Real Exchange Rate Volatility and US Exports: An ARDL Bounds Testing Approach

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ABSTRACT

This paper examines the impact of exchange rate volatility on US exports to the rest of the world, and to each of its five main markets of destination by means of the recently developed ARDL bounds testing approach to cointegration, which is applicable irrespective of whether the regressors are I(1) or I(0). Using a long-term measure of volatility that captures persistence and mean-reversion in the movements of the real exchange rate, we find that in most of the cases considered export volume is significantly affected by volatility, although the sign and magnitude of this effect varies across markets of destination.

1. Introduction

SINCE THE BREAKDOWN of the Bretton-Woods agreement, the question of the impact of exchange rate volatility on the volume of international trade has been a major empirical concern to economists. Yet, despite the abundance of research on this issue,[2] to date no consensus has emerged as to whether exchange rate volatility exerts a significant, positive or negative, effect on the demand for exports. We argue that the absence of a clear pattern of conclusive results in this line of inquiry may be attributed to two main shortcomings in the extant empirical literature.

The first problem concerns the use of inappropriate estimation methods given the time series properties of the variables typically entering the conventional export function. Early studies employed OLS estimation under the erroneous presupposition of stationarity of all the series (see, for example, Pozo, 1992). More recent contributions have used the Johansen procedure to cointegration (Johansen, 1988; 1991) under the equally restrictive assumption that all the variables entering the export function contain a unit root (see among others Chowdhury, 1993). The problem here is that in the presence of a mixture of I(1) and I(0) regressors, statistical inference based on conventional likelihood ratio statistics for cointegration is no longer valid.[3] Both intuition and some previous evidence, however, suggest that the standard export equation is characterised by a mixture of I(1) and I(0) regressors. Foreign income and relative price are expected, and usually found to be integrated of order one, while the volatility measure is, often by construction, level stationary.[4]

The second issue concerns the choice of proxy for exchange rate risk or uncertainty. Most studies have focused on short-term volatility (see, among others, Kenen and Rodrik, 1986; Lastrapes and Koray, 1990; Arize, 1997). Yet, it has long been recognised that it is long-term swings in real exchange rates or currency misalignments (see Williamson, 1983, and De Grauwe, 1988), which are more likely to generate uncertainty and exert a significant impact on export volumes.[5] As argued by Bleaney (1992, p. 564) 'conventional short-run measures of volatility are inadequate, because they fail to take account of either persistence or mean-reversion in the dynamics of the real exchange rate'.

In this paper, we provide additional empirical evidence on the impact of exchange rate volatility on US exports using the recently developed bounds testing procedure to cointegration within an autoregressive distributed lag (ARDL) framework (Pesaran and Shin, 1999; Pesaran *et al.*, 2001). The advantage of this approach over traditional econometric techniques to the analysis of level relationships, is that it allows testing for cointegration regardless of whether the regressors are individually I(1) or I(0). Given the ambiguity surrounding the order of integration of the volatility series, we consider this methodology to be the most congenial in this context. Unlike previous studies, we also make use of a long-term measure of volatility that, being based on the level rather than the rate of change of the real exchange rate, is able to capture the effects of drift as well as persistence in the movements of the variable of interest.

The remainder of this paper is organised as follows. Section 2 describes the specification of the model, the data used, and the ARDL bounds testing procedure. In section 3, we discuss the empirical results. Finally, section 4 provides a summary of the main findings.

