JOURNAL OF ECONOMIC DEVELOPMENT
Volume 18, Number 2, December 1993

Does Exchange-Rate Volatility Deter Trade Volume of LDCs?

Mohsen Bahmani-Oskooee
and
Sayeed Payesteh*

In this paper we investigate the response of the trade flows of six less developed countries to exchange rate volatility using quarterly data over 1973-1990 period. Just like previous studies when standard econometric techniques were used, there was evidence of adverse effects of exchange rate uncertainty on the trade volume. However, application of cointegration technique rejected the notion of any long-run relationship between imports and exports and their determinants.

I. Introduction

One of the major economic issues that has received a great deal of attention by many researchers is the effects of exchange rate volatility on the trade volumes. According to published research, the empirical literature provides a mixed conclusion. Examples of studies that have shown that exchange rate volatility has adversely effected the trade flows, includes Cushman (1983, 1986, 1988), Akhtar and Hilton (1984), Kenen and Rodrik (1986), Thursby and Thursby (1987), Brada and Mendez (1988), and Peree and Steinherr (1989). On the other hand, Makin (1976), Hooper and Kohlhagen (1978), Gotur (1985), Bailey et al. (1986, 1987), Koray and Lastrapes (1989), and Lastrapes and Koray (1990) are examples of studies who have concluded that trade volumes are not sensitive to exchange-rate volatility. All these studies have investigated the response of industrial countries' trade volumes to exchange-rate volatility.

Less developed countries (LDCs) have received little attention.

* Professor and Associate Professor of Economics, respectively, University of Wisconsin-Milwaukee, Milwaukee, WI 53201, U.S.A.

Caballero and Corbo (1989) and Bahmani-Oskooee (1991) are two published works on the issue. The former authors estimated export demand equation for Chile, Colombia, Peru, Philippines, Thailand, and Turkey and concluded that exchange rate volatility has significantly negative effects on the exports of all these countries. The latter author estimated import as well as export demand equations for Brazil, Greece, Korea, Pakistan, Philippines, Thailand and Turkey and showed that exchange rate uncertainty had an adverse effect on the imports of Korea, Pakistan, Thailand, and Turkey. As for the response of exports, he showed that exchange rate volatility had negative effect on the exports of Greece and Turkey and positive effect on the exports of Brazil and Korea. In addition to these two published papers, there are three unpublished manuscripts that deserve to be mentioned. Kenen and Rodik (1986), who investigated the response of import volumes of industrial countries to exchange-rate volatility, have made a reference to three studies. Coes (1981) has looked at the Brazilian exports and has found that after adopting the crawling peg in 1968, a decrease in exchange rate uncertainty had positive effect on the Brazilian exports. The second study is an unpublished Ph.D. thesis by Gupta (1980) who analyzed the export supply of five LDCs and found that short-term volatility had negative effects on the export supply in two cases. Finally, there is a reference to Rana (1981) who found that exchange rate volatility had negative effects on the import volumes of South Korea, Taiwan, and the Philippines.

The issue is very important and deserves further attention. Especially, when a specific model is introduced into the literature, it is usually tested by drawing data from industrial countries due to availability of the data. It will be interesting to see how does that model perform for a less developed country (LDC). One such model is an import demand equation that was introduced by Kenen and Rodrik (1986). Therefore, it is the purpose of this paper to investigate the impact of exchange-rate volatility on the import volume of as many LDCs as the data permits, using the Kenen and Rodrik's model. Due to symmetry between import and export demand models, we also investigate the response of LDCs exports. Section II provides a brief explanation of the models. In Section III we present our empirical results. Section IV concludes. Data definition and sources are cited in the Appendix.

II. The Methodology

The common practice to analyze the effects of exchange-rate volatility on the import volume is to speficy an import demand equation which relates the volume of imports to a price term, a scale variable, and a measure of exchange-rate variability. To this end, we adopt the most simple model from the literature developed by Kenen and Rodrik (1986). Kenen and Rodrik (1986) who investigated the effects of exchange-rate volatility on the import volume of 11 industrial countries employed the following model:

$$\text{(1)} \qquad \text{Log } M_t = \beta_0 + \beta_1 \text{ Log } R_{t-1} + \beta_2 \text{ Log } Y_t + \beta_3 \text{ T} + \beta_4\, \sigma_t^R$$

where M_t = import volume at time t,
R_{t-1} = import-weighted real effective exchange-rate at time t-1,
Y_t = index of economic activity of the importing country at time t,
T = trend term,
σ_t^R = measure of exchange-rate volatility.

