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Mariam Camarero

Cecilio Tamarit

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A panel cointegration approach to the estimation of the peseta real exchange rate.[*]

Mariam Camarero[†]
Jaume I University and FUNCAS

Cecilio Tamarit
University of Valencia and FUNCAS

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Abstract

In this paper we estimate different specifications of a model for the determination of the bilateral real exchange rate of the peseta relative to nine European Union members. The model is based on Meese and Rogoff (1988) monetary approach as extended by MacDonald (1998). The applied econometric techniques are the recent panel cointegration tests developed by Kao (1999), McCoskey and Kao (1998) and Pedroni (1999) for homogeneous and heterogeneous panels. The results are favorable to a model containing relative productivities in tradables and non-tradables and the real interest rate differentials as explanatory variables.

Keywords: real exchange rate, European Monetary Union, panel cointegration

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

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[†]Corresponding author: Department of Economics. University Jaume I. Campus del Riu Sec. E-12071 Castellón (Spain). e-mail: camarero@eco.uji.es.

1 Introduction.

The aim of this paper is to study the main determinants of the Spanish peseta real exchange rate during the ‡oat and the …rst years of European Monetary System (EMS) membership. The study of the real exchange rate is particularly relevant under the European Monetary Union (EMU), once the member countries have …xed their bilateral nominal exchange rates.

The peseta and the Spanish economy o¤er an interesting case study to analyze the e¤ects of a process of opening to international competition in a context of regional integration. The Spanish experience can be relevant for the assessment of the most adequate path to liberalization in emerging economies in the light of the European Union enlargement process.

Following the 1986 entry into the European Community, Spain dismantled the majority of the restrictions on trade and on international capital ‡ows. The entry of foreign capital tended to appreciate the currency. Both domestic and foreign investment increased to accommodate the domestic market to the new competitive conditions. A natural outcome of this opening process would be the increase in productivity, mainly in the manufacturing sector, more exposed to international competition. According to the Balassa-Samuelson e¤ect, this would cause an appreciation of the domestic currency.

During this same period, democracy in Spain introduced drastic changes in …scal policy, not only in the form of new income taxes and the VAT imported from the European Community, but also with new orientations of government expenditure, that adopted an active and expansionary approach.

Finally, from 1986 (and specially since the 1989 entry in the EMS) the Spanish monetary authorities followed the so-called "competitive disin‡ation" strategy. In this framework, by keeping locked the nominal exchange rate, either the domestic prices are kept at the same (or lower) level as the competitors or a rise in the terms of trade causes a progressive loss of external markets. This would provoke a more costly adjustment in terms of recession and, consequently, unemployment, but would …nally improve the competitiveness of the economy.

Bearing this strategy in mind, the Spanish authorities adopted, specially during the period 1989-1992, a "hard peseta" policy in order to maintain the nominal exchange rate target. These measures also provoked a considerable appreciation of the real exchange rate of the peseta challenging, at the same time, the whole strategy.

This paper investigates the factors, from the demand and the supply-side of the economy that may have been in‡uential in the process of appreciation that su¤ered the peseta during the eighties and the beginning of the nineties. Although there has been an increase in traded-goods productivity in Spain during the period analyzed[1], the Balassa-Samuelson effect alone may not explain all the peseta appreciation, so that the demand conditions should play an important role in the Spanish economy. Hence, the study of these factors is of crucial importance for the future developments in the EMU, where another episode of real appreciation can be very dangerous for the competitive position of Spain, due to the absence of the nominal exchange rate as a policy instrument.

The empirical literature of real exchange rate determination has traditionally obtained mixed evidence when using structural monetary models, so that more eclectic formulations were also adopted in many empirical studies. Recently, the traditional models have been revisited using cointegration techniques. However, the short spans of data available make it difficult to extract reliable conclusions. In this paper we have made an attempt to overcome these problems by applying new panel cointegration tests that allow us to extend the number of countries for a total of 20 years (from 1973 to 1992) in order to estimate the Spanish peseta bilateral real exchange rate relative to 9 of its European Union partners.

The theoretical formulation adopted in this paper is based on MacDonald's (1998) eclectic approach, that encompasses several structural models traditionally tested in the applied literature. This will permit to discriminate between different models: from the Meese and Rogoff (1988) real interest differential model to the Rogoff (1992) intertemporal model based on productivity differentials and relative public expenditure, and to the traditional Balassa-Samuelson effect.

The analysis will be performed with annual data for 10 European Union member countries for the period 1973-1992. Concerning the econometric techniques, we use panel data cointegration methods, as proposed by Kao (1999), McCoskey and Kao (1998) and Pedroni (1999). These recent tests combine two approaches: first, the use of panel data that traditionally has measured relationships between changes in exchange rates and changes in its determinants and, second, the use of cointegration techniques that measures long-run relationships between the level of the exchange rate and the level of its determinants. The panel cointegration tests add, consequently, new insights to both the time series techniques and the cross-section studies, because previous cross-section analysis could not obtain long-run or equilibrium models for the exchange rates, whereas the short spans of time-series data available make difficult to extract reasonable estimates of the long-run cointegrating vectors[2]. This kind of tests may be useful for many cases in empirical work where there is a limited availability of data. Pooling the data across individual members of a panel makes available more information concerning the cointegration hypothesis.

The paper is organized as follows: the second section reviews the theoretical and applied literature and presents the set of variables that may explain the behavior of the peseta real exchange rate; the third section is devoted to the discussion of the empirical results, and the fourth concludes.

2 Theoretical model: finding the determinants of the real exchange rate.

There is a wide academic agreement on the difficulties commonly found in applied work to model exchange rates. The monetary structural models of the seventies failed in the eighties when their forecasts were compared to a simple random walk model in the seminal paper of Meese and Rogoff (1983). A similar debate exists concerning the fulfillment of PPP. In both cases, new econometric techniques, such as cointegration, have boosted a new generation of applied research in this field. However, the results are still not conclusive and cannot give complete support to the traditional theories.

Trying to improve the results of the structural monetary models, Meese and Rogoff (1988) studied the link between real exchange rates and real interest rate differentials. Starting from a monetary model for the determination of the real exchange rate, q_t, this variable is defined as:

$$q_t \equiv e_t + p_t^* - p_t \tag{1}$$

where e_t is the logarithm of the domestic price of one unit of foreign currency and p_t and p_t^* are the logarithms of domestic and foreign prices. Thus, when the real exchange rate increases, the domestic currency depreciates. 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:

$$E_t(q_{t+k} - q_t) = R_t - R_t^* \tag{2}$$

where R_t and R_t^* are, respectively, the real domestic and foreign interest rates for an asset of maturity k:

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

$$q_t = \hat{q}_t - \varphi(R_t - R_t^*) \tag{3}$$

where φ is a positive parameter larger than unity.

This leaves relatively open the question of which are the determinants of $\hat{q}_t$, that is a non-stationary variable.

