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Are Devaluations Contractionary in LDCs?*

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Devaluation is said to stimulate the aggregate demand by increasing its net export component. On the other hand, it is said to discourage the aggregate supply by raising cost of imported inputs. The ultimate effect on output has to be settled through empirical analysis. In this paper we employed quarterly data on the measure of domestic output and real as well as nominal effective exchange rate of 23 LDCs and the cointegration technique to show that devaluations have no long-run effect on output in most LDCs.

I. Introduction

Some studies in the Literature have investigated the effects of devaluation on the output of LDCs. The standard text book argument is that by making a country's exports more competitive, devaluation should lead to an increase in exports, therefore, to an increase in aggregate demand. On the other hand, as Gylfason and Risager (1984, p. 53) have shown, a devaluation can decrease aggregate supply because of an increase in the prices of imported inputs. The net effect on domestic output will depend on the extent to which the aggregate demand and aggregate supply change.

The literature on the effects of devaluation on domestic output includes some theoretical and some empirical studies. At the theoretical level, by a reference to income redistribution effect of devaluation, Diaz-Alejandro (1963, p. 577) pointed out that an observer of devaluation could be puzzled to see that devaluations that resulted in an improvement of the trade balance, were accompanied by a decline in the level of total output. Recently, in an influential theoretical paper Krugman and Taylor (1978), using a Keynesian as well as a monetarist framework showed that under certain conditions devaluation could be contractionary. They showed that under Keynesian model, devaluations are contractionary if (i) imports initially exceeds exports; (ii) there are differences in consumption propensities from profits and wages; (iii) government revenues are increased by devaluation. They argued that similar effects exist under monetarist model as well. They then showed several channels of contractionary influence specially relevant to the case of LDCs. Extending Krugman and Taylor (1978) model to incorporate the factor market, Hansen (1983) derived a condition for contractionary effects of devaluation that depended upon the sign of one plus the weighted sum of the price elasticities of the demand for imported consumer goods and of derived demand for imported inputs. However, when additional channels and additional features of LDCs were identified and incorporated into a

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general and theoretical framework by Lizondo and Montiel (1989, p. 182), they concluded that "the direction of the impact effects of devaluation on real output is ambiguous on analytical grounds."[1]

On the empirical ground, Gylfason and Schmid (1983) estimated parameters of a simple macro-economic model for a group of five industrial countries and five less developed countries and showed that for most of them devaluations are expansionary in the sort to medium run. Gylfason and Risager (1984) introduced a model that emphasized the effects of devaluation on interest payments on foreign debt. Using parameter data for 8 developing countries and 7 developed countries, they showed that while devaluations are generally contractionary for the first group, they are expansionary for most of the countries in the second group. Solimano (1986) investigated the experience of Chile and concluded that devaluations are contractionary in the short to medium run. Edwards (1986) has estimated a reduced form real output equation by pooling data from 12 less developed countries and concluded that devaluations have a negative short-run effect on output. However, Edward's results showed that after one year, devaluations are expansionary with no effect on output in the long-run. Finally, Gylfason and Radetzki (1991) present simulation results for 12 least developed countries that suggest that for most of the countries in their sample, devaluations are contractionary in the short to medium run and this contractionary effect is even stronger when wages are indexed.

The review of empirical studies indicate that most of them have investigated the short-run effects of devaluation. In this paper we intend to discover whether devaluations have any long-run effect on LDCs domestic output using cointegration technique. Our finding suggests that in most cases, devaluations have no long-run effect on LDCs domestic production. Section II introduces and discusses the cointegration technique. Section III presents our empirical results. Section IV concludes. Finally, data definition and sources are reported in the appendix.

I. The Method

The long-run equilibrium relationship between two variables could be detected by cointegration method. Define a non-stationary time series X_t as integrated of order d if it achieves stationarity after being differenced d times, denoted by $x_t \sim I(d)$. According to Engle and Granger (1987) two $I(d)$ variables will be cointegrated if a linear combination between them is integrated at any order less than d. Usually, as an estimate of a linear combination, residuals from the simple OLSQ regression of one variable on the other are tested for stationarity. For example, if $X_t \sim I(1)$ and $Y_t \sim I(1)$, in order for X_t and Y_t to be cointegrated, the residuals from OLSQ regression of X_t on Y_t or vice versa should be $I(0)$.

