

Real Effective Exchange Rates and Export Adjustment in the U.S.

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The persistence of trade deficits despite significant depreciation of the U.S. dollar since the mid-1980s has raised questions about the impact of exchange rates on international trade. This article investigates how U.S. export is dynamically associated with foreign income and the real effective exchange rate of the U.S. using cointegration and error correction models. Cointegration results indicate a direct relationship between exports and foreign income and an inverse relationship between U.S. exports and real exchange rates. The error correction model indicates a significant short-run relationship between changes in exports and changes in foreign income.

INTRODUCTION

As the U.S. trade deficit reached unprecedented levels in the first half of the 1980s, different explanations emerged for the cause and persistence of the deficit. The traditional elasticities approach ascribed the widening of the deficit to the appreciation of the dollar and differences in the growth rates between the U.S. and its trading partners. Bryant and Holtman (1987), Helkie and Hooper (1987), and Krugman and Baldwin (1987), for example, suggest that the rise in the value of the dollar accounted for most of the deterioration in the trade balance.

An important challenge to the conventional analysis of trade deficits was mounted by Bergsten and Cline (1985), Mundell (1987), and Mackinnon and Ohno (1986). They relate the trade balance to the difference between national income and national expenditure or, equivalently, to the difference between saving and investment. This approach suggests the irrelevance of exchange rates as a factor in determining the current account equilibrium. The results of these studies, however, rest on the strong assumption that changes in nominal exchange rates do not have a lasting influence on relative prices.

Feldstein (1987), Hutchinson and Piggot (1984), and Laney (1984) suggest that the primary reason for the deteriorating trade balance is the U.S. budget deficit. The persistence of large budget deficits is thought to have caused the long term real interest rate differential and an appreciation of the dollar. The appreciation of the dollar dramatically

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increased the price of American products relative to foreign products leading to a decline in the volume of U.S. exports and an increase in imports.

An alternative explanation rooted in the modern theory of trade balance determination (Greenwood, 1984; Razin, 1984; and Hill, 1990) suggests that the main reason for the external trade imbalances is intertemporal shocks that shift the time distribution of consumption and production. It is argued that these disturbances affect both the exchange rate and trade balance. One implication of the modern theory is that the conventional theory may exaggerate the importance of exchange rates as factors initiating change in the size of the trade balance. According to this theory, the real exchange rate initially changes to accommodate the existing imbalances in trade. This in turn induces subsequent changes in the trade balance. Thus, as Hill (1990) argues, the relationship between the exchange rate and the trade balance may be bidirectional.

The persistence of trade deficits despite the sharp depreciation of the dollar in the second half of the 1980s posed serious questions to the conventional exchange rate explanation of trade balance. Helkie and Hooper (1987), Krugman and Baldwin (1987), and Rosenweig and Koch (1988) address the persistence of the deficit and conclude that it reflects, for the most part, normal lags in the adjustment to a depreciation of the dollar that followed a long period of appreciation.

Others suggest that the current empirical dollar indices are flawed in their construction and hence have overstated the depreciation of the dollar. They argue that while the dollar had depreciated sharply against the currencies of Japan and western Europe, the currencies of many LDCs had depreciated against the dollar. To address these perceived shortcomings, numerous new dollar indices have been proposed. For example, Rosensweig (1986), Cox (1987), and Harvey and Strauss (1987) have constructed alternative indices to measure the foreign exchange value of the dollar.

The wide disparity of views on the relationship between exchange rate changes and trade balance points to the need for further study of the issue. Most of the earlier studies specify trade models in levels or in log level form. These models have been criticized because the levels and log levels of many economic variables in trade models are nonstationary. The regression equation relating such variables could lead to spurious regressions, a phenomenon first described in Granger and Newbold (1986). This phenomenon refers to the possibility that inferences based on the ordinary least square parameter estimates in such models are invalid because *t*- and *F*-ratio test statistics do not converge to their limiting distribution as the sample size increases. In this case the null hypothesis of no relation would be rejected wrongly as discussed in Engels and Granger (1987).

This paper examines the relationship between the real effective exchange rate and U.S. exports using cointegration and error correction approaches.

