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Abstract

This paper examines the bilateral trade dynamics of the EU with its major trade partners. Previous studies on the bilateral trade dynamics of the EU have been based on estimations without the consideration of the presence of structural breaks. This paper examines the impacts of the real exchange rate and real income on the trade balance of the EU with its major trade partners in the presence of structural breaks. The empirical analysis includes ten major trade partners of the EU for 1980-2012, on a quarterly basis. The paper applies the Bai and Perron (1998) structural break test to determine the presence of structural breaks in series. In order to test the cointegration relationships of series, three different cointegration techniques were applied to the data. First, the Gregory and Hansen (1996) cointegration test was applied, which allows for one structural shift; then, for cases where two breaks were detected, the Hatemi-J (2008) cointegration test was employed. Finally, for countries where more than two breaks are detected, the Maki (2012) cointegration test was applied, which allows for an unknown number of breaks. The parameters of the model were estimated using the Bai and Perron (1998) procedure, which allows for structural breaks, and the OLS procedure without consideration of structural breaks. The paper investigates how the different dynamics of the bilateral trade balance of the EU appear after possible structural breaks consideration.

JEL: F14; F31.

Key Words: Bilateral trade balance, J-curve, cointegration, structural breaks, EU.

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Introduction

The issue of trade elasticities estimation has received a great deal of attention in the literature. Even though many studies are based on aggregate data (Kreinin, 1967; Khan, 1974; Arize, 1987; Bahmani-Oskooee and Niroomand, 1998; Sinha,2001; Imbs and Mejean, 2010). Many studies at the same time admit that trade elasticities on a bilateral basis are more important to and appropriate for develop trade policies and international linkages compared to aggregate trade elasticities (Marquez 1990; Bahmani-Oskooee and Kantipong, 2001; Hatemi-J and Irandoust, 2005; Bahmani-Oskooee and Ratha 2008; Ketenci and Uz 2011; Murad 2012; Marquez 1990). They estimate income and price elasticities for bilateral trade among developed countries, such as Canada, Germany, Japan, the UK, the US, and a group of OPEC countries. Non-OPEC developing countries are also included in estimations, as their roles are becoming increasingly important in international trade. Bahmani-Oskooee and Ratha (2008) assess the impact of currency depreciation on bilateral trade flows of the US with its 19 trading partners that are developed countries. Ketenci and Uz (2011) investigate the dynamics of the bilateral trade of the EU with its major partners, eight countries, and six regions, where developed as well as developing countries are investigated.

The share of developing countries in the world trade has increased significantly in the last decades. The major partners of developed countries frequently do not consist of developed countries, but of developing countries as well. Therefore, studies on the bilateral trade of developed countries very often include developing countries as their major partners. Data on developing countries due to their changing nature very often carry structural breaks. Cointegration relationships among investigated economic variables have structural changes that can be explained by domestic or external shocks, or by political or economic changes. However, lately, not only the economies of developing countries, but those of developed countries as well are more often become influenced by these types of shocks. The consideration of structural breaks in investigation of cointegration relationships and bilateral trade elasticities in export and import demand functions, however, have received little attention in empirical literature. For example, Mah (1993) in his study emphasized the importance of examining the structural stability of Korean import demand. As an example of a developing country, the Korean import demand is highly affected by liberalization measures in the country. Aziz and Li (2007) investigate the reasons for the changing trade elasticities of

China, employing tests to detect structural shifts. The CUSUM and the Bai and Perron (2003) tests failed to detect any breaks in either the export or import demand equations, even though the Chow tests showed a breakpoint in both equations. One of explanations for the failure of the former tests to detect a structural shift is an aggregate data employment that does not have a large abrupt one-time shift. Uz (2010) in her research tests cointegration relationships in the presence of structural breaks in the case of Turkey, employing the Gregory and Hansen (1996) cointegration test. The bilateral trade elasticities the author estimated by using cointegration techniques that do not allow for structural breaks.

There is a limited number of studies related to developed countries and particular the European Union on bilateral trade with consideration of structural breaks. One of reasons for this is the relatively stable economic and political position of the European Union countries in the past. However, lately, the European Union countries have become the subject to a greater degree of influence by external and internal shocks. Furthermore, the largest partners of the EU are developed countries as well as developing countries. Even though developing countries by their nature are sources of economic and political shocks that spread easily to other countries, most of largest trade partners of the EU, despite of their level of development, for a long time or recently, have undergone numerous structural shifts in their economies and have been deeply influenced by external shocks. Therefore, uncounted breaks in cointegrating relationships may bring spurious estimation results.

This study investigates the bilateral trade dynamics of the EU with its major partners. The largest export partners of the EU are the US, China, Switzerland, Russia, Turkey, Japan, Norway, India, Brazil, South Korea; and its largest import partners, China, Russia, the US, Norway, Switzerland, Japan, Turkey, India, Brazil, and South Korea. The share of the investigated ten largest export partners in total EU trade consisted 58.7 percent in 2012, while the share of the ten largest import partners of the EU made up 63.6 percent.[2] The trade of the EU includes the EU15 member countries. The data for selected countries were extracted from the official statistical site of the EU, Eurostat. The quarterly data are used in this research and cover the period from 1980 to 2012.

The novelty of this study is the consideration of structural breaks in the investigation of bilateral trade dynamics. To our knowledge there are no similar studies in the literature. The rest of the paper is organized as follows. In the next section, the applied methodological

[2] European Commision, Chief Economics Statistics Sector.

approach is presented. In section 3, the obtained empirical results are reported, and finally, the last section concludes.

Methodology

This study estimates the determinants and their elasticities of bilateral export and import demand functions. The aggregate estimations of basic export and import demand functions employ variables of export and import prices and income variables; however indexes of export and import prices are not available on the bilateral basis. Therefore, following Dornbush (1980), who has adopted real exchange rate in the estimation of the import demand function, real exchange rates are employed in this study as a proxy of export and import prices. Thus, we assume that the bilateral export and import demand functions take the following forms:

$$\ln X_{\mathrm{j,t}} = a_0 + a_1 \ln\left(\frac{Pxe}{P_j}\right)_t + a_2 \ln Y_{j,t} + \varepsilon_{1,t} \tag{1}$$

$$\ln M_{\mathrm{j,t}} = b_0 + b_1 \ln\left(\frac{Pxe}{P_j}\right)_t + b_2 \ln Y_t + \varepsilon_{2,t} \tag{2}$$

where X and M are the values of export and import, respectively. P_t is domestic, EU, price level, $P_{j,t}$ is the price for the *j*th country, *e* is the nominal bilateral exchange rate represented in foreign currency per Euro. Finally, $Y_{j,t}$ is the foreign output at period *t*, and Y_t is the domestic output. An increase in real exchange rate, the appreciation of domestic, EU, currency, is expected to decrease its exports and increase imports. Thus, α_1 is expected to have negative sign and b_1 is expected to have positive sign. When the foreign economy growth level of exports to that country increases, it is expected that the coefficient α_2 has a positive sign. Similarly, the growth of the domestic economy leads to an increase in the level of imports. Therefore, coefficient b_2 is expected to have positive sign as well.

