Re-examining Turkey's trade deficit with structural breaks: Evidence from 1989-2011

Erten, Irem and Okay, Nesrin

Financial Engineering Program, Bogazici University, Department of Management, Financial Engineering Program, Bogazici University

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Irem Erten
Program of Financial Engineering, Boğaziçi University
Bebek 34342, İstanbul – Turkey
lirem.erten@gmail.com, Tel: +90 (530) 2638064

Nesrin Okay
Department of Management, Program of Financial Engineering, Boğaziçi University
Bebek 34342, İstanbul – Turkey
okay@boun.edu.tr, Tel: +90 (212) 3596508, Fax: +90 (212) 2877851

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The goal of this paper is to examine the sustainability of the trade deficit of Turkey with cointegration techniques allowing for structural breaks. We follow Husted (1992) model, which shows that if a country's exports and imports are cointegrated, and if the cointegrating vector is (1,-1), then its trade deficit is sustainable. First, classical cointegration tests indicate that exports and imports are cointegrated, but the cointegrating vector significantly differs from (1,-1). Next, the existence of cointegration is confirmed with an alternative method proposed by Silvestre and Sansó (2006) controlling for structural breaks. Our analysis detects two breaks in the cointegration relationship on 2001:01, and 2008:09, coinciding with economic crises. We show that since 2001, Turkish trade deficit has not been sustainable in the strong form, and after 2008, the country has moved away from sustainability. Based on our analysis, the sustainability of Turkey's widening trade deficit is highly doubtful.

Keywords: Exports, Imports, Trade Deficit, Sustainability, Cointegration, Structural Breaks
JEL classification: C22; G14; G15

I Introduction

Sustainability of trade deficit is one of the major concerns for emerging countries and has been drawing increasing attention from economists. In economic literature, sustainability of current account deficit essentially refers to a country's solvency and capacity to service its external debt (Crockett and Goldstein, 1987). Baharumshah *et al.* (2003) explain that continuous trade imbalances eventually lead to high domestic interest rates, excessive borrowing and trigger financial crises. Since a persistent trade deficit can expand foreign debt and threaten a country's growth and solvency, sustainability of trade balance is considered one of the key indicators of macroeconomic stability.

Since trade liberalization in 2002, Turkey's trade deficit has been consistently rising, and now it raises serious concerns regarding its sustainability. The foreign deficit spiked from \$1.84 billion in 2001 to \$8.12 billion in 2011, according to the figures published by the Central Bank of the Republic of Turkey (CBRT). This record rise is problematic because Turkey is a developing country where international trade is a vital component of economic growth. Moreover, a growing deficit implies that the country becomes more dependent on external financing and more vulnerable to global shocks. The current trade deficit of Turkey seems to be mainly caused by raw and processed materials used in production. This is due to the fact that Turkey relies on imports for most of its energy needs and on foreign intermediate goods for industrial production.

Husted (1992) has developed a cointegration-based framework to analyse the dynamics of a country's trade balance. He has shown that the existence of cointegration between exports and imports with parameters (1, -1) indicates that a country's trade balance is sustainable and that international budget constraint is not violated. Following his framework, many researchers have examined the sustainability of Turkey's trade deficit (Utkulu, 1998; Bozdağloğlu, 2007; Erbaykal and Karaca, 2008; Ucan and Putun, 2011; Göktaş *et al.*; 2011). These studies have applied cointegration methods in different periods, leading to different conclusions regarding the sustainability of the Turkish trade deficit.

In this paper, we investigate whether Turkey's trade deficit is sustainable with robust cointegration methods, including techniques allowing for structural breaks. Previous studies that have applied cointegration techniques with structural breaks to Turkey's international trade data have concluded that exports and imports are cointegrated, and that the Turkish current account deficit is sustainable only in the weak form (Erbaykal and Karaca, 2008; Göktaş *et al.*; 2011). We contribute to the existing literature by employing an alternative methodology developed by Silvestre and Sansó (2006) in order to obtain stronger evidence of sustainability (or lack thereof). Furthermore, we extend the time frame of previous studies beyond 2008, which allows us to analyse whether Turkey's export-import dynamics have shifted after the global financial crisis.

In our analysis, we use aggregate monthly data of exports and imports from 1989-2011.

