# An empirical model for the Turkish trade balance: new evidence from ARDL bounds testing analyses

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# AN EMPIRICAL MODEL FOR THE TURKISH TRADE BALANCE: NEW EVIDENCE FROM ARDL BOUNDS TESTING ANALYSES

# TÜRK TİCARET DENGESİ İÇİN UYGULAMALI BİR MODEL: ARDL SINIRLAR TESTİ ÇÖZÜMLEMELERİNDEN YENİ BULGULAR

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## ABSTRACT

In this paper, the determinants of the Turkish trade balance are tried to be analyzed in an empirical modelling approach. For this purpose, the contemporaneous ARDL-based bounds testing has been used to examine the existence of a long run co-integration relationship between the variables of our interest. The estimation results indicate that real exchange rate depreciations improves the trade balance in a strong and significant way, that domestic real income affects the trade balance negatively, and that trade balance is strongly improved due to an increase in foreign real income. No significant effect of crude oil prices can be observed on trade balance. The error correction modeling gives results in line with the long run findings of the co-integration analysis.

*Key Words*: Trade Balance; ARDL Bounds Testing Approach; Turkish Economy;

*JEL Classification*: C32 ; F10 ; F41 ;

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**ÖZET**

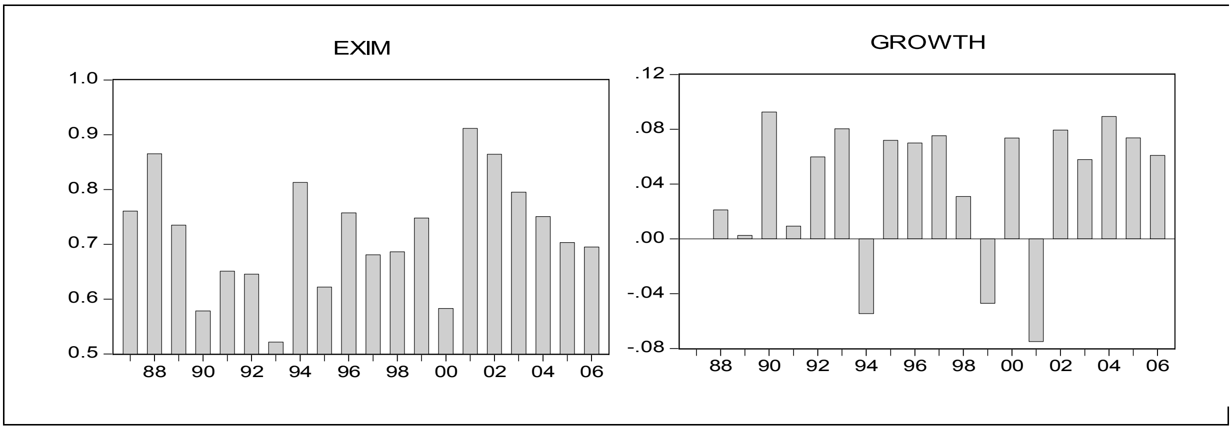
Bu çalışmada Türk ticaret dengesinin belirleyicileri uygulamalı bir modelleme yaklaşımı içerisinde çözümlenmeye çalışılmaktadır. Bu amaçla çağdaş ARDL temelli sınırlar testi ilgi alanımıza giren değişkenler arasındaki uzun dönemli eş-bütünleşik bir ilişkinin varlığının incelenmesi için kullanılmıştır. Tahmin sonuçları reel döviz kuru değer kayıplarının ticaret dengesini güçlü ve anlamlı bir şekilde iyileştirdiğini, yurtiçi reel gelirin ticaret dengesini negatif olarak etkilediğini ve ticaret dengesinin yabancı reel gelirdeki bir artış sonucu güçlü bir şekilde iyileştiğini göstermektedir. Ham petrol fiyatlarının ticaret dengesi üzerinde anlamlı bir etkisi gözlenememektedir. Hata düzeltme modellemesi eş-bütünleşim çözümlemesinin uzun dönem bulguları doğrultusunda sonuçlar vermektedir.

*Anahtar Kelimeler*: Ticaret Dengesi; ARDL Sınırlar Testi Yaklaşımı; Türkiye Ekonomisi;
*JEL Sınıflaması*: C32 ; F10 ; F41 ;

## 1. INTRODUCTION

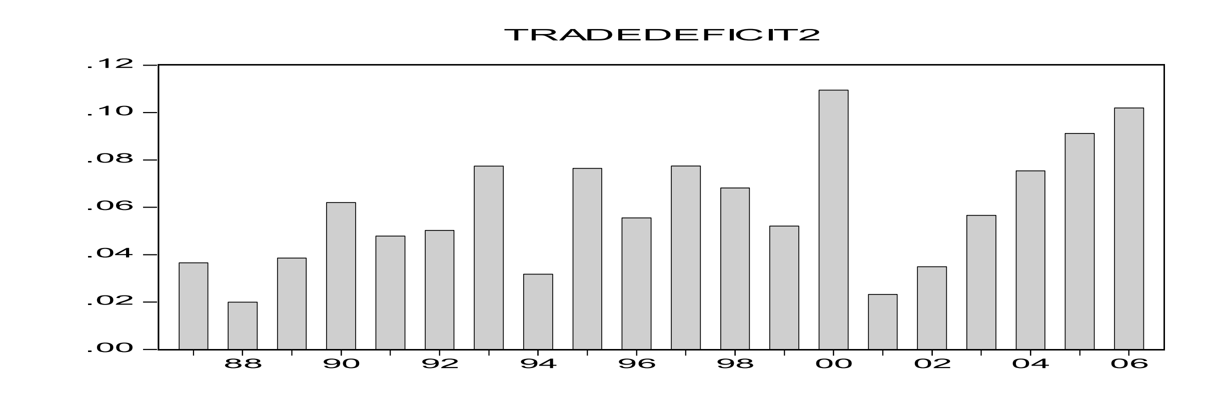
By the post-1989 capital account liberalization period, the Turkish economy can mainly be characterized with a highly volatile real domestic income growth process which seems to be in a close relation with the courses of both real exchange rate and trade balance. An ever-increasing trade balance deficit except the two substantial economic crisis periods in 1994 and 2001 coincides also with the increasing openness ratio of external trade volume to gross domestic product (GDP), and such developments which resulted in trade imbalances cast some doubts as to whether improvements in trade balance must be attributed to real macroeconomic income growth process and whether devaluations of real exchange rate are expansionary. These all in turn, to the great extent, shed some light upon the trade balance-based business cycle properties of the Turkish economy. These stylized facts are given below in a cursory way for the 1987 – 2006 period of annual observations.

