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SHORT-RUN AND LONG-RUN EFFECTS OF CURRENCY DEPRECIATION ON THE BILATERAL TRADE BALANCE BETWEEN PAKISTAN AND HER MAJOR TRADING PARTNERS

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Previous studies that investigated the short-run (J-curve) and the long-run effects of currency depreciation on the trade balance of Pakistan used aggregate trade data between Pakistan and the rest of the world and provided no evidence of any significant impact. We wonder whether lack of the relation is due to aggregation bias. In this paper, therefore, we go one step further by employing disaggregated data at bilateral level between Pakistan and her 13 major trading partners to determine if we can discover partners whose trade balances react to changes in the real bilateral exchange rate. The results from bounds testing approach are still inconclusive and show that only in half of the cases the real bilateral exchange rate plays a role.

Keywords: Bilateral J-Curve, Bounds Testing, Cointegration, Pakistan
JEL classification: F31

1. INTRODUCTION

A large number of studies that examine short run and long run relationships between exchange rate and trade balance have been conducted in the last few decades. Magee (1973) was one of the very first attempts in the literature to outline possibility of the J-curve phenomenon. He put forward two possible reasons for the existence of a deteriorating trade balance in the wake of currency depreciation. First, there are contract rigidities that take time to wear off. Second, there is a pass-through effect of currency depreciation on domestic prices which may not take place until some time has passed after such depreciation. As a result, favorable effects of exchange rate depreciation may not be immediately visible even though the long run elasticities satisfy the Marshall -Lerner condition.

Over the next decades several studies sought to gather evidence for Magee's

[*] Valuable comments of an anonymous referee are appreciated. Any error, however, is ours.

hypotheses. Important works from this period include Miles (1979), Bahmani-Oskooee (1985), Flemingham (1988), Meade (1988), Rosenweig and Koch (1988), Noland (1989), Marquez (1991), and Marwah and Klein (1996) among others. These studies experimented with various econometric models, introduced new definitions for the endogenous and exogenous variables, covered different time periods and included a wide range of countries in their analysis. The empirical evidence however remained mixed. Almost all of the works discussed above relied on either the ordinary least squares (OLS), instrumental variables (IV) or the two-step least squares (2SLS) techniques, all of which are subject to the hazards of spurious correlation unless the time series data under consideration is stationary, thus making their predictions somewhat untenable.[1]

By late 1980s advancements in econometric theory had allowed researchers to estimate short run and long run relationships in the presence of non-stationary time series data. The ground-breaking econometric advances in this direction are due to Sims (1980) who pioneered the vector autoregressive (VAR) technique, Engle and Granger (1987) who introduced a two step cointegration test in an error correction modeling framework and, Johansen and Juselius (1990), and Johansen (1991) who proposed cointegration tests for VAR models based on the maximum likelihood method. This availability of advanced cointegration techniques in time series analysis ushered in a new round of empirical testing from early 1990s to early 2000s. Again, however, the empirical evidence was mixed with some studies supporting the existence of a short run J-curve while others rejecting it. The cointegration techniques discussed above require that all time series variables included in the analysis be integrated of the same order. For models that contain both stationary and non-stationary variables, transformation of the data may not be a trivial task, not to mention that the results for such transformed variables can be notoriously difficult to interpret.

In the last few years, a new approach to error correction modeling introduced by Pesaran, Shin and Smith (2001) called the bounds testing approach has been employed in time series analysis. This technique can be applied to models in which exact order of integration of variables, though unknown, is not greater than I(1). In other words, the bounds testing approach can be used in any of the following situations: when all variables are I(0), when all variables are I(1), and when the variables are either I(0) or I(1) or combination of the two. In addition, the bounds testing approach makes it relatively simple to derive the short-run and the long-run effects of one variable on the other. Bahmani-Oskooee and Ratha (2004a) provide a comprehensive review of the literature showing who has applied which technique in testing the J-Curve.

[1] For some other studies that deal with trade-related issues see Agbola and Damoense (2005), Alse and Bahmani-Oskooee (1995), Charos *et al.* (1996), Du and Zhu (2001), King (1993), Love and Chandra (2005), Narayan and Narayan (2005), Truett and Truett (2000), Seyoum (2007), and Bahmani-Oskooee and Hegerty (2007).

Since this paper is about Pakistan, a brief review of the literature about Pakistan's experience is in order. The number of studies that have included Pakistan in their investigation of the J-curve is small. Many studies investigate only the long run relationship between exchange rate and trade balance while ignoring the J-curve altogether. Furthermore, almost all of the studies that include Pakistan in their analysis either use the OLS, IV or the 2SLS techniques, all of which are prone to the problem of spurious correlation unless the time series under consideration are stationary, thus casting considerable doubt on their findings. Because of the stationarity problem, at the very least the empirical results in those studies cannot be directly compared with the ones that employ recently developed econometric methods such as the VAR and error-correction modeling technique.

Gylfason and Schmid (1983) used aggregate data on five developed and five less developed countries and incorporated both demand and supply side effects of real exchange rate depreciation into their model. They found support for a long run relationship between exchange rate and trade balance with an expected increase in trade balance due to a 10% devaluation of Pakistan's rupee to be equal to 1.3% of Pakistani GNP. However, since the data used were not tested for stationarity, their empirical results are somewhat biased.

Bahmani-Oskooee and Alse (1994) formulated their model following the error correction specification proposed by Engel and Granger (1987). They used aggregate quarterly data from 1970I to 1990IV for 19 developed and 22 less developed countries. Once they controlled for stationarity properties of regression variables, support for co-integration between real exchange rate and trade balance for Pakistan could not be found. It should be noted however that the model used in this study regressed trade balance directly on the real exchange rate without controlling for other variables such as income. Short run dynamics for countries that failed the co-integration test which included Pakistan were ignored.

Bahmani-Oskooee (1998) employed the Johansen and Juselius maximum likelihood co-integration technique to estimate the well-known Marshall-Lerner condition for six countries using quarterly data over the period 1973I-1990IV. In the results for Pakistan, the Marshall-Lerner condition which implies that the sum of import and export demand elasticities must add up to more than one was not met. This finding using aggregate data to test the Marshall-Lerner condition is in line with that of Bahmani-Oskooee and Alse (1994) who could not establish cointegration between Pakistani trade balance and the real exchange rate. These findings were, however, contradicted by Aftab and Aurangzeb (2002) who used Johansen and Juselius method and quarterly data over the period 1980-2000 to show that the long-run Marshall-Lerner condition for Pakistan is satisfied. Although their method is an improvement over Bahmani-Oskooee (1998), it may still suffer from aggregation bias as bilateral data for individual trading partners were not employed.