In equation (1), β_1 is expected to be negative. This is due to the fact that as the Appendix indicates, we have defined R as number of units of domestic currency per unit of foreign currency. An increase in R which reflects depreciation of domestic currency is expected to discourage imports. Since a growing economy is expected to import more, the Y variable is expected to carry a positive coefficient, i.e., $\beta_2 > 0$. However, as discussed by Khan (1974) and Khan and Ross (1975), if increase in domestic activity is due to production of import-substitute goods, then imports may decline as Y increases, yielding a negative β_2. The trend term could carry a positive or negative coefficient. Finally, if volatility in exchange rate, σ, is to depress the LDCs imports, β_4 is expected to be negative.

When we tried to estimate equation (1), we encountered some cases for which the real exchange rate variable did not carry significant coefficient. In such circumstances, following the methodology of Kenen and Rodrik (1986), we too imposed an Almon lag structure on the real effective exchange rate and estimated the following model:

$$\text{(2)} \qquad \text{Log } M_t = \alpha_0 + \sum_{k=0}^{n} \delta_k \text{ Log } R_{t-k} + \alpha_2 \text{ Log } Y_t + \alpha_3 \text{ T} + \alpha_4\, \sigma_t^R$$

Furthermore, when Almon lag procedure outlined by equation (2) did not yield a significant real effective exchange-rate elasticity, we replaced R variable by the ratio of import price to domestic price level and estimated one of the following two equations.

(3) $$\text{Log } M_t = b_0 + b_1 \text{ Log}(PM/PD)_{t-1} + b_2 \text{ Log } Y_t + b_3 T + b_4 \sigma_t^R$$

or

(4) $$\text{Log } M_t = a_0 + \sum_{k=0}^{m} \gamma_k \text{Log}(PM/PD)_{t-k} + a_2 \text{ Log } Y_t + a_3 T + a_4 \sigma_t^R$$

where PM is the index of import price and PD, the index of domestic prices. Indeed, most other studies like Houthakker and Magee (1969), Warner and Kreinin (1983), and Bahmani-Oskooee (1986) have used the price ratio in their models.

As indicated in the introductory section, we also try to apply Kenen and Rodrik's methodology to the export volume. Due to symmetry between import and export demand equations, we replace the import volume M in equation (1) by the volume of exports X and the domestic income Y, by world income YW. Thus, the export demand equation that we propose to estimate takes the following form:

(5) $$\text{Log } X_t = \beta_0' + \beta_1' \text{ Log } R_{t-1} + \beta_2' \text{ Log } YW_t + \beta_3' T + \beta_4' \sigma_t^R$$

In equation (5) β_1' is expected to be negative. As the Appendix indicates, in this export demand equation R is defined as number of foreign currencies per unit of domestic currency. Thus, a decrease in R or a depreciation of domestic currency is expected to stimulate exports. World income elasticity is also expected to be positive indicating that an increase in world income should help LDCs to export more. Finally, if exchange rate uncertainty is to deter export volume, β_4' is expected to be negative.

Like import demand equation if one period lagged exchange rate did not yield significant elasticity, we imposed an Almon lag structure on the real effective exchange rate and estimated the following model:

(6) $$\text{Log } X_t = \alpha_0' + \sum_{k=0}^{n} \delta_k' \text{ Log } R_{t-k} + \alpha_2' \text{ Log } YW_t + \alpha_3' T + \alpha_4' \sigma_t^R$$

Furthermore, when Almon lag procedure outlined by equation (6) did not yield a significant real effective exchange-rate elasticity, we replaced R variable by the ratio of export price to world export price level and estimated one of the folloiwng two equations.

(7) $$\text{Log } X_t = b_0' + b_1' \text{ Log}(PX/PXW)_{t-1} + b_2' \text{ Log } YW_t + b_3' T + b_4' \sigma_t^R$$

or

$$\text{(8)} \qquad \text{Log } X_t = a'_0 + \sum_{k=0}^{m} \gamma'_k \text{ Log}(PX/PXW)_{t-k} + a'_2 \text{ Log } Y_{wt} + a'_3 \text{ T} + a'_4 \ \sigma_t^R$$

where PX is the index of export price and PXW, the index of world export price. It is expected that b'_1 and a'_k to be negative.

