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# Does Exchange Rate Variability Depress Trade Flows ? Evidence From Error Correction Models

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Published version. *Review of Economics and Statistics,* Vol. 75, No. 4 (November 1993): 700-706.
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# DOES EXCHANGE RATE VOLATILITY DEPRESS TRADE FLOWS? EVIDENCE FROM ERROR-CORRECTION MODELS

Abdur R. Chowdhury*

*Abstract*—This paper examines the impact of exchange rate volatility on the trade flows of the $G$-7 countries in the context of a multivariate error-correction model. The error-correction models do not show any sign of parameter instability. The results indicate that the exchange rate volatility has a significant negative impact on the volume of exports in each of the $G$-7 countries. Assuming market participants are risk averse, these results imply that exchange rate uncertainty causes them to reduce their activities, change prices, or shift sources of demand and supply in order to minimize their exposure to the effects of exchange rate volatility. This, in turn, can change the distribution of output across many sectors in these countries. It is quite possible that the surprisingly weak relationship between trade flows and exchange rate volatility reported in several previous studies are due to insufficient attention to the stochastic properties of the relevant time series.

## I. Introduction

The high degree of volatility and uncertainty of exchange rate movements since the beginning of the generalized floating in 1973 have led policy makers and researchers to investigate the nature and extent of the impact of such movements on the volume of trade.[1] However, these studies dealing with the effects of exchange rate volatility on trade flows have yielded mixed results. On one hand, a number of studies have argued that exchange rate volatility will impose costs on risk averse market participants who will generally respond by favoring domestic to foreign trade at the margin. The argument views traders as bearing undiversified exchange risk; if hedging is impossible or costly and traders are risk-averse, risk-adjusted expected profits from trade will fall when exchange risk increases. Akhtar and Hilton (1984), Coes (1981), Cushman (1983, 1986), Kenen and Rodrik (1986), Koray and Lastrapes (1989), Thursby and Thursby (1987) provide evidence in support of this view. On the other hand, Franke (1992), Giovannini (1988), and Sercu and Vanhulle (1992) have shown that trade benefits from exchange rate volatility or risk.[2] According to these studies, trade can be considered as an option held by firms. Like any other option, such as, stocks, the value of trade can rise with volatility. Franke (1992) develops a model in which a firm evaluates the exit (entry) costs associated with leaving (entering) a foreign market against losses (profits) created by exports. Under a variety of behavioral assumptions, it is possible that any given firm will on average enter sooner and exit later when exchange rate volatility rises, thus increasing the average number of trading firms.[3] Empirical evidence supporting this view is given in IMF (1984) and Asseery and Peel (1991). Gotur (1985) and Bailey, Tavlas, and Ulan (1986), on the other hand, do not find conclusive evidence that exchange rate volatility have any significant effect on trade flows.

In summary, there is conflicting evidence in the literature about the relationship between exchange rate volatility and trade flows. No clear pattern of results or consistent conclusions emerge from these studies. This paper reexamines the impact of exchange rate volatility on the trade flows of the $G$-7 countries—Canada, France, Germany, Italy, Japan, the United Kingdom, and the United States—in the context of an error-correction model. The paper intends to shed additional light on a topic characterized by conflicting empirical evidences. This paper is different from previous papers in this area in several ways. First, most of the previous studies fail to recognize that real exports and some of its determinants, e.g., foreign real income or real world trade, are, potentially non-stationary integrated variables.[4] This study focuses upon the appropriate representation of the nature of nonstationarities apparent in various time series across different countries. Second, there is much evidence that a lagged relationship may exist between the volume of exports and its various determinants. Standard trade models tend to ignore

Received for publication January 24, 1992. Revision accepted for publication August 31, 1993.

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The paper was presented at the Allied Social Science Association Meetings at Anaheim, January 1993. I am indebted to the referees of this *Review* for extremely helpful comments and suggestions. The usual caveats apply.

[1] Akhtar and Hilton (1984) provides an excellent theoretical discussion of how exchange rate volatility can affect trade flows. For an empirical analysis, see, Makin (1976), Hooper and Kohlhagen (1978), Cushman (1983), Thursby and Thursby (1987), Koray and Lastrapes (1989), and Lastrapes and Koray (1990). Kumar and Whitt (1992) provides a summary of the results of most of these studies.

[2] Edison and Melvin (1990) have also discussed the plausible arguments in favor of a positive relationship between trade and exchange rate volatility.

[3] Moreover, the substitution effect of resources shifted away from the traded-goods sector due to volatility increases may be dominated by an income effect working in the opposite direction. De Grauwe (1988) has argued that this will be the case if producers are sufficiently risk averse and increase resources in the export sector to offset the drop in expected utility of export revenue caused by the increase in exchange rate volatility.