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 (3). The simplest model will assume that the expected real exchange rate is constant[3], while the rest of the models will be specified using other determinants.

The second approach relaxes the assumption that the expected real exchange rate is constant and 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[4].

An additional factor that has also been considered in the literature, together with the previous determinants, is the productivity differential[5], considered a major source of supply shocks affecting the real exchange rate and also a proxy variable for the Balassa-Samuelson effect.

MacDonald (1998) follows a similar approach, dividing the real exchange rate determinants into two components: the real interest rate differential and a set of fundamentals which include productivity differentials, a demand side bias, the effect of relative fiscal balances on the equilibrium real exchange rate, the private sector savings and the real price of oil.

Finally, 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, in contrast to the more traditional monetary approaches. Rogoff (1992) develops an intertemporal model for exchange rate determination for the case of relatively closed capital markets and factors that are not perfectly mobile across sectors.

Recently, the use of panel cointegration and unit root tests for panels of countries, such as in Chinn and Johnston (1996), Papell (1997), Anthony and MacDonald (1998), Cecchetti, Mark and Sonora (1998), Papell and Theodoridis(1998), and MacDonald and Nagayasu (2000) have permitted to obtain more encouraging results.

Therefore, in this paper, we will adopt an eclectic view closely related to MacDonald (1998) that includes the majority of the variables suggested in the literature and we will estimate a model for the bilateral real exchange rate of the peseta relative versus a group of European Union partners using panel cointegration techniques.

MacDonald (1998) discusses a real exchange rate decomposition and defines the real exchange rate in a similar fashion as in equation (1). He arrives at the same expression that can be obtained from the Meese and Rogoff (1988) monetary approach in equation (3). Thus, the actual equilibrium exchange rate has two components: the unobservable expectation of the exchange rate, $\hat{q}_t$; driven by the fundamentals and the real interest rate differential, that may depend on other determinants.

Then, he further decomposes the real exchange rate in order to explain which are the group of fundamentals that may determine $\hat{q}_t$. He also proposes using the cointegration framework to estimate the static relationship given in expression (3).

Thus, a similar relationship to the one for the real exchange rate in equation (1) may hold for the price of traded goods:

$$q_t^T \equiv e_t + p_t^{T*} - p_t^T \qquad (4)$$

where a T superscript indicates that the variable is defined for traded goods. If the prices in (4) are composite terms, for q_t^T to be constant it has to be assumed that each of the goods prices which enters p_t^T has an equivalent counterpart in p_t^{T*}; and the weights used to produce these composite price levels are the same.

Then, the general price indexes can be decomposed into traded and non-traded components as:

$$p_t = (1 - \alpha_t)p_t^T + \alpha_t p_t^{NT} \qquad (5)$$

$$p_t^* = (1 - \alpha_t^*)p_t^{T*} + \alpha_t^* p_t^{NT*} \qquad (6)$$

where the α_t denote the shares of nontradable goods sectors in the economy, and are assumed to be time-varying, and NT denotes the non-traded goods. By substitution, MacDonald (1998) obtains an expression for the long-run equilibrium real exchange rate, $\hat{q}_t$:

$$\hat{q}_t \equiv q_t^T + \alpha_t(p_t^T - p_t^{NT}) - \alpha_t^*(p_t^{T*} - p_t^{NT*}) \qquad (7)$$

Equation (7) permits to distinguish three potential sources of long-run real exchange rate variability: first, nonconstancy of the real exchange rate for traded goods, that may arise if the goods entering international trade are imperfect substitutes; second, movements in the relative prices of traded to non-traded goods between the home and foreign country, due to productivity differentials in the traded goods sectors; finally, differences in the weights used to construct the prices in the home and foreign country. However, these weights will be considered constant, due to the difficulties found to assess clearly their importance.

- The first factor considered by MacDonald (1998) relates to the relative price of traded to non-traded goods across countries, reflected in the term $(p_t^T - p_t^{NT}) - (p_t^{T*} - p_t^{NT*})$. The differences may be affecting the relative price of non-traded goods, without necessarily affecting the relative price of traded goods. He points out at two separate sets of effects. The first one would be the Balassa-Samuelson effect, while the second one has to do with a demand-side bias towards non-traded goods.

  According to the Balassa-Samuelson effect, the nominal exchange rate moves to ensure the relative price of traded goods is constant over time, so that $q_t^T = c$. However, productivity differences in the production of traded goods across countries can cause a bias into the overall exchange rate because productivity advances tend to concentrate in the traded goods sector. Then, if prices of goods are linked to wages, wages to productivity and wages linked across industries, then the real exchange rate tends to appreciate for fast growing countries, even if the law of one price holds for traded goods. Consequently, if the home country is a fast growing economy, the term $(p_t^T - p_t^{NT}) - (p_t^{T*} - p_t^{NT*})$ will be negative and this pushes $\hat{q}_t$ below q_t^T.

  The second effect is related to the existence of a non-traded good bias in demand, that pushes the exchange rate away from its PPP level defined using traded goods prices. If the income elasticity of demand for non-traded goods is greater than unity, the relative price of non-traded goods will rise as income rises[6]. This effect is reinforced by the fact that the share of government expenditure devoted to non-traded goods is larger than the share of private expenditure. Consequently, the demand bias will also tend to appreciate the real exchange rate.

- Despite the assumption made above in which q_t^T is constant, this fact is not uncontroverted. In fact, the imperfect substitutability of the traded goods produced by industrial countries makes that the existing price differences may not be quickly arbitraged away. According to MacDonald (1998), two factors may introduce variability in q_t^T: international differences in savings and investment, and the real price of oil.

  Concerning the first factor, the relative price of non-traded goods is a major determinant of the goods and nonfactor services component of the current account. The current account depends on the determinants of national savings and investment, and the fiscal balance is one key component of national savings. Thus, private savings and the relative fiscal policies will influence the equilibrium real exchange rate.

The literature on whether fiscal policy consolidation strengthens or weakens the domestic currency is wide and has not solved the question yet. In the traditional Mundell-Fleming model, a tightening of fiscal policy increases national savings and, thus, lowers the domestic real interest rate and depreciates the currency. In contrast, the portfolio balance models and the asset market/balance of payments synthesis model of Frenkel and Mussa (1988) consider that any interest earnings on net foreign assets are offset by a corresponding trade imbalance. Hence, a permanent fiscal consolidation will increase net foreign assets and appreciate the long-run real exchange rate.

The other factor, private sector savings, tend to be very constant over time. However, recent research, such as Masson et al (1993) suggest that demographic factors may influence differences across countries in the savings rate and, thus, affect net foreign asset positions[7].

Changes in the real price of oil can also have an effect on the relative price of traded goods through their effect on the terms of trade. This issue has already been used by Rogoff (1992) and by Amano and van Norden (1998) who find that the real price of oil captures the majority of the terms of trade shocks. For countries that are oil importers, the real exchange rate would depreciate relative to oil producer countries[8].