III. The Estimation Results

The models outlined in Section II were estimated for six LDCs using quarterly data over 1973-1990 period. These are the only countries for which data on all variables for the 1973-90 period were available. As the Appendix indicates, following Lanyi and Suss (1982), the variability measure of real effective exchange-rate is defined as the standard deviation of quarterly percentage changes in the real effective exchange rate. To generate time series data on this variable, for each quarter we constructed it over the 8 previous quarters ending at the current quarter. Following Kenen and Rodrik, the Almon lags, whenever necessary, were generated from a third degree polynomial without end point constraint. Maximum of 15 lags were tried and again following Kenen and Rodrik (1986, p. 313) optimum number of lags, thus, equations were chosen for the quality of their price and activity terms, rather than the sign or significance of their volatility terms, and for overall goodness of fit. Tables 1 and 2 show that results and name of the countries involved.

Concentrating on Table 1 which displays the estimates of import demand equations, we gather that exchange-rate volatility exerts significantly negative effect on the import volume of Pakistan, the Philippines, and Singapore. Kenen and Rodrik (1986) who applied equation (1) or (2) for 11 developed countries (DCs), only obtained significantly negative coefficient for exchange-rate variability in four cases. It appears that the model and the methodology introduced by Kenen and Rodrik, has a higher success rate for LDCs (50%) than DCs (36%). Note that in two cases (Greece and Korea) the real effective exchange rate did not yield significant coefficient estiamtes, whereas, the relative prices did. The differential response of trade flows to a change in relative prices and to a change in exchange rate goes back to Orcutt (1950). Junz and Rhomberg (1973) interpreted Orcutt's conjecture by arguing that the response of trade flows to a change in exchange rate could be faster than their response to a change in relative prices

Table 1

COEFFICIENT ESTIMATES OF THE IMPORT DEMAND EQUATION

	Exogenous Variables							
		Price Term						
Country	Constant	Simple Lag	Almon Lag[a]	Domestic Income	Trend	Exch.-Rate Variability	$\bar{R}^2$	D-W
Greece	−1.7128	−0.8094[b]	—	1.0131	0.0040	0.1321	0.61	1.9618
	(0.71)	(3.91)		(2.98)	(3.25)	(0.11)		
Korea	5.8063	−0.8064[b]	—	0.1847	0.0204	1.7153	0.96	2.1897
	(12.2)	(4.18)		(3.61)	(5.99)	(1.14)		
Pakistan	−3.0352	−0.0023	—	1.1421	−0.0049	−1.3291	0.87	1.7143
	(0.91)	(2.14)		(3.81)	(0.96)	(2.92)		
Philippines	7.8806	—	−1.1780	0.4680	0.0080	−4.8461	8.93	1.4396
	(7.44)		(3.37)	(2.22)	(3.67)	(3.44)		
Singapore	6.5773	0.0067	—	0.1899	0.0144	−1.7749	0.98	1.9027
	(5.28)	(6.92)		(1.27)	(5.41)	(1.70)		
S. Africa	3.4182	−0.3058	—	0.6360	0.0030	−0.9863	0.96	1.7661
	(2.17)	(1.61)		(4.69)	(1.32)	(11.18)		

Notes: Numbers inside the parentheses are the absolute value of t-ratios.

a. Sum of Almon lag coefficients and t-ratio for the sum.

b. The ratio of import price to domestic price is used in these cases.

because of publicity that surrounds around the change in exchange rates. On the other hand, it could be slower because of the large resource shifts necessary to correct any disequilibrium. While these arguments could be applied for a given country, the differential response across countries could be attributed to the structure as well as different macro policies. For example, more trade oriented countries may respond to a change in exchange rate faster than less trade oriented countries.

Four of these countries were also included in Bahmani-Oskooee (1991) who investigated the effects of exchange-rate volatility on the trade flows of seven LDCs, using somewhat different model and data over shorter period of time (1975-1985). Our findings for Greece and Pakistan are consistent with those of Bahmani-Oskooee (1991). However, for Korea and the Philippines the results are inconsistent. The negative and significant results for the Philippines in this paper versus insignificant results in Bahmani-Oskooee (1991) cold be due to larger sample size and due to the use of Almon lag procedure to obtain a significant price term. The insignificant coefficient of exchange rate variability for Korea in this paper versus the significant coefficient in Bahmani-Oskooee could only due to larger sample size and the differences in the models.