Following Engle and Granger (1987), we also investigate causality between output and exchange rate using an error-correction model as outlined by Equations (1) and (2):

1. The channels through which devaluation could affect the aggregate demand and aggregate supply, as identified by Krugman and Taylor (1978) and by Lizondo and Montiel (1989) are through the effects of devaluation on consumption, relative prices, real income, imported inputs, income distribution, real tax revenue, wealth, investment, nominal interest rates, nominal wages, and cost of capital.

$$(1-L)X_t = a_0 + b_0\varepsilon_{t-1} + \sum_{i=1}^{M} c_{0i}(1-L)X_{t-i} + \sum_{i=1}^{N} d_{0i}(1-L)Y_{t-i} + \mu_t, \tag{1}$$

and

$$(1-L)Y_t = a_1 + b_1\varepsilon'_{t-1} + \sum_{i=1}^{M} c_{1i}(1-L)Y_{t-i} + \sum_{i=1}^{N} d_{1i}(1-L)X_{t-i} + \mu'_t. \tag{2}$$

where L is the lag operator and the error correction terms ε and ε' are the stationary residuals from OLSQ regression of X_t on Y_t and Y_t on X_t respectively. In Equation (1), Y is said to Granger cause X either the d_{0i}'s are jointly significant or b_0 is significant. If two variables are not cointegrated, the error-correction terms are excluded from (1) and (2) and simple Granger causality test is carried out.

III. Empirical Results

In this section we try to apply the cointegration analysis between domestic output (real GNP) and the nominal effective exchange rate (E) of 23 LDCs using quarterly data over 1973-1988 period.[2] In order to determine whether real devaluations have different effects than nominal devaluations, we also carry out the whole exercise between GNP and the real effective exchange rate. Following Bahmani-Oskooee and Malixi (1987) we define the real effective rate as (P^*E/P) where P^* is the foreign price level, E is the nominal effective exchange rate (number of units domestic currency per unit of foreign currency) and P is the domestic price level.

In applying the cointegration technique, we first need to verify that for each country GNP, (P^*E/P), and E series are integrated of the same order. To this end, we rely upon Augmented Dickey-Fuller (ADF here after) test which includes a trend term.[3] The results of the ADF test for real GNP, (P^*E/P), and E variables are reported in Table 1. Table 1 not only reports the results of ADF test for level of each variable but, for the first differenced variables and if necessary for the second differenced variables as well.

A closer look at the calculated ADF statistics reported in Table 1 leads us to classify the 23 countries into four groups. The first group includes six countries, i.e., Cameroon, Pakistan, South Africa, Thailand, Turkey, and Venezuela for which the cointegration technique cannot be applied. This is due to the fact that variables are not integrated of the same order. For example, for Cameroon real effective rate is $I(0)$, whereas, real GNP is $I(2)$. For Pakistan real GNP is $I(0)$ and real or nominal effective rate is $I(1)$ and etc. The second group includes 9 countries, i.e., Bangladesh, Barbados, Egypt, Ethiopia, Honduras, Indonesia, Korea, the Philippines, and Tunesia for which we can apply cointegration technique to the relation between their real GNP and their real as well as nominal effective rates. This is

2. It should be mentioned that due to unavailability of data the study period for Brazil was 1973I-1985III; for Mexico it was 1973I-1987III; and for Portugal it was 1973I-1986IV. The availability of the data is the major reason for selecting the 23 LDCs. If we were to include additional LDCs, number of observations would be too small.
3. For a detailed formulation of the ADF test see Bahmani-Oskooee and Payesteh (1993, p.197). In selecting the number of lags in the ADF test as well as in error-correction models, following Bahmani-Oskooee and Rhee (1992) we minimized the Akaike's Final Prediction Error (FPE) statistic.

because all three variables are $I(1)$ for these 9 countries. The Ecuador, Mexico, and Portugal for which we can investigate the long-run relation between their real GNP and only their real effective exchange rates. This is due to the fact that their GNP and real exchange rates are both $I(1)$, whereas, their nominal effective rates are $I(2)$. Finally, the fourth group includes India, Malaysia and Singapore for which we can only apply the technique to the relation between their real GNP and their nominal effective rates. For these three countries, the real effective exchange rate is $I(2)$, whereas, the other two variables are both $I(1)$. Note that all 17 countries included in the last three groups do have one thing in common. The appropriate variables that will be employed in the cointegration analysis are all $I(1)$. Therefore, what remains to be determined is whether the residuals from the cointegration equations are on a stationary process, i.e., whether they are $I(0)$.

