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# THE EXPORT PERFORMANCE OF THE EURO AREA: A PANEL QUANTILE REGRESSION APPROACH

**ABSTRACT:** *This paper examines the impact of relevant factors on merchandise exports, volume growth, and the competitiveness of euro-area countries. It uses panel regressions to explain the development of merchandise exports in the euro-area countries by various price/cost competitiveness indicators, development of foreign and domestic demand, and the structure of merchandise exports. A cointegration analysis of panel time series was established by applying econometric tests and the cointegration equation for the period 1999–2018 was estimated using FMOLS (Fully Modified OLS) and DOLS (Dynamic OLS) estimators. We show that the increase in the share of information communication technology (ICT) product exports in the total merchandise exports of euro-area members had a positive impact on their export performance. The empirical results show that foreign demand has a positive impact on real merchandise exports, while the estimated coefficients decreased from the lower to upper quantiles. The results regarding price and cost competitiveness differ depending on the choice of indicators, but in general they are less robust. Therefore, we conclude that in the long run, non-price factors will play an increasingly important role in strengthening the competitive position of euro-area countries in international markets.*

**KEY WORDS:** *Exports, real effective exchange rate, competitiveness, foreign demand, internal demand, panel data, cointegration, quantile regression*

**JEL CLASSIFICATION:** E32, F32, F41, F44.

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## 1. INTRODUCTION

The establishment of the European Central Bank (ECB), which is responsible for monetary policy in the euro area, and the introduction of the euro as the official currency of 19 EU member countries created the framework for the emergence of a monetary union. The path to monetary union entailed harmonising economic conditions and regulations in a number of European Union countries. The coordination of macroeconomic policies is an important prerequisite for the normal functioning of the euro area and for the convergence of its members. A key pillar of coordination is the Stability Pact, which should allow the budget deficit to be kept under control (at no more than 3% of GDP). This threshold was first crossed by Germany, France, and especially Greece, which casts doubt on the eurozone's ability to uphold this rule. Thus, over the course of time, the differences in eurozone members' productivity and inflation levels have increased the pressure on public debt growth in countries that have lost competitiveness within the euro area. The single currency and the common monetary policy for the eurozone have abolished the possibility of using the exchange rate as an instrument of economic policy; that is, as an instrument for raising the export competitiveness of the economically weaker members of the euro area. In some member states the current account deficit has increased. Although the integration of the financial and capital markets has led to an equalisation of interest rates, the rise in government debt has created a gap in the interest rates on eurozone members' government bonds.

When the global financial crisis erupted in 2008 it quickly shifted to eurozone members, and shortly thereafter it was followed by a eurozone sovereign debt crisis.[1] This slowed the economic recovery of the eurozone, and additional difficulties arose with the rise of protectionism in the world economy. The current economic forecasts for the euro area's economic growth are much lower than for US and emerging markets (see IMF 2019 and OECD 2019). Although all euro-area members have made significant progress in reducing their current account imbalance and public debt, with the outbreak of the global crisis the constraints of

1 Some authors believe that the eurozone crisis is not a sovereign debt crisis, but rather that major intra-eurozone capital flows occurred prior to the crisis (see Baldwin and Giavazzi 2015, p.23). The cost of resolving the crisis has been mostly referred to over-indebted eurozone countries (Frieden and Walter 2017).

the eurozone became apparent. The absence of a centralised fiscal policy, low labour mobility within the eurozone, and growing productivity differences after the outbreak of the crisis have all become visible. This makes it difficult to respond to asymmetric shocks. The ECB's non-standard monetary policy significantly eased the euro area's internal tensions but could not remove its inherent constraints.

Many countries' trading performance is influenced by Global Value Chains (GVCs). These production models are widespread in manufacturing processes worldwide, especially in the manufacturing sector. The existence of GVCs greatly affects international trade between large countries and eurozone export performance is significantly influenced by GVCs.[2] The eurozone countries are heavily involved in cross-border production chains, and their participation in GVCs is large compared to the world average, and significantly higher than in the US and China (ECB 2019, p.12). In addition, euro-area countries are more involved in regional than in global GVCs. As a result, an increasing share of foreign value added is emerging in euro-area exports (ECB 2017, p.77) so that the countries are competing with each other at certain stages of global value creation. The participation in GVCs can increase foreign demand elasticity outside the eurozone.

Prior to the outbreak of 2008 global financial crisis the current account balance in the euro-area countries varied between countries. The indebted euro-area countries (Greece, Portugal, Spain, Ireland) which run large current account deficits entered an adjustment process after the 2008 financial crisis. The deficits that had been accumulated prior to the outbreak of the crisis have led to these countries having large external indebtedness. Efstathiou and Wolff (2017) studied the adjustment processes in these countries for the period 2007–2016 and found that in most countries adjustment was achieved through exports, while imports remained flat or increased more slowly. According to their survey, the trade balance of Spain and Portugal was adjusted by about 8.5 percentage points of GDP, exports contributed 8 and 11 percentage points respectively, while the contribution of imports (increase in imports) was 0.5 and −2.5 percentage points, respectively. In Ireland, according to the same survey, exports contributed to the improvement of the trade balance by 60pp. and imports by 43pp. Only in Greece

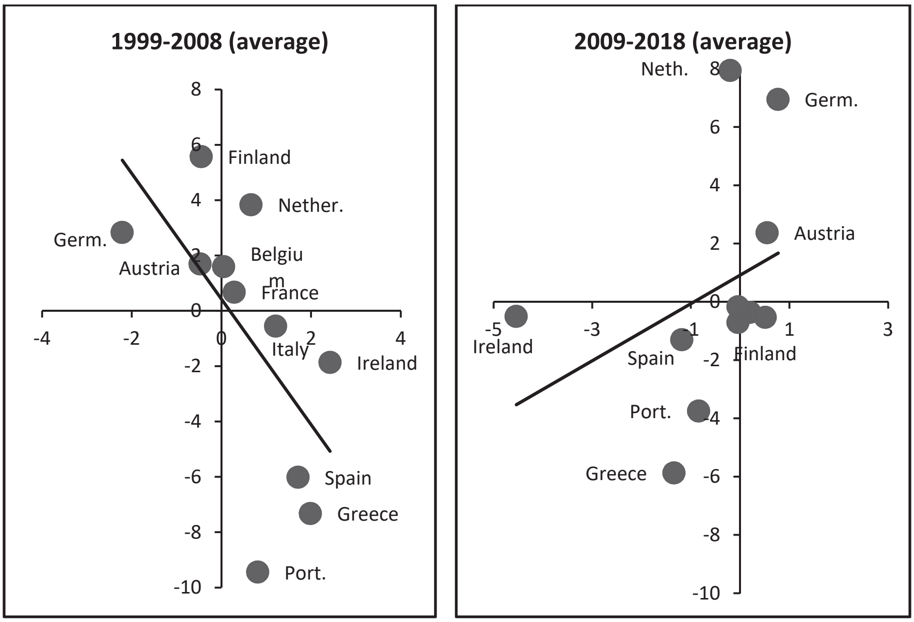
2 Although the growth of global value chains has slowed since the global financial crisis of 2008–2009, it has not actually been interrupted. Data for 2017 show that complex GVCs grew faster than GDP (Li, Meng, and Wang 2019, p.9).

has a massive contraction in imports contributed to reduction in the deficit, while exports did not play a significant role in the adjustment. Thus, the research shows that export growth is the main channel of current account adjustment in the indebted euro-area countries.

The export competitiveness of euro-area countries is significantly influenced by labour costs, which are recorded in the development of their current account balances (Figure 1).

The countries which experienced improvements in their current account (as % of GDP) for the period 1999–2008 were also the countries where the unit labour costs (ULC) were stagnant (Germany, Austria, Finland), while the countries with a rise in ULC experienced worsening current accounts (Greece, Portugal, Spain, Ireland, and Italy). In the post-crisis period (2009–2018), Greece, Spain, and Portugal reduced labour costs but still had a current account deficit, which was lower than before the outbreak of the crisis.

**Figure 1:** Current account and unit labour costs

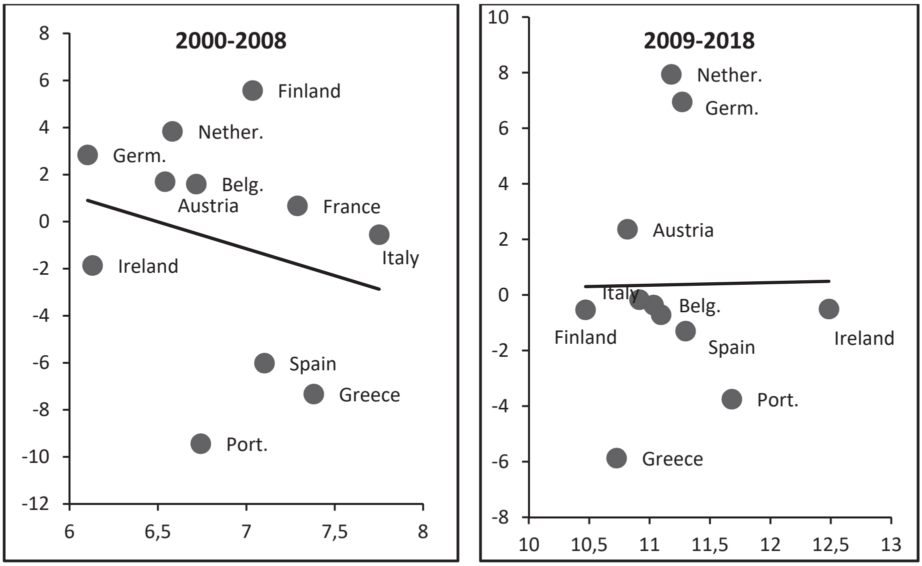



**Note**: X axis – unit labour costs (year on year); Y axis – current account (GDP percentage points).
**Source**: Autor's calculation based on Eurostat data.

The increase in exports has significantely contributed to the reduction of the current account deficit in most euro-area countries. The relationship between exports and the current account balance of the selected euro-area countries is given in Figure 2.

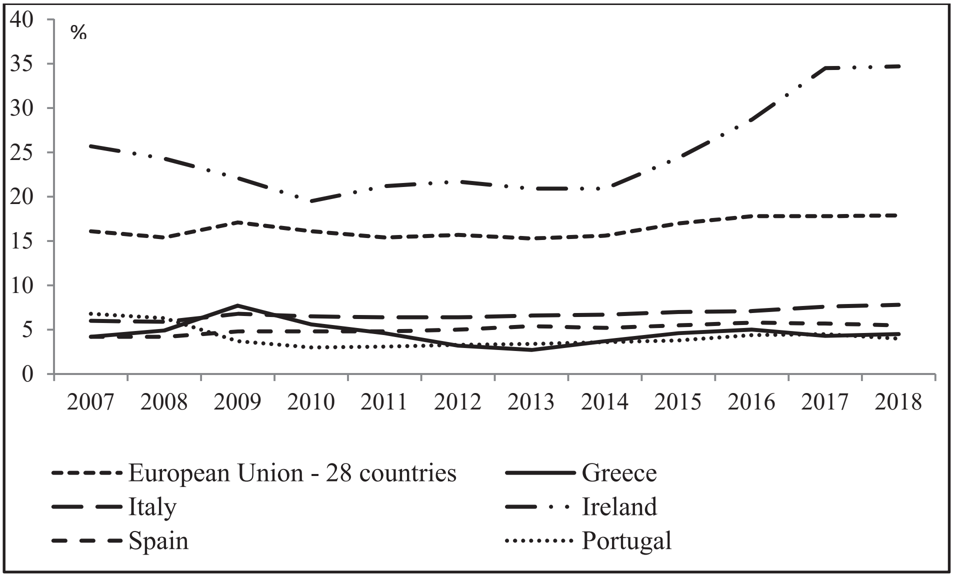
According to Figure 2, countries that lost cost competitiveness in the period before 2008 had a significant increase in volume exports of goods and services (the costs expressed through unit labour costs are given in Figure 1). At the same time, these countries accumulated a larger current account deficit (as % of GDP) than the other countries. In the period 2009–2018 the current account balance of all periphery EU members improved. In some countries, such as Greece, this improvement was the result of a reduction in imports due to a reduction in domestic demand.

**Figure 2:** Current account and exports




**Note**: X axis – exports of goods and services, chain-linked volumes, index 2010=100 (cumulative); Y axis – current account (GDP percentage points) (average).
**Source**: Author's calculation based on Eurostat data.

**Figure 3:** High-tech exports – Exports of high technology products as a share of total exports




**Note**: The data shows the share of exports of all high technology products in total exports. High technology products are defined according to SITC Rev. 4 (2007) as the sum of the following products: Aerospace, Computers–office machines, Electronics–telecommunications, Pharmaceuticals, Scientific instruments, Electrical machinery, Chemicals, Non-electrical machinery, Armaments. The total exports do not include intra-EU trade. Data for years prior to 2007 is not available.
**Source**: Author's calculation based on Eurostat data.

The structure of merchandise exports significantly influences the export dynamics and trends of certain euro-area countries (see Figure 3). Figure 3 shows that in the export of high-technology products the peripheral euro-area members are lagging behind the EU average, as the main market for their merchandise exports.

