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Effects of Devaluation on the LDCs Demand for International Reserves

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In a recent article Bahmani-Oskooee and Niroomand (1988) investigated the effects of a depreciation on the industrial countries' demand for international reserves. They demonstrated that real effective exchange rate depreciation had a significantly negative effect on the developed countries' (DCs) demand for international reserves. This conclusion was reached by estimating a reserve demand function for a sample of 13 DCs using pooled quarterly data over 1973-85 period.[1] Could be depreciation of the real effective exchange rate exert significantly negative effect on the less developed countries' (LDCs) demand for international reserves? This is the issue we attempt to investigate in this paper.

The primary purpose of this paper is to assess the effects of a change in the real effective exchange rate on the demand for international reserves by LDCs. In addition to this practice, we also try to tackle some of the issues that were raised against Bahmani-Oskooee and Niroomand's work by Sephton (1989), though they have already been discounted by the former authors.[2] Since the model and methodology is similar to that of Bahmani-Oskooee and Niroomand, only a short account of the model will be provided in section I. Section II presents our estimation results. Section III concludes the study. Finally, data definition and sources are provided in the appendix.

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1 The criteria for selecting the 13 DCs was basd on the fact that effective exchange rate is only published for these countries in the International Financial Statistics of the IMF.

2 See Bahmani-Oskooee and Niroomand (1989).

I. The Model

Following Bahmani-Oskooee and Niroomand, we assume that the long-run reserve demand function takes the following form:

$$(1)\quad \text{Log}(R/P)_{it} = c_0 + c_1\text{Log}(Y/P)_{it} + c_2\text{Log}\sigma^{BP}_{it} + c_3\text{Log}(P^*E/P_d)_{it} + u_{it}$$

where R = level of nominal reserves in US dollars,
P = US implicit price deflator for GNP,
Y = nominal GNP in US dollars,
σ^{BP} = variability measure of balance of payments,
P^* = foreign price level,
E = effective exchange rate,
P_d = domestic price level,
u = error term.

It is expected that estimate of c_1 and $c_2 > 0$ and as the appendix indicates, since the effective exchange rate, E, is defined as units of domestic currency per unit of foreign currency, an increase in it, i.e., depreciation of domestic currency is supposed to lower the trade deficit leading to a decline in the demand for international reserves. Thus, it is expected that $c_3 < 0$.[3]

II. The Estimation Results

In this section we try to estimate the reserve demand equation (1) by pooling time-series and cross-sectional data from 15 developing countries.[4] Due to this pooling, one common practice is to assume that the disturbance terms in the model are corss-sectionally heteroskedastic and time-wise autoregressive. Moreover, those who have made such assumptions, have relied solely upon first order serial correlation.[5] Now that computer programs are available, one can investigate the incidence of higher order serial correlation. Indeed, this was one of the main criticisms

[3] For more on the signs of these coefficients see Bahmani-Oskooee and Niroomand (1988) and for a comprehensive survey of literature on the demand for reserves see Bahmani-Oskooee (1985).

[4] See Table 1 for the list of countries.

[5] Frenkel and Hakkio (1980), Edwards (1980), and Bahmani-Oskooee and Niroomand (1988) are some examples on this point.

made by Sephton (1989) on Bahmani-Oskooee and Nirromand's (1988) model. To determine whether the disturbance terms in equation (1) follow, say, third order serial correlation of the following type:

$$(2) \qquad u_{it} = \rho_{i1}u_{it-1} + \rho_{i2}u_{it-2} + \rho_{i3}u_{it-3} + \epsilon_{it}$$

we estimated the reserve demand equation (1) for each country separately, taking account of specification (2). We then looked at the size and significance of estimated ρ's to determine the order of the serial correlation for each country. Table 1 reports the estimated ρ's along with the corresponding t-statistics.