2. Model specification and econometric methodology

The standard, or most generally used, long-run export demand model is specified as follows:

$$ex_t = \beta_0 + \beta_1 p_t + \beta_2 ic_t + \beta_3 V_t + u_t \tag{1}$$

where ex_t is real export volume, p_t represents relative price, ic_t is foreign income, and V_t is a measure of volatility. We estimate (1) for US exports of goods to the rest of the world, and to each of its five main destination markets: Canada, Mexico, Japan, the UK and Germany.[6] Relative price is defined as the ratio of US export prices to foreign prices denominated in US dollars. For substitute goods, foreign prices were measured using manufactured goods producer price indices. Foreign income is measured using real GDP for all cases except Germany for which, due to the inexistence of published real GDP data prior to 1991, industrial production was utilised. Following Bleaney (1992), our long-term measure of volatility is based on a moving average standard deviation of the level of the real exchange rate (in log form), where the lag length is constructed using quarterly observations of the previous six years.

Data were obtained from the US Bureau of Economic Analysis and the OECD main economic indicators for the period 1987Q1 to 2001Q2. The start date of the sample period was dictated by data availability with respect to US export price indices for goods, which are only published from 1987Q1. The end date, 2001Q2, was chosen to avoid any potential structural breaks arising from the economic consequences of the terrorist attacks of 9/11.

To implement the ARDL bounds testing approach, we begin by defining y_t as an I(1) dependent variable, and x_t as a vector of I(d) regressors, (where $0 \le d \le 1$). Δy_t is modelled as a conditional ECM:

$$\Delta y_t = c_0 + c_1 t + \pi_y y_{t-1} + \pi_x x_{t-1} + \sum_{i=1}^{p-1} \varphi_i \Delta y_{t-i} + \sum_{j=1}^{q-1} \delta_j' \Delta x_{t-j} + \gamma' \Delta x_t + \varepsilon_t \tag{2}$$

where c_0 and t are drift and trend components, and π_y and π_x are the long-run coefficient matrices for y_{t-1} and x_{t-1}. The short-run dynamic structure of Δy_{t-i} and Δx_{t-j} is set to ensure the residuals, ε_t, are white noise errors.

Cointegration between y_t and x_t, is tested through OLS estimation of (2) and by calculating an F-statistic for the joint significance of the coefficients of the lagged levels, so that $H_0 : \pi_y = 0, \pi_x = 0$. Pesaran *et al.* (2001) prove that, under the null hypothesis, the asymptotic distribution of the F-statistic is non-standard regardless of whether the regressors are I(0) or I(1), and provide two adjusted critical values that establish lower and upper bounds of significance. If the F-statistic exceeds the upper critical value, we can conclude that a long-run relationship exists. If the F-statistic falls below the lower critical value, we cannot reject the null hypothesis of no cointegration. A value of the F-statistic that lies within the bounds makes the test inconclusive. Critical values are also made available to encompass a range of different deterministic components: no drift and no trend; unrestricted intercept and no trend; restricted intercept and no trend; unrestricted intercept and unrestricted trend; and unrestricted intercept and restricted trend.

Following Pesaran and Shin (1999), in the presence of cointegration, the long-run model derived from estimation of the conditional ECM given in (2) is obtained as follows:

$$y_t = \Theta_0 + \Theta_1 t + \Theta_2 x_t + \nu_t \qquad (3)$$

where $\Theta_0 = -c_0/\pi_y$, $\Theta_1 = -c_1/\pi_y$, and $\Theta_2 = -\pi_x/\pi_y$.

The vector x_t is assumed to consist of long-run forcing variables for y_t. Given this assumption, the cointegrating rank is restricted to unity. To test for the absence of feedback from the level of y_t, we use a variant of the bounds test suggested originally by Banerjee *et al.* (1998), which is based on the t-test for H_0: $\pi_y=0$, from OLS estimation of the following:

$$\Delta x_t = c_0 + c_1 t + \pi_x x_{t-1} + \pi_y y_{t-1} + \sum_{i=1}^{p-1} \varphi_i \Delta x_{t-i} + \sum_{j=1}^{q-1} \delta_j' \Delta y_{t-j} + \gamma' \Delta y_t + \varepsilon_t \qquad (4)$$

If the null hypothesis cannot be rejected then x_t is confirmed to be long-run forcing.