According to Figure 3, the peripheral EU members have a lower share of high-tech products in their exports than the EU average (the exception is Ireland, which has a higher share). This means that products with a lower level of technological intensity dominate in the exports of peripheral countries. This export structure suggests that exporters from these countries are facing strong export competition from low-income countries (mostly exports from China).[3] Countries with a higher share of high-tech products than the euro-area average have higher export opportunities in the world market.[4]

The aim of this paper is to examine the impact of price competitiveness and foreign and domestic demand on euro-area merchandise exports. This will determine whether the most indebted euro-area countries divert domestic production into exports when faced with a fall in domestic demand, thereby compensating for the losses resulting from the reduced domestic demand.

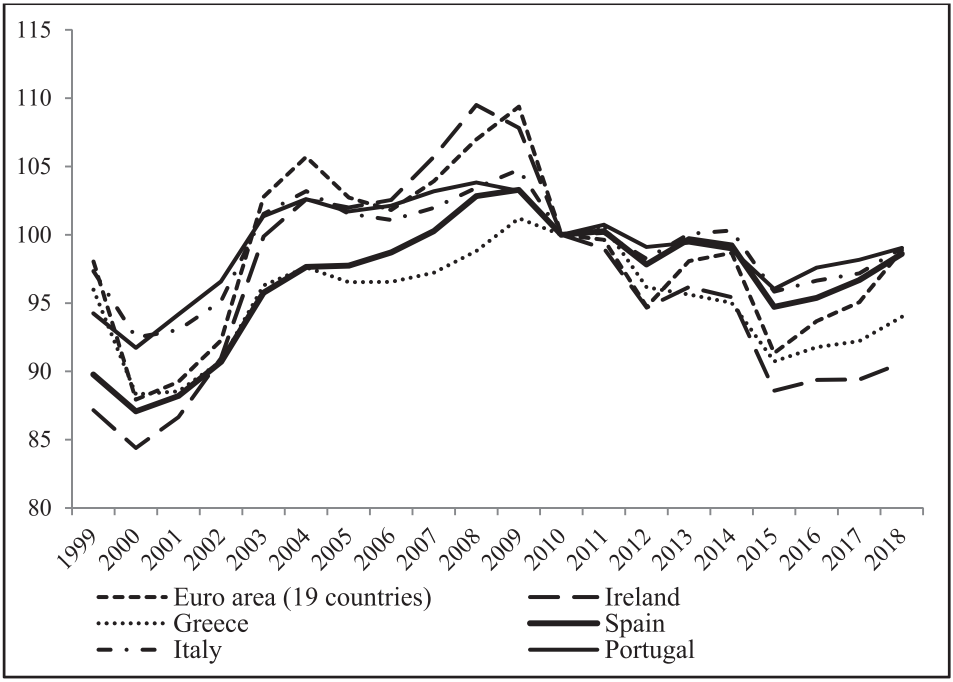
A significant indicator of international price competitiveness is the movement in the real effective exchange rate (REER). The tendencies of the REER of euro-area countries are given in Figure 4. It shows the periphery countries, as they experienced the greatest loss of competitiveness.

Figure 4 shows that the REER indices of peripheral euro-area members had an upward trend after the introduction of the euro. The REER of Italy and Portugal entered the appreciation area in 2003, joined by Ireland in 2004, Spain in 2007, and Greece in 2009. This REER growth trend pushed competitiveness downwards, both in relation to third countries and within the Economic and Monetary Union (EMU). After 2012 all countries (except Italy in 2013 and 2014) experienced REER depreciation. Greece and Ireland had the best competitiveness performance. The growth trend of this index has been registered in all the countries since 2015. REER based on export unit values has the same trend, appreciating steadily in Greece and Italy by some 20% between 2000 and 2003. Moving into the REER appreciation area meant a loss in price competitiveness of exports (see ECB 2005, p.15).

3 Wierts, van Kerkhoff and de Haan (2011) find that the euro area countries (core EMU countries) have a higher technological intensity of exported goods and lower export elasticity with respect to the real exchange rate, while the opposite occurs in peripheral members. Higher technological intensity provides the opportunity for higher profit margins.

4 The ECB (2019) study analyses the dynamics of the share of total euro-area merchandise exports in world merchandise exports. The research finds that in the period 2000–2014 the old members lost and the new members increased global market share. The changes in market share are similar for gross exports and exports based, on the value added concept.

**Figure 4**: Real effective exchange rate of eurozone peripheral members




**Note**: Real effective exchange rate (deflator: consumer price index – 42 trading partners – industrial countries) Index, 2010=100. Double export weights are used to calculate REERs, reflecting not only competition in the home markets of the various competitors, but also competition in export markets elsewhere. A rise in the index means a loss of competitiveness.
**Source**: Author's calculation based on Eurostat data.

The deep recession in 2009 led to a deterioration in the eurozone's macroeconomic performance. There was a physical decline in euro-area merchandise exports in 2009. Although recovery in all countries had already taken place in 2010, euro-area merchandise exports failed to return to the pre-crisis trend track. The euro-area economies hardest hit by the crisis managed to increase their physical export volumes in the post-crisis period, thereby reducing their current account deficits (Ireland, Greece, Cyprus). Estonia, Lithuania, Latvia, and Slovenia also achieved a dynamic increase in the volume of their goods exports.

Our paper aims to examine the impact of relevant factors on euro-area merchandise exports. Given that the eurozone was hit by a deep recession in 2009 and that there was a structural break in the time series representing merchandise exports and economic growth – as with most other variables, in our paper we will estimate the parameters for two periods: the period before the crisis in 2008 (1999–2008), and the period after the crisis (2009–2018). The appearance of a structural break in the time series in 2009 will be modelled by the introduction of a 'step' dummy variable, since there has been a change in the intercept and slope of the time series trend. This dummy variable will also include the impact of the sovereign debt crisis, which began with the collapse of Iceland's banking system in 2008 and then spread to Portugal, Italy, Ireland, Greece, and Spain in 2009.

Although there is a large body of literature that analyses export demand equations, this paper expands the analysis of export competitiveness in the euro-area countries. It contributes to the existing literature by including an additional regressor in the export equation – the share of ICT product exports in the total merchandise exports of each euro-area member – as well as the traditional explanatory variables: price and income. In addition, this paper follows Esteves and Rua (2013) in including domestic demand as an additional explanatory variable. The empirical analysis in this paper uses panel methodology. This approach is less sensitive to outliers and the multicollinearity between explanatory variables (Baltagy 2008). In this paper we look at two models that are estimated using FMOLS and DOLS estimators (see Phillips and Hansen 1990). The analysis in this paper covers the period 1999–2018. The main contribution of this work is that it uses panel analysis and quantile regression to assess the impact of selected export factors on different points of merchandise exports' conditional distribution in euro-area countries. To the best of the author's knowledge, this paper is the first attempt at analysing export performance in euro-area countries using quantile regression.[5]

The rest of the paper is structured as follows. A more detailed literature review is given in section 2. The methodology and data for the empirical analysis in this

5 Fugazza (2004) analyses the determinants of export performance for several countries by quantile regression. Wagner (2004) performs a microeconometric analysis of foreign trade by quantile regression in the case of the influence in the industrial sector of German companies' characteristics on export activities.

paper are described in section 3. Section 4 presents the empirical results and their interpretation. Section 5 summarises the main results and limitations and gives suggestions for further research.

## 2. LITERATURE REVIEW

Anderton, di Mauro, and Moneta (2004) analyse the elasticity of extra-euro-area exports, assessing the impact of demand and relative prices. They calculate that an 1% increase in foreign demand leads to a 1% increase in extra-euro-area export volumes in the long run, while a 1% improvement in relative export price competitiveness increases export volumes by 0.5%. Developments in the trade balance significantly affect the euro zone current account. Marsh and Tokarick (1996) assess the impact of three indicators of price and cost competitiveness (real exchange rate deflated by consumer prices, export unit values of manufacturing goods, and normalised unit labour costs in manufacturing) on the trade flows of G7 member countries.

Many papers assess the external competitiveness of the eurozone, based on the development of the real effective exchange rate and foreign demand. Bayoumi, Harmsen, and Turunen (2011) assess the impact of several variants of the real effective exchange rate (REER)[6] on manufactured eurozone exports (REER based on wholesale/producer price indices, the consumer price index, and unit labour costs in manufacturing).[7] The obtained elasticity coefficients with REER differ in terms of statistical significance and range between –0.6 and 0.01. The elasticity of export volume to foreign demand is estimated to be in the range between 1.8 and 1.9, and the coefficients are highly significant in each specification.[8]

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6 More details on the methodology of calculating REER can be found in Zanelo and Desruelle (1997).

7 For a more complete analysis of export competitiveness, see the ECB (2005). For an overview of the impact of non-price factors on the export competitiveness of the euro area see Di Mauro and Forster (2008). The export unit value-based REERs are particularly sensitive to changes in export structure (see more details at Silver, 2007).

8 Berthou and Dhyne (2018) analyse the heterogeneous effects of competitiveness within the euro area and find that the elasticity of export prices varies significantly across firms with low and high productivity, so that the distribution of productivity across firms in different countries can affect how changes in competitiveness affect external balance and other macroeconomic variables.

Ca' Zorzi and Schnatz (2007) estimate alternative export equations in which they use several different cost and price competitiveness euro-area indicators. Although each indicator has different advantages and disadvantages, the authors estimate that improving cost and price competitiveness by 1% is associated with an increase in extra-euro-area export volumes in the range of 0.3%–0.4% for most indicators. The exception is the indicator of relative export prices, where the reaction was stronger and amounted to 0.6%. These authors also find that a 1% growth in foreign demand generally implies a growth in extra-euro-area export volumes in the range of 0.75%–0.80%. This finding suggests that the share of the euro area in world exports declined in the analysed period (1992–2006). The deviation of export elasticity from the unit in relation to foreign demand was the least when export prices were used as an indicator of competitiveness.

Böing and Stadtmann (2016) empirically assess the impact of competitiveness, measured through unit labour costs, on the euro-area current account. They estimate a panel of annual data for 2000–2013. They find that higher unit cost increases lead to a reduction in the current account balance. Splitting the growth of unit costs into wage growth and productivity growth, these authors find that wage growth has a negative impact on the current account, while productivity growth has a positive impact. They consider that wage cuts are relatively ineffective in dealing with a current account deficit, as productivity growth triggers wage growth.

Staehr and Vermeulen (2019) analyse the impact of competitiveness on the macroeconomic performance of 11 euro-area countries. The VAR models for individual countries are estimated using quarterly data for the period 1995Q4 to 2013Q4. In addition to unit labour costs as a competitiveness indicator, the model includes GDP, current account, and domestic credits. Empirical analysis shows that unit labour cost changes do not have the power to explain current account balance or domestic credit in most countries. In addition, the effects of unit labour costs vary greatly between euro-area countries. On the basis of the obtained results the authors conclude that it is not possible to formulate a general policy measure for current account correction, but country-specific characteristics must be taken into account.

Baumann and di Mauro (2007) estimate that the globalisation of the economy has a significant impact on the adjustment in euro-area trade. They find a positive impact of the increasing euro-area specialisation in some medium–high and high-tech sectors, such as the pharmaceutical industry, where strong demand and dynamic productivity growth have been observed, while a decline in specializsation has been observed in lower-tech sectors such as textiles. The authors also find that euro-area imports of semi-finished goods from low-cost countries are increasing in order to reduce production costs and to focus on the higher value added stages of the production chain.

Allard, Catalan, Everaert, and Sgherri (2005) find that the characteristics of the external sector varied considerably between the four largest euro-area countries in the period 2001–2004. The paper econometrically assesses the contribution of domestic and foreign demand and cost and price competitiveness to the evolution of trade volumes (exports and imports) during the observed period. Although it has been established that these factors have a long-term impact on the differences between countries, significant unexplained residuals remain. On the export side, all countries benefitted from the growth of global demand, with Spain profiting the most and France the least. All countries were burdened by the appreciation of the real exchange rate, with Italy suffering the most and Germany the least.

Eltető (2018) comes to the conclusion that the fall in domestic demand due to the international crisis forced companies to export to foreign markets. This has increased the importance of exports, which has contributed to increasing countries' openness. The paper aims to analyse trends in the previous decade in order to determine whether the long-term structural and regional changes in exports have been realised. The analysis focuses on Iberian, Baltic, and Visegrád countries. The results show that the redirection of trade towards non-EU areas was temporary, while the production structure of exports remains largely similar to that before the crisis. The author demonstrates that the share of high-tech products in exports depends on the activities of foreign multinational companies and not on domestic innovations and R&D developments.

Tressel and Wang (2014) analyse the results of external and internal rebalancing in indebted euro-area countries. These authors find clear evidence of relative price and current account adjustments, but limited evidence of internal reallocation of production from non-tradable to tradable sectors. In addition, they find that export performance improvement depends significantly on external demand, including demand from within the euro area. The prolonged weakness of demand in the euro area has contributed significantly to the slowdown in the export growth of countries such as Italy and Portugal, which are more dependent on intra-euro-area demand. The authors also conclude that current account adjustment is partly influenced by cyclical factors.

## 3. DATA AND METHODOLOGY

### 3.1. Data

In this paper the empirical analysis of eurozone members' merchandise exports was carried out using the annual panel time series for 19 euro-area members. The data used is from the EUROSTAT and UNCTAD databases for the period 1999–2018. The text below presents the basic variables of the model.

*Merchandise exports (EXPVOL)*. The data for euro-area merchandise exports are chain-linked volumes, index 2010=100. We focused on merchandise exports to look for opportunities to increase export competitiveness. The export volume indices respond directly to changes in price competitiveness. The data are taken from Eurostat.