Structural change presence

In the long run macroeconomic series such as export, import, output and price levels may contain a variety of structural changes within a country or at the international level, such as economic or political changes or some other shocks. Therefore in order to examine the regression models (1) and (2) in the presence of multiple structural breaks, Bai and Perron

(1998) methodology was employed in this study. The methodology considers the multiple linear regression in the presence of *m* breaks, which means *m+1* regimes.

$$y_t = x_t'\beta + z_t'\delta_j + e_t \qquad (3)$$

where $t = T_{j-1} + 1, \ldots, T_j$ is the time period with $j = 1, \ldots, m+1$ regimes. y_t is dependent variable of the regression, x_t and z_t are vectors of covariates with sizes of ($px1$) and ($qx1$), respectively, β and δ_j are vectors of coefficients, where the parameter vector β is not subject to change, while δ_j is changing across regimes. Finally, e_t is the disturbance term of the regression. The purpose of this methodology is to estimate the unknown coefficients of the regression together with treated as unknown *m* number of break points. For every *m* partition ($T_1, \ldots, T_m$), estimates of coefficients β and δ_j are generated by minimizing the sum of squared residuals which is represented by the following equation:

$$S_T(T_1, \ldots, T_m) = \sum_{i=1}^{m+1} \sum_{t=T_{i-1}+1}^{T_i} \left[y_t - x_t'\beta - z_t'\delta_t \right]^2 \qquad (4)$$

Substituting estimates $\hat{\beta}(\{T_j\})$ and $\hat{\delta}(\{T_j\})$ into equation (3) the estimators of break locations will be obtained, which are the global minimum of the sum of squared residuals objective function, and can be expressed by the following equation:

$$(\hat{T}_1, \ldots, \hat{T}_m) = \arg\min_{T_1, \ldots, T_m} S_T(T_1, \ldots, T_m) \qquad (5)$$

The minimization of the sum of squared residuals is obtained in all partitions ($T_1, \ldots, T_m$), that $T_i - T_{i-1} \geq q$. The estimates of regression parameters are least-squares estimates associated with *m*-partition $\{\hat{T}_j\}$, i.e. $\hat{\beta} = \hat{\beta}(\{T_j\})$ and $\hat{\delta} = \hat{\delta}(\{T_j\})$. Bai and Perron (2003) proposed the efficient algorithm of obtaining the locations of break points, which is based on the principle of dynamic programming.

The procedure for the specification of the number of breaks proposed by Bai and Perron (1998) is as follows. Firstly, the statistics for *UD*max and *WD*max tests have to be calculated. *UD*max and *WD*max tests are double maximum tests that examine for the hypothesis of no structural break against an unknown number of breaks with the given upper bound of breaks *M*, and can be calculated by the following formulas:

$$UD\max F_T(M, q) = \max_{1 \leq m \leq M} \sup_{(\lambda_1, \ldots, \lambda_m) \in \Lambda_\varepsilon} F_T(\lambda_1, \ldots, \lambda_m; q). \qquad (6)$$

where $F_T(\lambda_1, \ldots, \lambda_m; q)$ is the sum of *m* dependent chi-square random variables, each one divided by *m*, with *q* as degree of freedom.

$$WD\max F_T(M,q) = \max_{1\le m\le M} \frac{c(q,\alpha,1)}{c(q,\alpha,m)} x \sup_{(\lambda_1,\ldots,\lambda_m)\in\Delta_\varepsilon} F_T(\lambda_1,\ldots,\lambda_m;q). \tag{7}$$

where *c(q, α, m)* is the asymptotic critical value of the individual tests with *α* as significance level.

Next, Wald type tests have to be applied, where the sup F(0|1) test examines for the hypothesis of no breaks against 1 break existence. If the statistics of this test reject the hypothesis of no breaks, the sup F(*l+1*| *l*) has to be applied to specify the number of breaks in series. The number of breaks in series can be chosen as well on the basis of the Bayesian Information Criteria (BIC), and the modified version of BIC proposed by Liu et al. (1997) (LWZ).

Unit root tests

Before proceeding to cointegration tests, the stationarity of employed variables has to be examined. In order to test integration properties of variables two different unit root tests were applied. The first test is the unit root test proposed by Ng and Perron (2001), which has maximum power against *I*(0) alternatives. In order to generate efficient versions of the modified tests of Perron and Ng (1996), Ng and Perron (2001) employed the generalized lest squares detrending procedure proposed by Elliot, Rothenberg and Stock (1996). Ng and Peron stressed that the choice of the lag length of a regression is extremely important for the good size and power properties of an efficient unit root test. Therefore, Ng and Perron proposed modified AIC and recommended the use of a minimized value of modified Akaike information criterion (AIC) for selecting the regression's lag length.

An additional unit root test is employed in this study is a test proposed by Zivot and Andrews (1992), which is the sequential break point selection test with the null hypothesis of unit root without structural break against the alternative that series are trend-stationary with one break point. Zivot and Andrews considered three different models: model A allows for a break in the intercept; model B allows for a break in the slope; and model C allows for a single break in the intercept and in the slope of the function. In this study, model C was employed.

Cointegration

In order to test for cointegration characteristics between variables under the consideration of a structural break presence, the Gregory and Hansen (1996) test was employed for countries where one structural shift was detected. This test allows for the break in the three alternative models, such as a break in the level (model C), in the level with trend (model C/T), and in the level and slope coefficients (model C/S). For cases where the Bai and Perron (1998) test detected two breaks, the Hatemi-J (2008) test was employed. The Hatemi-J (2008) test is an extended procedure of the Gregory and Hansen (1998) method to allow for two structural shifts in three different models: model C, model C/T and model C/S. In order to apply the Hatemi-J (2008) test, export and import equations (1 and 2) have to be reformulated in the following form:

$$\ln X_{\mathrm{j,t}} = \alpha_0 + \alpha_1 D_{1t} + \alpha_2 D_{2t} + \beta_0 \ln P_t + \beta_1 D_{1t} \ln P_t + \beta_2 D_{2t} \ln P_t + \\ + \gamma_0 \ln Y_{jt} + \gamma_1 D_{1t} \ln Y_{jt} + \gamma_2 D_{2t} \ln Y_{jt} + \varepsilon_{1t} \tag{8}$$

$$\ln M_{\mathrm{j,t}} = \alpha'_0 + \alpha'_1 D_{1t} + \alpha'_2 D_{2t} + \beta'_0 \ln P_t + \beta'_1 D_{1t} \ln P_t + \beta'_2 D_{2t} \ln P_t + \\ + \gamma'_0 \ln Y_{jt} + \gamma'_1 D_{1t} \ln Y_{jt} + \gamma'_2 D_{2t} \ln Y_{jt} + \varepsilon_{1t} \tag{9}$$

Where D_{1t} and D_{2t} are binary indicator variables that are identified as follows:

$$D_{1t} = \begin{cases} 0 & if \quad t \le T_1 \\ 1 & if \quad t > T_1 \end{cases}, \text{ and } \qquad D_{2t} = \begin{cases} 0 & if \quad t \le T_2 \\ 1 & if \quad t > T_2 \end{cases}$$

T_1 identifies the period before the first break point, and T_2 identifies the period before the second break point, where $T_1 + T_2 = T$ is the sample size. Coefficients β_0 and γ_0 in equation (8) and β'_0 and γ'_0 in equation (9) denote slope coefficients before break points. Coefficients β_1 and γ_1 in equation (8) and β'_1 and γ'_1 in equation (9) denote changes in the slopes at the time of the first structural shift and finally coefficients β_2 and γ_2 in equation (8) and β'_2 and γ'_2 in equation (9) denote changes in the slopes at the time of the second structural shift.

