Real Exchange Rate Effects on the Balance of Trade: Cointegration and the Marshall-Lerner Condition.

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Abstract.

A typical finding in the empirical literature is that import and export demand elasticities are rather low, and that the Marshall-Lerner (ML) condition does not hold. However, despite the evidence against the ML condition, the consensus is that real devaluations do improve the balance of trade, though after a lag because of J-curve effects. The aim of this paper is to try and measure the effects of the real exchange rate on the balance of payments using structural cointegrating Vector Autoregressive Distributed Lag (VARDL) models for domestic and foreign output, the balance of trade and the real exchange rate. Small systems are estimated for eight OECD countries to investigate long-run cointegration. Generalised impulse response functions are calculated to investigate the response to shocks. Unlike conventional orthogonalised impulse functions these are invariant to ordering. The VARDL estimates suggest a single cointegrating vector and that output and the real exchange rate can be treated as weakly exogenous for the parameters of the balance of payment equation. This allows estimation using a single equation ARDL. Although there is considerable heterogeneity, overall the results suggest that the Marshall-Lerner condition is satisfied in the long run with there being J-curve effects in the short run.

1 Introduction

The "elasticity approach", also known as the "imperfect substitutes" model, is still the most commonly used in balance of trade analysis (see Goldstein and Khan 1985, for a clear exposition). A key issue is the extent to which trade flows are responsive to relative price changes, more specifically whether or not a devaluation improves the trade balance, i.e. whether the well-known Marshall-Lerner (ML) condition holds[1]. The classic analysis is Houthakker and Magee (1969), who reported that the estimates of the price elasticities were not precise enough "to permit a definite stand in the 'elasticity pessimism' controversy"[2]. Several studies followed (see, e.g., Khan 1974, Goldstein and Khan 1978, Wilson and Takacs 1979, Warner and Kreinin 1983, Bahmani-Oskooee 1986, Krugman and Baldwin 1987), all of which essentially employed least-squares methods to estimate price elasticities using import and export equations, though not settling this controversy. Marquez (1990) used band spectrum analysis as well and estimated bilateral trade elasticities, which she then aggregated to obtain multilateral ones. She also reported mixed results, the evidence from the spectral estimates being less supportive of the ML condition, and concluded that, although her multilateral estimates were broadly consistent with those from aggregate studies, valuable additional information could be obtained by using bilateral data.

Subsequent advances in econometrics made it clear that the evidence accumulated until then was possibly "spurious", as non-stationarities in the data may not have been properly handled. Recent studies invariably use a cointegration approach to estimate long-run trade elasticities, both in OECD and non-OECD countries, and are generally more supportive of the ML condition (Rose 1991 being a notable exception). However, they differ in the type of equation estimated, and the relative price measure and techniques used. The majority of the literature estimates import and export demand functions, and then checks whether or not the sum of the absolute value of the respective price elasticities implies that the ML condition holds (see, e.g., Bahmane-Oskooee 1998, Bahmani-Oskooee and Niroomand 1998, Caporale and Chui 1999). The terms of trade are the appropriate price variable in this case. However, such a method requires identifying the structural parameters, a difficulty which can be avoided by estimating instead a reduced-form equation for the balance of trade, a method which allows direct testing for the response of trade flows to relative price movements without any knowledge of the "deep" parameters (see, e.g., Rose 1991, and Lee and Chinn 1998) - this is the approach also taken in

[1]Note that some recent literature has reexamined the ML condition from a theoretical point of view. In particular, Ide and Takayama (1991) have argued that the relevant stability condition should be based on the Marshallian output adjustment process rather than the Walrasian price adjustment one, and that this condition implies that the conventional results hold under constant returns to scale, but not under variable returns to scale. Furthermore, Bughin (1996) has shown that the original ML condition need to be modified if pricing-to-market (PTM) strategies are adopted because of imperfect product competition.

[2]Estimates of the elasticities obtained from disaggregate data tend to be larger than those obtained from aggregate data, but in this paper we are concerned with the aggregate results.

the present paper. The real exchange rate (as opposed to the terms of trade) is then used.

Concerning the estimation and testing techniques, some papers rely mainly on the Engle and Granger (1987) approach (see, e.g., Rose 1991, though he also uses non-parametric techniques), others adopt Johansen's (1991) maximum likelihood framework (see Bahmani-Oskooee 1998, and Bahmani-Oskooee and Niroomand 1998), others, such as Caporale and Chui 1999, carry out Johansen's (1991) cointegration tests, but then use single-equation estimators, which have more desirable small sample properties (see Caporale and Pittis 1999).

In the present paper, we estimate a cointegrating VAR, which is preferable to a structural VAR, such as the one estimated by Chenn and Lin (1998), for the reasons spelt out below (see also Pesaran and Smith 1998). Further, we use a single-equation method, namely the ARDL approach, which outperforms other estimators in small samples (see Pesaran and Shin 1995).

2 Literature

Theories of real exchange rate determination focusing on supply-side factors, such as productivity differentials (see Balassa, 1964, and Samuelson, 1964), depend on a number of crucial assumptions, namely that the country is small, there is international capital mobility, both factors of production (capital and labour) are perfectly mobile across sectors of the economy, and there are constant returns to scale in these factors. If any of these conditions are not met, then there is scope for demand factors to play a role. For instance, Froot and Rogoff (1991) find that government spending differentials have highly persistent effects, though factor mobility causes them to die out in the long run. A similar conclusion is reached by De Gregorio et al (1994a,b), and De Gregorio and Wolf (1994), the latter also reporting that terms of trade shocks are important empirically. A number of papers are based on the idea that there is a relationship between cumulated current account deficits and long-run real exchange rate depreciation. Obstfeld and Rogoff (1995), for instance, find a significant correlation between trade-weighted real exchange rate changes and changes in net foreign asset positions, and Bayoumi et al (1994) show that it matters whether the driving factor is monetary or fiscal policy. Various theoretical explanations have been offered for such a correlation. Krugman (1990), e.g., argues that it is through transfers of wealth across countries with different consumption patterns that current account changes lead to real exchange rate movements. [3]