*Real Effective Exchange Rate (REER)*. The paper uses three panel time series for REER. The first, REERCPI, is the real effective exchange rate when the consumer price indices relative to a panel of 42 leading eurozone trading partners from the industrial country group are used as the deflator [9] (double export weights are used to calculate REERs, reflecting not only competition in the home markets of the various competitors, but also competition in export markets elsewhere). The data are expressed as an index with base year 2010. The country's price competitiveness relative to its main competitors in foreign markets is estimated based on REER movement. An increase in the REER index above 100 indicates a loss in export competitiveness. The data are downloaded from Eurostat.

9 The panel of 42 countries includes EU-28 Member States and 14 other industrial countries: Australia, Canada, United States, Japan, Norway, New Zealand, Mexico, Switzerland, Turkey, Russia, China, Brazil, South Korea, and Hong Kong.

The second indicator of real effective exchange rate, REERULC, represents REER based on unit labour cost in the total economy (ULCT).[10] The country's cost competitiveness relative to its principal competitors in international markets is assessed based on the movement of REERULC. An increase in this REER index above 100 also indicates a loss in export competitiveness, while a decrease in REER (depreciation) means gains in cost competitiveness. The data for this variable were downloaded from Eurostat.

The price competitiveness index variable RELATIVEPRICE is equivalent to the real effective exchange rate that is calculated using export prices. It is calculated by dividing the unit value of each eurozone member's exports by the unit value of world exports, indices 2010 = 100. This is a kind of overflow index and shows the price competitiveness of exports. A negative sign is expected in front of this indicator. The data are downloaded from the UNCTADStat database.

*Foreign demand (WORLDGDP).* This variable uses world GDP data, indices 2010 = 100, constant prices (the parameter $\beta 2$ in the export equation is expected to be positive, indicating that exports rise as world income increases). Data are downloaded from the UNCTADStat database.

*Internal demand (FINALDEM).* This variable represents internal demand as the volume of final consumption expenditure, 2010 = 100. The sign in front of this variable is expected to be negative, as the growth of internal demand keeps goods in the domestic market. Data for this variable are from EUROSTAT.

[10] This indicator for REER is calculated as a weighted average for 37 eurozone trading partners (IC37 = EU28 + 9 other industrial countries (Australia, Canada, United States, Japan, Norway, New Zealand, Mexico, Switzerland, and Turkey). The unit labour costs of the total economy include the costs of producing non-tradable goods. They need to be taken into account because changes in their prices also affect decisions about redirecting production to the export sector. The limitation of this performance indicator is that it does not include other costs such as distribution costs, taxes, etc.

*Share of ICT products in total merchandise exports of euro-area countries (ICTEXPORT).* These shares are converted into indexes, base 2010 = 100, and transformed in logarithms.[11] The data are taken from UNCTATstat.

*Dummy variable (DUMMY).* As the financial crisis since 2008 has significantly worsened the eurozone's economic performance, the dummy variable in our export equation equals 1 after the year 2008, and 0 otherwise.

### 3.2. Methodology

The purpose of this section is to determine which factors affect the exports of euro-area countries, using panel data regression (annual data for the 19 euro-area countries in the period 1999–2018). The applied methodology follows the approach developed by Goldstein and Khan (1985) and Bayoumi, Harmsen, and Turunen (2011). Our aim is to assess the importance of the structural component of merchandise exports in euro-area countries as a non-price competitiveness factor.

#### 3.2.1. Regression specification

The assessment of the export equation in our study provides an insight into the impact of the selected explanatory variables on euro-area merchandise export volumes. The volumes of exported merchandise are the dependent variable. The explanatory variable are three sets of data for the real effective exchange rate, and indicators which represent world export demand for euro-area goods, the internal demand of each euro-area member, and the share of high technology products in the merchandise exports of euro-area members. The impact of the 2008 crisis on euro-area merchandise exports is included in our model through the DUMMY variable. The time series of merchandise exports has a structural break in 2009; therefore, we estimate the coefficients of the regression model for two subperiods, 1999–2008 and 2009–2018. All variables in the model except the dummy variable are shown in the indexes 2010 = 100 and have been transformed into logarithms. The paper uses panel analysis. The estimated coefficients in the model are the elasticity coefficients. We used the dependent variable in our model

11 According to OECD methodology, 93 products at the 6-digit level of the harmonised system code, 2012 revision, have been classified in the ICT sector. See: http://unctad.org/en/PublicationsLibrary/tn_unctad_ict4d02_en.pdf Retrieved 12.03.2019.

as an explanatory variable with one lag,[12] while the independent variables are used without lags. The model was constructed according to Comunale and Hessel (2014). The lagged dependent variable enables the process dynamics to be controlled. A similar approach to constructiong the model can be found in recent work (e.g., Bayoumi, Hansen, and Turunen 2011). Like most economic phenomena, exports are a dynamic category in which its previous values affect the current value. Therefore, to empirically analyse export performance in the euro-area countries we use a dynamic panel model, expressed by the following equation:

$$\begin{aligned} logEXPVOL_{i,t} &= \alpha_i + \delta_i logEXPVOL_{i,t-1} + \beta_1 logREER_{i,t} + \beta_2 logWORLDGDP_{i,t} + \\ &\beta_3 logFINALDEM_{i,t} + \beta_4 logICTEXPORT_{i,t} + \beta_5 DAMMY_{i,t} + \varepsilon_{it} \end{aligned},$$

$$i=1,..., N, t=1,..., T \quad (1)$$

where N represents the number of euro area countries; T is the number of periods; $EXPVOL_{i,t}$ is the value of the dependent variable (in our case euro-area merchandise exports) of country *i* in period *t*; parameter α is a constant;

$EXPVOL_{i,t-1}$ is a one-year lagged dependent variable; $\delta$ is the export parameter; REER, WORLDGDP, FINALDEM, and ICTEXPORT are independent variables;

$\beta_1$, $\beta_2$, ..., $\beta_5$ are parameters of the exogenous variables; and ε is an error term. It is assumed that all independent variables are strictly exogenous and uncorrelated with any one $\varepsilon_{it}$.

### 3.2.2. Estimation strategy

Before we estimated a regression model we checked the stationarity of the panel time series. As the literature (e.g., Bayoumi, Harmsen, and Turunen 2011) suggests that panels with aggregate trade data are non-stationary, we examined whether the conditions for applying cointegration in our panel had been met.

[12] The Akaike Information Criterion (AIC) and Bayesian Information Criterion (BIC) are applied to select a lag order. Several different lag combinations were applied in the export equation, and the difference between the estimated coefficients is not significant. Accordingly we decided to show the estimated variables with one lag.

### 3.2.3. Panel unit root tests

In order to check the stationarity of the panel time series we performed the following unit root tests: (1) Levin-Lin-Chu test (LLC) (2002), (2) Im-Pesaran-Shin test (IPS) (2003), (3) Breitung test (2000), (4) Fisher-ADF and Fisher-PP tests (see Maddala and Wu 1999 and Choi 2001), and (5) Hadri test (2000). It is necessary to emphasize that the LLC, Breitung, and Hadri tests assume a common unit process, while the IPS, Fisher-ADF, and Fisher-PP tests assume an individual unit root process. Depending on whether these unit root tests revealed that the panel time series had a unit root, we checked for the existence of co-integration.

### 3.2.4. Cointegration tests and regression model estimation

In our study we applied several cointegration tests: Pedroni (2004), Kao (1999), and the Johansen-Fisher Panel Cointegration Test (see Maddala and Wu 1999). There are two versions of the Pedroni residual cointegration test, which differ from each other according to alternative hypotheses. One hypothesis is that there is a homogeneous cointegration in the panel (a unique equilibrium for all panel units), while the other hypothesis claims that there are different equilibrium relations across countries (heterogeneous cointegration). Both variants of the test have the ability to use individual effects or individual effects and trends as a deterministic component in the cointegration equation. Using the Pedroni test, 11 test statistics were used to test the null hypothesis that claims that there are no cointegration links among the variables in the model.

The Pedroni test statistics were divided into two groups: ‘group’ and ‘panel’ statistics. The ‘group statistics’ category contained three test statistics, representing the average values of the autoregressive coefficients, which were estimated individually for each country in the panel. Hence, these statistics are also referred to as ‘between-dimension’ statistics. These are Group rho-Statistic, Group PP-Statistic, and Group ADF-Statistic. Unlike group statistics, the panel statistics are ‘within-dimension’ statistics, where a common autoregressive coefficient is estimated. These are Panel v-Statistic, Panel-PP statistic, Panel rho-Statistic, and Panel ADF-Statistic. The Kao Cointegration test and Johansen Fisher Panel Cointegration test were applied as supplemental tests to check the robustness of the findings. The Pedroni and Kao tests are based on Engle-Granger (1987) two-step (residual-based) cointegration tests. The Fisher test is a combined Johansen test. In order to estimate the cointegration equation we used FMOLS and DOLS estimators (see Phillips and Hansen 1990).

The impact of the explanatory variables on euro-area merchandise exports was also estimated using quantile regression. The advantage of this approach is that it is possible to estimate the impact of the explanatory variables at different points of the conditional distribution of merchandise exports. In this way it is possible to see the potential heterogeneous effect of the selected determinants on merchandise exports across different quantiles. In this study we used the quantile regression model based on Koenker (2005). Our quantile regression model is estimated with econometric software package EViews 11, and the setting and theoretical overview of the quantile function in the next section of the paper is presented according to the Eviews 11 Background for Quantile regression (http://www.eviews.com/help/helpintro.html#page/content%2Fquantreg-Background.html%23). In order to estimate coefficients for this model, the least absolute deviations (LAD) estimator was included, which allows fitting the conditional median of the response variable.

#### 3.2.5. The quantile function

The Quantile regression model (Koenker 2005) assumes that $Y$ is a random variable with probability distribution function

$$F_y = Prob\left(Y \leq y\right) \tag{2}$$

so 0<τ<1, where the τ-th quantile of $Y$ can be defined as the lowest value of $y$ that meets the condition $F_y \geq \tau$ :

$$Q_r = inf\left\{y : F_{(y)} \geq \tau\right\} \tag{3}$$

Assuming that there is a set of $n$ observations for $Y$, the traditional empirical distribution function is given as:

$$F_n\left(y\right) = n^{-1}\sum_{i=1}^{n} I\left(Y_i \leq y\right) \tag{4}$$

that is

$$F_n(y) = \sum_k 1(Y_i \leq y) \tag{5}$$

where 1(z) is a function indicator that takes the value of 1 if the argument $z$ is true, and 0 otherwise. The empirical quantile is given by:

$$Q_n(\tau) = inf\{y : F_n(y) \geq \tau\} \tag{6}$$

or equivalently, in the form of a simple optimisation problem:

$$Q_n(\tau) = argmin_\xi \left\{ \sum_{i:Y_i \geq \xi} \tau |Y_i - \xi| + \sum_{i:Y_i < \xi} (1-\tau)|Y_i - \xi| \right\} = argmin_\xi \{\sum_i \rho_\tau (Y_i - \xi)\} \tag{7}$$

where $\rho_\tau(u) = u(\tau - 1(u < 0))$ is a so-called check function which weighs positive and negative values asymmetrically.

The quantile regression extends this simple formulation to the regressor $X$. A linear specification for the conditional quantile of dependent variable $Y$ is given values for the p-vector of the explanatory variables $X$:

$$Q(\tau | X_i, \beta(\tau)) = X_i^{'}\beta(\tau) \tag{8}$$

where $\beta(\tau)$ is a vector of the coefficient that relates to the τ quantile. An analog to the above unconditioned quantile minimization is the conditional quantile regression estimator.

$$\hat{\beta}_n(\tau) = argmin_{\beta(\tau)} \{\sum_i \rho_\tau (Y_i - X_i^{'}\beta(\tau))\} \tag{9}$$

The quality of the quantile regression model can be assessed using goodness-of-fit criteria. The Koenker and Machado (1999) definition of goodness-of-fit statistics for quantitative regression is analogous to $R^2$ from conventional regression analysis. Starting from the linear quantile specification $Q(\tau | X_i, \beta(\tau)) = X_i^{'}\beta(\tau)$ we accept the assumption that we can split data and coefficient vector as

$X_i = \left(1, X_{i1}^{'}\right)'$ and $\beta(\tau) = \left(\beta_0(\tau), \beta_1(\tau)'\right)'$, so that

$$Q\left(\tau \middle| X_i, \beta(\tau)\right) = \left(\beta_0(\tau)\right) + X_{i1}^{'}\beta_1(\tau). \tag{10}$$

Then we can define:

$\hat{V}(\tau) = min_{\beta(\tau)} \sum_i \rho_\tau (Y_i - \beta_0(\tau) - X_{i1}^{'}\beta_1(\tau))$, that is

$$\hat{V}(\tau) = min_{\beta_0(\tau)} \sum_i \rho_\tau (Y_i - \beta_0(\tau)) \tag{11}$$

the minimized objective functions. The Koenker and Machado goodness-of-fit criterion is given by:

$$R^1(\tau) = 1 - \hat{V}(\tau) / V(\tau) \tag{12}$$

This statistic is an obvious substitute for conventional

$R^2 . R^1(\tau)$ is located between 0 and 1, and shows the relative success of the model in fitting the data for the $\tau -$th quantile.

Koenker and Bassett (1982) propose testing whether the slopes between quantiles are the same as a robust test of heteroskedasticity. The null hypothesis is defined as:

$$H_0 : \beta_1(\tau_1) = \beta_1(\tau_2) = \ldots = \beta_1(\tau_K) \tag{13}$$

which imposes $(p-1)(K-1)$ restrictions on the coefficients. Based on this, corresponding Wald statistics, distributed as $\chi^2_{(p-1)(K-1)}$, can be obtained .