DATA AND ESTIMATION METHOD

The export demand model is basically like any other demand model. Price and quantity are inversely related, *ceteris paribus*, with equilibrium price and quantity determined by the interaction of supply and demand. In most empirical studies own price is assumed exogenous, i.e., supply is perfectly elastic. Thus, the export supply equation is not explicitly considered in many trade models. (See Murray and Ginnman, 1976; Houthakker and Magee, 1969; Warner and Kreinen, 1983; and Krugman, 1989.)

The assumption of infinite supply elasticity reduces the export model to a single equation. Following the conventional demand theory, the traditional export demand equation relates the quantity of exports demanded to income and relative prices for U.S. goods and foreign goods (the real exchange rate). The notation is as follows:

- $\ln x_t$ = Log of an export volume index;
 $\ln e_t$ = Log of a real effective exchange rate index; and
 $\ln y_t^f$ = Log of foreign real income.

The export model is specified as:

$$(1) \ln x_t = \beta_0 + \beta_1 \ln e_t + \beta_2 \ln y_t^f + \mu_t$$

The data used for this study are from the International Monetary Fund's *International Financial Statistics* (IFS) CD-ROM, September 1994. Because there are no real GDP data for some of the U.S. trading partners, the weighted industrial production index of 13 major importers of U.S. goods is used as a proxy for real income. The countries included in the index are Australia, Canada, U.K., Belgium, France, Germany, Italy, Netherlands, Japan, Hong Kong, Korea, Singapore, and Mexico. The weight for an individual country is based on its share of U.S. exports. The data for export shares are obtained from the International Monetary Fund's *Direction of Trade Statistics Year Book*, 1994 (pp. 420-422).

Determining the relevant set of economic variables that determine the sustainable exports of a country involves identifying and estimating the long-run relationship between the volume of exports and its determinants. The long-term comovement between a set of time series variables is usually detected by cointegration techniques suggested by Granger (1986) and Engel and Granger (1987). A more powerful test that allows for the detection and estimation of the number of cointegrating vectors, however, was developed by Johansen (1988) and Johansen and Juselius (1990) in the context of a vector error correction model (VECM).

In the Johansen procedure, maximum likelihood is applied to an autoregressive representation of the form given by:

$$(2) \begin{bmatrix} \Delta \ln x_t \\ \Delta \ln e_t \\ \Delta \ln y_t^f \end{bmatrix} = \Gamma(L) \begin{bmatrix} \Delta \ln x_{t-1} \\ \Delta \ln e_{t-1} \\ \Delta \ln y_{t-1}^f \end{bmatrix} + \Pi \begin{bmatrix} \ln x_{t-1} \\ \ln e_{t-1} \\ \ln y_{t-1}^f \end{bmatrix} + \begin{bmatrix} \mu^x \\ \mu^e \\ \mu^{y^f} \end{bmatrix}$$

$\Gamma(L)$ is a 3×3 matrix of polynomials in the lag operator, which shifts the series back in time, that is $Ly_t = y_{t-1}$.

The cointegration test focuses on the properties of the matrix coefficient Π . In the absence of cointegration, Π is a singular matrix (i.e., its rank $r = 0$). Hence, in the export model the rank Π could be anywhere between zero, if no cointegrating vector exists, and three, the number of variables in the system. The two commonly used tests of cointegration are $\lambda_{\max}$ and Trace tests. According to $\lambda_{\max}$, the null hypothesis is that there are r or

less cointegrating vectors. The alternative hypothesis is that there are $r + 1$ cointegrating vectors. If $\lambda_{\max}$ exceeds the critical value tabulated under the null hypothesis, we can reject the null hypothesis in favor of the alternative. The Trace test has the same null hypothesis as the $\lambda_{\max}$ test; however, the alternative hypothesis is that the rank $\Pi \geq r + 1$ (Dickey *et al.*, 1991; Mehra, 1993).

The basic idea of cointegration is that two or more variables may be regarded as defining a long-run relationship even though they may drift apart in the short run. This long-run relationship is referred to as a cointegrating vector. Because there is a long-run relationship between the variables, a regression containing levels of all variables of cointegrating vector will have stationary error term, even if none of the variables taken alone is stationary.