Due to conflicting findings by previous research, as reviewed above, we would like to reconsider the short-run (i.e., the J-curve) and the long-run effects of real depreciation of Pakistani rupee on her trade balance. However, unlike previous research we employ trade data at bilateral level between Pakistan and her 13 major trading partners, a practice originally introduced by Rose and Yellen (1989) for the trade balance between U.S. and her seven trading partners. These 13 partners account for almost 70% of overall trade activity of Pakistan in 2003. In order to get some insight about the relative importance of each partner, we provide these trade shares in Table 1.[2]

Table 1. Bilateral Trade Flow Between Pakistan and Her Major Trading Partners in 2003

Trading Partners	Value of Exports (Millions of U.S. dollars)	Value of Imports (Millions of U.S. dollars)
China	447	1,858
France	310	318
Germany	598	672
Hong Kong	472	117
Italy	394	352
Japan	145	838
Korea	259	371
Kuwait	70	918
Malaysia	74	590
Saudi Arabia	448	1,492
U.A.E.	1,013	1,632
U.K.	795	525
U.S.A.	2,528	924
World total	11,283	14,825

Source: Direction of Trade Statistics 2004, International Monetary Fund.

The remaining of the paper is composed of three additional sections. Section 2 presents the model and the method that is based on bounds testing approach to cointegration and error-correction modeling. Empirical results are presented and discussed in sections 3, and formal concluding remarks summarizing the overall findings are presented in section 4. Finally, data sources and definition of variables are cited in an appendix.

[2] There is now a growing literature on testing the J-Curve at the bilateral level. Some examples are: Shirvani and Wilbratte (1997), Bahmani-Oskooee and Kanitpong (2001), Wilson (2001), Bahrumshah (2001), Bahmani-Oskooee and Goswami (2003), Bahmani-Oskooee and Ratha (2004b, 2004c), Bahmani-Oskooee, *et al.* (2005), Bahmani-Oskooee, *et al.* (2006), and Bahmani-Oskooee, *et al.* (2008).

2. THE MODEL AND METHOD

In assessing the short-run and the long-run effects of changes in the exchange rate on the trade balance, whether at the aggregate or at the bilateral level, it is a common practice to regress a measure of trade balance directly on real exchange rate while controlling for real income at home and in foreign country. In specifying such a trade balance model between Pakistan and her trading partner *i*, we follow the specification by Bahmani-Oskooee and Brooks (1999) and Arora *et al.* (2003) as outlined by equation (1):

$$\log TB_{i,t} = \alpha + \beta \log Y_{Pakistan,t} + \gamma \log Y_{i,t} + \varphi \log REX_{i,t} + \varepsilon_t \,. \tag{1}$$

This specification expresses trade balance between Pakistan and trading partner *i* (TB_i) defined as the ratio of Pakistan's nominal imports from trading partner *i* to her nominal exports to the same trading partner as a function of Pakistan's income, $Y_{Pakistan}$, income of trading partner *i*, Y_i, and the real bilateral exchange rate. We expect an estimate of β to be positive as an increase in domestic (Pakistan) income generally leads to an increase in imports. A negative estimate for β is possible if increase in domestic income reflects expansion in the production of import-substitute goods (Bahmani-Oskooee (1986)). An estimate of γ is expected to be negative as an increase in trading partner's income leads to higher exports by Pakistan. However, a positive estimate of γ is possible if increase in foreign income comes from an expansion in foreign production of substitutes for Pakistani export goods. Finally, as the appendix shows REX_i is defined in a way that a decrease reflects a real depreciation of Pakistani rupee. If depreciation is to decrease imports and increase exports, hence improve the trade balance, an estimate of φ would be positive

Since the model given in (1) is a long run relationship and the J-curve phenomenon occurs in the short run, it is necessary to modify (1) in order to incorporate the short-run dynamics. A common practice is to express (1) in an error-correction modeling format. We do this by following Pesaran *et al.*'s (2001) bounds testing approach as in (2):

$$\begin{aligned}\Delta \log TB_{i,t} = {} & \alpha + \sum_{k=1}^{n} \eta_k \Delta \log TB_{i,t-k} + \sum_{k=0}^{n} \beta_k \Delta \log Y_{c,t-k} + \sum_{k=0}^{n} \gamma_k \Delta \log Y_{i,tk} \\ & + \sum_{k=0}^{n} \varphi_k \Delta \log REX_{i,t-k} + \delta_1 \log TB_{i,t-1} + \delta_2 \log Y_{Pakistan,t-1} \\ & + \delta_3 \log Y_{i,t-1} + \delta_4 \log REX_{i,t-1} + \upsilon_t .\end{aligned} \tag{2}$$

Pesaran *et al.* (2001) show that for the error correction specification in (2) there is no need to test for unit roots as long as all the variables involved are either I(0) or I(1) or combination of the two. In order to justify the retention of lagged level variables, we need to test whether their coefficients are jointly significant. In other words, the null hypothesis of no cointegration, i.e., $H_0: \delta_1 = \delta_2 = \delta_3 = \delta_4 = 0$ is tested against the alternative of $H_1: \delta_1 \neq 0, \delta_2 \neq 0, \delta_3 \neq 0, \delta_4 \neq 0$. Pesaran *et al.* (2001) propose applying the familiar *F* test with new critical values that they tabulate. An upper bound critical value is tabulated if all variables are I(1) and a lower bound critical value is tabulated if all variables are I(0). An acceptance of the null hypothesis would thus provide evidence against co-integration while its rejection would provide evidence in support of co-integration. In this set up, the short-run effects of real depreciation is judged by the estimates of φ_k's. Negative values for lower lags followed by positive values for higher lags will indeed support the J-curve. The long-run effects of real depreciation are inferred by the estimate of δ_4 that is normalized on δ_1.

