

# Real Exchange Rate and Trade Balance in Turkey: Evidence from Heterogeneous Panel Data

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**Summary:** In this study, the effect of real exchange rate on bilateral trade balance between Turkey and its 25 main trade partners is investigated for the period of 1996 – 2015 with heterogeneous panel data techniques. Trade balance model is estimated by using Mean Group (MG) estimator, which allows parameter heterogeneity, Common Correlated Effects Mean Group (CCEMG), and Augmented Mean Group (AMG) estimators, which both allow cross-section dependency and heterogeneity. Results indicate that the real exchange rate elasticity of the trade balance ranges between -0.40 and -0.45 and Marshall-Lerner (ML) condition is valid for Turkey. According to the results, the foreign income elasticity of trade balance ranges between 1.54 and 2.84, while for domestic income elasticity, it is found between -0.75 and -1.38. Country-specific results show that ML condition is valid for the USA, Belgium, Spain, Switzerland, Romania, and Russia at the bilateral level according to both CCEMG and AMG estimators.

**Keywords:** Marshall-Lerner condition, panel data, cross-section dependence, mean group estimators

**JEL Codes:** C23, F14, F31

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Turkey has significantly transformed its economy and moved to export-oriented growth model with the help of reforms and policies implemented since 1980. Although macroeconomic crises and political instabilities slowed this outward looking reform process in the 1990s, the Customs Union with the European Union (EU) in 1995 has underpinned Turkey's integration into the European economy. After suffering a major economic crisis mainly stemming from the unsustainable public debt in the early 2000s, Turkey has experienced stable growth rates and strong export performance partly owing to the comprehensive reform program following to the crisis. This export performance, however, has also led to a rapid import growth and the foreign trade deficit has risen significantly in recent years. The high level of foreign trade deficit makes exchange rate an important policy tool for Turkey. However, the effect of exchange rates on the trade balance is controversial in the literature.

This study aims to test whether Marshall-Lerner (ML) condition holds for Turkey at bilateral level after the Customs Union. To this end, the relationship between bilateral real exchange rates and bilateral trade balance is investigated for Turkey's 25 main trade partners in the period of 1996:Q1 – 2015:Q2. Mean Group (MG), Common Correlated Effects Mean Group (CCEMG), and Augmented Mean Group (AMG) estimators are used to estimate the price and income elasticities of trade balance. All of these three techniques allow heterogeneous parameter estimates among Turkey's trade partners. Furthermore, the last two relax the cross-section independence assumption underlying traditional panel data estimators.

The rest of the paper is organized as follows: Section 1 summarizes the theoretical and empirical literature; Section 2 describes the model, data, and empirical strategy; Section 3 discusses empirical results, and the last section concludes the findings of this study.

## **1. Theoretical Background and Empirical Literature**

Theoretical foundations of Marshall-Lerner (ML) condition are dated back to C. F. Bickerdike (1920), Joan Robinson (1947) and Lloyd Metzler's (1948) studies. According to Bickerdike-Robinson-Metzler (BRM) approach, a depreciation or devaluation of local currency makes imports expensive and exports cheaper by changing relative prices of tradable goods. Thus, the demand for imported products decreases in domestic markets while the demand for exported products increases in international markets. ML condition suggests that under the assumption that elasticities of export and import supply are infinite, if the absolute sum of the price elasticities of export and import is greater than one, the depreciation of local currency

improves the trade balance (Robert Stern 1973). This can be shown mathematically with the help of the export and import demand functions that have been used in earlier studies by David M. Gomez and Guadalupe F. Álvarez-Ude (2006), in which:

$$X_t = \left(\frac{P}{P^*E}\right)_t^\eta (Y_t^*)^\varepsilon \quad (1)$$

$$M_t = \left(\frac{P^*E}{P}\right)_t^\gamma (Y_t)^\Pi \quad (2)$$

where X and M are the exports and imports, E is the nominal exchange rate, Y and Y\* are domestic and foreign incomes, and P and P\* are domestic and foreign price levels, respectively.  $\eta$  and  $\gamma$  indicate the real exchange rate elasticities of exports and imports, respectively, while  $\varepsilon$  and  $\Pi$  stand for income elasticities. By using logarithmic transformation, equations (1) and (2) become:

$$\ln X_t = \eta[\ln P_t - \ln P_t^* - \ln E_t] + \varepsilon \ln Y_t^* \quad (3)$$

$$\ln M_t = \gamma[\ln P_t^* + \ln E_t - \ln P_t] + \Pi \ln Y_t \quad (4)$$

where  $[\ln P_t - \ln P_t^* - \ln E_t]$  is the natural logarithm of real exchange rate ( $\ln RER_t$ ). The ratio of exports to imports is defined as the trade balance, which can be written as:

$$\ln TB_t = \varepsilon \ln Y_t^* - \Pi \ln Y_t + \theta \ln RER_t \quad (5)$$

where  $\theta$  shows the real exchange rate elasticity of trade balance:  $\theta = (\eta + \gamma - 1)$ . According to equation (5), the ML condition holds if  $(\eta + \gamma) > 1$ .

