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# Current Account Imbalances and Real Exchange Rates in the Euro Area†

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**Abstract**

Global current account imbalances have been one of the focal points of interest for policymakers during the last few years. Less attention has been paid, however, to the diverging current account balances of the individual euro area countries. In this paper we consider the dynamics of current account adjustment and the role of real exchange rates in current account determination in the EMU. After controlling for the effects of income growth, we find the relationship between real exchange rates and the current account to be substantial in size and subject to non-linear effects. Overall, we argue that real exchange rates can offer further insights, beyond the effects of the income catch-up process, relevant to current account determination in the EMU.

**Keywords:** current account, real exchange rate, EMU, nonlinearities

**JEL classification:** C51; C52; F31; F32; F41

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## 1. INTRODUCTION

The global imbalances, as manifested by the current account positions of some of the major industrial countries, possibly constitute at present the most challenging issue in international macroeconomics with the main focus being on the US current account deficit.[1] Another type of current account imbalances, however, emerges that has been less intensively scrutinized, namely those within the euro area. While the aggregate euro area current account is currently close to balance, a number of the member states exhibit large current account deficits with a worsening trend. Figure 1 presents the seasonally adjusted current account balance as per pent of GDP in ten EMU countries (quarterly frequency). For example, starting from a balanced current account in the mid-1990s, by 2005 the current accounts of Greece, Portugal and Spain were in deficits equal to 7.9%, 9.2% and 7.6% of GDP respectively. The current accounts of France and Italy display a similar, though less pronounced, deterioration pattern. By contrast, a number of euro area countries display positive current accounts, including Austria, Belgium, Finland, the Netherlands and most notably Germany (4.1% in 2005).

The workhorse approach in assessing current account imbalances focuses on the determinants of saving and investment.[2] In the context of the euro area, Blanchard and Giavazzi (2002) consider how increased economic integration in the euro area may have led to a decrease in saving and an increase in investment which are reflected in a large current account deficit. This effect can be particularly relevant for the poorer EMU countries that are catching up such as Greece and Portugal. Besides high levels of investment and shortfalls in savings, however, the current account deficits may reflect a loss of structural or price competitiveness. The loss of the exchange rate implied by monetary union participation may have direct consequences for the latter. On the one

[1] For example, see Obstfeld and Rogoff (2004, 2005).
[2] For example, see Masson (1998), Chinn and Prasad (2000), IMF (2005).

hand, such developments may not be a cause for concern. In the long run it is expected that euro area members' competitiveness levels may converge as the laggard countries will be catching up, in which case the currently observed increased external borrowing will be offset in the future by higher income levels. Moreover, given that the interest rate and exchange rates are determined at the euro area level, member countries do not face a number of typical balance of payments financing problems (currency attacks, risk premiums, and so on) which outside the EMU could slow down the convergence process. On the other hand, however, the net borrowing of a nation cannot grow indefinitely, even if it takes place within the monetary union. Furthermore, the idea of two groups of countries displaying persistently "wining" and "losing" current account positions within the monetary union cannot be very comforting.

To better understand the dynamics, the sources, and the implications of the diverging EMU current account positions, one need to characterize, among other things, the role of relative prices, that is real exchange rates. While it is widely accepted in theory that shifts in real exchange rates cause changes in the current account, surprisingly, limited recent empirical evidence has been produced explicitly focusing on this relationship in general and no evidence exists for the euro area.[3] The existing literature for the euro area mainly focuses on intertemporal considerations and income growth differences to explaining the eurozone current accounts. In this paper we examine if in addition to those factors the real exchange rate is important for the current account in the long run and for its adjustment dynamics in the short-run.

The relationship between real exchange rates and the current account emerges in the context of traditional approaches (e.g., Friedman, 1953; Mundell, 1962; Dornbusch,

[3] Earlier studies that consider the real exchange rate as the main explanatory variable in estimating current account equations include Edwards (1989), Khan and Knight (1983), and more recently Cline (2003). Another set of studies examines the relationship between real exchange rates and current accounts within a VAR framework (e.g., Lee and Chinn, 2006; Leonard and Stockman, 2001).

1976; Branson, 1983) as well as in the context of the recent new open economy macroeconomics literature (e.g. Obstfeld and Rogoff, 1995). The main channel through which real exchange rate shifts cause current account changes is an "expenditure-switching" effect captured by the IS curve in the variations of the traditional Fleming-Mundell model and the relative price changes in Friedman (1953). The "expenditure-switching" mechanism retains its validity in the Obstfeld and Rogoff (1995) Redux model provided that nominal prices are fixed in the producer country and the exchange rate pass-through is complete.[4] Moreover, this causal link is central in the analysis of the Theory of Optimum Currency Areas (TOCA) on the potential costs of joining a monetary union (see e.g. Mundell, 1961).

Table 1 presents some prima-facie evidence on the potential links between real income growth, real exchange rates and current account balances. We report the average values of the three variables during the post-euro period and a pre-EMU window of equal duration (1992-1998 versus 1999-2005 respectively), as well as the difference between the two periods.[5],[6] The reported correlation coefficients suggest that higher income growth and real exchange appreciation are associated, to a reasonable degree, with movements of the current account. On the other hand, higher income growth does not appear to be correlated with changes in real exchange rates, as postulated by the Balassa-Samuelson hypothesis. Overall, Table 1 suggests that both the intertemporal and the TOCA arguments discussed above may be relevant in explaining the growing intra-EMU imbalances. At the same time, Table 1 reveals the existence of significant differences across individual countries, implying that the links between the variables

[4] For a detailed discussion of the expenditure switching effect in the context of new open economy macroeconomic models see Engel (2002)

[5] Note that the pre-Euro window 1992-1998 corresponds to the period between the signing of the Maastricht Treaty in December 1991 and the introduction of the Euro in January 1999. As such, it is the period covered by the convergence programs implemented by national governments in preparation for the adoption of the single currency in 1999.

must be studied in more depth and on a country-specific basis. This consideration is one of two reasons underlying our choice to work within a country-specific time-series rather than a panel framework of analysis. The second is that working on a country-by-country basis allows us to test for and model non-linear effects in the process of current account adjustment.

Our main findings can be summarised as follows: First, current account balances in the euro area are determined by shifts in domestic and foreign income as well as real exchange rates. Second, there exist important differences across countries regarding the significance of each variable in equilibrium current account determination. Third, adjustment of the current account towards its equilibrium is gradual, with the disequilibrium term being in most countries the main determinant of current account dynamics. Finally, in seven out of ten countries examined, current account adjustment is found to be a non-linear process, with the speed of adjustment being a function of the sign (in six cases) and the size (in one case) of the disequilibrium term. Overall, our findings suggest that both the intertemporal and the TOCA arguments are relevant in explaining diverging current account balances in the euro area.

The remainder of this paper is structured as follows: Section 2 outlines our methodology and discusses our data. Section 3 investigates the determinants of long-run current account determination, while section 4 examines the process of short-run current account adjustment. In particular, section 4.1 presents estimates of linear adjustment models; section 4.2 tests for non-linear effects in the process of current account adjustment and section 4.3 estimates non-linear current account models. Finally, Section 5 offers a discussion and conclusion.

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[6] Real exchange rates are quoted using the indirect quotation convention so that an increase (reduction) in the rates' values denotes a real appreciation (depreciation).