Concentrating on the other features of the results, Table 1 reveals that the price term carries its significant negative coefficient in all but one case. This is perhaps the only study that produces a negative and significant price elasticity for all LDCs. Again, one has to attribute this to Kenen and Rodrik's suggestion of using lagged terms. The domestic income also carries its positive and significant coefficient in all but one case.

Turning now to Table 2 which displays the coefficient estimates of export demand equations, we gather that the results are similar to those of Table 1. More precisely exchange rate variability carries significantly negative coefficient in three out of six countries. Once again we observe a reversal in the results for Greece, Korea, and the Philippines in this paper as compared to the results reported in Bahmani-Oskooee (1991, Table 2) which could be attributed to different models and different time span over which data are collected. Two other features of the results in Table 2 also deserve mention. Just like import demand equation, the price term carries its expected significant negative coefficient in all but one case. However, the world income carries a significant coefficient only in the cases of Korea and Singapore (two of the four NICs countries) indicating that an increase

Table 2

COEFFICIENT ESTIMATES OF THE EXPORT DEMAND EQUATION

	Exogenous Variables							
		Price Term						
Country	Constant	Simple Lag	Almon Lag[a]	World Income	Trend	Exch.-Rate Variability	$\bar{R}^2$	D-W
Greece	8.2255	-1.2104[b]	—	0.5002	0.0046	0.5418	0.82	1.9289
	(3.59)	(3.00)		(0.91)	(1.24)	(0.51)		
Korea	8.7927	-1.2588[b]	—	1.1001	0.0194	4.0085	0.98	1.9601
	(5.31)	(11.0)		(2.38)	(5.39)	(3.61)		
Pakistan	10.765	—	-0.6048	0.1282	0.0173	-0.9011	0.85	1.7684
	(2.28)	—	(2.90)	(0.13)	(3.21)	(0.31)		
Philippines	8.9738	-0.2347	—	0.3824	0.0111	-2.0526	0.89	1.4147
	(5.96)	(2.31)	—	(1.14)	(5.06)	(3.34)		
Singapore	3.9271	-0.3660	—	1.3921	0.0811	-0.9742	0.99	2.0453
	(2.66)	(3.29)		(4.76)	(10.2)	(0.85)		
S. Africa	6.8834	0.0992	—	0.2812	0.0103	-1.7803	0.96	2.1109
	(2.22)	(0.54)		(0.43)	(2.38)	(2.09)		

Notes: Numbers inside the parentheses are the absolute value of t-ratios.

a. Sum of Almon lag coefficients and t-ratio for the sum.

b. The ratio of export price to world export price is used in these cases.

in world income stimulates exports of Korea and Singapore.

With the unit roots that exist in almost all macroeconomic variables, one could criticize our findings in this paper as well as all others in the literature on the ground that they all suffer from what has become known as "spurious regression problem." The problem is that since most macro time series variables contain a unit root or they are non-stationary, the t-ratios cannot be used to judge their significance. The cointegration analysis developed by Engle and Granger (1987) and others tries to remedy this problem. In applying the cointegration technique, we first need to determine the degree of integration of each variable. A variable is said to be integrated of order d or I(d), if it achieves stationarity after being differenced d times. Engle and Granger (1987) show that two or more I(d) variables could be cointegrated if the residuals (as a proxy for a linear combination among variables) in the simple OLS regression of one variable on the others are integrated at any order less than d. For example two or more I(1) variables are cointegrated if the residuals are I(0).

To provide some preliminary results we first test for the stationarity of each variable using the ADF test. The ADF test for a time series Z_t is formulated by equation (9) below:

$$(9) \qquad \Delta Z_t = \alpha + \beta t + \sigma Z_{t-1} + \sum_{i=1}^{k} \tau_i \Delta Z_{t-i} + w_t$$

where $\Delta Z_t = Z_t - Z_{t-1}$, t is a trend term and w is an error term. The ADF test statistic is calculated as the ratio of the estimate of σ to its standard error. The cumulative distribution of the ADF test statistic with and without the trend term for any sample size is provided by Mackinnon (1991). If the calcualted statistic is less than its critical value, then Z is said to be stationary or I(0). In selecting the number of lags in (9) one common practice is to rely upon the significance of lagged first differenced terms using standard t-test. The results of the ADF tests for the stationarity of the level of each variable as well as for their first differences are reported in Table 3.