The positive effect of exchange rate depreciation on trade balance may not come into effect in the short-run. On the contrary, depreciation initially worsens the trade balance while the improvement impact realizes after a while. Since the graphical representation of the relationship between exchange rate and trade balance resembles letter “J”, this phenomenon is labeled as “J-curve” (Stephen Magee 1973). In the short-run, since the exchange rate elasticities of exports and imports are relatively small due to both supply and demand side reasons, J-curve effect may be observed. Exporters and importers may not quickly respond to relative price changes because of the consumption habits and contracts in the short-run. In the long-run, however, these barriers lessen, economic agents accommodate new prices, thus, the depreciation of exchange rates improves the trade balance. Helen B. Junz and Rudolph R. Rhomberg (1973) identified five reasons that

cause J-curve effect including *recognition, decision, delivery, replacement* and *production* lags and they also argue that the complete impact of relative prices on trade flows realizes about four or five years.

This theoretical framework has been tested many times for Turkey and the great majority of these studies used aggregate trade data of Turkey to estimate trade elasticities. However, since the dynamics of trade differs structurally among Turkey's trade partners at the bilateral level, studies using aggregate data ignore this kind of country specific heterogeneities. In recent years, there is a growing literature that investigates the ML condition by disaggregating Turkey's trade flows between her trade partners. The disaggregated approach began with a study by Andrew K. Rose and Janet L. Yellen (1989) in which they investigated the J-curve effect between the US and its six major trading partners. They suggested that when depreciation occurs in a country's exchange rates, the impact on trade balance may vary considerably with its trade partners. In other words, exchange rate depreciation may improve the trade balance with one country, but at the same time, it may worsen the trade balance with another country (Ferda Halicioğlu 2008). Using aggregate data to measure ML condition or J-curve effect suppress these kinds of trading partner specific heterogeneities (Mohsen Bahmani-Oskooee and Gour Goswami 2004; Muhittin Kaplan and Huseyin Kalyoncu 2011). The disaggregated approach provides researchers more detailed analysis since it can give trade partner specific results for a country's trade elasticities.

Abdelhak Senhadji (1998) investigated import elasticities of 77 countries along with Turkey in the period of 1960-1993. By using OLS and Fully Modified OLS techniques, he concluded that income elasticity of import is greater than the exchange rate elasticity in the long run. Claudio Montenegro and Senhadji (1998) performed the same analysis for export elasticities of 75 countries in the same period and found that exchange rate elasticity of Turkey is more determinative than its income elasticity compared to other countries. M. Doganlar, Harun Bal, and Mehmet Ozmen (2004) and M. Faruk Aydin, Hulya Saygili, and Mesut Saygili (2007) found the high price elasticities of Turkish exports. Saygili and Saygili (2011), on the other hand, found that the importance of real exchange rate elasticity of Turkish exports is decreasing as Turkey integrates into the global production networks. Some studies investigate the ML condition by using trade balance model. In this regard, Rahmi Yamak and Abdurrahman Korkmaz (2005), Murat Karagoz and Cetin Dogan (2005), Osman Peker (2008), Cengiz Aktas (2010) and Celal Kizildere, Burhan Kabadayi, and O. Selcuk Emsen (2014) did not find enough evidence for ML condition for Turkey in the long run,

while Elif Akbostanci (2004) and Mehtap Aksu (2007) confirmed the validity of ML condition hypothesis for Turkey.

Abovementioned studies used aggregate trade data of Turkey. There are also studies that investigated the elasticities of Turkish trade by disaggregating the data across countries. Halicioğlu (2008) applied ARDL technique to estimate the bilateral trade elasticities between Turkey and 13 countries by using quarterly data from 1987 to 2005 and did not find J-curve effect in the short-run. In the long-run, however, ML condition is found valid for Turkey's bilateral trade with the USA and the UK. Mehmet Yazici and M. Qamarul Islam (2014) performed the same technique to Turkey's trade with EU15 countries in the 1982-2001 period, and found evidence in favor of ML condition in cases of Austria, Denmark, France, Ireland, Italy, Sweden, and the UK. Yazici and Islam (2011) also investigated the effect of exchange rate on trade balance between Turkey and EU15 for 21 different sectors. Short-run results of their analysis indicate that the effect of exchange rate on the trade balance is found significant for 13 sectors, while in the long-run there is no significant effect of exchange rate on the trade balance.

Hassan Kimbugwe (2006) examined the relationship between exchange rates and bilateral trade balances of Turkey and its nine main trade partners in the 1960-2000 period with Johansen-Juselius cointegration analysis and impulse-response functions. Results show that the cointegration relationship exists between exchange rates and bilateral trade balances in the long run, although J-curve effect is not observed in the short-run. Kaplan and Kalyoncu (2011) investigated the trade elasticities between Turkey and 12 trade partners in the 1987-2005 period, and found that devaluation has a positive effect on Turkey's trade balance although the magnitude of this effect differs by trade partners. Aydin, Yusuf Soner Baskaya, and Ufuk Demiroğlu (2015) performed panel regression to calculate the elasticities for Turkey and 91 of her trade partners in 1995-2012 and concluded that ML condition is valid.

In addition, some studies estimate the export elasticities at the bilateral level. In this regard, Evren Erdogan Cosar (2002) found that foreign income elasticity is greater than the real exchange rate elasticity for Turkey's exports to its six main partners in 1989-2000. Olcay Y. Culha and Mustafa Koray Kalafatçılar (2014) estimated the regional export elasticities of Turkey and concluded that income elasticities are greater for the exports to advanced countries compared to the emerging countries. Price elasticities, on the other hand, gradually increased for the exports to Middle East and North Africa (MENA) region according to their study. Finally, Ihsan Bozok, Bahar Sen Dogan, and Caglar Yunculer (2015) estimated the export elasticities of Turkey by using Dynamic OLS, MG and CCEMG estimators for the period

of 2005-2013. Their results confirmed the fact that exchange rate elasticity of exports is greater than the foreign income elasticity.

