The J-Curve Dynamics of Turkey: An Application of ARDL Model

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21 January 2008

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Abstract

This article seeks an empirical evidence for the existence of the J-curve phenomenon both in the short-run and long-run for Turkey over the period 1980-2005. The bounds testing cointegration approach is employed to estimate the trade balance model. An augmented form of Granger causality analysis is implemented between trade balance, real effective exchange rates, foreign income and domestic income. The stability of the short-run as well as long-run coefficients in the trade balance model is tested too. The empirical results suggest that the J-curve phenomenon is supported only in the short-run. Whilst causality tests reveal mix results, the parameter stability tests seem to be inconclusive.

Keywords: J-curve, trade balance, cointegration, causality, stability tests, Turkey
JEL: C22, F14, F31

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I. INTRODUCTION

A deficit in the trade balance of a country may be eliminated by a real devaluation in the domestic currency. Success of devaluation, however, depends on whether or not the sum of import and export elasticities exceeds unity, which is also known as the Marshall- Lerner (ML) condition. Bahmani-Oskooee (1985) argued that there have been cases under which the ML condition was satisfied yet the trade balance continued to deteriorate. Therefore, he concludes that the focus of a trade policy should be on the short-run dynamics that trace the post devaluation time path of the trade balance implying the J-curve phenomenon. The J-curve effect suggests that the trade balance of a nation initially worsens after a real depreciation of the home currency, and then it gets better. Krueger (1983) has argued that at the time an exchange occurs, goods, which are already in transit and under contract, have been purchased, and the completion of those transactions dominates the short-term change in the trade balance. Arndt and Dorrance (1987) indicates that this so called J-curve effect occurs if the domestic currency prices of exports are sticky. There has been growing interest in the J-curve phenomenon in the last three decades. Bahmani-Oskooee and Ratha (2004a) provides a very comprehensive survey on the J-curve literature for the period of 1973-2003 from thirty-seven articles. The recent examples of the J-curve studies include Bahmani-Oskooee *et al*. (2006), Bahmani-Oskooee *et al*. (2005), Bahmani-Oskooee and Ratha (2004b), Hacker and Hatemi (2004), Narayan (2004), and Narayan and Narayan (2004). The empirical literature on the evidence of the J-curve, however, is fairly ambiguous.

In regards to the empirical evidence for the Turkish J-curve, one can identify readily a few previous studies with inconclusive results. Rose (1990) finds that the real exchange rates have no impact on the trade balance. The work of Bahmani-Oskooee and Malixi (1992), which is based on Almon lag structure on real exchange rate, has found no support for it either. On employing the Engle-Granger cointegration approach, Bahmani-Oskooee and Alse (1994) asserts that the long-run impact of devaluation on the trade balance model is positive. Brada *et al*. (1997) who divided the data set into two sub-samples reports no long-run relationship between the variables of the trade balance function in the 1970s but they have revealed reverse results for the 1980s. Kale (2001) points out that a real depreciation of the domestic currency helps to improve the trade balance with a lag of about one-year and the impacts of devaluations on the trade balance are positive in the long-run. In a recent study, Akbostanci (2004) also presents empirical evidence of the J-curve phenomenon in the long-run. As a passing note, one should point out that the last two studies are based on that of Johansen-Juselius (1990) multivariate cointegration procedure with relatively short data spans derived from the 1980s and 1990s.

Turkey has adopted a policy of trade liberalization since 1980 and has pursued a successful export-led growth policy. The ratio of total exports to gross domestic product (GDP) increased from 4.1 to 13.3 percent during the period 1980-1988 and the real GDP grew by 5.8 percent in the same period. In 1989, there was a policy reversal, which slowed the depreciation of the Turkish lira (TL), in part to control inflation, but mainly to be able to easily borrow from the domestic markets, which led to five recessions in 1989, 1991, 1994, 1999, and 2001. The recessions of 1991, 1994 and 2001 were preceded by substantial increases in the real exchange rates. The impacts of devaluations on the trade balance, however, seem to be short lived as the early improvements in the trade balance are reversed steadily after a while. To combat the spiralling twin deficits in 2001, the IMF-led stabilization policy was put into effect once more. As a result, the internal imbalance has improved considerably but at the

same time the external balance has got worse. Despite having the free-floating exchange regime, the TL has been steadily appreciating against the major world currencies in real terms since 2002. This situation is being attributed to excess real domestic interest rates that are intentionally set at high levels to prevent inflation rising again. As a direct consequence of the overvalued TL, the current account deficit, which stems largely from the trade account deficit, has currently exceeded 6% of that Turkish GDP. Ertugrul and Selcuk (2001) provides a detailed account of the causes and consequences of the Turkish twin deficits in the 1980s and 1990s. A similar account of the Turkish economy beyond 2000 is given in Akyurek (2006).