## 2. METHODOLOGY AND DATA

### 2.1. Methodology

We investigate the potential links between the current account to GDP series ($ca$), the domestic and foreign output levels ($y$ and $y^*$ respectively), and the real exchange rate ($q$) within the context of a VAR($k$) model described by equation (1) below:

$$X_t = f + A_1 X_{t-1} + \ldots + A_k X_{t-k} + u_t \quad (1)$$

In (1) $X_t$ is a $(4\times1)$ vector $X_t = [ca_t, q_t, y_t, y^*_t]$ where $q_t$, $y_t$ and $y^*_t$ are expressed in logs; $A_i$ is a $(4\times4)$ matrix of parameters with $i = (1\ldots k)$; $f$ is a constant term; and $u_t$ a $(4\times1)$ matrix of Gaussian errors. Johansen and Juselius (1990) have shown that if $X_t$ consists of $n$ terms integrated of order one, (1) can be reformulated as a linear vector error-correction model (VECM) given by equation (2) below:

$$\Delta X_t = G_1 X_{t-1} + \ldots + G_{k-1} X_{t-k+1} + P X_{t-k} + u_t \quad (2)$$

In (2) $\Delta$ is the first difference operator, $G_i = -(I - A_1 - \ldots - A_i)$, $P = -(I - A_1 - \ldots - A_k)$, $I$ is the identity matrix, and $i = (1, .., k-1)$. If $P$ includes $r$ linearly independent columns where $r<n$ and $n$ is the number of variables in $X_t$, equation (1) convergences to a long-run equilibrium described by $P = \alpha\beta'$, where $\alpha$ and $\beta$ are both $(4\times r)$ matrices. Matrix $\beta$ includes the coefficients defining the long-run equilibrium, and matrix $\alpha$ the coefficients of the speed of adjustment towards the latter. In that case, the VECM in (2) can be re-written as

$$\Delta X_t = G_1 X_{t-1} + \ldots + G_{k-1} X_{t-k+1} + \alpha(\beta' X_{t-k}) + u_t \quad (3)$$

where $\beta X_{t-k}$ yields a maximum of $(n\text{-}1)$ cointegration relationships ensuring that $X_t$ converges to its long-run steady-state solution. The number of cointegrating vectors $r$ is given by the rank of $P$. Johansen and Juselius determine $r$ using the Likelihood Ratio Maximal-Eigenvalue ($l\text{-}max$) and Trace Statistic tests, calculated using the maximum-likelihood estimates of the cointegrating vectors.

In summary, our methodological steps have as follows: First, we test for cointegration and identify any long-run relationships between $ca_t$, $q_t$, $y_t$ and $y^*_t$. Then, we investigate the process of short-run current account adjustment using the linear VECM model (equation 3) and test for non-linear effects. Finally, when such effects are found to exist, we model them formally using suitable non-linear models.

**2.2. Data**

Our main data source is IMF's International Financial Statistics (IFS) Databank available by Datastream. We use data of quarterly frequency, except from Greece for which lack of a quarterly GDP data series extending prior to 1990 obliges us to use annual observations.[7] To calculate the current account to GDP series we multiply the quarterly current account balance series expressed in current US dollars by the average national currency to US dollar series, and then divide by current GDP. The resulting current account-to-GDP series exhibit strong seasonality patterns for which we account through seasonal adjustment.[8] For real exchange rates and domestic national income, we respectively use the IFS' CPI-based real effective exchange rate[9] and the seasonally-

[7] For Ireland and Luxembourg, we found no consistent GDP and real exchange rate series of any frequency prior to 1997 and 1995 respectively. As a result, these countries are excluded from our analysis.

[8] We adjust the series using the Census X11 multiplicative seasonal adjustment method, used by the US Bereau of Census to seasonally adjust publicly released data; the X11 routine is available in EViews.

[9] For Greece the IFS databank does not offer a consistent series for the CPI-based real exchange rate; as a result, we use the unit-labour-cost based real effective exchange rate offered by IFS.

adjusted real GDP volume index series.[10] Finally, to approximate foreign income we use the seasonally-adjusted real GDP volume index series of the G7 area, provided by OECD's Main Economic Indicators databank.

Data availability defines our sample periods as 1975(1)-2005(3) for Austria, Finland and Germany (123 observations); 1977(1)-2005(3) for the Netherlands and Portugal (115 observations); and 1980(1)-2005(3) for Belgium, France, Italy and Spain (103 observations). For Greece, for which annual observations are used, our sample covers 1978-2005 (28 observations). Preliminary data analysis suggests that all series are integrated of order 1.[11] This allows us to investigate the links between $ca_t$, $q_t$, $y_t$ and $y^*_t$ within the Johansen-Juselius cointegration framework described in section 2.1.

## 3. LONG-RUN CURRENT ACCOUNT DETERMINATION

Table 2 presents the results of the cointegration tests calculated for the system of equations in (1). For each VAR we determine $k$ using the Akaike information criterion. Both the $l$-max and the Trace statistic provide evidence of cointegration for all countries. At the 1% level, both statistics suggest the existence of one cointegrating vector ($r$ = 1) for all countries, with the exception of Finland for which we obtain $r$ = 2. At the 5% level, the $l$-max produces identical results, whereas the Trace statistic yields $r$ = 2 for Belgium, Finland, Germany and Greece. Overall, and taking into account our theoretical priors, we accept for all countries the existence of one cointegrating vector.

Table 3 reports the estimated cointegrating vectors normalised on $ca_t$, as well as the p-values of the Johansen-Juselius (1992) Chi-square tests imposing a zero restriction on each variable's coefficient in the beta (long-run coefficients') matrix. As a test of

[10] For some countries, the real GDP volume index series provided by IFS is not seasonally adjusted. In those cases, we have adjusted the series ourselves using the X11 seasonal adjustment filter.
[11] To save space these results are not reported here but are available upon request.

robustness, we also present the cointegrating vectors estimated using the Engle and Granger (EG, 1987) single-equation cointegration methodology.[12] The discussion in section 1 implies that $ca_t$ should be negatively related to $y_t$ and $q_t$ and positively related to $y_t$*.[13] The reported results are largely consistent with our theoretical priors as 27 and 26 out of 30 reported coefficients in the JJ and EG cointegrating vectors respectively present the correct sign; the remaining (wrongly-signed) coefficients are not statistically different from zero.

A number of interesting observations emerge from Table 3. First, the zero-coefficient restriction is rejected for $ca_t$ in all cases. Second, there exist differences with regards to the significance of the rest of the variables across countries. Using the JJ estimates, all three variables enter the cointegrating vector with statistically non-zero coefficients in France, Germany, Portugal and Spain; for Finland, Italy and the Netherlands the only variable with a statistically non-zero coefficient is the real exchange rate; finally for Austria, Belgium and Greece the non-zero coefficients are those of domestic and foreign income. We note, however, that the results for Greece may be affected by the limited number of annual observations; indeed, the EG estimates suggest that the $q$ coefficient is very close to be significant at the 5% level; it is thus possible for the movements of $q$ to affect the equilibrium value of $ca$ in that country too.

Third, with the exception of Finland and Italy, the absolute values of the coefficients of $y_t$ and $y_t$* are in all countries higher than those of $q_t$. This suggests that relative incomes have been playing a more prominent role than real exchange rates in long-run current account determination. This, in turn, implies that the current account deterioration observed in countries such as Greece and Spain following the introduction

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[12] For the EG vectors p-values are available only for the regression's right-hand side variables, i.e. $q$, $y$ and $y$*. The reported p-values refer to Chi-square tests imposing zero restrictions on these variables.
[13] In Table 4, given the indirect definition used for $q$, this corresponds to positive signs for the $q$ and $y$ coefficients and a negative sign for $y$*.

of the Euro is mainly due to higher than EMU average income growth rather than other factors contributing towards real exchange rate appreciation, a finding consistent with Blanchard and Giavazzi's (2002) view.