Newey and Powell (1987) constructed a symmetry test, which can be applied to the quantile regression. The assumption of this test is that if the distribution of $Y$ for the given $X$ is symmetric, then:

$$\frac{\beta(\tau)+\beta(1-\tau)}{2}=\beta\left(\frac{1}{2}\right). \tag{14}$$

This restriction can be evaluated using the Wald test on the quantile process. Suppose that there is an odd number, *K*, of sets of estimated coefficients ordered by $\tau_k$.

It is assumed that the mean value $\tau_{(K+1)/2}$ is equal to 0.5, and that the remaining τ are symmetric around 0.5, with $\tau_j = 1-\tau_{K-j+1}$, for *j* = 1, ..., (*K*-1)/2. Then the null hypothesis can be set up as:

$$H_0 : \frac{\beta(\tau_j)+\beta(\tau_{K-j-1})}{2}=\beta\left(\frac{1}{2}\right) \tag{15}$$

for *j* = 1, ..., (*K*-1)/2.

The Wald test defined for this null hypothesis is equal to zero under the null hypothesis of symmetry. The null has *p*(*K*-1)/2 restrictions, so the Wald statistics are distributed as $\chi^2_{p(K-1)/2}$. Newey and Powell point out that if it is known in advance that the errors are independent and identically distributed but possibly asymmetric, it is possible to restrict the null to apply the restriction for the intercept only. This restrictive assumption imposes only (K-1)/2 constraints on the process coefficients.

We expect that the applied quantile approach will give a more complete explanation of the impact of the selected determinants on euro-area real merchandise exports.

## 4. EMPIRICAL RESULTS AND DISCUSSION

In order to check the stationarity of the panel time series we applied the following unit root tests: LLC, IPS, ADF-Fisher Chi-square, PP-Fisher Chi-square, and the Hadri test. It was useful to apply several different unit root tests because of the short time series. Testing was performed at the level of the time series and their first differential, and in the test equations two options were used, constant and constant with trend. The results of the unit root tests are given in Table I.

The results of the unit root tests show that most variables are not stationary at level, where the statistical significance is 5%. However, the first differential test showed all variables to be stationary, both with a constant and with a constant and trend. Thus, the model variables are integrated of order 1 ((Xt~I (1)). Only the Hadri test did not confirm the stationarity of the first difference series in every case. This was to be expected because this test often rejects the null hypothesis of stationarity.

Since the test results confirme that the panel time series are integrated of order 1, the next step was to determine whether there was cointegration between exports as the dependent variable and the explanatory variables. To check the existence of cointegration we used the following tests: Pedroni residual cointegration test, Kao residual cointegration test, and the Johansen Fisher panel cointegration test. We used three indicators of price competitiveness (LREERULC, LREERCPI, and LRELATIVEPRICE).[13] For each of these we will show separate cointegration tests with other variables that are common to all three cointegrating equations.

Table 1 gives the results of 11 Pedroni statistics, systematised in two groups. The first group contains three test statistics, which belong to the option of heterogeneous cointegration. The other tests belong to the category of homogeneous cointegration. In the Pedroni test we used intercept and trend as a deterministic component. In Table 1, LREERULC is used as an indicator of price competitiveness.

Based on 6 out of 11 Pedroni test statistics, the null hypothesis of no cointegration was rejected at a standard 5% significance level. Although a large number of Pedroni test statistics confirm the existence of cointegration, the other tests need to be applied in order to check the robustness of this assessment. Therefore, the results of the Kao Residual Cointegration Test and the Johansen Fisher Panel Cointegration Test are given in Table 1. Although the Kao test does not confirm the existence of cointegration, the Johansen Fisher cointegration test confirms the existence of cointegration at a statistical significance level of 1%, according to

13 The first letter L in the name of each variable is a substitute for Log.

Trace and Maximum Eigenvalue statistics. Bearing in mind that most of the applied cointegration tests confirmed the presence of cointegration in the given combination of macroeconomic variables, the conditions for estimating the specified cointegration equation were met. However, before this assessment we checked for cointegration in the remaining two specifications.

**Table 1**: Panel Cointegration Tests

| **Pedroni Residual Cointegration Test**<br>Series: *LEXPVOL LREERULC LWORLDGDP LFINALDEM LICTEXPORT*<br>*(The first letter (L) in the series labels means log data)*<br>Null Hypothesis: No cointegration<br>Trend assumption: Deterministic intercept and trend | | | | |
|---|---|---|---|---|
| **Alternative hypothesis: common AR coefs. (within-dimension)** | | | | |
| | **Statistic** | **Prob.** | **Weighted Statistic** | **Prob.** |
| Panel v-Stat. | 0.694578 | 0.2437 | 0.857558 | 0.1956 |
| Panel rho-Stat. | 3.240170 | 0.9994 | 2.891053 | 0.9981 |
| Panel PP-Stat. | −2.209156 | 0.0136 | −3.804713 | 0.0001 |
| Panel ADF-Stat. | −3.552880 | 0.0002 | −4.767459 | 0.0000 |
| **Alternative hypothesis: individual AR coefs. (between-dimension)** | | | | |
| | **Statistic** | **Prob.** | | |
| Group rho-Statistic | 4.233365 | 1.0000 | | |
| Group PP-Statistic | −5.771874 | 0.0000 | | |
| Group ADF-Statistic | −5.383473 | 0.0000 | | |

<table>
<tr><td colspan="5">Kao Residual Cointegration Test<br>Series: LEXPVOL LREERULC LWORLDGDP LFINALDEM LICTEXPORT<br>Null Hypothesis: No cointegration<br>Trend assumption: No deterministic trend</td></tr>
<tr><td colspan="2" rowspan="2">ADF</td><td colspan="2">t-Statistic</td><td>Prob.</td></tr>
<tr><td colspan="2">–0.646134</td><td>0.2591</td></tr>
<tr><td colspan="5">Johansen Fisher Panel Cointegration Test<br>Series: LEXPVOL LREERULC LWORLDGDP LFINALDEM LICTEXPORT<br>Null Hypothesis: No cointegration<br>Trend assumption: Linear deterministic trend<br>Lags interval (in first differences): 11</td></tr>
<tr><td>Hypothesized No. of CE(s)</td><td>Fisher Stat.* (from trace test)</td><td>Prob.</td><td>Fisher Stat.* (from max-eigen test)</td><td>Prob.</td></tr>
<tr><td>None</td><td>690.0**</td><td>0.0000</td><td>449.8**</td><td>0.0000</td></tr>
<tr><td>At most 1</td><td>361.5**</td><td>0.0000</td><td>229.1**</td><td>0.0000</td></tr>
<tr><td>At most 2</td><td>174.5**</td><td>0.0000</td><td>117.5**</td><td>0.0000</td></tr>
<tr><td>At most 3</td><td>96.34**</td><td>0.0000</td><td>88.00**</td><td>0.0000</td></tr>
<tr><td>At most 4</td><td>55.04***</td><td>0.0363</td><td>55.04***</td><td>0.0363</td></tr>
</table>

**Note**: * Probabilities are computed using asymptotic Chi-square distribution. ** Test statistics are significant at the 1% level. *** Test statistics are significant at the 5% level. Automatic lag length selection based on SIC with a max lag of 2. Newey-West automatic bandwidth selection and Bartlett kernel.
**Source**: Author's calculation.

The variable LREERCPI is used as an indicator of price competitiveness in Table 2. The existence of cointegration with other macroeconomic variables was checked.

**Table 2**: The Panel Cointegration Tests

| **Pedroni Residual Cointegration Test**<br>Series: *LEXPVOL LREERCPI LWORLDGDP LFINALDEM LICTEXPORT*<br>Null Hypothesis: No cointegration<br>Trend assumption: Deterministic intercept and trend | | | | |
|---|---|---|---|---|
| **Alternative hypothesis: common AR coefs. (within-dimension)** | | | | |
| | **Statistic** | **Prob.** | **Weighted Statistic** | **Prob.** |
| Panel v-Stat. | 0.252583 | 0.4003 | 0.192495 | 0.4237 |
| Panel rho-Stat. | 3.736720 | 0.9999 | 3.500357 | 0.9998 |
| Panel PP-Stat. | 0.268766 | 0.6059 | −2.937141 | 0.0017 |
| Panel ADF-Stat. | −0.529324 | 0.2983 | −4.159706 | 0.0000 |
| **Alternative hypothesis: individual AR coefs. (between-dimension)** | | | | |
| | **Statistic** | **Prob.** | | |
| Group rho-Statistic | 4.573796 | 1.0000 | | |
| Group PP-Statistic | −6.00584 | 0.0000 | | |
| Group ADF-Statistic | −5.63157 | 0.0000 | | |

| **Kao Residual Cointegration Test**<br>Series: *LEXPVOL LREERCPI LWORLDGDP LFINALDEM LICTEXPORT*<br>Null Hypothesis: No cointegration<br>Trend assumption: No deterministic trend | | |
|---|---|---|
| ADF | t-Statistic | Prob. |
| | −0.59063 | 0.2774 |

| **Johansen Fisher Panel Cointegration Test**<br>Series: *LEXPVOL LREERCPI LWORLDGDP LFINALDEM LICTEXPORT*<br>Null Hypothesis: No cointegration<br>Trend assumption: Linear deterministic trend<br>Lags interval (in first differences): 11 | | | | |
|---|---|---|---|---|
| **Hypothesized No. of CE(s)** | **Fisher Stat.* (from trace test)** | **Prob.** | **Fisher Stat.* (from max-eigen test)** | **Prob.** |
| None | 693.6** | 0.0000 | 430.2** | 0.0000 |
| At most 1 | 366.2** | 0.0000 | 229.8** | 0.0000 |
| At most 2 | 179.6** | 0.0000 | 121.5** | 0.0000 |
| At most 3 | 98.78** | 0.0000 | 87.95** | 0.0000 |
| At most 4 | 62.26** | 0.0078 | 62.26** | 0.0078 |

**Note**: * Probabilities are computed using asymptotic Chi-square distribution. ** Test statistics are significant at the 1% level. Automatic lag length selection based on SIC with a max lag of 2. Newey-West automatic bandwidth selection and Bartlett kernel.
**Source**: Author's calculation.

According to 4 of the 11 Pedroni test statistics, the null hypothesis of no cointegration among the observed panel time-series is rejected at the standard level of 5% significance. The Kao Residual Cointegration Test does not confirm cointegration, but the Johansen Fisher Panel Cointegration Test rejects the null hypothesis of no cointegration among the observed variables, with a statistical significance of 1%. Since most tests confirm the presence of cointegration among the observed panel time series, a cointegration vector can be estimated.

The third indicator of price competitiveness in this paper is the LRELATIVEPRICE variable. In order to check whether cointegration between this and other variables exists, we ran the same cointegration tests as in the previous two cases. The results are given in Table 3.

**Tabel 3**: Panel Cointegration Tests

| **Pedroni Residual Cointegration Test**<br>Series: *LEXPVOL LRELATIVEPRICE LWORLDGDP LFINALDEM LICTEXPORT*<br>Null Hypothesis: No cointegration<br>Trend assumption: Deterministic intercept and trend | | | | |
|---|---|---|---|---|
| **Alternative hypothesis: common AR coefs. (within-dimension)** | | | | |
| | **Statistic** | **Prob.** | **Weighted Statistic** | **Prob.** |
| Panel ν-Stat. | 0.395507 | 0.3462 | 1.621133 | 0.0525 |
| Panel rho-Stat. | 3.935694 | 1.0000 | 3.353359 | 0.9996 |
| Panel PP-Stat. | 0.873267 | 0.8087 | −3.036286 | 0.0012 |
| Panel ADF-Stat. | −0.787708 | 0.2154 | −4.105030 | 0.0000 |
| **Alternative hypothesis: individual AR coefs. (between-dimension)** | | | | |
| | **Statistic** | **Prob.** | | |
| Group rho-Statistic | 4.364813 | 1.0000 | | |
| Group PP-Statistic | −8.12922 | 0.0000 | | |
| Group ADF-Statistic | −6.47853 | 0.0000 | | |

| **Kao Residual Cointegration Test**<br>Series: *LEXPVOL LRELATIVEPRICE LWORLDGDP LFINALDEM LICTEXPORT*<br>Null Hypothesis: No cointegration<br>Trend assumption: No deterministic trend | | |
|---|---|---|
| ADF | t-Statistic | Prob. |
| | −0.857543 | 0.1956 |

| **Johansen Fisher Panel Cointegration Test**<br>Series: *LEXPVOL LRELATIVEPRICE LWORLDGDP LFINALDEM LICTEXPORT*<br>Null Hypothesis: No cointegration<br>Trend assumption: Linear deterministic trend<br>Lags interval (in first differences): 11 | | | | |
|---|---|---|---|---|
| **Hypothesized No. of CE(s)** | **Fisher Stat.* (from trace test)** | **Prob.** | **Fisher Stat.* (from max-eigen test)** | **Prob.** |
| None | 734.5** | 0.0000 | 455.3** | 0.0000 |
| At most 1 | 415.3** | 0.0000 | 255.7** | 0.0000 |
| At most 2 | 208.8** | 0.0000 | 158.4** | 0.0000 |
| At most 3 | 93.21** | 0.0000 | 79.80** | 0.0000 |
| At most 4 | 65.31** | 0.0038 | 65.31** | 0.0038 |

**Note**: * Probabilities are computed using asymptotic Chi-square distribution. ** Test statistics are significant at the 1% level. Automatic lag length selection based on SIC with a max lag of 2. Newey-West automatic bandwidth selection and Bartlett kernel.
**Source**: Author's calculation.