The motivation of this study is two fold: persistent trade balance deficits of Turkey since the 1980s provide a good rationale to revisit the J-curve phenomenon and the cointegration procedure in this paper has been used for many countries in testing the J-curve phenomenon, except for Turkey.

The objectives of this study are as follows: i) to investigate the existence of the J-curve phenomenon both in the short-run and long-run using recent advances in time-series econometrics; ii) to establish the direction of causal relationships between trade balance, real effective exchange rates, foreign and domestic incomes variables; and iii) to implement parameter stability tests of Brown *et al*. (1975) to ascertain stability or instability in the trade balance model.

The remainder of this paper is organized as follows. Section II describes the study's model and methodology. Section III discusses the empirical results, and finally Section IV concludes.

II. THE MODEL AND METHODOLOGY

The trade balance model employed in this study adopted the form of Rose and Yellen (1989) and it takes the following long-run (cointegrating) form:

$$\ln TB_t = a_0 + a_1 \ln REER_t + a_2 \ln YW_t + a_3 \ln YT_t + \varepsilon_t \qquad (1)$$

where the measure of the trade balance, TB is the ratio of imports to exports; REER is the real effective exchange rate; YW is the industrial production index of industrial countries; YT is the industrial production index of Turkey. Ln is the natural logarithm transformation and ε_t is the random error term. According to the J-curve phenomenon, it is expected that $a_1 < 0$ since an increase in real effective exchange rate initially reduces the demand for the home country's export but increases its demand for imports. As a result, the balance of trade worsens initially but it will improve after a while as export and import volumes adjust to price changes. While there are no *apriori* expectations about the signs of a_2 and a_3, however, one asserts tentatively that a_2 is negative and a_3 is positive.

On investigating the cointegrating trade balance model with a view of testing the J-curve phenomenon, several econometric methods were implemented in the last two decades. In regards to univariate cointegration approaches, there are several examples including Engle-Granger (1987) and the fully modified OLS procedures of Phillips and Hansen's (1990). In terms of multivariate cointegration, Johansen (1988), Johansen and Juselius (1990), and Johansen's (1996) full information maximum likelihood procedures are widely employed. See Bahmani-Oskooee and Ratha (2004a) for a full account of the econometric procedures applied in the J-curve studies. A recent single cointegration approach, known as autoregressive-distributed lag (ARDL) of Pesaran

and Shin (1999) and Pesaran *et al*. (2001), has become popular amongst the researchers. Pesaran *et al.,* cointegration approach, also named as bounds testing, has certain econometric advantages in comparison to other single cointegration procedures. Firstly, endogeneity problems and inability to test hypotheses on the estimated coefficients in the long-run associated with the Engle-Granger method are avoided. Secondly, the long and short-run parameters of the model in question are estimated simultaneously. Thirdly, all variables are assumed to be endogenous. Fourthly, the econometric methodology is relieved of the burden of establishing the order of integration amongst the variables and of pre-testing for unit roots. The ARDL approach to testing for the existence of a long-run relationship between the variables in levels is applicable irrespective of whether the underlying regressors are purely *I*(0), purely *I*(1), or fractionally integrated. Finally, according to Narayan (2004), the small sample properties of the bounds testing approach are far superior to that of multivariate cointegration.
An ARDL representation of equation (1) is formulated as follows:

$$\begin{aligned}\Delta \ln TB_t &= a_0 + \sum_{i=1}^{m} a_{1i}\Delta \ln TB_{t-i} + \sum_{i=0}^{m} a_{2i}\Delta \ln REER_{t-i} + \sum_{i=0}^{m} a_{3i}\Delta \ln YW_{t-i} + \sum_{i=0}^{m} a_{4i}\Delta \ln YT_{t-i} \\ &+ a_5 \ln TB_{t-1} + a_6 \ln REER_{t-1} + a_7 \ln YW_{t-1} + a_8 \ln YT_{t-1} + v_t\end{aligned} \tag{2}$$