3. EMPIRICAL RESULTS

The error-correction model outlined by Equation (2) is estimated between Pakistan and each of her 13 partners using quarterly data over the period 1980-2003. The first step is to select the number of lags of first differenced variables. Bahmani-Oskooee and Brooks (1999) have shown that the results of the *F* test will depend on the number of lags. In order to see how the *F* test reacts to number of lags selected, as a starting exercise two, four, six and eight lags are introduced. The calculated *F* values for these models are presented in Table 2. As can be seen from the results, the *F* values are sensitive to the number of lags imposed. As we move from 2 lags to 8 lags, the number of significant cases at 10% level of significance drops from 11 to just 2.

Table 2. The Results of the *F* Test at Different Lags

Trading Partners	2 lags	4 lags	6 lags	8 lags
China	12.0122	3.8139	3.5187	1.9880
France	6.4872	3.6668	3.7595	3.5733
Germany	4.5402	4.2837	2.0405	1.6201
Hong Kong	4.2470	2.9708	1.9853	1.4031
Italy	2.8284	1.8739	1.1373	1.8288
Japan	4.1240	3.7934	1.2314	1.1215
Korea	2.3109	1.0972	2.7453	1.3999
Kuwait	6.6424	5.4359	3.2146	1.8568
Malaysia	4.1733	2.7942	1.9837	1.2673
Saudi Arabia	4.3051	3.3158	4.0083	3.5524
U.A.E.	4.5701	2.8070	0.9683	1.1089
U.K.	6.8756	4.5153	3.6117	3.3482
U.S.A.	6.3129	3.2869	3.9560	2.2607

Note: Critical values of *F* statistic at 5% and 10% levels of significance are 4.01 and 3.52 respectively.

Source: These are taken from Pesaran, Shin and Smith (2001).

Given such sensitivity of test results to number of lags, following Bahmani-Oskooee and Gelan (2006) we rely on some information criterion in order to select the optimal number of lags and carry the *F* test at optimum lags. The two criteria considered in this study are the Akaike Information Criterion (AIC) and the Schwarz Bayesian Criterion (SBC). *F* test results and number of optimal lags for models using AIC and SBC are presented in Table 3. These results show that the number of significant cases at 10% level of significance for models with AIC-selected lags is 11 while this number increases to 12 for models using SBC-selected lags. Furthermore, all the *F* values that are significant at 10% level of significance under both AIC and SBC, unlike Table 2 results, are also significant at 5% level of significance.

Table 3. *F* Statistics at AIC and SBC-Selected Optimal Lags

Trading Partners	AIC Optimal lags	*F* Statistic at AIC-Selected Optimal Lags	SBC Optimal Lags	*F* Statistic at SBC-Selected Optimal Lags
China	3,0,3,0	8.2941	1,0,0,0	30.3398
France	0,1,0,0	19.4358	0,1,0,0	19.4358
Germany	0,3,0,0	10.2310	0,0,0,0	9.7009
Hong Kong	0,3,7,0	9.6113	0,0,0,0	10.0022
Italy	6,0,0,2	1.3540	0,0,0,0	5.3217
Japan	0,3,0,0	11.2341	0,0,0,0	6.9143
Korea	2,0,1,0	2.4253	2,0,0,0	2.3155
Kuwait	2,0,4,3	7.2734	0,0,1,0	18.5362
Malaysia	2,0,1,0	5.0955	0,0,0,0	9.9226
Saudi Arabia	6,7,8,0	4.6881	0,0,3,0	6.8545
U.A.E.	1,0,0,1	4.7006	0,0,0,0	12.7654
U.K.	0,0,0,1	13.5272	0,0,0,1	13.5272
U.S.A.	4,0,0,0	6.0962	0,0,0,0	15.2139

Notes: a) Critical values of *F* statistic at 5% and 10% levels of significance are 4.01 and 3.52 respectively. These are taken from Pesaran, Shin and Smith (2001). b) The number of lags follows the specification in model (2). Thus 3,0,3,0 for China means that three lags were imposed on $\Delta \log TB_i$, zero lags on $\Delta \log Y_{Pakistan}$, three lags on $\Delta \log Y_i$, and zero lags on $\Delta \log REX_i$.

It should be noted here however that Table 3 results are sensitive to lag selection criterion. This can be observed in case of Italy for which the calculated *F* value is not significant when AIC is used but becomes significant under SBC. For this reason, and because of additional co-integration analysis presented later in this work, it is decided to keep the lagged values in all cases even where *F* test is not significant.

(Table 4 is here.)

In order to assess the J-curve or the short-run effects of real depreciation, we next report the coefficient estimates obtained for $\Delta \log REX_{t-1}$ variables. Since AIC criterion chooses longer lags, here we restrict ourselves to reporting AIC based results. These results are reported in Table 4.

As indicated before, existence of the J-curve can be inferred by looking at the coefficient estimates of $\Delta \log REX_{t-i}$. Negative coefficients followed by positive ones would support the J-curve. The results in Table 4 suggest that these coefficient estimates follow the J-curve pattern only for Italy even though some of these estimates are not significant. Although results in Table 4 do not support the existence of J-curve, there is at least one significant coefficient at the 10% level in the cases of China, Italy, Korea, Kuwait, U.A.E., and the U.K., suggesting that importance of real exchange rate as a determinant of trade balance in the short run cannot be completely ignored.[3] The next question is in how many of these countries, the short-run significant effects last into the long run. To this end, we report in Table 5 and 6 estimates of δ_2, δ_3 and δ_4 normalized on δ_1 from both AIC based and SBC based models.

Table 5 results show that the coefficient on real exchange rate is positive in twelve out of thirteen cases, and is significant and positive for six cases at 5% level of significance. These results signal that a long run relationship between real exchange rate and trade balance cannot be ignored. The coefficient on Pakistani income is positive for six of the trading partners but is significant for only five at 5% level of significance. For Hong Kong and Japan, real Pakistani income has a negative but significant coefficient. These results provide support for a long term relationship between real Pakistani income and her trade balance. Coefficients on incomes of trading partners have the correct sign and significance at 5% level of significance for only two partners, Germany and U.K. For Kuwait and U.A.E. the coefficients are significant but have a positive sign.