Before estimating the cointegration parameters, the order of integration of each series should be examined. The common practice recommended by Engel and Granger (1987) is to use the augmented Dickey Fuller test (ADF). The ADF test for a unit root is performed by regressing the first difference of the variable on its own level lagged one period, a constant, and lagged first difference of the series:

$$(3) \Delta Z_t = A_0 + A_1 \text{trend} + A_2 Z_{t-1} + \sum_{j=1}^k \psi_j \Delta Z_{t-j} + \varepsilon_t$$

where:

- Z_t = The pertinent variable;
- ε = A random disturbance term; and
- k = The number of lags of the first difference on Z necessary to make ε serially uncorrelated.

If $A_2 = 0$, Z_t has a unit root. The null hypothesis $A_2 = 0$ is tested using t-statistics. The cumulative distribution of the ADF test statistic is provided by Mackinnon (1991). The optimal lag length for ADF regression is determined by adding lags until the Ljung-Box test does not reject the null hypothesis of no autocorrelation.

The Granger representation theorem states that if a cointegrating relationship exists, then a dynamic error correction representation of the data also exists. According to Engel and Granger (1987), the following error correction model (ECM) is estimated with all nonstationary variables in the cointegration equation 1:

$$(4) \Delta \ln x_t = \alpha + \sum_{i=1}^{k_1} \beta_i \Delta \ln x_{t-i} + \sum_{i=0}^{k_2} \delta_i \Delta \ln c_{t-i} + \sum_{i=0}^{k_3} \psi_i \Delta \ln y_{t-i} + \lambda \mu_{t-1} + v_t$$

Where all the variables are defined above, the disturbance term is v_t ; Δ is the first difference operator; k_j ($j = 1, 2, 3$) represents the various lags on regressors and μ_{t-1} is the error correction (one lagged error) generated from Johansen multivariate procedure (see Table 4) as in Arize (1995) and Sedgley and Smith (1994).

Equation (4) gives both the short-run and long-run relationships. The long-run relationship in the model is captured by the lagged value of the long-run random disturbance

Table 1—Unit Root and Mean Stationary Test Results

Variable	ADF (k) Test Statistics	
	Without Trend	With Trend
$\ln x_t$	-0.1814 (k = 2)	-1.8028 (k = 4)
$\Delta \ln x_t$	-15.0795**	-5.9901**
$\ln e_t$	-0.6243	-1.1659
$\Delta \ln e_t$	-6.2170**	-6.2674**
$\ln y_t$	-1.8503	-1.3328
$\Delta \ln y_t$	-5.1390**	-5.2655**

Note: The null hypothesis in the ADF test is that the time series contains a unit root (i.e., it is nonstationary) against an alternative of mean stationarity. The null hypothesis is rejected if the value of the test statistic is greater than the critical value calculated by Mackinnon (1991). The ADF (k) test statistics for a variable Z_t is given by the t-statistics on the estimated coefficient A_2 , in the following auxiliary regression:

$$\Delta Z_t = A_0 + A_1 \text{trend} + A_2 Z_{t-1} + \sum_{j=1}^k \psi_j \Delta Z_{t-j}$$

where Z is the pertinent variable and k is determined by the lowest order of lag length for which the Ljung-Box does not reject the null hypothesis of no autocorrelation in the residuals. Unless otherwise specified $k = 0$

** Significant at 1 percent

* Significant at 5 percent

term (μ_{t-1}). If this term were omitted, all variables would be in the first difference, and only short-run effects would be detected. (See Harvey, 1988.) The error correction factor plays a crucial role in the model. The coefficient of the lagged error term is an adjustment coefficient and represents the proportion by which the long-term disequilibrium in the dependent variable is corrected each short period (Harvey, 1988; Mehra, 19991). The F-test of the differencing explanatory variables indicates the short-term causal effects, whereas the long-term causal relation is implied through significance of the t-test of the lagged error-correction term.

EMPIRICAL RESULTS

The long-run and short-run interaction between a trivariable system consisting of exports ($\ln x_t$), the real effective exchange rates ($\ln e_t$), and foreign income ($\ln y_t$) is examined. The data are quarterly and cover the sample period 1975Q1 to 1993Q2. The sample period is restricted to the flexible exchange rate period for which data are available for all relevant variables.

Tests for unit roots and mean stationarity are presented in Table 1. The null hypothesis of the ADF test is that the series is nonstationary. It is evident from the ADF test results that the series $\ln x_t$, $\ln e_t$, and $\ln y_t$ are nonstationary in levels and stationary in first differences. These results indicate that all the variables are $I(1)$. This satisfies the condition that all the variables should have the same order of integration to be cointegrated.