The eclectic model proposed by MacDonald encompasses, thus, other theoretical models and empirical specifications widely used in the literature, as it has been summarized above. In particular, the intertemporal approach of Rogoff (1992) can be easily tested in this context.

- Rogoff (1992) model is based on the dynamic micro-foundation models of the real exchange rate. 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 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.

Assuming that PPP holds for tradables, (so that as stated earlier $q_t^T = c$), Rogoff (1992) arrives to one expression of the type:

$$\hat{q}_t = -\Gamma[(a_{Tt} - a^*_{Tt}) - \zeta_N(a_{Nt} - a^*_{Nt}) + (\zeta_N - 1)(g_t - g^*_t) + p_0] \quad (8)$$

where Γ is a parameter related to the weights of the tradables and non-tradables in the aggregate price index, $(a_{Tt} - a^*_{Tt})$ and $(a_{Nt} - a^*_{Nt})$ are the difference in traded and non-traded goods productivity in the home and foreign country, respectively, $(g_t - g^*_t)$ is the difference in real public expenditure to GDP between the two countries, ζ_N is the ratio of non-traded goods output to private non-traded goods consumption and p_0 represents the initial conditions from the intertemporal model.

This model also encompasses the Balassa-Samuelson model, if one assumes that the parameter corresponding to government spending $(\zeta_N - 1)$ is zero.

Rogoff (1992) and Chinn (1997) use this expression to test the model, and the former also recommends to include the real price of oil to account for possible shifts in the terms of trade[9].

Consequently, the models to be estimated in the paper will include combinations of the following set of explanatory variables:

$$q_t = f[\underset{(-)}{(R_t - R^*_t)}; \underset{(-)}{(a_{Tt} - a^*_{Tt})}; \underset{(+)}{(a_{Nt} - a^*_{Nt})}; \underset{(-/+)}{(g_t - g^*_t)}]$$

where $(R_t - R^*_t)$ is the real interest rate differential, $(a_{Tt} - a^*_{Tt})$ and $(a_{Nt} - a^*_{Nt})$ the productivity differentials in the traded and non-traded goods sectors, and $(g_t - g^*_t)$ the difference in the real public expenditure to GDP ratio.

To sum up, according to the expression above, an increase in the real interest rate differential would cause the national currency to appreciate, from the assumptions of the monetary model summarized in equation (3): larger domestic real interest rates will attract capitals and increase the demand for the domestic currency.

Second, when domestic productivity in the traded goods sector increases relative to foreign productivity, there is a tendency of the domestic currency to appreciate. Higher productivity in the tradables makes both labor and capital more productive in that sector. Thus, the domestic output in the traded sector also increases and the domestic …rms are able to reduce traded prices. This will cause a reduction in the relative price of traded/non-traded goods and, as a result, the domestic currency appreciates in real terms.

In contrast, when the domestic productivity of non-tradables increases relative to foreign nontraded productivity, this induces workers to move from the traded to the more productive nontraded sector[10], so that the domestic output of nontraded goods would increase. This makes it possible for the domestic …rms to reduce the price of nontraded goods, so that the relative price of traded/non-traded goods will increase and the real exchange rate would depreciate.

Finally, according to Rogo¤ (1992) public expenditure is biased towards non-traded goods, as it has been stated in the theoretical comments above. When the domestic real public expenditure to GDP is larger than the foreign variable, the real exchange rate appreciates, due to the increase in non-tradables demand that causes the traded/non-traded price ratio to decrease. However, it should be also taken into account that a positive sign could also be found, according to Rogo¤ (1992) if public spending is not so clearly biased towards non-traded goods or, alternatively, if it is the traditional Mundell-Fleming model approach that works.

3 Empirical results.

For the empirical analysis of the paper, the set of variables …nally chosen and combined are the following:

$$rer_{it} = f[\underset{(i)}{drr_{it}}; \underset{(i)}{rprotra_{it}}; \underset{(+)}{rprontra_{it}}; \underset{(+=i)}{dpexp_{it}}] \quad (9)$$

where rer_{it} is the bilateral real exchange rate of the peseta based on CPIs; drr_t is the real interest rate di¤erential, $rprotra_{it}$ and $rprontra_{it}$ are the relative productivity of the traded and non-traded goods in the domestic economy, respectively; and $dpexp_t$ is the di¤erential in real domestic public expenditure over GDP relative to the foreign country[11].

We have considered 10 European Union countries, having excluded Luxembourg, Greece, Ireland, Portugal and Finland due to problems of data availability. This permits us to obtain 9 relative variables, that give the …nal dimension of the panel.

Previous to the study of the order of integration of the variables, we have included in Figure 1 as an example the plots of the variables corresponding to the French case. From the observation of the graphs it can be derived that the relationships existing between the real exchange rate and its determinants in the case of France closely follow what theory predicts.

Insert Figure 1.

There are some general trends that can be drawn from the graphical inspection of the variables. First, the real exchange rate of the peseta experiences a progressive appreciation from the seventies, that became even sharper at the end of the eighties and beginning of the nineties to suffer later four devaluations. Thus, the competitive position of the Spanish economy has deteriorated continuously during the period analyzed. This deterioration has been only partially offset by the devaluations, although the ones in 1993 and 1995 are beyond the sample period.

The time-path of traded-goods relative productivity is shown in the upper left graph of Figure 1. The main pattern in tradables productivity differential with France is the steady growth of this variable that started at the beginning of the eighties. Only at the end of the decade this differential slightly decreases. From the graph, its relation with the real exchange rate is generally negative for the period analyzed, although it may change at the end of the sample. In contrast, the lower left graph shows a very different pattern. The non-tradables productivity differential seems to be positively related with the real exchange rate and no significant changes in this behavior can be observed.

The upper right panel shows the time-path of the real interest rate differential and the dependent variable. In this case, the relation between the two series is negative, so that an increase in the real interest rate differential would cause a real appreciation of the currency, as theory predicts.

Finally, in the lower right panel, the public expenditure differential shows a positive slope. This trending behavior is due to the small size of the Spanish public sector at the beginning of the sample. Although public expenditure has grown in all European countries, the Spanish one grew faster, due to its initial low level. It should be noted that the sharp increase experienced at the end of the sample (versus France, in this case) is contemporaneous to the real exchange rate appreciation of the end of the eighties. The negative relationship observed in the graph would give support to the portfolio balance interpretation of the linkage between the two variables, as well as to the negative relation postulated by the Rogoff (1992) model.

3.1 Panel cointegration.

Previous to the cointegration analysis, the Kwiatkowski, Phillips, Smith and Shin (1992) stationarity tests pointed at the non-stationarity of the variables[12].