For countries where more than two breaks were detected, the Maki (2012) test was applied. The Maki (2012) test is based on the Bai and Parron (1998) test for structural breaks, and on the unit structural breaks proposed by Kapetanios (2005). The Maki (2012) proposes cointegration tests allowing for an unknown number of breaks. The null hypothesis of the test is no cointegration, with the alternative hypothesis of cointegration with unspecified number of breaks i that are smaller or equal to the maximum number of breaks ($i \le k$). The Maki (2012) test has an advantage over standard cointegration tests that allow for one or two structural changes in cointegration relationships when multiple unknown numbers of breaks exist.

Empirical Results

Unit root tests

Table 1 presents the results of the Ng and Perron (2001) unit root tests for variables employed in the model: values of export and import, real exchange rate and foreign and domestic income. The results of all tests in Table 1 are consistent with each other. The null hypothesis of the unit root was not rejected for any of the series by any of Ng and Perron tests. The results of the unit root test demonstrate the non-stationarity of the variables in use.

Next, the Zivot and Andrews (1992) unit root test was applied in Table 2, which allows for the structural break allocation. Only a few of the countries exhibited the absence of unit root in their series. Table 2 displays the *t* statistics of the test and possible break locations. Thus, the unit root hypothesis was rejected for export series only in the case of South Korea. Import series displayed the stationarity in the cases of Brazil, India, and Russia. At the same time, the hypothesis of non-stationarity was not rejected for real exchange rate and income series in any case.

The non-stationarity of the series under observation is verified for all countries except South Korea in the case of export series, and Brazil, India, and Russia in the case of import series only when structural break is taken into account. Estimation results will be interpreted on the basis of unit root results. Having verified the non-stationarity of the series under observation by the Ng and Perron (2001) unit root test, with mixed results using the Zivot and Andrews (1992) unit root test, structural change presence and cointegration tests were conducted.

Structural change presence

The Bai and Perron (1998) procedure allows for the presence of non-stationary as well as stationary variables; however, it is developed for cointegrated regression models. Therefore, before proceeding to the test of structural changes presence, it is important to estimate the cointegrating relationships of considered variables. Table 3 presents the results of the Johansen cointegration test estimations. The Johansen cointegration test presents two statistics, Trace and Max-Eigenvalue. In most cases, the results of the Trace likelihood ratio test statistic and of the Max-Eigenvalue likelihood ratio test statistic are consistent with each other. The results report, at least on the cointegration relationship, in both the export and import equations. The results of the Johansen cointegration test estimations provided enough evidence to conclude that cointegration relationships exist in export and import demand equations when structural breaks are not considered.

Table 4 reports the results of the Bai and Perron (1998) tests for detecting structural changes in series for export and import demand equations. The Sup F(*k*) tests are significant for at least one value of *k* in all cases. The last two columns of the table present statistics for *UD*max and *WD*max tests. The null hypothesis of no structural breaks in all countries was rejected by both tests. As a result, it can be concluded that there is strong evidence of structural changes in the employed series of export and import demand equations.

Table 5 reports the results for the sequential test of l versus $l+1$ structural changes proposed by Bai and Perron (1998). In this study, the Bayesian information criterion (BIC) and the modified Schwarz criterion (LWZ) are used for the detection of the number of breaks, which are presented in last two columns. In both cases of export and import demand equations, at least one break was detected by one of employed information criterion with an upper bound of five breaks.

Cointegration

Table 6 presents the results of the Gregory and Hansen (1996) cointegration test applied to countries where at least one of the information criterion for model selection - BIC or LWZ - detected one structural break. Three different models were applied in running the cointegration test, (C) a structural shift in the intercept, (C/T) a structural shift in the slope, and (C/S) a structural shift in both intercept and slope of the regression. The Gregory and Hansen test is employed to investigate the relation among series in the export equations of South Korea, Switzerland, Turkey, and the US; and in the import equations of Brazil and Turkey. The results of the cointegration test statistics of ADF*, Z_t*, and Z_α* provide enough evidence for cointegration in all cases where at least one of the test statistics of the Gregory and Hansen cointegration test suggest the existence of cointegration relations in exports and imports models. The choice of model C, C/T, or C/S does not significantly affect the results of the cointegration test, except in the case of import demand function in Brazil. Cointegration is found only when a structural shift is allowed in both intercept and the slope.

Next, the Hatemi-J (2008) test is applied to cases where the Bai and Perron (1998) test detected two structural shifts. The test is employed for the export demand function for Canada and Russia, and for the import demand function for Canada and the United States. The results are reported in Table 7. As the results reveal, at least two of the tests in each case reject the hypothesis of no cointegration at the one percent significance level in both export and import models.

The Maki (2012) test was employed to cases where the Bai and Perron (1998) test detected more than two structural shifts. The results of the Maki (2012) test are demonstrated in Table 8, where MB*k* presents t-statistics of the Maki test where *k* denotes the maximum number of breaks. The results imply that the export demand functions of Brazil, India, Japan and import demand functions of China, India, Russia, South Korea and Switzerland have cointegration relationships when multiple unknown numbers of breaks are allowed. The test statistics failed to reject the null hypothesis of no cointegration in cases of export demand function for China and Norway, and in the cases of import demand function of Japan and Norway.

Coefficients estimates

Table 9 reports the results of the parameters estimations of regression (3) in the presence of structural breaks, where dependent variable y_t is the value of export in the case of export demand function and the value of import in case of import demand function. The covariate x_t is the vector of dependent variables: real exchange rate and foreign income in the case of export demand function, and real exchange and domestic income in the case of import demand function.

Estimates of break locations are given in the last four columns $\{\hat{T}_j\}$ of the table based on a 95 percent confidential level. Estimates of the coefficients $\hat{\beta}_1$ and $\hat{\beta}_2$ in the presence of structural breaks are given in the second and third columns. In all cases of export and import demand functions coefficients were found significant. In the case of the export demand function estimates of real exchange rate coefficient carry the expected negative sign for all countries except Turkey. Appreciation of the Euro increases the relative prices of European goods and leads to a decrease in European exports to partner countries. The highest estimate of the coefficient in absolute value was found in the United States, 1.03, where three breaks were detected by BIC and LWZ. The lowest estimate of the coefficient was found in the case of Brazil, where three breaks were detected as well. Only in the case of Turkey was the real exchange rate coefficient found positive, indicating that appreciation of the Euro does not lead to a decrease in its exports to Turkey, but leads to slight increase. The lack of alternatives to the exported products may explain the positive sign of the coefficient. The EU exports to Turkey are dominated by the industrial sector, at 95.9 percent. The Turkish industrial sector is highly dependent on European raw materials and parts.[3] Therefore, appreciation of the Euro leads to an increase in the value of its exports to Turkey, because its industrial sector does not have enough alternatives for European raw materials and industrial sector parts. Nevertheless, the value of the estimate, 0.19, is too low to make a conclusion on the significant effect of the currency appreciation on the bilateral trade with Turkey.