The net export component of the current account is usually modelled as a function of the real exchange rate (or competitiveness), and other exogenous factors. By imposing the balance of payments equilibrium condition that changes in reserves be zero and using the definition of the current account as the sum of trade balance and net overseas income, one can obtain an expression for the real exchange rate as a function of exogenous variables, net overseas income

[3] Froot and Rogoff (1994) and Rogoff (1996) provide an extensive discussion of the various possible reasons for deviations from PPP.

and the capital account. In the long run net capital flows are expected to be zero, but deviations from PPP will still occur if either net overseas income or the exogenous variables have non-zero values. PPP would be satisfied if trade elasticities were infinitely large, which is not a typical empirical finding. On the contrary, they are usually estimated to be rather low, with the Marshall-Lerner condition not holding. Consequently, changes in either supply or demand factors could result in movements in the real exchange rate away from its absolute PPP level. Even if the exogenous variables have zero values, violations of PPP will still be observed if net foreign assets have accumulated during the dynamic adjustment (a point emphasised in the portfolio balance approach), or if there is high capital mobility resulting in further capital flows (see MacDonald, 1995, for more details).

It is possible, however, to reconcile the standard approach to trade balance determination (which suggests low price elasticities and implies that differences in income elasticities and/or in growth rates will result in secular trends in real exchange rates) with PPP within a "new trade theory" framework based on increasing returns and product differentiation (see Krugman, 1989). This can be shown to be consistent with the empirical regularity known as the "45-degree rule", i.e. the fact that faster growing economies have high income elasticities of demand for their exports but lower import elasticities, which implies that faster growth can be observed without any marked secular trend in real exchange rates (see Caporale and Chui, 1999 for some evidence).

The econometric issues which arise when testing for PPP are analysed by Boyd and Smith (1999). They emphasise that, although PPP does appear to hold in the long run (in the developing countries, in their case), there are major difficulties in measuring the speed of convergence in cross-country panels, unless data for very long time periods are available and the structural parameters are constant, both across countries and over time. In the case of industrialised countries, it is plausible to assume that the supply of exports, as well as the supply of imports, is infinitely elastic. Under these circumstances, one obtains the standard Marshall-Lerner condition according to which a devaluation will improve the balance of payments if and only if the sum of the demand elasticities exceeds one. However, the available econometric evidence suggests that in fact these elasticities are rather low, with the implication that the ML condition is not satisfied. Still, devaluations are normally associated with an improvement in the trade balance, even though trade volumes typically adjust to price changes with a lag, thereby generating the phenomenon known as the J-curve.

The vast majority majority of papers focus alternatively on analysing the determinants of the equilibrium real exchange rate (or deviations from it, in order to establish to what extent PPP holds), as, e.g., in Baxter (1994) or De Gregorio and Wolf (1994), or the "sustainability" of the current account, as in Sheffrin and Woo (1990). In the absence of a clearly defined link between the real exchange rate and external equilibrium, which is implied by theory, the validity of the empirical results reported in such studies is questionable. An encompassing framework focusing on the macroeconomic determinants of the real exchange rate is introduced by Stein (1994). His NATREX (Natural Real

Exchange Rate) model includes both goods market equilibrium and portfolio balance, and endogenises capital stock and foreign debt. Such a model can be used to rationalise the simultaneous improvement in the current account balance and real appreciation observed in the long run in countries such as Japan. This could not be explained in terms of the J-curve, namely the lagged response of exports and imports to real exchange rate changes, since these are by definition short-run effects (see Song, 1997).One of the few other papers attempting to provide a simultaneous explanation of exchange rate and current account behaviour for the G7 is a recent study by Lee and Chinn (1998), which we discuss in more detail below.

3 The Model

The balance of trade is usually measured by the difference between exports and imports, here it will prove more convenient to work with the ratio of exports to imports, since in a logarithmic model this gives the Marshall-Lerner condition exactly rather than as an approximation. The results can be transformed back to the difference if required. The ratio of nominal exports to nominal imports, B, is given by the ratio of the volume of exports, X, multiplied by domestic prices, P, to the volume of imports M, multiplied by foreign prices, P^*, and the nominal spot exchange rate S:

$$B_t = (P_t X_t)/(P_t^* S_t M_t)$$

or using lower case letters for logarithms:

$$b_t = x_t - m_t - (s_t - p_t + p_t^*) = x_t - m_t - e_t \tag{1}$$

where $e_t = (s_t - p_t + p_t^*)$ is the real exchange rate. Notice that this is defined in terms of the real exchange rate (using general price indices) rather than the terms of trade (using import and export price indices).

Long run import and export demand are given by:

$$\begin{aligned} x_t &= \alpha_x + \beta^* y_t^* + \eta_x e_t + \gamma_x t \\ m_t &= \alpha_m + \beta y_t - \eta_m e_t + \gamma_m t. \end{aligned} \tag{2}$$

The trends capture terms of trade effects (e.g. for primary product producers), unmeasured quality improvements or policy measures such as trend liberalisation. The long run balance of trade is

$$b_t = (\alpha_x - \alpha_m) + \beta^* y_t^* - \beta y_t + (\eta_x + \eta_m - 1)e_t + (\gamma_x - \gamma_m)t. \tag{3}$$

The coefficient on e_t gives the familiar Marshall-Lerner condition for a devaluation (increase in e) to improve the balance of payments. Solvency requires $b_t = 0$ in the long run, PPP requires $e_t = e$, a constant, and in steady state

$y_t = y_0 + gt$ and $y_t^* = y_0^* + g^* t$, then

$$\begin{aligned} 0 &= (\{\alpha_x + \beta^* y_0^*\} - \{\alpha_m + \beta y_0\}) & (4) \\ & +(\eta_x + \eta_m - 1)e & (5) \\ & +(\{\gamma_x + \beta^* g^*\} - \{\gamma_m + \beta g\})t. & (6) \end{aligned}$$

Taking first differences and rearranging gives the familiar Kaldor condition for the steady state growth rate.

$$g = \frac{\beta^*}{\beta} g^* + \frac{\gamma_x - \gamma_m}{\beta} \tag{7}$$

The growth rate of a country will be a linear function of the world growth rate (if there are no exogenous trends, a proportional function) with *slope* given by the ratio of export to import income elasticities of demand. Empirically this relationship between relative growth rates and relative elasticities works quite well.

