Long-run and Short-run Effects of Exchange Rate Movements for Major EU Countries: Cointegration and Error-Correction Modeling

Manuel Cantavella-Jordá
Universitat Jaume I and Instituto de Economía Internacional

Celestino Suárez-Burguet
Universitat Jaume I and Instituto de Economía Internacional

Abstract

This paper examines the long-run and short-run effects of depreciation/ devaluation for major European Union countries (Germany, France, the United Kingdom, and Italy) over the 1975-1997 period. The approach is based on cointegration techniques proposed by Johansen [1988] and uses quarterly data. The empirical results indicate the existence of a positive relationship between the exchange rate and the trade balance for each country although long-run effects are rather moderate. According to the short-run analysis, there is a finding of a J-curve for Italy and the United Kingdom. The costs of relinquishing individual exchange rates may be rather small for major EU countries. (JEL Classification: F31, F41)

I. Introduction

Economists, very often, do not reach a common agreement as to whether exchange rate changes can be useful instruments to improve a deteriorated trade balance. Empirical research on the conventional view that exchange rate changes have a positive impact on the trade balance has provided different results ⟨Miles [1979], Gylfasson and Rissager [1984], Bahmani-Oskooee [1985, 1991], Rose and Yellen [1989], Rose [1991], Mahdhavi and Sohravian [1993], Bahmani-Oskooee and Alse [1994], Arize [1994]⟩. Governments are not always willing to devalue. Some problems are believed to be caused by devaluation such as stagflation, negative real income effects and inflation. Furthermore, it is widely accepted that the external debt burden increases whenever this is measured in foreign currency and trade accounts do not improve even when accompanied by appropriate monetary or fiscal policies perhaps as a result of low price elasticities. These problems were reflected to a great extent in the 1973 and 1979 oil crises along with the debt crises of the 1980s. Nevertheless, failing to devalue in time, when such an action is necessary, might also have negative consequences thereafter. All in all, the argument about the effectiveness of a devaluation is still questionable.

This argument may be quite relevant in the European Union (EU) framework. Monetary integration has not yet been achieved. Most EU members enjoy a "quasi-floating" exchange rate system since the obligatory bilateral marginal intervention limits were widened to ±15%, in August 1993.[1] This means that exchange rates can fluctuate and therefore they may depreciate or appreciate within the new broader bands. These movements in the exchange rate along with changes in relative prices may affect a country's competitiveness and consequently its trade account.

In the absence of wage flexibility and labour mobility within the EU, a negative demand shock on a member country's products might be corrected by devaluating its currency. This action would increase foreign demand by reducing relative prices and, thereby, alleviating the shock. A devaluation

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1. The exception is the Dutch guilder and the German mark which maintain their bilateral bands of 2.25%.

in this case would produce the expected outcomes. On the other hand, if devaluation is ineffective in correcting economic shocks, European policy makers would have to rely on other instruments to ensure adjustment of their trade balance. However, the costs for each member, of relinquishing individual exchange rates, can be inferred from the degree of effectiveness that a devaluation or depreciation might have. Therefore, finding out whether a devaluation/depreciation of exchange rates leads to a reduction of a trade balance deficit or not, and to what extent, can be very interesting.

It is our purpose to examine econometrically the relationship between the trade balance and exchange rates for the EU countries over the 1975-1997 period, both in the long and short run and evaluate its effects on a possible monetary union. A great part of existing empirical studies on trade relationships have run regressions with data in levels. However, given the possibility that most of the underlying series had non-stationary residuals, the simple application of ordinary least squares (OLS) methodology might have led to spurious regression results. Cointegration deals with this phenomenon and aims at correcting it. In line with this methodology, several studies have been carried out focusing on the long-run relationship between the trade balance and exchange rates ⟨Rose [1991], Bahmani-Oskooee [1991], Bahmani-Oskooee and Alse [1994], Arize [1994]⟩ and the short-run link between the two variables ⟨Rose and Yellen [1989], Rose [1990], Bahmani-Oskooee and Alse [1994], Mahdhavi and Sohravian [1993]⟩. There are some mixed results as far as long-run effects are concerned; for example, Rose [1991] finds no evidence of a cointegrating relationship whereas Bahmani-Oskooe [1991] and Arize [1994] do. There is more consensus on short-run effects as to negative findings of J-curves. Nevertheless, the utilisation of different sample periods, countries, data frequency, and even cointegration techniques make it difficult to directly compare final results. The approach used in this paper is based on Johansen [1988] multivariate cointegration technique. Our study involves then, cointegration as well as error correction models so that not only long-run (static) effects will be inferred but also short-run (dynamic) effects will be focused on, emphasising, at the same time, their corresponding elasticities.