The situation in Table 6 is almost similar with the only difference that Saudi Arabia now has a negative coefficient on real exchange rate and the coefficient for Japan is no longer significant. For five countries the coefficients are positive and significant, again suggesting that a long run relationship between real exchange rate and trade balance cannot be ruled out. The coefficient on real Pakistani income is positive for more than half of the trading partners but is significant and positive for only three at 5% level of significance. Two trading partners have a negative but significant coefficient on real Pakistani income. Overall these results provide support for a long run relationship between Pakistani income and her trade balance. Coefficients on income of trading partner have the correct sign and significance at 5% level of significance for three partners. For four other partners the coefficients are significant but have a positive sign. Again, these results provide support for a long run relationship between partner's income and own trade balance.

[3] Results from SBC-selected models showed a similar story with scarce support for the J-curve.

Table 5. Long-run Coefficient Estimates of Models Using AIC Lag Selection

Trading partners	Constant	$\log Y_{Pakistan}$	$\log Y_{Partner}$	$\log REX$
China	-12.01	0.84	0.87	2.86
	(1.19)	(1.96)	(1.53)	(5.10)
France	-12.07	1.46	-0.18	-0.12
	(2.10)	(3.42)	(0.78)	(0.99)
Germany	1.64	0.83	-0.71	0.00
	(0.52)	(2.60)	(3.40)	(0.01)
Hong Kong	28.38	-3.41	0.77	3.59
	(2.16)	(5.44)	(0.64)	(3.71)
Italy	-23.38	-0.06	1.92	0.01
	(0.43)	(0.03)	(0.50)	(0.07)
Japan	1.62	-1.38	0.79	1.71
	(0.20)	(5.11)	(1.47)	(2.50)
Korea	-80.00	0.63	4.62	6.86
	(1.46)	(0.67)	(1.38)	(1.52)
Kuwait	-19.87	-0.54	4.12	3.11
	(1.34)	(0.52)	(7.15)	(2.69)
Malaysia	-14.64	1.44	0.44	0.29
	(1.21)	(2.00)	(0.65)	(0.19)
Saudi Arabia	-9.58	-0.29	1.02	1.06
	(0.90)	(0.67)	(1.60)	(1.32)
U.A.E.	-7.28	-0.91	2.11	1.42
	(0.66)	(1.22)	(3.43)	(2.89)
U.K.	24.50	-0.15	-1.61	0.9
	(5.35)	(0.63)	(7.25)	(4.63)
U.S.A.	6.49	1.68	-1.5	0.31
	(0.42)	(3.24)	(1.89)	(0.88)

Note: Figures in parentheses are *t*-values.

Table 6. Long-run Coefficient Estimates of Models Using SBC Lag Selection

Trading partners	Constant	$\log Y_{Pakistan}$	$\log Y_{Partner}$	$\log REX$
China	3.02	0.50	-0.15	2.60
	(0.41)	(1.44)	(0.39)	(5.40)
France	-9.06	1.26	-0.26	-0.06
	(1.90)	(3.57)	(1.32)	(0.57)
Germany	3.63	0.53	-0.59	0.21
	(1.26)	(2.15)	(3.04)	(1.13)
Hong Kong	8.75	-3.07	2.52	4.77
	(0.68)	(4.22)	(2.13)	(4.57)

Italy	6.18	-0.41	-0.18	0.07
	(0.60)	(0.62)	(0.28)	(1.37)
Japan	-1.92	-1.18	0.91	1.66
	(0.15)	(3.08)	(1.08)	(1.59)
Korea	-79.46	0.25	4.71	7.49
	(1.37)	(0.26)	(1.32)	(1.50)
Kuwait	-34.99	0.42	4.6	2.36
	(2.82)	(0.48)	(9.92)	(2.31)
Malaysia	-3.78	0.92	0.14	1.43
	(0.37)	(1.49)	(0.23)	(1.06)
Saudi Arabia	3.77	-0.83	0.36	-0.59
	(0.31)	(1.73)	(0.45)	(0.58)
U.A.E.	-6.92	-0.91	2.1	1.5
	(0.72)	(1.40)	(3.91)	(3.45)
U.K.	25.22	-0.18	-1.65	0.9
	(6.85)	(0.89)	(9.27)	(5.72)
U.S.A.	11.03	1.58	-1.78	0.17
	(0.94)	(3.88)	(2.94)	(0.61)

Note: Figures in parentheses are *t*-values.

In the next stage of this analysis, estimates of δ_2, δ_3 and δ_4 normalized on δ_1 are used to calculate the linear combination of $\log TB_{i,t-1}$, $\log Y_{Pakistan,\, t\text{-}1}$, $\log Y_{i,t-1}$, and $\log REX_{i,t-1}$ as a new series, denoted by $ECM_{i,t-1}$. After replacing the linear combination of lagged level variables by ECM_{t-1} and after imposing the optimum lags, we estimate (2) one more time. A significantly negative coefficient obtained for ECM_{t-1} not only supports cointegration but also adjustment toward equilibrium. These results are also reported in Table 4. As can be seen, all $ECM_{i,t-1}$ coefficients are negative and significant with Italy being the only exception.[4]

Although $ECM_{i,t-1}$ coefficient estimates reported in Table 4 are highly significant and constitute enough evidence to support endogeneity of $\log TB_i$, it is worthwhile to test the possibility of any of the other three variables in (1) being endogenous.[5] Following Boyd *et al.* (2003) and Bahmani-Oskooee and Wang (2006), this is done by

[4] It should be mentioned that SBC-based models yielded the same results.