**Figure 1: Trade Balance and Real Economic Growth**



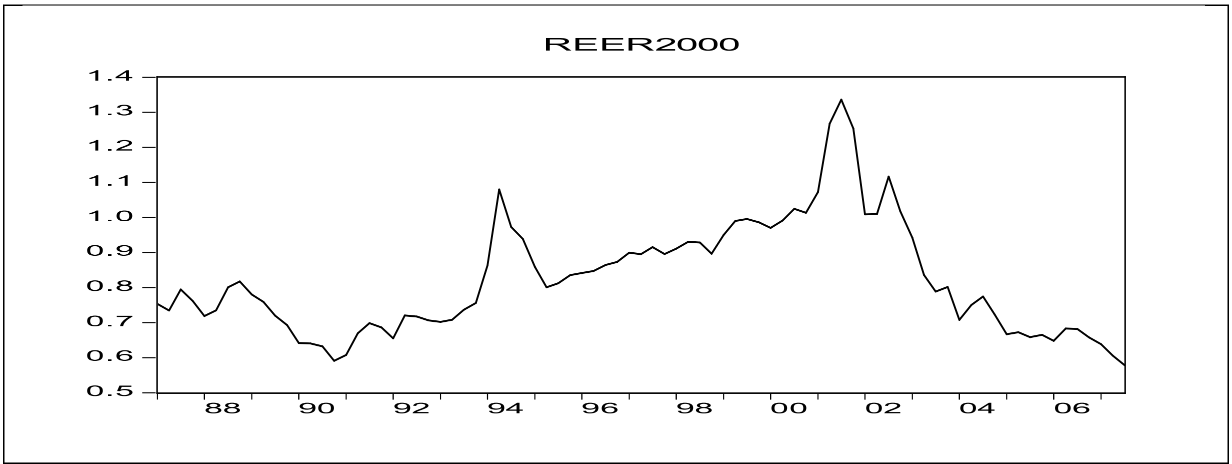

In Figure 1, the ratio of aggregate exports to imports in millions of US$s (EXIM) and the real GDP growth rates (GROWTH) using annual observations have been compared. All the data are taken from the electronic data delivery system of the Central Bank of the Republic of Turkey (CBRT). We can easily notice that, as of the early-1990s, the trade deficit ratio which is represented by the ratio of exports to imports in million U.S. dollar terms decreases steadily and takes critical values below 0.52 just before the 1994 economic crisis. As a result of the 1994 crisis conditions leading to both an enormous depreciation in domestic real income and decreasing imports and increasing exports volumes through the real depreciations in domestic currency, this ratio has a value larger than 0.81 in 1994 but begins to decline by the subsequent periods again to the values between 0.60 – 0.75. The year 2000 witnesses that this ratio comes back to the margin of 0.58 such as just before the burst of the 1994 crisis, but following the crisis conditions it increases significantly above the threshold value 0.91. However, trade balance perpetuates to be deteriorating for the post-2002 period.

**Figure 2: Trade Deficit Ratio**




In Figure 2, what is of special interest supporting the above explanations is that the larger the depreciation of trade balance the larger would be the real income growth rates, whereas real income depreciation periods such as years 1994 and 2001 do not indicate huge depreciation of trade balance as opposed to the earlier periods nor do they coincide with the increasing appreciation of domestic currency when compared with the former periods as can be seen in Figure 3 below. These are highly explicit especially for the 2002 - 2006 period in that there exists an upward trend in the trade deficit ratio but these years have an about in average of 7.2% real income growth rate with a steadily appreciating real exchange rate.

**Figure 3: U.S. Dollar Based Real Exchange Rate**



Following Berument and Dinçer (2005), the real exchange (REER2000) data in Figure 3 is calculated as US$ / Turkish lira times the US consumer price index, which uses all items, devided by the producer price index of Turkey, which is based on manufacturing products. Then, this final series is divided to the average of the real exchange rate estimate for the year 2000. The time series data are also used in the empirical model of the paper in the later sections. Both price indices used for estimating real exchange rate have the base 2000: 100 and are taken from the OECD online statistical data base, http://www.oecd.org. An increase in the real exchange rate such calculated means a depreciation, while a decrease means appreciation. It is highly explicit that the 1994 and 2001 periods witness upward jumps in real exchange rate, that is, depreciation of domestic currency against U.S. dollars.

Note that the exports, imports and real GDP data are obtained from the electronic data delivery system of the Central Bank of the Republic of Turkey (CBRT), http://www.tcmb.gov.tr. Trade deficits in Figure 2 are calculated as [(exports – imports) / GNP], where GNP data represent gross domestic product at market prices in million U.S. dollars and are obtained from the electronic data delivery system of the Turkish Republic Prime Ministry State Planning Organization, http://www.dpt.gov.tr/.

Having documented some general stylized facts of the Turkish economy, in this paper, the determinants of the Turkish trade balance have been tried to be analyzed by testing any possible long run equilibrium relationship as well as short run dynamics. For this purpose, the next section gives a well-constructed reduced form model in the international trade balance literature. A survey of empirical papers for the Turkish economy follows this theoretical section. Data and time series characteristics are presented in the fourth section, while a methodological discussion for the estimation process is carried out in the fifth section. The sixth section of the paper represents our main contribution to the existing literature and aim at giving an empirical essay for the Turkish trade balance. The last section concludes the paper.

## 2. A REDUCED FORM MODEL

Most of the studies on trade balance for the post-1990 period are based on the imperfect substitutes models of Goldstein and Khan (1985) and Rose and Yellen (1989) where the reduced form of trade balance is developed. The key underlying assumption of the imperfect substitutes model is that neither imports nor exports are perfect substitutes for domestic goods. They reveal that main underlying reasons leading to imperfect substitutes model can be considered in the sense that, if domestic and foreign goods were perfect substitutes, then one should observe (i) either the domestic or foreign good swallowing up the whole market when each is produced under constant or decreasing costs, (ii) each country as an exporter or importer of a traded good but not both. Since both of these predictions are counter to fact at the aggregate and disaggregated level, i.e. one normally observes the co-existence of imports and domestic output and the flourishing of two-way trade, the perfect substitutes hypothesis can be rejected. Following the informative modeling approach of Stučka (2004) with a standard two-country imperfect substitutes model, let imports, exports and trade balance refer to the merchandise component, and neither imports nor exports are perfect substitutes for domestic goods, so that finite elasticities for demand and supply can be estimated for most traded goods. The volume of imports demanded domestically, $M_d$, and the quantity of imports by the rest of the world, $M_d^*$, are given below:

$$M_d = f_1(Y, P_m, P),\ (\delta M_d / \delta Y) > 0\ ,\ (\delta M_d / \delta P_m) < 0,\ (\delta M_d / \delta P) > 0 \tag{1}$$