Having said that, the statistical significance of the *q* terms in Italy, Spain and Portugal (and according to the EG estimates, perhaps Greece), suggest that other factors beyond income growth may explain the current account positions of these countries. With regards to the recent period of current account deterioration (1999-2005), these factors may relate to consistently higher inflation rates relative to EMU average. Ultimately, these positive inflation differentials and the consequent competitiveness losses are most likely reflecting structural rigidities/weaknesses in the real sector of these economies. The role of such weaknesses seem to be even more prominent in the cases of Italy and Portugal, two countries in which income growth has been particularly slow since 1999 and for which the coefficients of *q* reported in Table 3 indicate that the long-run current account effects of exchange rate appreciation, net of income growth, are more pronounced as compared to Greece and Spain.

Finally, Figure 2 presents the estimated cointegrating vectors obtained by both the JJ and EG methodologies, revealing that at the end of our sample periods (2005) the majority of the EMU countries have had current account positions close to their long-run equilibrium values. This suggests that for Italy, Portugal and possibly Greece the current account deterioration experienced since 1999 is equilibrium rather than a transitory phenomenon. On the other hand, Spain, Belgium and France appear to have current account deficits larger than those justified by their long-term determinants. This implies that at least part of the recently observed high current account deficits in those countries is due to slow adjustment towards an otherwise healthier long-run position, an issue to which we turn our attention immediately below.

## 4. SHORT-RUN CURRENT ACCOUNT ADJUSTMENT

### 4.1. Linear current account adjustment models

We now estimate the VECM system given by (3) and report the results in Table 4. This presents the current account ($\Delta ca_t$) equations of the VECM system estimated using the Full Information Maximum Likelihood (FIML) method accompanied by two sets of restrictions' tests.[14] The first set reports the p-values of the Chi-square tests on the joint significance of the lagged values of $\Delta ca_t$, $\Delta q_t$, $\Delta y_t$, and $\Delta y^*_t$ in the reported $\Delta ca_t$ equation. The purpose of these tests is to examine whether shocks to $q_t$, $y_t$, and $y^*_t$ cause any short-run current account noise independent of the systematic correction of any pre-existing disequilibrium captured by the cointegrating vector $cv_t = \beta X_{t-k}$. The second set reports the p-values of the Chi-square tests for the statistical significance of $cv_{t-1}$ in each of the four equations constituting the VECM system. These effectively test the hypothesis of weak exogeneity for each of the variables entering matrix $X$ in system (1). Clearly, for our analysis on current account adjustment to be meaningful, $ca_t$ must not be weakly exogenous.

The results reported in Table 4 suggest that the $cv_{t-1}$ coefficients are correctly signed and statistically significant at the 5% level or lower, the only exceptions being France and Portugal. Note, however, that for these countries too, the $cv_{t-1}$ coefficients are highly significant in the non-linear models presented in section 4.3 below. The size of the error correction coefficients ranges from -0.058 for Portugal to -0.300 for Italy, suggesting slow to moderate speed of adjustment, with some noteworthy differences across countries. For Greece, for which annual data is used, it takes the value of -0.489.

With regards to the rest of the variables, the reported weak exogeneity tests suggest that $q$ is not weakly exogenous only in Austria, Finland, France and Germany; $y$

is weakly exogenous only in Austria, Belgium and Germany; and $y^*_t$ is weakly exogenous in all but two countries. These imply that in the majority of EMU countries, and in particular those presenting high current account deficits in recent years, deviations from equilibrium are corrected through current account and income adjustments. The lack of adjustment of $q$ to the disequilibrium captured by $cv_{t-1}$ can be interpreted as an indication of structural rigidities in the economies of the EMU countries, as the real exchange rate is a measure of a country's international competitiveness. Interestingly, in the EMU's two largest economies, France and Germany (but also in Austria and Finland) $q$ is not weakly exogenous. This indicates that these countries possess a higher degree of adaptability to changing external sector conditions, giving them a competitive edge relative to the rest of the EMU members. Finally, in the majority of the countries the lagged values of $\Delta q_t$, $\Delta y_t$, and $\Delta y^*_t$ are jointly non-significant, implying that the major determinant of current account adjustment is the latter's tendency to return to its steady-state equilibrium.[15]

### 4.2. Tests of non-linear current account adjustment

We now test the hypothesis of non-linear current account adjustment following the procedure proposed by Saikkonnen and Luukkonen (1988), Luukkonen et al (1988), Granger and Teräsvirta (1993) and Teräsvirta (1994). This involves estimating equation (4) below:

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[14] The results are robust to estimating the VECMs using alternative estimation methodologies, including OLS equation-by-equation, 2- and 3-stage instrumental variables. The results are available upon request.

[15] An interesting exception is Greece, where the $\Delta y_{t-1}$ term is statistically significant with the theoretically expected negative sign and a coefficient greater in absolute terms than that of the error correction term. This indicates that income changes create significant short-run noise in the current account of that country. This, in turn, implies that the higher-than-EMU average income growth observed in recent years has contributed to the observed high current account deficits not only through its effect on the latter's equilibrium value but also by means of magnifying the latter's effect through their short-run dynamic effects. A similar argument may apply with regards to the significant real exchange rate appreciation experienced by Italy over the post-Euro period, as the $\Delta q_t$ terms are jointly significant at the 8 per cent level. Finally, for Belgium, Finland and Portugal we obtain statistical significance for the lagged values of $\Delta y_t^*$, which suggests that international economic conditions create current account noise in these countries that is not present in the rest of the EMU.

$$cv_t = g_{00} + \sum_{j=1}^{f} \left( g_{0j} cv_{t-j} + g_{1j} cv_{t-j} cv_{t-d} + g_{2j} cv_{t-j} cv_{t-d}^2 + g_{3j} cv_{t-d} cv_{t-d}^3 \right) +$$

$$+ g_4 \, cv_{t-d}^2 + g_5 \, cv_{t-d}^3 + n_t \tag{4}$$

In (4), $cv_t$ is the JJ cointegrating vector estimated in Table 3; $f$ is the order of the autoregressive parameter $g_i$, determined through the partial autocorrelation function of $\hat{u}_t$;[16] $d$ is the delay parameter of the transition function; and $v_t$ a random error term. Equation (4) is estimated for all plausible values of $d$. Given the quarterly frequency of our data we consider values of $d$ up to 8. For each value of $d$, we test the null of linear current account adjustment, described by $H_0: g_{1j} = g_{2j} = g_{3j} = g_4 = g_5 = 0$, $j = (1,2...f)$, against the alternative of general non-linear adjustment. We do so by employing an LM-type test denoted by $LM^G$. A statistically significant $LM^G$ implies the rejection of the null of linearity with the optimum value of $d$ determined by the highest $LM^G$ score. Provided that $LM^G$ is significant, further tests can be undertaken to determine the exact form of non-linearity (logistic versus quadratic). To that end, we first test the null of linear or non-linear quadratic adjustment, defined as $H_0: g_{3j} = g_5 = 0, j \in (1,2...f)$, against the alternative of logistic non-linear adjustment. We denote the LM score testing this null as $LM^L$. A significant $LM^L$ implies logistic non-linear adjustment and terminates the testing process. If $LM^L$ is insignificant, we compute a third statistic, $LM^Q$, which tests the null of linearity $H_0: g_{1j} = g_{2j} = g_4 = 0 \,|\, g_{3j} = g_5 = 0$, $j \in (1,2...f)$ against the alternative of quadratic non-linear adjustment. Given an insignificant $LM^L$, a significant $LM^Q$ score implies quadratic non-linearity.