As in the first two specifications, the results in Table 3 show that several Pedroni cointegration tests rejected the null hypothesis of no cointegration, the Kao test does not confirm cointegration, but the Johansen Fisher test confirmed the existence of cointegration with a statistical significance of 1%.

Since most of the tests confirmed the presence of cointegration, the conditions for estimating the cointegration vector have been met. The long-term relationship between the macroeconomic variables in the starting model for all three specifications was assessed using FMOLS and DOLS estimators. The deep recession in the eurozone in 2009 caused the appearance of structural breaks in the time series of the macroeconomic variables, so we estimated the starting export equation for two subperiods, 1999–2008 and 2009–2018, and for the full 1999–2018 period. The results are given in Table 4.

**Table 4**: Export equation of euro area (dependent variable: log of merchandise export volume) – FMOLS and DOLS estimates

| Subperiod 1999–2008 | | | | | | |
|---|---|---|---|---|---|---|
| | ULC | | CPI | | RELATIVEPRICE | |
| | FMOLS | DOLS | FMOLS | DOLS | FMOLS | DOLS |
| **LEXPVOL(–1)** | 0.63(0.00) | 0.66(0.00) | 0.63(0.00) | 0.66(0.00) | 0.64(0.00) | 0.67(0.00) |
| **LREER** | –0.13(0.00) | –0.14(0.28) | –0.03(0.32) | –0.04(0.78) | 0.07(0.00) | 0.09(0.48) |
| **LWORLDGDP** | 0.58(0.00) | 0.54(0.00) | 0.58(0.00) | 0.54(0.00) | 0.57(0.00) | 0.51(0.00) |
| **LFINALDEM** | 0.40(0.00) | 0.40(0.00) | 0.31(0.00) | 0.32(0.01) | 0.29(0.00) | 0.29(0.02) |
| **LICTEXPORT** | 0.11(0.00) | 0.11(0.00) | 0.11(0.00) | 0.11(0.00) | 0.11(0.00) | 0.11(0.00) |
| **R-squared** | 0.953 | 0.959 | 0.953 | 0.959 | 0.953 | 0.959 |
| | | | | | | |
| **Subperiod 2009–2018** | | | | | | |
| | ULC | | CPI | | RELATIVEPRICE | |
| | FMOLS | DOLS | FMOLS | DOLS | FMOLS | DOLS |
| **LEXPVOL(–1)** | 0.79(0.00) | 0.78(0.00) | 0.78(0.00) | 0.78(0.00) | 0.77(0.00) | 0.77(0.00) |
| **LREER** | –0.37(0.00) | –0.37(0.00) | –1.01(0.00) | –1.04(0.00) | –0.49(0.00) | –0.51(0.00) |
| **LWORLDGDP** | 0.50(0.00) | 0.53(0.00) | 0.44(0.00) | 0.45(0.00) | 0.53(0.00) | 0.53(0.00) |
| **LFINALDEM** | –0.17(0.00) | –0.18(0.20) | –0.13(0.00) | –0.15(0.26) | –0.29(0.00) | –0.31(0.02) |
| **LICTEXPORT** | 0.01(0.15) | 0.01(0.76) | 0.03(0.00) | 0.02(0.47) | 0.02(0.07) | 0.01(0.66) |
| **R-squared** | 0.895 | 0.895 | 0.900 | 0.900 | 0.894 | 0.894 |
| | | | | | | |
| **Full 1999–2018 period** | | | | | | |
| | ULC | | CPI | | RELATIVEPRICE | |
| | FMOLS | DOLS | FMOLS | DOLS | FMOLS | DOLS |
| **LEXPVOL(–1)** | 0.88(0.00) | 0.87(0.00) | 0.90(0.00) | 0.87(0.00) | 0.89(0.00) | 0.86(0.00) |
| **LREER** | –0.25(0.00) | –0.23(0.00) | –0.27(0.00) | –0.21(0.03) | –0.07(0.13) | –0.04(0.65) |
| **LWORLDGDP** | 0.36(0.00) | 0.41(0.00) | 0.39(0.00) | 0.45(0.00) | 0.41(0.00) | 0.47(0.00) |
| **LFINALDEM** | 0.09(0.03) | 0.14(0.09) | –0.01(0.78) | 0.04(0.64) | –0.09(0.01) | –0.03(0.69) |
| **LICTEXPORT** | 0.05(0.00) | 0.06(0.00) | 0.05(0.00) | 0.05(0.00) | 0.04(0.00) | 0.05(0.00) |
| **DUMMY** | –0.04(0.00) | –0.06(0.00) | –0.05(0.00) | –0.06(0.00) | –0.06(0.00) | –0.07(0.00) |
| **R-squared** | 0.932 | 0.939 | 0.930 | 0.938 | 0.929 | 0.937 |

**Note**: Panel method: Pooled estimation; Cointegration equation deterministics: Constant (C); Coefficient covariance computed using default method; Long-run covariance estimates (Bartlett kernel, Newey-West fixed bandwidth) used for coefficient covariances; First-stage residuals use heterogeneous long-run coefficients (FMOLS); Trend is additional regressor in FMOLS; p-values are given in brackets.
**Source**: Author's calculation.

Most of the estimated coefficients of the export equation in Table 4 have the expected sign and are statistically significant at the level of 1%. Since the equation has a logarithmic form, the estimated coefficients represent the elasticity.

Table 4 shows that the lagged dependent variable is statistically significant at the 1% level. The elimination of this variable would cause autocorrelation in the residuals.

The results of the estimation of the cointegration equation for the subperiod 19992008 show that ULC- and CPI-based REER are better indicators of price competitiveness than relative export prices (Table 4). The estimated coefficients of export elasticity in relation to the changes in the real effective exchange rate (ULC- and CPI-based REER) have similar values and the expected sign. On the basis of the estimated coefficient with ULC-based REER (FMOLS estimation), a real effective appreciation of 1% leads to a decrease in real goods exports by 0.13%, or by 0.14% according to DOLS estimation (the DOLS-estimated coefficient is not statistically significant at the standard level). However, the estimated coefficient for CPI-based REER is not statistically significant. Also, the export prices (DOLS estimation) are not statistically significant at the standard level, which is small in value and has the opposite sign to that predicted by the theory.[14] Therefore, it is an inferior indicator of price competitiveness than ULC- and CPI-based REER. When a country changes its merchandise export structure in the direction of high-technology products, the impact of price competitiveness on real exports is reduced. This is in line with the idea that price competitiveness is stronger in low-technology industries.[15]

In the period after 2009 the estimated coefficients of export elasticity for the three REER indicators are higher than in the 1999–2008 subperiod, and all are statistically significant at the 1% level. Also, all have the expected sign. The CPI-based price competitiveness indicator (FMOLS estimation) shows that an effective appreciation of the euro by 1% reduces real euro-area exports by about

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14 This indicator, as the ratio of unit values to exports, is affected by changes in the merchandise export structure and reflects the existence of pricing-to-market strategies.

15 Wierts, van Kerkhoff, and de Haan (2011) have analysed the impact of the structure of merchandise exports on trade imbalances. Based on the regression analysis, they conclude that the elasticity of euro-area merchandise exports with respect to the real exchange rate decreased from –1 to –0.3 along the merchandise spectrum from low-tech to high-tech exports.

1%. The estimated price elasticity coefficients in Table 4 confirm that the overestimated value of REER negatively affects real goods exports, which is particularly evident in the peripheral euro-area members, which have experienced appreciation of the real effective exchange rate.

The foreign demand elasticity of exports is estimated at around 0.44 to 0.53 and appears to be highly significant. It shows the importance of rising foreign demand for euro-area merchandise exports. The importance of export demand is indicated by the fact that the global economic recovery after the 2008 financial crisis contributed significantly to the increase in real exports of goods in the euro area. The resurgence of protectionism in world trade slows down the growth of the world economy and can hurt the eurozone's merchandise exports.

The estimated domestic demand coefficients for the period 1999–2008 are statistically significant but do not have the expected sign. A positive sign is not in line with theoretical expectations, using either the FMOLS or the DOLS estimators. In this case the estimated coefficients are also small in size.

However, our results for internal demand in the 2009–2018 subperiod show that the estimated coefficients have the expected sign, but their statistical significance varies depending on which REER indicator is included in the regression equation. The small estimated coefficients value in all specifications in Table 4 indicates that the impact of this variable on real euro-area exports is modest. If we use CPI-based REER in the export equation specification we find an internal demand elasticity of –0.13 (FMOLS) and in the specification with relative export prices the elasticities are –0.29 and –0.31 (estimated by FMOLS and DOLS, respectively). These estimates are in line with the theoretical approach according to which an increase in internal demand diverts some potential exports to the domestic market. However, the small quantitative value of the estimated coefficients for the variables that represent internal demand indicates that this factor has no great impact on the slowdown in real merchandise exports, nor strongly affects any trade deficits of euro-area members. The specific characteristic of the estimated coefficients for the internal demand variable is high *p*-values of the obtained statistics for ULC- and CPI-based coefficients, which diminishes the statistical significance of the estimated coefficients. Generally, the stronger link between the three different REER indicators and the volume of euro-area commodity exports after 2009 indicates an increase in price competitiveness for euro-area merchandise exports. This is particularly true for the periphery eurozone members that have taken measures to reduce the trade deficit by increasing exports.

In order to understand the significance of structural changes in exports for the increased share of higher technological intensity products, we estimated the elasticity coefficient of a variable that represents the share of information and communication technology (ICT) products in the total merchandise exports of the euro-area countries. The estimated coefficients for this variable have the expected positive sign and are statistically significant at the 1% level for the subperiod 1999–2008. However, the estimated coefficients for the subperiod 2009–2018 are small (0.01–0.03) and are not statisticaly significant.[16] Nevertheless, a positive sign indicates that countries with an increasing share of these products in their merchandise exports can realise an increase in total real goods exports. The bad news for the euro area is that the share of ICT declined from 10.3% of their merchandise exports in 2000 to 4.8% in 2017.[17]

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16 The share of ICT products in the merchandise exports of the leading euro-area countries has been decreasing since 2000. This can be seen in the example of several countries: Germany (2000–8.4%, 2017–5.0%); France (2000–10.8%, 2017–3.9%); Italy (2000–4.4%, 2017–1.9%); Netherlands (2000–17.9%, 2017–11.0%). At the same time, the share of ICT products in the exports of several new euro area members has increased: Latvia (2000–0.9%, 2017–9.3%); Slovakia (2000–3.3%, 2017–16.3%). This data is taken from https://unctadstat.unctad.org/wds/TableViewer/tableView.aspx Retrieved 19.11.2019.

17 UNCTADstat, accessed on 21 March 2019. Today, IT products account for aproximately 15% of world merchandise exports (WTO 2018, p.88). Product classification according to technological export intensity shows that high technology exports as a percentage of euro-area manufactured exports also fell, from 20% in 2000 to 13.9% in 2017. The decline of this share is registered not only in the core euro-area countries (Germany – a decline from 18.6% in 2000 to 13.7% in 2017; France,– a decline from 24.6% in 2000 to 23.5% in 2017) but also in the peripheral members (Italy – a decline from 8.0% in 2000 to 6.8% in 2017 ; Spain – a decline from 8.0% in 2000 to 7.0% in 2017). Data downloaded from: https://data.worldbank.org/indicator/TX.VAL.TECH.MF.ZS?locations=FR-IT-ES-CN Accessed 21.03.2019. The share of high-tech products in core euro-area merchandise exports is significantly higher than the share in peripheral members, which indicates that core countries are less sensitive to changes in the real effective exchange rate as an indicator of price competitiveness. The growth of their merchandise exports is influenced significantly by the structure of exports. On the other hand, the lower value of the technological intensity

In most cases, the estimated coefficients for the whole period from 1999 to 2018 are statistically significant and have the expected signs. The coefficient values vary between the estimated values for the two subperiods and confirm the robustness of the applied regression model.

In REER models that include the role of GVC and are based on the bilateral trade value-added weights, appreciation of the exchange rate against the currencies of the trading partners from which a country imports a large amount of intermediate products can enhance a country's competitiveness by reducing its manufacturing costs. By contrast, depreciation vis-à-vis such trading partners may be detrimental to a country's competitiveness, as it increases the cost of importing intermediate goods. Starting from the methodology developed by Bems and Johnson (2015), conventional euro-area REERs (based on gross value trade weights) and REERs involving the role of GVC were calculated in an ECB (2019) study for the euro area. A high correlation was found between these indicators. Given the small quantitative differences between these effective rates, they send consistent messages regarding eurozone competitiveness.

In most cases the R-squared value (Table 4) was above 0.9 (it was in the interval 0.894–0.959). This suggests that the equation regressors explain a large part of the variability of the dependent variable, while the remaining part is being influenced by unwritten factors. More recently the general opinion has been that non-price factors as determinants of export competitiveness should be included in the export equation, which could increase $R^2$.

In order to check the stationarity of the residuals from the estimated cointegration equation, the same tests were carried out as for unit roots. The results show that the residuals of regression are stationary.