In this section we apply various recent panel cointegration tests to the model specified in section 2. More specifically, the panel tests that have been implemented in this paper are: first, the DF and ADF-type tests proposed by Kao (1999) for the null hypothesis of no cointegration in homogeneous and heterogeneous panels; second, the panel cointegration test proposed by McCoskey and Kao (1998) for the null of cointegration in heterogeneous panels, based on Harris and Inder (1994) LM test developed for time series; finally, we compare the results with the Pedroni (1999) heterogeneous panel and group tests[13].

Due to the variety of explanatory variables that we have presented in section 2, we have proposed five different specifications that we test in this paper, following the theoretical aspects discussed in the previous section:

M1: $rer_{it} = \alpha_i + \beta_{1i} rprotra_{it} + \beta_{2i} rprontra_{it}$ (the Balassa-Samuelson model)

M2: $rer_{it} = \alpha_i + \beta_{1i} rprotra_{it} + \beta_{2i} rprontra_{it} + \beta_{3i} drr_{it}$ (a partial version of MacDonald's model)

M3: $rer_{it} = \alpha_i + \beta_{1i} rprotra_{it} + \beta_{2i} rprontra_{it} + \beta_{3i} dpexp_{it}$ (the Rogoff (1992) model)

M4: $rer_{it} = \alpha_i + \beta_{1i} drr_{it}$ (the Meese and Rogoff (1988) model)

M5: $rer_{it} = \alpha_i + \beta_{1i} dpexp_{it}$ (a restricted version of the Rogoff (1992) model)

The first model is a productivity-version[14] of the Balassa-Samuelson model, that is encompassed by models M2 and M3. The former is a specification that includes the Meese and Rogoff (1988) real interest differential model and the Balassa-Samuelson effect, that are three of the variables proposed by MacDonald (1998). In the specification M4 the simplest version of Meese and Rogoff (1988), when the long-run real exchange rate is assumed to be constant, is tested. Model M3 is the specification presented in Rogoff (1992), whereas M5 is a restricted version of this model, where the productivity effects are assumed not to be significant.

These five models are tested for the 9 bilateral real exchange rates in the sample covering the period 1973-1992. In addition, for each model we will first assume that all the elements of the panel share the same slope parameters. This implies that $\beta_{11} = \beta_{12} = \dots = \beta_{19} = \beta_1$; the same restriction would apply for the rest of the slope parameters in the five specifications. This is what we call the "homogeneous" model. If we relax this strong assumption and allow for the slope parameters to differ across the panel, we will be testing the "heterogeneous" model. Consequently, we will present both cointegration tests and parameter estimates for the homogeneous and the heterogeneous models. However, in the two types of models the intercept is allowed to be different for the cross-sections.

The ordering of the panel cointegration results is based on this duality. In table 1 we first present the homogeneous Kao (1999) cointegration tests. The next table is devoted to the corresponding parameter estimates by OLS and bias corrected OLS. In table 3 we present the results of the panel and group tests proposed by Pedroni (1999) for heterogeneous panels. Finally, tables 4 and 5 contain the LM and ADF tests for cointegration in heterogeneous panels for the five models specified, table 6 presents some cointegration tests allowing for structural changes, and tables 7 to 10 the parameter estimates corresponding to the five models.

3.1.1 Homogeneous panel cointegration tests and estimates.

The analysis of this section starts with the tests proposed by Kao (1999), based on the OLS residuals and assuming as the null hypothesis the absence of cointegration. The DF_{ρ}^{*} and the DF_{t}^{*} statistics are not dependent on the nuisance parameters, and are computed under the assumption of endogeneity of the regressors. Alternatively, he defines a bias-corrected serial correlation coefficient estimate and, consequently, the bias-corrected test statistics and calls them DF_{ρ} and DF_{t}. Finally, he also proposes an ADF-type test for the null of no cointegration. The results of applying these tests to the five model specifications are presented in table 1. For all the models and tests the null of no cointegration is rejected at 1%.

Insert table 1.

Tables 2A and 2B show the OLS and bias corrected parameter estimates for all the specifications. The results of the two estimation techniques do not differ significantly, with the exception of model M3. For the Balassa-Samuelson model presented first, that includes the two productivity differentials as explanatory variables, the parameters are highly significant and exhibit the expected signs: negative for traded-goods productivity and positive for the non-traded one. It should be noted, however, that the first parameter is larger in the bias-corrected estimates, whereas the second one is smaller. Models M4 and M5 include as explanatory variables the real interest rate differential and the relative public expenditure, respectively. In other words, these two specifications include each one demand-side factor. In the two models and with the two estimators, the parameters have the correct

Insert tables 2A and 2B.

sign and are very significant. Consequently, the four variables considered seem to be adequate to explain the behavior of the real exchange rate during the period analyzed.

The two other specifications combine the productivity variables with drr_{it} (in model M2, the so-called MacDonald specification) and with dp-exp_{it} (model M3, as in Rogoff (1992)). In the case of model M2, the homogeneous specification maintains the values of the coefficients' estimates for the productivity variables and the real interest rate differential presents the correct sign. All the parameters are significant both for the OLS and the bias-corrected estimates. The outcome is different in the case of model M3: the coefficient of government spending is negatively signed (being compatible with Rogoff's model and the portfolio approach) and significant, whereas the other two variables loose their explanatory power, specially in the bias corrected estimation.

As a conclusion, the M2 specification seems to be, for the homogeneous model, the most representative of the relationships linking the real exchange rate and the supply and demand determinants. From this estimated relation, an increase in the traded-good sectors productivity in Spain relative to its competitors tends to appreciate the currency, whereas when this increase occurs in the non-traded good sectors, it follows a depreciation. Finally, from the demand-side, positive real interest rate differentials cause an appreciation of the domestic currency.

3.1.2 Heterogeneous panel tests and estimates.

In this section we present the heterogeneous panel tests and estimates for the five models specified above using the Dynamic OLS (DOLS) estimation technique. The choice of this method relies on the Monte Carlo results of Kao and Chiang (2000), who conclude that the OLS estimator has non-negligible bias in finite samples and the Fully Modified (FMOLS) estimator does not improve the results of the former in general. In contrast, the DOLS estimator seems to be more promising for cointegrated panel regressions. The estimated parameters are efficient and the t-statistics can be used for inference.

Pedroni (1999) proposed several tests for the null hypothesis of non-cointegration for panels with more than one regressor. We present in table 3 the results of these tests for models M1, M2 and M3 (the models with two and three regressors) with intercepts but no trends, that is the specification used here for the rest of the tests[15]. From the seven tests proposed by Pedroni (1999) four are called panel tests, that presume a common value for the unit root coefficient, whereas the group tests allow for differences in this parameter and give an additional source of potential heterogeneity. The tests are

Insert table 3.

di¤erent panel versions of the Phillips and Perron rho and t-statistic, as well as panel versions of the ADF test. In fact, the parametric panel t-statistic is a panel cointegration version of the Levin and Lin (1993) panel unit root test statistic, whereas the group equivalent is analogous to the Im, Pesaran and Shin (1997) group mean unit root statistic[16].