First, we apply the conventional cointegration techniques to the data set and next, we extend the analysis allowing for multiple structural breaks with the methodology of Carrion-i-Silvestre and Sansó (2006). The aspect of considering structural breaks is particularly important, because the time span of the study is long, and Turkey experienced several financial crises and institutional changes during this period. Hence, the trade balance is likely to be subject to variation due to economic events and reforms, which needs to be considered to better assess the issue of deficit sustainability.

Based on a statistical analysis of Turkey's export-import dynamic relationship, we find that Turkey's trade deficit fulfills the weak form sustainability. Existence of cointegration with a two-break structure, the implications of the break dates, and sub-sample analysis provide insight into the stability and structure of Turkey's trade balance. We organize the rest of the paper as follows. The underlying theoretical framework is described in Section II; the econometric methodology is explained in Section III, and the data are presented in Section IV. Section V discusses the empirical results, and Section VI concludes.

II Model

Husted (1992) has developed a cointegration-based framework to analyse the dynamics of a country's trade balance. He starts his model by setting up the intertemporal budget constraint for a given economy:

$$B_{t+1} = Y_{t+1} - C_{t+1} - I_{t+1} + (1 + r) \times B_t \tag{1}$$

where C, Y, B, and I denote consumption, output, net borrowing and investment, respectively; r is one-period interest rate.

Husted then proceeds by deriving a test model from Equation 1, which can be used to determine whether a country satisfies its intertemporal budget constraint[1]:

$$exports_t = \theta \times imports_t + \epsilon_t \tag{2}$$

Following Arize (2002), this model can also be written as:

$$imports_t = \theta \times exports_t + \epsilon_t \tag{3}$$

According to this test model, there exist two conditions for a country to maintain its intertemporal budget constraint. First, the error term of Equations 2 and 3 should be stationary. This means that exports and imports are cointegrated and have a long run relationship. Failure to fulfill this condition shows that the economy does not function as required and has not succeeded in maintaining its intertemporal budget constraint (Erbaykal and Karaca, 2008).

Second, the slope coefficient θ should be statistically equal to 1, and the cointegrating vector (1, -1). If exports and imports are cointegrated with a cointegrating vector (1, -1), then $imports_t$-$exports_t$ (or trade deficit) becomes a stationary process. If the slope coefficient is lower than 1 in Equation 2 or higher than 1 in Equation 3, imports exceed exports and trade deficit can grow without bounds. Similarly, if the slope coefficient is higher than 1 in Equation 2 or if it is lower than 1 in Equation 3, exports exceed imports, and trade surplus can permanently grow. (Erbaykal and Karaca, 2008).

III Econometric Methodology

According to the Husted model (1992), the primary condition for sustainability is a long-run relationship (cointegration) between exports and imports. Since the methods for determining the presence of cointegration require that the variables be integrated of the same order, we start

[1] See Husted, 1992 for the derivation

by analyzing the time-series properties of the series. We employ the Augmented Dickey-Fuller (1981) and Philips-Peron (1988) tests, in addition to the Montañés-Clemente-Reyes Additional Outliers and Innovational Outliers Models (1988), which allow for structural breaks.

After verifying that the series are I(1), we analyse the cointegration relationship between exports and imports by performing the standard Engle-Granger (1987) and Johansen (1990) Trace and Maximum EigenValue Tests. However, breaks may introduce spurious unit root behavior in the cointegrating relationship and yield misleading results for standard cointegration tests (Gregory *et al.*, 1996). For that reason, we focus on the possibility of structural shifts and employ the methodology proposed by Carrion-i-Silvestre and Sansó (2006). The approach of Carrion-i-Silvestre and Sansó tests the null hypothesis of a cointegrating relation in the possibility of one or many structural breaks against the alternative of no cointegration (with possible breaks). The technique is a complementary method to ensure the presence of cointegration around a break-cointegrating relationship. The test statistic is an extension of of the KPSS test (Kwiatkowski *et al.*, 1992) and corrects for the presence of endogeneous regressors. The model allows for structural breaks in both deterministic and cointegrating vector and estimates the following:

$$y_t = \alpha + \theta DU_t + x_t'\beta_1 + x_t'\beta_2 DU_t + \sum_{i=-n}^{n} \gamma_i \Delta x_{t-1} + \epsilon_t \tag{4}$$

