

The Effects of Competitiveness on Trade Balance: The Case of Southern Europe

Oscar Bajo-Rubio, Burcu Berke, and Vicente Esteve

Abstract

According to conventional wisdom, “peripheral” Southern European members of the euro area (Greece, Italy, Portugal and Spain) suffer from a problem of competitiveness. Since their membership of the euro area renders devaluation impossible, adjustment should come through decreasing wages and prices in these countries, which, by improving the trade balance, should lead to a recovery of previous levels of employment and growth. In this paper, the authors estimate trade balance equations for the Southern European countries, both for total trade and for the trade performed with the European Union, taking three alternative measures of the real exchange rate, based on consumption price indices, export prices and unit labour costs, respectively. Their main conclusion is that demand seems to be more relevant than relative prices when explaining the evolution of the trade balance.

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Keywords Trade balance; real exchange rate; competitiveness

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1 Introduction

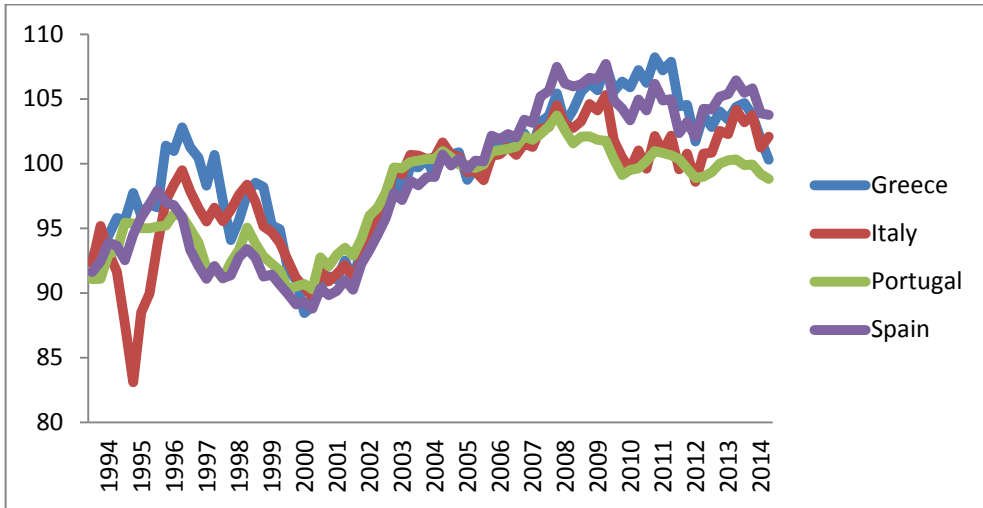
The evolution of real exchange rates is an important issue when analyzing the price competitiveness of a country. Many countries around the world use devaluation as a tool to improve their trade balances, and this happens not only in fixed exchange rate regimes but also under the managed fluctuation regime in existence since 1973. In addition, for a country joining a monetary union (as, e.g., the Economic and Monetary Union of the European Union, EU), real exchange rate developments reflect inflation differentials of that country versus the other members of the monetary union.

In discussions at the EU level regarding the crisis that started in 2008, the conventional wisdom is that “peripheral” Southern European members of the euro area (Greece, Italy, Portugal and Spain) suffer from a problem of competitiveness. Specifically, it has been argued that, prior to the crisis, these countries experienced a greater increase in wages and prices as compared to those of Northern Europe, leading to a loss of competitiveness vis-à-vis the latter. Since their membership of the euro area renders devaluation impossible, it has been argued that adjustment has to come through decreasing wages and prices in Southern Europe. In other words, these countries should apply an “internal devaluation”.

Figures 1a and 1b show the evolution of the competitiveness of Greece, Italy, Portugal and Spain, measured by the real effective exchange rate (REER) computed using consumption price indices (CPIs), vis-à-vis 37 industrialized countries and the EU, respectively, for the period 1994–2014. As can be seen, an appreciation of the real exchange rate (reflecting a higher relative increase in their prices) appears for the four countries throughout the period, somewhat softened in recent years. Notice, however, that such a loss of competitiveness seems to be more nuanced with respect to the EU. Similar conclusions can be derived from Figures 2a and 2b, in which REERs are now computed using export prices instead of total prices, as measured by the CPI, and from Figures 3a and 3b, in which REERs are computed using unit labour costs (ULCs).