In a Monte Carlo experiment, Pedroni (1997) compares the performance of the seven statistics in terms of size distortion and power, …nding …rst that the panel variance statistic was normally dominated by the other tests. In addition, the panel rho statistic exhibited the least distortions, whereas the group ADF had the worst and the group PP, panel PP and panel ADF fell in between. Concerning power and short samples (T=20, as in our case), the group ADF generally performs best, followed by the panel ADF and the panel rho.

Bearing in mind these considerations, the results of the tests do not di¤er signi…cantly between the three models analyzed. The null of non-cointegration is rejected at 1% for the three speci…cations using both the panel and group ADF tests. However, it seems again that there is slightly more support for model M2, that is, MacDonald's extended version of Meese and Rogo¤ (1988) model, than for the other two alternatives: the only test that does not permit to reject the null for this model is the panel rho.

Tables 4 and 5 present the individual and panel LM and ADF tests for the heterogeneous case. It should be noted, before analyzing the results, that in the …rst test the null hypothesis is cointegration, whereas in the ADF test, the null is absence of cointegration.

Insert tables 4, 5 and 6.

For the Balassa-Samuelson model (M1) the LM test by McCoskey and Kao (1998) of the null of cointegration cannot be rejected either for the individual countries or for the panel. The ADF test results, however, allow for the rejection of the null at 5% for the panel, Austria and France, whereas the cases of the Netherlands, Sweden and the UK are only rejected at 10%. For the rest, the absence of cointegration cannot be rejected. However, the Monte Carlo experiments reported by McCoskey and Kao (1999) point at the larger power of the LM test if compared with other residual based tests for cointegration in heterogeneous panels under the null of no cointegration (the average ADF test and Pedroni's pooled tests).

This same pattern is present in models M2 and M3, where the LM tests, both individual and panel, support the cointegration hypothesis, whereas the evidence depends on the country for the ADF-type individual tests. However, the panel ADF tests reject in the two cases the null of no cointegration.

In contrast, the results for models M4 and M5, that only include demand-side variables, are less favorable to the existence of cointegration. The ADF individual and panel tests do not permit to reject the absence of cointegration. The LM tests are more supportive of the cointegration hypothesis, although it is rejected for France and Italy in model M4.

Concerning the parameter estimates, we present the results for the five models and the nine countries in tables 7 to 10. For model M1, the estimated coefficients for traded and non-traded relative productivity are presented in table 7. For the first variable, 7 out of 9 coefficients are significant and 5 of them present the "correct" sign. For non-tradables, 4 are the significant coefficients, all of them positive, as the theory predicts. A similar pattern can be observed in table 8 for model M2. However, the number of significant coefficients is now larger for the non-traded goods: 6 out of 9 and all of them positive. Concerning the variable that has been added to this specification, drr_{it}, it is significant for 6 countries and the coefficient is negative. The parameter values are also very close to those obtained in the homogeneous estimation using OLS. Finally, in the specification that includes the relative productivities and the public expenditure differential (model M3 in table 9), there are only three significant coefficients for the two productivities and only one for the fiscal variable.

Insert table 7.

Insert table 8.

Insert table 9.

The other two model estimation results, presented in table 10, also support the explanatory power of the two demand variables alone. In model M4, the real interest rate differential is significant in all the countries analyzed with the only exception of Italy. In the case of $dpexp_{it}$ 5 out of 9 coefficients are significant, the exceptions being Sweden, Austria, Italy and the UK.

Insert table 10.

If we combine the test results and the estimated parameters, also the heterogeneous estimation seems to give support to model M2 as the most complete and robust specification from the five proposed[17]. In addition, it should be emphasized that the parameters do not differ substantially across the cross-sections, specially for the demand variables. This permits us to conclude that the performance of the real exchange rate of the peseta presents similar patterns for all the countries in the sample. Begum (2000) has found similar parameter estimates for productivity differentials in the case of the G-7 countries using the Johansen cointegration method.

3.2 Stability tests.

As an additional test to the specification finally chosen (M2), we have applied the Gregory and Hansen (1996) tests for cointegration with regime changes that have been computed for each individual country. The results are presented in table 7, and should be interpreted together with the ADF tests results in table 6 for the heterogeneous panel. Two tests and two models have been estimated: the ADF^* and the Z_t^* tests in model C where there is a change in the intercept, as well as in model C/T where a trend is also present.

According to Gregory and Hansen (1996), both the ADF and the ADF^* statistics test the null of no cointegration, so that rejection of the null using any of them can be considered evidence in favor of cointegration and the existence of a long-run relationship in the data. However, the ADF^* test should be used more as a pre-test (previous to applying the conventional ones) than as specification test. The Hansen (1992) is a better alternative for specification purposes, because the null hypothesis is the existence of a shift and the alternative is absence of this shift.

In addition, if the standard ADF statistic does not reject but the ADF^* does, this implies that structural change in the cointegrating vector may be important. Finally, if one finds rejection from both tests, no inference about structural change can be derived. The analysis should be complemented, in this case, using the Hansen (1992) tests.

The results presented in table 7 point to the general rejection of the null hypothesis of no-cointegration (if considered separately from the results of the individual ADF tests in table 6). In the case of Austria using the two specifications (models with intercept and with intercept and trend) and Belgium for model C, the null can only be rejected at 10%. In addition, the conventional results for model 2 (see the third column in table 6) do not allow the rejection of the null hypothesis. As Gregory and Hansen (1996) point out, rejection with any of the two tests would indicate the existence of a long-run relation, and this is the general pattern observed in our results.

The empirical evidence found using the Z_t^* is less favorable to the rejection of the null hypothesis, with the exception of France, Sweden and the UK.

Therefore, if one considers the possibility of a structural change, there is evidence of instabilities (as detected by the Gregory and Hansen (1996) statistics) that occur (according to the ADF^* test results) at the end of the seventies or beginning of the eighties (being the exceptions France and Italy). Even if this was the case, the simultaneity of the majority of the instability episodes would offset, at least partially, the economic consequences of possible structural changes.

4 Conclusions.

In this paper we have studied the determinants behavior of the peseta real exchange rate in relation to 9 of its European partners for the period 1973-92, using panel cointegration tests. The theoretical formulation adopted in this paper, based on MacDonald's (1998) eclectic approach, has permitted to compare several structural models traditionally tested in the applied literature.

The empirical methodology applied allows for di¤erent degrees of ‡exibility in the speci…cation of the long-run parameters linking the real exchange rate and its determinants. The panel results for the homogeneous models impose the restriction of common slope coe¢cients for all the countries in the sample, whereas the heterogeneous speci…cation allows for di¤erent estimates for every country. These two levels of results are crucial to discriminate between the models proposed.