Pesaran *et al.* cointegration procedure is briefly outlined as follows. The bounds testing procedure is based on the F or Wald-statistics and is the first stage of the ARDL cointegration method. Accordingly, a joint significance test that implies no cointegration, (H_0: $a_5 = a_6 = a_7 = a_8 = 0$), should be performed for equation (2). The F test used for this procedure has a non-standard distribution. Thus, Pesaran *et al*. compute two sets of critical values for a given significance level. One set assumes that all variables are *I*(0) and the other set assumes they are all *I*(1). If the computed F-statistic exceeds the upper critical bounds value, then the H_0 is rejected. If the F-statistic falls into the bounds then the test becomes inconclusive. Lastly, if the F-statistic is below the lower critical bounds value, it implies no cointegration.
Once a long-run relationship has been established, equation (2) is estimated using an appropriate lag selection criterion. At the second stage of the ARDL cointegration procedure, it is also possible to perform a parameter stability test for the selected ARDL representation of the error correction model.
A general error correction model (ECM) of equation (2) is formulated as follows:

$$\begin{aligned}\Delta \ln TB_t &= b_0 + \sum_{i=1}^{k} b_{1i}\Delta \ln TB_{t-i} + \sum_{i=0}^{k} b_{2i}\Delta \ln REER_{t-i} + \sum_{i=0}^{k} b_{3i}\Delta \ln YW_{t-i} \\ &+ \sum_{i=0}^{k} b_{4i}\Delta \ln YT_{t-i} + \lambda EC_{t-1} + u_t\end{aligned} \tag{3}$$

where λ is the speed of adjustment parameter and EC is the residuals that are obtained from the estimated cointegration model of equation (1).
The Granger representation theorem suggests that there will be Granger causality in at least one direction if there exists a cointegration relationship among the variables in equation (1), providing that they are integrated order of one. Engle-Granger (1987) cautions that the Granger causality test, which is conducted in first difference via a vector autoregression (VAR), will be misleading in the presence of cointegration.

Therefore, an inclusion of an additional variable to the VAR system, such as the error correction term would help us to capture the long-run relationship. To this end, an augmented form of Granger causality test involving the error correction term is formulated in a multivariate *p*th order vector error correction model.

$$\begin{bmatrix} \Delta \ln TB_t \\ \Delta \ln REER_t \\ \Delta \ln YW_t \\ \Delta \ln YT_t \end{bmatrix} = \begin{bmatrix} c_1 \\ c_2 \\ c_3 \\ c_4 \end{bmatrix} + \sum_{i=1}^{p} \begin{bmatrix} d_{11i} d_{12i} d_{13i} d_{14i} \\ d_{21i} d_{22i} d_{23i} d_{24i} \\ d_{31i} d_{32i} d_{33i} d_{34i} \\ d_{41i} d_{42i} d_{43i} d_{44i} \end{bmatrix} \begin{bmatrix} \Delta \ln TB_{t-i} \\ \Delta \ln REER_{t-i} \\ \Delta \ln YW_{t-i} \\ \Delta \ln YT_{t-i} \end{bmatrix} + \begin{bmatrix} \lambda_1 \\ \lambda_2 \\ \lambda_3 \\ \lambda_4 \end{bmatrix} \left[EC_{t-1} \right] + \begin{bmatrix} \omega_{1t} \\ \omega_{2t} \\ \omega_{3t} \\ \omega_{4t} \end{bmatrix} \qquad (4)$$

EC_{t-1} is the error correction term, which is derived from the long-run relationship, and it is not included in equation (4) if one finds no cointegration amongst the variables in question. The Granger causality test may be applied to equation (4) as follows: i) by checking statistical significance of the lagged differences of the variables for each vector; this is a measure of short-run causality; and ii) by examining statistical significance of the error correction term for the vector that there exists a long-run relationship.

The existence of a cointegration derived from equation (2) does not necessarily imply that the estimated coefficients are stable as argued in Bahmani-Oskooee and Brooks (1999). Hence, stability tests of Brown *et al*. (1975), which are also known as cumulative sum (CUSUM) and cumulative sum of squares (CUSUMSQ) tests based on the recursive regression residuals, may be employed to that end. These tests also incorporate the short-run dynamics to the long-run through residuals. The CUSUM and CUSUMSQ statistics are updated recursively and plotted against the break points of the model. Providing that the plot of these statistics fall inside the critical bounds of 5% significance, one assumes that the coefficients of a given regression are stable. These tests are usually implemented by means of graphical representation.