The capital account has been completely ignored in this story - it could be introduced by including a real UIP condition with a risk premium dependent on the balance of trade:

$$\Delta e = (i_t - i_t^*) + \delta b_t \tag{8}$$

where i and i^* are real interest rates, then the long-run equilibrium condition would be

$$b_t = -(i_t - i_t^*)/\delta \tag{9}$$

and the real interest rate differential would appear in (4).

This is a demand side story assuming infinite elasticities of supply at home and abroad. With finite elasticities of supply y_t would appear in the export equation and y_t^* in the import equation as proxies for capacity utilisation or other supply side constraints. Empirically, identifying these supply side effects has not proved easy, partly because of a lack of suitable exogenous supply side instruments. When equations like (2) have been estimated, the general result is that the measured import and export demand elasticities have been rather low and the Marshall-Lerner condition has not held. In earlier times, this 'elasticity pessimism' prompted a move towards the absorption approach. Despite the empirical evidence against the ML condition, the consensus is that real devaluations do improve the balance of trade, though after a lag because of J-curve effects.

We can write the long-run relationship (3) as

$$b_t = \alpha + \beta^* y_t^* - \beta y_t + \eta e_t + \gamma t \tag{10}$$

where $a = (\alpha_x - \alpha_m)$; $\eta = (\eta_x + \eta_m - 1)$; and define the deviation from long-run equilibrium as

$$z_t = \alpha + \beta^* y_t^* - \beta y_t + \eta e_t + \gamma t - b_t$$

Although the number of cointegrating vectors among these four variables is an empirical matter which will be investigated, for the moment treat this as the only cointegrating vector. Define $x_t^* = (b_t, e_t, y_t, y_t^*)'$, we can then model the dynamic adjustment by a pth order Vector Autoregression (VAR), which can be written as a Vector Error Correction Model (VECM) in the cointegrating relationship z_t:

$$\Delta x_t^* = \mu + \alpha z_{t-1} + \sum_{i=1}^{p} \Gamma_i \Delta x_{t-i}^* + u_t \tag{11}$$

where μ includes unrestricted deterministic elements and α is a 4×1 vector. This VECM can be regarded as the reduced form for a structural cointegrating Vector Autoregressive Distributed Lag (VARDL) model, which includes exogenous variables. In particular, regard world output as exogenous and define $x_t = (b_t, e_t, y_t)'$, then the VARDL takes the form:

$$\Delta x_t = \mu + \alpha z_{t-1} + \sum_{i=1}^{p} \Gamma_i \Delta x_{t-i} + \sum_{i=0}^{p} \delta_i \Delta y_{t-i}^* + u_t. \tag{12}$$

Writing (12) out explicitly, ignoring the short-run dynamics, we get

$$\begin{aligned} \Delta b_t &= \alpha_b z_{t-1} + \mu_{10} + u_{1t} \\ \Delta e_t &= \alpha_e z_{t-1} + \mu_{20} + u_{2t} \\ \Delta y_t &= \alpha_y z_{t-1} + \mu_{30} + u_{3t}. \end{aligned}$$

If in addition $\alpha_e = 0$ and $\alpha_y = 0$, then we can also treat e_t and y_t as exogenous and condition on them to give a single equation Autoregressive Distributed Lag (ARDL) model. .

This framework allows us to trace through the dynamic effects of deviations from the long-run solvency condition on output, real exchange rate and the balance of trade, within a manageable system. If the effects were primarily on the balance of payments α_b should be negative and large in absolute value; and if they worked through deflationary effects on output, α_y should be large. This framework assumes that b_t, y_t, y_t^* and e_t are each I(1) and that there is a single cointegrating vector given by the solvency condition. The integration properties of these three variables have been extensively analysed, though in different literatures. The stationarity of the balance of payments b_t has been extensively discussed in the literature on the Feldstein-Horioka puzzle, e.g. Coakley, Kulassi and Smith (1996), the stationarity of the real exchange rate e_t in the Purchasing Power Parity literature, e.g. Boyd and Smith (1999), and the stationarity of deviations of domestic from foreign output, $y_t - y_t^*$ in the convergence literature, e.g. Lee, Pesaran and Smith (1997). In all three cases, while there are good theoretical reasons to expect them to be bounded in the very long run and they may be I(0) on very long spans of data, they look very much like I(1) variables over the type of time-series sample that will be used here.

The exercise in this paper is closely related to Lee and Chin (1998). They use a 'structural VAR' rather than a cointegrating VAR (identifying the model by assumptions about the nature of the shocks[4]); start from a different theoretical framework (open-economy, sticky-price, IS-LM rather than trade balance); and only look at the real exchange rate and current account, treating productivity shocks (which we proxy by output) as unobserved. Despite the different frameworks, the reduced forms of their model and ours are identical; our equation (10) corresponds to their equation (7), though the interpretation of the parameters is different. The big difference comes from their assumption that the real exchange rate is non-stationary while the current account (as a share of GDP) is stationary, thus they estimate a bivariate VAR in the change in the real exchange rate and the current account. This raises both theoretical and empirical problems. Firstly their equation (7) (which relates the balance of payments, assumed I(0), to the real exchange rate, assumed I(1), and differential productivity shocks) only balances if differential productivity shocks are I(1) and cointegrated with the real exchange rate, making the whole right hand side I(0). But from their equation (5) the differential productivity shocks are clearly intended to be I(0). Empirically, their asymmetrical treatment of the real exchange rate and the current account is unsatisfactory, since they have very similar time series properties. As discussed above, in theory both should be I(0) but it is hard to reject the I(1) null. They test the I(0) null for the current account using the KPSS test and cannot reject stationarity. However, the KPSS test is sensitive to the choice of window size (see Kwiatkowski et al 1992), and their VAR estimates suggest a root close to unity for the current account in a number of countries. They do not test for the non-stationarity of the real exchange rate. While their results are driven by the asymmetric treatment of the two variables, in a cointegrating VAR the issues can be data determined.