$$M_d^* = f_2(Y^* e, P_m^*, P^*),\ (\delta M_d^* / \delta Y^* e) > 0\ ,\ (\delta M_d^* / \delta P_m^*) < 0,\ (\delta M_d^* / \delta P^*) > 0 \tag{2}$$

where $Y$ is domestic income, $P_m$ the domestic currency price paid by domestic importers, $P$ the overall domestic price level, $Y^*$ the foreign income, $e$ the exchange rate as the domestic currency price of foreign exchange, $P_m^*$ the foreign currency price paid by domestic importers and $P^*$ the overall foreign price level. In this functional form of external balance, the quantity demanded is a function of the level of money income in the importing region, the imported goods' own price and the price of domestic substitutes, where domestic income and foreign income elasticities as well as cross price elasticities of demand are assumed positive, while own-price elasticities of demand are assumed to be negative. Homogeneity of demand function is accepted, so that the consumer would not suffer from money illusion. For instance, demand would remain constant when doubling money income and prices. This homogeneity assumption is expressed by dividing the explanatory variables on the right hand side by $P$, allowing to use real income and relative prices of imports to domestically produced goods:

$$M_d = f_1(Y_r, RP_m),\ (\delta M_d / \delta Y_r) > 0\ ,\ (\delta M_d / \delta RP_m) < 0,\ Y_r = (Y / P),\ RP_m = (P_m / P) \tag{3}$$

$$M_d^* = f_2(Y_r^*, RP_m^*),\ (\delta M_d / \delta Y_r) > 0\ ,\ (\delta M_d / \delta RP_m^*) < 0,\ Y_r^* = (Y^* / P^*),\ RP_m^* = (P_m^* / P^*) \tag{4}$$

Since the relative price of imports is equivalent to the foreign currency price of foreign exports adjusted for exchange rate, relative prices of imports can be defined as follows:

$$RP_m = (P_m / P) = (eP_x^* / P) = (eP^* / P)(P_x^* / P^*) = Q(P_x^* / P^*) = Qp_x^* \tag{5}$$

where $p_x^*$ represents the real foreign currency price of exports, while $Q$ denotes the real exchange rate and an increase in $Q$ refers to a domestic currency depreciation:

$$Q = eP^* / P \tag{6}$$

The quantity of imports supplied by the rest of the world to the domestic country and the quantity of exports domestically supplied to the rest of the world are given below.

$$X_s = f_3(P_x, P) \text{ and } X_s^* = f_4(P_x^*, P^*) \tag{7}$$

where $P_x$ is the domestic currency price received by domestic exporters. In equilibrium conditions:

$$M_d = X_s^* e \text{ and } M_d^* = X_s \tag{8}$$

By defining the trade balance as, $TB = p_x M_d^* - Qp_x^* M_d$ , and solving for the levels of domestic exports and imports as well as the relative price of imports as a function of real exchange rate, we obtain in Eq. (9) the partial reduced form of the domestic trade balance that we use for empirical purposes:

$$TB = f(Y_r, Y_r^*, Q),\ (\delta TB / \delta Y_r) < 0\ ,\ (\delta TB / \delta Y_r^*) > 0\ ,\ (\delta TB / \delta Q) > 0 \tag{9}$$

Therefore, real foreign income and real exchange rate are expected to be positively related and domestic income is assumed negatively related to the course of the trade balance.

## 3. A BRIEF LITERATURE SURVEY UPON THE TURKISH ECONOMY

Using the Turkish data, Rose (1990) finds out no impact of real exchange rate on trade balance for the 1970 – 1988 period. Domaç (1993) investigates the validity of the *so-called* J-curve effect, which simply requires initially worsening and then gradually improvement of trade balance following a devaluation, for the Turkish economy by imposing Almon lag structure on exchange rate. The results indicate that long run devaluation does not improve the trade balance of Turkey and that in the short-run trade balance initially deteriorates and then starts to be improved. Brada et al. (1997) using Engel-Granger procedure of the co-integration methodology and polynomial curve analysis examine the balance of trade data for the pre-and-post 1980 based on the changes in the trade policies inside the period. Their findings indicate the rejection of the existence of any functional relationship between exchange rate and trade balance for the pre-1980 period, while trade balance is found responsive to changes in exchange rate for the post-1980 period, suggesting that exchange rate policy was able to create and maintain a satisfactory balance of trade position in the 1980s and early-1990s. Kale (2001) examines the relationship between the balance of trade and real exchange rate using co-integration analysis. She finds that a real depreciation would improve the Turkish trade balance in the long run. Akbostancı (2004) using co-integration / vector error correction modeling and dynamic generalized impulse response analysis finds that however a real depreciation of domestic currency would improve the Turkish trade balance in the long run in a way supporting the findings in Bahmani-Oskooee (2001), results do not support the short run worsening of trade balance. Short run dynamic behavior of trade balance in response to real exchange rate shocks indicate an S-pattern rather than a J-curve pattern, that is, trade balance would be initially improved, then worsened and then improved in response to real exchange rate shocks. Berument and Dinçer (2005) consider the currency denomination of exports and imports when analysing the Turkish trade balance given that exports are mostly denominated in Euros and imports are mostly denominated in US$s. By including US$ / Euro into the analysis of trade balance, they find that parity effects in favor of appreciation of Euro against US$ would increase output and appreciate the local currency while improving the trade balance. Finally, Zortuk and Durman (2008) investigate the long run relationship between trade balance and terms of trade in Turkey. Their estimation results reveal a long run relationship between trade balance and income terms of trade. However, the authors cannot observe a long run relationship between trade balance and commodity terms of trade.

## 4. DATA AND TIME SERIES CHARACTERISTICS

In this paper, a similar model specification to Brada et al. (1997) and Akbostancı (2004) is used for modeling purposes. The quarterly frequency data are considered for the post-1990 data realizations and the period used for estimation purposes cover the time span of 1990:Q1-2007:Q3. We must specify that non-inclusion of the years 1987, 1988 and 1989 is due to the fact that adding these observations leads our overall estimation results to giving econometrically inconsistent results dealing with a possible stationary combination of the variables that we consider in this paper. These detailed findings not reported here to save space will of course be presented to the readers and researchers if requested from the authors. For trade balance, the ratio of exports to imports in natural logarithms (*TBAL*) is used. Bahmani-Oskooee (1991; 2001) provide two justification for such a variable specification and indicate that this ratio is not sensitive to units of measurement and that it could be interpreted as nominal or real trade balance. We tend to use as the real exchange rate variable (*RE*) time series the same data given in Figure 3 above. These data are constructed as in Eq. (10).