Table 1 reveals that while estimates of ρ_1 are significant for most countries, the estimates of ρ_2 and ρ_3 are not. Indeed, this justifies the assumption of first order serial correlation in all previous studies.[6] Therefore, following Bahmani-Oskooee and Niroomand (1988) we assume that the error term in equation (1) is subject to the following assumptions:

$$(3) \qquad E(u^2_{it}) = \sigma^2_i \text{ (heteroskedasticity)}$$

$$(4) \qquad E(u_{it}u_{jt}) = 0 \text{ for } i \quad j \text{ (cross-section independence)}$$

$$(5) \qquad u_{it} = \rho_i u_{it-1} + e_{it} \text{ (autoregression)}$$

where $e_{it} \sim N(0, \sigma^2_{ei})$,

$u_{i0} \sim N\{O, \sigma^2_{ei}/(1-\rho^2_i)\}$,

and $E(u_{it-1}e_{jt}) = 0$ for all i, j.

Estimating the reserve demand equation (1) where assumptions (3)-(5) are incorporated into estimation process, amounts to obtaining estimates of ρ_i and σ_i and transforming the data by $(X_{it}-\hat{\rho}_i X_{it-1})/\hat{\sigma}_i$ where X is the vector of all endogenous and exogenous variables. Application of OLSQ to the transformed data results in unbiased and consistent estimates of the coefficients.[7] Table 2 presents the estimates of ρ and σ for each country.

We are now in a position to use the estimates of ρ's and σ's from Table 2 to estimate equation (1). Following Bahmani-Oskooee and Niroomand (1988), we included country-specific dummy variables in equation (1) and

[6] As a further exercise, we also tested for second order serial correlation among disturbance terms. Once again, estimates of ρ_1 were highly significant for almost all countries, whereas, estimates of ρ_2 were not.

[7] For more on this estimation procedure see Kmenta (1971, p. 508).

Table 1

ESTIMATES OF ρ_1, ρ_2, AND ρ_3 FROM AN AR(3) PROCESS

Country	$\hat{\rho}_1$	$\hat{\rho}_2$	$\hat{\rho}_3$
Argentina	1.0429	0.0867	−0.2493
	(6.77)	(0.38)	(1.54)
Brazil	1.4997	−0.5791	0.0020
	(8.75)	(1.99)	(0.01)
Colombia	1.1762	0.0240	−0.2562
	(7.66)	(0.10)	(1.59)
Ecuador	0.8504	0.2585	−0.3186
	(5.61)	(1.32)	(2.15)
Egypt	0.5428	−0.3330	0.3220
	(3.73)	(2.08)	(2.40)
Korea	0.8881	0.0271	−0.0731
	(5.60)	(0.12)	(0.43)
Malaysia	1.0439	−0.1460	−0.0475
	(6.57)	(0.66)	(0.29)
Morocco	0.1596	0.5705	0.2214
	(1.04)	(4.36)	(1.44)
Nigeria	1.2164	−0.0269	−0.2422
	(7.22)	(0.09)	(1.35)
Philippines	0.2954	0.2746	0.3019
	(1.92)	(1.66)	(1.91)
Portugal	0.7433	−0.0882	0.2361
	(4.83)	(0.44)	(1.51)
Singapore	0.8182	−0.2013	0.2682
	(5.08)	(1.03)	(1.92)
Tunisia	0.9374	−0.5632	0.7076
	(7.23)	(3.13)	(4.44)
Yugoslavia	0.9546	−0.2579	−0.0174
	(6.15)	(1.24)	(0.12)
Zambia	0.5715	0.4906	−0.1934
	(3.38)	(2.86)	(1.22)

Note: Absolute values of the t-ratios are in parentheses.