Section II sets the analysis of the relationship between the trade balance and exchange rates for major EU countries in the imperfect substitutes

model as termed by Goldstein and Khan [1985]. Section III presents the uni-variate analysis of trade variables. Section IV copes with the long-run analy-sis and its trade elasticities. The short-run analysis and its trade elasticities are undertaken in section V. Finally, some concluding remarks are provided in section VI.

II. Model Specification and Methodology

In order to examine the effects of the exchange rates on the trade balance an extensive part of empirical literature on foreign trade equations has worked with export and import demand equations.[2] The objective was to investigate whether the Marshall-Lerner condition would hold or not. Our approach, however, deviates from checking the elasticities condition and essentially concentrates on a reduced-form model expressing the trade bal-ance as a function of supposedly exogenous variables: exchange rates, domestic income, and foreign income, that is,

$$TB_i = f(q_i, Y_i, Y^*) \quad i: \text{holds for each country} \tag{1}$$

This is the equation of interest where,

TB_i : trade balance for country i
q_i : real effective exchange rate for country i
Y_i : domestic income for country i
Y^* : foreign income

This non-structural approach is obtained from the combination of export demand and import demand equations and it allows us to directly examine the impact, if any, of exchange rates on the trade balance, taking also into account the effects of domestic and foreign income so that relevant vari-ables in the model are not omitted.[3]

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2. Examples of studies that have used this method include Houthakker and Magee [1969], Khan [1974], Warner and Kreinin [1983], Gylfason and Risager [1984], Bah-mani-Oskooee [1986], among others.

3. A group of studies have also relied upon estimating some reduced-form models such as Miles [1979], Bahmani Oskooee [1985], Himarios [1989]. The findings of those papers are cast into doubt because of possible spurious problems when dealing with

In econometric terms, equation (1) becomes,

$$\log TB_t = \alpha_0 + \alpha_1 \log q_t + \alpha_2 \log Y_t + \alpha_3 \log Y_t^* + u_t$$

$$t = 1975Q1 \ldots\ldots 1997Q1 \tag{2}$$

Note: 1975Q1 stands for the first quarter of 1975; analogously for 1997Q1.

where all the variables are expressed in natural logarithms so that elasticities can also be interpreted; α's are the parameters of the model that have to be estimated; TB is a ratio of exports over imports; q is the real effective exchange rate; Y is the GDP proxy variable for domestic income; Y^* is the OECD GDP proxy variable for foreign income; u is the error term which represents omitted factors left out by the deterministic part of the model (see Appendix A for more details about variables). The model is applied to four major EU countries, Germany, France, the United Kingdom, and Italy.

The problem with equation (2) could arise as a consequence of the spurious regression phenomenon first described by Granger and Newbold [1974]. This is due to non-stationary tendencies in time series data. The mean, variance, and autocorrelation of the series are in general non-constant through time, the coefficient of determination (R^2) may simply capture correlated trends and low Durbin-Watson (DW) statistics may reflect non-stationary residuals. In this case, as Phillips [1986] argues, OLS estimates do not converge to constants and the standard t and F statistics do not even have the limiting distributions. In view of this concern, one has to investigate whether a series is stationary in levels, $I(0)$, or stationary in differences, $I(1)$, $I(2)$, ... $I(n)$, in order to apply the correct methodology, avoiding any spurious inferences.[4]

Cointegration becomes an issue when one has to deal with non-stationary data. If TB and q are, for example, $I(1)$ variables and therefore non-stationary in levels, one cannot simply regress ΔTB (stationary in first differences) on Δq (stationary in first differences) to avoid problems of non-stationarity

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non-stationary variables. Recent studies that account for that problem and that also use reduced-form models are, for instance, Rose [1991], Bahmani-Oskooee [1991], Mahdhavi and Sohravian [1993], Bahmani-Oskooee and Alse [1994], and Arize [1994].