$$y_t = \alpha + \theta DU_t + \epsilon t + \gamma DT_t^* x_t'\beta_1 + x_t'\beta_2 DU_t + \sum_{i=-n}^{n} \gamma_i \Delta x_{t-1} + \epsilon_t \tag{5}$$

where $DU_t = 1$ for $t > T_b$ and zero otherwise, DT_t*=1 for $t > T_b$ and 0 otherwise, $\alpha + \theta DU_t$ *level shift*, *t time trend*, and γDT_t* *change in slope of the time trend*. In the above specification, DU is a dummy variable that captures the stuctural break in the long-run relationship between the series. The difference between Equations 4 and 5 is that Equation 5 considers a time trend. To test the null hypothesis of co-integration, the test employs the LM statistic, which is given as:

$$SC(\lambda) = T^{-2} \times \hat{w}_1^{-2} \times \sum_{t=1}^{T} S_t^2 \tag{6}$$

where $\lambda = \frac{T_b}{T}$, $\hat{w_1^2}$ denotes a consistent estimator of the long-run variance of $\hat{e}_{k}{}_{k=1}^{T}$, $S_t = \sum_{k=1}^{t} \hat{e}_k$, and $\hat{e}_{k}{}_{k=1}^{T}$, are the estimated residuals derived from Equations 4 and 5. If the computed LM statistic is greater than the critical value, then the null hypothesis of cointegration is rejected. Silvestre and Sansó (2006) prove that the LM statistic in Equation 6 converges asymptotically to:

$$SC(\lambda) = T^{-2} \times \hat{w_1}^{-2} \times \left[\sum_{t=1}^{Tb} (\sum_{j=1}^{t} \hat{e}_j)^2 + \sum_{t=1}^{Tb} (\sum_{j=1}^{t} \hat{e}_j)^2 \right] \Rightarrow \lambda^2 \int_0^1 V_{k,i}^2(b_1)db_1 + (1-\lambda)^2 \int_0^1 V_{k,i}^2(b_2)db_2 \tag{7}$$

where V_k^i are functions of Wiener process.

The primary advantage of this approach is that it allows for a shift in the cointegrating relation. Furthermore, it does not require a model that assumes covariance stationary variables or a cointegrated system a priori (Beyer, Haug and Dewald, 2009).

IV Data

The data set is obtained from the Central Bank of the Republic of Turkey (CBRT), and it consists of the monthly observations of Total Exports and Total Imports (measured in millions of dollars and at current exchange rate) between 1989:01 and 2011:12. In this study, we employ nominal exports and imports measured in millions USD[2], and we do not use seasonal correction because the time series do not exhibit seasonal behavior[3]. The data series are plotted in Figs. 1 to 3.

[2] We conduct the analysis in U.S. dollars because the focus of our study is the sustainability of trade deficit, which mainly concerns indebtedness, and much of Turkey's indebtedness is denominated in dollars and other non-Turkish currencies.

[3] In order to check for seasonality in the data, we employ the HEGY test (1990), and we detect no seasonal unit roots and no seasonal cycles in the data. Results are available upon request.