I now examine whether a long-run equilibrium relationship between the series exists using the Johansen-Juselius multivariate procedure. As can be seen from Table 2, the null hypothesis of zero cointegration ($H_0: r = 0$) is rejected both by Trace and $\lambda_{\max}$ statistics. The hypothesis of at most one cointegrating vector ($H_0: r \leq 1$) cannot be rejected.

Table 2—Johansen Multivariate Cointegration Test Results

Trace				λ_{max}			
H_0	H_1	Stat	90%	H_0	H_1	Stat	90%
$r = 0$	$r \geq 1$	29.32	26.70	$r = 0$	$r = 1$	19.82	13.39
$r \leq 1$	$r \geq 2$	9.42	13.31	$r \leq 1$	$r = 2$	5.96	10.60
$r \leq 2$	$r \geq 3$	3.5	2.7	$r \leq 2$	$r = 3$	3.53	2.71

Note: H_0 is the null hypothesis, H_1 is the alternative hypothesis, and r is the number of cointegration vectors. 90 percent critical values for the Johansen test are from Hensen and Juselius CATS in RATs computer procedure for cointegration

Table 3—Johansen Cointegration Equation Parameter Estimates and Exclusion Restriction Tests

$$\text{Model: } c + \beta_1 \ln x_t + \beta_2 \ln e_t + \beta_3 \ln y_t^f + \mu_t \sim I(0)$$

Variable	Parameter Estimates		Test for Exclusion		
	Unrestricted	Normalized	H_0	LR(1) ¹	LR(2)
c	0.680	0.488			
β_1	15.5082	1.0000	$\beta_1 = 0$	7.25**	9.47**
β_2	13.5045	0.871	$\beta_2 = 0$	12.18**	13.01**
β_3	-22.0675	-1.423	$\beta_3 = 0$	7.43**	9.54**

Note: $\ln x_t$ is the measure of U.S. export volume index, 1990 = 100, $\ln e_t$ is the measure of effective real exchange rates, and $\ln y_t^f$ is real foreign income. The likelihood ratio test statistic LR(k) is distributed as $\chi^2(rk)$ where k is the number of restrictions and r is the number of cointegrating vectors

** Significant at 1 percent

The cointegrating vector corresponding to the dominant long-run relationship is reported in Table 3. Cointegration results indicate a direct relationship between U.S. exports and foreign income and an inverse relationship between exports and real exchange rates in the U.S. The cointegrating vector is normalized with respect to exports. Parameter estimates of the normalized vector indicate that exports are elastic with respect to foreign income. The volume of exports, however, is inelastic with respect to exchange rates.

To test the significance of the elasticities I employ the likelihood ratio test statistics. Both real exchange rate and foreign income elasticities are significant at the 1 percent level with $\chi^2(1)$ test statistics of 12.18 and 7.43, respectively.

Having established that the foreign income, real effective exchange rate, and export volume index are cointegrated, I next examine the short-run dynamic interaction between these variables using the error correction model. Implementing an error correction model is necessary to examine the short-run dynamics of the model. Akaike's (1969) FPE criterion is used to determine the lag length. The results of the error correction model are reported in Table 4.

It is necessary to ensure that the residuals in the error correction models are serially uncorrelated. A Lagrangian multiplier (LM) test statistic is used to test the null hypothesis that the first "p" autocorrelations are zero. The LM test is recommended when the regression involves, as all error correction models do, a lagged dependent variable (Harvey, 1991). This test statistic indicates there is no serial correlation.

Table 4—Error Correction Model Estimates

$$\Delta \ln x_t = \alpha + \sum_{i=1}^k \beta_i \Delta \ln x_{t-i} + \sum_{i=0}^k \delta_i \Delta \ln e_{t-i} + \sum_{i=0}^k \psi_i \Delta \ln y_{t-i}^f + \lambda \mu_{t-1}$$

α	0.404 (4.590)
$\Sigma \delta \ln x$ (k = 8)	0.3788 [11.508]**
$\Delta \ln e$ (k = 0)	0.0504 (0.679)
$\Delta \ln y_t^f$ (k = 0)	1.1029** (2.716)
μ_{t-1}	-0.240** (-4.621)