4. An $I(n)$ variable means that the original series has been differenced n times to become stationary (n is called *order of integration*).

because this way valuable long-term information between the two variables would be lost. It is important then to deal with levels. Equation (2) is expressed in levels and it reflects the long-run relationship among the trade balance, real effective exchange rate, domestic income, and foreign income. Thus, if these variables are cointegrated some linear combination of them will have a lower order of integration.[5]

Another question of interest concerns the short-run dynamic response of the net trade balance to movements in the real exchange rate, taking also into account the effects of real income on the trade balance. The answer to this question is directly obtainable from the error correction model (ECM) derived from the cointegrating vector that we obtain in the previous analysis. As was mentioned before, if the variables in consideration are cointegrated there is a long-run equilibrium relationship. However, it may be possible that in the short run there exists disequilibrium. The cointegrating vector $\hat{u}_t$ (derived from equation (2)), also called the disequilibrium term in the ECM, can be used to tie the short-run behaviour of the endogenous variable (*TB*) to its long-run value. Thus, a group of cointegrated variables can be represented in an ECM. This concept, first used by Sargan [1964] and then popularised by Engle and Granger [1987], corrects for any disequilibrium.

Nevertheless, our approach uses the Johansen procedure [1988] which has got the advantage of being able to model the multivariate nature of the estimation problem. This means that an ECM could incorporate one, two, or even more cointegrating vectors just depending on the number of existing cointegrating vectors. The ECM equation turns out to be in terms of differenced variables with the error-correction component measured in terms of level variables. From an economic point of view, one would be expecting some of these variables to be exogenous such as in the case of q. From a statistical point of view, we want to see whether the data support the property of exogeneity. This is the reason why, since there are four variables (*TB*, q, Y, Y^*) in our model, the ECM derived from the cointegrating vector(s) should be a four-by-four model, since all the variables are treated as poten-

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5. Engle and Granger [1987] state that two series Y_t and X_t of the same order of integration $I(\underline{d})$ are cointegrated of order $(\underline{d}, \underline{b})$ if there exists a vector $(1, -\beta)'$ such that the combination $u_t = Y_t - \beta X_t$ is $I(\underline{d} - \underline{b})$, $b > 0$.

tially endogenous. The long-run equilibria imposed by cointegrating vectors enter the disequilibrium terms in the dynamic form. Thus the choice of these cointegrating vectors is very important. The advantage of this model following the Johansen approach is that the dynamics are much richer and they allow us to investigate the elasticities and the short-run responses more easily.

III. Univariate Analysis of Trade Variables

Testing stationarity of times series and even more important, the type of trends involved, (whether deterministic or stochastic), leads us to the implementation of the econometric model using the appropriate methodology. Thus, the econometric modeling depends on the nature of our trade variables. Many macroeconomic time series display trends when observations are plotted against time. Although most of them appear to be $I(1)$ or difference stationary processes (DSPs) as Nelson and Plosser [1982] demonstrate, it is fundamental to the cointegration analysis that we distinguish between a DSP which contains a stochastic trend,

$$y_t = \beta + y_{t-1} + \varepsilon_t \quad \text{where } \varepsilon_t \sim I(0) \tag{3}$$

and a trend stationary process (TSP) or an $I(0)$ + trend process with a deterministic trend,

$$y_t = \alpha + \beta t + \varepsilon_t \quad \text{where } \varepsilon_t \sim I(0) \tag{4}$$

Stochastic trended (random walk) processes, equation (3), are called difference stationary because differencing should yield an uncorrelated stationary error process. On the other hand, deterministic trended processes, equation (4), are also called trend stationary because, although the first difference is stationary, it is not appropriate to difference them to achieve stationarity.[6]

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6. In order to see this, take first differences, (denoted by Δ) of equation (4), as if the data generation process were a DSP, then,

$$\Delta y_t = y_t - y_{t-1} = (\alpha + \beta t + \varepsilon_t) - [\alpha + \beta(t-1) + \varepsilon_{t-1}] = \beta + \varepsilon_t - \varepsilon_{t-1}$$

Thus, differencing a variable that is stationary around a trend, causes negative autocorrelation in the error term.