Estimates of the real exchange rate coefficient in the import demand function were found significant, however, with unexpected negative sign in most countries, indicating that appreciation of the Euro leads to a decrease in its imports from the estimated countries. The highest estimate of the coefficient in absolute value was found in the case of Norway, 1.41. In all other cases, the real exchange rate coefficients were estimated at levels below unity with the lowest level 0.09 in the case of China and highest level 0.074 in the case of South Korea. An increase in the relative prices of the European Union leads to a decline in its imports, or vice versa, a decrease in the relative prices of the European Union leads to an increase in its imports. Therefore, a negative sign of the real exchange rate coefficient can be explained by the J curve existence, when depreciation of the domestic currency makes imported goods more expensive. The volume of imported goods does not change immediately for several reasons, such as existing contracts or difficulty in finding cheaper alternatives immediately. Therefore, the value of imports in short run, even in the medium run, increases. On the other hand, a unexpected negative sign of the real exchange rate coefficient in import demand function can be explained by the types of products the EU imports, indicating the shortage of domestic alternatives for imported products. Exchange rate elasticity is more than unity only in the case of Norway, where a slight depreciation of the Euro means an increase in the relative price of foreign products; however, it leads to significant increases in the value of imports from Norway. The EU's imports from Norway mainly consist of energy-related products, which composed 55.6 percent[4] of total EU's imports from Norway in 2012. After Russia, Norway is the largest supplier of energy products to the EU. Therefore, a lack of domestic alternatives does not allow to Europe to decrease immediately the value of imports from Norway.

Only in the cases of Russia and Turkey were the sign of the estimated coefficient found with the expected positive sign, where the appreciation of euro stimulated imports (or the depreciation of the euro led to a decrease in imports) from Russia and Turkey. Russia is the main primary energy supplier of the EU's imports, such as crude oil, natural gas and hard

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[3] European Commission, Directorate-General for Trade

[4] European Commission, Directorate-General for Trade.

coal. Thus Russia is a leading supplier of hard coal, where its share of EU-27 imports of hard coal increased from 13.1 percent in 2002 to 27.1 percent in 2010, with a peak in 2009 when Russia's share consisted 30.2 percent. The share of EU-27 imports of crude oil from Russia steadily increased from 29.2 percent in 2002 to 34.5 percent in 2010.[5] However, at the same time, the share of European imports of natural gas from Russia significantly decreased from 45 percent in 2002 to 31.8 percent in 2010,[6] while the share of natural gas from countries such as Qatar, Nigeria, Libya, and others increased substantially. Lately, Europe has been trying to weaken its energy dependence on Russia for political reasons and for high prices policy as well, which has led to a decline in the profits of Gazprom, the largest Russian state gas company. Therefore, the expected positive sign of the exchange rate coefficient demonstrates the EU policy to decrease energy imports from Russia for cheaper alternatives from other countries. The EU imports from Turkey are predominantly industrial products, which composed 91.9 percent[7] in 2012. Thus, the dominant share of the EU imports from Turkey belongs to textiles and transport equipment, and account for about 24 percent,[8] while machinery accounts for 17.7 percent of total EU imports from Turkey. A positive sign of the exchange rate coefficient indicates that relative prices affect decisions on import demand from Turkey.

Income coefficients are significant and positive in all cases of both export demand and import demand functions. The altitude of income coefficient elasticities is significantly higher compared to the real exchange rate elasticities, and in all cases appeared highly elastic, except for the case of China, where estimation results indicated an inelasticity of the income coefficient. The results of estimations in the presence of structural breaks provide evidence of income being the determining factor in the bilateral trade of Europe rather than the real exchange rate, which are similar to studies which do not consider breaks, for example, Hatemi-J and Irandoust (2005), Ketenci and Uz (2010).

Estimated regression coefficients under breaks are compared with estimated coefficients using the OLS procedure, Table 10. The coefficients of real exchange rate in export demand function were estimated using the OLS procedure appeared with the same sign as estimated coefficients under breaks, except in the case of China, for which the coefficient appeared with positive sign. The absolute values of OLS exchange rate coefficients in export function were estimated and appeared lower compared to the estimated coefficients under

[5] Eurostat
[6] Eurostat
[7] European Commission, Directorate-General for Trade.
[8] Statistics on the EU imports is extracted from Eurostat.

breaks, except for the case of Norway. Comparing the estimated exchange rate coefficients of import demand function, the results for two countries came out with opposite sign. The exchange rate coefficient of Brazil appeared with a positive sign, while the exchange rate coefficient of Russia appeared with a negative sign in the OLS estimations. The absolute values of estimated exchange rate coefficients do not have a tendency of increase or increase in estimations of a particular procedure. Income coefficients in export and import demand functions appeared with the expected positive sign in both cases with breaks and without breaks consideration. In both cases, income coefficients were estimated as highly elastic, except for China, in export demand function. The increase in Chinese GDP led only to a slight increase in EU exports to China.

Conclusion

This study investigated the bilateral trade flows of the EU with its major partners. The largest export partners of the EU are the US, China, Switzerland, Russia, Turkey, Japan, Norway, India, Brazil, South Korea; and the largest import partners are China, Russia, the US, Norway, Switzerland, Japan, Turkey, India, Brazil, South Korea. First of all, in order to test the existence of structural breaks in export and import demand functions, the Bai and Perron (1998) procedure was employed. The results of this procedure, Table 4 and 5, detected at least one break in both export and import demand equations in all country cases. After detecting structural breaks in the series, the employment of ordinary cointegration tests that do not allow for a break can bring spurious results. Therefore, in order to investigate cointegration characteristics between variables employed cointegration tests allow for structural breaks existence. Three different cointegration tests were employed. First was the Gregory and Hansen (1996) test, applied to series where only one structural break was detected, Table 6. One break was detected in the export equations of South Korea, Switzerland, Turkey, and the US, and in the import equations of Brazil and Turkey. The results of the test provided enough evidence to suggest the existence of cointegration relations in export and import demand equations in all countries with one structural break. Secondly, the Hatemi-J (2008) test was employed to cases where two structural breaks were detected by the Bai and Perron (1998) procedure, Table 7. Two structural breaks were detected in the export equation of Canada and Russia and in the import equation of Canada and the US. The results of the test provided evidence of cointegration in all cases. Finally, the Maki (2012) was employed to series where more than two structural breaks were detected, as in the export equations of Brazil, China, India, Japan, and Norway, and in import equations of China, India, Japan, Norway, Russia, South Korea and Switzerland. Cointegration relations were found in all cases except export demand function for China and Norway and in cases of import demand function of Japan and Norway. As a result, the employed cointegration tests that allow for the existence of structural breaks provided enough evidence to conclude that cointegration relations exist in all considered cases of export and import demand functions, except for China and Norway in export demand function, and except for Japan and Norway in import demand function.

The results of estimations in the presence of structural breaks provide evidence for income being determining the factor in the bilateral trade of Europe rather than real exchange rate. Similar results were found in studies that were run for EU trade without consideration of breaks; for example, Hatemi-J and Irandoust (2005), and Ketenci and Uz (2011). This study found that cointegration relations between variables exist in both export and import demand function in all countries when structural breaks are not taken into account (Table 3). However, estimations of cointegration tests that allow for structural breaks revealed the absence of cointegrating relations between export value, real exchange rate, and foreign income in the export equations of China and Norway. In the case of import demand equations, the existence of cointegration relations was not supported by tests in Japan and Norway.

The increased number of crises in the preceding decades damaged not only developing countries, but developed countries as well. Moreover, the origin of crises lately has moved to developed countries also. The main outcome of this study is that the consideration of structural breaks in research is of great importance. Estimations with uncounted structural breaks do not bring entirely different results; however, important details can be missing and misinterpreted.