[4] Koray and Lastrapes (1989), Lastrapes and Koray (1990), and Asseery and Peel (1991) are exceptions.

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the possibility of such a lagged relationship. This paper explicitly takes this possibility into account. It studies the long run relationship between exports and its various determinants and also considers the short-run dynamics by which exports converge on their equilibrium long-run values. Third, several different exchange rate volatility measures have been used in the literature, e.g., average of absolute changes, standard deviations, and deviations from trend. In contrast, this paper uses the standard deviation of the growth rate of the exchange rate to measure volatility. This measure captures the general movements in exchange rate volatility, and therefore exchange rate risk, over time. Finally, it employs a longer sample period than any of the previous studies in this area.

The variables and the data sources are given in section II. The estimation procedure and initial results are presented and discussed in section III while section IV reports the short-run dynamic results. The last section contains a summary and conclusions.

## II. Variables and Data Source

Drawing upon the empirical literature in this area and the implications for dynamic specification of the possible existence of error-correcting mechanisms in the data generating process,[5] a simple, standard long-run relationship between real exports, the level of real activity, competitiveness and exchange rate volatility is specified [Gotur (1985), Asseery and Peel (1991)]:

$$\ln X_t = \delta_0 + \delta_1 \ln Y_t + \delta_2 \ln P_t + \delta_3 V_t \qquad (1)$$

where $X_t$ is real export volume, $Y_t$ is a measure of real foreign economic activity, $P_t$ represents relative price which is a measure of competitiveness, and $V_t$ is the measure of exchange rate volatility. Gotur (1985) has shown that equation (1) can be derived as a long run solution of behavioral demand and supply functions for exports. The reduced form approach used in this paper is consistent with much of the literature in this area.

The export variable is measured by the export volume index of each country. The data on export is taken from the *International Financial Statistics Tape*. Economic theory suggests that income in trading-partner countries is a major determinant of a nation's exports. For each of the seven countries considered in the paper, we construct a series defined as the weighted average of the GDP series of each country's ten most important trading partners.[6] Quarterly GDP data for the major trading partners of the seven countries considered in this paper are available in the *OECD Main Economic Indicators*. The individual GDP series are converted to a common currency—US dollar—for aggregation purposes.

The second explanatory variable in the export equation measures competitiveness ($P_t$), where $P_t$ is defined as the ratio of export prices of country $i$ to those of its major trading partners. Hence for each country, the variable is constructed as the ratio of that country's dollar-denominated export unit value to the aggregate export unit value for its ten major trading partners.[7] Data for individual country's export-unit values are taken from *International Financial Statistic Tape* while aggregate export unit value data for the trading partners are taken from *OECD Main Economic Indicators*. Bailey et al. (1986) have shown that since the components of $P_t$ are denominated in terms of U.S. dollar, changes in $P_t$ are changes in real terms of trade, reflecting (i) variations in nominal exchange rates; (ii) differing rates of inflation among countries; and (iii) changes in relative prices in each country between its non-traded goods and its exports. Hence the ratios reflect real exchange rates in terms of traded goods.

A time varying measure of exchange rate volatility is included in the model in order to account for periods of high and low exchange rate uncertainty.[8] The variable is constructed by the moving sample standard deviation of the growth rate of the real exchange rate

$$V_t = \left[ (1/m) \sum_{i=1}^{m} (\log Q_{t+i-1} - \log Q_{t+i-2})^2 \right]^{1/2} \qquad (2)$$

where $m = 8$ is the order of the moving average.[9] This measure is similar to those used in much of the literature (for example, Kenen and Rodrik (1986), Koray and Lastrapes (1989), Lastrapes and Koray (1990)).[10] Koray and Lastrapes (1989) have shown that this measure captures the temporal variation in the absolute magnitude of changes in real exchange rates, and

[5] See Davidson, Hendry, Srba, and Yeo (1978) for a more detailed discussion on this issue.

[6] Following Savvides (1992), the ten most important trading partners are determined on the basis of average trade (exports and imports) weights over the 1973–90 period. Weights are then computed as each trading partner's trade share in the total.

[7] The weights used for constructing the price variable are similar to those used in the construction of the GDP variable.

[8] Since there is no unique way to measure exchange rate uncertainty, empirical research on its effects has generally used some measure of exchange rate variability as a proxy for uncertainty. See Akhtar and Hilton (1984) for a discussion on this issue.

[9] Estimations have also been performed using two different values of $m$, $m = 4$ and $m = 12$. The conclusion appears to be robust irrespective of the value of $m$.