Before testing for this distinction, it should be first established the non-stationarity nature of our trade variables. The univariate analysis is carried out through the implementation of the Dickey and Fuller (DF) [1979, 1981] tests and, in particular, the augmented Dickey-Fuller (ADF) test. Aside from telling us about the existence of unit roots in the variables, it recognises whether the data generation process (DGP) is a DSP or a TSP. This analysis is completed with the performance of the tests proposed by Durbin and Hausman ⟨see Choi [1992] and Appendix B for more details⟩ which have better power properties in finite samples, especially when the model includes an intercept and a linear time trend.

Table 1 reports, in columns 2 and 3, standard DF unit root tests results. Column 2 shows that all series are non-stationary after first differencing since the null hypothesis of non-stationarity cannot be rejected for each of them. Considering, then, that all series are stationary in first differences, the data could be generated by either $I(1)$ or $I(0)$ + trend processes. Accord-

Table 1
Unit Root Tests (quarterly data)

Variables	ADF statistics $I(1)$ vs. $I(0)$	ADF statistics $I(2)$ vs. $I(1)$	ADF statistics DSP vs. TSP	DH statistics DSP vs. TSP
GETB	–2.18 (2)	–12.11 (0)	–2.16 (2)	12.32
GEXR	–0.045 (0)	–7.96 (0)	–2.10 (0)	9.87
GEY	–0.40 (4)	–3.19 (3)	–2.29 (4)	38.78
FRTB	–1.96 (4)	–15.01 (0)	–3.14 (4)	33.52
FRXR	–1.64 (0)	–9.21 (0)	–3.46 (3)	26.56
FRY	–1.78 (1)	–7.35 (0)	–2.87 (2)	9.12
UKTB	–1.71(3)	–9.50 (2)	–1.72 (3)	52.87
UKXR	–1.69 (0)	–7.60 (0)	–1.94 (0)	7.06
UKY	–0.45 (2)	–5.01 (1)	–1.87 (2)	15.13
ITTB	–2.49 (1)	–10.07 (0)	–2.30 (2)	51.32
ITXR	–1.75 (0)	–8.67 (0)	–1.97 (0)	16.45
ITY	–2.39 (1)	–5.41 (0)	–2.66 (1)	12.62
Y^*	–1.14 (1)	–6.55 (0)	–2.01 (1)	8.24
Critical Values (5%):	–2.88	–2.88	–3.44	76.14

Note: Numbers in brackets are number of lags used in the ADF test in order to remove serial correlation in the reiduals.

ing to the ADF results (column 4), trade time series are shown to be differ-ence stationary processes. Only France real effective exchange rate rejects the null hypothesis of a difference stationary process in favour of a trend stationary process. However, it is well known the low power of DF tests, especially in the presence of a deterministic trend, thus, Durbin-Hausman (DH) tests have been also applied. These indicate that absolutely all vari-ables are DSPs or $I(1)$, that is, they contain a stochastic and not a determin-istic trend. The next step is , then, to examine the order of integration. Col-umn 3 reflects the fact that the null hypothesis of being $I(2)$ variables is rejected in favour of $I(1)$ variables. Thus, the data generation process exami-nation suggests that the use of cointegration techniques will be suitable to proceed with the long-run analysis.

IV. Long-Run Specification

Cointegration provides the appropriate tools to work with non-stationary variables, and particularly with $I(1)$ variables. Aside from this, the technique also allows for a useful and meaningful link between the long- and short-run approach to econometric modeling as we shall see. The next step is then to specify our multivariate model and apply the Johansen [1988] methodology. This approach estimates long-run or cointegration relationships between non-stationary variables using a maximum likelihood procedure which tests for the number of cointegrating relationships and estimates de parameters of those cointegrating relationships. The general vector autoregression is,

$$y_t = a + \sum_{i=1}^{3} \phi_i Q_{it} + \sum_{i=1}^{k} \pi_i y_{t-i} + \varepsilon_t \tag{5}$$

where either y_t and y_{t-i} include the logarithms of the four trade variables (TB, q, Y, Y^*); a is the intercept; Q_{it} represents the deterministic seasonal (quarterly) dummies, and ε_t is a disturbance term independently and identi-cally distributed with zero mean and constant variance.[7]

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7. In Table 2 only the trade balance as a function of q, Y, and Y^* is shown. The construc-tio of error correction models (available from the authors) for each of the variables except for foreign income which was considered exogenous beforehand, indicated that both real effective exchange rates and domestic income were exogenous.