## 2. Data and Methodology

### 2.1. Model and Data

As we have shown in the previous section, our model based on Gomez and Ude (2006) can be presented in equation (6).

$$\ln TB_{it} = \beta_{1i} + \beta_{2i} \ln TRGDP_t + \beta_{3i} \ln GDP_{it} + \beta_{4i} \ln RER_{it} + \varepsilon_{it} \quad (6)$$

In the above equation,  $\ln TB_{it}$  indicates the trade balance between Turkey and  $i^{\text{th}}$  country, illustrated as the ratio of exports to imports.  $\ln TRGDP_t$  and  $\ln GDP_{it}$  stand for the real gross domestic product of Turkey and  $i^{\text{th}}$  country, respectively. Finally,  $\ln RER_{it}$  shows the bilateral real exchange rates between Turkey and  $i^{\text{th}}$  country. Bilateral real exchange rates are calculated using the following formula:

$$RER_{it} = \frac{P_t}{P_{it}^* E_{it}} \quad (7)$$

where  $P_t$  and  $P_{it}^*$  indicate the consumer prices index of Turkey and  $i^{\text{th}}$  country, respectively, while  $E_{it}$  shows the bilateral nominal exchange rates between Turkey and  $i^{\text{th}}$  country. According to this exhibition, an increase (decrease) in  $RER$  signifies the appreciation (depreciation) of the local currency; it means that prices of domestic goods increase (decrease) compared to the foreign goods (Yahya Kocakale and Hakan H. Toprak 2015).

As for the expected signs of trade balance model, the  $\beta_{2i}$  coefficient is expected to be negative if the increase in domestic income is accompanied with an increase in imports. However, it would be positive if the increase in income stems from the increase in the production of import-substitute goods (Swarnjit Arora, Mohsen Bahmani-Oskooee, and Gour Goswami 2003). Considering the high economic growth rates in Turkey and strong import growth brought with it especially after 2001 crisis, we expect  $\beta_{2i}$  coefficient to be negative. Similarly,  $\beta_{3i}$  coefficient, which shows the foreign income elasticity of trade balance, could be either positive or negative. For Turkey, directing nearly half of its exports to advanced EU countries, we expect this coefficient to be positive. Finally, we expect  $\beta_{4i}$  coefficient to be negative assuming that depreciation of local currency – decrease in real exchange

rates – would improve the trade balance, which is evidence of ML condition is holding. In addition, it should not be forgotten that these coefficients may differ across Turkey's trade partners.

The data in this study covers the quarterly data in the period of 1996:Q1 – 2015:Q2 and contains the 25 main trade partners of Turkey having over 1 percent share in Turkey's total trade in 2014. Bilateral trade data is from the Turkish Statistical Institute (TURKSTAT). Bilateral nominal exchange rates are retrieved from Bloomberg with direct quotation method. Consumer price indices of countries are derived from the International Monetary Fund (IMF) – International Financial Statistics (IFS) database. Seasonally adjusted real gross domestic product data of countries are compiled from the World Bank and Oxford Economics databases. Finally, in order to control the effects of crises, which Turkey faced in 2002 and 2008 on trade balance, we included dummy variables that correspond to these years. All data are transformed into indices reflecting 2010=100 and are measured in logarithms. Model estimations and econometric tests are performed by using Gauss 10 and Stata 12 statistical software programs.

## **2.2. Preliminary Analysis**

### *Cross-Section Dependence Tests*

Cross-section dependence in panel data can occur from the unobservable common factors or spillover effects across the panel units (Badi Baltagi and M. Hashem Pesaran 2007). Especially for the macro panel data sets containing countries which that are highly dependent on each other, global economic shocks or spillovers could affect countries together and cause cross-section dependence problem. Ignoring this problem may lead to biased and inconsistent results for both unit root tests and estimators (Vasilis Sarafidis and Donald Robertson 2009). Therefore, it is important to identify if the series are cross-sectionally dependent before performing unit root test. Pesaran (2004) CD test and Pesaran, Aman Ullah, and Takashi Yamagata (2008) bias-adjusted LM test are used to check the cross-sectional dependency in this study. Pesaran (2004) CD test is robust to non-normality of errors, non-stationarity, structural breaks and coefficient heterogeneity. However, it is best suitable when the cross-sectional dimension (N) is greater than the time dimension (T) in the panel. Bias-adjusted LM test, on the other hand, produces consistent results even if CD test is inconsistent (Pesaran, Ullah, and Yamagata 2008). Therefore, we used the bias-adjusted LM test to support the CD test's results and to reach more robust inference.

CD test and bias-adjusted LM test statistics can be calculated respectively as (8) and (9):

$$CD = \sqrt{\frac{2T}{N(N-1)}} \left( \sum_{i=1}^{N-1} \sum_{j=i+1}^N \hat{\rho}_{ij}^2 \right) \quad (8)$$

$$LM^* = \sqrt{\frac{2T}{N(N-1)}} \left( \sum_{i=1}^{N-1} \sum_{j=i+1}^N \hat{\rho}_{ij}^2 \right) \frac{(T-k)\hat{\rho}_{ij}^2 - E(T-k)\hat{\rho}_{ij}^2}{\text{Var}(T-k)\hat{\rho}_{ij}^2} \quad (9)$$

where  $T$  and  $N$  show the time and cross-section dimensions of panel data, respectively, while  $\hat{\rho}_{ij}$  reflects pairwise correlation of residuals estimated by ordinary least squares (OLS) technique. The results of the CD and bias-adjusted LM test presented in Table 1 indicate that we can reject the null hypothesis of cross-section independence at 1 percent significance level.