[16] Granger and Teräsvirta (1993) and Teräsvirta (1994) advise against choosing $f$ using information criteria, which may induce a downward bias.

We present our results in Table 5. For Spain, linearity is clearly maintained. In six of the remaining countries the $LM^G$ test rejects linearity at the 5 per cent or lower and in the other three cases at the 7% per cent. Out of these nine countries, the $LM^L$ test is statistically significant in six; and in the remaining three the $LM^Q$ test is significant in one. Overall, we conclude that current account adjustment is a linear process in the cases of Italy, Netherlands and Spain; non-linear of logistic type in Austria, Finland, France, Greece, Germany and Portugal; and non-linear of quadratic type in Belgium.

### 4.3. Non-linear current account adjustment models

We now model formally the non-linear adjustment effects found in the previous sub-section. For the countries that display non-linear behaviour of the logistic type we estimate the Logistic Smooth Threshold Error Correction Model (L-STECM). For Belgium, for which quadratic non-linearity has been found, we estimate the Quadratic Logistic Smooth Threshold Error Correction Model (QL-STECM).[17] The L-STECM is given by equations (5) to (8) below:

$$\Delta ca_t = q_t M_{1t} + (1-q_t) M_{2t} + e_t \qquad (5)$$

$$M_{1,t} = a_1 + \sum_{i=1}^{k} b_{1i}\Delta ca_{t-i} + \sum_{i=1}^{k} g_{1i}\Delta q_{t-i} + \sum_{i=1}^{k} d_{1i}\Delta y_{t-i} + \sum_{i=1}^{k} x_{1i}\Delta y^*_{t-i} + z_1 cv_{t-1} + u_{1t} \qquad (6)$$

$$M_{2,t} = a_2 + \sum_{i=1}^{k} b_{2i}\Delta ca_{t-i} + \sum_{i=1}^{k} g_{2i}\Delta q_{t-i} + \sum_{i=1}^{k} d_{2i}\Delta y_{t-i} + \sum_{i=1}^{k} x_{2i}\Delta y^*_{t-i} + z_2\, cv_{t-1} + u_{2t} \qquad (7)$$

$$q_t = pr\left\{t \geq \hat{u}_{t-d}\right\} = 1 - \frac{1}{1+e^{-s[\hat{u}_{t-d}-t]}} \qquad (8)$$

The L-STECM distinguishes between a lower and an upper regime, respectively denoted by $M_1$ and $M_2$ and given by equations (6) and (7). $M_1$ and $M_2$ are linear current account adjustment models similar to those estimated in section 4.1, defined according to whether the transition variable $cv_{t-d}$ takes values below or above a critical threshold $t$. Equation (5) models $\Delta ca_t$ as a weighted average of $M_1$ and $M_2$, with the regime weight $q$ modelled in equation (8) as the probability that $cv_{t-d}$ takes value in the lower regime (below $t$). The delay parameter and speed of transition between the two regimes are denoted by $d$ and $\sigma$ respectively.[18] The difference between the L-STECM and the QL-STECM is that in the latter $M_1$ and $M_2$ respectively describe current account adjustment within an outer and an inner regime, defined by two critical thresholds, $t^L$ and $t^U$. In that case, $q$ is defined as the probability that $cv_{t-d}$ takes values within the inner regime, modelled using the quadratic function given by equation (9) below:

$$q_t = pr \{ t^L \leq (\hat{u})_{t-d} \leq t^U \} = 1 - \frac{1}{1+e^{-s[(\hat{u}_{t-d})_{t-d}-t^L][(\hat{u}_{t-d})_{t-d}-t^U]}} \quad (9)$$

Table 6 presents the results of our non-linear models. Given the quarterly frequency of our data, we set $k$ = 4 in (6) and (7), with the exception of Greece for which we use annual data and set $k$ = 1.[19] We follow a general-to-specific estimation approach and report statistically significant dynamic terms at the 5% level or lower; we also report terms statistically significant at the 10% level in case their inclusion results in a reduction in the model's regression standard error.[20] Table 6 suggests that with the exception of Portugal, for all countries for which an L-STECM model has been estimated, the absolute value of the error correction term is significantly higher in the

[17] For a detailed discussion of these models see van Dijk et al (2002).

[18] In practise, the parameter $\sigma$ is usually estimated very imprecisely as the likelihood function in (8) is very insensitive to this parameter. This is also the case for our estimations. For a detailed discussion on this point, see van Dijk et al. (2002).

[19] In the case of Greece we also experimented by setting the value of $k$ = 2; however, almost certainly due to the limitations imposed by our small sample, we could not obtain model convergence.

[20] The only exception is Greece, for which the small number of available annual observations does not allow the estimation of a non-linear model with all coefficients being well-defined; hence the results presented for that country can only be described as indicative.

lower rather than the upper regime.[21] For Belgium, the speed of adjustment is faster in the outer than in the inner regime. These findings are consistent with a state of the world in which macroeconomic variables adjust more rapidly to high-magnitude current account imbalances as opposed to small ones.[22]

Compared to the linear models presented in Table 4, we obtain a higher number of statistically significant lagged $\Delta y_t$ and $\Delta q_t$ terms, suggesting that the short-run current account noise caused by shifts in national income and real exchange rates may be higher than what suggested in Section 4.1. All but one $\Delta y_t$ terms reported in Table 6 have a negative sign, indicating that income increases result in short-run current account deterioration higher than the long-run one suggested by the cointegrating vectors reported in Table 3. Finally, for Austria, France and Germany the lagged $\Delta q_t$ terms have a positive sign, which is consistent with the presence of J-curve effects.

## 5. DISCUSSION AND CONCLUSION

The euro area countries' current accounts display increasingly diverging patterns during the last few years. This diverging performance has typically been attributed to the different rates of growth in the context of the convergence process. Nevertheless, as some central banks hint to,[23] competitiveness considerations can be relevant as well. To understand the nature and the implications of the current account imbalances in the individual EMU countries we model their determination and equilibrium adjustment process. Our approach captures not only the income convergence process repercussions

[21] For Portugal, our findings may reflect the unusually positive deviations from equilibrium observed during the early years of our sample (see Figure 2).
[22] The only country for which we find adjustment to be higher in the upper rather than the lower regime is Portugal, a fact that may reflect the persistent positive disequilibrium values of the late 1970s and early 1980s. Further support towards this hypothesis is provided by the fact that the critical threshold estimated for Portugal is significantly positive.

but the potential role of the real exchange rate as well. We adopt a "back-to-the-basics" modelling approach, whereupon changes in current account, among other factors, are modelled on real exchange rate shifts. The causal link we emphasize constitutes a standard feature of all mainstream models of international macroeconomics. Surprisingly, however, limited recent evidence exists on this issue in general and not at all (to our knowledge) in the context of the EMU.

Our empirical findings show that a negative relationship exists between the movements of the real effective exchange rates and the current account in the majority of the EMU-member countries after controlling for the role of income growth. This relationship is of non-linear nature for the majority of the euro area countries. The speed with which current accounts adjust towards equilibrium appears to be a function of the sign (and in one case the size) of the disequilibrium term. We find that the two groups of countries with systematically improving/deteriorating current account balances during the post-EMU era correspond to the two groups of countries that experience persistent real exchange rate depreciation/appreciation. Interestingly, these groups largely correspond to those that previous research has identified as respectively belonging and not belonging to a European Optimum Currency Area.