Table 2

ESTIMATES OF ρ_i AND σ_i FROM THE LONG-RUN MODEL

Country	$\hat{\rho}_i$	$\hat{\sigma}_i$
Argentina	0.856681	0.299453
Brazil	0.826442	0.201366
Colombia	0.947177	0.205679
Ecuador	0.664631	0.261038
Egypt	0.809276	0.299897
Korea	0.949671	0.167334
Malaysia	0.846610	0.173072
Morocco	0.857667	0.597190
Nigeria	0.960870	0.2922033
Philippines	0.796131	0.300496
Portugal	0.874692	0.230662
Singapore	1.015597	0.129508
Tunisia	0.570447	0.313270
Yugoslavia	0.961995	0.211064
Zambia	0.909585	0.462727

then applied OLSQ to the transformed data.[8] The LSDV (least squares with dummy variables) results on the transformed data are reported in equation (6) with the t-ratios enclosed in parentheses below each coefficient. The estimated country specific dummy variables are reported in Table 3.

$$(6) \quad \underset{}{\mathrm{Log}(R/P)} = c'_0 + \underset{(0.19)}{0.0054\mathrm{Log}(Y/P)} + \underset{(5.11)}{0.2059\mathrm{Log}\sigma^{BP}}$$

$$-\underset{(2.69)}{0.1666\mathrm{Log}(P^{*}E/P_d)}$$

$$R^2 = 0.91 \qquad F = 467 \qquad D\text{-}W = 1.8674$$

The results reported in (6) together with those in Table 3 show that all estimated elasticities carry their theoretically expected signs and they are all highly significant except for real income elasticity.

Concentrating on the central theme of this paper, we gather that the

8 This method of least squares with dummy variables is what Edwards (1984) and Bahmani-Oskooee and Niroomand have labeled LSDV method.

Table 3

ESTIMATES OF COUNTRY-SPECIFICCONSTANTS

Country	Estimated Coefficient	t-statistics
Argentina	3.2918	13.8731
Brazil	6.6094	15.1577
Colombia	1.7421	11.1385
Ecuador	7.1747	13.7711
Egypt	3.6575	12.4058
Korea	2.0161	10.8685
Malaysia	6.2344	15.8587
Morocco	1.1403	7.6229
Nigeria	0.8992	6.9895
Philippines	4.3644	14.1808
Portugal	3.5287	13.9098
Singapore	−0.6865	−5.6059
Tunisia	7.3179	13.4510
Yugoslavia	1.0593	7.4558
Zambia	0.7278	5.3761

real effective exchange rate elasticity is −0.1666 for LDCs in this paper. The comparable figure for developed countries obtained by Bahmani-Oskooee and Niroomand was −0.2634. This shows that depreciation lowers DCs demand for reserves more than LDCs. This is in line with collective wisdom, due to the fact that DCs have enjoyed more flexibility in their exchange rates than LDCs.

Our finding for LDCs provides indirect support for Edwards (1983). After classifying LDCs into those that maintained fixed exchange rate and those that devalued occasionally during 1964-1972 period Edwards concluded that fixed exchange rate countries tend to correct discrepancies between actual and desired reserves more slowly than countries that devalue their currency.

Before concluding this study we would like to consider another issue raised about Bahmani-Oskooee and Niroomand's work by Sephton concerning the stability test of the reserve demand function. Previous studies that have provided some evidence of structural instability, have concentrated on the demand for nominal reserves. More precisely, while Heller and Khan (1978) and Frenkel (1978, 1980) showed that there was structural shift in the demand for nominal reserves in 1973 due to a change in the international monetary system from fixed to floating exchange rates, Bahmani-Oskooee (1988a) showed that the structural shift toward the end of 1979 was mainly due to an oil price hike. Since our period of analyais is inclusive of 1979 during which oil prices were almost doubled, we attempted to test the stability of our estimates for the period under study. To this end, we first use Quandt's likelihood ratio statistics to detect the switching point and then use the Chow test to determine the significance of that switch.[9] Quandt's likelihood ratio statistics is shown to be:[10]

$$(7) \qquad \lambda_r = \frac{1}{2} r\mathrm{Log}\sigma^2{}_1 + \frac{1}{2}(T-r)\mathrm{Log}\sigma^2{}_2 - \frac{1}{2}T\mathrm{Log}\sigma^2$$

where $\sigma^2{}_1$ and $\sigma^2{}_2$ denote the ratio of sum of squared residuals to the number of observations when the regression is fitted to the first r and the remaining T–r observations, respectively. σ^2 then carries the same definition when the regression is fitted to the total number of observations, T. The switching point is determined by that value of r for which λ_r attains its minimum.