The LLC, ADF Fisher, and PP Fisher tests confirmed the stationarity of the residuals in the estimated cointegration equation at 1% significance, as the effect

coefficient in the peripheral euro-area members, compared to the core countries, indicates that their merchandise exports are influenced by price competitiveness. The declining value of the effective exchange rate in this case can contribute to an increase in their merchandise exports. The difficulty is that the exchange rate can only be reduced if wages are reduced or if productivity in these countries increases.

of the constants was excluded (excluding the constants is the usual procedure when estimating the residual stationarity). The test results conclude that the panel time series are co-integrated. All three export equation specifications confirm that the selected variables have a significant impact on merchandise exports volume in euro-area countries.

The estimated coefficients for the 2009–2018 subperiod can also serve to assess the robustness of the estimated coefficients for the whole 1999–2018 period. This test confirms the robustness of the estimates for the whole sample period (1999–2018). We performed an additional robustness test by repeating the regressions in Table 4 several times, temporarily excluding one country from the sample. In some cases the negative values of the estimated coefficients for the real effective exchange rate were lower than in the full sample, confirming the robustness of estimates for the entire period. The sign of the estimated coefficients for the CPI-based REER variable changed in several cases when the sample changed.

The impact of the selected regressors along the conditional distribution of the euro-area export volumes as the dependent variable was estimated in all export equation specifications using the quantile regression method. Cointegration equations were estimated for the entire 1999–2018 period because the time series were short, so estimates of the impact of the relevant factors on the conditional distribution of euro-area export volumes were more reliable for the whole observed period (Table 5).

**Table 5:** Estimates of real merchandise exports with LEXPVOL as the dependent variable (Panel Least Squares and Quantile Regression Estimates), 1999–2018

**MODEL 1**

| Independent Variable | Panel Least Squares estimate | Quantile Regression Estimates | | | | |
|---|---|---|---|---|---|---|
| | | 0.10 | 0.25 | 0.50 | 0.75 | 0.90 |
| | (1) | (2) | (3) | (4) | (5) | (6) |
| C | –0.766 | –1.420 | –0.413 | –0.930 | –0.548 | –0.340 |
| | (0.011) | (0.001) | (0.063) | (0.000) | (0.218) | (0.682) |
| LEXPVOL(–1) | 0.916 | 0.918 | 1.000 | 0.953 | 0.841 | 0.835 |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) |
| LREERULC | –0.148 | –0.125 | –0.043 | –0.096 | –0.136 | –0.323 |
| | (0.013) | (0.031) | (0.229) | (0.173) | (0.014) | (0.002) |
| LWORLDGDP | 0.406 | 0.598 | 0.173 | 0.293 | 0.282 | 0.477 |
| | (0.000) | (0.000) | (0.001) | (0.000) | (0.005) | (0.046) |
| LFINALDEM | –0.007 | –0.105 | –0.038 | 0.056 | 0.134 | 0.110 |
| | (0.918) | (0.241) | (0.414) | (0.392) | (0.085) | (0.598) |
| LICTEXPORT | 0.016 | 0.025 | 0.003 | 0.009 | 0.015 | 0.004 |
| | (0.123) | (0.030) | (0.572) | (0.463) | (0.413) | (0.851) |
| DUMMY | –0.073 | –0.130 | –0.046 | –0.068 | –0.030 | –0.025 |
| | (0.000) | (0.000) | (0.008) | (0.001) | (0.171) | (0.334) |
| Number of cases | 361 | 361 | 361 | 361 | 361 | 361 |
| R-squared | 0.927 | | | | | |
| Pseudo R-squared | | 0.811 | 0.784 | 0.733 | 0.690 | 0.668 |

**MODEL 2**

| Independent Variable | Panel Least Squares estimate | Quantile Regression Estimates | | | | |
|---|---|---|---|---|---|---|
| | | 0.10 | 0.25 | 0.50 | 0.75 | 0.90 |
| | (1) | (2) | (3) | (4) | (5) | (6) |
| C | –0.648 | –1.289 | –0.431 | –0.952 | –0.956 | 0.959 |
| | (0.096) | (0.028) | (0.065) | (0.000) | (0.145) | (0.590) |
| LEXPVOL(–1) | 0.915 | 0.922 | 0.996 | 0.969 | 0.848 | 0.879 |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) |
| LREERCPI | –0.135 | –0.144 | –0.032 | –0.028 | –0.041 | –0.346 |
| | (0.100) | (0.433) | (0.418) | (0.661) | (0.755) | (0.161) |
| LWORLDGDP | 0.432 | 0.606 | 0.181 | 0.275 | 0.319 | 0.228 |

| | | | | | | |
|---|---|---|---|---|---|---|
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.605) |
| LFINALDEM | −0.073 | −0.136 | −0.049 | −0.005 | 0.079 | 0.075 |
| | (0.244) | (0.273) | (0.278) | (0.917) | (0.245) | (0.759) |
| LICTEXPORT | 0.017 | 0.035 | 0.003 | 0.008 | 0.021 | −0.015 |
| | (0.097) | (0.079) | (0.572) | (0.381) | (0.287) | (0.562) |
| DUMMY | −0.076 | −0.125 | −0.048 | −0.065 | −0.033 | −0.025 |
| | (0.000) | (0.000) | (0.004) | (0.000) | (0.150) | (0.557) |
| Number of cases | 361 | 361 | 361 | 361 | 361 | 361 |
| R-square | 0.927 | | | | | |
| Pseudo R-squared | | 0.811 | 0.784 | 0.731 | 0.686 | 0.664 |

**MODEL 3**

| Independent Variable | Panel Least Squares estimate | Quantile Regression Estimates | | | | |
|---|---|---|---|---|---|---|
| | | 0.10 | 0.25 | 0.50 | 0.75 | 0.90 |
| | (1) | (2) | (3) | (4) | (5) | (6) |
| C | −1.022 | −0.867 | −0.433 | −1.027 | −1.475 | −1.088 |
| | (0.009) | (0.266) | (0.061) | (0.000) | (0.003) | (0.201) |
| LEXPVOL(−1) | 0.904 | 0.926 | 0.984 | 0.967 | 0.858 | 0.848 |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) |
| LRELATIVEPRICE | −0.021 | −0.141 | −0.031 | −0.005 | 0.084 | 0.113 |
| | (0.762) | (0.198) | (0.444) | (0.934) | (0.250) | (0.243) |
| LWORLDGDP | 4.452 | 0.560 | 0.202 | 0.283 | 0.307 | 0.426 |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.044) |
| LFINALDEM | −0.111 | −0.166 | −0.063 | −0.019 | 0.067 | −0.083 |
| | (0.055) | (0.076) | (0.142) | (0.615) | (0.310) | (0.643) |
| LICTEXPORT | 0.015 | 0.013 | 0.008 | 0.010 | 0.022 | −0.034 |
| | (0.138) | (0.144) | (0.193) | (0.299) | (0.252) | (0.142) |
| DUMMY | −0.0080 | −0.131 | −0.048 | −0.065 | −0.024 | 0.037 |
| | (0.000) | (0.000) | (0.005) | (0.001) | (0.323) | (0.186) |
| Number of cases | 361 | 361 | 361 | 361 | 361 | 361 |
| R-square | 0.926 | | | | | |
| Pseudo R-squared | | 0.810 | 0.784 | 0.731 | 0.687 | 0.661 |

**Notes**: Prob-values reported in parenthesis; Quantile regression component: Huber Sandwich Standard Errors & Covariance; Bandwidth method: Hall-Sheather (bw=0.048592 for tau 0.1; 0.094502 for tau 0.25; 0.13645 for Median; 0.094502 for tau 0.75, and 0.048592 for tau 0.9); Sparsity method: Kernel (Epanechnikov) using residuals; Unique optimal solution identified.
**Source**: Author's calculation.

The estimates of quantile regression coefficients for five quantiles (0.10, 0.25, 0.50, 0.75, and 0.90) are given in Table 5. The estimates are presented in three models that differ from each other according to the price/cost variable for calculating the REER. The correlation between the volume of goods exports and the corresponding explanatory variables is estimated in each model. We used empirical results from the regression model, estimated by the Panel Least Squares (PLS) estimator, as a benchmark for comparison.[18] The results are given in column 1 of Table 5. The estimated coefficients are statistically significant, except the estimates for internal demand and the ICT export variable. The application of the PLS estimator implies that the estimated relationship between the dependent variable and the explanatory variables are the same regardless of which conditional distribution point is involved.

Unlike this rather rigid assumption, quantitative regression allows estimating the significance of selected determinants of merchandise exports at different points of the conditional distribution of the dependent variable. The quantile regression was estimated for the five quantiles, and the results are presented in columns 2–6 of Table 5. The values of the estimated coefficients vary across the five selected quantiles, as well as in statistical significance. The estimated coefficients of the time-lagged dependent variable are high, have the expected positive sign, are statistically significant at the 1% level, and gradually decrease from lower to upper quantiles. The signs in front of the coefficients with LREERULC are negative and their value varies across quantiles, as does the level of their statistical significance. The negative impact of this variable becomes stronger over higher quantiles of conditional distribution of merchandise exports. This indicates that the effect of real effective exchange rate appreciation on real merchandise exports gets stronger towards the higher quantiles of conditional distribution of merchandise exports, as the dependent variable in the export equation. Since the ULC represents the total labour costs, the estimated coefficients for the real effective

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[18] We conducted a cross-sectional independence test to check whether the panel data model was cross-sectionally independent. Since *T* in our model is relatively small, we used the asymptotically standard normal Pesaran CD test (Pesaran 2004) for panel models 1, 2, and 3, described in Table 5. The null hypothesis of this test is no cross-section dependence (correlation) in residuals. The test statistics results of 25.193 (p-value of 0.000) for Model 1, 24.292 (p-value of 0.000) for Model 2, and 24.361 (p-value of 0.000) for Model 3 reported by the statistical test means that the Pesaran CD test strongly rejected the null at a significance level of 1%. According to these findings the panel data are cross-sectionally dependent.

exchange rate variable suggest that the growth of these costs leads to appreciation of the exchange rate and deterioration in euro area export competitiveness. Decreasing labour costs in some peripheral members after 2008 led to an increase in their price competitiveness, which encouraged an increase in merchandise exports and a reduction in their trade deficits.[19]

It is characteristic that the estimated coefficients with the internal demand variable have a negative sign in the lower quantiles, while in the upper quantiles they have a positive sign. Apart from changing signs, their statistical significance varies between quantiles. The greatest surprise is the transition of the modest positive impact of Information Technology Services (ITS) on the growth of exports, to the area of negative impact on higher quantiles of the conditional distribution of the dependent variable in Model 2 and Model 3. At the same time, the statistical significance of this coefficient for the upper quantiles is also reduced. The quantile process for all variables in Model 1 is shown in Chart 1 in the Annex.

Also in this paper we tested the hypothesis that the coefficients slope of the estimated model are equal across quantiles using quantile regression. The coefficients in Chart 1 suggest that the assumption of equality of slopes cannot be accepted. In order to arrive at a more precise conclusion we tested with a variant of the Wald test, formulated by Koenker and Bassett (1982). The null hypothesis that the slopes between quantiles are equals was tested ($H_0$: $\beta_1(0.25)=\beta_1(0.50)=\beta_1(0.75)=\beta_1(0.90)$). To save space, we only present the results on whether the slope is the same at the 0.9 quantile as at the median quantile. The test results are reported in Table 6.

[19] Greece, Portugal, Ireland, and Italy.

**Table 6**: Quantile Slope Equality Test and Symmetric Quantiles Test (Equation specification from Model 1, Table 5)

| **Quantile Slope Equality Test** | | | |
|---|---|---|---|
| Test Summary | Chi-Sq. Statistic | Chi-Sq. d.f. | Prob. |
| Wald Test | 35.31498 | 6 | 0.0000 |
| **Symmetric Quantiles Test** | | | |
| Test Summary | Chi-Sq. Statistic | Chi-Sq. d.f. | Prob. |
| Wald Test | 13.38980 | 7 | 0.0632 |

**Note**: Estimated equation quantile tau = 0.5; Test quantiles = 0.9; Test statistic compares all coefficients.
**Source**: Author's calculation.

Table 6 shows that the Wald test summary Chi-Sq. statistic value of 49.68559 is statistically significant at the 1% level, so the null hypothesis of slope equality across quantiles is rejected. This finding confirms the conclusion imposed by Chart 1 and serves as proof that the relationship between the explanatory variables and the dependent variable varies across quantile values. This is relevant because it shows that in cases when the research emphasis is on specific quantiles, linear models can lead to inadequate conclusions as to whether there is a link between the explanatory and dependent variables, and if a link exists these models may suggest a wrong conclusion about the strength of the link.

Table 6 also gives the results of the test for symmetry between quantiles. The null hypothesis of this test is that the distribution is symmetric. The test statistic is statistically significant at the 10% level, which shows significant asymmetry and rejects the hypothesis of null symmetry between quantiles. These findings confirm the heterogeneous impact of the explanatory variables on real euro-area merchandise exports.

## 5. CONCLUSION AND POLICY IMPLICATIONS

The empirical research in this paper explores the link between euro-area merchandise exports and selected explanatory variables. Variables representing the impact of price factors, external and internal demand, and the structure of merchandise exports were used as regressors. The starting export equations for the two subperiods 1999–2008 and 2009–2018 and for the full 1999–2018 period were estimated.