3. Empirical results

While the ARDL approach allows estimation of a cointegrating vector with both I(1) and I(0) series, it is still important to exclude the possibility that any of the series are I(2). Table 1 reports the ADF tests. In all cases, export volume is I(1) in levels. Relative prices are level stationary for US exports to the rest of the world but I(1) in the remaining cases. Foreign income is trend stationary for both Mexico and Japan but I(1) otherwise.

Table 1: Unit Root Tests

Variable	*Rest of the world*	*Canada*	*Mexico*	*Japan*	*Germany*	*UK*
ex	-2.479[a] (-5.312[a]*)	1.443 (-2.512*)	-2.969[b]	-2.561[a] (-4.818*)	1.951 (-9.942[a]*)	4.065 (-8.726[a]*)
p	-3.118* (-4.141*)	-2.037[a] (-2.618*)	-2.532[b]	-0.573 (-6.398*)	0.756 (-6.525*)	-1.066 (-6.833*)
ic	3.677 (-4.507[a]*)	2.032 (-2.280*)	-4.672[b]*	-3.468b* (-2.332*)	-2.538[b] (-6.612[a]*)	-2.610[b] (-3.462[a]*)
V	0.3242 (-2.862*)	-3.857[a]* (-3.802*)	-3.796[b]*	-3.311* (-2.158*)	-1.055 (-2.492*)	-1.772* (-2.664*)

The ADF tests for the first difference of each variable are shown in parentheses. * denotes significance at the ten per cent level, using the Mackinnon (1991) finite-sample critical values. Sequential hypothesis testing was used to ascertain whether drift and trend components should be included. [a] denotes the presence of a significant drift component but no trend term, [b] indicates both that drift and trend components are included. Because of the devaluation of the Mexican Peso in 1995Q1, Perron (1989) unit root tests were estimated to allow for this structural break. A structural break was also found in 1988Q1 for the relative price variables of US vs. Mexico. Pulse and level dummies were used, and the null hypothesis tested using Perron (1989) adjusted critical values.

For our purposes, interest centres on the order of integration of the volatility measure, which is found to be level stationary in four of the six cases considered.[7] This result is particularly important in that it confirms the use of the ARDL bounds testing approach that we apply as the most appropriate and useful cointegration procedure in the context of this paper.

To test for cointegration, bounds equations were estimated. For US exports to Mexico, the estimated equation incorporated both drift and trend components, and dummy variables to allow for structural breaks. For all other cases, the model was estimated both with and without drift but no trend term.[8] An 'optimal' dynamic structure was selected, firstly, to ensure an absence of serial correlation in the estimated residuals, and secondly, on the basis of the Akaike Information (AIC) and Schwarz Bayesian (SBC) model selection criteria. The estimated F-statistics reported in table 2 show evidence of cointegration in all cases.

Table 2: Bounds Tests

Importing country	*No intercept and no trend case*	*Unrestricted intercept and no trend case*	*Restricted intercept and no trend case*
Rest of the World	3.594^{a}	3.694^{a}	3.017^{b}
Canada	4.756^{a}	6.249^{a}	5.814^{a}
Japan	6.557^{a}	6.069^{a}	5.451^{a}
Germany	3.110^{a}	1.130^{c}	3.467^{a}
UK	4.584^{a}	7.212^{a}	6.079^{a}
	Unrestricted intercept and unrestricted trend case	*Unrestricted intercept and restricted trend case*	
Mexico	8.375^{a}	6.919^{a}	

The estimated F-statistics impose zero restrictions on the coefficients of the lagged levels from (2). The restricted intercept/trend cases test for the joint significance of the drift/trend components and the lagged levels. The bounds equation estimated for Mexico also includes pulse and level dummies to allow for structural breaks at 1995Q1 and 1988Q1. ^{a} indicates that the test statistic is above the ten per cent upper critical value; ^{b} indicates the value lies in the indeterminate region; ^{c} implies the test statistic falls below the lower critical value.