The Johansen estimation method is based on the error correction representation of the general vector autoregression. Thus, equation (5) can be rewritten as an ECM of the form:

$$\Delta y_t = a + \sum_{i=1}^{3} \phi_1 Q_{it} + \sum_{i=1}^{k-1} \Gamma_i \Delta y_{t-i} + \Pi y_{t-k} + \varepsilon_t \qquad (6)$$

As Δy_t and Δy_{t-1} are $I(0)$ and y_{t-k} variables are $I(1)$, equation (6) will be balanced if left-hand side and right hand-side have the same degree of integration. This will occur if $\Pi = 0$, in which case the y variables are not cointegrated or if the parameters of Π are such that Πy_{t-k} is also $I(0)$. The latter case applies when the y variables are cointegrated. The rank r (number of cointegrating vectors) of matrix Π should be less than the number of variables in y_t. Matrix Π can be decomposed as $\Pi = \alpha \beta'$, where β are the parameters in the cointegrating vector and α measures the strength of the cointegrating vectors in the ECMs.

The results of cointegration tests are reported in Table 2. The two test statistics, maximum eigenvalue (λMAX) and trace, are presented, where λMAX tests for at most r cointegrating vectors against the alternative of exactly $r+1$ cointegrating relationships, while Trace tests for at most r cointegrating vectors against the alternative of at least $r+1$ vectors.

A number of lags for each of the variables and countries have been included in order to capture the short-run dynamics of the model.[8] Up to four lags have been tried for each equation, which should provide a sufficient representation of the process generating the data given that we are dealing with quarterly time series. Every country seemed to show satisfactory results with four lags in its corresponding vector autoregression (VAR) according to serial correlation and normality diagnostics.[9]

From the economic point of view the existence of two cointegrating vectors

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8. In general, economic results are quite sensible with few lags.

9. We should remember that if we increase too much the value k (lag length) we may be capturing economic information that we do not wish, at this stage, to be specified in the model. Moreover, increasing too much the value of k (lag length) may cause overparameterisation of the model which affects the estimates of cointegrating vectors making it difficult to interpret economically the cointegration relations. On the other hand, setting too small a value of the lag length may distort the size of the tests.

Table 2
Johansen Maximum Likelihood Cointegration Tests (1975Q1-1997Q1)

r: number of cointegrating vectors (null hypothesis)	Germany		France		UK		Italy	
	λMAX	Trace	λMAX	Trace	λMAX	Trace	λMAX	Trace
$r=0$	37.51*	76.9*	40.14*	75.84*	44.25*	67.24*	33.99*	65.05*
$r\leq1$	20.59	39.18*	20.31	35.70*	14.05	22.98	24.20*	31.06
$r\leq2$	15.01	18.60	10.59	15.39	6.87	8.92	6.46	6.87
$r\leq3$	3.59	3.59	4.79	4.79	2.06	2.06	0.41	0.41

Parameter Estimates (normalised)

	Germany		France		UK	Italy	
	Vector 1	Vector 2	Vector 1	Vector 2	Vector	Vector 1	Vector 2
TB	−1	−1	−1	−1	−1	−1	−1
q (−)	−0.09	2.05	−1.22	−5.41	−0.14	−0.77	1.30
Y (−)	−1.75	−4.47	−1.34	−10.65	−1.46	−3.57	−7.95
Y^*(+)	1.38	5.40	1.15	9.43	0.76	3.18	6.18

Notes: λMAX and Trace are the likelihood ratio statistics for the number of cointegrating vectors. Estimation has been performed with Microfit 3.0.
* Indicates significance at 5 percent level; critical values are based on Osterwald-Lenum [1992].
In the second panel, *TB* denotes trade balance, *q*, real effective exchange rate, *Y*, domestic income, and Y^*, foreign income. In brackets are the expected signs for *q*, *Y*, and Y^*. Parameter estimates express different elasticities.