Using notations introduced so far we formulate the cointegration equations as below:

$$(GNP)_t = \alpha_1 + \beta_1 T + \gamma_1 (P^* E/P)_t + \varepsilon_{1t}, \tag{3}$$

$$(P^* E/P)_t = \alpha_2 + \beta_2 T + \gamma_2 (GNP)_t + \varepsilon_{2t}, \tag{4}$$

$$(GNP)_t = \alpha_3 + \beta_3 T + \gamma_3 (E)_t + \varepsilon_{3t}, \tag{5}$$

$$(E)_t = \alpha_4 + \beta_4 T + \gamma_4 (GNP)_t + \varepsilon_{4t}. \tag{6}$$

Cointegration Equations (3) and (4) employ the real effective exchange rate whereas, (5) and (6) employ nominal effective exchange rate. After estimating Equations (3)-(6) by OLSQ, the ADF test was applied to their residuals. Table 2 reports the calculated ADF test statistics for such residuals, only for the 17 countries for which the cointegration technique could be applied. Note that whenever the technique was not applicable, we do not report the ADF statistic.

From Table 2 we gather that for three countries, i.e., Bangladesh, Portugal and Tunesia $\varepsilon_1 \sim I(0)$. Since the degree of integration of ε_1 for these three countries is less than the degree of integration of the variable included in their cointegration equations, we conclude that for these three countries there exists a long-run relation between their real GNP and their real effective exchange rates. It is also evident that for Bangladesh, Barbados, India, and Tunesia $\varepsilon_3 \sim I(0)$ indicating that for these four countries there is a long-run relationship between their real GNP and their nominal effective exchange rate. For the remaining countries, all four ε's are $I(1)$ and so were their GNP and their $(P^* E/P)$ or E variables, indicating lack of any long-run relation. All in all, the results in Table 2 indicate that in most LDCs there is no long-run relation between their domestic output and their real or nominal effective exchange rate. This is an indication that in most LDCs, the expansionary effects of devaluation on aggregate demand is offset by the contractionary effects of devaluation on aggregate supply in the long-run. This finding supports Edwards (1986, p. 507) who concluded that "In the long-run devaluations were found to be neutral."

Table 2 The Calculated ADF Test Statistics for the Stationarity of ε_1, ε_2, ε_3, and ε_4[a]

Country	ε_1	ε_2	ε_3	ε_4
Bangladesh	-4.79[2][b]	-2.42[1]	-5.10[1]	-2.30[1]
Barbados	-3.17[2]	-1.49[4]	-4.23[2]	-2.06[4]
Brazil	-2.85[4]	-2.66[4]	-	-
Colombia	-2.49[3]	-0.42[3]	-	-
Ecuador	-2.58[4]	0.56[3]	-	-
Egypt	-1.18[3]	-2.49[1]	-2.26[3]	-2.28[1]
Ethiopia	-2.39[2]	-1.61[2]	-1.68[2]	-2.18[4]
Honduras	-1.99[4]	-2.22[4]	-2.00[4]	-2.30[4]
India	-	-	-3.97[4]	-2.08[1]
Indonesia	-1.69[3]	-1.30[4]	-2.16[4]	-1.49[4]
Korea	-1.51[4]	-2.20[4]	-0.93[4]	-3.03[1]
Malaysia	-	-	-2.38[4]	-0.88[4]
Mexico	-2.20[4]	-1.25[4]	-	-
Philippines	-3.21[4]	-2.12[4]	-3.35[4]	-1.98[1]
Portugal	-4.38[2]	-2.95[4]	-	-
Singapore	-	-	-2.10[4]	-2.06[4]
Tunesia	-6.11[1]	-2.02[4]	-6.47[1]	-1.45[4]

Notes: a. The Mackinnon (1991) critical value of the ADF statistic for cointegration between two variables when a trend term is included in the cointegration equation (for 50 observations) are -3.97 at the 5% level and -3.58 at the 10% level of significance.

b. Numbers inside the brackets are number of lags in equation 3, i.e., value of k.