[10] For an alternative measure of exchange rate volatility, see, for example, Puzo (1992). However, Puzo's results are not sensitive to the particular measure of exchange rate volatility employed.

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therefore exchange risk, over time.[11] The real effective exchange rate constructed and published by the Morgan Guaranty Bank are used in this paper. These rates are constructed by adjusting a country's nominal effective exchange rate by the differential movements between the manufacturers' segment of the trading partners' wholesale price indices (WPI) and that of the manufacturers' segment of the domestic nation's WPI.

### III. Estimation Procedure and Initial Results

This section reports the initial results of estimating the impact of exchange rate variability on exports in each of the *G*-7 countries—Canada, France, Germany, Italy, Japan, the United Kingdom, and the United States. The sample period runs from 1973:I to 1990:IV. The starting point of the sample period corresponds to the time of the formal abandonment of the fixed exchange rate regime. By the end of 1972, all the countries in the sample had begun to allow their currencies to float. Several observations at the beginning of the sample period are lost due to the construction of the volatility measure and use of lagged explanatory variables in the models. The estimation period is 1976:II–1990:IV.

For each country an error-correction model for exports is developed. However, a prerequisite for developing such a model is to test the long run relationship among the variables included in equation (1). The cointegration procedure developed in Johansen (1991) and Johansen and Juselius (1990) is employed to test the long-run relationship in equation (1).[12]

[11] There has been conflicting argument as to whether exchange rate uncertainty is better measured by nominal or real exchange rate volatility. Studies by Hooper and Kohlhagen (1978), Akhtar and Hilton (1984), among others, have reported significant trade flow effects of nominal exchange rate volatility, while Cushman (1983), Kenen and Rodrik (1986), and Thursby and Thursby (1985) have found significant trade flow effects of real exchange rate volatility. Consequently, the models have been estimated alternatively using the nominal and the real measures of exchange rate volatility. The results are qualitatively similar. Thursby and Thursby (1987) and Lastrapes and Koray (1990) reported similar findings. Addressing a different issue, Mark (1986) finds little difference in real and nominal exchange rate changes. Such a phenomenon may explain similar trade flow effects. Only results using the real exchange rate are reported in the paper. The results using the nominal exchange rate are available from the author upon request.

[12] The Johansen-procedure is preferred to Engle and Granger's (1987) regression-based technique because it fully captures the underlying times series properties of the data and provides estimates of all the cointegrating vectors that exist within a vector of variables. It clearly shows whether the system consists of a unique cointegrating vector or a linear combination of several cointegrating vectors. It has also been argued in the econometric literature that Johansen's technique is more discerning in its ability to reject a false null hypothesis (Ericsson (1991)).

Implementation of the Johansen procedure requires the determination of a lag length for the VAR model for each country and the orders of integration of the variables entering each of the VAR models. The order of integration of the individual time series are determined using the augmented Dickey-Fuller test (Dickey and Fuller (1979)) and the Kwiatkowski test (Kwiatkowski et al. (1992)). Irrespective of the country considered, the results (not reported here) indicate that the variables included in this study are integrated of order one.

Following Lutkepohl (1982), Akaike's AIC criteria is used to determine the lag length for the VAR model.[13] The unreported results indicate that the optimum lag length for the VAR models vary from a low of one in Japan to a high of three in France and the United Kingdom. The VAR models for Canada, Germany, Italy, and the United States have a lag length of two. The Ljung-Box $Q$-statistics indicate that the residuals from each VAR model is a white noise.

The next step is to test for the presence of common stochastic trends. Table 1 reports the results from the Johansen likelihood ratio tests for cointegration. The null hypothesis is that there are $r$ or fewer cointegration vectors. The alternative hypotheses are $r + 1$ and at least $r + 1$ cointegration vectors for the maximum eigenvalue and trace statistics, respectively. Among four variables there is the possibility of zero, one, two, or three cointegrating vectors. For the trace test, the hypotheses $r \leq 1$, $r \leq 2$, and $r \leq 3$ cannot be rejected for any of the countries in the sample, while the hypothesis $r = 0$ can be uniformly rejected. The 5% critical value for $H_0$: $r = 0$ is 48.42. The calculated test statistics vary from a low of 52.89 in Italy to a high of 74.35 in Japan. This implies that there is only one cointegrating vector.