To the extent that full convergence will be achieved in the future, the imbalances explained by the intertemporal approach will be removed. Nevertheless, it emerges that a focus on the real exchange rate can offer further insights given that relative price effects are capable of partially explaining current account developments. Moreover, the full convergence prospect may prove to be a long run process, thus rendering the current account effects of exchange rate shifts to be important for the short-to-medium run. In addition, the catch-up process itself is more likely to intensify the diverging current

[23] See for example the annual reports of the Bank of Greece (2005) and the Bank of Spain (2005).

account performance, as the faster growing countries within the eurozone will experience further real exchange rate appreciation. Finally, our analysis gives rise to some considerations about the implications of current account adjustment within a common currency area. Such considerations may be of relevance for those accession countries that already experience high current account deficits.

To summarize, by analyzing the diverging current positions of individual EMU member states we cover a topic that has been overlooked by the literature on global current account imbalances. At the same time we provide empirical evidence establishing the role of real exchange rates on current account determination, thus validating an important theoretical assumption of open macroeconomics literature for which little recent empirical evidence exists.

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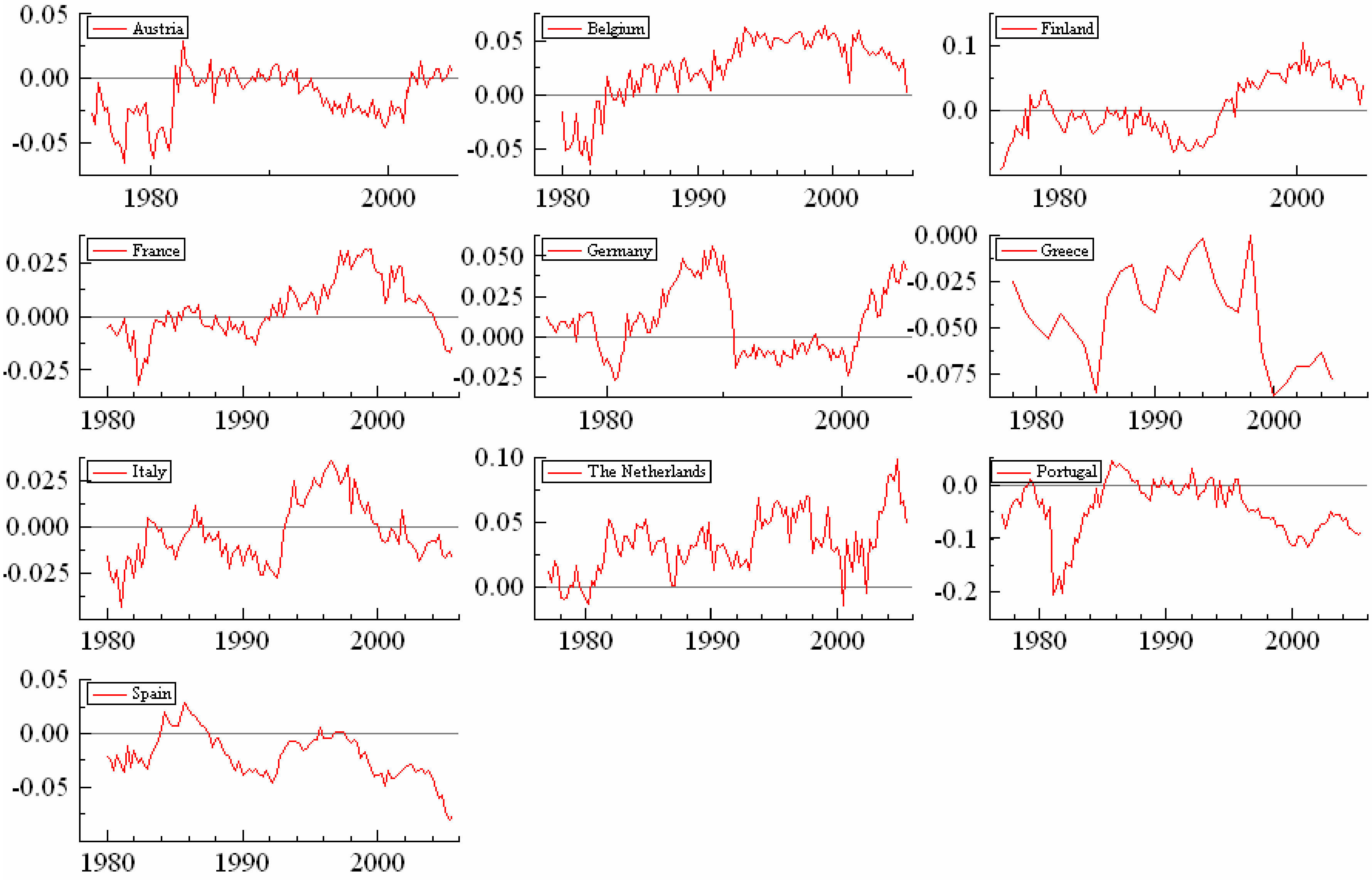
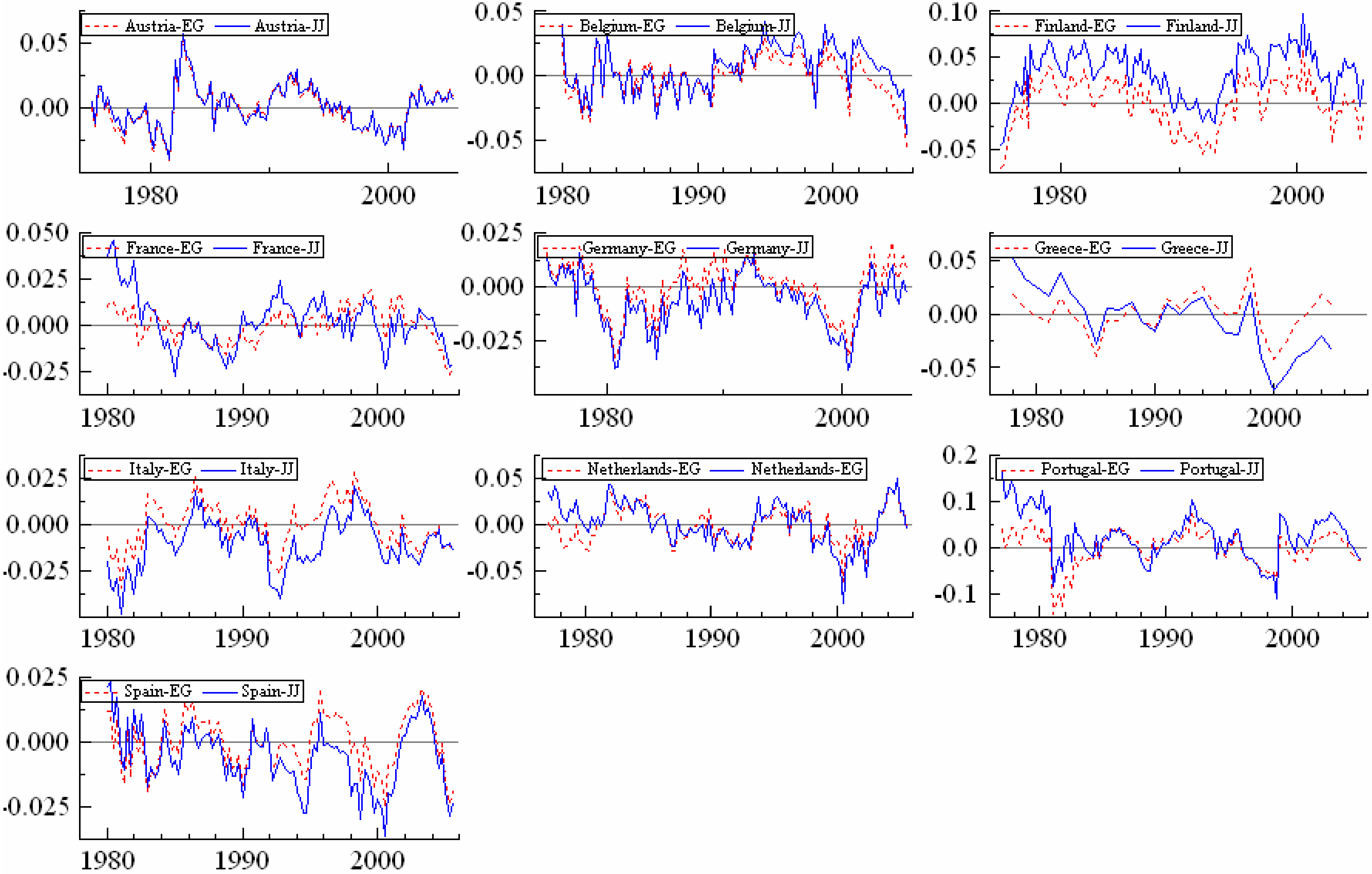
**Figure 1: Current Account Balance (% in GDP, seasonally adjusted)**