### Slope Homogeneity Tests

Traditional panel data estimators do not allow the cross-sectional heterogeneity or allow only heterogeneity of the intercepts across panel units. However, in macro panels containing countries structurally different from each other and having large  $N$  and large  $T$  dimensions, allowing only different intercepts do not properly reflect the heterogeneity across panel units (Pesaran and Ron Smith 1995; Kyong So Im, Pesaran and Yongcheol Shin 2003). In this study, we assumed trade elasticities could be different between Turkey and its trade partners at bilateral level due to the dynamics and motives of foreign trade being different structurally across countries. This assumption, i.e. the heterogeneity of coefficients, is tested with ‘‘Delta Test’’ presented by Pesaran and Yamagata (2008) as an alternative to Paravastu A. Swamy’s (1970) slope homogeneity test. Swamy test statistic, Delta test statistic and mean and variance bias adjusted version of Delta test statistic can be calculated as follows:

$$\hat{S} = \sum_{i=1}^N (\hat{\beta}_i - \hat{\beta}_{WFE})' \frac{\bar{X}'\bar{X}}{\hat{\sigma}_i^2} (\hat{\beta}_i - \hat{\beta}_{WFE}) \quad (10)$$

$$\tilde{\Delta} = \sqrt{N} \frac{N^{-1}\hat{S} - k}{\sqrt{2k}} \quad (11)$$

$$\tilde{\Delta}_{adj} = \sqrt{N} \frac{N^{-1}\hat{S} - E(\hat{Z}_{it})}{\sqrt{\text{Var}(Z_{it})}} \quad (12)$$

where  $\hat{\beta}_i$  is heterogeneous coefficients obtained by OLS estimation of equation (1) for each cross section.  $\hat{\beta}_{WFE}$  shows the coefficient of weighted fixed estimators (the weights are constructed using  $\hat{\sigma}_i^2$ ) and  $\bar{X}$  is the matrix of the deviations from the mean of explanatory variables. The estimated test statistics and corresponding p-values are presented in Table 1. Results

indicate that the null hypothesis of homogeneous slopes must be rejected and suggest that it is important to consider the slope heterogeneity.

**Table 1** Cross-Section Dependence and Slope Homogeneity Test Results

Test	Test Stat.	p-value	Null Hypothesis
CD-Test	7.98	0.00	Cross-section independence
LM*	197.40	0.00	$\hat{\rho}_{ij} = cor(u_{it}, u_{jt}) = 0$
$\tilde{\Delta}$	38.69	0.00	Homogeneous slopes
$\tilde{\Delta}_{adj}$	39.97	0.00	$\beta_1 = \beta_2 = \dots = \beta_n = \beta$

**Notes:** 1) First two rows show Pesaran (2004) and Pesaran, Ullah and Yamagata (2008) cross-section dependence tests, respectively. The last two denotes the Delta and Delta adjusted slope homogeneity test suggested by Pesaran and Yamagata (2008). 2) Since slope homogeneity tests require balanced panel, we excluded data of Egypt, Iran, Iraq, Saudi Arabia, UAE and Ukraine which make panel unbalanced and performed these tests by using data of other 19 countries.

**Source:** Author's calculations.

### Panel Unit Root Tests

In panel data literature, testing stationarity of series is categorized into two groups. The so-called first generation panel unit root tests such as G. S. Maddala and Shaowen Wu (1999), Kaddour Hadri (2000), Andrew Levin, Chien-Fu Lin, and Chia-Shang James Chu (2002) and Im, Pesaran and Shin (2003) are considered insufficient since they do not take into account the cross-section dependence across panel units, thus producing biased results (Baltagi and Pesaran 2007). Second generation panel unit root tests, on the other hand, allow the cross-section dependency of panel units. Since we have identified both cross-section dependence and slope heterogeneity in our data, we employed the cross-sectionally augmented panel unit root test (CIPS test) of Pesaran (2007). This test incorporates the cross-section mean of lagged levels and first differences of the individual series into the standard ADF regression to relieve the cross-section dependence problem. The cross-sectionally augmented ADF equation is given by:

$$\Delta Y_{it} = \alpha_i + \rho_i Y_{it-1} + d_0 \bar{Y}_{t-1} + \sum_{j=0}^p d_{j+1} \Delta \bar{Y}_{t-j} + \sum_{k=1}^p c_k \Delta Y_{i,t-k} + \varepsilon_{it} \quad (13)$$

where  $\alpha_i$  is a deterministic term,  $\bar{Y}_{t-1}$  is the cross-section mean of lagged variable,  $\Delta Y_{i,t-k}$  and  $\Delta \bar{Y}_{t-j}$  are augmenting terms of ADF equation and  $p$  is the lag order. Pesaran (2007) CIPS statistic is calculated by firstly running

equation (13) for each cross-section unit and then averaging t-stat of  $\rho_i$  from all of the individual regressions:

$$CIPS = N^{-1} \sum_{i=1}^N t_{p_i} \quad (14)$$

In addition to the stationarity of panel variables, we also checked the stationarity of cross-sectionally invariant *TRGDP* variable by using standard ADF test. Tables 2 and 3 present the results of both tests. Accordingly, *TB* and *RER* variables are found stationary in their level values, while *GDP* and *TRGDP* become stationary after taking first differences. So, the variables are integrated of different orders.

