant from the demand-side should be added (the relative public expenditure) whereas the real oil variable would be only significant in the short-run.

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

J.E.L. codes: C33, F31.

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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 “strengtheness” 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 propose a panel cointegration estimation both in an homogenous and heterogenous setup. This methodology allow 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 permit 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 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 and section 4 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\textsuperscript{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)\textsuperscript{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 Taugas (1999) find that the deviations from PPP can be explained largely by productivity differentials and interest rate differentials in an error correction model. Additionally, Clotsterman 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

\textsuperscript{2}For a complete overview of different empirical approaches, see Williamson (1994) or more recently, MacDonald (2000).

\textsuperscript{3}For the sake of simplicity we are omitting the NATREX and the PEER approaches. 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) model 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, this variable, $q_t$, as $q_t = \varepsilon_t - p_t + p_t^*$, where $\varepsilon_t$ 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, $\bar{q}_t$, 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:

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

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 $\bar{q}_t$ 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) can-
not 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 sig-
nificant long-run relationships and plausible point estimates using panel tests,
they conclude that the failure of previous researches may be due to the esti-
mation method used rather than to any theoretical deficiency.

In a second group of papers, the assumption that the expected real ex-
change 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 cur-
rent 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 \( \alpha \) and \( \alpha^* \) 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. Two separate groups of effects can be also distinguished:

   - 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. If the productivity improvement occurs in the non-traded goods sector, the real exchange rate depreciates.

   - A second effect can be associated to the presence of a non-traded good bias in demand, that would increase the relative price of these goods and appreciate the currency. The large share of public sector expenditure directed to non-traded goods also reinforces this effect.
2. Non-constancy of the real exchange rate for traded goods (the first term in (2)). Two additional factors may introduce variability in $q_t^*$:

- International differences in savings and investment, reflected in the current account of the economies and measured as the net foreign asset positions.
- Changes in the real price of oil, that tends to depreciate the currencies of the net oil importers or, in general, the relatively the currencies of the more 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:

$$
q_t = -\varphi(R_t - R_t^*) + \hat{q}_t = 
= f((R_t - R_t^*), (a_{TL} - a_{TL}^*), (a_{NL} - a_{NL}^*), (g_t - g_t^*), o_t, n f a_t)
$$

(3)

where $(a_{TL} - a_{TL}^*)$ is the difference between the domestic and foreign economies in traded goods productivity, $(a_{NL} - a_{NL}^*NTL)$ is its equivalent for the non-traded goods sector, $(g_t - g_t^*)$ is the public expenditure differential, that can be associated to the demand bias in non-traded goods, $o_t$ is the real oil price and $n f a_t$ is the net foreign asset position of the economy$^4$.

In contrast to the more traditional monetary approach, Rogoff (1992), Obstfeld (1993) and Asea and Mendoza (1994) emphasize the role of fiscal policy and other real variables (such as productivity shocks, for example) in real exchange rate models. Rogoff (1992) develops an intertemporal model for exchange rate determination with factors that are not perfectly mobile across sectors. In contrast to the Balassa-Samuelson approach, that considers that the characteristics of the individual's utility function and the level of government consumption spending have no effect, he stresses the role of demand factors on the long-run behavior of real exchange rates. He considers that in open capital markets, and under imperfect factor mobility across sectors, agents can smooth their consumption of tradables in the face of transitory traded goods productivity shocks. They cannot, however, smooth non-traded goods productivity shocks, normally caused by changes in government spending, although if they are small, traded-goods consumption smoothing will lead to also smoothing the intra-temporal price of traded and non-traded goods. Thus, according to this model, productivity shocks as well as changes in government spending will affect the real exchange rate. However, he considers that only the second type of shocks may have permanent effects. The critical issue that leads to predictions that do not support the traditional Balassa-Samuelson effect is the

$^4$Depending on the adopted theoretical approach, the sign associated with $n f a_t$ can be positive (portfolio balance models) or negative (monetary models).
existence of imperfections in factor mobility across sectors. He finds results favorable to this approach for the case of Japan versus the German Mark and the dollar. In addition to the productivity and fiscal variables, he includes the real oil price as a major terms of trade shock. The empirical specification takes the form:

\[ q_t = f((a_{\tau t} - a_{\tau t}^w), (a_{Nt} - a_{Nt}^w), (g_t - g_t^w), \omega_{il}t) \]  

(4)

Thus, although the assumptions associated with the two type of models are substantially different, equation (3) nests (4) and permits to develop a testing strategy.

4 Empirical results.

4.1 Panel cointegration 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 thirteen: the US as the domestic country, Japan, Canada and 10 European countries (those with information available for the sample period and variables of interest). Consequently, this first part of the analysis is not exactly a model for the dollar versus the Euro-area. We have chosen to include countries (such the UK, Denmark and Finland) not included in EMU, as well as Canada and Japan, in order to capture the behavior of the most important world currencies. The heterogeneous estimation will allow for a more individual study of each country, paying special attention to those in the Euro-zone. The breakdown of productivity into traded goods and non-traded goods sectors limits us to the data set from the OECD, and restricts our sample to 1972-1992.