After applying the LSDV method to the transformed data for the entire period as well as for different combinations of subperiods, λ was minimized when subperiods were divided as 1973I-1980I and 1980II-1985IV. To test the significance of this switch or the equality of the regression coefficients between the two subperiods, using the sum of squared residuals from the whole period results and from the subperiods results an F-statistics of –0.66 was obtained.[11] Since our calculated F is much lower than its critical value from the F-table at any level of significance, we conclude that the switch in 1980 was not significant which implies that there was no structural shift in the reserve demand function. This finding is in contrast with that of Bahmani-Oskooee (1988a) who concluded that both DCs and LDCs demand for reserves experienced structural instability in 1980 and 1981 respectively. Several factors may explain this inconsistency.

First, previous studies like Heller and Khan (1978) and Frenkel (1978, 1980) as well as Bahmani-Oskooee (1988a) investigated the stability of the demand for nominal reserves whereas, in this paper we assess the stability of the demand for real reserves. The US implicit price deflator used to deflate nominal reserves is already inclusive of a change in oil prices. Therefore, rather than acting as an exogenous shock, a change in oil price

9 Notice that we are following the methodology of Bahmani-Oskooee (1988a).

10 For Quandt's likelihood ratio statistics see Brown, Durbin, and Evans (1975).

11 Notice that a negative F-statistics is possible due to the fact that there is no constant term in any of the estimated equations. Rather than including a constant term we have included 15 country specific dummy variables.

is already included in the model, through a change in P.[12] Second, another contributing factor could be the fact that we have incorporated country specific dummy variables in our model whereas, previous studies failed to do so. An increase in oil price may have a different impact on the demand for real reserves by different countries. Country specific factors may capture these different effects, yielding an overall stable reserve demand function. Furthermore, part of these different impacts could also be captured by each country's domestic price level, P_d, also included in our model, but not in the previous ones.

Another point that was subject to debate between Sephton and Bahmani-Oskooee and Niroomand was equilibrium versus disequilbrium model specification. In order to see in what way the results will change, we estimated a dynamic reserve demand equation which includes a lagged dependent variable as one of the exogenous variables. In estimating this disequilibrium model we made sure that the disturbance term, once again, was subject to assumptions (3)-(5). The coefficient estimates are reported in equation (8) with the t-ratios below each estimated coefficient and $\hat{\rho}_i$'s and $\hat{\sigma}_i$'s along with the country specific dummies in Table 4.

$$(8) \qquad \text{Log}(R/P) = a'_0 + \underset{(1.17)}{0.034}\text{Log}(Y/P) + \underset{(1.61)}{0.062}\text{Log}\sigma^{BP} - \underset{(1.35)}{0.083}\text{Log}(P^{*}E/P_d) + \underset{(3.29)}{0.131}\text{Log}(R/P)_{t-1}$$

$$R^2 = 0.88 \qquad F = 314 \qquad D\text{-}W = 1.9944$$

From the reported results in (8) we gather that all estimated coefficients again carry their expected signs, though, at somewhat lower level of significance. More precisely, the exchange rate elasticity that was highly significant in the equilibrium model, is now only significant at the 10% level.[13] This is still an indication of the fact that depreciation of the real effective exchange rate could lower the LDCs demand for international reserves. If the size of adjusted R^2 is a criteria for model selection, then the equilibrium modeo, which provides strong evidence in support of the central theme of this paper, possesses a better predictive power than the disequilibrium model.