4 The Data

Quarterly data for 8 countries were taken from the IMF International Financial Statistics 1298_cd (IMF (1999) : Canada (1975Q1-1996Q4), France (1975Q1-1996Q4), Germany (1978Q3-1996Q4), Italy (1975Q1-1996Q4), Japan (1975Q1-1994Q4), Netherlands (1977Q1-1994Q4), United Kingdom (1975Q1-1994Q4), United States (1975Q1-1994Q4). The measure of e_t used is the IMF real effective exchange rate (reu). This is defined in the opposite way to e_t above, so that an increase represents an appreciation, and therefore its coefficient in the long-run solvency equation (10) should be negative for the Marshall-Lerner condition to hold. The trade data were exports and imports of goods and services (lines 90$c.c$ and 98$c.c$, respectively), the measure of real output was real GDP (99$b.r$) and the prices were consumer price indices (64). The annual world output series (99$bp.x$) came from the International Financial Statistics Yearbook 1993 and 1999. The method of interpolation used to derive quarterly figures from the annual series is that of Goldstein and Khan (1976) as reported in Weliwita and Ekanayake (1998).

[4] The relationship between the two approaches are discussed in Pesaran and Smith (1998).

Both real exchange rates and balance of payments ratios show the long persistent swings which give them their I(1) characteristics. Table 1 gives coefficients and ADF statistics for the log of the export-import ratio, b_t, the log of the real exchange rate, e_t, the log of relative income, ($y_t - y_t^*$) and for the residual of an OLS regression of (10). The ADF tests were conducted up to 4 lags with and without a trend and the reported values were chosen on the basis of the Aikaike information criteria (AIC). The results give very little evidence for stationarity, the balance of payments in the Netherlands and the real exchange rate in France being the only cases where the unit root null is rejected for the original variables. The Engle-Granger test suggests cointegration intwo cases, Italy and the UK. These results are not suprising given both the low power of unit root tests and the fact that the balance of payments and real exchange rate adjustment processes have been very noisy over this period.

5 Cointegrating VAR Analysis

Using an unrestricted VAR in the three variables, treating world output as exogenous, with a maximum lag length of six, model selection criteria and degrees of freedom adjusted Likelihood Ratio tests were used to investigate the appropriate order of the VAR, p (see Table 2). The Schwarz Bayesian Criterion, SBC, indicated 1 lag for each of the countries except for Canada where 2 lags were suggested. The Akaike Information Criterion (AIC) tended to indicate higher lags. The Adjusted Likelihood Ratio tests (ALR) tended to suggest zero lags, except for Canada and the USA where 1 lag is suggested and 4 for the Netherlands. There was little evidence of significant seasonality. The no-seasonality null was rejected at 5% in 6 of the 8 cases (see Table 2). This null was rejected in the cases of the Netherlands and was on the 5% margin for Italy. In 80% of the cases (19 out of 24) there was evidence of significant interdependence between the three variables, Granger Non-causality tests of one variable with respect to the other two (reported in the last three columns of Table 2) were generally rejected at standard significance levels, (though since the variables are I(1) the test statistics have non-standard distributions - see Caporale and Pittis, 1999).

Given the uncertainty about the lag length, cointegration analysis was carried out at various lag lengths. The reported results in Table 3 are only for $p = 4$ with seasonal dummies. The columns in the table show the sample period, the number of cointegrating vectors indicated by the Johansen trace test at the 5% and 10% levels, the estimates for a single cointegrating vector normalised with the coefficient of $b_t = -1$, the log-likelihood and the adjustment coefficients for the single cointegrating vector.

Overall, Johansen tests for cointegration suggested one cointegrating vector. The trace statistics (reported) suggested six cases with one cointegrating vector at 5% (Canada, France, Italy, Netherlands, United Kingdom and the United States) and five cases of one cointegrating vector at the 10% level. The eigenvalues (not reported) suggested one cointegrating vector at the 5% level for all the countries except Japan where the results suggest no cointegrating vector.

At the 10% level the eigenvalues results were the same as those for the trace except for Italy and Japan where the eigen results were no cointegration and a single cointegrating vector, respectively. We examined the cointegration properties with lag lengths $p = 1$ to $p = 4$, and the results remained fairly consistent irrespective of the order of the VAR.

We estimated the Johansen procedure with unrestricted ntercept and restricted trends since the data are trended and we wish to avoid the possiblity of quadratic trends in some of the variables. Virtually all the VECM feedback takes place through the balance of payments equation. Table 3 shows α_b, the error correction coeffiecient on the balance of payments equation to be significant and have the correct sign ($\alpha_b > 0$) in 7 out of the 8 cases. There are three significant feedback cases for e and a two significant case (France) for y but one with the wrong sign. The fact that most of the feedback from the cointegrating vector is through the balance of payments equation suggests that we can improve efficiency by treating output and the real exchange rate as weakly exogenous for the parameters of the cointegrating vector. This is investigated in the next section.