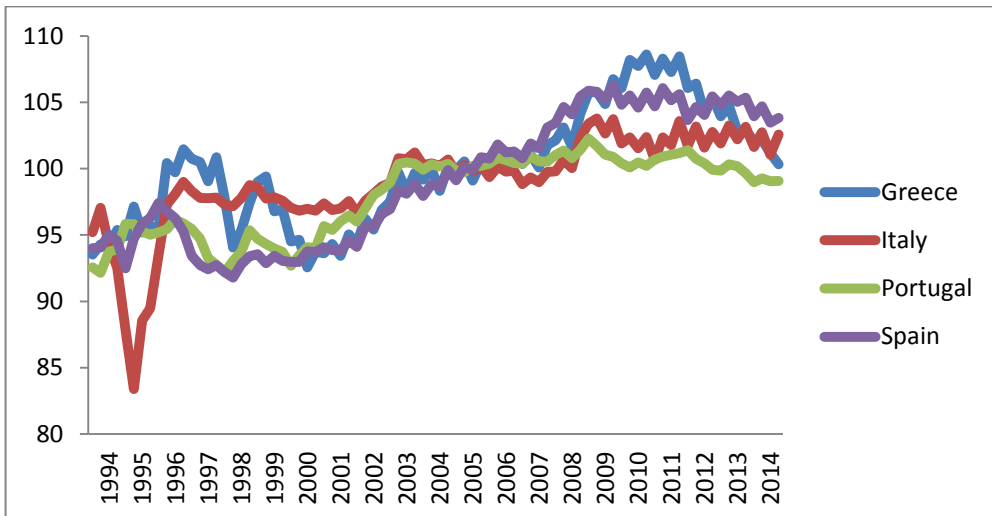
In any case, the resulting policy recommendation is that Southern European countries should implement contractionary fiscal policies aimed to reduce government budget deficits, coupled with structural reforms leading to a lower growth of wages and prices. In fact, the dominant policy stance has been mostly

Figure 1a: Real effective exchange rates computed using consumption price indices, vis-à-vis 37 industrialized countries, 1994–2014 (2005=100)



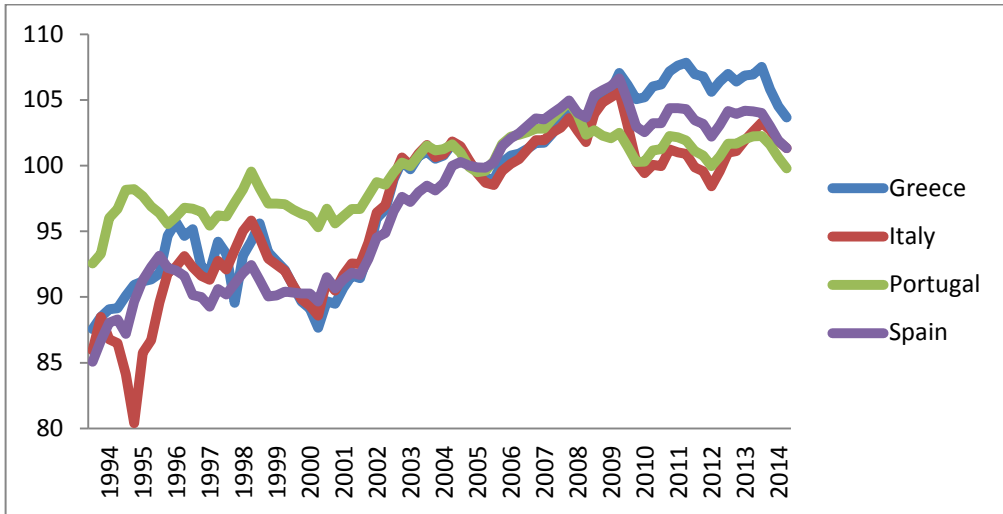
Source: Eurostat.