It is evident from Table 3 that using 5% critical value all variables are I(1) only in the cases of Greece, Singapore and South Africa. In the cases of Korea and Pakistan, exchange rate volatility, σ_R, is I(0) due to the fact that our calculated statistic is less than critical value when the ADF test is applied to the level of the variable. In the results for the Philippines Log X and Log (PX/PXW) are I(0) and σ_R is probably I(2). All in all since for these three countries all variables are not integrated of the same order, the cointegration analysis cannot be applied. For

Table 3

THE RESULTS OF ADF TEST APPLIED TO THE LEVEL AS WELL AS FIRST DIFFERENCED VARIABLES[a]

Variable	Greece	Korea	Pakistan	Philippines	Singapore	S. Africa
Log M	-2.41(3)[b]	-1.70(1)	-1.05(3)	-0.91(2)	-1.88(4)	-3.30(1)
Log R	-3.16(1)	-2.94(1)	-3.32(1)	-2.49(2)	-3.34(4)	-2.67(3)
Log(PM/PD)	-2.15(4)	-2.87(2)	-2.76(1)	-2.95(1)	-1.77(4)	-2.68(4)
Log Y	-2.04(4)	-2.39(4)	-1.85(4)	-2.16(4)	-1.97(4)	-2.15(3)
σ_R	-3.12(1)	-3.57(4)	-3.47(1)	-2.72(4)	-2.09(1)	-1.72(1)
Log X	-3.27(3)	-1.33(4)	-2.74(3)	-3.49(1)	-3.02(4)	-1.88(4)
Log(PX/PXW)	-3.24(4)	-2.39(1)	-1.74(1)	-6.18(3)	-2.07(1)	-2.59(2)
Log YW	-3.31(1)	-3.31(1)	-3.31(1)	-3.31(1)	-3.31(1)	-3.31(1)
ΔLog M	-9.50(2)	-5.84(4)	-13.2(2)	-9.87(1)	-3.75(3)	-5.40(1)
ΔLog R	-6.73(1)	-5.19(4)	-10.3(4)	-5.60(4)	-3.66(4)	-3.62(2)
ΔLog(PM/PD)	-6.20(3)	-5.44(4)	-5.13(3)	-6.31(1)	-4.93(3)	-3.74(3)
ΔLog Y	-5.75(3)	-4.76(3)	-5.58(4)	-3.99(3)	-4.64(3)	-7.93(1)
$\Delta\sigma_R$	-5.69(1)	-3.51(2)	-6.15(1)	-2.70(3)	-5.94(1)	-4.61(1)
ΔLog X	-12.1(2)	-3.63(4)	-5.59(3)	-8.60(1)	-6.16(2)	-3.93(3)
ΔLog(PX/PXW)	-7.86(2)	-5.18(3)	-5.41(3)	-8.35(4)	-4.86(1)	-9.04(1)
ΔLog YW	-4.58(1)	-4.58(1)	-4.58(1)	-4.45(1)	-4.58(1)	-4.58(1)

Notes: a. The Mackinnon (1991) critical value of the ADF statistic when a trend term is included in the procedure for 72 observations is -3.47 at the 5% level and -3.16 at the 10% level of significance.

b. Numbers inside the brackets are number of lags.

the remaining three countries, i.e., Greece, Singapore, and South Africa, we can apply the cointegration technique.

Before presenting cointegration results few points deserve mention. First, in cointegration analysis, equations such as (1) where lagged exchange rate is replaced by its current value, are called cointegration equations and coefficient estimates are referred to as cointegrating vectors. Second, in trying to use the residuals as a proxy for a linear combination among variables, depending upon which variable is used as the dependent variable, we could obtain different cointegrating vector. In case there is evidence of cointegration in more than one case or evidence of more than one cointegration vector, a common practice is to select the one that yields a highest adjusted R^2. This practice is followed, for example, by Miller (1991) and Bahmani-Oskooee and Rhee (1994) and will be adopted here if necessary. Thus, concentrating on the import demand equation, upon which of the four variables (imports, exchange rate, domestic income, and exchange rate volatility) is used as the dependent variable we need to estimate four cointegration equation. When exchange rate is replaced by relative prices, four additional cointegration equations are estimated for each country. Applying the same criteria on the export demand equation, we need to estimate eight more cointegration equation for each country. After estimating each cointegration equation, the ADF test was applied for their residuals and the results are reported in Table 4.