Some of the variables traditionally considered in the empirical real exchange rate models (see MacDonald, 1998), such as the net foreign asset position of the country, the relative public expenditure of the domestic and foreign country and the relative productivity in the traded goods sectors turned out not to be significant in any of the specifications tried following the theoretical models. Consequently, the results presented in this section both concerning the order of integration of the variables and the long-run relationships only include the

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5Hamilton (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.

6A detailed description of the variables can be found in Appendix A.

7See also Appendix A for a description of the way we have calculated these variables, excluded from the subsequent empirical analysis in this section.
variables that proved to play a role in the panel long-term relations\(^8\). In addition, 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 present four specifications where Model 1 nests the other three models.

Thus, Model 1, the most general specification, will take the form:

\[
\text{rerdol}_it = f(\text{pront}_{it}, \text{drre}_it, \text{oildep}_it)
\]

where \(\text{rerdol}_it\) 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; \(\text{pront}_{it}\) is the relative productivity of the US versus each of the other countries in the non-traded goods sector, so that (as explained in the theoretical section) an increase in the value of this variable tends to depreciate the currency; \(\text{drre}_it\) is the real interest rate differential between the US and the other countries analyzed: an increase in this differential appreciates the currency; finally, \(\text{oildep}_it\) 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 dependency of the foreign countries is increasing.

The specifications proposed are the following:

**Model 1:** \(\text{rerdol}_it = \alpha_i + \beta_{1i}\text{pront}_{it} + \beta_{2i}\text{drre}_it + \beta_{3i}\text{oildep}_it\)

**Model 2:** \(\text{rerdol}_it = \alpha_i + \beta_{1i}\text{pront}_{it} + \beta_{2i}\text{drre}_it\)

**Model 3:** \(\text{rerdol}_it = \alpha_i + \beta_{1i}\text{drre}_it + \beta_{2i}\text{oildep}_it\)

**Model 4:** \(\text{rerdol}_it = \alpha_i + \beta_{1i}\text{pront}_{it} + \beta_{2i}\text{oildep}_it\)

The three first models can be considered different versions of the Meese and Rogoff (1988) real interest rate differential model, augmented with different variables that may account for the determinants of the long-run equilibrium real exchange rate from the supply-side of the economy. The fourth model is a partial version of the Balassa-Samuelson approach, including the real oil price as an additional variable, representing the most important terms of trade shock, real oil prices.

### 4.1.1 Order of integration of the variables.

Bearing all these considerations in mind, we should start the analysis by the study of the order of integration of the variables. Several tests to test for unit roots 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, in the long-run heterogeneous relationships we will be using a LM test that maintains the

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\(^8\)The stationarity analysis and the results from the alternative specifications are available from the authors upon request.

\(^9\)Finally published as Levin, Lin and Chu (2002).
null hypothesis of cointegration due to its better power. Equivalently, 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:

\[
H_0 : \sigma^2_{u1} = ... = \sigma^2_{uN} = 0
\]  

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

4.1.2 Cointegration relationships: homogeneous and heterogeneous panels.

When studying the long-run relationships in panels, two different hypotheses can be maintained about the estimated slope parameters. Let us consider the following empirical model, similar to the one to be estimated:

\[
y_{it} = \alpha_i + \beta_{1i}x_{1it} + \beta_{2i}x_{2it} + e_{it}
\]

where \( x_{1it} \) and \( x_{2it} \) are the two explanatory variables and \( e_{it} \) are the errors. The first hypothesis that can be tested is that of homogeneity, that is, all the cross-section share the same coefficient value for the same variable: \( \beta_{11} = \beta_{12} = ... = \beta_1 \) and \( \beta_{21} = \beta_{22} = ... = \beta_2 \). In contrast, in the heterogeneous panel, each cross-section estimate is allowed to differ from the others. No restriction is imposed concerning their value.

In the rest of this section we briefly summarize the tests, for both homogeneous and heterogeneous panels, used to assess the existence of long-run relations linking the variables in the model.

Panel cointegration test results: homogeneous panel. Concerning the long-run analysis, we will first apply the panel cointegration tests and estimation procedures for homogeneous panels. In this framework, that means that we allow for fixed specific effects for each country but restrict the slope
coefficients to be equal for all the members of the panel. Kao (1999) proposed $DF$-type panel cointegration tests based on the $OLS$ residuals from the homogeneous panel regression.

The $DF$ test from Kao (1999) follows the model:

$$y_{it} = \alpha_i + \beta x_{it} + \epsilon_{it}, \quad i = 1, \ldots, N, \quad t = 1, \ldots, T$$

(6)

where both $y_{it}$ and $x_{it}$ are random walks. Thus, under the null hypothesis of no cointegration, the residual series $\epsilon_{it}$ should be non-stationary.