**Table 2** Pesaran (2007) CIPS Test Results

Lag	TB				GDP				RER			
	Intercept		Int.+Trend		Intercept		Int.+Trend		Intercept		Int.+Trend	
	Zt-bar	p-val.	Zt-bar	p-val.	Zt-bar	p-val.	Zt-bar	p-val.	Zt-bar	p-val.	Zt-bar	p-val.
0	-8.5	0.00	-9.5	0.00	-0.6	0.27	1.6	0.94	-10.3	0.00	-9.4	0.00
1	-4.5	0.00	-3.6	0.00	-1.9	0.03	1.2	0.89	-5.5	0.00	-4.7	0.00
2	-2.0	0.02	-0.6	0.27	-1.5	0.07	2.0	0.98	-5.1	0.00	-4.0	0.00
3	-1.2	0.11	0.7	0.75	-2.0	0.02	0.3	0.63	-4.3	0.00	-3.6	0.00
4	-1.3	0.09	0.6	0.72	-1.1	0.13	1.3	0.91	-2.4	0.01	-0.7	0.23

**Note:** Zt-bar is the test statistic for Pesaran (2007) CIPS test. Ho: Series are I(1).

**Source:** Author's calculations.

**Table 3** ADF Unit Root Test Results for *TRGDP*

Specification	Test Statistics	p-values
Intercept	-0.198	0.933
Intercept + Trend	-2.844	0.187

Ho: Series are I(1).

**Source:** Author's calculations.

### Panel Cointegration Tests

Similar to the panel unit root tests, panel cointegration tests are also separated into two groups as first and second generation cointegration tests with regard to whether it allows the cross-section dependence across panel units or not. We performed the second generation Joakim Westerlund (2007) and Westerlund (2008) Durbin-Hausman cointegration tests that yield robust results for both heterogeneous coefficients and cross-section dependence. Westerlund (2007) calculates four different statistics based on the error-correction model across panel units and panel as a whole.  $G_t$  and  $G_a$  indicate cointegration relationship for at least one of the cross-section unit, while  $P_t$

and  $P_a$  show this relationship for whole panel (Damiaan Persyn and Westerlund, 2008).

Although Westerlund (2007) panel cointegration test allows heterogeneity and cross-section dependence across panel units, it assumes that all variables used in the panel are integrated of order one. Durbin-Hausman test, on the other hand, calculates two statistics as  $DH_p$  for the panel and  $DH_g$  for cross-section units depending on the similar methodology and can be used even if the variables are integrated of different order (Westerlund, 2008). Since we have found that our variables are integrated of different order according to our panel unit root test results, we performed both tests to check the cointegration relationship among variables for robustness. Both test results presented in Table 4 reject the null hypothesis of no cointegration for cross-section units and panel as a whole, which means that there is a stable and meaningful long-run relationship between these variables.

**Table 4** Panel Cointegration Test Results

Statistic	Value	Z-value	p-value	Rob. p-value
Gt	-3.266	-4.776	0.000	0.000
Ga	-21.432	-6.485	0.000	0.010
Pt	-12.093	-3.522	0.000	0.030
$P_a$	-16.087	-5.735	0.000	0.100
$DH_g$	2514.76	-	0.000	-
$DH_p$	755.16	-	0.000	-

**Notes:** 1) Average lag lengths are selected as 3 by the Akaike Information Criteria (AIC). 2) Since panel cointegration tests require balanced panel, we excluded the data of Egypt, Iran, Iraq, Saudi Arabia, UAE and Ukraine which make panel unbalanced and performed these tests by using data of other 19 countries.  $H_0$ : No cointegration.

**Source:** Author’s calculations.

### 2.3. Estimation Methodology

As we have shown in the preliminary analysis, we have both cross-section dependence and slope heterogeneity in our data. Therefore, instead of using traditional panel data estimators, which do not allow slope heterogeneity and cross-section dependence across panel units, we use three heterogeneous panel data estimators to estimate the trade elasticities given in equation (6). Firstly, we use Mean Group (MG) Estimator that allows the heterogeneous coefficients across panel units developed by Pesaran and Smith (1995). Subsequently, to control the cross-section dependence problem, we perform Pesaran (2006) Common Correlated Effects Mean Group (CCEMG) and Markus Eberhardt and Stephen Bond (2009) and Eberhardt and Francis Teal

(2010) Augmented Mean Group (AMG) estimators. These estimators can be explained briefly with the help of equation (15). For  $i = 1, 2 \dots N$  and  $t = 1, 2 \dots T$ ;

$$\begin{aligned} Y_{it} &= \alpha_i d_t + \beta_i X_{it} + u_{it} \\ u_{it} &= \lambda_i f_t + \varepsilon_{it} \end{aligned} \quad (15)$$

where  $Y_{it}$  ve  $X_{it}$  are dependent and independent variables, respectively,  $\beta_i$  is country-specific coefficients and  $u_{it}$  is the sum of unexplained effects and error term ( $\varepsilon_{it}$ ). In the above equation,  $d_t$  represents time-invariant observed common effects across panel units while  $f_t$  stands for the unobserved common effects capturing the time-variant heterogeneity and cross-section dependence across panel units.  $\alpha_i$  and  $\lambda_i$  are the heterogeneous factor loadings for  $d_t$  and  $f_t$  respectively. For all of the estimators performed in this study, the estimation procedure is same. As a first step, the model is estimated with OLS technique for each panel members. Afterwards, the estimated coefficients are averaged to reach the panel coefficients (Eberhardt 2012).