The maximum eigenvalue test results are similar to the trace statistics results. For each country, the null hypothesis $r = 0$ is uniformly rejected in favor of the alternative hypothesis $r = 1$. On the other hand, none of the null hypotheses $r = 1, r = 2$, or $r = 3$ can be rejected in favor of the alternative hypotheses $r = 2$, $r = 3$, and $r = 4$, respectively. Consequently, these test results indicate that, irrespective of the country considered, real exports is cointegrated with the measure of foreign economic activity, competitiveness measure, and exchange rate volatility. Moreover, the results confirm the presence of one cointegrating vector for each country in the sample. The maximum eigenvalue test of $r = 1$ versus $r = 2$ fails to reject the null hypothesis of

[13] Sawa (1978) has argued that the AIC criterion tends to choose models of higher order than the true model but states that the bias is negligible when $k < T/10$, as it is here. The maximum lag length considered in this paper is three. Considering a maximum lag length of more than three quarters would greatly reduce the degrees of freedom.

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TABLE 1.—RESULTS FROM COINTEGRATION TESTS

| Country | Trace Statistics | | | | Maximum Eigenvalue | | | |
|---|---|---|---|---|---|---|---|---|
| | $r = 0$ | $r \leq 1$ | $r \leq 2$ | $r \leq 3$ | $r = 0$ | $r = 1$ | $r = 2$ | $r = 3$ |
| Canada | 73.15 | 26.16 | 14.39 | 3.74 | 46.99 | 11.77 | 10.65 | 3.74 |
| France | 59.38 | 29.66 | 15.60 | 4.88 | 29.72 | 14.06 | 10.72 | 4.88 |
| Germany | 61.30 | 27.33 | 14.20 | 2.19 | 33.97 | 13.13 | 12.01 | 2.19 |
| Italy | 52.89 | 19.75 | 10.66 | 4.28 | 33.14 | 9.09 | 6.38 | 4.28 |
| Japan | 74.35 | 26.34 | 12.59 | 6.55 | 48.01 | 13.75 | 6.04 | 6.55 |
| U.K. | 69.22 | 18.93 | 10.44 | 4.10 | 50.29 | 8.49 | 6.34 | 4.10 |
| U.S.A. | 58.16 | 24.60 | 13.78 | 4.36 | 33.56 | 10.82 | 9.42 | 4.36 |
| Critical Value | 48.42 | 31.26 | 17.84 | 8.08 | 27.34 | 21.28 | 14.60 | 8.08 |

Note: $r$ denotes the number of cointegrating vectors: The critical values are for the 5% level of significance (Johansen and Juselius (1990, table A2)).

TABLE 2.—COINTEGRATING VECTORS AND LIKELIHOOD RATIO TESTS

| Country | Normalized Cointegrating Vector | Test[a] $H_0$: $\delta_3 = 0$ |
|---|---|---|
| Canada | $X_t = 0.178Y_t - 0.462P_t - 0.684V_t$ | 15.63[a] |
| France | $X_t = 0.340Y_t - 0.371P_t - 0.502V_t$ | 4.10[a] |
| Germany | $X_t = 0.373Y_t - 0.005P_t - 0.518V_t$ | 3.91[a] |
| Italy | $X_t = 0.013Y_t + 0.233P_t - 0.077V_t$ | 3.70[b] |
| Japan | $X_t = 0.348Y_t - 0.224P_t - 0.820V_t$ | 8.24[a] |
| United Kingdom | $X_t = 0.270Y_t - 0.180P_t - 0.677V_t$ | 7.11[a] |
| United States | $X_t = 0.174Y_t - 0.260P_t - 0.598V_t$ | 12.36[a] |

Note: The test $H_0$: $\delta_3 = 0$ in the equation $X_t = \delta_0 + \delta_1 Y_t + \delta_2 P_t + \delta_3 V_t$ has a $\chi^2(1)$ distribution under the null hypothesis.
[a] Significant at the 5% level.
[b] Significant at the 10% level.

$r = 1$. This suggests that there is one linear long-run equilibrium relationship among the variables, and that any departure from this relationship may be due to temporary disequilibrating forces.

The cointegrating vectors are given in table 2. To give economic meanings to the estimated vectors, they are normalized on $X_t$. This is done by setting the estimated coefficient on $X_t$ equal to $-1$ and dividing each cointegrating vector by the negative of the estimated $X_t$ coefficient. The results of this normalization yield estimates of the long-run elasticities. For each country, foreign economic activity $(Y_t)$ is positively related to export volume $(X_t)$. The long term elasticity ranges from a low of 0.013 in Italy to a high of 0.348 in Japan. The competitiveness variable $(P_t)$ has the expected negative sign in all countries except Italy. More interestingly, the volatility measure has a negative sign for all countries. The long term elasticity appears to be relatively high for all countries, except one. The last column of the table report test statistics for the null hypothesis that the volatility elasticity is zero. The calculated test statistics, distributed as a $\chi^2(1)$, is large enough to reject the null hypothesis. The hypothesis is rejected at the 5% level for all countries except Italy, where it is rejected at the 10% level.