III. THE EMPIRICAL RESULTS

Quarterly data over 1980I-2005IV period were used to estimate equation (2). Data definition and sources of data are cited in the appendix. Equation (2) was estimated in two stages. In the first stage of the ARDL procedure, the order of lags on the first–differenced variables for equation (2) was obtained from unrestricted VAR by means of Akaike Information Criterion (AIC) and Schwarz Bayesian Criterion (SBC), which indicated the optimal lag level as four and six quarters respectively for this study. The results of this stage are not reported here for brevity. Then an F deletion test was applied to equation (2) in order to test the existence of a long-run relationship by using lags from four to six on following Bahmani-Oskooee and Goswami (2003). As they have shown that the results of this stage are sensitive to the order of VAR. Equation (2) was also estimated three more times in the same way but the dependent variable each time was replaced by one of the explanatory variables in search of other possible long-run relationship in any other form than it had already been described in equation (1). The summary results of bounds tests are presented in Table 1. Table 1 indicates only one plausible a long-run relationship in which lnTB is dependent variable.

Table1. F-statistics for cointegration relationship

	Critical value bounds of the F-statistic			
	90% level		95% level	
	$I(0)$	$I(1)$	$I(0)$	$I(1)$
	2.721	3.773	3.232	4.351
Calculated F-statistics				
F(lnTB | lnREER, lnYW, lnYT) =5.175*				
F(lnREER | lnTB, lnYW, lnYT) =1.412				
F(lnYW | lnREER, lnTB, lnYT) =3.283				
F(lnYT | lnREER, lnYW, lnTB) =3.104				

The relevant critical values are obtained from Table C1.iii (with an unrestricted intercept and no trend with three regressors) in Pesaran *et al.* (2001).* indicates the statistical significance at the 5 % level. The optimal lag length is four.

In the second stage of ARDL cointegration procedure, equations (2) and (3) were estimated respectively on the basis of AIC and SBC criteria with six lags in search of selecting best-fitted model. The results of the model selection stages are reported in Table 2 and 3 respectively.

Table 2. Long-run coefficients of the trade balance of Turkey

Panel A: the long-run results Dependent variable $\ln TB$				
	Model Selection Criterion			
Regressors	AIC ARDL (1,2,3,0)		SBC ARDL (1,1,3,0)	
$\ln REER$	0.468	(1.449)	0.656*	(2.068)
$\ln YW$	-3.123*	(2.326)	-3.440*	(2.477)
$\ln YT$	1.328*	(2.371)	1.491*	(2.572)
Constant	7.387	(2.081)	7.186	(1.989)
Panel B: the short-run diagnostic test statistics				
$\chi^2_{SC}(4)$=2.324, $\chi^2_{FC}(1)$=5.731			$\chi^2_{SC}(4)$=2.144, $\chi^2_{FC}(1)$=4.948	
$\chi^2_{N}(2)$=8.176, $\chi^2_{H}(1)$=0.206			$\chi^2_{N}(2)$=7.192, $\chi^2_{H}(1)$=0.325	

Notes: The absolute value of t-ratios is in parentheses. χ^2_{SC}, χ^2_{FC}, χ^2_{N}, and χ^2_{H} are Lagrange multiplier statistics for tests of residual correlation, functional form mis-specification, non-normal errors and heteroskedasticity, respectively. These statistics are distributed as Chi-squared variates with degrees of freedom in parentheses. * indicates statistical significance at the 5% level.

Considering the long-run coefficients in the panel A of Table 2, one concludes that the SBC based trade balance model seems to be a slightly better fit than the AIC model since all the coefficients in the former model are statistically significant at the 5% level. Apart from that difference, both model selection criteria reveal almost identical results.

Table 3. Short-run coefficients of the trade balance of Turkey-estimates from ECM
Dependent variable $\Delta \ln TB_t$

	Model Selection Criterion			
Regressors	AIC ARDL (1,2,3,0)		SBC ARDL (1,1,3,0)	
$\Delta \ln REER_t$	0.907*	(3.759)	0.952*	(3.890)
$\Delta \ln REER_{t-1}$	0.494*	(2.044)	-	-
$\Delta \ln YW_t$	-2.323*	(4.011)	-2.225*	(3.787)
$\Delta \ln YW_{t-1}$	0.831	(1.276)	1.116	(1.723)
$\Delta \ln YW_{t-2}$	2.845*	(5.138)	2.939*	(5.233)
$\Delta \ln YT_t$	0.467*	(2.525)	0.522*	(2.804)
Constant	2.600	(2.251)	2.517	(2.143)
EC_{t-1}	-0.351*	(4.938)	-0.350*	(4.830)
$\overline{R}^2$	0.481		0.507	
F-statistics	14.168		15.287	
DW-statistics	2.034		2.041	
RSS	2.071		2.131	

Notes: The absolute values of t-ratios are in parentheses. RSS stands for residual sum of squares.
* indicates statistical significance at the 5% level.