#### Mean Group (MG) Estimator

Pesaran and Smith (1995) found that when  $T$  dimension is large and model coefficients differ across panel units, traditional panel data estimators such as fixed effect and random effect regressions or cross-section regressions of averages over time and time-series regressions of group averages yield inconsistent estimations (Baltagi 2005:224). They showed that when  $T$  is large enough, instead of carrying out traditional estimations, performing OLS estimations for each panel members and then averaging the estimated coefficients yield consistent estimations. So, the MG estimator estimates equation (15) for each panel member and considers the heterogeneity across panel units. On the other hand, it either rules out the  $\lambda_i f_t$  term which indicates the possible cross-section dependence in equation (15) or includes model as a linear trend.

#### Common Correlated Effects Mean Group (CCEMG) Estimator

Pesaran (2006) solves the possible cross-section dependence problem imposed by the  $\lambda_i f_t$  term in equation (15) by including cross-section averages of dependent and independent variables ( $\bar{Y}$  and  $\bar{X}$ ) into the model.  $\bar{Y}$  and  $\bar{X}$  together constitute the  $f_t$  variable indicating unobserved common factor. Thus, CCEMG transforms equation (15) into the following model and estimate this for each panel member:

$$Y_{it} = \alpha_{1i} + \beta_i X_{it} + \gamma_{1i} \bar{Y}_t + \gamma_{2i} \bar{X}_t + \varepsilon_{it} \quad (16)$$

### Augmented Mean Group (AMG) Estimator

AMG estimator, which also takes into account the coefficient heterogeneity and cross-section dependence, is developed by Eberhardt and Bond (2009) and Eberhardt and Teal (2010) as an alternative to CCEMG. The main difference between AMG and CCEMG is the procedure of including unobserved common factor ( $\lambda_i f_t$ ) into the model. The AMG estimator follows a two-stage procedure to include the unobserved common factor into the regression. In the first stage, equation (15) is augmented with time dummies as depicted in equation (17) and estimated with first differenced OLS. Then the coefficients of time dummies are collected.

$$\Delta Y_{it} = \alpha_{1i} + \beta_i \Delta X_{it} + \sum_{t=2}^T \gamma_t DUMMY_t + \varepsilon_{it} \quad (17)$$

In the second stage, the unobserved common factor ( $f_t$ ) is replaced with the coefficients of time dummies ( $\hat{\gamma}_t$ ) and transformed model in equation (18) is estimated by OLS for each panel member.

$$Y_{it} = \alpha_i d_t + \beta_i X_{it} + \lambda_i \hat{\gamma}_t + \varepsilon_{it} \quad (18)$$

To reach the panel coefficients, as mentioned above, all of these estimators take the simple arithmetic mean of estimated coefficients for each panel members. For  $\tilde{\beta}_i$ ,  $\hat{\beta}_i$  and  $\check{\beta}_i$  are OLS estimates of coefficients in equations (15), (16) and (18) respectively;

$$MG = N^{-1} \sum_{i=1}^N \tilde{\beta}_i \quad (19)$$

$$CCEMG = N^{-1} \sum_{i=1}^N \hat{\beta}_i \quad (20)$$

$$AMG = N^{-1} \sum_{i=1}^N \check{\beta}_i \quad (21)$$

### **3. Empirical Results**

The estimation results for the whole panel are presented in Table 5. As we can see from the table, the real exchange rate elasticity of the trade balance is found statistically significant at 1 percent level and ranges between -0.40 and -0.45 by different estimation methods. Accordingly, 1 percent increase in real exchange rate – in other words, 1 percent appreciation of Turkish Lira – worsens the trade balance of Turkey by about 0.4 percent. So, according to three different heterogeneous panel estimation methods, ML condition is valid for Turkey as of its main trade partners in 1996:Q1 – 2015:Q2 period.

**Table 5** Panel Estimation Results

	<b>MG</b>	<b>CCEMG</b>	<b>AMG</b>
<b>TRGDP</b>	-1.384*** (0.376)	-0.120 (0.450)	-0.754* (0.438)
<b>GDP</b>	2.776*** (0.792)	2.836*** (0.684)	1.544*** (0.532)
<b>RER</b>	-0.446*** (0.104)	-0.405*** (0.138)	-0.438*** (0.100)
<b>Constant</b>	0.201 (3.331)	0.061 (3.983)	2.785 (2.300)
<b>Observation</b>	1793		
<b>Countries</b>	25		
<b>Period</b>	1996:Q1 - 2015:Q2		

**Notes:** Dependent variable is TB in all estimations. Standard errors are in parenthesis. Dummy variables representing 2001 and 2008 crisis are found insignificant in all estimations. \*\*\* $p < 0.01$ , \*\* $p < 0.05$ , \* $p < 0.1$

**Source:** Author's calculations.

In line with our expectations, *GDP* coefficient showing the income of Turkey's trade partners is found statistically significant at 1 percent level and ranges between 1.54 and 2.84 by different methods. Specifically, 1 percent increase in income of Turkey's trade partners improves the trade balance of Turkey by about 1.54 – 2.84 percent according to our results. As for the cross-section invariant *TRGDP* coefficient showing the Turkey's domestic income, the significance of coefficient differs in terms of estimation methods. While results indicate that the domestic income elasticity of trade balance ranges between 0.75 and 1.38, the significance levels are 1 percent in MG estimation, 10 percent in AMG estimation and finally not significant in CCEMG estimation.