[5] In any error-correction model, one interpretation of a significant lagged error-correction term is that the right-hand-side variables in the long-run model cause the dependent variable, implying that the right-hand-side variables are exogenous where as the dependent variable is endogenous. For more on this see Jones and Joulfaian (1991, p. 150).

re-estimating (2) after interchanging $\Delta \log TB_i$, the dependent variable in (2) with the remaining three exogenous variables one by one. The ECM_{t-1} coefficient estimates from these additional models are reported in Table 7. These results show strong support for endogeneity of real exchange rate with all 13 coefficients being negative and significant at 5% level of significance. There is also mixed support for the endogeneity of real income of Pakistan and that of her trading partners. In short, the assumption of right hand side variables in (1) being completely exogenous seems untenable.[6]

Table 7. Coefficient Estimates of $ECM_{i,t-1}$ with Different Dependent Variables Under AIC

Trading partners	Dependent variable $\Delta \log TB$	Dependent variable $\Delta \log REX$	Dependent variable $\Delta \log Y_{Pakistan}$	Dependent variable $\Delta \log Y_{Partner}$
China	-1.11	-0.38	0.04	0.01
	(5.17)	(3.73)	(0.53)	(0.30)
France	-0.82	-0.12	-0.08	-0.07
	(7.56)	(2.73)	(0.72)	(2.40)
Germany	-0.67	-0.18	-0.24	-0.07
	(6.34)	(3.31)	(2.59)	(1.63)
Hong Kong	-0.71	-0.18	-0.14	-0.38
	(6.28)	(3.15)	(1.33)	(4.90)
Italy	-0.14	-0.11	0.06	-0.07
	(0.92)	(2.49)	(0.81)	(1.82)
Japan	-0.52	-0.25	-0.00	-0.40
	(5.96)	(3.18)	(0.01)	(6.40)
Korea	-0.17	-0.25	-0.13	-0.18
	(2.47)	(3.16)	(2.70)	(2.06)
Kuwait	-0.91	-0.32	-0.60	-0.18
	(4.96)	(4.57)	(4.03)	(1.55)
Malaysia	-0.53	-0.35	-0.22	-0.05
	(4.11)	(4.30)	(3.82)	(1.54)
Saudi Arabia	-0.56	-0.38	-0.02	-0.06
	(4.35)	(3.47)	(0.29)	(1.40)

[6] Note that Bahmani-Oskooee and Tanku (2006, p. 261) have argued that the issue is not as serious as it sounds, mostly because lagged values in the model could be considered as instruments for current values which amounts to treating each variable as endogenous.

U.A.E.	-0.59	-0.12	-0.29	-0.02
	(4.86)	(2.71)	(2.13)	(0.34)
U.K.	-0.80	-0.36	-0.15	-0.15
	(7.41)	(4.02)	(1.66)	(3.73)
U.S.A.	-0.61	-0.43	-0.15	-0.10
	(4.36)	(4.65)	(1.12)	(2.06)

Note: Figures in parentheses are t-values.

Table 8. Johansen's Maximum Likelihood Results with AIC-selected VAR Order

	□ $\lambda-\max$				*Trace*			
Null	$r = 0$	$r <= 1$	$r <= 2$	$r <= 3$	$r = 0$	$r <= 1$	$r <= 2$	$r <= 3$
Alternative	$r = 1$	$r = 2$	$r = 3$	$r = 4$	$r = 1$	$r = 2$	$r = 3$	$r = 4$
China	20.98	18.60	6.64	3.27	49.50	28.52	9.92	3.27
France	20.42	9.66	7.21	2.25	39.55	19.13	9.46	2.25
Germany	19.45	12.27	3.74	1.84	37.30	17.85	5.58	1.84
Hong Kong	35.62	14.60	9.11	4.09	63.42	27.80	13.20	4.09
Italy	27.09	16.59	6.11	1.93	51.71	24.62	8.03	1.93
Japan	18.75	11.78	7.51	6.29	44.34	25.58	13.80	6.29
Korea	14.15	8.71	4.68	3.25	30.80	16.64	7.93	3.25
Kuwait	21.29	15.81	8.14	3.59	48.83	27.54	11.73	3.59
Malaysia	16.19	11.44	6.28	3.06	36.96	20.77	9.34	3.06
Saudi Arabia	23.54	11.72	4.69	2.48	42.43	18.89	7.17	2.48
U.A.E.	22.04	19.59	10.63	4.52	56.79	34.74	15.15	4.52
U.K.	17.66	15.71	7.79	1.00	42.15	24.49	8.79	1.00
U.S.A.	32.87	17.69	11.77	3.19	65.53	32.66	14.96	3.19
90% critical value	25.80	19.86	13.81	7.53	49.95	31.93	17.88	7.53
95% critical value	28.27	22.04	15.87	9.16	53.48	34.87	20.18	9.16

Note: Number of co-integrating vectors is given by r.

In order to allow the feedback effects among the variables in (1), Johansen's co-integration approach is adopted. After confirming through various tests, including the ADF unit root test, the I(1) property of our variables, the next step in the Johansen's co-integration analysis is to calculate $\lambda-\max$ and *trace* statistics that will help us identify the number of co-integrating vectors.[7] Once again in selecting the optimum lags,

[7] China, Italy and U.A.E. have at least one I(0) series and thus the results for these countries should be interpreted with some caution.

we rely upon the AIC criterion. Note that following Cheung and Lai (1993, p. 317), the two statistics are adjusted for the number of observations *T*, number of lags *n*, and the number of variables in the co-integrating space *m*. The adjustment factor is *(T-nm)/T*. Thus, $\lambda-\max$ and *trace* statistics that are reported in Table 8 are the original figures multiplied by the adjustment factor.

Table 9. The Maximum Likelihood Estimate of Each Co-integrating Vector Under AIC

Trading Partners	*LTB*	$\log Y_{Pakistan}$	$\log Y_{Partner}$	$\log REX$	*Intercept*
China	-1.00	1.10	0.99	2.90	-15.92
	(2.16)	(0.72)	(0.48)	(0.87)	(0.41)
France	-1.00	0.86	-0.20	0.11	-5.38
	(13.45)	(2.07)	(0.83)	(0.55)	(0.59)
Germany	-1.00	0.29	-0.24	0.60	2.61
	(6.72)	(0.27)	(0.41)	(2.64)	(0.55)
Hong Kong	-1.00	-0.09	-10.71	-6.78	99.68
	(2.88)	(0.00)	(24.66)	(8.70)	(9.69)
Italy	-1.00	-0.93	-0.44	0.12	14.39
	(9.88)	(1.05)	(0.56)	(2.86)	(1.65)
Japan	-1.00	-0.46	0.90	1.28	-8.01
	(4.57)	(0.57)	(0.70)	(1.06)	(0.20)
Korea	-1.00	0.27	-43.28	-65.93	701.99
	(0.20)	(0.00)	(5.86)	(5.35)	(5.57)
Kuwait	-1.00	-1.71	3.35	3.73	-0.45
	(3.67)	(0.30)	(1.38)	(2.35)	(0.00)
Malaysia	-1.00	1.55	0.48	1.52	-12.61
	(5.69)	(4.10)	(0.47)	(0.44)	(0.44)
Saudi Arabia	-1.00	-1.21	-0.27	-0.58	13.97
	(16.66)	(4.10)	(0.16)	(0.11)	(0.99)
U.A.E.	-1.00	-0.19	5.71	3.91	-41.08
	(0.21)	(0.00)	(1.62)	(2.08)	(0.70)
U.K.	-1.00	1.30	-1.16	-0.33	0.78
	(0.20)	(1.24)	(0.07)	(0.03)	(0.00)
U.S.A.	-1.00	2.65	0.65	0.99	-31.65
	(6.65)	(8.10)	(0.14)	(1.68)	(1.37)