References

Arize, A. (1987). The supply and demand for imports and exports in a simultaneous model. *Applied Economics, 19(9),* 1233-1247.

Aziz, J. and Li, X. (2007). China's changing trade elasticities. *IMF Working Paper, November,* WP/07/266.

Bahmani-Oskooee, M. and Kantipong, T. (2001). Bilateral J-curve between Thailand and her trading partners. *Journal of Economic Development 26*, 107-117.

Bahmani-Oskooee, M. and Niroomand, F. (1998). Long-run price elasticities and the Marshall-Lerner condition revisited. *Economics Letters, 61(1),* 101-109.

Bahmani-Oskooee, M. and Ratha, A. (2008). Exchange rate sensitivity of US bilateral trade flows. *Economics System, 32*, 129-141.

Bai, J. and Perron, P. (1998). Estimating and testing linear models with multiple structural changes. *Econometrica, 66,* 47-68.

Bai, J. and Perron, P. (2003). Computation and analysis of multiple structural change models. *Journal of Applied Econometrics, 18*, 1–22.
Dornbusch, R. (1980). *Open Economy Macroeconomics*. New York. Basic Books.
Elliot, G., Rothenberg, T.J. and Stock, J.H. (1996). Efficient tests for an autoregressive unit root. *Econometrica, 64*, 813-836.
Hatemi-J, A. and Irandoust, M. (2005). Bilateral trade elasticities: Sweden versus her major trading partners. *American Review of Political Economy, 3(2),* 38-50.
Hatemi-J, A. (2008). Tests for cointegration with two unknown regime shifts with an application to financial market integration. *Empirical Economics, 35(3)*, 497-505.
Gregory, A.W. and Hansen, B.E. (1996). Tests for cointegration in models with trend and regime shifts. *Oxford Bulletin for Economics and Statistics, 58(3)*, 555-560.
Johansen, S. (1988). Statistical analysis of cointegrating vectors. *Journal of Economic Dynamics and Control*, *1*2, 231-54.
Imbs, J. and Meajean, I. (2010). Trade elasticities: A final report for the European Commission. *Economic Papers, 432*, December.
Kapetanios, G. (2005). Unit-root testing against the alternative hypothesis of up to m structural breaks. *Journal of Time Series Analysis, 26*, 123–133.
Ketenci,N and Uz, I. (2011). Bilateral and regional trade elasticities of the EU. *Empirical Economics, 40(3),* 839-854.
Khan, M. (1974). Import and export demand in developing countries. *IMF Staff Papers, September*, 678-693.
Kreinin, M.E. (1967). Price elasticities in the international trade. *The Review of Economics*
Liu, J., Wu, S., & Zidek, J.V. (1997). On segmented multivariate regressions. *Statistica Sinica, 7*, 497-525.
Mah, J.S. (1993). Structural change in import demand behavior: The Korean experience. *Journal of Policy Modelling, 15(2)*, 223-227.
Marquez, J. (1990). Bilateral trade elasticities. *The Review of Economics and Statistics*, *February*, 70-77.
Murad, S.M.W. (2012). Bilateral export and import demand functions of Bangladesh: A cointegration approach. *Bangladesh Development Studies*, *XXXV(1)*, 43-60.
Maki, D. (2012). Tests for cointegration allowing for an unknown number of breaks. *Economic Modelling*, *29(5)*, 2011-2015.
Ng, S. and Perron, P. (2001). Lag selection and the construction of unit root tests with good size and power. *Econometrica*, *69*, 1519–1554.
Perron, P. and Ng, S. (1996). Useful modifications to some unit root tests with dependent errors and their local asymptotic properties. *Review of Economic Studies, 63*, 435-463.
Sinha, D. (2001). A note on trade elasticities in Asian countries. *The International Trade Journal, 15(2)*, 221-237
Uz, I.(2010). Testing for structural change in the bilateral trade elasticities of Turkey. *Middle East Technical University Studies in Development, 37(1),* 53-72.
Zivot, E. and Andrews, D. (1992). Further evidence of great crash, the oil price shock and unit root hypothesis. *Journal of Business and Economic Statistics*, *10(3)*, 251-270.

Table 1. Unit Root Tests Ng and Perron (2001)

Country	MZ_{α}^{GLS}	MZ_{t}^{GLS}	MSB^{GLS}	MP_{T}^{GLS}	MZ_{α}^{GLS}	MZ_{t}^{GLS}	MSB^{GLS}	MP_{T}^{GLS}
	Export				***Import***			
Brazil	1.49	1.89	1.27	119.28	0.73	0.67	0.92	57.07

Canada	0.90	1.08	1.19	94.85	1.65	1.89	1.14	100.86
China	1.45	1.57	1.09	88.27	-0.82	-0.39	0.47	15.39
India	1.19	1.24	1.04	77.95	1.38	1.55	1.12	92.93
Japan	0.79	1.01	1.29	105.97	0.04	0.04	0.96	52.73
Norway	1.46	1.99	1.37	135.99	1.50	2.28	1.52	167.52
Russia	1.61	1.59	0.99	76.87	-2.05	-0.88	0.43	10.76
South Korea	1.11	1.30	1.18	96.07	0.52	0.62	1.18	85.63
Switzerland	1.74	2.58	1.48	167.79	1.16	1.12	0.96	66.91
Turkey	1.30	2.35	1.81	225.98	0.98	1.23	1.26	106.26
US	0.82	0.99	1.22	96.44	0.69	0.68	0.98	63.81
	RER				***Income***			
Brazil	-0.0002	-0.0002	0.83	40.69	1.43	1.18	0.82	53.30
Canada	-8.93	-2.09	0.24	2.81	1.32	1.81	1.37	134.13
China	-0.38	-0.30	0.79	34.49	2.15	4.50	2.09	351.79
India	-1.54	-0.81	0.53	14.67	1.46	1.38	0.94	68.42
Japan	-3.70	-1.32	0.36	6.64	0.25	0.21	0.85	44.74
Norway	-6.71	-1.83	0.27	3.65	0.46	0.34	0.73	36.71
Russia	-1.24	-0.64	0.52	15.48	-2.76	-0.86	0.31	7.99
South Korea	-11.86	-2.39	0.20	2.21	1.04	1.48	1.42	134.09
Switzerland	-7.07	-1.75	0.25	3.93	1.61	2.27	1.41	148.31
Turkey	-0.84	-0.58	0.69	24.62	1.62	3.36	2.08	314.96
US	-7.72	-1.96	0.25	3.19	-0.17	-0.09	0.57	21.83
EU					0.94	1.19	1.27	106.84

Notes: MZ_α^{GLS} is the modified Phillip-Perron test MZ_α; MZ_t^{GLS} is the modified Phillip-Perron MZ_t test; MSB^{GLS} is the modified Sargan-Bhargava test; MP_T^{GLS} is the modified point optimal test, for details see Ng and Perron (2001). The order of lag to compute the test has been chosen using the modified AIC (MAIC) suggested by Ng and Perron (2001). ** denote the rejection of the null hypothesis at the 1% significance level. The critical values for the above tests have been taken from Ng and Perron (2001).