The limiting distributions are asymptotically normally distributed at mean zero. However, they contain nuisance parameters because of possible long-run weak exogeneity and serial correlation in the errors. Thus, it would be necessary to have good estimates of the long-run parameters. Kao constructs new statistics whose limiting distributions are $N(0, 1)$ and do not depend on the nuisance parameters, that are called $DF_{\rho}^*$ and $DF_{t}^*$. Alternatively, he defines a bias-corrected serial correlation coefficient estimate and, consequently, the bias-corrected test statistics and calls them $DF_{\rho}$ and $DF_{t}$. According to Baltagi and Kao (2000), the main difference between the two groups of tests is that whereas the $DF_{\rho}$ and $DF_{t}$ tests are based on the strong exogeneity of the regressors and errors, the $DF_{\rho}^*$ and $DF_{t}^*$ are more adequate for cointegration with endogenous relationships between regressors and errors. Finally, he also proposes an $ADF$ version of the test.

In table 2 we present the results of the different versions of the $DF$ test and the $ADF$ tests described above. For the four models considered and the different tests it is possible to reject the null hypothesis of non-cointegration at 1% significance levels. Thus, this information will not allow us to appropriately compare the models or extract any conclusion on the adequacy of any of them.

Table 3 offers some more information related to the homogeneous estimation results: in the first panel we present the $OLS$ estimates, whereas panel B displays the bias corrected estimates. In the case of Model 1 (M1 hereafter), the $OLS$ estimates have the signs predicted by the theory and are all significant. However, the bias-corrected estimate (although no very different in magnitude) is non-significant in the case of the relative productivity variable. The same happens in model M2, whereas in M4 the productivity variable is also non-significant in the $OLS$ estimation. However, we maintain this variable in the models due to its importance in the heterogeneous estimation presented in the next subsection\(^{10}\). An alternative specification, if we dropped $\text{prout}_{it}$ would be the one presented in M3. In this case, both the $OLS$ and bias corrected $OLS$ estimates are significant and the signs are correct. This alternative would include both demand and supply-side variables, supporting the Meese and Rogoff (1988) real interest differential model, although augmented by the oil dependency variable.

A word of caution should be given before progressing: due to the heterogeneity of the countries involved in the analysis, the homogeneous analysis may

\(^{10}\)In addition, Kao (1999) recommends using other more efficient estimation techniques, unavailable in this stage of the study due to lack of degrees of freedom.
introduce too strong restrictions in the parameters, not necessarily supported by the data, so that the heterogeneous analysis should be carried out. Non-significant parameters may be the consequence of large discrepancies between the different countries in the cross-sections.

**Panel cointegration tests: heterogeneous panel.** In this section, the parameters are allowed to differ across the cross-sections, so that we will analyze the so-called heterogeneous panel. We will apply McCoskey and Kao (1998) residual-based panel test of the null hypothesis of cointegration. This test is an extension of the Lagrange Multiplier (LM) test and the Locally Best invariant (LBI) test for a MA unit root in the time series literature. They follow a similar approach to the one that has been proposed for the time series case by Harris and Inder (1994) and Shin (1994). For this test it is necessary to use an efficient estimation technique of the cointegrated variables, such as the Dynamic OLS that is the one recommended by McCoskey and Kao (1998).

The model, that allows for varying slopes and intercepts:

$$y_{it} = \alpha_{it} + x'_{it} \beta_i + e_{it}, \quad i = 1, ..., N, \quad t = 1, ..., T, \quad (7)$$

with $x_{it} = x_{it-1} + \varepsilon_{it}$, $e_{it} = \gamma_{it} + u_{it}$, and $\gamma_{it} = \gamma_{it-1} + \theta u_{it}$, where $u_{it} \sim iid N(0, \sigma_u^2)$.

By backward substitution:

$$y_{it} = \alpha_{it} + x'_{it} \beta_i + \theta \sum_{j=1}^{t} u_{ij} + u_{it} = \alpha_{it} + x'_{it} \beta_i + e_{it}$$

where $e_{it} = \theta \sum_{j=1}^{t} u_{ij} + u_{it}$.

If we estimate the DOLS of $\beta_i$, we have to run the following regression using $q$ leads and lags of the explanatory variables to correct for correlation and endogeneity effects:

$$y_{it} = \alpha_{it} + x'_{it} \beta_i + \sum_{j=-q}^{q} c_{ij} \Delta x_{it+j} + u_{it}$$

The null hypothesis of cointegration is equivalent to $\theta = 0$. In addition, if $\theta = 0$ then $e_{it} = u_{it}$ and is stationary.

The asymptotic distribution of the test is not dependent on any nuisance parameters.