As the appendix indicates since effective exchange rate is defined as number of units of domestic currency per unit of foreign currency, an increase in it will be an indication of depreciation of domestic currency. Therefore, if depreciation is to be expansionary, the slope coefficient in cointegration Equations (3) and (5) must be positive. Indeed, this was the case in the results for all countries for which there was evidence of cointegration, indicating that for those countries devaluations are expansionary.[4]

Let us now turn to the estimates of the error-correction models in order to determine the direction of causation between the two variables. Specifically, in the cases of Bangladesh, Portugal, and Tunesia for which we have established that the two variables are cointegrated and there is a long-run relation between them, we would like to determine which variable causes the other and provides the short-run dynamic adjustment toward the long-run equilibrium. Note that the detection of short-run causality between real GNP and the real or nominal effective exchange rate could be investigated for all countries for which the three variables were found to be integrated of the same order, as outlined in Tables 1. Table 3 reports the results of causality detection between real GNP and real effective exchange rate of 14 LDCs for which both variables were $I(1)$. Table 4 reports the same exercise between real GNP and nominal

4. As a matter of fact, all slope coefficients for all countries were positive and highly significant (measured by standard t-statistics). But since for most of the countries, there is no evidence of cointegration between variables involved, these highly significant coefficients cannot be relied upon.

effective rate, only for 12 countries for which these two variables were found to be $I(1)$.[5]

Table 3 Causality Detection between Real GNP and Real Effective Exchange Rate (P^*E/P)

		Independent Variables		
Country	Dependent Variable	t-statistic for EC_{t-1}	F-statistic for $\sum(1-L)GNP_{t-i}$	F-statistic for $\sum(1-L)(P^*E/P)_{t-i}$
Bangladesh	(1-L)GNP	-3.78*	0.87[2]a	2.14[1]
	(1-L)(P*E/P)	-	2.30[1]	3.50**[1]
Barbados	(1-L)GNP	-	6.06*[7]	1.39[8]
	(1-L)(P*E/P)	-	1.51[7]	2.31*[8]
Brazil	(1-L)GNP	-	4.85*[8]	0.04[1]
	(1-L)(P*E/P)	-	5.16*[2]	2.44[1]
Colombia	(1-L)GNP	-	27.3*[3]	7.27*[2]
	(1-L)(P*E/P)	-	1.96**[6]	4.83*[4]
Ecuador	(1-L)GNP	-	2.96*[6]	2.14**[5]
	(1-L)(P*E/P)	-	4.84*[2]	3.86*[8]
Egypt	(1-L)GNP	-	4.99*[4]	1.46[2]
	(1-L)(P*E/P)	-	1.78[3]	0.76[1]
Ethiopia	(1-L)GNP	-	5.56*[2]	3.64*[3]
	(1-L)(P*E/P)	-	1.25[1]	2.47*[5]
Honduras	(1-L)GNP	-	28.9*[4]	0.07[1]
	(1-L)(P*E/P)	-	0.91[5]	4.36*[1]
Indonesia	(1-L)GNP	-	28.6*[3]	4.01*[1]
	(1-L)(P*E/P)	-	0.08[1]	5.29*[1]
Korea	(1-L)GNP	-	24.4*[4]	0.68[1]
	(1-L)(P*E/P)	-	2.24[2]	1.36[8]
Mexico	(1-L)GNP	-	42.9*[4]	3.81*[4]
	(1-L)(P*E/P)	-	1.41[1]	2.47[1]
Philippines	(1-L)GNP	-	3.05**[2]	4.34*[5]
	(1-L)(P*E/P)	-	2.43*[6]	3.02*[5]
Portugal	(1-L)GNP	-2.75*	0.85[8]	2.34**[5]
	(1-L)(P*E/P)	-	1.78[1]	1.39[4]
Tunesia	(1-L)GNP	-2.82*	0.96[3]	2.13[2]
	(1-L)(P*E/P)	-	3.22*[7]	2.38**[3]

Notes: a. Numbers inside the brackets are the number of lags selected by the FPE criteria.
* Significant at the 5% level.
** Significant at the 10% level.
EC denotes the error-correction term.

Concentrating on Table 3 and three countries for which there was evidence of cointegration, we gather that in the cases of Bangladesh and Portugal, the causality is from real exchange

5. It should be indicated that in selecting the number of lags, again we minimized the FPE statistic by imposing all possible lag combination on both variables with a maximum of 8 lags on each.