To examine the dynamic responses of the VECM in more detail, Generalised Impulse Response (GIR) Functions were calculated. Unlike the traditional orthogonalised impulse response functions, these are invariant to the ordering of the variables. Their advantages are discussed further in Pesaran and Smith (1998). Since the GIR use the information in the historical covariance matrix, to calculate them we treated foreign output as endogenous, despite the earlier evidence on its possible exogeneity. We examine the balance of payments GIR functions for a one standard error appreciation shock in the real exchange rate. With a competitive real depreciation an initial worsening in the bop followed by an improvement would give rise to J-curve effects. These effects arise because in the short run price effects prevail, whereas quantity adjustments dominate in the longer run. Thus (assuming the ML condition holds) a devaluation first worsens then eventually improves the balance of payments. In this particular case of a one standard error premanent appreciation in the real exchange rate, the converse would occur: an initial improvement in the balance of payments followed by a worsening in the long run. This (converse to the J-effect, an L-effect) can be observed in six of the country balance of payments reponses: Canada, Italy, Japan, Netherlands, UK and the USA, see Figure 1. Italy shows a long-run worsening of the balance of payments due to quantity adjustments that we would expect from a J-curve effect, but it does not show the initial price effect improvement we would expect from an appreciation shock. Rather the balance of payments keep worsening over the 6 quarters, before settling to its long-run level at -0.018. In the other five cases the classic J-curve (or rather, its converse) is discernible, therefore, six of the eight countries showed strong responses to the exchange rate shock. In the two remaining cases, France has the opposite movement: a short run decline followed by a long run improvment. Germany showed an initial improvement followed by a decline but this was followed by a long run improvement after 7 quarters.

6 ARDL analysis

The feedback taking place solely through the balance of payments suggests that we treat e and y as long-run forcing variables. If there is a single cointegrating vector which only appears in the balance of payments equation, then there are no efficiency gains in estimating the system as a whole. Estimation can be done using the balance of payments equation alone. The ARDL approach (see Pesaran and Shin, 1995), which can be used only if there exists a unique cointegrating vector, is valid for both trend-stationary and difference stationary regressors, and in small samples it appears to perform better than other procedures such as the FM-OLS method put forward by Phillips and Hansen (1990). Another advantage is that, in the case of difference-stationary variables, appropriate augmentation of the order of the regressors is sufficient to simultaneously correct for residual serial correlation and the endogeneity problem in the estimation of the long-run parameters.

Starting with a fourth order ARDL(4444), with a trend

$$A_1(L)b_t = \alpha + \beta^* A_2(L)y_t^* + \beta A_3(L)y_t + \eta A_4(L)e_t + \gamma t \tag{13}$$

where $A_i(L) = (1 + a_1 L + ... + a_p L^p)$ a general to specific search was conducted guided by the SBC model selection criterion. In all cases the trend was insignificant and the SBC selected ARDL without a trend results are reported in Table 4. Where routine diagnostics (for serial correlation, functional form, normality and heteroscedascity) were failed and the AIC chosen order provided better diagnostic results these are reported, as in the cases of Germany and the Netherlands.

The Marshall-Lerner condition requires that $\theta_e > 0$ and in all cases the Marshall-Lerner condition holds, see Table 4: the coefficient on the log real exchange rate is negative and significant in five of the eight cases (France, Germany, Japan, Netherlands and the USA). Real income effects have the correct sign except for Japan and the USA (neither of which are significant) and are significant for France, Germany, Italy and the United Kingdom. World Income has the correct sign and is also significant in the four instances recorded for domestic income. World income has the wrong sign in the case of the USA but is not statistically significant. The coefficient on the Δe_t variable in the error correction representation of the ARDL estimations that should be positive for J-curve effects to occur, showed only two cases were J-curve effects were statistically significant (Japan and the United Kingdom). Germany and the USA had the correct sign but were not significant; Canada and Italy had the wrong sign but were also not significant; France and the Netherlands had the wrong sign and were statistically significant. Only Japan showed significant Marshall-Lerner and J-curve effects. For the United Kingdom $\theta_e < 0$ was not significant, although there were significant J-curve effects.

Overall the equations showed little sign of structural instability or misspecification. Only Canada failed the Cusum test while passing the Cusumsq test. All the Cusumsq tests were failed except for Canada and the Netherlands (see Table 4). Except for Japan all the countries passed Serial Correlation, Functional

Form, Normality and Heteroscedasticity diagnostic tests. The only instance of reported diagnostic failure is functional form for Japan. The proportion of the change in the log balance of payments ratio that is explained is not high, indicating that a lot of other influences are operating.

A Likelihood Ratio test of the joint hypothesis that all the real exchange rate long-run coefficients were zero is strongly rejected with a test statistic of 68.6.

7 Conclusion

The responsiveness of trade flows to relative price changes is a crucial issue in designing trade or exchange rate policies. Economists have long known that there are no theoretical reasons why devaluations should be associated with an improvement in the trade balance, as this depends on whether or not trade elasticities satisfy the Marshall-Lerner condition. Despite the large body of empirical literature analysing this issue, the controversy on the size of these elasticities has yet to be settled, the available evidence being mixed, and often based on inappropriate statistical methods and/or theoretical frameworks. In particular, many studies, especially earlier ones, do not handle properly possible non-stationarities in the data, and do not take into account the links between relative prices and external equilibrium which are implied by theory. Lee and Chinn (1998) is one of the few exceptions.

In the present paper we use three econometric models, which differ in the degree to which they condition on exogenous variables. First we use a cointegrating VAR which treats all four variables as endogenous. This is used to calculate Generalised Impulse Response Functions, which, unlike conventional impulse responses, are invariant to ordering. Secondly we use a cointegrating VARDL which treats world output as exogenous. Thirdly we use a single equation ARDL which treats all the variables but the balance of payments as exogenous. We estimate each model for eight OECD countries. The evidence suggests that there is a single cointegrating vector and the restrictions involved in moving from the VAR to the ARDL cannot be rejected: output and the real exchange rate can be treated as weakly exogenous for the parameters of the balance of payments equation. Increasing the degree of conditioning (how many variables are treated as exogenous) tends to increase the evidence for the ML condition. Although there is considerable heterogeneity, overall the results suggest that the ML condition holds in the long run with some degree of J-curve effects in the short run. If in addition nominal devaluations can shift the real exchange rate, this would support the case that devaluations can improve the trade balance, with obvious implications for policy makers.