In this framework it is possible to obtain locally optimal invariant tests for the null hypothesis. The $LBU1$ test statistic proposed by McCoskey and Kao is an $LM$ statistic:

$$LM = \frac{1}{N} \sum_{i=1}^{N} \frac{1}{T^2} \sum_{t=1}^{T} S_{i,t}^{2}$$

where $S_{i,t}^{2}$ is the partial sum process of the residuals, and $s^{2} = \frac{1}{N} \sum_{i=1}^{N} \sum_{t=1}^{T} e_{it}^{2}$. 

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The application of McCoskey and Kao LM test to the four models specified above is presented in Table 4, both for the individual countries and for the whole panel. Only in the case of Italy in M2 we can reject the null hypothesis of cointegration, whereas the panel tests support the existence of cointegration in the four cases.

In Tables 5 and 6 we present the actual estimates of the slope parameters in the four models. We have considered significant coefficients those with t-statistics around the 1.60 value.

1. From these results, the first pattern that we can extract is that the real oil dependency seems to be the variable that appears to be significant for most of the countries in the three specifications where it is present. Nine cases in M4, eight in M1 and five in M3. We partly attribute the significance of this variable to the information it contains so that not only captures the evolution of the real price of oil but also includes the relative efficiency of each country if compared to the US.

2. Second, although the real oil dependency variable is absent from model 2, the other supply variable, \textit{prompt}, is also significant in many cases (seven). Thus, one supply factor or other turns out to be important for the explanation of the real exchange rate behavior.

3. Third, concerning the productivity variable, although it is only significant in three cases with the correct sign in model M1 (and in the case of Sweden only marginally), we have not excluded it from the analysis due to its significance in models M2 and M4 (seven and five non-zero coefficients, respectively) with the correct sign and large in magnitude.

4. Fourth, the real interest rate differential turns out to be significant for five countries in model M3 and six in model M2. However, in M1 there are only three significant cases with correct sign, one of them quite marginal (Finland). A closer look reveals that these countries are Germany and Japan. A plausible explanation of this result can be that monetary policy is transmitted internationally between the key international currencies. In the case of Europe, the Germany could be the channel of transmission towards the rest of the countries in the European Monetary System at that time. It should be noted that, as mentioned previously, in other specifications (models M2 and M3) there are more significant country-cases, although the majority of them did not belong to the EMS at that time or had very recently joined (Finland, Sweden, Spain).

\footnote{In Camarero and Tamarit (2002b) we have also found this variable to be relevant in the explanation of the real exchange rate of the peseta during the period 1970-1997, together with the real interest rate differential.}

\footnote{See in Table 7 the evolution of the variable during the sample period in all the countries considered.}

\footnote{Canada also presents a significant coefficient, but the sign is positive instead of negative.}

\footnote{This would give support to the so-called "German Dominance Hypothesis".}
5. Fifth, the case of Canada deserves special attention. In model M1 there are two reverse signs: the real interest rate differential and the real oil price. It is hard to find an explanation for the first one\(^{15}\) but the second one is consistent in all the estimations and its magnitude important. Canada is the only important commodity exporting country in the sample. Amano and van Norden (1993) have documented also a sign reversal for this variable. In a recent paper, Chen and Rogoff (2002) suggest that accounting for the commodity prices may help to solve part of the exchange rate puzzles in the developed countries that are heavy commodity-exporters.

Thus, from the estimations carried out in the context of panel techniques, we would propose to choose M1\(^{16}\), that supports the Meese and Rogoff's (1988) real interest rate differential, although augmented using supply-side variables that account for two potentially important effects: first, the terms of trade shocks caused by the changes in oil prices; and, second, the traditional Balassa-Samuelson effect, that would associate productivity gains in non-tradables with the depreciation of the real exchange rate through its negative impact on prices of the non-tradables. If this is the case, the non-significance of the relative productivity of traded goods would mean that the real exchange rate is mainly affected by changes in the internal relative prices, rather than by the relative price of traded goods in the domestic and foreign country. This pattern could be plausible when studying developed countries, as in this case: with similar technologies and the progressive openness in trade, the most important differences may reside in the non-traded goods, those sectors not affected by international competition\(^{17}\).

### 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 either the monetary model of Meese and Rogoff (1988), nor in the Rogoff (1992) intertemporal approach or in the traditional Balassa-Samuelson framework alone. In this section, we are interested in the aggregate behavior of the Euro-area variables. This analysis is a complement to the previous one. Our purpose is to compare the results using “true” variables for the European countries individually obtained in the heterogeneous part of the panel (although for a shorter time-period), with those using “synthetic” Euro-area variables. 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 method-

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\(^{15}\)In M1, although only marginally significant, the sign is correct.

\(^{16}\)In Camarero and Tamarit (2002a), using this methodology for the real exchange rate of Spanish peseta and for the period 1972-1992, we find a similar result, although including both traded and non-traded goods productivities as well as the real interest rate differential.

\(^{17}\)A very interesting and broad study of the relationship between the real exchange rate, the Balassa-Samuelson effect, and different measures of productivity and competitiveness can be found in MacDonald and Ricci (2002).
ologies do not show important discrepancies, we could feel more confident when using the aggregate series for inference and policy analysis.