may appear somewhat confusing. This feature derives from the fact that a number of variables may be tied together in the long run. According to Muscatelli and Hurn [1992], applied economists should use only that cointegrating vector that makes economic sense. This means that signs and magnitudes of coefficients should agree with the expected ones. From the econometric point of view one should choose the first of the cointegrating vectors as the relationship is stronger. In other words, the first vector contains residuals that are closer to white noise than the second cointegrating vector. Bearing in mind both aspects, vector 1 should probably be chosen for every country.[10]

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10. It should be born in mind that for a number of reasons, already pointed out, the

The cointegration results provide evidence that the real effective exchange rate affects the trade balance in the long run for each of the four EU countries in the expected direction (Table 2). While exchange rates have a predictable and systematic impact on trade, price elasticities tend to be low, in most instances below unity.[11] These low estimates, with probably the exception of France, indicate that an external adjustment in the face of movements in exchange rates are quite difficult for major EU countries. The costs of foregoing the exchange rate as an instrument of economic policy may be inferred from these results. In general, these costs would be rather moderate and depending on the country in question. In the case of Germany, for example, they would be practically insignificant, whereas for France they would be more important.

Income effects, on the other hand, play an important role in the attainment of cointegration relationships. In spite of the tendency for imports to rise more rapidly than exports which may entail significant trade imbalances, according to the different income elasticities (domestic and foreign), one cannot conclude that this circumstance appears as an external restriction to growth because this would mean that international trade is restrictive for a country's growth. The magnitude of the estimates for domestic income variables implies that strong domestic activity may provide a relevant expansionary impulse to other countries, thus confirming a potential engine role of the EU.

In order to examine the robustness of the above results one is referred to Table 3 which reports the outcomes of the parameter restriction tests for exchange rates and income variables. In general, all trade variables are significant with the exception of the Italian real effective exchange rate. However, the omission of this variable in the model prevents obtaining a cointegrat-

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empirical results obtained in this paper should not be considered directly comparable to outcomes derived from other studies.

11. The inclusion of oil products in these countries' trade balance, as well as pass-through and pricing behaviour of foreign firms considerations, may account for obtaining low elasticities in general. Moreover, as De Grauwe [1989] argues, in the current period of flexible exchange rates where there is high exchange rate variability, the sensitivity of trade flows to given movements of the exchange rate declines as a result of the uncertainty that these exchange rate movements will last.

Table 3
Tests of Parameter Restrictions on *TB*, *q*, *Y* and *Y**

Country	Chi-squared test statistic (*TB*)	Chi-squared test statistic (*q*)	Chi-squared test statistic (*Y*)	Chi-squared test statistic (*Y**)
Germany	8.74	20.23	24.16	20.16
France	6.47	12.97	14.50	9.32
The UK	7.40	8.57	14.26	12.32
Italy	7.62	3.16*	20.11	24.17

Note: * is not significant at 5 percent level

ing vector (tested but not shown) between the trade balance and income variables. Therefore, for this link it is important to include this variable.

A long-run analysis including these four countries is also carried out by Bahmani-Oskooee and Alse [1994]. Their results differ from the ones obtained here. Besides the reasons previously mentioned about reaching different outcomes, there is one more in this case that should be important. Bahmani-Oskooee and Alse [1994] uses just two variables in their model, trade balance and real effective exchange rates. They cannot find a statistical impact in the long run between those two variables. It is possible that the omission of relevant variables in the model such as domestic and foreign income may have affected the final results. As a matter of fact, in this paper it has been also tested (although not shown) the possibility of obtaining a cointegrating relationship between the trade balance and the real effective exchange rate. Only for the UK there was cointegration between those two variables but the real effective exchange rate was not even significant.

V. Short-Run Analysis: Error Correction Modeling

Obtaining long-run estimates from the Johansen procedure has been the first step to estimating the complete model. The next step is to specify the short-run VAR model in error-correction form, that is to include, not only those variables which contain the short-run information, but also the cointegration relationships (long-run information) previously calculated. In order

to assess the short-run effects of movements in the exchange rate before they achieve the long-run equilibrium we estimate ECMs. It should be noted that the specification of a VAR (4) on quarterly data for all the countries allows us to obtain, quite satisfactorily, the information derived from the dynamics of equivalent ECMs. These require that all terms in equation (6) are stationary so that one can apply standard OLS and interpret t-ratios. The number of lags finally included in the ECMs is consistent with those added to the cointegration analysis.