## IV. Short-run Dynamic Adjustments

Based on the Representation theorem developed in Engle and Granger (1987), it can be shown that the following error correction model (ECM) exists for a cointegrating vector $(X, Y, P, V)$:

$$\Delta X_t = \alpha_0 + \alpha_1 R_{t-1} + \sum \beta_i \Delta X_{t-i} + \sum \gamma_i \Delta Y_{t-i} + \sum \delta_i \Delta P_{t-i} + \sum \tau_i \Delta V_{t-i} + e_t \tag{3}$$

where $R_{t-1}$ is the lagged error correction term and is the residual from the cointegrating regression equation (1). It should be noted that if the variables have a cointegrating vector, then $R_t \sim I(0)$ represents the deviation from equilibrium in period $t$. The ECM shows how the system converges to the long-run equilibrium implied by the cointegrating regression. The coefficient $\alpha_1$ in equation (3) represents the response of the dependent variable in each period to departures from equilibrium. This approach makes it easier to distinguish between the short-run and long-run real exports function.

The ECMs are estimated for each of the $G$-7 countries for the sample period 1976:II–1990:IV. The re-

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TABLE 3.—REGRESSION RESULTS FOR ERROR CORRECTION MODELS: 1976:II–1990:IV

| Country | lag | Variables: $R$ | $\Delta X$ | $\Delta Y$ | $\Delta P$ | $\Delta V$ | Summary Statistics |
|---|---|---|---|---|---|---|---|
| Canada | 1 | −0.275 | 0.066 | 0.313 | 0.154 | −0.245 | $\bar{R}^2 = 0.55$ DW = 2.04 |
| | | (−2.6) | (2.9) | (1.9) | (2.3) | (−3.6) | AR $F[4, 41] = 1.05$ |
| | | | | | | | ARCH $F[4, 37] = 0.71$ |
| | 2 | | 0.034 | 0.456 | 0.321 | −0.201 | RESET $F[1, 44] = 0.19$ |
| | | | (3.16) | (2.4) | (1.9) | (−2.8) | NORMALITY $\chi^2[2] = 2.15$ |
| | | | | | | | COV $F[14, 31] = 1.14$ |
| | 3 | | 0.166 | 0.330 | −0.003 | | |
| | | | (3.0) | (2.4) | (−2.0) | | |
| | 4 | | −0.124 | | | | |
| | | | (−4.1) | | | | |
| France | 1 | −0.354 | 0.344 | 0.666 | −0.159 | −0.077 | $\bar{R}^2 = 0.61$ DW = 2.01 |
| | | (−3.5) | (1.9) | (2.3) | (−3.9) | (−3.2) | AR $F[4, 47] = 1.09$ |
| | | | | | | | ARCH $F[4, 43] = 0.53$ |
| | 2 | | 0.300 | | | −0.106 | RESET $F[1, 50] = 0.16$ |
| | | | (2.1) | | | (−4.3) | NORMALITY $\chi^2[2] = 2.60$ |
| | | | | | | | COV $F[8, 43] = 1.03$ |
| Germany | 1 | −0.598 | 0.166 | | | −0.288 | $\bar{R}^2 = 0.72$ DW = 2.09 |
| | | (−4.1) | (3.8) | | | (−5.0) | AR $F[4, 46] = 1.60$ |
| | | | | | | | ARCH $F[4, 42] = 0.60$ |
| | 2 | | | 0.134 | 0.006 | −0.177 | *RESET* $F[1, 49] = 0.25$ |
| | | | | (3.9) | (2.1) | (−3.8) | *NORMALITY* $\chi^2[2] = 1.95$ |
| | | | | | | | COV $F[9, 41] = 0.78$ |
| | 3 | | −0.109 | 0.199 | | | |
| | | | (−2.3) | (6.1) | | | |
| Italy | 1 | −0.351 | | 0.496 | 0.115 | −0.060 | $\bar{R}^2 = 0.49$ DW = 1.96 |
| | | (−3.2) | | (2.1) | (5.6) | (−5.0) | AR $F[4, 47] = 1.00$ |
| | | | | | | | ARCH $F[4, 43] = 0.75$ |
| | 2 | | | 0.294 | | | RESET $F[1, 50] = 0.39$ |
| | | | | (2.0) | | | NORMALITY $\chi^2[2] = 2.41$ |
| | | | | | | | COV $F[8, 43] = 0.66$ |
| | 3 | | 0.036 | | | −0.045 | |
| | | | (1.9) | | | (−4.2) | |
| Japan | 1 | −0.441 | −0.411 | | 0.138 | −0.254 | $\bar{R}^2 = 0.60$ DW = 2.13 |
| | | (−5.3) | (−2.1) | | (2.3) | (−3.9) | AR $F[4, 46] = 1.60$ |
| | | | | | | | ARCH $F[4, 42] = 0.66$ |
| | 2 | | | | −0.066 | −0.204 | RESET $F[1, 49] = 0.11$ |
| | | | | | (−1.9) | (−3.6) | NORMALITY $\chi^2[2] = 1.73$ |
| | | | | | | | COV $F[9, 41] = 0.63$ |
| | 3 | | | | | −0.101 | |
| | | | | | | (−2.7) | |
| | 4 | | −0.155 | | | | |
| | | | (−1.9) | | | | |
| United Kingdom | 1 | −0.365 | −0.188 | 0.225 | | −0.096 | $\bar{R}^2 = 0.53$ DW = 2.07 |
| | | (−2.41) | (−3.6) | (3.1) | | (−2.0) | AR $F[4, 45] = 0.42$ |
| | | | | | | | ARCH $F[4, 41] = 2.01$ |
| | 2 | | −0.241 | 0.214 | −0.168 | −0.070 | RESET $F[1, 48] = 0.43$ |
| | | | (−4.1) | (2.0) | (−5.9) | (−1.9) | NORMALITY $\chi^2[2] = 1.84$ |
| | | | | | | | COV $F[10, 39] = 0.45$ |
| | 3 | | −0.266 | | | | |
| | | | (−4.4) | | | | |
| United States | 1 | −0.455 | 0.302 | 0.103 | −0.255 | −0.187 | $\bar{R}^2 = 0.58$ DW = 1.90 |
| | | (−4.0) | (2.2) | (4.1) | (−2.0) | (−3.6) | AR $F[4, 43] = 1.30$ |
| | | | | | | | ARCH $F[4, 39] = 0.96$ |
| | 2 | | 0.266 | 0.098 | | −0.215 | RESET $F[1, 46] = 0.50$ |
| | | | (2.1) | (3.6) | | (−4.8) | NORMALITY $\chi^2[2] = 1.90$ |
| | | | | | | | COV $F[12, 35] = 0.40$ |
| | 3 | | | 0.077 | | | |
| | | | | (3.4) | | | |
| | 4 | | | | −0.366 | 0.088 | |
| | | | | | (−2.8) | (2.0) | |