It should be noted that, in order to compute aggregate variables\textsuperscript{18} we have used a different definition for the productivity and the real oil dependency variable. The first one, obtained from the Bureau of Labor Statistics, has been homogeneously calculated for the US and Europe, but does not allow for the breakdown between traded and non-traded goods\textsuperscript{19}. In the case of the real oil dependency variable, as the individual series had been calculated using net imports of petroleum as a percentage of GDP and in order to avoid problems related to aggregating trade balances, we have instead used a measure of real oil prices corrected by the relative oil consumption\textsuperscript{20}.

For this part of the analysis we use the Johansen (1995) methodology for the estimation and identification of cointegrated systems. The variables in the system are: the short-run differential of interest rates ($drr_t$) between United States and Europe, a measure of the energetic dependence of Europe relative to the United States ($dep_t$), the differential of public expenditure over GDP ($pex_t$) and the differential of the productivity measured as apparent labor productivity between United States and Europe. The sample period is quarterly running from 1970:Q1 to 1998:Q4.

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 $I(1)\textsuperscript{21}$. Next, table \ref{tab:1} shows the $\lambda_{\text{max}}$ and Trace test statistics for the determination of number of cointegration relationships. 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), for the existence of cointegration. According to these results, the Trace test statistic allows us to reject the absence of cointegration at a 5\% significance level, suggesting the choice of $r = 1$. In contrast, the $\lambda_{\text{max}}$ test statistic rejects such possibility so that according to this test the right choice is $r = 2$. To complement this evidence we have also analyzed the roots of the companion matrix: five of them are close to unity, implying that this is the number of common stochastic trends. Moreover, when $r = 1$ the largest roots are removed, leaving no near unit root in the model, suggesting therefore that this is the appropriate choice for $r$. From this evidence, the most feasible outcome is the existence of one cointegration vector, that is, $p - r = 2$, where $p$ is the number of common stochastic trends.

The recursive analysis of the system also provides useful information regarding the existence of cointegration: the recursive time path of the Trace statistics and the behavior of the estimated cointegration relations (where only the first vector seems to be stationary) also supports this conclusion.

The above results jointly considered permit us to accept the existence of

\textsuperscript{18}See Appendix A for a detailed explanation.
\textsuperscript{19}An alternative, that will be the considered in future research, is to use the prices in traded and non-traded goods sectors as proxy-variables for the Balassa-Samuelson effect.
\textsuperscript{20}See Appendix A for details.
\textsuperscript{21}The results are available upon request.
one cointegration vector.

Then, the cointegration vector is identified imposing the overidentifying restriction that the variable for energetic dependence \((dep)\) is excluded from the long-run: the LR statistic is \(\chi^2(1)=2.04\) with a probability value of 0.15. The resulting cointegration vector takes the form \((t\text{--values in parentheses})\):

\[
q_{t-1} = -0.07 \quad p_{ex_{t-1}} - 0.29 \quad drr_{t-1} + 7.92 \quad pr_{ot_{t-1}}
\]

At this stage of the analysis, if compared with the panel results, two are the main differences: first, the above mentioned long-run exclusion of the oil dependency variable\(^{22}\) and, second, the presence of the public expenditure differential. Coming back to the theoretical aspects described in section 3, the cointegration vector \((9)\) would include the majority of the fundamental factors compatible with the “augmented” monetary approach: the real interest differential is present, together with the relative public expenditure, a second demand-side factor and relative productivity. The two first variables have the correct signs: an increase in the two differentials tends to appreciate the dollar. Concerning the productivity variable, a broad measure of labor productivity, presents a positive sign. This result would be in accordance with those obtained in the previous section, where relative productivity in the tradables sector was insignificant and only differences in non-traded goods productivity affected the real exchange rate. A positive coefficient in this aggregate productivity measure gives support to the hypothesis that internal relative prices are the ones behind the real exchange rate developments.

The relative public expenditure is a variable that was not significant in the panel regressions but appears in the long-run equation in the Euro-aggregate model. A plausible explanation for this outcome can be found in the increase in fiscal-policy coordination that started around 1992 (the last year included in the panel analysis), due to the convergence programs implemented to satisfy the two fiscal Maastricht criteria. Although during this same period the US administration also followed a policy of fiscal consolidation, the process may have been stronger in Europe. In fact, this process can be considered one of the causes explaining the recession that lasted until the mid-nineties.