Notes: The critical values of χ^2 statistic are given in parentheses. At 1%, 5% and 10% levels of significance with 1 degree of freedom these critical values are 6.63, 3.84 and 2.71.

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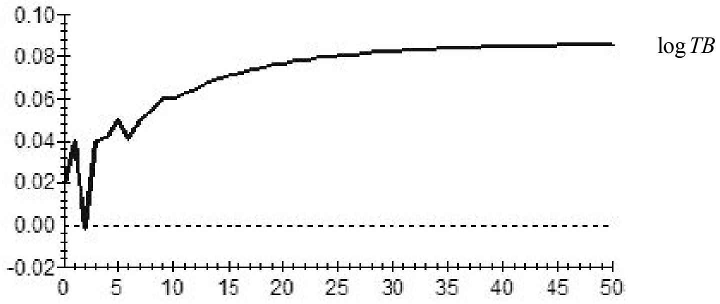
Generalized impulse response of $\log TB$ to one standard error shock in the equation for $\log REX$



Horizon

Figure 1-m. U.S.A

Figure 1. Generalized Impulse Response of $\log TB$ to One Standard Error Shock in the Equation for $\log REX$ Under AIC

Table 8 figures show that null of no co-integration is rejected at 10% level of significance only for Hong Kong, Italy, U.A.E. and U.S.A. using either $\lambda - \max$ or *trace* test. Also, using either test, there is no evidence for the existence of more than one vector except in case of U.A.E. and U.S.A. Based on these results it is assumed that for all countries at least one co-integrating vector exists. The results for the first vector are reproduced in Table 9 along with calculated value of the χ^2 statistic for likelihood ratio test of exclusion of corresponding variable in parentheses below the maximum likelihood estimates. Notice that the maximum likelihood estimates are normalized on the coefficient of $\log TB_i$ which is set equal to -1.

It can be seen from Table 9 that $\log REX_i$ has a significant coefficient for Hong Kong and Korea at 5% level of significance, with Italy added to the list at 10% level. However, only the coefficient for Italy carries the correct sign. With SBC lag selection criterion (results not reported here but available upon request) we can further expand the list of trading partners to include China, Hong Kong, Japan, U.A.E. and U.K with both significant and positive coefficients on $\log REX_i$, and Saudi Arabia with a significant but negative coefficient at 10% level of significance. These results are similar to those obtained from the bounds testing approach.

Following Halicioglu (2007), as a final step, the existence of J-curve is tested under this feedback scenario using Johansen's full information estimates for each trading partner by tracing out the generalized impulse response function of $\log TB_i$ to one standard error shock in the equation for $\log REX_i$. Since increase in $\log REX_i$ represents real exchange rate appreciation, an inverse J shape of the impulse response function would constitute evidence for the existence of the J-curve. Graphs of impulse response functions for all trading partners are presented in Figure 1. None of these plots clearly support the inverse J-curve.

4. SUMMARY AND CONCLUSION

In the last few decades numerous studies have sought evidence in support of the J-curve. The results however have not been conclusive. A recent trend in literature has been to employ disaggregated bilateral data between a country and her major trading partners in order to avoid the aggregation bias that can be present when aggregated data is used.

Previous research on Pakistan relied on aggregated data for J-curve related studies. This study went one step further and employed disaggregated bilateral data between Pakistan and her 13 largest trading partners in order to test the short-run (J-curve) and the long-run effects of real depreciation of Pakistani rupee. The two econometric techniques used for this purpose were the bounds testing approach and Johansen's co-integration approach. Two information criteria (Akaike Information Criterion and Schwarz Bayesian Criterion) for model selection were employed and a number of diagnostic tests were conducted in order to ensure the appropriateness of econometric results.

The bounds testing approach provided some evidence of short run effect of real exchange rate on trade balance. However these short run dynamics were inconsistent with the J-curve hypothesis. The long run results showed evidence of a positive and significant relationship between real exchange rate and trade balance in almost half of the trading partners in the sample using the bounds testing approach. The list included China, Hong Kong, Japan, Kuwait, and U.A.E. One policy implication of our findings is that not all trading partners are affected by real depreciation of Pakistani currency. The two largest trading partners, i.e., China and U.A.E. will be hurt by depreciation of Pakistani rupee. Johansen's co-integration approach did not provide much evidence in support of the J-curve nor any evidence of a significant long-run impact of the real exchange rate on bilateral trade balance.

APPENDIX. Data Definition and Sources

Quarterly data over 1980Q1-2003Q4 period was used for empirical analysis and came from World Development Indicators (World Bank, 2004 CD-ROM), International Financial Statistics (International Monetary Fund; online CD-ROM) and Direction of Trade Statistics (International Monetary Fund; 2004 CD-ROM).

Variables:

1. $TB_i =$ Pakistan's trade balance with her trading partner *i*. It is calculated as Pakistan's nominal imports from trading partner *i* divided by her nominal exports to the same trading partner.

2. $Y_i =$ measure of real GDP for country *i*. Where unavailable, quarterly GDP figures were generated following Bahmani-Oskooee (1986). $Y_{Pakistan}$ is Pakistani real GDP.

3. $REX_i =$ bilateral real exchange rate between Pak Rs. and the trading partner *i*'s currency. $REX_i = (P_{Pakistan} / P_i) \times NEX$, where $P_{Pakistan}$ and P_i are the price levels (CPI used as proxy) in Pakistan and in trading partner *i*, and NEX_i is the bilateral nominal exchange rate expressed as number of units of trading partner *i*'s currency per Pak Rs.