Figure 1b: Real effective exchange rates computed using consumption price indices, vis-à-vis the European Union, 1994–2014 (2005=100)



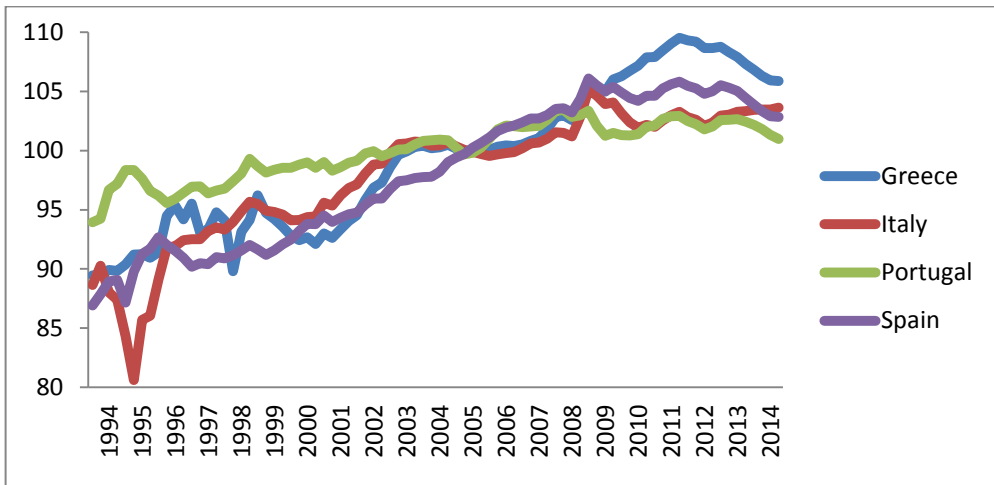
Source: Eurostat.

Figure 2a: Real effective exchange rates computed using export prices, vis-à-vis 37 industrialized countries, 1994–2014 (2005=100)



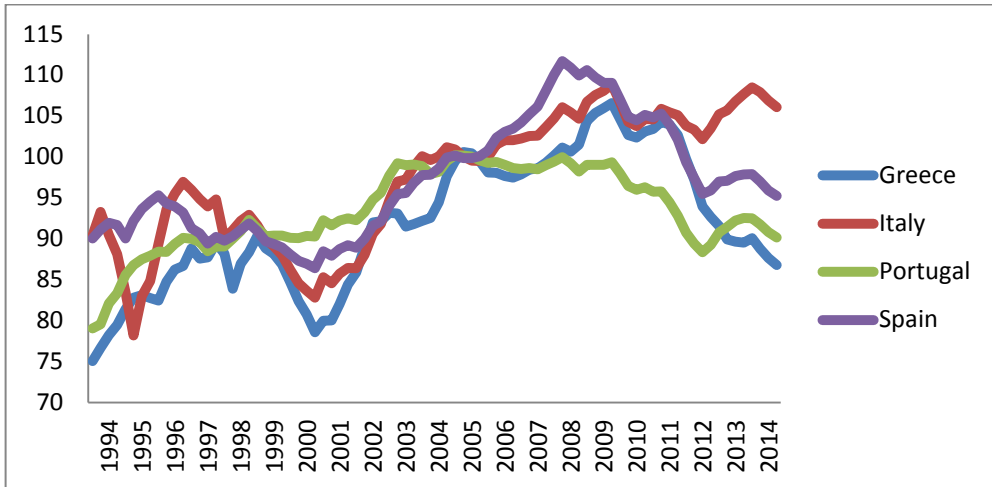
Source: Eurostat.