*Source: International Financial Statistics*

**Figure 2: Estimated cointegrating vectors**




Note: EG and JJ denote cointegrating vectors estimated using the Engle-Granger (1987) and Johansen-Juselius (1990) methodologies respectively

**Table 1: Output growth, real exchange rates and current account balance in the euro area, 1992-2005**

*Average values*

| | Real GDP growth | | | CPI-based Real Exchange Rate (index, 100 in 2000) | | | Current Account Balance (% in GDP) | | |
|---|---|---|---|---|---|---|---|---|---|
| | (a) | (b) | (c) | (d) | (e) | (f) | (g) | (h) | (i) |
| | 1992-1998 | 1999-2005 | Change (b) – (a) | 1992-1998 | 1999-2005 | Change (f) – (e) | 1992-1998 | 1999-2005 | Change (i) – (h) |
| Austria | 2.2 | 2.6 | 0.4 | 108.2 | 102.8 | -5.4 | -1.9 | -0.9 | 1.0 |
| Belgium | 2.1 | 2.0 | -0.1 | 111.1 | 104.5 | -6.6 | 5.0 | 4.5 | -0.5 |
| Finland | 2.6 | 2.8 | 0.2 | 106.0 | 96.3 | -9.7 | 2.0 | 5.6 | 3.6 |
| France | 1.6 | 2.2 | 0.6 | 111.9 | 104.1 | -7.8 | 1.3 | 0.8 | -0.5 |
| Germany | 1.5 | 1.2 | -0.3 | 115.7 | 104.1 | -11.6 | -0.9 | 1.3 | 2.2 |
| Greece | 1.8 | 4.2 | 2.4 | 100.3 | 107.8 | 7.5 | -1.9 | -7.5 | -5.6 |
| Italy | 1.4 | 1.1 | -0.3 | 88.2 | 106.1 | 17.9 | 1.4 | -0.6 | -2.0 |
| Netherlands | 2.7 | 1.7 | -1.0 | 108.8 | 107.7 | -1.1 | 4.6 | 4.5 | -0.1 |
| Portugal | 2.7 | 1.6 | -1.1 | 113.7 | 105.7 | -8.0 | -2.7 | -8.5 | -5.8 |
| Spain | 2.4 | 3.6 | 1.2 | 107.5 | 105.7 | -1.8 | -1.1 | -4.4 | -3.8 |

*Correlation coefficients (columns in parentheses)*

| | Real GDP growth (c) | Real Exchange Rate (f) | Real Exchange Rate (f) |
|---|---|---|---|
| Real GDP growth (c) | 1 | | |
| Real Exchange Rate (f) | 0.06 | 1 | |
| Current Account (i) | -0.27 | -0.45 | 1 |

*Source: International Financial Statistics*

Note: An increase (reduction) in the value of the real exchange rate denotes a real appreciation (depreciation)

## Table 2: Johansen – Juselius cointegration tests

| | | $l$-Max | | | | Trace | | | |
|---|---|---|---|---|---|---|---|---|---|
| | VAR lag-length ($k$) | $H_0 : r = 0$<br>$H_1 : r = 1$ | $H_0 : r \leq 1$<br>$H_1 : r = 2$ | $H_0 : r \leq 2$<br>$H_1 : r = 3$ | $H_0 : r \leq 3$<br>$H_1 : r = 4$ | $H_0 : r = 0$<br>$H_1 : r = 1$ | $H_0 : r \leq 1$<br>$H_1 : r = 2$ | $H_0 : r \leq 2$<br>$H_1 : r = 3$ | $H_0 : r \leq 3$<br>$H_1 : r = 4$ |
| Austria | 2 | 50.73 [0.00]** | 15.93 [0.31] | 5.92 [0.79] | 4.53 [0.35] | 77.11 [0.00]** | 26.39 [0.33] | 10.45 [0.60] | 4.53 [0.35] |
| Belgium | 2 | 35.68 [0.00]** | 21.16 [0.07]+ | 11.47 [0.23] | 4.83 [0.31] | 73.14 [0.00]** | 37.46 [0.03]* | 16.30 [0.16] | 4.83 [0.31] |
| Finland | 2 | 38.39 [0.00]** | 34.33 [0.00]** | 10.90 [0.27] | 6.67 [0.15] | 90.29 [0.00]** | 51.90 [0.00]** | 15.57 [0.11] | 6.67 [0.15] |
| France | 2 | 41.10 [0.00]** | 15.99 [0.31] | 4.70 [0.90] | 3.78 [0.46] | 65.57 [0.00]** | 24.46 [0.44] | 8.48 [0.78] | 3.78 [0.46] |
| Germany | 2 | 36.66 [0.00]** | 19.12 [0.13] | 15.43 [0.06]+ | 2.04 [0.77] | 73.25 [0.00]** | 36.59 [0.03]* | 17.47 [0.12] | 2.04 [0.77] |
| Greece | 1 | 49.89 [0.00]** | 20.58 [0.09]+ | 10.69 [0.29] | 5.19 [0.27] | 86.35 [0.00]** | 36.46 [0.04]* | 15.88 [0.18] | 5.19 [0.27] |
| Italy | 2 | 39.10 [0.00]** | 18.07 [0.18] | 6.49 [0.73] | 2.77 [0.63] | 66.45 [0.00]** | 27.34 [0.28] | 9.27 [0.71] | 2.77 [0.63] |
| Netherlands | 2 | 32.60 [0.01]** | 18.37 [0.17] | 10.06 [0.34] | 4.29 [0.38] | 65.32 [0.00]** | 32.72 [0.09]+ | 14.36 [0.27] | 4.29 [0.38] |
| Portugal | 2 | 45.63 [0.00]** | 11.64 [0.69] | 8.22 [0.53] | 2.69 [0.65] | 68.19 [0.00]** | 22.55 [0.56] | 10.91 [0.56] | 2.69 [0.65] |
| Spain | 3 | 30.08 [0.03]* | 17.28 [0.23] | 4.65 [0.91] | 4.39 [0.37] | 56.40 [0.03]* | 26.32 [0.33] | 9.04 [0.73] | 4.39 [0.37] |

Notes: +,* ,** respectively denote statistical significance at the 10, 5 and 1 per cent level; the numbers in square brackets denote p-values calculated using the small-sample correction of critical values provided by PcGive. The lag structure of the estimated VAR systems ($k$) has been determined using the Akaike information criterion. The reported $k$ parameters are consistent with those suggested by the Schwarz and the Hannan-Quinn information criteria provided by PcGive.