Note: Figures in parentheses are the $t$-statistics.
DW tests first-order residual autocorrelation.
AR $F[q, T-k-q]$ test $q^{\text{th}}$ order residual autocorrelation.
ARCH $F[q, T-k-2q]$ is the $F$-test for autoregressive conditional heteroskedasticity.
*RESET* $F[q, T-k-Q]$ is the $q^{\text{th}}$ order Ramsey-Reset Statistics.
*NORMALITY* $\chi^2[2]$ is the Jarque-Bera test for skewness and excess kurtosis. It has a chi-square distribution with 2 degrees of freedom.
COV $F[k, T-2k]$ tests the null hypothesis that the parameters are not constant over subsamples.
There are $T$ observations and $k$ regressors in each model under the null. The value of $q$ may differ across statistics, as may those of $k$ across models; $T = 59$.

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sults from the ECM's are summarized in table 3.[14] Four lagged differences of each variable are initially included in each regression equation. Starting with the longest lagged difference, lagged terms that are jointly insignificant by a conventional $F$-test are eliminated from the specification, unless their elimination introduced serial correlation.[15] The statistical adequacy of these error correction equations are checked using a battery of tests. These tests indicate that the ECM for each country is adequate for further analysis. The adjusted $R^2$'s range from a low of 0.49 in Italy to a high of 0.72 in Germany. These $R^2$ values are normal for regressions based on first differences in variables.

For all countries, the error correction term, $R_{t-1}$, is statistically significant and has the expected negative sign. This coefficient gives a measure of the average speed at which export volume adjusts to a change in equilibrium conditions. The absolute values of the error correction terms indicate that the movement of real exports towards eliminating disequilibrium within a quarter varies from one country to another. For example, in Canada only about 28% of the adjustment occurs in one quarter while the figure is 60% for Germany. The coefficients on the foreign income ($Y$) and the relative price ($P$) variables in the ECMs show how the average speed of export adjustment may differ depending on whether the adjustment is in response to foreign income or relative price shocks. In almost all cases, the income coefficients are larger than the relative price coefficients, indicating a faster response of export volume to foreign income changes than to relative price changes.[16]

Of particular interest is the finding that the estimated volatility measure is negative and statistically significant for each of the seven countries. The coefficients on the volatility terms are relatively high in all the countries except Italy and the United Kingdom. However, even in these two countries, the coefficients are statistically significant. These results can be generalized to argue that risk averse market participants react to exchange rate volatility by favoring domestic to foreign trade. This confirms earlier findings reported in Akhtar and Hilton (1984) and Kenen and Rodrik (1986), while contradicts the results given in Gotur (1985) and Asseery and Peel (1991).