$$\frac{CPI^{US} * \$^{US}}{PPI^{TURKEY}} \qquad (10)$$

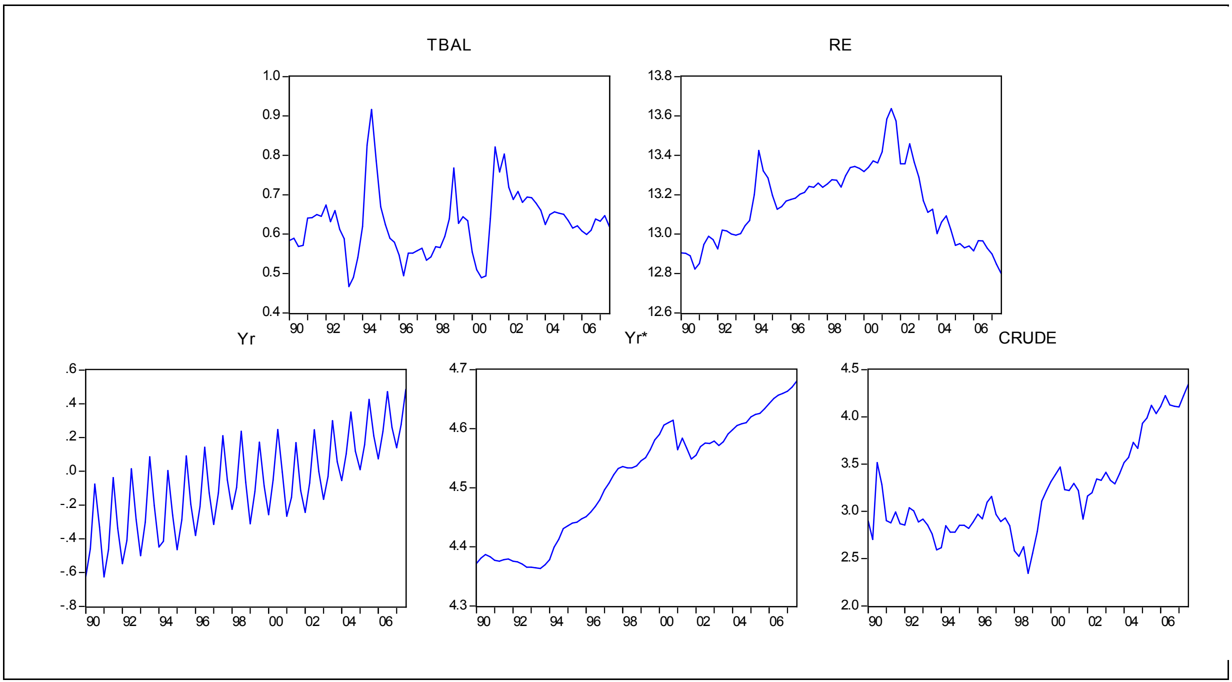
$CPI^{US}$ and $PPP^{TURKEY}$ are the US consumer price index, which uses all items, and Turkish producer price index, which is based on manufacturing products, respectively. Both price indices have the base 2000: 100. Such calculated data, then, are divided to the four-quarter average of the real exchange rate estimate for the year 2000.[1] An increase in real exchange rate means a depreciation, while a decrease means appreciation. The real GDP data ($Y_r$) to proxy domestic real income level use the 1987: 100 base year. In a similar way to the final real exchange rate series calculation, the real income data are divided to the four-quarter average of the real GDP for the year 2000. The foreign real income level ($Y_r^*$) is represented by the G-7 countries industrial production index data with the base 2000: 100. These variable specifications serve us to empirically use Eq. (9). Additionally, we have added into our model construction the crude oil prices (*CRUDE*) to account for any other possible effects on trade balance resulted from the developments in the world markets. The data for constructing trade balance and domestic income are obtained from the electronic data delivery system of the

[1] The author would like to thank Gökhan Karabulut of the Istanbul University Department of Economics for informative explanations on this issue.

CBRT, http://www.tcmb.gov.tr. The nominal exchange data used in real exchange rate series are also compiled from the same source. The price indices data used for estimating real exchange rate are taken from the OECD online statistical data base, http://www.oecd.org. The data for $Y_r^*$ and CRUDE variables come from the International Monetary Fund (IMF) International Financial Statistics bulletins.

All the data used indicate seasonally unadjusted values and are in their natural logarithms which enable us to explain them in a constant elasticity form, except the trade balance variable for which no logarithmic transformation has been applied. Following the suggestions of an anonymous referee, to take account of seasonality in the estimation process we include a set of seasonal dummies into the model evaluation process. The time series representation of the variables can be seen in Figure 4.

**Figure 4: Time Series Used for Modeling Purposes**




Using a functional representation, the trade balance equation of our interest in this paper can be indicated as in Eq. 11 below.

$$TBAL = f(Y_r, Y_r^*, RE, CRUDE) \tag{11}$$

Instead of such a reduced form of trade balance, for the case of the Turkish economy, Şahinbeyoğlu and Ulaşan (1999) interest in real export function and Kotan and Saygılı (1999) estimate a nominal import demand function. Aydın et al. (2004) estimate both a real export supply and an import demand function using co-integration analysis and a dynamic VAR model of impulse responses of trade deficit. Similarly, Yavuz and Güriş (2006) give a well-organized paper on aggregate import demand function for the Turkish economy using a similar estimation methodology followed in this paper.

For *a priori* signs of the variables, real income elasticity of trade balance is expected to be negative since increasing domestic real income would stimulate more imports through increasing domestic absorption which initially deteriorates the trade balance. But as Domaç (1993) states, when real income increases, the production of import substitute goods may reduce the volume of imports, and in this case the sign of real income would be positive instead.

The sign of foreign output with respect to trade balance is expected to be positive. An increase in world income would stimulate the demand for home country goods. But also, an increase in world real income may mean increasing real income and degree of absorption for domestic residents, and in this case the net effect of an increase in world real income on domestic trade balance would be uncertain.

For the real exchange rate variable, we can assume that a real depreciation, which means an increase in the real exchange rate series used in this paper, should improve the trade balance in a long run perspective through price effects, and this results in a negative relationship between trade balance and real exchange rate. Because, a decrease in real exchange rate means that domestic goods would have been cheapened to the foreigners in real terms and this should stimulate exports. However, since foreign goods to the domestic residents are now more expensive in real terms, imports of domestic economic agents should be discouraged.

Finally, we tend to appreciate the effect of crude oil prices on trade balance through our estimation findings.