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Table 1

ADF Tests

Country	b_t	e_t	$(y_t - y_t^*)$	**Eqn 10, EC**
			-	**cv=-4.6069**
Canada	-2.286 DF,nt	-2.083 DF,nt	1.004 ADF(1),t	-3.1544
	-9.067, DF,nt*	-10.977 DF,nt*	-7.061 DF,t*	
France	-3.144 ADF(3), t	-4.325 ADF(3),t*,[1]	-2.346 ADF(2), t	-4.2028
	-5.312 ADF(4), nt*		-3.177 ADF(4), nt*	
Germany	-1.611 DF, nt	-2.833 ADF(1), t	-2.112 ADF(4), nt	-1.7587
	-9.058 DF,nt*	-6.798 DF,nt*	-7.645 DF,nt*	*cv= -4.6437*
Italy	-2.176 DF,t	-2.084 ADF(4), nt	-1.049 ADF(1),nt	-4.8045*
	-9.657 DF,nt*	-4.827 ADF(2), nt*	-6.708 DF, t*	
Japan	-2.453 ADF(2), nt	-2.315 ADF(1),t	-1.189 DF,nt	-1.9485
	-4.350 ADF(4),nt*	-6.888 DF,nt*	-6.379 DF,nt*	
Netherlands	-3.678 DF,t*	-3.394 ADF(4),t[2]	-2.568 DF,nt	-4.5353
		-3.497 ADF(4),t*	-8.095 ADF(1),nt*	*cv= -4.6262*
United Kingdom	-2.036 DF,nt	-1.966 ADF(1),nt	-2.351 ADF(3),t	-5.2198*
	-9.773 DF,nt*	-7.681 DF,nt*	-3.617 ADF(2),nt*	
United States	-2.386 ADF(4),nt	-1.028 ADF(1),nt	-1.994 ADF(1),t	-3.4118
	-4.117 ADF(2),nt*	-7.032DF,nt*	-7.819DF,nt*	

DF ~ indicates Dickey Fuller test with specification: $\Delta y_t = \alpha + \beta y_{t-1}$

ADF(.) ~ indicates Augmented Dickey Fuller test with the order shown in the brackets, eg. ADF(1) specification is: $\Delta y_t = \alpha + \beta y_{t-1} + \gamma \Delta y_{t-1}$

nt ~ indicates no trend,

t ~indicates inclusion of a trend

* indicates rejection of the unit root null at 5%.

Table2

Testing for Lag length, Seasonality and Granger Non-Causality

Lag Length Seasonality GNC p values, $p = 6$

Country	**AIC**	**SBC**	**ALR5%**	$\mathbf{p(SC = 0)}$	**b**	**e**	**z**
Canada	3	2	1	0.182	0.049	0.000	0.000
France	1	1	0	0.939	0.039	0.015	0.011
Germany	1	1	0	0.107	0.109*	0.002	0.001
Italy	2	1	0	0.050	0.253*	0.266*	0.013
Japan	3	1	0	0.674	0.003	0.001	0.232*
Netherlands	5	1	4	0.028	0.005	0.006	0.000
United Kingdom	2	1	0	0.943	0.000	0.390*	0.000
United States	2	1	1	0.848	0.002	0.000	0.001

* indicates that the non-causality null is **not** rejected at 5%. Maximum lag length $p = 6$.

Table 3 **VAR(4) Cointegration Tests and Estimates**

Cointegrating equation: $[lb \quad le \quad ly; \quad ly^* \quad \& \quad dly^* \quad s1 \quad s2 \quad s3]$

	Sample	**Trace**	**Trace**	**e**	**y**	**y***	**LL**	$\boldsymbol{\alpha}_b$	$\boldsymbol{\alpha}_e$
		5%	10%						
Canada	75:1	1	1	-1.23	-1.09	-2.97	712.8	0.133	0.064
	96:4			(0.36)	(0.61)	(1.06)		[3.6]	[1.4]
France	75:1	1	2	20.76	20.69	-8.91	772.9	-0.012	-0.028
	96:4			(55.1)	(56.5)	(27.8)		[-1.1]	[-4.4]
Germany	75:3	2	2	-2.28	-2.24	5.55	556.8	0.247	0.069
	96:4			(1.09)	(0.66)	(2.32)		[4.0]	[1.9]
Italy	75:1	1	1	0.49	-3.47	2.71	671.7	0.218	-0.210
	96:4			(0.48)	(1.73)	(1.12)		[3.1]	[-3.2]
Japan	75:1	0	0	-0.88	0.22	0.98	585.6	0.137	-0.057
	96:4			(0.49)	(0.62)	(2.7)		[4.2]	[-1.1]
Netherlands	75:1	1	1	0.23	-0..28	-1.27	687.3	0.423	-0.287
	96:4			(0.14)	(0.24)	(0.50)		[3.5]	[-2.7]
United Kingdom	75:1	1	1	-0.04	-2.03	1.31	641.4	0.592	-0.224
	96:4			(0.07)	(0.18)	(0.60)		[4.4]	[-1.1]
United States	75:1	1	1	-0.58	-0.45	0.65	676.1	0.238	-0.058
	96:4			(0.15)	(1.6)	(1.56)		[5.1]	[-0.9]

Table 4

Error Correction ARDL: Long Run Coefficients

Country	**ARDL**	$\boldsymbol{\theta_e}$	θ_y	θ_y^*	R^{2*}	**cusum/sq**	**SC**
Canada	1000	*-0.31 [-.90]*	-1.14 [-1.29]	0.78 [1.18]	0.10	f / p	p
France	1000	-1.46 [-2.57]	-1.67 [-2.88]	0.93 [3.10]	0.18	p / f	p
Germany	2110[1]	-0.88 [-2.55]	-2.18 [-6.19]	2.27 [5.79]	0.47	p / f	p
Italy	1000	-0.03 [-.15]	-2.19 [-3.36]	1.73 [3.92]	0.18	p / f	p
Japan	1300	-1.18 [-2.6]	0.36 [0.44]	0.54 [0.63]	0.47	p / f	p
Netherlands	1440[1]	-0.21 [1.96]	-0.50 [-1.48]	0.47 [1.90]	0.51	p / p	p
United Kingdom	1200	-0.01 [-.20]	-1.93 [-7.48]	1.04 [6.31]	0.33	p / f	p
United States	1120	-0.63 [-4.68]	0.09 [0.06]	-0.28 [0.76]	0.40	p / f	p

* R^2 of the error correction VARDL specification.