As far as the ECMs are concerned, it can be seen from Table 3 that the results of equation (3) are also very close in both model selection criteria. The long-run and short run results of the SBC version of the trade balance model, however, provide again relatively better results in terms of the expected signs and econometric diagnostics. The speed of adjustment coefficient, -0.35 is considerably low indicating a slow convergence to equilibrium in the case of a shock to the cointegrating relationship and that long-run equilibrium is attained only after three quarters. Based on the AIC and SBC results, however, one can suggest that there are no J-curve effects both in the long and short-runs since the coefficient of the real effective exchange rates does not alter from positive to negative in either case. Considering that the focus of this paper is essentially on the dynamics of devaluation, the impact of the lags of the real exchange rates on the trade balance from error correction version of the SBC model was estimated. The results are reported in Table 4. The evidence of the J-curve is now apparent in the REER variable as the positive coefficients are followed by the negative coefficients. This pattern of the coefficients shows an initial deterioration followed by an improvement in the trade balance.

Table 4. Coefficient estimates of $\Delta \ln REER_{t-i}$ and error correction term

Dependent variable $\Delta \ln TB_t$

Regressors	Coefficients	t-statistics
$\Delta \ln REER_t$	0.890*	3.144
$\Delta \ln REER_{t-1}$	0.370	1.308
$\Delta \ln REER_{t-2}$	0.323	1.115
$\Delta \ln REER_{t-3}$	0.327	1.141
$\Delta \ln REER_{t-4}$	-0.094	0.319
$\Delta \ln REER_{t-5}$	-0.240	0.838
$\Delta \ln REER_{t-6}$	0.403	1.407
Constant	-0.018	0.961
EC_{t-1}	-0.327*	4.102
$\overline{R}^2$	0.421	
Diagnostic test statistics		
$\chi^2_{SC}(4)$=5.129 $\chi^2_{FC}(1)$=2.623, $\chi^2_N(2)$=4.143, $\chi^2_H(1)$=0.003		

Notes: t-ratios are in absolute values. χ^2_{SC}, χ^2_{FC}, χ^2_N, and χ^2_H are Lagrange multiplier statistics for tests of residual correlation, functional form mis-specification, non-normal errors and heteroskedasticity, respectively. These statistics are distributed as Chi-squared variates with degrees of freedom in parentheses. * indicates statistical significance at the 5% level.

Granger causality test requires that all variables are in equation (1) should be in integrated of order one. To this end, the traditional unit root tests such as the augmented Dickey and Fuller (1979, 1981) and the Phillips and Peron (1988) were employed. The unit root test results suggest that all variables in their first level differences are *I* (1). For brevity of presentation, they are not reported here. Having a cointegrating relationship among [TB_t, $REER_t$, YW_t, YT_t] on the basis of the results of the bounds test in Table 1, Granger causality test was conducted to equation (4) as such that only the trade balance vector was estimated with an error correction term. Similarly the Granger causality tests were applied to other models without the error correction terms since one could not ascertain any long-run relationship for the other vectors. Table 5 summarises the results of long-run and short-run Granger causality.

Table 5. Results of Granger Causality

	F statistics (probability)				
Dependent Variable	$\Delta \ln TB_t$	$\Delta \ln REER_t$	$\Delta \ln YW_t$	$\Delta \ln YT_t$	EC_{t-1} (t-statistics)
$\Delta \ln TB_t$	-	2.94* (0.01)	4.82* (0.01)	1.54 (0.18)	-0.26* (3.24)
$\Delta \ln REER_t$	3.32* (0.01)	-	1.42 (0.22)	1.46 (0.20)	-
$\Delta \ln YW_t$	0.79 (0.55)	0.65 (0.66)	-	2.18 (0.06)	-
$\Delta \ln YT_t$	1.64 (0.15)	3.40* (0.00)	2.97* (0.01)	-	-
Causality inference: REER⇔TB, YW⇒TB, YW⇒YT, REER⇒YT					

Notes:* indicates statistical significance at the 5% level. The optimal length is 4 and is based on SBC.