Then, we formally test for weak exogeneity of the variables in the system. According to our results, both short-term interest rates and public expenditure differentials appear to be weakly exogenous. The joint hypothesis of weak exogeneity and the identifying restrictions on the cointegration space \(\beta\) are clearly accepted: the LR statistic value is \(\chi^2(2)=0.84\) with a probability of 0.36. We present next the error correction model (ECM hereafter) for the real exchange rate equation:\(^{23}\)

\[
\Delta q_t = 0.1423 + 0.248\Delta d_{q_{t-1}} - 0.0095 \Delta dep_{t-1} + 2.770\Delta pr_{ot_{t-2}} - 0.085 \Delta ecmt_{t-1} + \varepsilon_t
\]

\(^{22}\)However, as we will see below, this variable turns out to be significant in the short-run.

\(^{23}\)The ECM for the remainder equations of the system are available upon request.
Misspecification tests\textsuperscript{24}:

- Residual correlation: $\chi^2(4) = 0.21 \quad p-value = 0.99$
- ARCH: $\chi^2(1) = 0.94 \quad p-value = 0.33$
- Normality: $\chi^2(2) = 2.75 \quad p-value = 0.25$
- Functional form: RESET $\chi^2(1) = 0.56 \quad p-value = 0.45$

where $\varepsilon_t$ is a vector of disturbances, $ecm_{t-1}$ is the cointegration vector (1) and the t-values are in parentheses.

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 whether the cointegration space is stable subject to the existence of one cointegration vector. Thus, this test is tantamount to assess the stability of the $\beta$ coefficients of our cointegration relation. We also test for the stability of the loading parameters. If both $\alpha$ and $\beta$ appear to be stable, we can conclude that our error correction model is well specified for the period analyzed.

Figure 1, 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 $\chi^2$-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 is 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. Although stability using the $Z$-representation may apparently be rejected, it is not the case, since this test statistic requires ten years approximately to achieve stability.

In Figure 1, panel (b), we show the time path of the loadings to the cointegration vector. The recursive plot of $\alpha$ lies 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.

As for the real exchange rate ECM, presented in equation (10), 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, three are the variables that appear in the dynamics of the real exchange rate. First, the one-lagged real exchange rate. Second, with four lags, the real oil dependency variable ($dep_t$) that, although excluded from the long-run, has a significant influence on the exchange rate in the short run. Moreover, the negative parameter, as in the panel analysis, is the one expected from the theory. Finally,

\textsuperscript{24}The functional form statistic is Ramsey’s RESET test, that has been computed using the square of the fitted values. These statistics have been obtained using the econometric package MICROFIT (Pesaran and Pesaran, 1997).
the productivity differential measure, lagged two periods, with the same positive sign found in the long-run time series analysis and in the panel section of the paper. It should be also emphasized that the contemporaneous values of the variables are not significant reflecting the sluggishness of the adjustment towards equilibrium.

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 the panel cointegration methodology for both homogeneous and heterogeneous panels, we study the long-run relationship between the bilateral real exchange rate of the dollar versus 10 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 panel cointegration analysis, we find that both supply and demand-side factors should be accounted for to explain the bilateral real exchange rate of the US dollar. The specification including relative productivity in the non-traded goods sectors, the real interest rate differential and the real oil dependency passes all the cointegration tests and the parameters obtained are in line with the theoretical propositions. However, the chosen model is an eclectic one, that supports the Meese and Rogoff’s (1988) real interest rate differential model, although augmented using supply-side variables that account for two potentially important effects: first, the terms of trade shocks caused by the changes in oil prices; and, second, the traditional Balassa-Samuelson effect.

The main empirical findings obtained for the Euro-area variables support those from the panel. However, there are two differences concerning the long-run specification. First, the variable real oil dependency is excluded from the long-run equation but turns out to be significant in the error correction model. Second, the relative public expenditure, that was not significant in the panel, is now included in the long-run aggregate model. The process of fiscal consolidation associated to the fulfillment of the Maastricht criteria may be the most plausible explanation for its aggregated importance. In addition, the cointegration vector passes all the stability tests applied. These results are in line with the recent empirical literature about the euro-dollar exchange rate.
References


A Data sources.

A.1 Panel analysis

The data in this section is annual and covers the period 1972-92. The panel consists of the following 10 European countries: Belgium, Denmark, Finland, France, Germany, Italy, the Netherlands, Spain, Sweden 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). The productivity variables have been obtained from the International Sectoral Database (OECD) and from non-published Bank of Spain data provided by Paco de Castro.

$\text{rerdoll}_{it}$: 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 $j$ to purchase a unit of USD. Source: IFS.

$$\text{rerdoll}_t = \log \left( \frac{p_t}{s_t \times p^{usa}_t} \right)$$

$\text{drre}_{it}$: 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 (source: IFS) using the Hodrick and Prescott filter.

$$\pi_t = \frac{IPC_t - IPC_{t-1}}{IPC_{t-1}} \times 100$$

$$\pi^e_t = \pi_t - \pi^t_t$$

$$rr^t_t = r_t - \pi^e_t$$

$$\text{drrre}_t = rr^{USA}_t - rr^j_t$$

where $\pi^e_t$ is expected inflation filtered using the HP filter; $\pi^t_t$ is the transitory component of inflation; $rr^{USA}_t$ is the American real interest rate and $rr^j_t$ the foreign rate.