Having defined the data, we appreciate below the time series stationarity properties of the variables. Since these methodologies are well-known in today's economics literature, no discussion has been given for the econometrical details of these tests. The related Turkish readers are suggested to apply to Nemlioğlu (2005) and Göktaş (2005) to be highlighted for some excellent knowledge upon these issues of interest. Briefly to say, the spurious regression problem analyzed by Granger and Newbold (1974) indicates that using non-stationary time series steadily diverging from long run mean leads to unreliable correlations within the regression analysis leading to unbounded variance process. However, for the mean, variance, and covariance of a time series to be constant over time, conditional probability distributions of the series must be invariant with respect to the time. Such a case means that the variables of the model must be differenced (*d*) times to obtain a covariance-stationary process. Dickey and Fuller (1979; 1981) suggest the use of one of the commonly applied test methods known as augmented Dickey-Fuller (ADF) test to detect whether the time series is of stationary form. However, Dickey-Fuller type tests may have low estimation power against the plausible stationary alternative hypothesis and the null hypothesis of a unit root may tend to be accepted unless there is strong evidence against it. Considering these facts, Kwiatkowski et al. (1992) develop an alternative approach known as the KPSS tests which are designed to test the null hypothesis of stationarity against the unit root alternative. The related reader can find a brief comparative analysis of these tests in a highlighting paper yielded by Yavuz (2004). In our study, we apply to the ADF ve KPSS tests for the times series considered. The estimation results are reported in Table 1. $\tau_c$ and $\tau_t$ are the test statistics with allowance for only constant and constant&trend terms in the unit root tests, respectively. The numbers in parentheses are the lags used for the ADF test, which are augmented up to a maximum of 10 lags, and bandwiths for the KPSS tests. The choice of optimum lag for the ADF test was decided on the basis of minimizing the Schwarz information criterion. '*' denotes that the variable is of stationary form.

The unit root results yield some contradictory estimates for variable time series. We can say definitely that the stationary characteristic of the trade balance variable and the unit root charactestic of the crude oil prices in the level form cannot be rejected by the data. However, we can infer that the real domestic and foreign income, and real exchange rate data are I(0) or I(1), in a way depending on the arbitrary choice of the researcher as to the inclusion of the deterministic terms into the unit root testing equations. Neither ADF nor KPDS tests give satisfactory estimations when considered as a whole. Thus we are unable to reject on this point that the variables are integrated at different orders of time series. Such a case is highly crucial in order to be able to determine the appropriate methods for testing the long run economic relationships that must produce stationary econometric counterparts in contemporaneous economics and econometrics thinking.

**Table 1. ADF and KPSS Unit Root Tests**

| | $\tau_c^{ADF}$ | $\tau_{c\&t}^{ADF}$ | $\tau_c^{KPSS}$ | $\tau_{c\&t}^{KPSS}$ | Inference |
|---|---|---|---|---|---|
| Variables | | | | | |
| *TBAL* | -5.06 (2)$^*$ | -5.14 (2)$^*$ | 0.13 (4)$^*$ | 0.05 (4)$^*$ | I(0) |
| Δ*TBAL* | -7.32 (0)$^*$ | -7.28 (0)$^*$ | 0.04 (3)$^*$ | 0.03 (3)$^*$ | I(0) |
| $Y_r$ | -0.32 (8) | -2.40 (8) | 1.19 (5) | 0.10 (16)$^*$ | I(0) or I(1) |
| $\Delta Y_r$ | -3.03 (7)$^*$ | -2.98 (1) | 0.13 (12)$^*$ | 0.10 (12)$^*$ | I(0) |
| *RE* | -1.46 (1) | -0.78 (0) | 0.25 (6)$^*$ | 0.24(6) | I(0) or I(1) |
| Δ*RE* | -6.49 (0)$^*$ | -6.82 (0)$^*$ | 0.31 (3)$^*$ | 0.08 (6)$^*$ | I(0) |
| $Y_r^*$ | 0.06 (0) | -1.67 (0) | 1.06 (6) | 0.13 (6)$^*$ | I(0) or I(1) |
| $\Delta Y_r^*$ | -7.36 (0)$^*$ | -7.33 (0)$^*$ | 0.09 (4)$^*$ | 0.08 (4)$^*$ | I(0) |
| *CRUDE* | -0.56 (0) | -2.14 (0) | 0.81 (6) | 0.23 (6) | I(1) |
| Δ*CRUDE* | -9.23 (0)$^*$ | -9.35 (0)$^*$ | 0.22 (3)$^*$ | 0.03 (5)$^*$ | I(0) |

5% critical values are $\tau_c^{ADF}$ =-2.90, $\tau_{c\&t}^{ADF}$ =-3.47, $\tau_c^{KPSS}$ = 0.46 and $\tau_{c\&t}^{KPSS}$ = 0.15

Notes: Estimations are carried out in EViews 6.0.

## 5. METHODOLOGICAL DISCUSSION

We find in the former section that the variables have different orders to be integrated. Therefore, we cannot apply to the Engle and Granger (1987) or widely popular maximum likelihood based Johansen (1988; 1995) and Johansen and Juselius (1990) multivariate co-integration techniques. Instead of these approaches, let us follow the methodologies developed in Pesaran and Shin (1999) and Pesaran et al. (2001). These estimation techniques, namely autoregressive distribued lag (ARDL) Bounds Testing procedures, can allow us to consider our I(0) and I(1) variables together in a co-integrating equation, and thus, to perpetuate our empirical analysis of the Turkish trade balance. Let us consider the vector error correction model in Eq. (12):

$$\Delta Y_t = \mu + \lambda Y_{t-1} + \sum_{j=1}^{p-1} \gamma_j \Delta Y_{t-j} + \varepsilon_t \tag{12}$$

In Eq. (12), $Y_t$ = $[y_t\ x_t]'$ is defined as the variable vector in which $y_t$ represents the endogeneous variable $TBAL_t$, that is, the trade balance, and $x_t$ represents the explanatory variables vector *a priori* assumed affecting the trade balance which includes domestic and foreign real income levels, real exchange rate and the crude oil prices. $\mu = [\mu_y\ \mu_x]'$ is a vector of constant terms and $\Delta = (1 - L)$ indicates the difference operator. The vector of error terms is assumed to satisfy $\varepsilon_t = [\varepsilon_y\ \varepsilon_x]' \sim N(0, \Omega)$, and $\Omega$ is positive definite. The variance matrice of error terms can be given as follows:

$$\Omega = \begin{bmatrix} \omega_{yy}\omega_{yx} \\ \omega_{yy}\omega_{xx} \end{bmatrix} \tag{13}$$

In Eq. (12), $\lambda$ is the long-run multiplier matrix and $\gamma$ is the short-run reaction matrix, shown in Eq. (14) and Eq. (15).

$$\lambda = \begin{bmatrix} \lambda_{yy}\lambda_{yx} \\ \lambda_{xy}\lambda_{xx} \end{bmatrix} = -\left(I - \sum_{j=1}^{p} \phi_j\right) \tag{14}$$

$$\gamma_j = \begin{bmatrix} \gamma_{yy,j}\gamma_{yx,j} \\ \gamma_{xy,j}\gamma_{xx,j} \end{bmatrix} = -\sum_{k=j+1}^{p} \phi_k \tag{15}$$