Table 4 displays the final trade balance ECMs. It should be read as indicated below the second panel.[12] This table also contains diagnostic tests (serial correlation, functional form, normality, and heteroskedasticity) that allow us to measure to what extent these models are valid. The numbers in brackets, in the third panel, express the significance level at which the corresponding null hypothesis would be rejected.

ECMs for each of the four countries (Germany, France, the UK, and Italy) validate all the diagnostic tests shown in the second panel of Table 4. Practically in all cases, the error correction term denoted by Z_{-1} carries a significant coefficient. This provides further evidence on the long-run effects of exchange rates and income variables on the trade balance. The adjustment of the trade balance toward the long-run equilibrium is quite gradual for France and Germany and quicker for the UK and Italy. Between 11 and 39 per cent of the disequilibrium is corrected during the first quarter. It is worth noting that the trade balance for the UK and Italy worsens (the sign of the exchange rate is positive, that is, contrary to the one in the long-run analysis) before getting better. This outcome is consistent with the J-curve phenomenon. The exchange rate impact on the trade balance is immediate for both countries (first quarter). Over time, the trade balance improves (medium and long run) for the UK and Italy probably as new contracts reflect not only the initial value effect but also the volume effect which now can be adjusted to the new situation. Germany and France do not experience any significant short-run effect of the real effective exchange rate. In the case of Germany it is likely that relative prices are not significant due, among other factors, to the general perception of foreign importers about

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12. Notation of symbols are below the second panel in Table 2.

Table 4

Final Estimate of Trade Balance ECMs

	Germany	France	UK	Italy
c	0.037	0.017	−0.015	−0.025
	(3.45)	(2.35)	(−0.66)	(−2.57)
ΔTB_{-1}	0.07	−0.18	−0.17	0.14
	(2.57)	(−1.85)	(−1.75)	(2.64)
ΔTB_{-2}	0.32			
	(3.09)			
ΔTB_{-4}		0.19		
		(2.07)		
Δq_{-1}			0.34	0.33
			(2.89)	(2.21)
ΔY_{-1}		−0.38		
		(−2.11)		
ΔY_{-3}	−0.17	−0.65	−0.47	
	(−2.66)	(−2.67)	(−2.41)	
ΔY^{*}_{-4}	0.28			0.47
	(2.34)			(3.38)
Z_{-1}	−0.21	−0.11	−0.38	−0.39
	(−2.74)	(−2.61)	(−4.25)	(−4.82)
R^2	0.63	0.54	0.59	0.66
F	9.44	14.05	16.01	30.42
$s.e.$	0.028	0.026	0.042	0.073

Note: ΔTB_{-1} represents the trade balance in differences, lagged one period; Δq_{-1} denotes the real effective exchange rate, lagged one period; ΔY_{-1} expresses domestic income, lagged one period; ΔY^{*}_{-1} represents foreign income, lagged one period; Z_{-1} is the disequilibrium term with 1 lag; R^2 is the coefficient of determination; F is the joint test; all significant at 5 percent level; $s.e.$ is the standard error of the regression. Numbers in parentheses beneath each coefficient are t-ratios (at 5 and 10 per cent significance level). Only significant variables at 5 and 10 percent levels are reported, except for the constants which are all shown.

$s.c.$	(0.929)	(0.556)	(0.866)	(0.639)
$f.f.$	(0.934)	(0.698)	(0.148)	(0.371)
$n.$	(0.678)	(0.502)	(0.699)	(0.663)
$h.$	(0.212)	(0.104)	(0.240)	(0.643)

$s.c.$ stands for serial correlation; $f.f.$ for functional form; $n.$ for normality; $h.$ for heteroskedasticity. The numbers between parentheses are the percentages at which the null hypothesis is rejected.

the high quality of German products.[13]

Income variables are a major determinant of the behaviour of the German and France trade balance in the short run. Foreign-income elasticity is larger than domestic-income elasticity following the same pattern given in the long-run analysis for Germany whereas for France only domestic income seems to be significant in the short-run dynamics. As to the UK and Italy, changes in domestic and foreign income respectively along exchange rate movements are determinants of their trade balance adjustment in the short run.