The impact of price factors was assessed using several indicators of the real effective exchange rate. The ULC- and CPI-based REER estimates have similar numerical values for the subperiod 1999–2008, with the expected negative sign. The estimated coefficient of elasticity with ULC-based REER (FMOLS estimator) shows that a real effective appreciation of 1% will result in a 0.13% decline in the volume of merchandise exports, while this decline according to the DOLS estimator will be 0.14%. The results of the regression analysis confirm that real effective exchange rate appreciation diminishes the competitiveness of real merchandise exports.

The estimated export elasticity coefficients for REER variables in the 1999–2008 subperiod are smaller than the coefficients in the 2009–2018 subperiod. All estimated coefficients for the REER variable in the period 2009–2018 are statistically significant at the 1% level and have the expected sign. The CPI-based REER ranges between −1.01 (FMOLS estimation) and −1.04 (DOLS estimation). This means that real effective depreciation of the euro reduces euro-area merchandise exports by near 1%. This is particularly pronounced in the periphery euro-area members that face real effective exchange rate appreciation due to domestic wage growth and productivity decline. According to the estimated model, the impact of appreciation was stronger in the 2009–2018 subperiod than in the 1999–2008 subperiod.

The estimated regression coefficients for the real merchandise export equation for the euro area reveal that the impact of foreign demand on real merchandise exports was significantly stronger than the impact of price competitiveness. The recovery of the world economy in the years following the 2008 financial and economic crisis significantly affected the real growth of merchandise exports in euro-area countries. The growing sensitivity of exports to fluctuations in world demand is partly related to structural changes that were encouraged by the globalisation of the world economy due to vertical specialisation.

The estimated regression coefficients for the internal demand variable vary in value and by sign depending on the REER indicator used. However, the small numerical value of this coefficient in all specifications for the subperiod 2009–

2018 shows that domestic demand in the euro area does not have a significant impact on the merchandise exports of these countries.

An attempt to include the impact on real merchandise exports of the structure of merchandise exports by technological intensity was made in an empirically estimated export equation by introducing a variable that represents the share of ICT products in the total merchandise exports of the euro-area countries. In most export equation specifications for the subperiod 2009–2018 the estimated coefficients with this variable are not statistically significant at the standard level (an exception is the specification with CPI-based REER in which the estimated coefficient is statistically significant at the 0.01% level) and have a positive sign. However, their numerical value is negligible and ranges between 0.04 and 0.05. These values are indicative because they show that an increase in the share of technologically intensive products in merchandise exports affects the increase in real merchandise exports. However, the tendency of a reduction in the share of ITS products in euro-area merchandise exports from 10.3% in 2000 to 4.8% in 2017 is not encouraging news for this integration. The core eurozone members (Germany and France) face this tendency as well as the peripheral members.

The regression model was also estimated by quantile regression, thus testing the potential heterogeneity of the regressor's impact at various points on the conditional distribution of real merchandise exports as the dependent variable. The quantile regression coefficients were estimated for five quantiles (010, 025, 050, 075, and 0.90). The coefficient estimates for different quantiles vary in both size and level of statistical significance. The estimated coefficients with the export demand variable are statistically significant at the 1% level for lower quantiles and 5% level for upper quantiles, have the expected positive sign, and are characterised by a decreasing trend from lower to higher quantiles.

The signs of the estimated coefficients with ULC-based REER are negative and the value of the coefficients and the level of their statistical significance varies across quantiles. In this case the real effective exchange rate appreciation shows an increasing negative impact on real merchandise exports with upper quantiles.

Price competitiveness, represented in quantile regression with ULC-based REER, decreases from the lower to the higher quantiles of the conditional distribution of merchandise exports as the dependent variable. Practically, this means that the effect of real effective exchange rate appreciation on real merchandise exports is higher for the upper quantiles of the conditional distribution of merchandise exports. Since the ULC represents total labour costs, it suggests that the growth of these costs in peripheral euro-area members may lead to an appreciation of the real effective exchange rate, with negative effects on real merchandise exports. Inelastic wages and low labour-market flexibility compromise the export competitiveness of peripheral euro-area members.

The models estimated in this paper indicate that euro-area merchandise exports were influenced by price and non-price competitiveness, and that these factors have contributed to peripheral eurozone members reducing current account imbalances.[20] Internal devaluation also works to increase merchandise exports and reduce external imbalances. The synchronisation of market-based reforms in the euro area can significantly reduce transition costs. Labour market reform leads to a reduction in real wages, and the ECB (2012, p.53) findings show that improving price competitiveness in deficit countries can lead to trade rebalancing in the euro area in the medium term.

Protectionist trade policies have decreased the growth of euro-area merchandise exports. The involvement of eurozone members in the GVC and the high openness of the economy also affect exports, as demand for intermediate products is declining. The increased integration in global supply chains has made the eurozone more vulnerable to the slowdown in world trade.

Countries with a public debt-to-GDP ratio of more than 80%[21] are expected to align their public debt and budget deficits with convergence criteria. Strong economic policy coordination is needed at the eurozone level to boost domestic demand and increase merchandise export growth rates. More dynamic growth in domestic demand can enhance the functioning of the euro area and make crisis management easier. The poor internal demand hinders rebalancing between

[20] The models used in the ECB study (2012, p.36) suggest that improving the current account balance by 1% of GDP requires a temporary wage cut (relative to competitors) of about 5% to 10%.

[21] In 2018 the public debt to GDP of Greece, Italy, Portugal, Cyprus, Belgium, Spain, and France was greater than 80%:. The data is from https://ec.europa.eu/eurostat/tgm/table.do?tab=table&init=1&language=en&pcode=sdg_17_40&plugin=1 Accessed 11.12.2019.

eurozone members. Economic growth based on investment in productive areas of the economy is an important prerequisite for maintaining competitiveness in the long run.

Coordination is also needed in the fiscal area. After many years of support from the ECB through the quantitative easing programme to sustain post-crisis economic growth, changes are needed in the fiscal area. Coordinated support for investment through fiscal stimulus could support the eurozone's economic growth and exports after monetary policy normalisation. Interest rates are at a historically low level, so a new stimulus is needed.

Peripheral eurozone members have significantly reduced merchandise deficits in the post-crisis period (Ireland has had a surplus since 1999, Italy switched to a surplus in 2012, and Greece, Spain, and Portugal have significantly reduced their deficit).[22] Improving the export competitiveness of peripheral eurozone members by reducing domestic prices entails greater costs than the devaluation of the national currency, a measure often taken before the introduction of the euro. However, this measure is no longer available to these countries because the euro-area has a single currency. The common monetary policy and the existence of a single currency are not an adequate solution for countries affected by asymmetric shocks. As a rule, the common exchange rate is overestimated for less competitive members and undervalued for more competitive countries. One member's competitiveness is only part of the eurozone’s overall competitiveness, so the common exchange rate reflects the competitiveness of the monetary union as a whole. Restrictive fiscal policy in the peripheral eurozone countries has contributed to a decline in domestic demand in those countries, leading to sharp deceleration.

Although they have significantly reduced their current account deficits (Italy and Ireland are in surplus), the peripheral members of the euro area face the challenge of implementing structural changes in order to increase export competitiveness. Restructuring their exports with a greater share of higher value-added and high

22 The peripheral eurozone members have faced a boom in demand from both private and public sources due to their entry into the eurozone, as well as a fall in interest rates and excess credit (Martin and Philippon 2017). This has led to a rise in unit labour costs and has deteriorated their competitiveness.

technological intensity products would reduce the sensitivity of their exports to price competition, thus avoiding the need for the socially painful measure of basic wage cuts. Also, wage cuts are more feasible in flexible labour markets. Although increasing cost-competitiveness can be useful for increasing merchandise exports it is difficult to sustain in a global market where competition from low-cost economies is increasing. A greater reliance on non-price competitiveness factors would allow peripheral eurozone members to improve export competitiveness by raising product quality rather than reducing prices.

Germany and several other eurozone members (Netherlands, Belgium, Ireland, Italy, Slovenia, and Slovakia)[23] achieved a trade surplus in 2018. These countries are more competitive in exports than the peripheral member countries, which generate a deficit. Thus, there is an imbalance in the eurozone, because some countries have trade deficits and others have surpluses. This shows that convergence is not yet complete and that trade deficit countries are the most vulnerable in crises and are exposed to pressure to make adjustments.

The limitation of this research and similar studies is the use of standard indicators of the real effective exchange rate. Our research, like many other studies, is also based on Armington's assumption (Armington 1969) of the substitutability of all products. In our case, the assumption of elastic substitution is partly justified, since a significant part of euro-area merchandise exports are realised in intra-euro-area trade. However, this assumption should be checked in a more detailed analysis of the merchandise trade where high-value-added exports are compared to low-value-added exports, or export of capital goods to export of consumer products.

Due to the increasing impact of non-tariff and tariff measures on international trade, further analysis of the effects of these measures on the competitiveness of euro-area exports is necessary.

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23 According to EUROSTAT data, these countries achieved the following trade surpluses in 2018: Germany–233.3 billion euro, Netherlands–68.8, Ireland–48.3, Italy–39.3, Belgium–13.1, Slovenia–1.6, and Slovakia–0.4. https://ec.europa.eu/eurostat/tgm/table.do?tab=table&init=1&language=en&pcode=tet00047&plugin=1 Accessed 12.12.2019.

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## APPENDIX

**Table I**. Results of the panel unit root tests of euro area countries for the period 2000–2018

| Variable | Method | | Determin. terms | Statist. | Prob. | Obs. | Non Stationary (NS) Stationary (S) |
|---|---|---|---|---|---|---|---|
| **LEXPVOL** | LLC | Level | Constant | –0.08 | 0.47 | 357 | NS |
| | | | Constant and trend | –1.51 | 0.07 | 353 | NS |
| | | First Diff. | Constant | –14.42 | 0.00 | 339 | S |
| | | | Constant and trend | –13.08 | 0.00 | 339 | S |
| | IPS W-stat. | Level | Constant | 2.56 | 0.99 | 357 | NS |
| | | | Constant and trend | –0.70 | 0.24 | 353 | NS |
| | | First Diff. | Constant | –11.41 | 0.00 | 339 | S |
| | | | Constant and trend | –8.97 | 0.00 | 339 | S |
| | ADF-Fisher Chi-squared | Level | Constant | 29.36 | 0.84 | 357 | NS |
| | | | Constant and trend | 47.20 | 0.15 | 353 | NS |
| | | First Diff. | Constant | 185.36 | 0.00 | 339 | S |
| | | | Constant and trend | 137.33 | 0.00 | 339 | S |
| | PP-Fisher Chi-squared | Level | Constant | 23.42 | 0.97 | 361 | NS |
| | | | Constant and trend | 33.48 | 0.68 | 361 | NS |
| | | First Diff. | Constant | 247.05 | 0.00 | 342 | S |
| | | | Constant and trend | 208.95 | 0.00 | 342 | S |
| | Breitung t-stat. | Level | Constant | – | – | – | – |
| | | | Constant and trend | –1.11 | 0.13 | 334 | NS |
| | | First Diff. | Constant | – | – | – | – |
| | | | Constant and trend | –11.0 | 0.00 | 320 | S |
| | Hadri Z-stat. | Level | Constant | 11.85 | 0.00 | 380 | NS |
| | | | Constant and trend | 6.18 | 0.00 | 380 | NS |
| | | First Diff. | Constant | 0.32 | 0.37 | 361 | S |
| | | | Constant and trend | 3.77 | 0.00 | 361 | NS |
| **LREERCPI** | LLC | Level | Constant | –3.08 | 0.00 | 359 | S |
| | | | Constant and trend | –0.13 | 0.45 | 354 | NS |

| | | | | | | | |
|---|---|---|---|---|---|---|---|
| | | First Diff. | Constant | –10.45 | 0.00 | 338 | S |
| | | | Constant and trend | –7.61 | 0.00 | 333 | S |
| | IPS W – stat. | Level | Constant | –1.90 | 0.03 | 359 | S |
| | | | Constant and trend | –0.11 | 0.45 | 354 | NS |
| | | First Diff. | Constant | –11.94 | 0.00 | 338 | S |
| | | | Constant and trend | –10.37 | 0.00 | 333 | S |
| | ADF-Fisher Chi-squared | Level | Constant | 52.81 | 0.06 | 359 | NS |
| | | | Constant and trend | 40.20 | 0.37 | 354 | NS |
| | | First Diff. | Constant | 192.86 | 0.00 | 338 | S |
| | | | Constant and trend | 196.56 | 0.00 | 333 | S |
| | PP-Fisher Chi-squared | Level | Constant | 46.50 | 0.16 | 361 | NS |
| | | | Constant and trend | 27.88 | 0.86 | 361 | NS |
| | | First Diff. | Constant | 193.32 | 0.00 | 342 | S |
| | | | Constant and trend | 159.75 | 0.00 | 342 | S |
| | Breitung | Level | Constant | – | – | – | – |
| | | | Constant and trend | –0.56 | 0.28 | 335 | NS |
| | | First Diff. | Constant | – | – | – | – |
| | | | Constant and trend | –6.64 | 0.00 | 314 | S |
| | Hadri Z-stat. | Level | Constant | 9.14 | 0.00 | 380 | NS |
| | | | Constant and trend | 8.54 | 0.00 | 380 | NS |
| | | First Diff. | Constant | 1.64 | 0.05 | 361 | NS |
| | | | Constant and trend | 2.04 | 0.02 | 361 | NS |
| **LREERULC** | LLC | Level | Constant | –1.83 | 0.03 | 352 | S |
| | | | Constant and trend | –1.04 | 0.15 | 354 | NS |
| | | First Diff. | Constant | –10.46 | 0.00 | 336 | S |
| | | | Constant and trend | –6.54 | 0.00 | 327 | S |
| | IPS W – stat. | Level | Constant | –0.07 | 0.47 | 352 | NS |
| | | | Constant and trend | 0.69 | 0.75 | 354 | NS |
| | | First Diff. | Constant | –11.13 | 0.00 | 336 | S |
| | | | Constant and trend | –9.84 | 0.00 | 327 | S |
| | ADF-Fisher Chi-squared | Level | Constant | 33.92 | 0.67 | 352 | NS |
| | | | Constant and trend | 36.91 | 0.52 | 354 | NS |