## Table 3: Cointegrating vectors

| | *ca* | *a* | *q* | *y* | *y** |
|---|---|---|---|---|---|
| Austria – JJ | 1.000<br>[0.00]** | -0.549 | 0.264<br>[0.62] | 1.072<br>[0.05]* | -1.058<br>[0.05]* |
| Austria – EG | 1.000 | -0.407 | 0.196<br>[0.15] | 0.842<br>[0.00]** | -0.831<br>[0.00]** |
| Belgium – JJ | 1.000<br>[0.02]* | -0.277 | 0.018<br>[0.14] | 1.904<br>[0.00]** | -1.798<br>[0.00]** |
| Belgium – EG | 1.000 | 0.111 | -0.028<br>[0.77] | 1.302<br>[0.00]** | -1.350<br>[0.00]** |
| Finland – JJ | 1.000<br>[0.00]** | -0.048 | 0.257<br>[0.00]** | -0.008<br>[0.46] | -0.229<br>[0.47] |
| Finland – EG | 1.000 | 0.022 | 0.243<br>[0.00]** | -0.048<br>[0.85] | -0.229<br>[0.34] |
| France – JJ | 1.000<br>[0.07]+ | -2.149 | 0.719<br>[0.00]** | 1.802<br>[0.00]** | -1.458<br>[0.00]** |
| France – EG | 1.000 | -0.629 | 0.167<br>[0.05]* | 0.902<br>[0.00]** | -0.760<br>[0.00]** |
| Germany – JJ | 1.000<br>[0.00]** | -0.192 | 0.027<br>[0.00]** | 1.196<br>[0.00]** | -1.134<br>[0.00]** |
| Germany – EG | 1.000 | -0.203 | 0.037<br>[0.65] | 1.016<br>[0.00]** | -0.956<br>[0.00]** |
| Greece-JJ | 1.000<br>[0.00]** | -0.124 | 0.169<br>[0.73] | 0.459<br>[0.04]* | -0.558<br>[0.01]** |
| Greece – EG | 1.000 | -0.637 | 0.210<br>[0.07]+ | 0.480<br>[0.00]** | -0.349<br>[0.00]** |
| Italy – JJ | 1.000<br>[0.00]** | -0.273 | 0.251<br>[0.00]** | -0.315<br>[0.12] | 0.195<br>[0.18] |
| Italy – EG | 1.000 | -0.224 | 0.173<br>[0.00]** | -0.223<br>[0.95] | -0.040<br>[0.87] |
| Netherlands – JJ | 1.000<br>[0.00]** | -0.308 | 0.354<br>[0.08]+ | 0.139<br>[0.69] | -0.374<br>[0.35] |
| Netherlands – EG | 1.000 | 0.492 | -0.124<br>[0.46] | 0.143<br>[0.15] | -0.284<br>[0.00]** |
| Portugal – JJ | 1.000<br>[0.05]* | 0.134 | 0.985<br>[0.00]** | 1.699<br>[0.00]** | -2.692<br>[0.00]** |
| Portugal - EG | 1.000 | 0.117 | 0.374<br>[0.06]+ | 0.934<br>[0.00]** | -1.321<br>[0.00]** |
| Spain – JJ | 1.000<br>[0.00]** | -0.196 | 0.285<br>[0.05]* | 1.295<br>[0.00]** | -1.475<br>[0.00]** |
| Spain –EG | 1.000 | 0.022 | 0.227<br>[0.00]** | 0.974<br>[0.00]** | -1.050<br>[0.00]** |

NOTES: JJ and EG respectively denote cointegrating vectors using the Johansen-Juselius and Engle and Granger cointegration methodologies; +, *, ** respectively denote statistical significance at the 10, 5 and 1 level significance. Numbers in square brackets denote p-values of Chi-square tests imposing zero restrictions on the coefficients of the beta matrix. For the reported EG cointegrating vectors the Chi-Square tests have been calculated using the Andrews (1991) autocorrelation- and heteroscedasticity-consistent correction.

# Table 4: VECM current-account adjustment equations

| | Austria | Belgium | Finland | France | Germany | Greece | Italy | Netherlands | Portugal | Spain |
|---|---|---|---|---|---|---|---|---|---|---|
| *Estimated $\Delta ca_t$ equations* | | | | | | | | | | |
| $a$ | 0.005 (0.002)* | 0.005 (0.002)* | 0.004 (0.004) | 0.004 (0.004) | -0.001 (0.001) | 0.007 (0.012) | -0.002 (0.002) | 0.002 (0.003) | 0.004 (0.005) | 0.000 (0.059) |
| $\Delta ca_{t-1}$ | -0256 (0.101)* | -0.417 (0.110)** | -0.494 (0.099)** | -0.289 (0.109)** | -0.084 (0.107) | -0.041 (0.255) | -0.266 (0.104)* | -0.314 (0.108)** | -0.170 (0.097)+ | -0.100 (0.120) |
| $\Delta ca_{t-2}$ | -0.098 (0.093) | -0.152 (0.095) | -0.139 (0.093) | -0.074 (0.108) | 0.044 (0.101) | | -0.029 (0.097) | -0.021 (0.105) | 0.072 (0.097) | 0.097 (0.110) |
| $\Delta ca_{t-3}$ | | | | | | | | | | 0.197 (0.107)+ |
| $\Delta q_{t-1}$ | 0.007 (0.208) | -0.228 (0.208) | -0.053 (0.046) | -0.138 (0.092) | -0.099 (0.082) | -0.021 (0.230) | 0.070 (0.050) | -0.175 (0.224) | -0.395 (0.253) | -0.099 (0.098) |
| $\Delta q_{t-2}$ | 0.162 (0.207) | 0.082 (0.210) | -0.021 (0.046) | 0.066 (0.096) | -0.015 (0.082) | | 0.092 (0.050)+ | 0.242 (0.223) | 0.019 (0.277) | -0.098 (0.095) |
| $\Delta q_{t-3}$ | | | | | | | | | | 0.116 (0.096) |
| $\Delta y_{t-1}$ | -0.476 (0.247) | 0.312 (0.310) | -0.081 (0.292) | -0.113 (0.318) | -0.160 (0.160) | -1.207 (0.610)+ | -0.319 (0.300) | -0.236 (0.179) | -0.081 (0.207) | -0.315 (0.198) |
| $\Delta y_{t-2}$ | -0.159 (0.246) | 0.071 (0.289) | -0.233 (0.289) | -0.551 (0.308)+ | -0.079 (0.155) | | -0.133 (0.300) | 0.052 (0.172) | -0.124 (0.188) | -0.328 (0.203) |
| $\Delta y_{t-3}$ | | | | | | | | | | 0.159 (0.203) |
| $\Delta y^*_{t-1}$ | -0.518 (0.537) | 1.584 (0.818)+ | 2.034 (0.774)* | 0.076 (0.402) | 0.463 (0.433) | -0.040 (1.034) | -0.087 (0.512) | -0.658 (0.801) | -3.079 (1.262) | -0.443 (0.487) |
| $\Delta y^*_{t-2}$ | -0.440 (0.541) | -2.966 (0.776)** | -0.880 (0.779) | 0.353 (0.408) | -0.105 (0.445) | | 0.219 (0.457) | 0.338 (0.800) | 2.555 (1.262) | 0.736 (0.500) |
| $\Delta y^*_{t-3}$ | | | | | | | | | | -0.241 (0.464) |
| $cv_{t-1}$ | -0.204 (0.067)** | -0.232 (0.091)* | -0.150 (0.071)* | -0.058 (0.556) | -0.141 (0.070)* | -0.490 (0.230)* | -0.300 (0.073)** | -0.159 (0.071)* | -0.077 (0.051) | -0.182 (0.082)* |
| *$\Delta ca_t$ equation restrictions (p-values): Joint zero restrictions on coefficients of:* | | | | | | | | | | |
| $\Delta ca_{t-i}$ | 0.04* | 0.00** | 0.00** | 0.03 | 0.59 | 0.87 | 0.03* | 0.00** | 0.10+ | 0.16 |
| $\Delta q_{t-i}$ | 0.73 | 0.55 | 0.47 | 0.31 | 0.44 | 0.93 | 0.08+ | 0.47 | 0.29 | 0.34 |
| $\Delta y_{t-i}$ | 0.15 | 0.57 | 0.71 | 0.17 | 0.58 | 0.05* | 0.51 | 0.40 | 0.76 | 0.15 |
| $\Delta y^*_{t-i}$ | 0.27 | 0.00** | 0.03* | 0.61 | 0.54 | 0.97 | 0.89 | 0.71 | 0.03* | 0.51 |
| *VECM system restrictions (p-values): Zero restriction on the coefficient of $cv_{t-1}$ in the equation of:* | | | | | | | | | | |
| $\Delta ca_t$ | 0.00** | 0.01* | 0.04* | 0.30 | 0.05* | 0.03* | 0.00** | 0.03* | 0.13 | 0.03* |
| $\Delta q_t$ | 0.02* | 0.55 | 0.04* | 0.00** | 0.00** | 0.98 | 0.36 | 0.19 | 0.64 | 0.81 |
| $\Delta y_t$ | 0.94 | 0.41 | 0.00** | 0.06+ | 0.63 | 0.00** | 0.06+ | 0.00** | 0.00** | 0.05* |
| $\Delta y_t^*$ | 0.03* | 0.62 | 0.39 | 0.38 | 0.21 | 0.95 | 0.02* | 0.30 | 0.91 | 0.03* |