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Table 4. Coefficient Estimates for $\Delta \log REX_{t-i}$ Under AIC Lag Selection

Trading Partners	$\Delta \log REX_{t-0}$	$\Delta \log REX_{t-1}$	$\Delta \log REX_{t-2}$	$\Delta \log REX_{t-3}$	$\Delta \log REX_{t-4} \cdots$	$\Delta \log REX_{t-8}$	ECM_{t-1}
China	3.18						-1.11
	(4.23)						(5.17)
France	0.23						-0.82
	(1.14)						(7.56)
Germany	0.38						-0.67
	(1.62)						(6.34)
Hong Kong	0.66						-0.71
	(0.51)						(6.28)
Italy	-0.05	-0.01	0.10				-0.14
	(0.97)	(0.17)	(2.07)				(0.92)
Japan	-0.52						-0.52
	(1.05)						(5.96)
Korea	1.19						-0.17
	(2.11)						(2.47)
Kuwait	0.33	0.21	-4.33	-3.11			-0.91
	(0.21)	(0.14)	(2.94)	(2.01)			(4.96)
Malaysia	0.15						-0.53
	(0.19)						(4.11)
Saudi Arabia	-0.59						-0.56
	(1.23)						(4.35)
U.A.E.	1.32	-1.85					-0.59
	(1.58)	(2.47)					(4.86)
U.K.	0.21	-0.54					-0.80
	(0.95)	(2.56)					(7.41)
U.S.A.	0.19						-0.61

	(0.94)	(4.36)

Note: Figures in parentheses are *t*-values.

Generalized impulse response of $\log TB$ to one standard error shock in the equation for $\log REX$

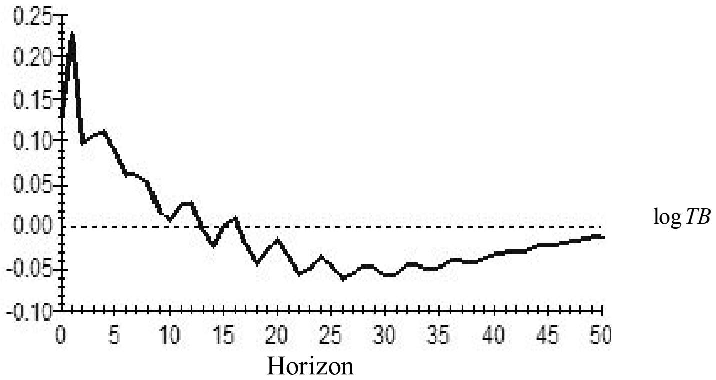


Figure 1-a. China

Generalized impulse response of $\log TB$ to one standard error shock in the equation for $\log REX$

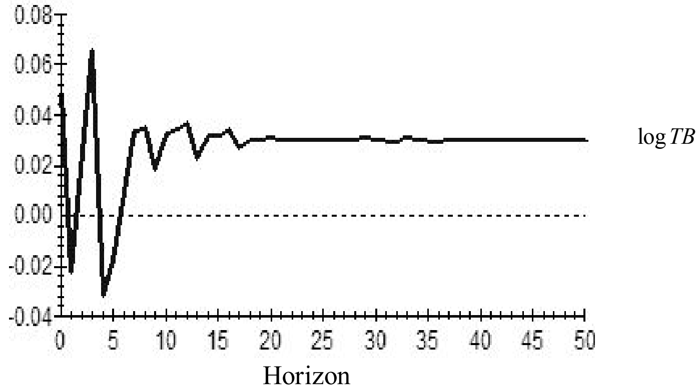
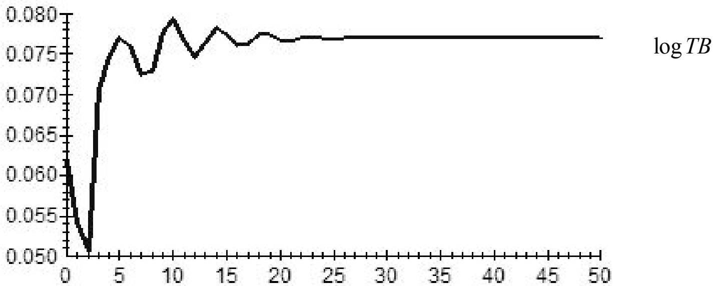


Figure 1-b. France

Generalized impulse response of $\log TB$ to one standard error shock in the equation for $\log REX$



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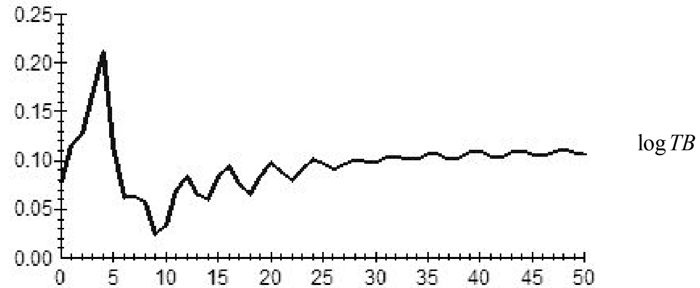
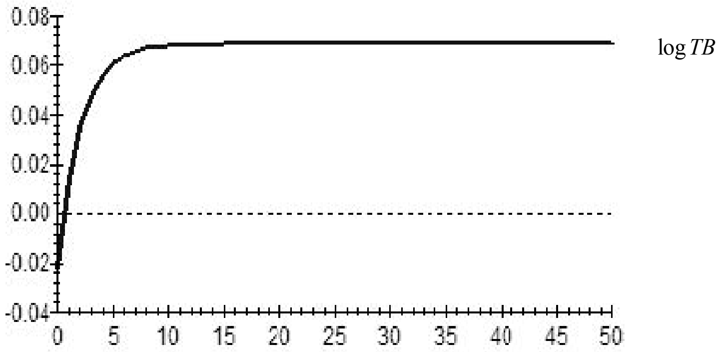


Horizon

Figure 1-c. Germany

Generalized impulse response of $\log TB$ to one standard error shock in the equation for $\log REX$