As can be seen from Table 4, at the usual 5% level of significance there is only evidence of cointegration in the fifth equation for Greece and in the 7th and 11th equation for South Africa. Concentrating in equations where Log M and Log X are the dependent variables, only in the case of Greece (5th equation) there is evidence of cointegration. By and large, these results are indicative of lack of any long-run relationship among the variables involved. It should, however, be mentioned that lack of cointegration could be interpreted as lack of any stable relationship among the variables. Indeed, McCallum (1993, p. 37) argues that when the residuals are nonstationary, "it is highly misleading to conclude that in any practical sense long-run relationships are therefore nonexistent." McCallum attributes his conjecture to the fact that technical progress cannot be captured by any measurable variable (not even by the trend term). Therefore, if technical progress is not captured by a variable in any model, it must be influencing the disturbance term. According to McCallum, since technological progress (or shocks) do not usually reverse themselves, one would expect the disturbance term to contain a permanent component making it nonstationary. In light of these new developments in the literature one

Table 4

THE RESULTS OF ADF TEST APPLIED TO THE RESIDUALS OF COINTEGRATION EQUATIONS[a]

Cointegration Equation	Greece	Singapore	S. Africa
1. Log M = F (T, Log R, Log Y, σ_R)	–2.50(3)[b]	–2.16(4)	–3.11(4)
2. Log R = F (T, Log M, Log Y, σ_R)	–3.01(1)	–3.06(4)	–3.11(3)
3. Log Y = F (T, Log R, Log M, σ_R)	–1.65(3)	–3.27(4)	–2.83(3)
4. σ_R = F (T, Log R, Log Y, Log M)	–3.30(4)	–2.37(1)	–3.23(3)
5. Log M = F (T, Log PM/PD, Log Y, σ_R)	–5.48(1)	–1.43(3)	–3.81(1)
6. Log PM/PD = F (T, Log M, Log Y, σ_R)	–1.99(2)	–2.91(1)	–3.36(4)
7. Log Y = F (T, Log PM/PD, Log M, σ_R)	–2.56(4)	–2.54(3)	–6.82(1)
8. σ_R = F (T, Log PM/PD, Log Y, Log M)	–3.54(4)	–2.94(1)	–3.54(4)
9. Log X = F (T, Log R, Log YW, σ_R)	–3.16(3)	–2.86(4)	–1.79(4)
10. Log R = F (T, Log X, Log YW, σ_R)	–3.00(1)	–3.94(4)	–4.09(3)
11. Log YW = F (T, Log R, Log X, σ_R)	–2.87(1)	–4.34(4)	–5.13(3)
12. σ_R = F (T, Log R, Log YW, Log X)	–3.10(1)	–2.08(1)	–3.43(3)
13. Log X = F (T, Log PX/PXW, Log YW, σ_R)	–3.37(3)	–2.84(4)	–1.82(4)
14. Log PX/PXW = F (T, Log X, Log YW, σ_R)	–2.69(3)	–2.87(2)	–3.17(4)
15. Log YW = F (T, Log PX/PXW, Log X, σ_R)	–3.48(4)	–3.73(4)	–3.37(1)
16. σ_R = F (T, Log PX/PXW, Log YW, Log X)	–3.29(1)	–2.27(1)	–2.02(1)

Notes: a. Mackinnon critical value of the ADF statistic for 72 observations when a trend term is included in the cointegration equation is –4.63 at the 5% level and –4.28 at the 10% level.
b. Number inside the bracket is the number of lags in the ADF procedure.

should consider our cointegration results preliminary. More cointegration analysis are recommended for future research on this issue, perhaps using different techniques such as Johansen and Juselius (1990) method.

IV. Concluding Remarks

Since the advent of current floating exchange-rates many authors have analyzed the effects of exchange-rate volatility on the trade flows. Less devleoped countries have received little attention. Therefore, any study that deals with the experience of LDCs will be an addition to the literature. In this paper we tried to investigate the response of six LDCs import and export volumes to exchange-rate volatility using quarterly data over 1973I-1990IV period using traditional econometrics method as well as cointegration technique.

When standard estimation method was used to estimate the import and export demand equations, we were able to show that for 50% of the countries in the sample volatility of real effective exchange-rates has depressed the import and export volumes of LDCs, a result consistent with the experience of DCs. However, when we relied upon cointegration technique, there was not much evidence of cointegration among the variables of import and export demand equations. After reconciling the conflicting findings more work on the subject for future research using different cointegration techniques was suggested.