NOTES: +, *, ** respectively denote statistical significance at the 10, 5 and 1 per cent level respectively; numbers in parentheses denote standard errors; numbers in square brackets denote p-values; standard errors for Austria and France have been estimated using White's (1980) heteroscedasticity-consistent methodology; standard errors for Belgium, Germany, Netherlands, Portugal and Spain have been calculated using Andrews (1991) autocorrelation- and heteroscedasticity-consistent methodology. Estimating standard errors without these corrections does not affect the qualitative nature of our statistical inference.

**Table 5: Tests for non-linear current account adjustment**

| | *f* | *d* | $LM^{G}$ | $LM^{L}$ | $LM^{Q}$ |
|---|---|---|---|---|---|
| Austria | 4 | 2 | 3.95 [0.00]** | 2.55 [0.04]* | N/A |
| Belgium | 2 | 2 | 2.32 [0.04]* | 1.41 [0.24] | 4.12 [0.01]** |
| Finland | 2 | 8 | 3.72 [0.00]** | 3.58 [0.02]* | N/A |
| France | 1 | 6 | 2.91 [0.02]* | 4.35 [0.02]* | N/A |
| Germany | 1 | 4 | 5.12 [0.00]** | 3.47 [0.03]* | N/A |
| Greece | 1 | 2 | 2.43 [0.07]+ | 3.93 [0.04]* | N/A |
| Italy | 4 | 4 | 1.73 [0.07]+ | 0.40 [0.85] | 0.66 [0.71] |
| Netherlands | 2 | 7 | 1.87 [0.07]+ | 0.16 [0.92] | 1.30 [0.27] |
| Portugal | 2 | 1 | 3.82 [0.00]** | 2.97 [0.04]* | NA |
| Spain | 4 | 1 | 1.13 [0.35] | 1.26 [0.29] | 1.34 [0.24] |

NOTES: +, *, ** denote significance at the 10%, 5% and 1% level respectively.

**Table 6: Non-linear current account adjustment models**

| | Austria | Belgium | Finland | France | Germany | Greece | Portugal |
|---|---|---|---|---|---|---|---|
| $M_1$ | | | | | | | |
| $a_1$ | -0.004 (0.003) | 0.008 (0.004)* | 0.005 (0.003)+ | -0.006 (0.002)** | 0.000 (0.003) | 0.013 (0.024) | 0.008 (0.004)* |
| $\Delta ca_{t-1}$ | | -0.798 (0.249)** | | | | | -0.266 (0.100)* |
| $\Delta ca_{t-2}$ | | -0.365 (0.163)* | | | | | |
| $\Delta ca_{t-3}$ | | | | | | | 0.245 (0.068)** |
| $\Delta q_{t-1}$ | 1.050 (0.437)* | -0.977 (0.460)* | | | | | |
| $\Delta q_{t-2}$ | 1.480 (0.410)** | | | | 0.485 (0.216)* | | |
| $\Delta q_{t-3}$ | | | | 0.673 (0.290)* | | | |
| $\Delta y_{t-1}$ | | | 1.262 (0.056)* | | | -2.544 (1.951) | |
| $\Delta y^*_{t-1}$ | | | | | | | -2.146 (1.037)* |
| $cv_{t-1}$ | -0.781 (0.137)** | -0.541 (0.203)** | -0.598 (0.149)** | -0.646 (0.221)** | -0.316 (0.130)* | -0.778 (0.402)+ | -0.102 (0.056)* |
| $M_2$ | | | | | | | |
| $a_2$ | 0.001 (0.001) | 0.008 (0.002)* | 0.011 (0.004)* | 0.002 (0.001)+ | 0.002 (0.002) | 0.007 (0.006) | 0.021 (0.008)* |
| $\Delta ca_{t-1}$ | -0.221 (0.115)+ | -0.418 (0.116)** | -0.391 (0.114)** | -0.233 (0.104)* | | | |
| $\Delta ca_{t-1}$ | 0.149 (0.083)+ | | | | | | |
| $\Delta ca_{t-4}$ | | | | | | | -0.510 (0.249)* |
| $\Delta q_{t-1}$ | | | | | | | -1.061 (0.440)* |
| $\Delta y_{t-1}$ | | | | | | -0.993 (0.627) | |
| $\Delta y_{t-2}$ | | | | -0.621 (0.300)* | | | |
| $\Delta y_{t-3}$ | | | | | | | -0.334 (0.198)+ |
| $\Delta y_{t-4}$ | -0.592 (0.259)* | | | | | | |
| $\Delta y^*_{t-2}$ | | -2.319 (0.755)** | | | | | |
| $\Delta y^*_{t-3}$ | | | | | -0.915 (0.452)* | | |
| $cv_{t-1}$ | -0.145 (0.078)+ | -0.169 (0.116) | -0.262 (0.098)** | -0.119 (0.055)* | -0.167 (0.083)* | -0.552 (0.264)* | -0.226 (0.089)* |
| $t$ | -0.010 (0.001)** | | 0.014 (0.002)** | -0.011 (0.001)** | -0.015 (0.002)** | -0.019 (0.005)** | 0.059 (0.006)** |
| $t^U$ | | 0.021 (0.001)** | | | | | |
| $t^L$ | | -0.023 (0.011)* | | | | | |

+, *, ** respectively denote statistical significance at the 10, 5 and 1 per cent level respectively; numbers in parentheses denote standard errors.