To summarize the main empirical results, from models 1, 4 and 5, we can derive that both supply and demand variables have been important in the evolution of the peseta during the period studied. However, when combined, the speci…cation including the real interest rate di¤erential seems to be more supported by the data than the one containing the di¤erence in public expenditure. Moreover, in the case of Spain, the two demand factors (contractionary monetary policy and expansionary …scal policy) appear to lead to an increase of real interest rates and to an appreciation of the currency. Thus, model 2, where only the real interest rate is included from the demand-side of the economy, along with the productivity variables, can be considered the most suitable one. These results can interpreted as giving support to the Balassa-Samuelson model, although the non-tradables productivity differential is more signi…cant that its tradables equivalent. In addition, the hypotheses of Meese and Rogo¤ (1988) and Rogo¤ (1992) that stress the importance of the demand-side of the economy for the determination of the real exchange rate, is also supported by the data.

A Data sources.

The data in the paper is annual and covers the period 1973-92. The panel consists of 10 European countries, that is, all the EU members with the exceptions of Ireland, Luxembourg, Portugal, Finland and Greece due to data availability problems. The data has been obtained from the magnetic tapes of Cronos and the International Monetary Fund International Financial Statistics. 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.

q_{it} : bilateral real exchange rate of the peseta relative to the other European currencies considered. The nominal exchange rate, s_t; has been defined as units of domestic currency to purchase a unit of foreign currency. Source: Cronos.

$$q_t = \log\left(\frac{s_t \times p^*_t}{p_t}\right)$$

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 (source: Cronos) using the Hodrick and Prescott filter.

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

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

$$rr_t = r_t - \pi^e_t$$

$$drr_t = rr_t - rr^*_t$$

where π^e_t is expected inflation filtered using the HP filter; π^t_t is the transitory component of inflation; rr_t is the real interest rate and rr^*_t the foreign rate.

$protra_{it}$: productivity differential in tradables between Spain and each of the other countries in the sample. Tradables productivity includes just manufactured goods. The source is the ISDB of the OECD and the Bank of Spain (Paco de Castro kindly provided the Spanish data).

$prontra_{it}$: productivity differential in non-tradables between Spain and each of the other countries in the sample. We consider non-tradables goods all the sectors excluding manufacturing. 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.

$dpexp_{it}$: public expenditure differential. The government spending is calculated relative to GDP:

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

where $pexpn_t$ is nominal public expenditure, whereas $dpexp_t = pex_t - pex_t^*$: The sources are IMF and Cronos.

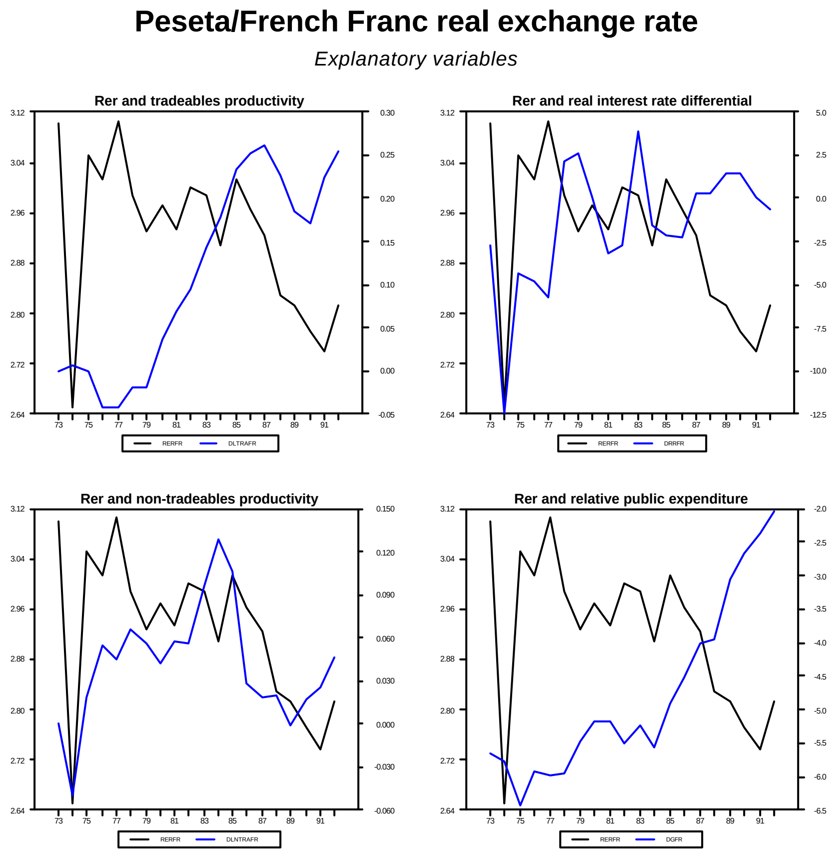

Figure 1:

B Graphs.

C Tables.

Table 1

Homogeneous panel cointegration.
Kao (1999) tests

Model specifications:

$$\text{M1: } rer_{it} = \alpha_i + \beta_1 rprotra_{it} + \beta_2 rprontra_{it}$$
$$\text{M2: } rer_{it} = \alpha_i + \beta_1 rprotra_{it} + \beta_2 rprontra_{it} + \beta_3 drr_{it}$$
$$\text{M3: } rer_{it} = \alpha_i + \beta_1 rprotra_{it} + \beta_2 rprontra_{it} + \beta_3 dpexp_{it}$$
$$\text{M4: } rer_{it} = \alpha_i + \beta_1 drr_{it}$$
$$\text{M5: } rer_{it} = \alpha_i + \beta_1 dpexp_{it}$$

1973-1992

Test / Model	M1	M2	M3	M4	M5
DF_ρ	-8.03^{***}	-10.29^{***}	-10.08^{***}	-9.08^{***}	-9.71^{***}
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
DF_t	3.64^{***}	2.28^{**}	2.44^{***}	3.07^{***}	2.67^{***}
	(0.00)	(0.01)	(0.007)	(0.001)	(0.003)
DF^*_ρ	-12.43^{***}	-15.75^{***}	-13.74^{***}	-14.96^{***}	-13.56^{***}
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
DF^*_t	-4.77^{***}	-5.89	-5.93	-5.09^{***}	-5.71^{***}
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
ADF	-3.97^{***}	-4.19^{***}	-5.43^{***}	-2.54^{***}	-5.09^{***}
	(0.00)	(0.00)	(0.00)	(0.005)	(0.00)

Note: Two and three asterisks denote rejection of the null hypothesis of no-cointegration at 5% and 1% respectively. The p-values are in parentheses. The test statistics are distributed as N(0,1).