Compared to the estimated coefficients, we see that the absolute value of real exchange rate elasticity of trade balance is lower than foreign income elasticities. This phenomenon implies that the effect of exchange rate policy is low for Turkey in terms of policies to be implemented to improve the trade balance. This result is not surprising for such a country that transforms its export composition from low value-added products such as textiles, apparels, and agricultural products to medium-high value-added products such as motor vehicles and machinery especially after the Customs Union. As Peter Wierts, Henk van Kerkhoff, and Jakob de Haan (2012) pointed out, price competition is important for the exports of low-medium technological products while factors such as quality and design are more determinant for the exports of medium-high technological products. The importance of exchange rate as a determinant of foreign trade is gradually decreasing for

Turkey which undergoes a significant structural transformation in exports after its entrance into Customs Union. Indeed, early studies like Montenegro and Senhadji (1998) found that the exchange rate elasticity is greater than the foreign income elasticity in Turkey's export function covering the period of 1960-1993. On the other hand, recent studies investigating this issue like Saygili and Saygili (2011) and Bozok, Dogan, and Yunculer (2015) found low elasticities of exchange rates on Turkish exports.

In addition, there is an ongoing debate in the literature on whether higher levels of integration between countries result in a decrease in the exchange rate elasticity of exports. Swarnali Ahmed, Maximiliano Appendino, and Michelle Ruta (2015) identified that the exchange rate elasticity is gradually decreasing in time and asserted that global value chains (GVCs) explain about 40 percent of the decreasing trend in exchange rate elasticity, especially in the manufacturing sector. In their study, the effect of exchange rates on exports is found lower for the countries with higher backward linkages (foreign value added embodied in exports) in GVCs. Another explanation to low exchange rate elasticity in our estimation results may be this phenomenon as Turkey has relatively higher backward linkages compared to forward linkages (domestic value added embodied in foreign exports). According to OECD-WTO Trade in Value Added (TIVA) database, backward participation of Turkey in GVCs is 25.7 percent as of total gross exports while forward participation in GVCs is only 15.3 percent in 2011 which is the latest figure. For comparison, backward participation in GVCs is 14.3 percent while forward participation is 19.4 for EU28 average in the same year.

**Table 6** Postestimation Cross-Section Dependence Test Results

Variable	CD-Test	p-value	Corr	Abs(corr)
CCE_res	-0.700	0.486	-0.006	0.158
AMG_res	0.940	0.345	0.006	0.157

**Notes:** *CCE\_res* and *AMG\_res* are residuals from the CCE and AMG regressions respectively. Corr shows the average correlation of residuals, while Abs(corr) denotes the average absolute correlation of the residuals.  $H_0$ : Cross-section independence.  $(\hat{\rho}_{ij} = cor(u_{it}, u_{jt}) = 0)$

**Source:** Author's calculations.

To see whether CCEMG and AMG methods solve the cross-section dependence problem, we collected the residuals in both estimations (*CCE\_res* and *AMG\_res* variables in Table 6) and performed the Pesaran (2004) CD test again. As seen from Table 6, the null hypothesis of cross-

section independence cannot be rejected for both variables, meaning that both CCEMG and AMG estimators annihilate the problem.

Turkey's trade partner specific CCEMG estimation results are presented in Table 7. The real exchange rate elasticities of bilateral trade balances are found statistically significant and negative, indicating the validity of the ML condition for 11 countries (Belgium, Bulgaria, Netherlands, Poland, Romania, Russia, Saudi Arabia, Spain, Switzerland, UAE and the USA) out of 25. Besides, similar to the panel results, here we found that the exchange rate elasticities are lower than the foreign and domestic income elasticities. Only for the Netherlands, Romania, Russia, and Saudi Arabia out of 11 countries that ML conditions hold, the exchange rate elasticities are greater than one according to CCEMG unit-specific estimation results.

Looking at the foreign income elasticities of bilateral trade balances, it is seen that elasticities are found statistically significant and positive for 11 countries (Bulgaria, Egypt, France, Germany, Greece, Iraq, Italy, Poland, Romania, the UK, and the USA) out of 25. Israel, among these countries, is the only one with which Turkey has negative income elasticity at 5 percent statistically significance level. Lastly, domestic income elasticities of bilateral trade balances are found statistically significant for 9 of the studied countries which are negative for 6 of those countries and positive for the other 3 countries. For Poland, Romania and Russia, which we have found positive domestic income elasticity in their trade with Turkey, these results may be interpreted as a sign that the income increase of Turkey is associated with an increase in exports to these countries. For other 6 countries (Bulgaria, China, France, Germany, the UK, and the USA), the results indicate that increase in income of Turkey raise the import demand from these countries, thus bilateral trade balances worsen.

Similar to the CCEMG estimation results, the trade partner specific AMG estimation results presented in Table 7 also show that the elasticities vary across countries. According to the AMG results, for 11 out of 25 countries, ML condition is valid at 5 percent of statistical significance level. Results also show that the exchange rate elasticities are greater than 1 for Egypt and Russia, meaning that Turkey's bilateral trade with these countries is highly influenced by real exchange rate developments. Additionally, for the Belgium, Romania, Russia, Spain, Switzerland, and the USA, bilateral exchange rate elasticities are significant and negative in common according to both CCEMG and AMG results which indicate the robustness of the validity of ML condition.