| | | | | | | | |
|---|---|---|---|---|---|---|---|
| | | First Diff. | Constant | 180.69 | 0.00 | 336 | S |
| | | | Constant and trend | 150.12 | 0.00 | 327 | S |
| | PP-Fisher Chi-squared | Level | Constant | 27.53 | 0.90 | 361 | NS |
| | | | Constant and trend | 37.35 | 0.50 | 361 | NS |
| | | First Diff. | Constant | 247.11 | 0.00 | 342 | S |
| | | | Constant and trend | 212.23 | 0.00 | 342 | S |
| | Breitung | Level | Constant | – | – | – | – |
| | | | Constant and trend | −0.81 | 0.21 | 335 | NS |
| | | First Diff. | Constant | – | – | – | – |
| | | | Constant and trend | −8.08 | 0.00 | 308 | S |
| | Hadri Z-stat. | Level | Constant | 8.67 | 0.00 | 380 | NS |
| | | | Constant and trend | 7.97 | 0.00 | 380 | NS |
| | | First Diff. | Constant | 1.18 | 0.12 | 361 | S |
| | | | Constant and trend | 2.51 | 0.01 | 361 | NS |
| **LRELATIVEPRICE** | LLC | Level | Constant | −1.33 | 0.08 | 358 | NS |
| | | | Constant and trend | −1.48 | 0.07 | 356 | NS |
| | | First Diff. | Constant | −12.03 | 0.00 | 340 | S |
| | | | Constant and trend | −9.96 | 0.00 | 340 | S |
| | IPS W – stat. | Level | Constant | −0.98 | 0.16 | 358 | NS |
| | | | Constant and trend | 0.38 | 0.65 | 356 | NS |
| | | First Diff. | Constant | −10.76 | 0.00 | 340 | S |
| | | | Constant and trend | −8.77 | 0.00 | 340 | S |
| | ADF-Fisher Chi-squared | Level | Constant | 40.68 | 0.35 | 358 | NS |
| | | | Constant and trend | 25.70 | 0.94 | 356 | NS |
| | | First Diff. | Constant | 174.26 | 0.00 | 340 | S |
| | | | Constant and trend | 134.70 | 0.00 | 340 | S |
| | PP-Fisher Chi-squared | Level | Constant | 40.25 | 0.37 | 361 | NS |
| | | | Constant and trend | 24.81 | 0.95 | 361 | NS |
| | | First Diff. | Constant | 169.72 | 0.00 | 342 | S |
| | | | Constant and trend | 131.63 | 0.00 | 342 | S |
| | Breitung | Level | Constant | – | – | – | – |
| | | | Constant and trend | −0.93 | 0.18 | 337 | NS |

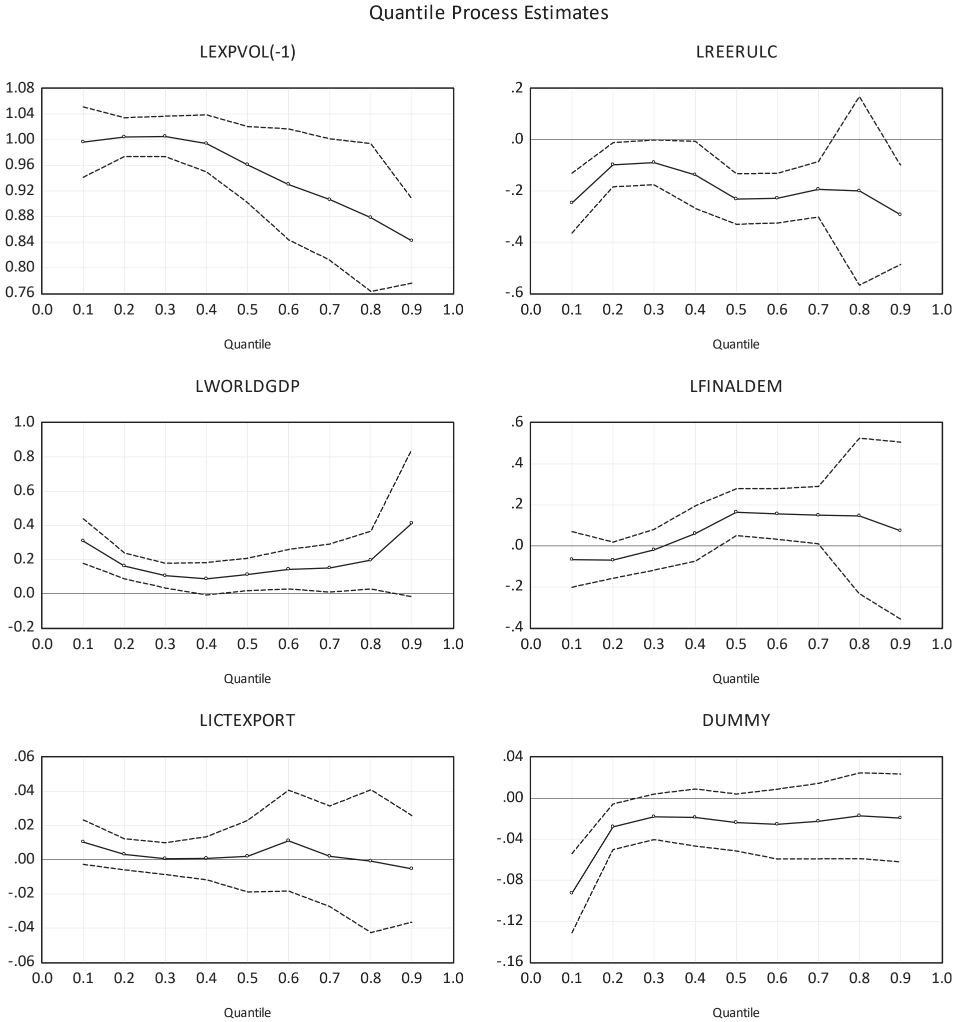
| | | | | | | | |
|---|---|---|---|---|---|---|---|
| | | First Diff. | Constant | – | – | – | – |
| | | | Constant and trend | −11.72 | 0.00 | 321 | S |
| | Hadri Z-stat. | Level | Constant | 6.27 | 0.00 | 380 | NS |
| | | | Constant and trend | 7.25 | 0.00 | 380 | NS |
| | | First Diff. | Constant | 0.05 | 0.58 | 361 | S |
| | | | Constant and trend | 5.23 | 0.00 | 361 | NS |
| **LWORLDGDP** | LLC | Level | Constant | −3.31 | 0.00 | 361 | S |
| | | | Constant and trend | −3.03 | 0.00 | 361 | S |
| | | First Diff. | Constant | −15.58 | 0.00 | 342 | S |
| | | | Constant and trend | −13.69 | 0.00 | 342 | S |
| | IPS W – stat. | Level | Constant | 3.29 | 0.99 | 361 | NS |
| | | | Constant and trend | 0.02 | 0.51 | 361 | NS |
| | | First Diff. | Constant | −11.27 | 0.00 | 342 | S |
| | | | Constant and trend | −7.95 | 0.00 | 342 | S |
| | ADF-Fisher Chi-squared | Level | Constant | 8.91 | 1.00 | 361 | NS |
| | | | Constant and trend | 27.84 | 0.89 | 361 | NS |
| | | First Diff. | Constant | 181.72 | 0.00 | 342 | S |
| | | | Constant and trend | 121.49 | 0.00 | 342 | S |
| | PP-Fisher Chi-squared | Level | Constant | 14.96 | 0.99 | 361 | NS |
| | | | Constant and trend | 32.36 | 0.73 | 361 | NS |
| | | First Diff. | Constant | 186.60 | 0.00 | 342 | S |
| | | | Constant and trend | 123.46 | 0.00 | 342 | S |
| | Breitung | Level | Constant | – | – | – | – |
| | | | Constant and trend | −4.55 | 0.00 | 342 | S |
| | | First Diff. | Constant | – | – | – | – |
| | | | Constant and trend | −11.31 | 0.00 | 323 | S |
| | Hadri Z-stat. | Level | Constant | 13.07 | 0.00 | 380 | NS |
| | | | Constant and trend | 5.99 | 0.00 | 380 | NS |
| | | First Diff. | Constant | −0.03 | 0.51 | 361 | S |
| | | | Constant and trend | 8.38 | 0.00 | 361 | S |
| **LFINALDEM** | LLC | Level | Constant | −3.14 | 0.00 | 341 | S |
| | | | Constant and trend | −1.58 | 0.06 | 344 | NS |

| | | | | | | | |
|---|---|---|---|---|---|---|---|
| | | First Diff. | Constant | –5.24 | 0.00 | 332 | S |
| | | | Constant and trend | –5.31 | 0.00 | 334 | S |
| | IPS W – stat. | Level | Constant | 1.26 | 0.90 | 341 | NS |
| | | | Constant and trend | 0.56 | 0.71 | 344 | NS |
| | | First Diff. | Constant | –4.79 | 0.00 | 332 | S |
| | | | Constant and trend | –3.53 | 0.00 | 334 | S |
| | ADF-Fisher Chi-squared | Level | Constant | 38.75 | 0.44 | 341 | NS |
| | | | Constant and trend | 32.87 | 0.71 | 344 | NS |
| | | First Diff. | Constant | 88.53 | 0.00 | 332 | S |
| | | | Constant and trend | 70.50 | 0.00 | 334 | S |
| | PP-Fisher Chi-squared | Level | Constant | 49.63 | 0.10 | 360 | NS |
| | | | Constant and trend | 21.32 | 0.99 | 360 | NS |
| | | First Diff. | Constant | 85.84 | 0.00 | 341 | S |
| | | | Constant and trend | 72.97 | 0.00 | 341 | S |
| | Breitung | Level | Constant | – | – | – | – |
| | | | Constant and trend | –1.43 | 0.08 | 325 | NS |
| | | First Diff. | Constant | – | – | – | – |
| | | | Constant and trend | –3.99 | 0.00 | 315 | S |
| | Hadri Z-stat. | Level | Constant | 10.72 | 0.00 | 379 | NS |
| | | | Constant and trend | 7.22 | 0.00 | 379 | NS |
| | | First Diff. | Constant | 1.64 | 0.05 | 360 | NS |
| | | | Constant and trend | 3.39 | 0.00 | 360 | NS |
| **LICTEXPORT** | LLC | Level | Constant | –3.10 | 0.00 | 354 | S |
| | | | Constant and trend | 0.18 | 0.57 | 347 | NS |
| | | First Diff. | Constant | –9.82 | 0.00 | 339 | S |
| | | | Constant and trend | –9.09 | 0.00 | 333 | S |
| | IPS W – stat. | Level | Constant | –0.53 | 0.27 | 358 | NS |
| | | | Constant and trend | 1.61 | 0.95 | 347 | NS |
| | | First Diff. | Constant | –7.97 | 0.00 | 339 | S |
| | | | Constant and trend | –6.87 | 0.00 | 333 | S |
| | ADF-Fisher Chi-squared | Level | Constant | 42.71 | 0.28 | 354 | NS |
| | | | Constant and trend | 25.66 | 0.94 | 347 | NS |

| | | | | | | | |
|---|---|---|---|---|---|---|---|
| | | First Diff. | Constant | 129.41 | 0.00 | 339 | S |
| | | | Constant and trend | 109.04 | 0.00 | 333 | S |
| | PP-Fisher Chi-squared | Level | Constant | 31.10 | 0.78 | 361 | NS |
| | | | Constant and trend | 9.76 | 1.00 | 361 | NS |
| | | First Diff. | Constant | 124.28 | 0.00 | 342 | S |
| | | | Constant and trend | 121.30 | 0.00 | 342 | S |
| | Breitung | Level | Constant | – | – | – | – |
| | | | Constant and trend | 0.12 | 0.55 | 328 | NS |
| | | First Diff. | Constant | – | – | – | – |
| | | | Constant and trend | –8.84 | 0.00 | 3.14 | S |
| | Hadri Z-stat. | Level | Constant | 10.10 | 0.00 | 380 | NS |
| | | | Constant and trend | 7.12 | 0.00 | 380 | NS |
| | | First Diff. | Constant | 0.90 | 0.18 | 361 | S |
| | | | Constant and trend | 5.90 | 0.00 | 361 | S |

**Note**: The following abbreviations are used in the table: LLC – Levin-Lin-Chu test; IPS – Im-Pesaran-Shin test. S – Stationary; NS – Non-stationary. The null hypothesis for all tests (except the Hadri test) is a unit root exists. (For the Hadri test, $H_0$ means there is no unit root). The level of significance is 5%. We used the Schwarz information criterion with a maximum of two lags to decide the lag length. Individual effects and both individual effects and individual trends have been used as deterministic components. In this paper we used a balanced panel data set.
**Source**: Author's calculation.

**Figure 1**: Quantile process for Model 1




**Note**: Coefficients plot. The full lines represent the coefficients values over τ levels, while the dashed lines form the areas that are the associated 95% confidence intervals.
**Source**: Author's.