Since the ARDL approach assumes that the size of the cointegrating space is unity, it is important to ascertain whether the regressors from (1) are in fact long-run forcing. With this aim in mind, (4) was estimated for each regressor, over all cases. As shown in table 3, the t-statistic for H_0: $\pi_y=0$ falls below the lower critical value in all of the 18 cases considered, therefore, the assumption of a unique cointegrating vector among ex_t, p_t, ic_t, and V_t cannot be rejected.

Table 3: Long-run Forcing Tests

Country	*p*	*ic*	*V*
Rest of the World	-0.878	1.126	1.442
Canada	1.632	1.753	-1.628
Mexico	0.794	4.736	-1.133
Japan	1.317	-0.624	1.041
Germany	1.899	2.678	-0.407
UK	0.550	-0.605	-0.328

t-statistics are reported for the null hypothesis H_0: $\pi_y=0$ from (3). The lower critical value (Pesaran *et al.*, 2001) is -2.57.

We then proceeded to derive the long-run estimates by means of the ARDL approach. In choosing the short-run dynamics of the ARDL-ECM, the lag structure was specified on the basis of the AIC and SBC model selection criteria used during the OLS estimation of the bounds tests. The results are presented in table 4.

Table 4: Long-run estimates

Importing country	*p*	*ic*	*V*
Rest of the World	-0.996	1.993	-0.364
(2,1,1,1)	(-9.766*)	(17.362*)	(-0.284)
Canada	-0.089	0.865	14.943
(5,0,2,0)	(-0.053)	(0.470)	(1.039)
Mexico[b]	-0.428	4.202	-0.093
(2,1,1,1)	(-1.823*)	(4.932*)	(-2.697*)
Japan	-0.773	1.736	1.824
(2,1,1,1)	(-5.590*)	(12.873*)	(2.961*)
Germany[a]	-0.514	3.761	-1.314
(4,3,3,3)	(-1.100)	(4.863*)	(-2.382*)
UK	-0.064	2.934	-0.705
(2,1,1,1)	(-0.249)	(10.139*)	(-1.967*)

The chosen lag structure is reported in parentheses (the value of p from (2)). [a] denotes the inclusion of a drift term, [b] indicates the use of drift and trend components as well dummy variables for structural breaks. The values in parentheses below the estimated coefficients are t-ratios. * indicates significance at the ten per cent level.

What is first apparent is that the dynamic structure varies between cases, though consistency appears for US exports to the rest of the world, Mexico, Japan and the UK. Of particular concern is the case of US exports to Canada, where the large dynamic specification necessary to remove serial correlation brings a loss of degrees of freedom which, in turn, appears to severely impair the precision of the point estimates.

The estimated foreign income coefficient is significant and positive in five out of the six cases. The estimates suggest that US exports are strongly influenced by the level of foreign economic activity. Export demand is income elastic in all but one case; US exports to Canada. All the parameters for relative price have the expected sign but are significant only with respect to US exports to Mexico, Japan and the rest of the world. Overall, it would appear that price factors are less important than foreign income in the determination of export demand.

The estimated volatility coefficient has a negative sign in four out of the six cases and it is statistically significant for US exports to Mexico, Germany and the UK. The magnitudes of these parameters vary from -0.093 (Mexico) to -1.314 (Germany).[9] Whilst doubt can be cast upon the accuracy of the estimated volatility coefficient for US exports to Canada (which is insignificant), a positive and significant volatility effect of exchange rate volatility is found in relation to exports to Japan. This stark finding lends itself to different interpretations. The most obvious one, would be to attribute this result to a strong and consistent appreciation of the Japanese yen vis-à-vis the US dollar. This explanation, however, is not supported by the data which indicate that, up to the mid 1990s, Japan experienced a significant deterioration of their terms of trade. It may be more likely that this result is due to the high degree of risk aversion characterising Japanese culture. The economic intuition behind this argument mirrors that made by De Grauwe (1988) who showed that if producers are sufficiently risk averse, an increase in exchange rate risk raises the expected marginal utility of export revenue and therefore induces them to increase their export activity. Since in the case of US exports it is the importer who typically bears the risk, and given the risk averse nature of Japanese traders, it is plausible to suggest that as volatility increases Japanese buyers will import more and build up stock to avoid the possibility of a sharp rise in their costs. Inspection of the data provides empirical support to this hypothesis by confirming that, over the sample period, the volume of Japanese imports of goods from the US grew by 147 per cent.