Table 2. Unit Root Tests Zivot and Andrews

Country	t statistics	break	t statistics	break
	Export		***Import***	
Brazil	-4.167	1994:4	-6.515**	1986:1
Canada	-3.901	1999:1	-3.245	1992:2
China	-4.027	1986:2	-4.346	2006:4
India	-4.909	1986:1	-5.499*	1986:1
Japan	-4.291	1993:4	-3.436	1998:4
Norway	-4.555	1987:1	-4.112	1991:1
Russia	-4.311	1992:1	-6.961**	1992:1
South Korea	-6.780**	1998:1	-4.078	2005:4
Switzerland	-3.634	1991:2	-2.659	2002:1
Turkey	-4.113	1996:2	-4.184	2004:2
US	-4.082	1999:2	-3.297	1996:4
	RER		***Income***	
Brazil	-4.601	1995:3	-4.643	1990:1
Canada	-4.621	1999:1	-3.449	1998:3
China	-4.002	1985:2	-4.507	2001:4

India	-4.114	1989:4	-4.677	2000:4
Japan	-4.006	2002:4	-4.692	1989:3
Norway	-4.754	1999:1	-2.922	1995:3
Russia	-4.056	1990:1	-4.733	1992:1
South Korea	-5.040	1985:2	-3.142	1994:1
Switzerland	-4.066	2006:1	-4.272	1992:3
Turkey	-5.059	1985:3	-4.603	1998:4
US	-4.021	1985:2	-3.630	2003:2
EU			-3.580	2006:1

Notes: The critical values for Zivot and Andrews test are -5.57, -5.08 and -4.82 at 1 %, 5 % and 10% levels of significance respectively.
* denotes statistical significance at 5% level. ** denotes statistical significance at 1% level.

Table 3. Johansen Cointegration Test

Country	Trace statistics		Max-Eigen Statistics	
Export equation	$r = 0$	$r \leq 1$	$r = 0$	$r \leq 1$
Brazil	36.73**	13.63	23.09**	9.74**
Canada	48.88**	25.11**	23.77**	16.53**
China	165.88**	20.76**	145.12**	17.65**
India	84.49**	25.99**	58.49**	20.09**
Japan	49.44**	26.27**	23.18**	15.89**
Norway	57.35**	20.15**	37.19**	13.58
Russia	38.22**	15.63	22.59**	11.09
South Korea	66.71**	28.29**	38.41**	15.92**
Switzerland	53.75**	16.19	37.57**	10.50
Turkey	65.23**	27.74**	37.48**	15.92**
US	70.27**	28.48**	41.78**	19.61**
Import equation				
Brazil	40.35**	17.09	23.26**	12.36
Canada	51.42**	23.66**	27.76**	17.63**
China	42.11**	11.07	31.04**	7.07
India	46.77**	6.44	40.33**	4.49
Japan	42.83**	17.56	25.27**	11.39
Norway	35.63**	17.06	18.58	13.42
Russia	68.04**	16.97	51.07**	15.06
South Korea	53.27**	27.59**	25.68**	19.57**
Switzerland	37.33**	11.46	25.87**	10.51
Turkey	59.04**	21.63**	37.41**	14.88
US	78.68**	19.57**	59.12**	12.32

Notes: ** denotes statistical significance at 5% level.

Table 4. Structural Break Tests of Bai and Perron (1998).

Country	Sup F(1)	Sup F(2)	Sup F(3)	Sup F(4)	Sup F(5)	UDmax	WDmax
Export equation							
Brazil	0.02	15.72**	431.94**	410.66**	87.27**	431.94**	655.56**

Canada	2.83	851.38**	501.99**	1670.27**	137.17**	1670.27**	2666.39**
China	0.44	3.99	1.83	6.31*	11.86**	11.86**	24.05**
India	0.00	2.01	12.03**	2.78	2.09	12.03**	17.32**
Japan	0.0001	14.77**	16.36**	187.89**	19.67**	187.89**	323.08**
Norway	0.00	5.62*	47.04**	33.00**	28.97**	47.04**	67.71**
Russia	0.00	197.94	36555**	210.12**	1037.37**	36555**	52642**
South Korea	0.001	13.79**	35.44**	129.70**	87.36**	129.70**	223.02**
Switzerland	0.00	6.96	32.15**	69.22**	58.41**	69.22**	128.18**
Turkey	0.0001	9.39*	34.38**	315.35**	5519.39**	5519.39**	12111.61**
US	0.0006	13.60**	172.96**	63.17**	926.90**	926.90**	2033.96**
Import equation							
Brazil	0.028	4.16	3345.79**	7.80**	6.16*	3345.79**	4816.60**
Canada	0.00	1394.57**	91537.87**	107.26**	6.32*	91537.87**	131777**
China	0.0005	1218.91**	2270.76**	62.38**	250.40**	2270.76**	3268.35**
India	0.00	5.10	78.45**	129.02**	212.75**	212.75**	466.84**
Japan	0.00	17.05**	616.15**	6264.27**	3319.00**	6264.27**	10771.04**
Norway	0.00	1326.59**	29.40**	25.94**	1.03	1326.59**	1576.47**
Russia	0.00	103.67**	470.05**	1281.80**	977.62**	1281.80**	2203.98**
South Korea	0.0002	1239.77**	718.63**	39.63**	75.74**	1239.77**	1473.30**
Switzerland	0.0003	22.21**	22.05**	32.15**	32.78**	32.78**	71.93**
Turkey	0.00	8.13*	15.42**	11.80*	940.67**	940.67**	2064.18**
US	0.0001	272.14**	337.23**	25.24**	4344.58**	4344.58**	9533.37**

Notes: * denotes statistical significance at 5% level. ** denotes statistical significance at 1% level.

Table 5. Sequential test of *l* versus *l+1* structural changes Bai and Perron (1998).