Table 2A

Homogeneous panel OLS cointegration estimates

Dependent variable: rer_{it}

Variables / Model	M1	M2	M3	M4	M5
rprotra	-0.4848	-0.3561	-0.1437	—	—
	(-5.36)	(-3.73)	(-1.48)		
rprontra	0.4590	0.4586	0.2239	—	—
	(3.80)	(3.91)	(1.95)		
drr	—	-0.0074	—	-0.0104	—
		(-3.43)		(-4.90)	
dpexp	—	—	-0.0285	—	-0.0335
			(-6.51)	—	(-9.31)

TABLE 2B

Homogeneous panel OLS bias corrected cointegration estimates

Dependent variable: rer_{it}

Variables / Model	M1	M2	M3	M4	M5
rprotra	-0.6466	-0.4704	-0.1957	—	—
	(-4.37)	(-3.73)	(-1.51)		
rprontra	0.3234	0.3303	0.1035	—	—
	(2.04)	(2.25)	(0.75)		
drr	—	-0.0081	—	-0.0130	—
		(-3.33)		(-4.74)	
dpexp	—	—	-0.0341	—	-0.0385
			(-5.60)	—	(-6.82)

Note: t-values in parentheses. Signi…cant coe¢cients in bold.

Table 3

Pedroni (1999) cointegration tests
for heterogeneous panels.

Model with intercepts but without trends

Test / Model	M1	M2	M3
Panel variance test	-1.42¤	-2.09¤¤¤	-1.35¤
	(0.07)	(0.01)	(0.08)
Panel ½ test	-0.05	0.13	0.61
	(0.47)	(0.44)	(0.26)
Panel t-test (non-p.)	-1.66¤¤	-2.17¤¤¤	-1.59¤¤
	(0.04)	(0.01)	(0.05)
Panel t-test (param.)	-152.25¤¤¤	-208.19¤¤¤	-176.24¤¤¤
	(0.00)	(0.00)	(0.00)
Group ½ test	1.04	1.53¤	2.13¤¤¤
	(0.14)	(0.06)	(0.01)
Group t-test (non-p.)	-8.28	-9.65¤	-9.38
	(0.11)	(0.06)	(0.11)
Group t-test (param.)	-4.02¤¤¤	-5.35¤¤¤	-4.98¤¤¤
	(0.00)	(0.00)	(0.00)

Note: One, two and three asterisks denote rejection of the null hypothesis of non-cointegration at 10, 5 and 1% respectively.. All the tests have been normalized, with the exception of the Group t-test in its non-parametric version. The probabilities are in parentheses.

Table 4

Heterogeneous individual and panel
LM cointegration tests results

1973-1992

Model	M1	M2	M3	M4	M5
Austria	0.025	0.008	0.009	0.079	0.044
Belgium	0.109	0.004	0.006	0.046	0.039
Denmark	0.057	0.013	0.016	0.062	0.111
France	0.027	0.006	0.005	0.102	0.071
Germany	0.021	0.006	0.009	0.338¤¤	0.043
Italy	0.051	0.005	0.018	0.326¤	0.153
Netherlands	0.022	0.007	0.010	0.134	0.071
Sweden	0.034	0.006	0.005	0.134	0.123
UK	0.009	0.004	0.005	0.209	0.140
Panel test	-2.86	-3.28	-3.24	-1.36	-2.23

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, 0.167, and 0.217; 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

Heterogeneous individual and panel
ADF cointegration tests results
1973-1992

Model	M1	M2	M3	M4	M5
Austria	-4.18¤¤	-4.71¤	-4.49¤	-3.51¤	-4.03¤¤
Belgium	-2.07	-1.90	-3.33	-1.75	-3.07
Denmark	-3.25	-2.66	-3.26	-2.20	-3.11
France	-3.15	-3.53	-5.31¤¤¤	-2.82	-4.62¤¤¤
Germany	-5.19¤¤¤	-3.96	-5.18¤¤¤	-1.17	-5.21¤¤¤
Italy	-3.11	-3.11	-6.12¤¤¤	-1.63	-1.61
Netherlands	-3.85¤	-3.88	-4.32¤	-2.20	-3.92¤¤
Sweden	-3.99¤	-3.91	-4.54¤¤	-0.60	-1.85
UK	-3.84¤	-4.37¤	-3.85	-2.10	-1.88
Panel test	-4.35¤¤	-4.20¤	-7.59¤¤¤	1.74	-2.97

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 lag order of the ADF tests is 1.

(c) (b) The critical values at 1% (***), 5% (**), and 10% (*) for the ADF tests are the following: with one regressor, -4.36, -3.80 and -3.51; with two regressors, -4.64, -4.15 and -3.84; with three regressors, -5.04, -4.48, and -4.19, and have been taken from Phillips and Ouliaris (1990).

Table 6
Gregory and Hansen (1996) tests for a structural change in the cointegration relationship

Model	Model 2 (C)				Model 3 (C/T)			
Country	ADF*	T_b	Z_t^*	T_b	ADF*	T_b	Z_t^*	T_b
Austria	-5.03* (K=1)[a]	1982	-3.98	1978	-5.54* (K=1)	1977	-4.62	1978
Belgium	-5.06* (K=1)	1980	-5.07*	1979	-5.76** (K=1)	1980	-4.30	1979
Denmark	-6.27*** (K=1)	1978	-5.17*	1978	-6.24*** (K=1)	1978	-4.90	1978
France	-6.36*** (K=0)	1987	-6.71***	1988	-6.61*** (K=2)	1984	-6.47***	1985
Germany	-6.22*** (K=1)	1982	-4.33	1978	-5.70** (K=1)	1982	-4.06	1978
Italy	-5.37** (K=1)	1989	-4.99	1989	-5.69** (K=1)	1989	-5.28	1982
Netherlands	-6.23*** (K=1)	1978	-4.96	1978	-6.16*** (K=1)	1978	-4.89	1978
Sweden	-6.71*** (K=0)	1977	-7.05***	1977	-5.86** (K=0)	1977	-6.04***	1977
UK	-6.99*** (K=2)	1981	-7.29***	1979	-6.78*** (K=2)	1981	-6.94***	1980

Notes:

(a) K stands for the number of lags in the AR for the ADF test.

(b) The critical values have been obtained from Gregory and Hansen (1996), table 1, and are for the case of three explanatory variables, -5.77, -5.28 and -5.02 at 1%, 5% and 10%, respectively in model 2 (C). For model 3 (C/T), the critical values are -6.05, -5.57 and -5.33, at the same significance levels. Rejection of the null hypothesis is marked with asterisks.

(c) The GAUSS codes to compute the tests have been obtained from B. Hansen web-page.