$I$ is an identity matrix and $\phi_j$ is the vector autoregression model coefficient matrix. The diagonal elements of matrix $\lambda$ are left unrestricted. Such a case allows for the possibility that the time series used can be either I(0) or I(1). For instance, $\lambda_{yy} = 0$ would imply that the variable $y$ is I(1) and $\lambda_{yy} < 0$ would imply that the variable is I(0). One of the non-diagonal elements of the long run multiplier matrix, $\lambda_{yx}$ or $\lambda_{xy}$, can take zero-value. The bounds testing co-integration approach, considering the above-explained methodology, enables researchers to use variables in testing single co-integrating relationship among the variables, no matter they are I(0) or I(1). In light of these explanations, we can write the possible co-integration relationship as follows:

$$\Delta y_t = \alpha + \varphi y_{t-1} + \delta x_{t-1} + \omega \Delta x_t + \sum_{j=1}^{p-1} \beta_{P,j} \Delta y_{t-j} + \sum_{j=1}^{q-1} \beta_{x,j} \Delta x_{t-j} + u_t \tag{16}$$

In Eq. (16), $\varphi$ and $\delta$ are the long run multiplier coefficients, while $\Delta y_{t-j}$ and $\Delta x_{t-j}$ express the short run dynamic structure of our error correction model. The bounds testing approach requires the ordinary least squares (OLS) estimation of Eq. (16) with or without trend component, and then we must test the absence of a long run relationship between the level values of $y_t$ and $x_t$ by use of the $F$-statistics in line with the below hypotheses:

$$\begin{aligned} &H_0 : \varphi = 0, \delta = 0 \\ &H_1 : \varphi \neq 0, \delta \neq 0 \end{aligned} \tag{17}$$

In Eq. (16), the rejection of $H_0$ hypothesis by the standart $F$- (or Wald-) tests leads to the acceptance of $H_1$ hypothesis and indicates a long run equilibrium relationship between the variables. The statistics such estimated, then, are compared with the non-standard distributed asymptotic critical value bounds reported in Pesaran et al. (2001). If estimated $F$-statistic falls outside of the critical value bounds, we can definitely infer whether or not there exists a co-integrating relationship between the variables, regardless of the order of integration of the variables. In this case, if $F$-statistic exceeds its respective upper critical values, this means rejection of the null hypothesis of no co-integration between the variables. If $F$-statistic is found below the lower critical value bounds, we cannot reject non-existence of a co-integrating relationship. If estimated statistic lies between the bounds, we cannot make any conclusive inference as to the existence of a possible co-integrating relationship and need to know the order of integration of the underlying regressors.

Having tested the existence of a potential co-integration relationship between the variables, the most appropriate lag specification of the variables in the ARDL model must be determined through the widely-used lag information criteria in the economics and econometrics literature, so that the long run equilibrium and short run dynamic error correction model coefficients can be estimated by way of employing the standard OLS methodology.

In addition to this estimation procedure, if we find that the value of the *t*-statistic of the one-period lagged coefficient of the dependent variable ($\varphi$) in Eq. (16) is greater than the critical values reported by Pesaran et al. (2001), this would also reflect the existence of a co-integrating relationship between the variables in the model. Note that Pesaran and Shin (1999) bring out that the ARDL-based bounds testing approach is able to yield consistent long run coefficient estimators even in small samples.

## 6. BOUNDS TESTING ESTIMATION RESULTS

As a next step in our study, the ARDL bounds testing co-integration and error correction modeling approaches have been used to examine the validity of the economic modeling issues constructed in the former sections. For this purpose, at first, the appropriate lag length ($p$) is tried to be determined. Following Pesaran et al. (2001), for $p = 1, 2, \ldots, 6$, the conditional error correction model in Eq. (16) is estimated by OLS methodology both with and without trend components in the regression. The results are given in Table 2:

**Table 2. Selection of the Lag Order for the Trade Balance Eq. (11)**

| | With deterministic trend | | | | Without deterministic trend | | | |
|---|---|---|---|---|---|---|---|---|
| $p$ | AIC | SC | $\chi^2_{SC}(1)$ | $\chi^2_{SC}(4)$ | AIC | SC | $\chi^2_{SC}(1)$ | $\chi^2_{SC}(4)$ |
| 1 | -2.89 | -2.28 | 0.47 | 0.40 | -2.84 | -2.27 | 0.87 | 0.42 |
| 2 | -2.91 | -2.12 | 1.85 | 1.19 | -2.80 | -2.04 | 0.20 | 0.36 |
| 3 | -3.36 | -2.40 | 0.01 | 0.66 | -3.11 | -2.18 | 1.12 | 0.78 |
| 4 | -3.43 | -2.32 | $4.60^{**}$ | 1.71 | -3.16 | -2.08 | $5.90^{**}$ | 2.12 |
| 5 | -3.65 | -2.37 | 2.04 | $3.34^{**}$ | -3.33 | -2.08 | 0.25 | $3.78^{**}$ |
| 6 | -3.66 | -2.22 | 0.25 | 1.94 | -3.24 | -1.82 | 1.16 | $3.99^{**}$ |

Notes: Estimations are carried out in EViews 6.0.

In Table 2, '$p$' is the lag order of the underlying VAR for the conditional error correction model in Eq. (16). AIC and SC represent Akaike and Schwarz information criterions, respectively. $\chi^2_{SC}(1)$ and $\chi^2_{SC}(4)$ are Breusch-Godfrey error terms Lagrange multiplier serial correlation test $F$-statistics under the null hypothesis of no serial correlation at orders 1 and 4, respectively. '*', '**' and '***' denote significance at 0.01, 0.05 and 0.10 levels, respectively. We can see that the best model with no serial correlation problem is the one that uses 3 lag lengths for the ARDL equation:

**Table 3. ARDL Unrestricted Error Correction Model of the Turkish Trade Balance**

| Dependent Var.: $\Delta TBAL$ | Coefficient | Std. Error | $t$-Statistic | Prob. |
|---|---|---|---|---|
| $C$ | c(1) = 2.8349 | 1.2192 | 2.3252 | 0.0255 |
| $\Delta TBAL_{t-1}$ | c(2) = 0.2180 | 0.1401 | 1.5565 | 0.1279 |
| $\Delta TBAL_{t-2}$ | c(3) = -0.1220 | 0.1243 | -0.9816 | 0.3325 |
| $\Delta TBAL_{t-3}$ | c(4) = -0.0974 | 0.1304 | -0.7471 | 0.4596 |
| $\Delta RE_t$ | c(5) = 0.1727 | 0.1081 | 1.5984 | 0.1182 |
| $\Delta RE_{t-1}$ | c(6) = -0.0154 | 0.1379 | -0.1116 | 0.9117 |
| $\Delta RE_{t-2}$ | c(7) = 0.1409 | 0.1054 | 1.3366 | 0.1893 |
| $\Delta RE_{t-3}$ | c(8) = -0.0403 | 0.1193 | -0.3381 | 0.7372 |
| $\Delta Y_{r,t}$ | c(9) = -0.9260 | 0.3267 | -2.8341 | 0.0073 |
| $\Delta Y_{r,\,t-1}$ | c(10) = -0.6635 | 0.2393 | -2.7727 | 0.0086 |
| $\Delta Y_{r,\,t-2}$ | c(11) = -0.4950 | 0.2701 | -1.8513 | 0.0719 |
| $\Delta Y_{r,\,t-3}$ | c(12) = -0.9542 | 0.2640 | -3.6148 | 0.0009 |
| $\Delta Y^*_{r,t}$ | c(13) = -0.4907 | 0.6965 | -0.7046 | 0.4853 |
| $\Delta Y^*_{r,t-1}$ | c(14) = 0.5981 | 0.7645 | 0.7823 | 0.4389 |
| $\Delta Y^*_{r,t-2}$ | c(15) = 2.1957 | 0.5419 | 4.0522 | 0.0002 |
| $\Delta Y^*_{r,t-3}$ | c(16) = 0.9746 | 0.5874 | 1.6593 | 0.1053 |
| $\Delta CRUDE_t$ | c(17) = -0.0540 | 0.0419 | -1.2864 | 0.2061 |
| $\Delta CRUDE_{t-1}$ | c(18) = -0.0684 | 0.0612 | -0.1.1175 | 0.2708 |
| $\Delta CRUDE_{t-2}$ | c(19) = -0.0645 | 0.0420 | -1.5350 | 0.1331 |
| $\Delta CRUDE_{t-3}$ | c(20) = -0.1458 | 0.0303 | -4.8100 | 0.0000 |
| $TBAL_{t-1}$ | c(21) = -0.6794 | 0.1265 | -5.3723 | 0.0000 |
| $RE_{t-1}$ | c(22) = 0.1213 | 0.0562 | 2.1565 | 0.0374 |
| $Y_{r,t-1}$ | c(23) = -0.9199 | 0.2425 | -3.7939 | 0.0005 |
| $Y^*_{r,t-1}$ | c(24) = 0.6934 | 0.2696 | 2.5711 | 0.0142 |
| $CRUDE_{t-1}$ | c(25) = -0.0150 | 0.0288 | -0.5218 | 0.6049 |
| TREND | c(26) = 0.0133 | 0.0026 | 5.0468 | 0.0000 |
| DUMMY2 | c(27) = 0.0520 | 0.0601 | 0.8660 | 0.3919 |
| DUMMY3 | c(28) = 0.0373 | 0.1003 | 0.3714 | 0.7124 |
| DUMMY4 | c(29) = 0.1423 | 0.0957 | 1.4872 | 0.1452 |

Notes: Estimations are carried out in EViews 6.0.

**Table 4. Regression Statistics in Table 3**

| | | | |
|---|---|---|---|
| $R^2$ | 0.7728 | Mean dependent var | 0.0007 |
| Adjusted $R^2$ | 0.6055 | S.D. dependent var | 0.0619 |
| S.E. of regression | 0.0389 | Akaike info criterion | -3.3579 |
| Sum squared resid | 0.0575 | Schwarz criterion | -2.4036 |
| Log likelihood | 141.49 | Hannan-Quinn criterion | -2.9803 |
| *F*-statistic | 4.6175 | Durbin-Watson statistic | 1.9903 |
| Prob(*F*-statistic) | 0.0000 | | |

Notes: Estimations are carried out in EViews 6.0.

The existence of a potential co-integration relationship between the variables has been examined by comparing our estimates with the critical values reported in Table CI(iv), Table CI(v) and Table CII(v) of Pesaran et al. (2001):

**Tablo 5. *F*- and *t*-statistics for Testing the Existence of Co-integration**

| $p$ | $F_{iv}$ | $F_v$ | $t_v$ |
|---|---|---|---|
| 3 | 6.19 | 7.27 | -5.37 |
| 0.05 Table Critical Values | | | |
| I(0) | 3.38 | 4.01 | -3.41 |
| I(1) | 4.23 | 5.07 | -4.16 |

Notes: $F_{iv}$ indicates the $H_0$ hypothesis that c(21) = c(22) = c(23) = c(24) = c(25) = c(26) = 0. $F_v$ indicates the $H_0$ hypothesis that c(21) = c(22) = c(23) = c(24) = c(25) = 0. $t_v$ indicates the $t$-statistic of the coefficient of one-period lagged dependent variable c(21) in Table 3. Critical values are taken from Pesaran et al. (2001). Estimations are carried out in EViews 6.0.

$F_{iv}$ is the $F$-statistic calculated by applying to Wald tests that impose zero value restriction to the one-period lagged level coefficient values and deterministic trend component. $F_v$ is the $F$-statistic calculated by applying to Wald tests that impose zero value to the only one-period lagged level coefficient values of the variables. $t_v$ is the $t$-statistic of the coefficient of one-period lagged level value of dependent variable, that is, *TBAL*, in Table 3. We can observe that estimation results of the $F$-statistics exceed the upper critical values, and thus, infer that there exists a co-integrating relationship between the time series in the level form, without considering whether they are I(0) or I(1). The $t$-statistic of the one-period lagged level value of the dependent variable also supports these findings in favor of co-integration. Following Bårdsen (1989), we can obtain the long run coefficients in Eq. (16) by dividing one period lagged level coefficient values of independent variables to the one period lagged level coefficient value of dependent variable, by multiplying this result with minus one, that is, - ($\delta$ / $\varphi$). However, as Atkins and Serletis (2003) specify, there is no reason why $p$ and $q$ in Eq. (16) should have the same value. Instead, we can assume for all the variables in Eq. (16) different lag structures taking values from one to six. Such a procedure requires estimation of too many ARDL models running regressions using all the possible lag lengths of variables to obtain a parsimonious model. Pesaran et al. (2001) apply to this estimation procedure when estimating their model. We use Microfit 4.0. software program in achieving this task and estimate that ARDL (1 0 2 0 0) model best fits in with the Turkish data. The long run coefficients of the model are given in Table 6:

**Table 6. Estimated Long Run Coefficients using the ARDL Approach**
**ARDL (1 0 2 0 0) Selected Based on Schwarz Bayesian Criterion**

| Dependent Variable: TBAL | Coefficient | Std. Error | $t$-Statistic | Prob. |
|---|---|---|---|---|
| $RE$ | 1.0885 | 0.4332 | 2.5129 | 0.015 |
| $Y_r$ | -2.5536 | 1.1719 | -2.1790 | 0.034 |
| $Y_r^*$ | 4.5664 | 2.0407 | 2.2377 | 0.030 |
| $CRUDE$ | 0.1080 | 0.1183 | 0.9127 | 0.366 |
| $C$ | 20.686 | 9.0660 | 2.2817 | 0.027 |
| D2 | 0.6734 | 0.3145 | 2.1410 | 0.037 |
| D3 | 1.1572 | 0.5259 | 2.2005 | 0.032 |
| D4 | 0.6709 | 0.4103 | 1.6349 | 0.108 |

Notes: Estimations are carried out in Microfit 4.0.