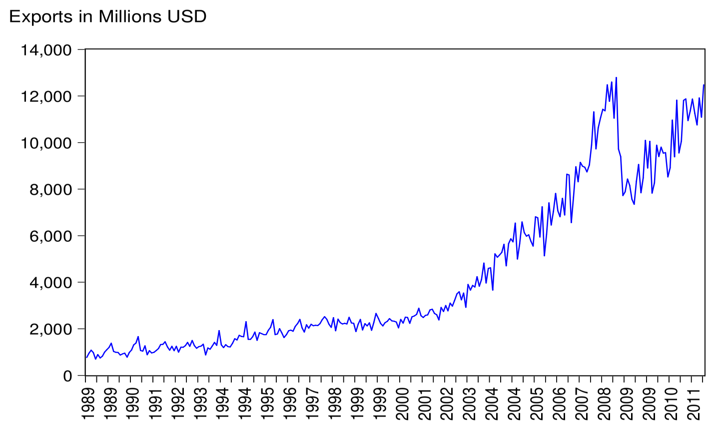

Fig. 1: Total exports

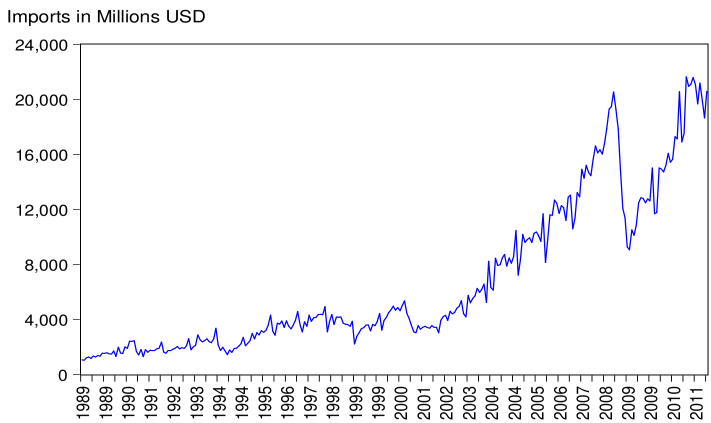

Fig. 2: Total imports

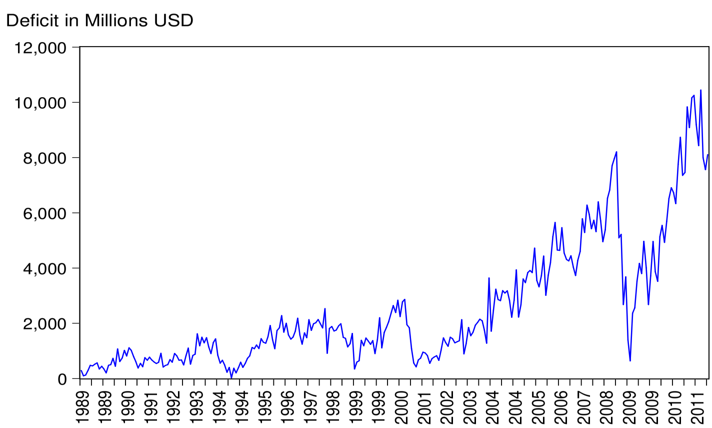

Fig. 3: Trade deficit

Table 1: ADF and PP Tests

	Variables	ADF Test Drift and Trend	PP Test Drift and Trend
Level	*Imports*	-2.639	$-$ 2.219
	Exports	-1.609	$-$ 2.851
Difference	Δ *Imports*	-3.939 ***	-21.335***
	Δ *Exports*	-5.833 ***	-30.697 ***

Notes: MacKinnon (1996) critical values for ADF and PP Tests with Drift and Trend at 1%, 5% and 10% significance level are -3.99,-3.43, and -3.14, respectively. * significant at level of 10%, ** significant at level of 5%, *** significant at level of 1%.

Table 2: Montañés-Clemente-Reyes Test

Additive Outliers Method		
Variables	min-*t*	Optimal Breakpoints
Imports	-4.684	2003:09, 2006:12
Exports	-3.211	2002:12, 2006:12
Innovative Outliers Method		
Variables	min-*t*	Optimal Breakpoints
Imports	-4.094	2003:01, 2006:01
Exports	-2.989	2003:01, 2005:12

Notes: Critical value for Montañés-Clemente-Reyes Test (1998) at 5% significance level is -5.490.

V Empirical Results

Results of Unit Root Tests

As shown in Table 1, Augmented Dickey-Fuller (1981) and Philips-Peron (1988) tests suggest that both *imports* and *exports* are I(1). Hence, the series are nonstationary on level, but first-order-differencing makes them stationary. However, this conclusion might be misleading if there are structural breaks in the variables (Lee and Chang, 2005). For that reason, we perform the Montañés-Clemente-Reyes AO and IO Tests (1988), which allow for two breaks. The results in Table 3 indicate that the null of a unit root cannot be rejected at 5% level of significance, with the breaks occuring generally in 2003 and 2006. Figs 4 and 5 plot the break dates detected with the IO Test.