Horizon

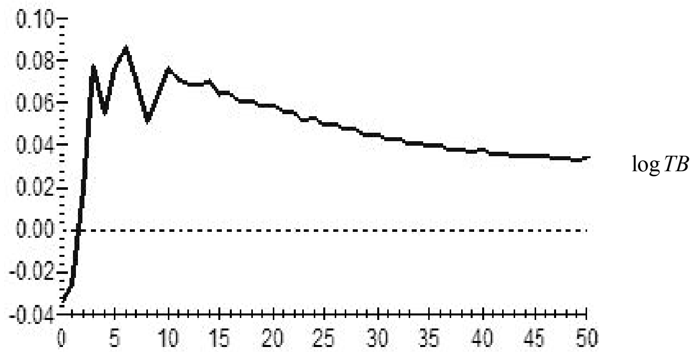
Figure 1-e. Italy

Generalized impulse response of $\log TB$ to one standard error shock in the equation for $\log REX$

Horizon

Figure 1-d. Hong Kong

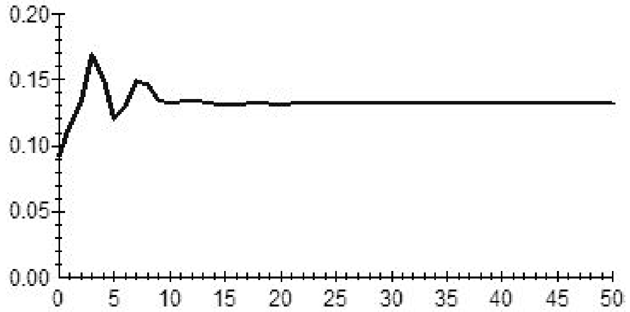
Generalized impulse response of $\log TB$ to one standard error shock in the equation for $\log REX$



Horizon

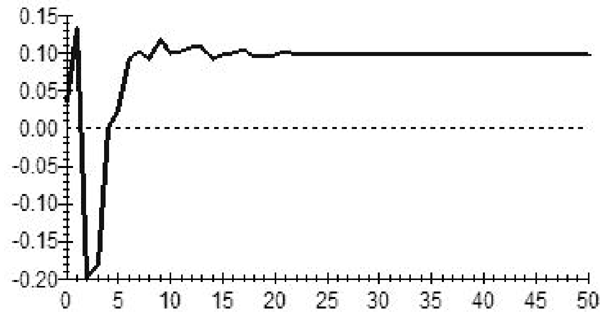
Figure 1-f. Japan

Generalized impulse response of $\log TB$ to one standard error shock in the equation for $\log REX$



Horizon

Figure 1-g. Korea

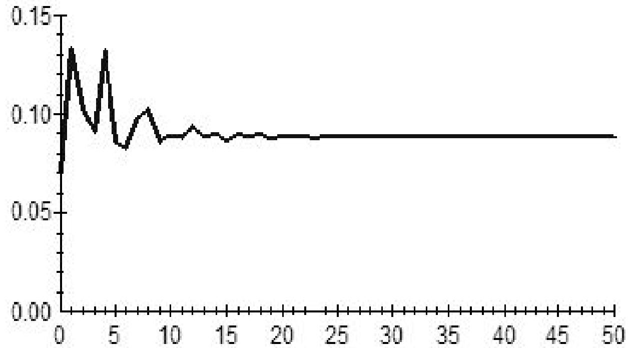


Horizon

Figure 1-h. Kuwait

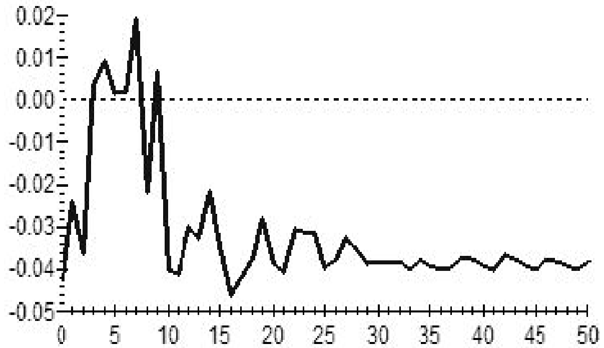
Generalized impulse response of $\log TB$ to one standard error shock in the equation for $\log REX$

Generalized impulse response of $\log TB$ to one standard error shock in the equation for $\log REX$



Horizon

Figure 1-i. Malaysia



Horizon

Figure 1-j. Saudi Arabia

Generalized impulse response of $\log TB$ to one standard error shock in the equation for $\log REX$

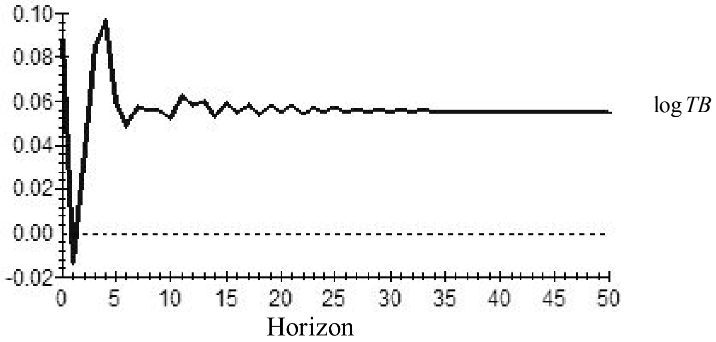


Figure 1-k. U.A.E

Generalized impulse response of $\log TB$ to one standard error shock in the equation for $\log REX$

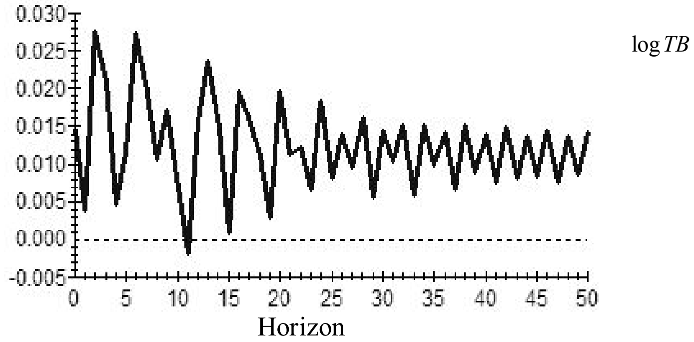


Figure 1-l. U.K.