Table 4 Causality Detection between Real GNP and Nominal Effective Exchange Rate, E.

		Independent Variables		
Country	Dependent Variable	t-statistic for EC_{t-1}	F-statistic for $\sum(1\text{-}L)GNP_{t-i}$	F-statistic for $\sum(1\text{-}L)E_{t-i}$
Bangladesh	(1-L)GNP	-5.19*	0.94[1][a]	8.35*[1]
	(1-L)E	-	4.52*[1]	12.6*[1]
Barbados	(1-L)GNP	-4.26*	3.05*[5]	0.01[1]
	(1-L)E	-	3.36*[1]	2.28[2]
Egypt	(1-L)GNP	-	5.94*[3]	2.21[2]
	(1-L)E	-	15.2*[1]	8.16*[1]
Ethiopia	(1-L)GNP	-	1.89[3]	4.02*[1]
	(1-L)E	-	0.45[1]	3.11*[3]
Honduras	(1-L)GNP	-	29.7*[4]	0.74[1]
	(1-L)E	-	2.18*[8]	3.99*[1]
India	(1-L)GNP	-	8.27*[7]	5.15*[6]
	(1-L)(P^*E/P)	-	1.23[1]	2.48*[6]
Indonesia	(1-L)GNP	-	27.4*[3]	2.32[1]
	(1-L)E	-	0.02[1]	3.61*[2]
Korea	(1-L)GNP	-	25.5*[4]	1.43[3]
	(1-L)E	-	0.70[2]	5.98*[1]
Malaysia	(1-L)GNP	-	5.11*[1]	2.42*[8]
	(1-L)E	-	0.41[2]	13.9*[1]
Philippines	(1-L)GNP	-	4.88*[2]	1.28[1]
	(1-L)E	-	4.06*[7]	1.93[4]
Singapore	(1-L)GNP	-	6.14*[4]	0.23[1]
	(1-L)E	-	3.25*[7]	5.00*[4]
Tunesia	(1-L)GNP	-5.40*	0.54[2]	2.27[1]
	(1-L)E	-	2.20*[7]	12.4*[1]

Notes: a. Numbers inside the brackets are the number of lags selected by the FPE criteria.
* Significant at the 5% level.
** Significant at the 10% level.
EC denotes the error-correction term.

rate to domestic output and in the results for Tunesia, there is evidence of bidirectional causality. While in the case of Bangladesh, the causality is through the error-correction term, in the case of Portugal, it is through the error-correction term as well as through the joint significance of lagged first differenced real exchange rate. For the remaining countries, real effective exchange rate Granger causes domestic output only in the results for Colombia, Ecuador, Indonesia, Mexico, and the Philippines. Output Granger causes real effective exchange rate in the cases of Brazil, Colombia, Ecuador, and the Philippines. However, since for these remaining countries there was lack of cointegration, these later findings are interpreted as short-run effects.

Table 4 which reports the causality results between real *GNP* and nominal effective exchange rate reveals that for Bangladesh, Barbados, and Tunesia, there is evidence of bidirectional causality between the two variables. For the remaining countries nominal effective rate Granger causes real output in the results for Ethiopia, India, and Malaysia and Real output Granger causes nominal effective exchange rate in the cases of Egypt, Honduras, the Philippines, and Singapore, however, only in the short-run. Furthermore, the results not reported but available from the author, showed that in each case, the estimated lagged coefficients took positive as well as negative signs, indicating that in the short-run devaluations could be expansionary or contractionary.

Before closing an issue that deserves investigation is the question of whether our finding of no cointegration for most LDCs could be due to omission of other policy variables from cointegration equations. In investigating this question, following Edwards (1986) we include a measure of fiscal policy, G; a measure of monetary policy, M; and a measure of commercial policy, terms of trade, TOT in the cointegration equation as in Equation (7) below:

$$GNP_t = F[\,T,\ G_t,\ M_t,\ TOT_t,\ E_t] + \varepsilon_{5t}. \tag{7}$$

Indeed, Engle and Granger (1987) have shown that if all variables in (7) are cointegrated, the OLSQ estimate of (7) will yield coefficient estimates (cointegrating vector) that are consistent and just as efficient as maximum likelihood estimator. It should be indicated that quarterly import and export prices to generate the terms of trade were not available for most LDCs. Thus, it was necessary to exclude it form our exercise and concentrate on the following formulation:

$$GNP_t = F[\,T,\ G_t,\ M_t,\ E_t] + \varepsilon_{6t}. \tag{8}$$

Alternatively, as Dornbush (1986, p. 58) has pointed out, since the terms of trade and the real exchange rate carry almost the same concept, we can substitute the real effective exchange rate for terms of trade. In this case, the nominal exchange rate, E should be dropped from the cointegration equation to avoid counting it twice. Thus, the second alternative is formulated as Equation (9) below:

$$GNP_t = F[\,T,\ G_t,\ M_t,\ (P^*E/P)_t\,] + \varepsilon_{7t}. \tag{9}$$

In applying the cointegration analysis to Equations (7)-(9), we first made sure that all variables are $I(1)$.[6] With few exceptions that will be explained later, the ADF test supported the $I(1)$ properties of G, M, TOT variables for countries for which the GNP, E, and P^*E/P were already found to be $I(1)$.[7] What remains to be shown for cointegration is whether the residuals in the cointegration equations (7)-(9) are $I(0)$. Thus, we estimated the linear formulation of each equation by OLSQ and applied the ADF test for their residuals. Table 5 reports the results. The critical values that depend on the number of variables are reported at the bottom of Table 5.

6. It should be indicated that due to lack of data on government expenditure or money supply figures the study period for Bangladesh was 1974I-1985IV, for Brazil it was 1973I-1985IV, and for Malaysia, it was 1973I-1987IV.

7. For list of these countries see Table 1.

Table 5 The Calculated ADF Test Statistics for the Stationarity of ε_5, ε_6, and ε_7 [a]

	ε_5	ε_6	ε_7
Bangladesh	-	-3.43[1]	-3.70[1][b]
Barbados	-	-5.97[1]	-5.47[3]
Brazil	-	-	-3.50[3]
Colombia	-	-	-1.91[2]
Ecuador	-	-	-4.25[4]
Egypt	-	-2.77[1]	-2.31[4]
Ethiopia	-	-2.78[4]	-2.47[2]
Honduras	-	-3.78[4]	-3.57[4]
India	-	-	-3.66[2]
Indonesia	-	-1.96[3]	-1.91[3]
Korea	-1.94[3]	-2.61[3]	-1.81[3]
Malaysia	-4.41[2]	-	-
Mexico	-	-	-2.25[3]
Philippines	-3.9[4]	-	-3.94[4]
Portugal	-	-	-3.20[3]
Singapore	-3.73[1]	-	-
Tunesia	-	-	-2.78[1]

a. The Mackinnon critical values when there are 4 variables in the cointegration equation (with trend) are -4.427 at the 5% level and -4.375 at the 10% level of significance. The comparable figures when there are 5 variables in the cointegration equation are -5.071 and -4.709.

b. Numbers inside the brackets are number of lags in equation 3, i.e., value of k selected by FPE criteria.

A glance through Table 5 indicates that only in the results for Barbados the ADF statistic is less than its critical value which is an indication of cointegration. For the remaining countries, there is no evidence of cointegration. It appears that inclusion of other policy variables have changed our previous conclusion of cointegration for Bangladesh, India, Portugal and Tunesia. All in all, our general conclusion that in most LDCs there is no long-run relation between domestic output and the exchange rate has remained unchanged. Few facts with regard to the reported results in Table 5 should be mentioned. First, it should be indicated that in the cases of Honduras, Mexico, Portugal, Tunisia, and Malaysia, the M variable was $I(2)$. Thus, the results should be viewed with caution. As a remedy, we included the first differenced M as $I(1)$ variable along with other $I(1)$ variables in the cointegration equations for the above countries and tested for the $I(0)$ property of the residuals. There was no change in our conclusion. Second, in case of Korea, her TOT variable was $I(0)$ which makes the results pertaining to Equation (9) somewhat biased. But, this should not be a matter of concern due to the fact that even after excluding TOT, there is no evidence of cointegration as in Table 5.