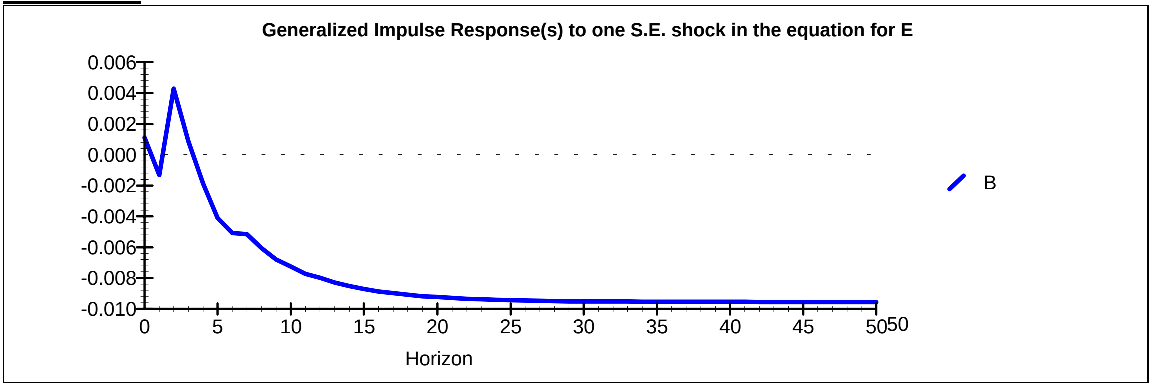
[1] Chosen on the basis of AIC, all others on the basis on SBC.

[..] t statistics in brackets

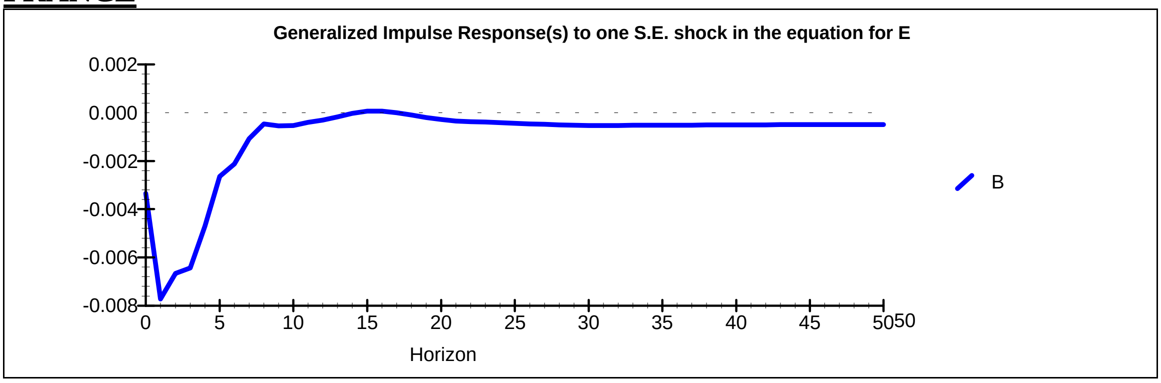
FIGURE 1

GENERALIZED IMPULSE RESPONSE OF COINTEGRATING VECTOR OF BALANCE OF PAYMENTS TO A ONE STANDARD ERROR SHOCK IN EQUATION FOR THE REAL EXCHANGE RATE