Our estimation results reveal that in a long run period satisfying a stationary relationship between the variables real exchange rate behaves in accordance with *a priori* model expectations in Eq. (9). The depreciation of real exchange rate seems to improve the trade balance in a strong and significant way. A 1% increase in real exchange rate leads to nearly 1.09% improvement on trade balance. The domestic real income affects the trade balance negatively. A 1% increase in domestic real income leads to 2.55% deterioration in trade balance. We can attribute this effect to the pressures of domestic absorption on trade balance. Indeed, such an inference is in line with the stylized facts of the Turkish economy outlined in Figure 1 and Figure 2. The foreign real income seems to be one of the most prominent determinants of the Turkish trade balance and takes an estimation value in line with our model formation in Eq. (9). The trade balance improves 4.57% as a result of 1% increase in foreign real income. We cannot find a significant effect of oil prices on trade balance. The error correction model has been given in Table 7. The real exchange rate depreciation improves the trade balance and the domestic real income has a negative effect with its dynamic lag on trade balance. An increase in foreign income has a negative impact on trade balance in the short run. We are unable to obtain a significant result for oil prices as is estimated in the long run model. The error correction term points out that nearly 24% of the disequilibrium conditions within our co-integration model is corrected within one period:

**Table 7. Error Correction Representation for the Selected ARDL Model**
**ARDL (1 0 2 0 0) Selected Based on Schwarz Bayesian Criterion**

| Dependent Var.: $\Delta TBAL$ | Coefficient | Std. Error | *t*-Statistic | Prob. |
|---|---|---|---|---|
| $\Delta RE_0$ | 0.2558 | 0.0608 | 4.2066 | 0.000 |
| $\Delta Y_{r,0}$ | -0.0314 | 0.0161 | -1.9469 | 0.057 |
| $\Delta Y_{r,1}$ | -0.5647 | 0.1540 | -3.6667 | 0.001 |
| $\Delta Y^*_{r,0}$ | -1.0730 | 0.2589 | -4.1446 | 0.000 |
| $\Delta CRUDE_0$ | 0.0254 | 0.0286 | 0.8879 | 0.379 |
| $C$ | 4.8607 | 1.1376 | 4.2728 | 0.000 |
| D2 | 0.1582 | 0.0403 | 3.9297 | 0.000 |
| D3 | 0.2719 | 0.0725 | 3.7496 | 0.000 |
| D4 | 0.1576 | 0.0839 | 1.8800 | 0.066 |
| $ecm_{t-1}$ | -0.2350 | 0.0800 | -2.9358 | 0.005 |

Notes: Estimations are carried out in Microfit 4.0.

In Table 7, $\Delta RE_0 = RE_t - RE_{t\text{-}1}$, $\Delta Y_{r,0} = Y_{r,t} - Y_{r,t\text{-}1}$, $\Delta Y_{r,1} = Y_{r,t\text{-}1} - Y_{r,t\text{-}2}$, $\Delta Y^*_{r,0} = \Delta Y^*_{r,t}$ - $\Delta Y^*_{r,t-1}$, $\Delta CRUDE_0 = \Delta CRUDE_t$ - $\Delta CRUDE_{t\text{-}1}$. The regression statistics of the error correction model and diagnostic test results are reported in Table 8 and Table 9, respectively:

**Table 8: Regression Statistics of the Error Correction Model**

| | | | |
|---|---|---|---|
| $R^2$ | 0.4587 | Mean dependent var | 0.4239 |
| Adjusted $R^2$ | 0.3546 | S.D. dependent var | 0.0508 |
| S.E. of regression | 0.0508 | Akaike info criterion | 93.366 |
| Sum squared resid | 0.1343 | Schwarz Bayesian criterion | 81.579 |
| Log likelihood | 104.37 | Durbin-Watson statistic | 2.0045 |
| *F*-statistic | 4.8962 | Prob(*F*-statistic) | 0.0000 |

Notes: Estimations are carried out in Microfit 4.0.

**Table 9: Diagnostic Tests**

| Tests Statistics | LM Version | *F* Version |
|---|---|---|
| A: Serial Correlation | Chi-square (4) = 4.0643 (0.397) | 0.8275 (0.514) |
| B: Functional Form | Chi-square (1) = 2.0246 (0.155) | 1.6934 (0.199) |
| C: Normality | Chi-square (2) = 1.0714 (0.585) | Not applicable |
| D: Heteroskedasticity | Chi-square (1) = 10.146 (0.001) | 11.709 (0.001) |

Notes: Estimations are carried out in Microfit 4.0.

## 7. CONCLUDING REMARKS

In this paper, we tried to analyze the determinants of the Turkish trade balance. For this purpose, we first documented some stylized facts of the Turkish economy for the post-1987 period till 2007. Then, we briefly examined a well-constructed reduced form model in the international trade balance literature. A survey of empirical papers for the Turkish economy followed these theoretical section. In the fourth and fifth sections, data and time series characteristics were presented and a methodological discussion for the estimation process were carried out. The sixth section of the paper formed our main contribution to the existing literature and aimed at giving an empirical essay for the Turkish trade balance.

The main results of the paper can be summarized as follows. First of all, we must emphasize that the time series course of the trade balance has a stationary form within the investigation period. Real exchange rate seems to be stationary, but no conclusive inference due to some popular unit root tests can be done for all other variables except the crude oil price variable, which was included into the paper to account for any other possible effects on trade balance resulted from the developments in the world markets. Therefore, we chose to apply to the contemporaneous ARDL based bounds testing approach to test the existence of a long run co-integration relationship between the variables. Our estimation results indicated that in a long run perspective which provides a stationary relationship between the variables, real exchange rate depreciations improves the trade balance in a strong and significant way, that domestic real income affects the trade balance negatively which reflects the pressures of domestic absorption on trade balance, and that trade balance is strongly improved due to an increase in foreign real income. No significant effect of oil prices can be observed on trade balance. The error correction modeling gives results in line with the long run findings of the co-integration analysis. Of course, these results are highly open to be criticized, and future studies using more analytical approaches to test the course of the trade balance must be constructed to examine the validity of the estimation findings obtained in this paper.

The usual disclaimer applies.

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