VI. Conclusions

In this paper, it has been found evidence that real effective exchange rates have a positive impact on the trade balance in the long run for major EU countries. This result sheds more light on the long-run statistical relationship between those two variables, at least in the Community context. The existence of that link is sustained by the effects that income variables have on the trade balance. The outcomes of this analysis in support of a long-run equilibrium relationship are consistent with the imperfect substitutes model confirming the validity of this model for economic policy implementation purposes.

Low long-run price elasticities indicate that a substantial change in relative prices should be made in order to considerably improve trade accounts. Costs of relinquishing individual exchange rates in the monetary union may, in general, be rather moderate given the estimates of price elasticities obtained. In any case, adjustment to external shocks will always have to be a real adjustment since one of the Maastricht commitments for EU countries, to reach the monetary union in 1999, is the stability of exchange rates, abandoning them as instruments of economic policy.

Short-run dynamic results provided by final ECMs along with those of cointegrating vectors for the UK and Italy, confirm the existence of a J-curve for these two countries. Income variables are major determinants in the

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13. In any case, one might want to more closely examine the reasons why there is no J-curve for Germany and France. For this purpose different hypotheses should be tested such as the response of import volume to import prices, the response of import prices to the exchange rate, and the response of export value to the exchange rate.

movements of the trade balance for Germany and France in the short run, being domestic income the variable that worsens their trade accounts according to the magnitude and sign of their elasticities.

Appendix

A. Data Sources and Definitions

The empirical work considers quarterly data and deals with four major European Union countries: Germany, France, the United Kingdom, and Italy. Quarterly data cover the first quarter of 1975 up to the first quarter of 1997.

The trade balance (*TB*) is defined on an aggregate basis, taking into account solely visible goods traded around the world. The construction of this variable has been carried out by use of a ratio of exports over imports. Thus, a rise of this ratio indicates an improvement of the trade balance and the contrary holds for a fall. It is measured in 1990 US dollars. Since we are dealing with global trade flows, the use of a global real effective exchange rate as one of the explanatory variables seems to be appropriate as the latter represents a summary measure of the value of a currency to the value of others, competitors and/or trading partners. The real effective exchange rate (q) is a weighted index that combines the exchange rates between a currency in particular and the currencies of seventeen other industrial countries (partner and/or competitor countries). It is adjusted for relative movements in labor unit costs and expressed on 1990 year base. Defined as units of foreign currency per unit of domestic currency, an appreciation of the real effective exchange rate is reflected by an increase of the index and a depreciation by a decrease of the index. Gross domestic product (GDP), in 1990 domestic currency units is used as a proxy for domestic income (Y). OECD GDP (PPP's of 1990) is employed as a proxy for foreign income (Y^*).

Trade balance and real effective exchange rate data are drawn from *International Financial Statistics*, CD-ROM, January 1998, published by the International Monetary Fund. GDPs are collected from various *OECD Quarterly National Accounts* bulletins.

B. Durbin Hausman Tests

Choi [1992] proposes Durbin-Hausman [1954, 1978] tests for a unit root based on the traditional parameterisation from which Dickey and Fuller [1979, 1981] derive their own tests. For this purpose, the OLS estimator and an instrumental variable (current variable) are used. Unlike ARMA models which usually work with lagged variables as instruments Choi [1992] employs y_t to instrument y_{t-1}. Thus, y_t is not a real instrument but a pseudo instrument.

The maintained model is the same as for DF tests,

$$y_t = \alpha + \beta y_{t-1} + \gamma t + \varepsilon_t$$

The null hypothesis is that of $\beta = 1$ against the alternative of $\beta < 1$, that is $I(1)$ against $I(0)$+trend. The test statistic is,

$$DH = \frac{(b_{iv} - b)^2}{\text{est } V(b)}$$

where b denotes the OLS estimate of β and b_{iv} indicates the pseudo-instrumental variables estimate of β using y_t to instrument y_{t-1}.

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