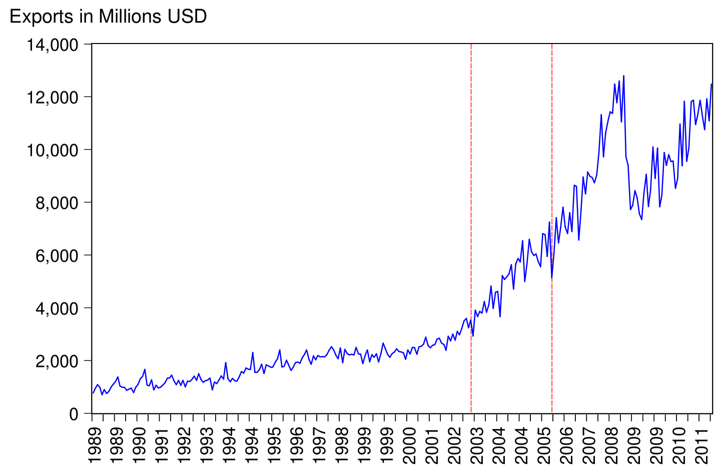

Fig. 4: Breaks in export series detected with the IO Test

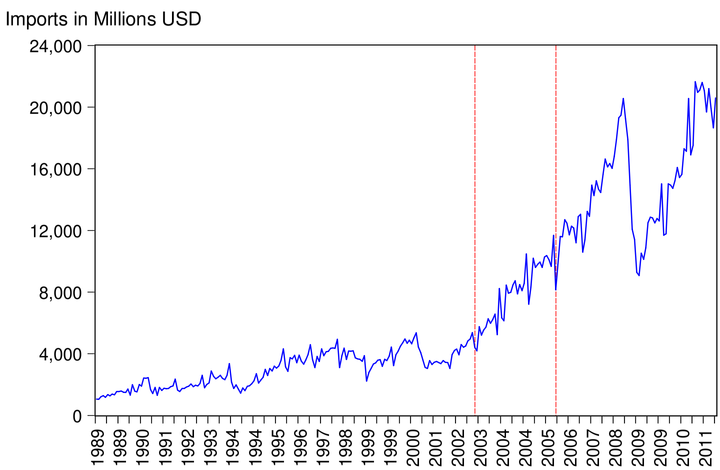

Fig. 5: Breaks in import series detected with the IO Test

Results of Standard Cointegration Analysis

Although the series are found to have unit root with breaks, we start our analysis with the standard cointegration analysis for the sake of completeness. Since the series are nonstationary and integrated of the same order I(1), we can employ the Engle-Granger (1987) and Johansen (1990) Trace and Maximum EigenValue Tests. Tables 3 and 4 report the results. All the tests reject the null hypothesis of no cointegration, indicating that Turkey's exports and imports have a long-term equilibrium relation. According to Husted (1992), this finding means that Turkey fulfills the first condition for sustainability.

Since *exports* and *imports* are found to be cointegrated, we estimate an error correction model (ECM) that describes their adjustments towards a long-run equilibrium. The model consists of one period cointegrating equation and the lagged first differences of the endogenous variables. We present the estimated ECM in Table 5 and report the standard errors in the brackets. While the adjustment coefficient on *exports* is positive and as high as 10.32%, the adjustment coefficient on *imports* is negative and quite small, -3.88%. Since a positive error correction term indicates a movement away from equilibrium, the estimated adjustment coefficients cast doubt on the cointegrating relationship. This result motivates further analysis to confirm the presence of cointegration.

Fig. 6 plots the cointegration residual from the estimated ECM. We see that although both series have stayed at reasonable proximity over the last 20 years, the tendency to drift away from the equilibrium has increased after the global crisis of 2008.

Finally, the LR tests are conducted on the VECM in order to determine whether the cointegrating vector is significantly different from (1, -1). The results in Table 6 indicate that the null hypothesis that the cointegrating vector is (1, -1) is clearly rejected.

Table 3: Engle-Granger Analysis

Dependent	*Tau*-Statistic	p-value	z-statistic	p-value
Imports	−4.7856	0.0005	−4.6447	0.0001
Exports	−3.9388	0.0099	−3.3354	0.0025

Table 4: Johansen Cointegration Analysis

Trace Test			
Null Hypotheses	Test Statistic	Critical Value at 5%	p-value
No Cointegrating Vector	20.8969	15.4947	0.0069
At Most 1 Cointegrating Vector	0.2116	4	0.6455
Max Eigenvalue Test			
Null Hypotheses	Test Statistic	Critical Value at 5%	p-value
No Cointegrating Vector	20.6852	14.2646	0.0042
Exactly 1 Cointegrating Vector	0.2116	3.8415	0.6455

Results of Cointegration Tests with Structural Breaks

Ignoring structural breaks can lead to misspecification errors concerning the existence of cointegration (Banerjee and Carrion-i-Silvestre, 2006). For that reason, we perform the methodology proposed by *Silvestre and Sansó*, allowing for structural breaks. We first consider one structural break and compute the test statistic 2.05×10^{-11} (Eviews Codes are available upon request). The test statistic is almost zero and does not exceed the critical values tabulated by Silvestre and Sansó (2006). Therefore, the null hypothesis of cointegration with one structural break cannot be rejected, and imports and exports are found cointegrated with one structural break.