 $R^2 = 0.68$

S.E.E = 0.0327

 $F(11, 53) = 11.467$ p = 0.000 $LM[\chi^2(2)] = 2.709231$ p = 0.3694Chow Test $F(12, 41) = 1.12913$ p = 0.6076

Note: An F-test is used to test the joint significance of lags in the first difference of export volume. The number in a bracket is the F- value. The numbers in parenthesis below the coefficient values are t-values. The values of k indicate the lag length selected using FPE criteria

** Significant at 1 percent

This table shows estimates of the coefficients of the error correction model. The symbols are defined as follows:

- Δ = The first difference
- $\ln x_t$ = Log of export volume index
- $\ln e_t$ = Log of real effective exchange rate index
- $\ln y_t^f$ = Log of foreign income

The data are from International Monetary Fund, *International Financial Statistics*, CD-ROM, September 1994. Weighted Industrial Production Index of 13 major importers of U.S. goods is used as a proxy for foreign income. The countries are Australia, Belgium, Canada, U.K., Italy, Netherlands, Japan, Hong Kong, Korea, Singapore, and Mexico. The weight of the individual country is determined by its share of U.S. exports. The data for export share are obtained from International Monetary Fund's *Direction of Trade Statistics Year Book*, 1994 (pp. 420-422)

The empirical results suggest that the statistical fit of the model to the data is satisfactory as indicated by the values of the adjusted R^2 , the standard error of the estimate (S.E.E), and the F value for testing the null hypothesis that all the right side variables as a group, except the constant term, have zero coefficient.

The results show a significant short- run relationship between changes in U.S. export volume and foreign income. The real exchange rate has the expected sign but it is not statistically significant. The error correction factor (λ) appears with a statistically significant coefficient and displays the appropriate negative sign, further confirming that the variables $\ln x_t$, $\ln e_t$, and $\ln y_t^f$ are cointegrated. Overlooking the cointegratedness of the variables may have introduced misspecification in earlier studies that use OLS regressions based on the levels or logs of the levels of the variables.

The magnitude of the error correction factor, λ , measures the single period response of the dependent variable to departure from equilibrium. The negative coefficient of the error correction factor also ensures the long run equilibrium is achieved. The adjustment toward equilibrium is not, however, instantaneous. The magnitude of the error correction factor ($\lambda = 0.24$) indicates that 24 percent of any quarter's deviation between the actual

export volume and the long-run equilibrium is incorporated into the next quarter growth of export volume.

The error correction model in Table 4 is further evaluated by examining the structural stability. The Chow test (Chow, 1960) is implemented using a dummy variable approach and a potential break point in 1985:III. The choice of the break point corresponds with the Plaza Accord during which the G-5 countries engaged in foreign exchange intervention. Table 4 shows that the F-statistic for the Chow test is not statistically significant and thus implies the regression reported in Table 4 does not depict parameter instability.

SUMMARY AND CONCLUSION

The results of the empirical analysis of the relationship between the exchange rate and U.S. exports from 1975:Q1 to 1993:Q2 may be summarized as follows:

- The cointegration tests based on the Johansen and Juselius (1990) reveal that the U.S. exports, the real effective exchange rate, and foreign income are cointegrated. This suggests that the traditional estimates based on the regression of levels or log levels of the series may not be reliable.
- The cointegration equation based on the Johansen and Juselius (1990) multivariate procedure indicates that both the real exchange rate and income have the expected signs. The real exchange rate is negatively related to the volume of exports while foreign income is positively related to exports. Moreover, exclusion tests on the parameters reveal that the coefficients of both exchange rate and foreign income are statistically significant. This is consistent with the results of Bryant and Holtman (1987), Helkie and Hooper (1987), and Krugman and Baldwin (1987) who argue that real exchange rate is an important determinant of trade.
- The short-run dynamics based on the error correction model show that the error correction term plays a significant role in modeling the U.S. exports. The results reveal that the effect of an exchange rate change on exports, however, is insignificant in the short run.

Implications of the findings are relevant to the current policy discussion on improving U.S. external balances. Real exchange rate depreciation, by making the products of the United States internationally more competitive, helps correct the trade deficit. Because persistent depreciation may lead to a lower standard of living (Feldstein, 1987), however, a lasting solution should focus more on increased labor productivity and quality improvements of U.S. goods and services.

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