How are our findings compared to those of the other authors in the literature? Out of 17 countries for which the cointegration technique could be applied (see Table 2), Brazil, India, the Philippines, and Turkey were the only ones included in Gylfason and Schmid (1983). For Brazil, while they found that devaluations are contractionary in the short to medium run, we found that devaluations have no long-run effect. For the other three countries, while they found that devaluations are expansionary, again in the short to medium run, we found no long-run effect. Six of the countries in our sample were also included in Gylfason and Risager (1984) who showed that for Korea and Pakistan devaluations have positive effect on output and for Brazil, the Philippines, Turkey, and Portugal they have negative effects, again in the short to medium run. According to our findings, however, these effects disappear in the long run. Our results for Bangladesh and Ethiopia also indicate that the negative short-to-medium run effects of devaluations reported by Gylfason and Radetzki (1991) vanishes in the long run. In addition to these two countries, they included many other least developed countries in their sample.

We examined the long-run effects of devaluation on domestic output whereas, most of the other studies concentrated on its short-run effects. Hence, we consider our findings to be an extension of the previous research. When our findings are coupled with those of Gylfason and Schmid (1983), Gylfason and Risager (1984), Edwards (1986), and Gylfason and Radetzki (1991), we are led to conclude that the short-run effects of devaluation on domestic output could be different than their long-run effects, indicating that in the long-run devaluations are ineffective, in most LDCs. Given the feedback effect that exists between exchange rate and output, future research must concentrate on a more appropriate cointegration technique such as that of Johansen-Juselius (1990).

IV. Summary and Conclusion

Besides the effects of devaluation on the trade balance, some studies have raised questions about their effects on domestic output. More precisely, the traditional view that devaluations tend to stimulate domestic output has been challenged in recent years. These studies, though at the theoretical level, have argued that devaluations reduce aggregate supply (due to an increase in cost of imported inputs) more than they expand the aggregate demand (due to increase in exports and decrease in imports), resulting in contraction in the total output, specially in less developed countries. The limited number of empirical studies have provided some support for the contractionary effects of devaluation, but, only in the short to medium-run.

Unlike most previous research, in this paper we made an attempt to discover whether devaluations in LDCs have any long-run effect on their domestic output. By using quarterly data on the measure of domestic output and real as well as nominal effective exchange rate of 23 LDCs, the cointegration technique was used to establish the long-run relation between their outputs and their effective exchange rates. After providing evidence that the cointegration technique could only be applied to 17 of those 23 LDCs, our statistical results showed that devaluations have no long-run effect on output in most LDCs. No evidence was found in favor of contractionary long run effects of devaluations indicating that the relationships between exchange rates and output in these countries is temporary.

Appendix

Data Sources and Definitions

All data are quarterly for the period 1973I-1988IV and collected from the following sources.

a. International Financial Statistics of IMF, various issues.

b. Direction of Trade, IMF.

Variables:

(P^*E/P) = real effective exchange rate defined by Bahmani-Oskooee and Malixi (1987) where P^*, E, and P are defined as follows.

P^* = index of foreign price level measured by the CPI of industrial countries (1985=100). This index is available from source a.

E = index of import-weighted effective exchange rate (1985=100). The period average bilateral exchange rates that are used to construct this index are defined as units of domestic currency per unit of trading partner's currency and collected from source a. The 1985 share of imports of each country from its trading partners are from source b. The trading partners for each country are Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, Netherlands, New Zealand, Norway, Spain, Sweden, Switzerland, U.K. and the U.S.

P = Index of domestic price level (1985=100) measured by CPI, source a.

GNP = real GNP or real GDP. Quarterly figures for this variable were not available, therefore, they were generated using the method in Bahmani-Oskooee (1986). Note that Bahmani-Oskooee used an inverse import demand equation to generate nominal GNP. Since in some cases we encountered some data problem, we relied upon quantity theory of money to generate quarterly nominal GNP figures. More precisely, annual GNP and money supply (M) data were used in $GNP=M.V$ to obtain an estimate of velocity, V. By then using quarterly data for M, we obtained quarterly figures for GNP. We then adjusted the estimated quarterly data such that for each year their sum added up to the actual annual nominal GNP in that year. The quarterly nominal GNP figures were deflated by domestic CPI to obtain a measure of real GNP. All annual data on nominal GNP, money supply and imports were obtained from source a.