4. Conclusions

Previous research on the impact of exchange rate volatility on export volume inadequately accounts for the statistical properties of the variables entering the conventional long-run export demand model. In this paper we paid specific attention to the correct treatment of the relevant time series by means of an econometric technique capable of testing for the existence of cointegration irrespective of the order of integration of the underlying regressors.

Using a long-term measure of volatility that accounts for persistence and mean-reverting tendencies in the dynamics of the real exchange rate, we estimated US exports to the rest of the world, and to each of its five main trading partners, over the 1987Q1-2001Q2 sample period. Three main conclusions can be derived from the empirical findings.

First, the standard export demand equation is confirmed to contain a mixture of I(1) and I(0) regressors, with the volatility series found to be mainly level stationary. Second, all the bounds test equations point to the existence of a unique cointegrating relationship among export volume, relative price, foreign income and real exchange rate volatility. Finally, our results indicate that volatility has a statistically significant impact on US exports to most of the markets of destination considered in this study, although the sign and magnitude of this effect varies across cases.

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ENDNOTES

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2. For a comprehensive review of this literature see McKenzie (1999).

3. The issue here is not whether the Johansen method can handle a mix of I(1) or I(0) variables, but how well it handles such a mix. As early as 1995, Harris highlighted that when using the Johansen technique, as a check in deciding the value of *r*, one should ideally compute also the eigenvalues of the companion matrix "*this is useful given the probable poor size and power properties of the λ-max and trace statistics, especially when additional I(0) variables are included*" (Harris, 1995, p.146). The problems inherent in regressions that try to explain the behaviour of a non-stationary series when stationary regressors are also included, are further discussed by Hassler (1996) and Rahbek and Mosconi (1999). The latter specifically demonstrates that when testing for cointegration rank, the presence of stationary explanatory variables as extra regressors leads to nuisance parameters (characterised as the canonical correlations between the common trends and the accumulated errors) in the asymptotic distribution of the trace statistic.

4. For example, Kroner and Lastrapes (1993) found a stationary volatility series in the case of UK and West German exports.

5. This distinction is important in that while traders can easily insure against short-term risk through forward market transactions, it is much more difficult and expen-

sive to hedge against long-term risk or uncertainty.

6. On average, over the sample period, these five markets of destination accounted collectively for 51.9 per cent of total US exports of goods.

7. It is worth pointing out that in the pre-testing phase we also investigated the statistical properties of commonly employed short-run measures of volatility, namely the moving average standard deviation of the growth rate of both the nominal and the real exchange rate, both of which were found to be level stationary in all cases.

8. These specifications were chosen following pre-testing for drift and trend components when calculating the unit root tests for export volumes.

9. As a check for robustness of our chosen measure of volatility, we also tested the stability of the estimated volatility coefficients on the shorter 1987-1999 subsample of our full set of observations. With the exception of the regression for US exports to Mexico (for which we obtained an insignificant parameter), all the other estimated volatility coefficients provided inferences consistent with those drawn from estimation based on the full sample period. Specifically, the estimated volatility coefficients for US exports to 'Rest of the World' and Canada were found to be insignificant, while the signs and magnitudes of the other statistically significant parameters were: 1.96 (Japan), -1.32 (Germany), and -0.67 (UK).

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