Country	0 BIC (LWZ)	1 BIC (LWZ)	2 BIC (LWZ)	3 BIC (LWZ)	4 BIC (LWZ)	5 BIC (LWZ)	BIC	LWZ
Export equation								
Brazil	-4.12 -4.11	-4.27 -4.19	-4.57 -4.43	-5.29 -5.08	-5.22 -4.93	-4.95 -4.59	3	3
Canada	-5.72 -5.71	-6.39 -6.32	-6.62 -6.48	-6.59 -6.38	-6.71 -6.43	-6.63 -6.28	4	2
China	-2.85 -2.84	-3.48 -3.41	-3.69 -3.54	-4.25 -4.04	-4.50 -4.22	-4.56 -4.21	5	4
India	-5.12 -5.11	-5.08 -5.01	-5.17 -5.02	-5.29 -5.08	-5.24 -4.96	-5.16 -4.81	3	0
Japan	-6.39 -6.38	-6.39 -6.32	-6.97 -6.82	-7.05 -6.84	-7.03 -6.75	-7.06 -6.71	5	3
Norway	-6.47 -6.46	-6.61 -6.54	-6.79 -6.65	-6.92 -6.71	-6.88 -6.59	-6.68 -6.33	3	3
Russia	-2.78 -2.77	-3.71 -3.63	-4.72 -4.57	-4.66 -4.44	-4.59 -4.31	-4.54 -4.19	2	2
South Korea	-6.26 -6.26	-6.49 -6.42	-6.53 -6.39	-6.49 -6.29	-6.53 -6.25	-6.43 -6.08	4	1
Switzerland	-7.29 -7.28	-7.37 -7.29	-7.36 -7.21	-7.41 -7.20	-7.42 -7.14	-7.35 -6.99	4	1
Turkey	-5.31 -5.30	-5.69 5.62	-5.68 -5.33	-5.76 -5.55	-5.74 -5.45	-5.67 -5.32	3	1
US	-6.74 -6.73	-6.98 -6.90	-6.98 -6.83	-6.97 -6.76	-6.94 -6.66	-6.86 -6.51	1	1
Import equation								
Brazil	-4.38 -4.37	-4.46 -4.38	-4.42 -4.28	-4.45 -4.24	-4.46 -4.18	-4.35 -4.00	1	1
Canada	-5.81 -5.81	-5.86 -5.79	-6.02 -5.88	-6.08 -5.87	-6.09 -5.81	-6.06 -5.71	4	2
China	-5.66 (-5.65)	-5.66 (-5.58)	-5.99 (-5.85)	-6.22 (-6.01)	-6.34 (-6.06)	-6.27 (-5.92)	4	4
India	-4.77 (-4.77)	-4.87 (-4.79)	-4.87 (-4.73)	-5.19 (-4.98)	-5.27 (-4.99)	-5.19 (-4.85)	4	4
Japan	-5.42 (-5.41)	-5.79 (-5.71)	-6.97 (-6.82)	-7.17 (-6.95)	-7.11 (-6.83)	-7.04 (-6.69)	3	3
Norway	-5.23 (-5.22)	-5.72 (-5.65)	-5.88 (-5.73)	-5.95 (-5.74)	-5.92 (-5.64)	-5.87 (-5.52)	3	3
Russia	-2.59 (-2.59)	-2.90 (-2.82)	-3.07 (-2.93)	-3.33 (-3.11)	-3.27 (-2.99)	-3.19 (-2.84)	3	3
South Korea	-5.06 (-5.05)	-5.15 (-5.08)	-5.23 (-5.08)	-5.34 (-5.13)	-5.39 (-5.11)	-5.37 (-5.02)	4	3
Switzerland	-6.51 (-6.49)	-6.59 (-6.51)	-6.69 (-6.55)	-6.79 (-6.58)	-6.84 (-6.56)	-6.82 (-6.47)	4	3
Turkey	-6.14 (-6.13)	-6.39 (-6.13)	-6.32 (-6.18)	-6.34 (-6.13)	-6.28 (-5.99)	-6.25 (-5.89)	1	1
US	-6.12 (-6.12)	-6.42 (-6.35)	-6.65 (-6.51)	-6.64 (-6.42)	-6.57 (-6.29)	-6.49 (-6.15)	2	2

Notes: S - sequential procedure, BIC - Bayesian Information Criteria, LWZ - the modified version of BIC proposed by Liu et al. (1997), are used for the selection of breaks number.

Table 6. Cointegration test with a structural break Gregory and Hansen

Country	Model	ADF*	Z_t^*	Z_α^*
EXPORTS				
South Korea	C	-3.74	7.24**	-77.41**
	C/T	-4.58	-7.84**	-84.80**
	C/S	-4.47	-7.56**	-80.98**
Switzerland	C	-4.57	-9.96**	-111.55**
	C/T	-4.73	-10.49**	-119.47**
	C/S	-4.75	-10.09**	-113.72**
Turkey	C	-6.07**	-10.62**	-120.02**
	C/T	-5.49*	-10.71**	-121.43**
	C/S	-6.17**	-11.27**	-128.59**
US	C	-4.75	-7.90**	-75.69**
	C/T	-5.03	-7.92**	-75.29**
	C/S	-5.77*	-9.43**	-101.78**
IMPORTS				
Brazil	C	-4.75	-5.23	-40.17
	C/T	-5.51*	-5.39	-42.33
	C/S	-4.54	-7.23**	-70.66**
Turkey	C	-3.63	-7.93**	-87.22**
	C/T	-4.27	-8.36**	-90.51**
	C/S	-4.16	-8.66**	-97.32**

Notes: * denotes statistical significance at 5% level. ** denotes statistical significance at 1% level.

Table 7. Cointegration test with two structural breaks Hatemi-J (2008)

Country	Model	ADF*	Z_t^*	Z_α^*
EXPORTS				
Canada	**C/S**	**-5.31**	**-9.46****	**-107.31****
	1	1988:Q4	1988:Q1	
	2	1997:Q2	1994:Q2	
Russia	**C/T**	**-5.16**	**-9.75****	**-110.04****
	1	1995:Q1	1986:Q2	
	2	1998:Q1	2002:Q4	
IMPORTS				
Canada	**C/S**	**-5.66**	**-7.96****	**-85.37***
	1	1995:Q3	1994:Q2	
	2	2001:Q4	2001:Q4	
US	**C/S**	**-5.54**	**-10.30****	**-118.62****
	1	1988:Q3	1987:Q4	
	2	1995:Q2	1996:Q1	

Notes: The critical values are collected from Hatemi-J (2008) and are -6.503, -6.015 and -5.653 (1%, 5% and 10%) for ADF and Zt tests, and are -90.794, 76.003 and 52.232 (1%, 5% and 10%).

Table 8. The cointegration test Maki (2012) with unknown number of breaks

Country	MB1	MB2	MB3	MB4	MB5
EXPORTS					
Brazil	-5.21	-7.87***	-8.16***	-8.27***	-8.27***
China	-4.71	-5.22	-5.44	-5.55	-5.83
India	-4.82	-7.33***	-7.33***	-7.33**	-7.33**
Japan	-5.74**	-5.74	-5.93	-5.93	-6.43
Norway	-4.23	-4.31	-4.66	-4.66	-6.85
IMPORTS					
China	-5.53*	-6.17**	-6.30*	-6.30	-6.69
India	-4.25	-5.17	-6.31*	-7.19**	-7.19*
Japan	-4.47	-5.55	-5.59	-5.76	-5.93
Norway	-4.75	-5.05	-5.23	-5.74	-5.74
Russia	-5.51*	-5.51	-5.6	-6.3	-6.61
South Korea	-4.36	-4.36	-8.48***	-8.48***	-8.62***
Switzerland	-5.45*	-5.45	-5.45	-5.99	-6.19

Notes: Critical values are taken from Maki (2012) – Table 1

Table 9. Estimated regression parameters under breaks.