[14] It should, however, be noted that since the ECM is treated as a univariate equation for exports, it entails a loss of efficiency compared to a systems approach (especially given the single equation pre-testing procedure).

[15] Dropping statistically insignificant variables or lags from an estimating equation is quite common in economic literature. This provides a model which is parsimonious and easy to interpret. See, for example, Turnovsky and Wohar (1987), for a discussion on this issue.

[16] Hickok, Hung, and Wulfekuhler (1991) report similar findings for the United States. This is also consistent with the results of standard trade models.

Finally, each of the real exports equation are further evaluated by examining their structural stability using the Chow test over the period 1973:I–1990:IV. The Chow test is implemented using the dummy variable approach and potential breakpoints covering the period 1979:IV to 1983:IV. The starting and the ending dates of the breakpoints are respectively eight quarters before and after the mid-point of the entire sample period. The test statistics for each country is the $F$ statistic that tests whether slope dummies when added to the ECM equations in table 3 are jointly significant. The results (not reported here) indicate that these $F$ statistics generally are not statistically significant and thus imply that the ECMs reported in table 3 do not show parameter instability.

## V. Summary and Conclusion

This paper analyzes the dynamic relationship between the volume of export and a measure of exchange rate volatility in the context of a multi-variate error-correction model. The model is estimated for each of the G-7 countries over the 1973–1990 sample period. Several conclusions can be derived from the results. First, each of the series considered in the study are non-stationary integrated variables. Nelson and Plosser (1982), among others, have shown that inferences based on time series which contain stochastic trend or unit root feature is less straightforward than in the case of stationary data. This raises some concern with several previous studies dealing with standard trade models. It is quite possible that the surprisingly weak relationship between trade flows and exchange rate variability reported in several previous studies are due to insufficient attention to the stochastic properties of the relevant time series.

Second, once the non-stationary behavior of the variables are taken into account, the error-correction results indicate that exchange rate volatility has a significant negative impact on the volume of exports in each of the G-7 countries. If market participants are risk averse, these results imply that exchange rate uncertainty causes them to reduce their activities, change prices, or shift sources of demand and supply in order to minimize their exposure to the effects of exchange rate volatility. This, in turn, can change the distribution of output across many sectors in these countries. Finally, the Chow test results show the absence of parameter instability in the estimated models.

REFERENCES

Akhtar, M. A., and R. Spence Hilton, "Effects of Exchange Rate Uncertainty on German and U.S. Trade," Federal Reserve Bank of New York *Quarterly Review* 9 (Spring 1984), 7–16.

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Asseery, A., and David A. Peel, "The Effects of Exchange Rate Volatility on Exports," *Economics Letters* 37 (1991), 173–177.

Bailey, Martin J., George S. Tavlas, and Michael Ulan, "Exchange Rate Variability and Trade Performance: Evidence for the Big Seven Industrial Countries," *Weltwirtschaftliches Archiv* 122 (1986), 466–477.

Coes, Donald V., "The Crawling Peg and Exchange Rate Uncertainty," in John Williamson (ed.), *Exchange Rate Rules: The Theory, Performance and Prospects of the Crawling Peg*, (New York: St. Martins, 1981).

Cushman, David O., "The Effects of Real Exchange Rate Risk on International Trade," *Journal of International Economics* 15 (Aug. 1983), 45–63.

Davidson, J. E. H., D. J. Hendry, F. Srba, and S. Yeo, "Econometric Modelling of the Time Series Relation between Consumers' Expenditure and Income in the U.K.," *The Economic Journal* 88 (1978), 661–692.

DeGrauwe, P., "Exchange Rate Variability and the Slowdown in Growth of International Trade," *IMF Staff Papers* 35 (1988), 63–84.

Dickey, D. and W. Fuller, "Distribution of the Estimators for Autoregressive Time Series with a Unit Root," *Journal of the American Statistical Association* 74 (1979), 427–431.

Edison, Hali, and Michael Melvin, "The Determinants and Implications of the Choice of an Exchange Rate System," in William Haraf and Thomas Willett (eds.), *Monetary Policy for a Volatile Global Economy* (Washington, D.C.: The AEI Press, 1990).

Engle, R., and C. W. J. Granger, "Cointegration and Error Correction: Representation, Estimation and Testing," *Econometrica* 55 (1987), 251–276.

Ericsson, Neil R., "Cointegration, Exogeneity, and Policy Analysis: An Overview," Board of Governors of the Federal Reserve System *International Finance Discussion Papers*, No. 415 (Nov. 1991).