Considering the possibility of multiple structural change, we also conduct the test of *Silvestre and Sansó* controlling for multiple breaks. We take the breakpoints exogenously as those obtained from the *Bai-Perron Test* (1998). (The test is performed with Bai-Perron Add-in in Eviews 7). Bai-Perron analysis estimates two breaks in the export-import relationship on 2001:01, and 2008:09. The dates of break coincide with the aftermath of the domestic and global financial crises, respectively.

Table 5: Vector Error Correction Model

D(Imports)=	−0.0388[*Imports(-1)*	−1.5777*Exports(-1)*	−71.42199]
	[−0.59687]	[−28.5579]	
	−0.1627*D(Imports(-1))*	−0.3135*D(Imports(-2))*	−0.2010*D(Exports(-1))*
	[−1.72362]	[−3.47701]	[−1.34544]
	+0.2989 *D(Exports(-2))*	+98.40	
	[2.14129]	[1.80062]	
D(Exports)=	0.1033[*Imports*(-1)	−1.5777*Exports*(-1)	−71.42199]
	[2.58841]	[−28.5579]	
	+0.1047*D(Imports(-1))*	−0.1566*D(Imports(-2))*	−0.6256(*D(Exports(-1)*
	[1.80893]	[−2.83303]	[−6.82823]
	−0.0735*D(Exports(-2))*	+72.06	
	[−0.85903]	[2.15087]	

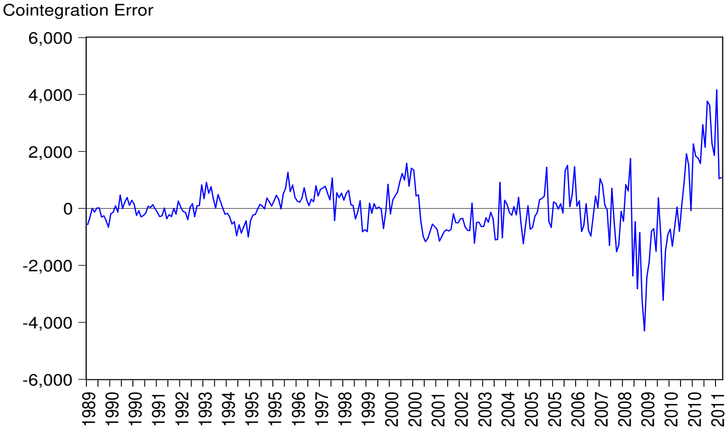

Fig. 6: Cointegration residual of the VECM

Table 6: Results of LR-Test

Cointegrating Vector	Null Hypothesis	*p*-value
Exports(−1)+(−0.6338)**Imports*(−1)+C=0	Coefficient on *Imports*=-1	0.000085
Imports(−1)+(−1.5777)**Exports*(−1)+C=0	Coefficient on *Exports*=-1	0.000085

Table 7: Engle-Granger Analysis

Sample	Dependent	*Tau*-Statistic	*p*-value	*z*-statistic	*p*-value
1989:01-	*Imports*	−5.0185	0.0015	−51.3234	0.0001
2001:01	*Exports*	−3.6293	0.0834	−26.6046	0.0431
2001:01-	*Imports*	−7.1333	0.0000	−65.2684	0.0000
2008:09	*Exports*	−7.1278	0.0000	−66.6254	0.0000
2008:09-	*Imports*	−6.6646	0.0001	−45.00382	0.0000
2011:12	*Exports*	−5.3873	0.0019	−34.8721	0.0011

To consider both breaks, we extend Silvestre and Sansó's model, since their original technique only allows for one structural break. We compute the model t-statistic with the two break dates estimated with Bai-Perron and estimate new critical values for the given break fraction vector. To this end, we simulate the test statistic with the estimated break dates 20000 times and compute the critical value as 41.28. Next, we find the SC test-statistic based on the estimated break dates as 6.60×10^{-05}. The test statistic is almost zero and obviously does not exceed the critical test statistic value. Hence, the null hypothesis of cointegration (given the estimated structural breaks) cannot be rejected. Therefore, the evidence of cointegration stays robust controlling for structural breaks.