$\text{prot}_{it}$: productivity differential in tradables between United States and each of the other countries in the sample. Tradables productivity includes manufactured goods and transport, storage and communication. The source is the ISDB of the OECD and the Bank of Spain.

$\text{pront}_{it}$: productivity differential in non-tradables between United States and each of the other countries in the sample. We consider non-tradables goods all the sectors excluding manufacturing and transport, storage and communication. The variable has been computed for each country as a weighted average where the weights depend on the relative importance of each sector in GDP.
\(d p e x_{it}:\) public expenditure differential. The government spending is calculated relative to GDP:

\[
pe_{xt} = \frac{pex_{Dt}}{gdp_{Dt}} \times 100
\]

where \(pex_{Dt}\) is nominal public expenditure, whereas \(d p e x_{it} = pe_{xt} - pe_{x_{t-1}}\). The sources is IMF.

\(dnfa_{it}:\) relative net foreign assets calculated as:

\[
nf_{At} = \frac{nfa^{US}_{It}}{gdp^{US}_{It}} - \frac{nfa^{j}_{It} \times s_{It}}{gdp^{j}_{It}}
\]

where \(s_{jt}\) is the nominal exchange rate (currency units of country \(j\) per USD), and \(nfa^{US}_{It}\) and \(nfa^{j}_{It}\) refers respectively to the american net foreign asset and the net foreign asset for country \(j\). Source: IFS.

### A.2 Time series analysis

The data in this section is quarterly and covers the period 1970:Q1-1998:Q4. The data has been obtained from the magnetic tape of the International Monetary Fund International Financial Statistics (IFS), from the database for European variables of the Banco Bilbao Vizcaya Argentaria (BBVA) and from the European Central Bank monthly bulletin (ECB). The productivity variables have been obtained from the Bureau of Labor Statistics of the US Department of Labor.

\(q_{it}:\) bilateral real exchange rate of the USD relative to the Euro. This variable is calculated as in the panel where now \(s_{jt}\) is defined as units of Euro require to purchase a unit of USD. Source: BBVA (from 1970:Q1 to 1997:Q4) and IFS for the rest of the sample.

\(drr_{it}:\) real interest rate differential between the american real interest rates and the interest rates for the Euro area. Source: BBVA, IFS and ECB.

\(pex_{it}:\) public expenditure differential between USA and the Euro area. The public expenditure for the USA is calculated as above. To compute the Euro-area public expenditure we first compute the public expenditure over GDP for each of the following individual countries: Finland, France, Germany, Italy and Spain. Then, the variable for the Euro area is calculated as the weighted average of the national values. The weights are the share of national GDP relative to the GDP for the Euro area. Source: IFS.

\(pro_{it}:\) productivity differential between USA and the Euro area. The productivity index for USA is calculated as real GDP per employed person. The productivity index for the Euro area is calculated as the weighted
average of the national values for the following countries: Austria, Belgium, France, Germany and the Netherlands. The weights are the share of national GDP relative to the GDP for the Euro area. The national values are calculated as real GDP per employed person converted to USD using the PPP for 1998. Thus the differential takes the form:

\[ p_{t} = \frac{gd_{t}}{e_{t}} - \left( \sum_{j} \alpha_{j} \frac{gd_{j}}{e_{t}} \right) \]

where \( gd_{t} \) and \( gd_{j} \) are respectively the real GDP for USA and the European countries. \( e \) refers to employment and \( \alpha_{j} \) are the weights. Since only annual data is available, the series are expanded to quarterly data assuming that the processes follow a random walk. Source: Bureau of Labor Statistics.

\textit{dep}_{it}: differential of energetic dependence based on petroleum consumption. For the USA is calculated as:

\[ dep_{t}^{USA} = \frac{\text{Total USA oil consumption in Brent barrels}}{\frac{gd_{t}^{USA}}{gd \text{ deflator}}} \times 100 \]

For the Euro area \( dep \) is calculated as the weighted average of the national values for the following countries: Belgium, Finland, France, Germany, Italy, the Netherlands and Spain. The weights are the share of national GDP relative to the GDP for the Euro area. The national values are calculated as in the case of USA. Then:

\[ dep = dep_{t}^{USA} - dep_{t}^{EURO} \]

Source: BBVA.
B Tables.