Appendix

Data Definition and Sources

All data are quarterly for the period 1973-1990 and are taken from the following sources.

a. International Financial Statistics of IMF, various issues.
b. Bahmani-Oskooee (1993).
c. OECD Statistics of Foreign Trade and OECD Main Economic Indicators, various issues.

Variables;

M = import volume. Nominal imports are deflated by import price index (1985 = 100) to obtain this measure. All data are from source a.

R = index of real effective exchange rate (1985 = 100) from source b. Note that Bahmani-Oskooee (1993) who constructed the real and nominal effective exchange rate for 22 LDCs for 1971-1990 period defined them as number of units of foreign currency per unit of domestic currency. While we employed this definition in *all export demand equations,* we used the *inverse* of Bahmani-Oskooee's measures in *all import demand equations.* Only with this transformation we could expect the elasticity of import with regard to exchange rate to be negative which is a consistent notion with any price elasticity.

Y = real GNP. Quarterly GNP figures were not available for the countries included in our study. Thus, we generated quarterly nominal GNP using the method in Bahmani-Oskooee (1986). We then deflated these nominal figures by the domestic price level (CPI, 1985 = 100) to obtain real GNP. The annual GNP and import figures used in this process are from source a.

σ = measure of exchange-rate variability. Following Lanyi and Suss (1982, p. 552), σ is computed as the standard deviation of quarterly percentage changes in the real effective exchange rate, R. To generate time series data for σ, following Bahmani-Oskooee (1991) for each quarter we defined it as the standard deviation of quarterly percentage changes in the real effective exchange rate over the preceding eight quarters as well as the current quarter. Quarterly data on real effective exchange rates which come from source b was available for 1971I-1990IV period. Due to method of construction of σ, the first eight observations were lost, thus, the entire study was performed by using data over 1973I-1990IV period.

PM = index of unit value of imports, 1985 = 100, source a.

PD = index of domestic price level measured by CPI, 1985 = 100, source a.

X = export volume. Nominal exports are deflated by export price index (1985 = 100) to obtain this measure. All data are from source a.

PX/PXW = the ratio of each country's dollar-denominated export unit value index (1985 = 100) to the dollar denominated export unit value index (1985 = 100) of the IMF's "industrial country" aggregate. This definition is used by Bailey et al. (1986) as well as by Bahmani-Oskooee (1991).

YW = world income. Following others in the literature this variable is proxied by the index of industrial production (1985 = 100) in OECD countries. Data come from source c.

References

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

Bahmani-Oskooee, M., "Determinants of International Trade Flows: The Case of Developing Countries," *Journal of Development Economics*, 20, Jan.-Feb. 1986, 107-123.

——, "Exchange-Rate Uncertainty and the Trade Flows of Developing Countries," *Journal of Developing Areas*, 25, July 1991, 497-508.

——, "Real and Nominal Effective Exchange Rates for 22 LDCs: 1971I-1990IV," Working Paper, Department of Economics, University of Wisconsin-Milwaukee, 1993.

—— and M. Malixi, "Effects of Exchange Rate Flexibility on the Demand for International Reserves," *Economic Letters*, 23, May 1987, 89-93.

—— and H.J. Rhee, "Long-Run Elasticities of the Demand for Money in Korea: Evidence from Cointegration Analysis," *International Economic Journal*, 1994, Forthcoming.

Bailey, M.J., G.S. Tavalas and M. Ulan, "Exchange-Rate Variability and Trade Performance: Evidence for the Big Seven Industrial Countries," *Weltwirtschaftliches Achiv*, 122, 1986, 466-477.

——, "The Impact of Exchange-Rate Volatility on Export Growth; Some Theoretical Considerations and Empirical Results," *Journal of Policy Modelling*, 9, Spring 1987, 225-243.

Brada, J.C. and J.A. Mendez, "Exchange Rate Risk, Exchange Rate Regime and the Volume of International Trade," *Kyklos*, 41, 1988, 263-280.

Caballero, R.J. and V. Corbo, "The Effect of Real Exchange Rate Uncertainty on Exports: Empirical Evidence," *The World Bank Economic Review*, 3, 1989, 263-278.

Coes, D.V., "The Crawling Peg and Exchange Rate Uncertainty," in John Williamson (ed.), *Exchange Rate Rules*, New York: St. Martins, 1981.