**Table 7** Country Specific Coefficient Estimates

Countries	CCEMG			AMG		
	TRGDP	GDP	RER	TRGDP	GDP	RER
Belgium	0.04	0.124	-0.404**	0.182	0.611	-0.307**
Bulgaria	-1.760***	2.270***	-0.425***	-2.328***	2.841***	-0.248
China	-6.774***	1.511	0.206	-6.911***	3.416***	-0.204
Egypt	-1.246	4.905**	-1.082	-3.930***	2.270**	-1.602***
France	-1.104***	2.326**	-0.104	-0.448***	3.356***	-0.280**
Germany	-0.910***	3.049***	0.001	-1.207***	2.149***	0.008
Greece	-0.766	4.428***	0.053	-2.151***	3.91***	-0.199
India	0	0.947	0.905	1.787	-1.923**	-0.109
Iran	1.89	0.222	-0.102	1.227*	-1.353*	-0.075
Iraq	-0.523	2.589*	0.028	-1.315	2.378**	-0.011
Israel	0.566	-2.418**	-0.256	0.492	-1.730***	0.420***
Italy	-0.415	3.602***	-0.069	-0.171	4.008***	0.089
Japan	1.065	3.389	-0.247	0.263	3.543	-0.873***
Netherlands	0.029	1.366	-1.063***	-0.141	2.214***	-0.197
Poland	3.951***	2.229*	-0.760*	1.691	-4.239***	-0.546
Romania	1.731***	1.938***	-1.144***	-0.739*	0.876*	-0.813***
Russia	3.308***	-0.229	-1.237***	1.864***	-2.869***	-1.413***
S. Arabia	-4.265	5.31	-2.354*	-3.219	1.203	-0.93
South Korea	0.467	3.627	0.974**	0.491	-0.242	-0.871**
Spain	-0.283	3.66	-0.795***	-0.886***	3.454***	-0.553**
Switzerland	3.188	12.415	-0.855**	2.651	3.814	-0.955***
UAE	0	-0.499	-0.161**	-2.729*	-0.002	0.093
UK	-1.233***	1.349*	-0.209	-0.488***	3.393***	0.089
Ukraine	1.352	0.274	-0.487	0.887***	-0.15	-0.635***
USA	-1.310***	12.518***	-0.532**	-3.731***	7.656***	-0.837***

**Notes:** Dependent variable is TB in both estimations. \*\*\*p<0.01, \*\*p<0.05, \*p<0.1.

**Source:** Author's calculations

As listed in Table 7, foreign income elasticities of trade balances obtained from AMG estimator are significantly positive for 13 countries (Bulgaria, China, Egypt, France, Germany, Greece, Iraq, Italy, Netherlands, Spain, Romania, the UK, and the USA), but significantly negative for 5 countries (India, Iran, Israel, Poland, and Russia). The results for these latter five countries show that these countries shift their imports from Turkey to other countries when their incomes increase. The countries with the highest foreign income elasticities are Greece, Italy, and the USA.

Finally, domestic income elasticities of trade balances are significantly negative for 11 countries (Bulgaria, China, Egypt, France, Germany, Greece, Romania, Spain, UAE, the UK, and the USA), but significantly negative for Iran, Russia and Ukraine according to AMG

estimation results. The countries with the highest domestic income elasticities are China, Egypt, and the USA.

#### **4. Conclusion**

The relationship between exchange rates and the trade balance is considered highly important and has been subjected to empirical investigations. Exchange rate developments may significantly affect the trade balances and external vulnerabilities especially for outward-oriented countries like Turkey. Depreciation of local currency may contribute to the improvement of trade balance of a country by changing relative prices when the ML condition is valid. For this reason, it is important to investigate whether this condition holds when examining foreign trade policies.

In this study, the ML condition between Turkey and its 25 main trade partners is examined with heterogeneous panel data techniques for the period of 1996:Q1 – 2015:Q2. In contrast to the previous studies on this topic, this study takes into account the country-specific heterogeneities and cross-section dependence in panel units when examining the exchange rates – trade balance relationship. To this end, we used three heterogeneous panel regression methods developed in recent years. The commonly used trade balance model is estimated by using Mean Group (MG) estimator, which allows parameter heterogeneity, Common Correlated Effects Mean Group (CCEMG) and Augmented Mean Group (AMG) estimators, which both allow cross-section dependency and parameter heterogeneity.

The results obtained from these three different estimators indicate that the real exchange rate elasticity of trade balance ranges between -0.40 and -0.45 and the Marshall-Lerner (ML) condition is valid for Turkey. According to the results, the foreign income elasticity of trade balance ranges between 1.54 and 2.84 while for the domestic income elasticity, it is found between -0.75 and -1.38. These results are compatible with the trade elasticities obtained in the previous literature and confirm that the exchange rate elasticity is lower and less important than the domestic and foreign income elasticities.

The country-specific results, on the other hand, show that ML condition is valid for Belgium, Romania, Russia, Spain, Switzerland, and the USA at the bilateral level according to both CCEMG and AMG estimators simultaneously. Moreover, the foreign income elasticity of trade balance is found significant and positive for 11 countries in CCEMG and for 12 countries in AMG estimators. In terms of the domestic income elasticity of the trade balance, it is found significant and negative for 6 countries in CCEMG and for 11 countries in AMG estimators.

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