Country	$\hat{\beta}_1$	$\hat{\beta}_2$	$\hat{\delta}_1$	$\hat{\delta}_2$	$\hat{\delta}_3$	$\hat{\delta}_4$	$\hat{\delta}_5$	$\hat{T}_1$	$\hat{T}_2$	$\hat{T}_3$	$\hat{T}_4$
EXPORT											
Brazil **(BIC, LWZ)**	-0.04** (0.01)	3.07** (0.17)	-9.26** (1.04)	-9.51** (1.04)	-9.22** (1.06)	-9.42** (1.08)	-	1986:Q4	1994:Q3	2002:Q3	-
Canada **(LWZ)**	-0.91** (0.06)	2.32** (0.06)	-4.24** (0.35)	-4.11** (0.35)	-4.18** (0.36)	-	-	1984:Q4	2004:Q4	-	-
China **(LWZ)**	-0.32** (0.11)	0.16** (0.03)	8.34** (0.12)	8.79** (0.14)	9.16** (0.16)	9.49** (0.16)	9.79** (0.17)	1984:Q4	1992:Q 4	2000:Q 3	2005:Q4
India **(BIC)**	-0.94** (0.07)	1.86** (0.09)	-2.81** (0.59)	-2.71** (0.61)	-2.83** (0.63)	-2.98** (0.66)	-	1984:Q3	1997:Q4	2008:Q1	-
Japan **(LWZ)**	-0.62** (0.04)	3.17** (0.06)	-9.54** (0.43)	-9.61** (0.43)	-9.54** (0.43)	-9.50** (0.44)	-	1990:Q4	1995:Q3	2000:Q2	-
Norway **(BIC, LWZ)**	-0.96** (0.13)	2.51** (0.08)	-2.48** (0.37)	-2.55** (0.38)	-2.63** (0.38)	-2.58** (0.39)	-	1987:Q4	1999:Q2	2006:Q1	-
Russia **(BIC, LWZ)**	-0.69** (0.05)	1.74** (0.10)	-0.39 (0.65)	-0.03 (0.64)	0.13 (0.65)	-	-	1992:Q4	1997:Q3	-	-
South Korea **(LWZ)**	-0.64** (0.06)	1.77** (0.03)	1.10** (0.16)	1.02** (0.17)	-	-	-	1998:Q2	-	-	-
Switzerland **(LWZ)**	-0.98** (0.092)	2.84** (0.06)	-5.00** (0.29)	-4.97** (0.30)	-	-	-	1996:Q3			
Turkey **(BIC)**	0.19** (0.05)	2.66** (0.11)	-5.59** (0.58)	-5.69** (0.60)	-5.59** (0.62)	-5.71** (0.64)	-	1985:Q4	1995:Q3	2005:Q4	
US **(BIC, LWZ)**	-1.03** (0.04)	2.36** (0.03)	-5.98** (0.21)	-5.91** (0.21)	-	-	-	1985:Q1			
IMPORTS											

Brazil **(BIC, LWZ)**	-0.12** (0.05)	3.28** (0.18)	-12.81** (1.28)	-13.39** (1.27)	-	-	-	1988:Q4			
Canada **(LWZ)**	-0.32** (0.08)	2.15** (0.09)	-5.41** (0.58)	-5.46** (0.59)	-5.37** (0.59)	-	-	1984:Q3	2006:Q4		
China **(BIC, LWZ)**	-0.09** (0.05)	6.90** (0.19)	-38.31** (1.28)	-38.19** (1.29)	-38.24** (1.30)	-38.17** (1.31)	-38.07** (1.31)	1991:Q2	1997:Q4	2003:Q2	2008:Q1
India **(BIC, LWZ)**	-0.71** (0.11)	5.62** (0.29)	-28.63** (1.95)	-28.80** (1.96)	-28.68** (1.97)	-28.89** (1.99)	-28.79** (2.00)	1985:Q4	1992:Q4	1998:Q2	2008:Q1
Japan **(BIC, LWZ)**	-0.45** (0.04)	3.14** (0.07)	-10.89** (0.52)	-10.82** (0.52)	-10.95** (0.54)	-11.08** (0.54)	-	1986:Q4	2001:Q2	2008:Q1	
Norway **(BIC, LWZ)**	-1.41** (0.19)	4.35** (0.19)	-19.01** (1.23)	-19.21** (1.24)	-19.31** (1.26)	-19.23** (1.27)	-	1985:Q4	1998:Q1	2004:Q3	
Russia **(BIC, LWZ)**	0.26** (0.09)	4.72** (0.55)	-22.54** (3.76)	-23.01** (3.80)	-23.53** (3.85)	-23.29** (3.89)	-	1986:Q1	1992:Q1	2004:Q4	
South Korea **(LWZ)**	-0.74** (0.11)	3.53** (0.23)	-13.15** (1.57)	-12.94** (1.58)	-13.01** (1.59)	-12.86** (1.62)		1984:Q3	1989:Q2	1997:Q3	
Switzerland **(LWZ)**	-0.69** (0.12)	2.71** (0.08)	-8.66** (0.53)	-8.72** (0.54)	-8.77** (0.55)	-8.71** (0.55)	-	1995:Q1	2001:Q4	2007:Q4	
Turkey **(BIC, LWZ)**	0.42** (0.03)	5.54** (0.07)	-29.10** (0.47)	-29.20** (0.48)	-	-	-	1988:Q1			
US **(BIC, LWZ)**	-0.44** (0.05)	2.93** (0.08)	-9.85** (0.53)	-9.92** (0.55)	-10.05** (0.55)	-	-	1985:Q2	2002:Q2		

Notes: The parentheses under the break points are 95% confidence intervals for the break dates.
**, * Denote statistical significance at the 1 and 5% level respectively.

Table 10. Estimated regression parameters OLS

Country	α_0	α_1	α_2			b_0	b_1	b_2
EXPORT					**IMPORTS**			
Brazil **(BIC, LWZ)**	-8.67** (0.93)	-0.03** (0.01)	2.95** (0.15)		Brazil **(BIC, LWZ)**	-13.44** (1.37)	0.09** (0.01)	3.28** (0.19)
Canada **(LWZ)**	-3.72** (0.32)	-0.53** (0.08)	2.24** (0.06)		Canada **(LWZ)**	-6.70 ** (0.46)	-0.63** (0.08)	2.33** (0.07)
China **(LWZ)**	5.56** (0.18)	0.68** (0.16)	0.76** (0.05)		China **(BIC, LWZ)**	-43.58** (057-8)	-0.09* (0.05)	7.67** (0.09)
India **(BIC)**	-0.26 (0.22)	-0.53** (0.06)	1.43** (0.04)		India **(BIC, LWZ)**	-23.34** (0.89)	-0.69** (0.09)	4.84** (0.14)
Japan **(LWZ)**	-10.71** (0.29)	-0.41** (0.05)	3.28** (0.04)		Japan **(BIC, LWZ)**	-6.48** (0.46)	-0.99** (0.09)	2.67** (0.07)
Norway **(BIC, LWZ)**	-0.53** (0.17)	-1.02** (0.15)	2.13** (0.05)		Norway **(BIC, LWZ)**	-14.59** (0.61)	-2.34** (0.27)	3.81** (0.11)
Russia **(BIC, LWZ)**	-3.20** (1.61)	-0.35** (0.13)	2.17** (0.26)		Russia **(BIC, LWZ)**	-2.51 (1.87)	-0.32** (0.14)	1.85** (0.27)
South Korea **(LWZ)**	1.73** (0.16)	-0.64** (0.07)	1.65** (0.02)		South Korea **(LWZ)**	-19.25** (0.57)	-0.45** (0.13)	4.31** (0.12)
Switzerland **(LWZ)**	-6.15** (0.18)	-0.93** (0.09)	3.06** (0.04)		Switzerland **(LWZ)**	-6.18** (0.27)	-0.84** (0.15)	2.35** (0.04)
Turkey **(BIC)**	-5.01** (0.21)	0.16** (0.06)	2.55** (0.04)		Turkey **(BIC, LWZ)**	-26.39** (0.32)	0.33** (0.04)	5.14** (0.05)
US **(BIC, LWZ)**	-6.74** (0.21)	-0.94** (0.05)	2.48** (0.03)		US **(BIC, LWZ)**	-5.38** (0.38)	-0.59** (0.06)	2.27** (0.06)

Notes: * denotes statistical significance at 5% level. ** denotes statistical significance at 1% level.
α_0, α_1, α_2 coefficients are from equation 1, b_0, b_1, b_2 coefficients are from equation 2.