Cushman, D.O., "The Effects of Real Exchange Rate Risk on International Trade," *Journal of International Economics*, 15, August 1983, 45-63.

——, "Has Exchange Risk Depressed International Trade? The impact of Third-Country Exchange Risk," *Journal of International Money and Finance*, 5, September 1986, 361-379.

_____, "U.S. Bilateral Trade Flows and Exchange Risk During the Floating Period," *Journal of International Economics,* 25, May 1988, 317-330.

Goutor, P., "Effects of Exchange Rate Volatility on Trade: Some Further Evidence," *IMF Staff Papers,* 32, 1985, 475-512.

Gupta, S., "Exchange Risk in International Trade under Alternative Exchange Systems: The Developing Countries' Experience," Unpublished Ph.D. Dissertation, Michigan State University, 1980.

Hooper, P. and S.W. Kohlhagen, "The Effects of Exchange Rate Uncertainty on the Prices and Volume of International Trade," *Journal of International Economics,* 8, November 1978, 483-511.

Houthakker, S.H. and S.P. Magee, "Income and Price Elasticities in World Trade," *The Review of Economics and Statistics,* 51, May 1969, 111-125.

Johansen, S. and K. Juselius, "Maximum Likelihood Estimation and Inference on Cointegration-With Application to the Demand for Money," *Oxford Bulletin of Economics and Statistics,* 52, May 1990, 169-210.

Junz, H. and R.R. Rhomberg, "Price Competitiveness in Export Trade among Industrial Countries," *American Economic Review, Papers and Proceedings,* 63, May 1973, 412-418.

Kenen, P.B. and D. Rodrik, "Measuring and Analyzing the Effects of Short-Term Volatility in Real Exchange Rates," *The Review of Economics and Statistics,* 68, May 1986, 311-315.

Khan, S.M., "Import and Export Demand in Developing Countries," *IMF Staff Papers,* 21, November 1974, 678-693.

_____ and K.Z. Ross, "Cyclical and Secular Income Elasticities of the Demand for Imports," *The Review of Economics and Statistics,* 57, August 1975, 357-361.

Koray, F. and W.D. Lastrapes, "Real Exchange Rate Volatility and U.S. Bilateral Trade: A VAR Approach," *The Review of Economics and Statistics,* 71, November 1989, 708-712.

Lanyi, A. and E.C. Suss, "Exchange Rate Variability: Alternative Measures and Interpretation," *IMF Staff Papers,* 29, December 1982, 527-560.

Lastrapes, W.D. and F. Koray, "Exchange Rate Volatility and U.S. Multilateral Trade Flows," *Journal of Macroeconomics,* 12, Summer 1990, 341-362.

Makin, J.H., "Eurocurrencies and the Evolution of the International Monetary System," in C.H. Stem, J.H. Makin, and D.E. Logue, (eds.), *Eurocurrencies and the International Monetary System,* American Enterprise Institute, Washington, D.C., 1976, 17-51.

McCallum, B.T., "Unit Roots in

Macroeconomic Time Series: Some Critical Issues," *NBER Working Paper No. 4368,* May 1993, 1-56.

MacKinnon, J.J., "Critical Values for Cointegration Tests," in *Long-Run Economic Relationships: Readings in Cointegration,* (ed.). R.F. Engle and C.W. Granger, Oxford, Oxford University Press, 1991, 267-276.

Miller, S.M., "Monetary Dynamics: An Application of Cointegration and Error-Correction Modeling," *Journal of Money, Credit and Banking,* 23, May 1991, 139-154.

Orcutt, G.H., "Measurement of Price Elasticities in International Trade," *Review of Economics and Statistics,* 32, May 1950, 117-132.

Peree, E. and A. Steinherr, "Exchange Rate Uncertainty and Foreign Trade," *European Economic Review,* 33, July 1989, 1241-1264.

Rana, R.B., *ASEAN Exchange Rates: Policies and Trade Effects,* Singapore: ASEAN Economic Research Unit, Institute of Southeast Asian Studies, 1981.

Thursby, J.G. and M.C. Thursby, "Bilateral Trade Flows, the Linder Hypothesis, and Exchange Risk," *The Review of Economics and Statistics,* 69, August 1987, 488-495.

Warner, D. and M. E. Kreinin, "Determinants of International Trade Flows," *The Review of Economics and Statistics,* 65, February 1983, 96-104.