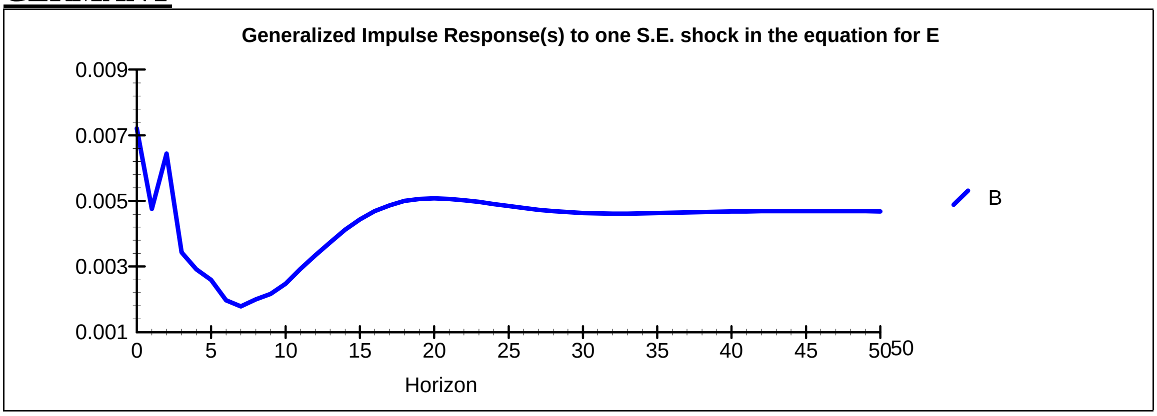
CANADA



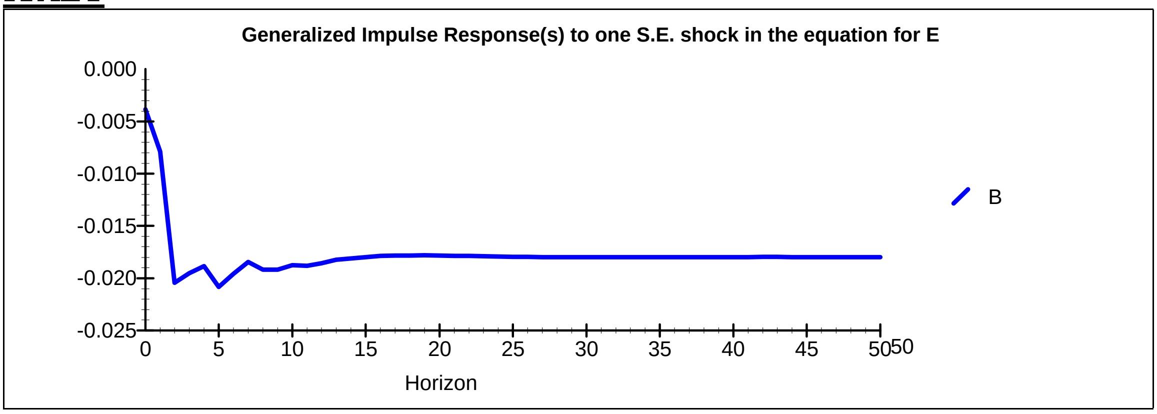
FRANCE



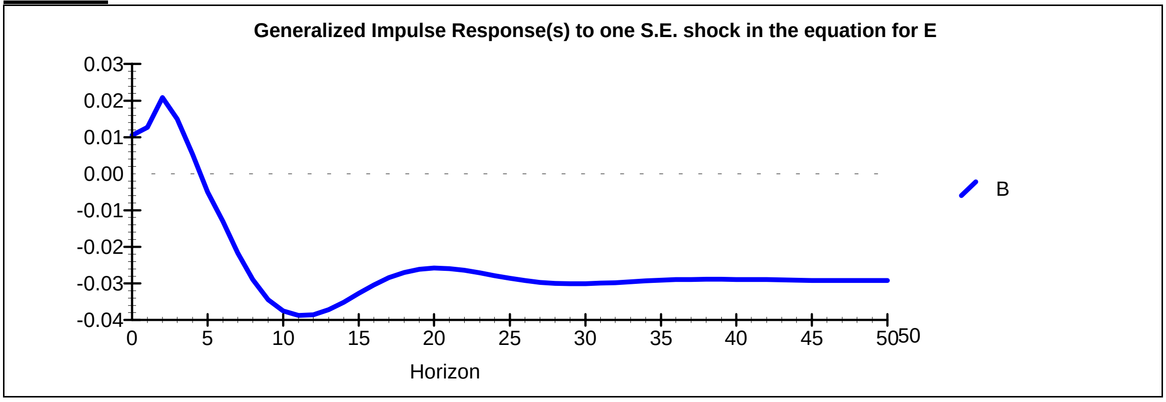
GERMANY



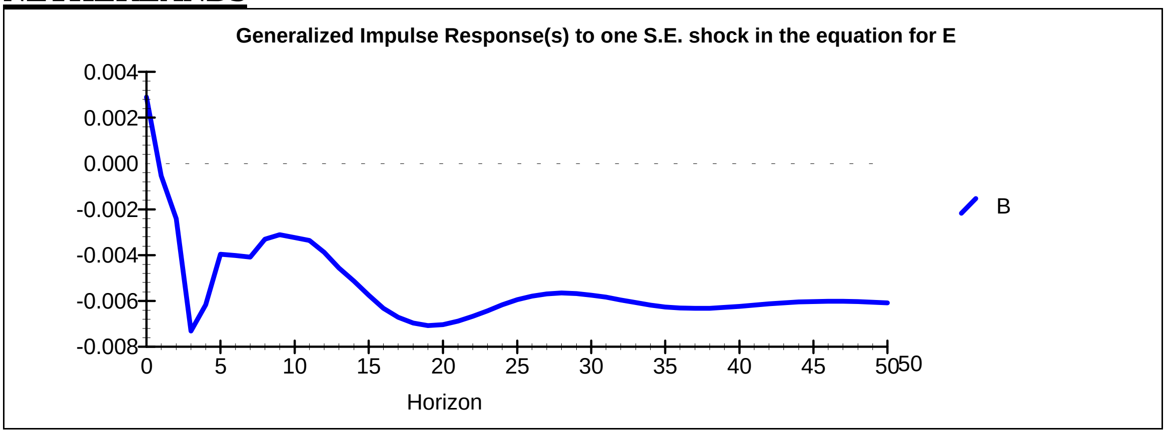
ITALY



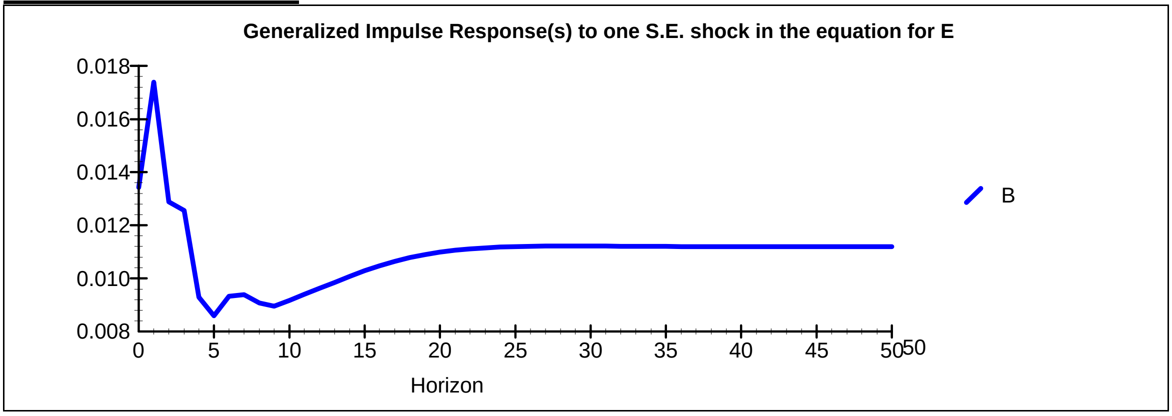
JAPAN



NETHERLANDS



UNITED KINGDOM



UNITED STATES OF AMERICA