Pre- and Post Break Analysis

Having detected the break point dates, we check for cointegration in *the pre- and post break samples*. Because Johansen method can give unreliable results in small samples (Stock and Watson, 1993), we use the Engle-Granger (1987) technique. Table 7 shows clear evidence of cointegration in all three subsamples. We then proceed by estimating the cointegrating parameters of export and import models to see the variation in trade deficit over time. We run DOLS regressions for each period and employ Wald test to check whether the slope coefficient significantly differs from unity. Results are provided in Table 8. In the first subsample (1989:01-2001:01), the slope coefficient is 1.43 and the null hypothesis that it equals one cannot be rejected. Therefore, from 1989 to 2001, Turkish trade balance is sustainable.

In the second subsample (2001:01-2008:09), the slope decreases to 1.40, but the restriction

Table 8: Pre- and Post-Break Sample Analysis

Estimated DOLS Equation:	*Imports*=C(1)**Exports*	+C(2)+C(3)*@TREND
Sample	DOLS slope estimates (SE)	Wald Test *t*-statistic (*p*-value)
1989 : 01 – 2001 : 01	1.427(0.469)***	0.911(0.3641)
2001 : 01 – 2008 : 09	1.402(0.109)***	3.673(0.0004)***
2008 : 09 – 2011 : 12	2.294(0.182)***	7.124(0.0000)***

Notes: *, **, and *** marks denote significance at 1%, 5%, and 10% level. Each DOLS regression is run with a constant and a trend and includes one lead and one lag. *Imports* are dependent and *Exports* independent variables.

that it equals one is rejected. Hence, the slope is significantly different from unity from 2001 to 2008, showing that trade deficit is no longer sustainable. However, it is noteworthy to mention that for this period, the coefficient on exports in the DOLS regression decreases. This shift is probably due to the economic and regulatory reforms in the early 2000s aimed at promoting export-led growth.

On the other hand, in the third subsample (2008:09-2011:12), the slope coefficient jumps to 2.3. The null hypothesis that it equals one is rejected again, showing that trade deficit is not sustainable. Hence, since 2008, Turkey imports more than twice than it exports. This shift can be related to the decrease in Turkey's exports with the fall in global demand, and suggests that Turkey's import-export relationship is highly vulnerable to global shocks. Moreover, the findings indicate that since 2008, Turkey has dramatically moved away from the condition of trade sustainability.

Finally, we re-estimate the DOLS equations by regressing exports on imports and find that import coefficients are much less than the export coefficients given in Table 9.[4] Hence imports grow much faster than exports. If this trend is due to domestic production dependency for intermediate goods imports, we may expect deficit to continue to expand even if external demand corrects.

[4]The results are not reported for sake of concision but are available upon request.

VI Conclusion

In this paper, we analyse the sustainability of the trade deficit of Turkey during 1989-2011. Following Husted (1992) model, we use the classical cointegration tests of Engle-Granger (1987) and Johansen (1990), as well as the technique of Sylvestre and Sansó (2006) which allows for structural breaks. We find that exports and imports have a long-run stable relationship from 1989 to 2011, no matter the analysis accounts for structural breaks. This means that Turkey fulfills the primary condition (weak form) of deficit sustainability, and macroeconomic policies are successful in that regard. However, the relationship between exports and imports is far from unity, and deficit is widening and remains persistent. The sub-sample analysis shows that since 2001, Turkish trade deficit has not been sustainable in the strong form. Moreover, the situation has significantly worsened after the 2008 financial crisis, with Turkey moving away from sustainability.

We conclude that the sustainability of the Turkish trade deficit is highly doubtful. This conclusion is consistent with the results of Göktas, Tunali and Hepsag (2011), who also apply a multiple structural change approach in cointegration models. The main implication is that Turkey should review its fiscal and monetary policies in order to address its trade imbalances. Specifically, Turkey should take steps in order to make its production less dependent on foreign supplies, and it must strengthen its productivity and export performance.

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