G = a measure of fiscal policy defined as real government expenditure. Except for Korea, this variable was not available on a quarterly basis. Thus, using annual nominal government expenditure and annual nominal GNP figures and a relation similar to quantity theory of money where money is replaced by G, we generated quarterly nominal government expenditures. More precisely, annual GNP and expenditure (G) data were used in $G=k(GNP)$ to obtain an estimate of k. By then using quarterly data for GNP, we obtained quarterly figures for G. We then adjusted the estimated quarterly data such that for each year their sum added up to the actual annual G in that year. The quarterly nominal G figures were deflated by domestic CPI to obtain a measure of real G. In few cases where government expenditures were not available, we employed government consumption. All annual data on government expenditure were obtained from source a.

M = a measure of monetary policy defined as real money supply. Nominal quarterly money supply (line 34 of IFS, source a) were deflated by domestic CPI to obtain this measure.

TOT = terms of trade defined as the ratio of export prices to import prices. Quarterly data were from source a.

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Table 1 The Calculated ADF Test Statistics for Stationarity of Real GNP, (P^*E/P), E, and Their First or Second Differences[a]

Country	GNP	(P^*E/P)	E	(1-L)(GNP)	(1-L)(P^*E/P)	(1-L)E
Bangladesh	-2.96[2][b]	-1.53[1]	-1.14[1]	-6.93[2]	-5.36[1]	-3.54[4]
Barbados	-2.49[2]	-0.87[1]	-1.43[2]	-7.99[2]	-6.50[1]	-4.33[1]
Brazil	0.83[3]	-1.22[4]	3.80[4]	-9.84[2]	-4.99[2]	-4.43[4][c]
Cameroon	-3.14[4]	-3.18[4]	-2.61[4]	-9.45[2][c]	-	-4.36[3]
Colombia	-1.39[3]	0.01[1]	0.12[4]	-8.96[2]	-4.94[1]	-8.25[2][c]
Ecuador	-2.53[4]	1.56[3]	6.91[2]	-7.08[2]	-7.32[1]	-4.39[2][c]
Egypt	-1.10[3]	-2.52[1]	0.41[1]	-6.76[2]	-4.82[2]	-4.34[1]
Ethiopia	-1.56[2]	-2.11[3]	-1.94[3]	-7.75[1]	-3.41[3]	-4.45[1]
Honduras	-1.63[4]	-2.34[4]	-2.25[4]	-3.95[3]	-4.51[1]	-4.31[1]
India	-2.37[4]	2.31[1]	-0.85[4]	-3.94[3]	-7.51[2][c]	-3.27[2]
Indonesia	-2.07[3]	-0.96[1]	-1.00[1]	-13.8[2]	-4.87[2]	-4.67[1]
Korea	-0.25[4]	-1.65[1]	-3.00[1]	-11.9[2]	-5.14[1]	-3.85[4]
Malaysia	-2.37[4]	-0.08[3]	0.84[2]	-4.49[1]	-5.74[4][c]	-4.71[1]
Mexico	-1.14[4]	-2.31[2]	3.09[4]	-3.48[3]	-3.58[4]	0.29[4][c]
Pakistan	-3.40[4]	-1.41[4]	0.51[1]	-	-6.60[1]	-3.62[4]
Philippines	-2.32[2]	-1.39[2]	-0.67[2]	-6.40[2]	-4.68[4]	-3.81[4]
Portugal	-1.11[3]	-2.89[4]	-2.15[4]	-9.17[2]	-6.38[1]	-8.13[2][c]
Singapore	-1.90[4]	-2.00[4]	-1.54[4]	-3.88[3]	-5.33[4][c]	-4.32[1]
S. Africa	-3.50[1]	-2.01[3]	-0.71[3]	-	-3.61[2]	-3.32[2]
Thailand	-2.38[3]	-1.69[4]	-1.67[4]	-7.83[2]	-5.42[4][c]	-8.89[2][c]
Tunesia	-2.46[3]	-1.63[4]	-0.15[1]	-5.53[4]	-3.74[4]	-3.53[4]
Turkey	-3.94[1]	-1.44[1]	3.00[3]	-	-4.94[3]	-1.17[3][c]
Venezuela	-3.39[3]	-1.71[2]	-0.52[2]	-	-6.38[1]	-5.77[1]

Notes: a. The Mackinnon (1991) critical value of the ADF statistic when a trend term is included in the procedure (for 50 observations) are -3.17 at the 10% level of significance.

b. Numbers inside the brackets are number of lags in equation 3, i.e., value of k.

c. These ADF statistics are obtained when second differenced data are used. The ADF statistics for the first diffirenced data were still larger than the critical values.