Franke, G., "Exchange Rate Volatility and International Trade," *Journal of International Money and Finance* 10 (June 1991), 292–305.

Giovannini, A., "Exchange Rates and Traded Goods Prices," *Journal of International Economics* 24 (1988), 45–68.

Gotur, Padma, "Effects of Exchange Rate Volatility on Trade: Some Further Evidence," International Monetary Funds *Staff Papers*, 32 (1985), 475–512.

Hickok, Susan A., Juann H. Hung, and Kurt C. Wulfekuhler, "An Expanded Cointegrated Model of U.S. Trade," Federal Reserve Bank of New York *Research Paper*, No. 9121 (June 1991).

Hooper, Peter, and Steven Kohlhagen, "The Effect of Exchange Rate Uncertainty on the Prices and Volumes of International Trade," *Journal of International Trade* 8 (Nov. 1978), 483–511.

International Monetary Fund, *Exchange Rate Volatility and World Trade*, Occasional Paper No. 28, Washington, D.C., (July 1984).

Johansen, Søren, "Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Regressive Models," *Econometrica* 59 (1991), 1551–1580.

Johansen, Søren and K. Juselius, "Maximum Likelihood Estimation and Inference on Cointegration—with Application to the Demand for Money," *Oxford Bulletin of Economics and Statistics* 52 (1990), 169–210.

Kenen, P. and D. Rodrik, "Measuring and Analyzing the Effects of Short-term Volatility in Real Exchange Rates," this REVIEW 68 (1986), 311–315.

Koray, Faik, and William D. Lastrapes, "Real Exchange Rate Volatility and U.S. Bilateral Trade: A VAR Approach," this REVIEW 71 (Nov. 1989), 708–712.

Kumar, Vikram, and Joseph Whitt, "Exchange Rate Variability and International Trade," Federal Reserve Bank of Atlanta *Economic Review* 77 (May/June, 1992), 17–32.

Kwiatkowski, D., P. Phillips, P. Schmidt, and Y. Shin, "Testing the Null Hypothesis of Stationarity Against the Alternative of a Unit Root," *Journal of Econometrics* 54 (1992), 159–178.

Lastrapes, William D., and Faik Koray, "Exchange Rate Volatility and U.S. Multilateral Trade Flows," *Journal of Macroeconomics* 12 (Summer 1990), 341–362.

Lutkepohl, Helmut, "Non-causality Due to Omitted Variables, *Journal of Econometrics* 44 (Aug. 1982), 367–378.

Makin, John H., "Eurocurrencies and the Evolution of the International Monetary System," in Carl H. Stem, John Makin, and Dennis Logue (eds.), *Eurocurrencies and the International Monetary System*, Washington, D.C. (1976), 17–53.

Mark, Nelson, "Real and Nominal Exchange Rates in the Short-Run and the Long-Run: An Empirical Investigation," Ohio State University, mimeo (1986).

Nelson, Charles R., and Charles I. Plosser, "Trends and Random Walks in Macroeconomic Time Series: Some Evidence and Implications," *Journal of Monetary Economics* 10 (Sept. 1982), 139–162.

Puzo, Susan, "Conditional Exchange Rate Volatility and the Volume of International Trade: Evidence from the Early 1900s," this REVIEW 74 (1992), 325–329.

Savvides, Andreas, "Unanticipated Exchange Rate Variability and the Growth of International Trade," *Weltwirtschaftliches Archiv*, 128 (1992), 446–463.

Sawa, Takamitsu, "Information Criteria for Discriminating among Alternative Regression Models," *Econometrica* 46 (Nov. 1978), 1273–1291.

Sercu, P., and C. Vanhulle, "Exchange Rate Volatility, Exposure and the Value of Exporting Firms," *Journal of Banking and Finance*, 16 (Feb. 1992), 155–182.

Thursby, Jerry G., and Marie C. Thursby, "Bilateral Trade Flows, the Linder Hypothesis, and Exchange Risk," *Review of Economics and Statistics* 69 (Aug. 1987), 488–495.

Thursby, Marie C., and Jerry G. Thursby, "The Uncertainty Effects of Floating Exchange Rates: Empirical Evidence on International Trade Flows," in Arndt et al., *Exchange Rates, Trade and the U.S. Economy* (Cambridge, MA: Ballinger Publishing Co., 1985), 153–166.

Turnovsky, Stephen J. and Mark E. Wohar, "Alternative Modes of Deficit Financing and Endogenous Monetary and Fiscal Policy in the U.S.A., 1923–82," *Journal of Applied Econometrics* 2 (1987), 1–25.

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