12 Similar point is also discussed in Bahmani-Oskooee (1988b, footnote 12).

13 Notice that in order to obtain the long-run elasticities outlined by equilibrium model (1) we need to divide the short-run estimates reported in (8) by one minue the estimated coefficient of the lagged reserves.

Table 4

ESTIMATES OF ρ_i, σ_i AND COUNTRY-SPECIFIC FACTORS FOR DISEQUILIBRIUM MODEL

Country	$\hat{\rho}_i$	$\hat{\sigma}_i$	Country-Specific Factors	
			Estimated Coefficient	t-ratio
Argentina	0.843913	0.294387	3.2332	11.07
Brazil	0.881597	0.174631	4.6065	11.64
Colombia	0.965153	0.195247	1.0859	6.89
Ecuador	0.634166	0.273159	6.5010	10.98
Egypt	0.790489	0.297344	3.5213	9.96
Korea	0.960667	0.165909	1.2061	6.93
Malaysia	0.951998	0.129629	2.3714	10.56
Morocco	0.859063	0.608662	0.9329	5.55
Nigeria	0.983927	0.286098	0.3314	2.42
Philippines	0.875718	0.275021	2.5189	10.09
Portugal	0.898409	0.086628	6.6765	12.11
Singapore	1.013636	0.123892	−0.5402	−3.85
Tunisia	0.698917	0.261663	5.3097	10.61
Yugoslavia	0.953824	0.206353	1.1527	6.57
Zambia	0.906518	0.434655	0.6531	4.11

III. Summary and Conclusion

A body of the literature in international finance includes some well known studies that have argued that one way to reduce a country's demand for international reserves is through the depreciation of that country's currency. Recently, a reserve demand function that included the real effective exchang rate as another determinant of the demand for reserves was introduced. Since the effective exchange rates were only available for some developed countries, that model was only estimated for DCs. The empirical results supported the common wisdom that indeed devaluation lowers the demand for reserves by DCs. In the absence of effective exchange rates for LDCs, the LDCs were excluded from empirical analysis.

The purpose of this paper was to assess the effects of a change in the real effective exchange rates on the LDCs demand for international reserves by relying upon the reserve demand equation that included real effective exchange rate. After constructing the quarterly effective exchange rates for 15 less developed countries for 1973I-1985IV period, the long-run and short-run reserve demand equations were estimated and it was concluded that indeed depreciation of LDCs' real effective exchange rates reduce their demand for international reserves.

Appendix
Data Definition and Sources

A. Data Sources: All Data are Quarterly for 1973-1985 Period and are the Following Sources:

a. International Financial Statistics of IMF.
b. Survey of Current Business
c. OECD Statistics of Foreign Trade, Series A.

B. Definition of Variables:

International Reserves (R): measured in terms of millions of SDRs, converted into millions of US dollars using SDR-Dollar rate and defined as the sum of gold, Special Drawing Rights, foreign exchange and reserve position at the fund. Quarterly data on this variable are from source a.

U.S. Implicit Price Deflator for GNP (P): 1975 = 100, source b.

Income (Y): measured as GNP (GDP in some cases) in millions of domestic currency, converted into US dollars using period average exchange rates from source a. Since the quarterly GNP figures were not available for countries in this study, they were generated using annual figures and the methodology of Bahmani-Oskooee (1986). The annual GNP and import figures used in this process are from source a.

Variability Measure of Balance of Payments (σ^{BP}): following Kelly (1970) for each quarter this variable is defined as the standard deviation of exports (in millions of US dollars) over the preceding eleven quarters and the current quarter. The exports data are from source a.

Foreign Price Level (P*): this variable is defined as the weighted average of CPI of industrial countries (1975 = 100). This index is available from source a.

Nominal Effective Exchange rate (E): for each country this variable is constructed as the import weighted effective rate where the weights belong to each country's major trading partners, mostly OECD countries. The bilateral exchange rates are from source a and the import weights from source c.

Domestic Price Level (P_d): for each country this variable is measured by CPI (1975 = 100), source a.

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