TABLE 7

Panel cointegration.
Individual DOLS parameter estimates
Model 1: 1973-92

Country	$intercept_i$	$rprotra_i$	$rprontra_i$
Austria	**2.096**	0.168	**2.428**
	(57.56)	(0.18)	(2.08)
Belgium	**1.004**	-0.385	0.546
	(6.57)	(-0.18)	(0.46)
Denmark	**2.795**	**-1.200**	0.221
	(31.61)	(-2.65)	(0.20)
Germany	**4.050**	**-0.820**	**4.308**
	(66.74)	(-3.92)	(3.16)
France	**3.013**	**-0.452**	**1.991**
	(152.38)	(-3.67)	(5.51)
Italy	**-2.452**	**2.574**	-0.125
	(-37.09)	(5.50)	(-0.54)
Netherlands	**4.077**	**-2.739**	**1.459**
	(56.97)	(-3.54)	(2.09)
Sweden	**3.124**	**-1.034**	0.504
	(43.10)	(-3.20)	(1.55)
UK	**5.149**	**2.015**	0.027
	(140.68)	(3.60)	(0.07)

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

TABLE 8

Panel cointegration.
Individual DOLS parameter estimates

Model 2. 1973-92

Country	$intercept_i$	$rprotra_i$	$rprontra_i$	drr_i
Austria	**2.092**	3.769	-1.877	-0.031
	(62.50)	(1.56)	(-0.58)	(-1.35)
Belgium	**0.829**	-0.334	**1.072**	**-0.061**
	(21.38)	(-0.72)	(3.88)	(-11.58)
Denmark	**2.405**	0.552	**1.948**	**-0.045**
	(13.71)	(0.56)	(1.75)	(-2.09)
Germany	**3.822**	**0.367**	**3.651**	**-0.035**
	(83.73)	(1.80)	(5.81)	(-5.77)
France	**2.935**	**-0.334**	**1.600**	**-0.014**
	(87.70)	(-3.27)	(4.36)	(-2.00)
Italy	**-2.517**	**2.372**	0.127	0.007
	(-40.56)	(4.94)	(0.61)	(0.92)
Netherlands	**2.907**	-0.200	**2.596**	**-0.058**
	(22.49)	(-0.11)	(3.54)	(-2.05)
Sweden	**2.907**	-0.208	**0.931**	**-0.035**
	(22.49)	(-0.39)	(2.77)	(-2.58)
UK	**5.126**	**2.273**	-0.129	0.009
	(71.05)	(2.19)	(-0.25)	(0.35)

Note:
(a) t-Students are reported in parentheses. Signi…cant coe¢ cients in bold.

Table 9

Panel cointegration.
Individual DOLS parameter estimates

Model 3. 1973-92

Country	$intercept_i$	$rprotra_i$	$rprontra_i$	$dpexp_i$
Austria	**2.161**	-0.885	2.666	0.003
	(4.95)	(-0.28)	(0.54)	(0.03)
Belgium	**1.005**	0.885	**1.567**	-0.030
	(10.28)	(0.55)	(1.72)	(-1.29)
Denmark	2.721	-0.698	1.004	-0.001
	(1.69)	(-0.39)	(0.43)	(-0.01)
Germany	**3.497**	0.451	2.227	-0.080
	(6.18)	(0.40)	(0.77)	(-1.20)
France	**2.353**	0.270	0.007	-0.114
	(3.46)	(0.35)	(0.00)	(-0.97)
Italy	**-2.437**	**3.364**	-0.511	-0.052
	(-12.48)	(4.64)	(-1.21)	(-1.15)
Netherlands	**3.778**	**2.215**	**2.179**	**-0.046**
	(49.26)	(4.22)	(12.86)	(-3.42)
Sweden	**3.579**	**-0.705**	**1.887**	0.045
	(14.05)	(-2.48)	(2.78)	(2.11)
UK	**4.936**	2.255	-0.388	-0.026
	(6.05)	(0.72)	(-0.143)	(-0.31)

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

TABLE 10

Panel cointegration.
Individual DOLS parameter estimates

Model 4 and model 5.
1973-92

MODEL	M4		M5	
Country	$intercept_i$	drr_i	$intercept_i$	$dpexp_{it}$
Austria	2.097	-0.012	2.062	-0.026
	(86.66)	(-2.03)	(22.79)	(-1.63)
Belgium	1.017	-0.045	1.107	-0.032
	(48.70)	(-6.97)	(16.72)	(-3.08)
Denmark	2.722	-0.019	2.342	-0.038
	(97.76)	(-3.60)	(13.34)	(-2.60)
Germany	4.040	-0.021	3.706	-0.060
	(155.31)	(-3.36)	(45.88)	(-5.46)
France	2.903	-0.026	2.526	-0.084
	(126.17)	(-2.61)	(29.37)	(-5.58)
Italy	-2.491	-0.004	-2.461	0.006
	(-85.96)	(-0.42)	(-17.67)	(0.09)
Netherlands	3.894	-0.033	3.936	-0.040
	(120.71)	(-3.63)	(73.52)	(-5.64)
Sweden	2.894	-0.041	2.507	-0.035
	(142.27)	(-6.09)	(15.87)	(-3.05)
UK	5.286	0.062	5.151	-0.020
	(208.32)	(3.20)	(32.97)	(-0.93)

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

References

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End notes

[1]According to De Grauwe (1994), the average annual productivity growth of manufacturing in Spain during the eighties was 3.5%, compared to 2.5% in France, 3.0% in Germany and 4.2% in Italy.

[2]Raymon and García-Greciano (1997) have estimated a model for the peseta real exchange rate using traditional panel methods.

[3]See, for example, Campbell and Clarida (1987), Meese and Rogo¤ (1988), Baxter (1994) and MacDonald and Nagayasu (2000).

[4]See Edison and Pauls (1993), Edison and Melick (1995) and Wu (1999).

[5]Strauss (1996), De Gregorio and Wolf (1994) and Lane and Milesi-Ferretti (2000) have studied this type of speci…cation.

[6]That is, as income rises, the demand for services increases.

[7]In the empirical part of the paper we will not consider this variable, that turned out to be stationary.

[8]Initially, we introduced the real oil prices in some of the speci…cations presented later in the paper, but the variable was not signi…cant (or it was but only for one or two of the countries in the sample), so that it was …nally discarded.

[9]It has been assumed that the terms of trade are constant, so that the inclusion of the real price of oil permits to consider a possible source of exogenous shocks in the model.

[10]If one accounts for the assumption of imperfect mobility across sectors as in Rogo¤ (1992), this e¤ect would be weaker.

[11]See appendix A for further details about the de…nition of the variables.

[12]These results have been omitted but are available from the authors upon request.

[13]The econometric procedures necessary to calculate the tests and estimate the coe¢cients have been kindly provided by S. McCoskey and C. Kao. In addition, we have used the program NPT 1.1 (see Chiang and Kao, 2000). All computations have been made in GAUSS 3.24.

[14]This model can also be tested using relative prices of traded and non-traded goods.

[15]Two other speci…cations, without constant and with constant and trend had been estimated, and are available upon request. The results did not di¤er from the ones presented in the text.

[16]See Pedroni (1999) for a detailed description of these statistics.

[17]Problems related to data availability for the breakdown of traded and non-traded goods productivity did not allow us to expand the sample period beyond 1992. However, in a previous version of the paper where we considered aggregate productivity instead, the sample ended in 1997. The results

were also favorable to a specification that included, as the main explanatory variables, the real interest rate differential and productivity differentials. Thus, from this evidence, we think that this pattern in the behavior of the peseta real exchange rate may have been mantained up to 1997.