On considering the long-run Granger causality, one suggests that the lagged error correction term with the expected sign is statistically significant which also confirms the results of the bounds test. The speed of the adjustment, however, is less than the ARLD estimations suggesting that the long-term convergence to equilibrium will take longer. Granger causality runs interactively via both exchange rates and world income to trade balance ratio. In regards to the short-run Granger causality tests based on the F statistics, one sees that there exists a bi-directional Granger causality between exchange rates and trade balance ratio, implying that as the trade balance improves it causes a real appreciation of the TL. Likewise, a real depreciation will have a positive impact on the trade balance. There are, however, uni-directional Granger causalities, which run from world income to the trade balance, from world income to domestic income, and from exchange rates to domestic income.

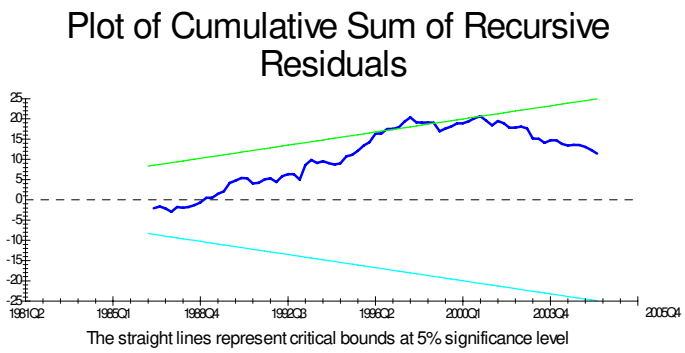

Figure 1. Plot of CUSUM

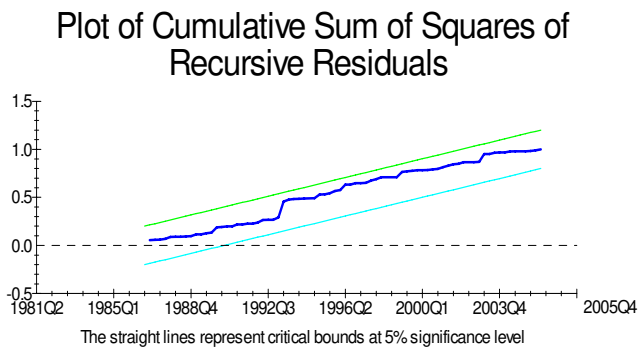

Figure 2. Plot of CUSUMSQ

The parameter stability tests were implemented via equation (3), which captures the short-run dynamics of equation (2) and the long-run impact of equation (1). Equation (3) was estimated by ordinary least squares with a lag length of four based on the SBC criterion. Figure 1 and 2 provides the plots of CUSUM and CUSUMSQ tests, respectively. Figure 1 and 2 display conflicting results in regards to the parameter stability tests. Whilst the former test indicates instability the later suggests stability in the parameters of the trade balance equation.

IV. CONCLUSION

This study has attempted to estimate the J-curve phenomenon through a reduced form trade balance model in search of providing fresh empirical evidence in the case of Turkish data. A recent single cointegration technique proposed by Pesaran *et al*. (2001) was employed to investigate the short-run and long-run responses of the trade balance to currency depreciation. The results from bounds tests have indicated only one possible long-run relationship in which the trade balance is the dependent variable. In contrast to the results of the previous results relating to Turkey, one finds no long-run impact of a real devaluation on the trade balance but there exists the J-curve phenomenon in the short-run. The existence of J-curve in the short-run confirms that the Turkish foreign trade balance deficit improves rather rapidly on following currency depreciations as it has happened in the TL depreciations of 1994, 1999 and 2004.Whilst, in the long-run, real effective exchange rates and world income Granger cause to the trade balance, in the short-run there is a feedback relationship between the real effective exchange rates and trade balance. The parameter stability tests on the long-run trade balance equation appear to be inconclusive.

APPENDIX

Data definition and sources

The data set used in this study cover the period 1980I to 2005IV. All data are collected from International Financial Statistics, (IMF) and Central Bank of Turkey (CBT).

TB is Turkey's trade balance. It is defined as the ratio of real imports to real exports. The raw data of exports and imports were deflated by the USA consumer price index. Source: IMF.
REER is the real effective exchange rates index. Source: CBT.
YW is industrial production index of industrial countries. Source: IMF.
YT is industrial production index of Turkey. Source: IMF

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