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

<table>
<thead>
<tr>
<th>Variables</th>
<th>$\eta_{\mu}$</th>
<th>$\eta_{\gamma}$</th>
</tr>
</thead>
<tbody>
<tr>
<td>rer dol it</td>
<td>3.09**</td>
<td>43.16**</td>
</tr>
<tr>
<td>prontas</td>
<td>16.12**</td>
<td>625.50**</td>
</tr>
<tr>
<td>drrec it</td>
<td>8.81**</td>
<td>89.90**</td>
</tr>
<tr>
<td>oildep it</td>
<td>4.34**</td>
<td>37.07**</td>
</tr>
</tbody>
</table>

Note: The statistic $Z_{\mu}$ does not include a time trend, whereas $Z_{\gamma}$ does, and are normally distributed. The two asterisks denote rejection of the null hypothesis of stationarity at 5%. The number of lags selected is $l = 1$. 
Table 2
Homogeneous Panel Cointegration Tests
Kao (1999) DF and ADF Tests

Model specifications:
Model 1: \( rerdol_{it} = \alpha + \beta_1pront_{it} + \beta_2drre_{it} + \beta_3oildep_{it} \)
Model 2: \( rerdol_{it} = \alpha + \beta_1pront_{it} + \beta_2drre_{it} \)
Model 3: \( rerdol_{it} = \alpha + \beta_1drre_{it} + \beta_2oildep_{it} \)
Model 4: \( rerdol_{it} = \alpha + \beta_1pront_{it} + \beta_2oildep_{it} \)

Sample: 1972-1992

<table>
<thead>
<tr>
<th>Test / Model</th>
<th>M1</th>
<th>M2</th>
<th>M3</th>
<th>M4</th>
</tr>
</thead>
<tbody>
<tr>
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<td>-3.97***</td>
<td>-3.84***</td>
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<td>(0.00)</td>
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<tr>
<td>(DF_t)</td>
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<tr>
<td>(DF^{*}_\rho)</td>
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<td>(ADF)</td>
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</table>

Note: the three asterisks denote rejection of the null hypothesis of non-cointegration at 1%, whereas probabilities are in parentheses. The tests statistics are distributed as \(N(0,1)\).
### Table 3A

**Homogeneous panel OLS cointegration estimates**

Dependent variable: \( \text{rerdol}_{it} \)

<table>
<thead>
<tr>
<th>Variables / Model</th>
<th>M1</th>
<th>M2</th>
<th>M3</th>
<th>M4</th>
</tr>
</thead>
<tbody>
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<td>(0.58)</td>
</tr>
<tr>
<td>( drre )</td>
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<td>-0.1934</td>
<td>-0.1356</td>
<td>—</td>
</tr>
<tr>
<td></td>
<td>(-5.60)</td>
<td>(-7.57)</td>
<td>(-5.12)</td>
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<tr>
<td>( oildep )</td>
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<tr>
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<td>(-3.64)</td>
<td>(-6.02)</td>
</tr>
</tbody>
</table>

### Table 3B

**Homogeneous panel OLS bias corrected cointegration estimates**

Dependent variable: \( \text{rerdol}_{it} \)

<table>
<thead>
<tr>
<th>Variables / Model</th>
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<th>M2</th>
<th>M3</th>
<th>M4</th>
</tr>
</thead>
<tbody>
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<td>—</td>
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<td>(-6.34)</td>
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</table>

*Note: t-values in parentheses. Significant coefficients in bold.*
Table 4
Heterogeneous individual and panel
LM cointegration tests results

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<th>Model</th>
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<td>0.0418</td>
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<tr>
<td>Canada</td>
<td>0.0114</td>
<td>0.0216</td>
<td>0.0594</td>
<td>0.0323</td>
</tr>
<tr>
<td>Denmark</td>
<td>0.0138</td>
<td>0.0365</td>
<td>0.0245</td>
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Notes:
(a) The tests and the models have been estimated using COINT 2.0 in GAUSS 3.24 using the procedures provided by S. McCoskey and C. Kao.
(b) The critical values at 1% (***) , 5% (**), and 10% (*) for the LM tests are the following: with one regressor, 0.549, 0.3202 and 0.233; with two regressors, 0.372, 0.21, and 0.167; with three regressors, 0.275, 0.159 and 0.120 (Harris and Inder, 1994). The critical value for the panel LM test is 1.64.
Table 5
Panel cointegration.
Individual DOLS parameter estimates\(^{25}\)

**Model 1:** \(rer dol_{it} = \alpha + \beta_1 pront_{it} + \beta_2 drre_{it} + \beta_3 oildep_{it}\)

**Model 2:** \(rer dol_{it} = \alpha + \beta_1 pront_{it} + \beta_2 drre_{it}\)

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Note:
(a) t-Students are reported in parentheses. Significant coefficients in bold.

\(^{25}\)The intercepts have been excluded to gain in clarity.
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Note:
(a) t-Students are reported in parentheses. Significant coefficients in bold.

The intercepts have been excluded to gain in clarity.
Table 7
Net oil imports as a percentage of GDP (toe per thousand 90 US$)

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Source: OECD.
Table 8
Cointegration Test Statistics

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Note: The critical values have been obtained from Osterward-Lemm (1992). An asterisk denotes rejection of the null hypothesis at 5% significance level.
Figure 1: Stability of the cointegration space

Test of known beta eq. to beta(t)

(a) Test of constancy of beta

(b) Stability of the adjustment coefficient
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