Figure 2b: Real effective exchange rates computed using export prices, vis-à-vis the European Union, 1994–2014 (2005=100)



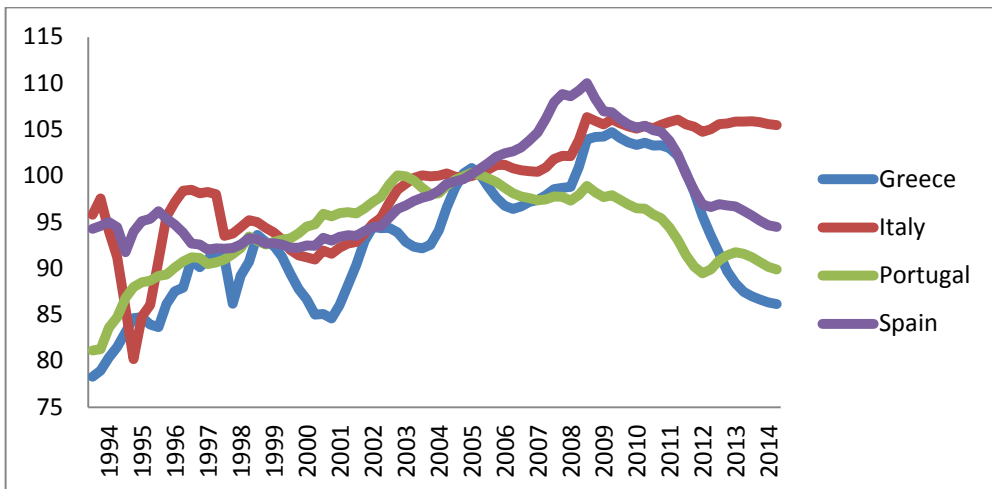
Source: Eurostat.

Figure 3a: Real effective exchange rates computed using unit labour costs, vis-à-vis 37 industrialized countries, 1994–2014 (2005=100)



Source: Eurostat.

Figure 3b: Real effective exchange rates computed using unit labour costs, vis-à-vis the European Union, 1994–2014 (2005=100)



Source: Eurostat.

centred on the labour market and is intended to reduce ULCs, which should result in lower relative prices vis-à-vis the rest of the euro area. By restoring external competitiveness, these measures should allow the trade balance to improve and hence recover the previous levels of employment and growth.

An initial problem with this argument lies in the definition of ULCs at the aggregate level. This is a misleading concept, since ULCs calculated with aggregate data are not simply the sum or a weighted average of the firms' ULCs; if anything, aggregate ULCs reflect the distribution of income between labour and capital (Felipe and Kumar, 2011). In addition, such policies have not been very successful in terms of reducing relative prices, given the simultaneous increase in profit margins (Uxó et al., 2014). Finally, these policies of austerity have resulted in a deeper recession in all Southern European countries, further complicating the situation (De Grauwe and Ji, 2013).

On the other hand, due to the expansion of global value chains, a debate has recently arisen on the possible disconnection between exchange rates and trade. The concept of global value chains (or, in other words, the international fragmentation of production) means that the different stages of the production process of a particular good or service are located across different countries; see Amador and Cabral (2014) for a review of the literature. The increasing participation of firms in global value chains could weaken the relationship between exchange rates and trade, since now an exchange rate depreciation would improve the competitiveness of just a fraction of the value of final exports.

This kind of argument has been suggested to explain why the recent depreciation of the yen has not led to a significant improvement of Japan's trade balance, since Japanese firms would have relocated the production of less differentiated goods to overseas subsidiaries, with the more differentiated goods being produced in Japan (Shimizu and Sato, 2015). From a more general point of view, several recent papers have empirically analyzed the possible disconnection between exchange rates and trade for a number of countries, although the evidence is not fully conclusive: whereas Ahmed et al. (2015) and Ollivaud et al. (2015) found evidence of a decreasing response in recent years of trade volumes to real exchange rate variations, this is not the case for Leigh et al. (2015), who obtained the opposite result. Notice, however, that in the latter paper, the average results for the whole sample are compatible with very different estimated elasticities for individual countries, which occasionally even have the theoretically wrong sign.

The aim of this paper is to assess whether changes in the real exchange rates, as a proxy of external competitiveness, affect the trade balance, therefore making the argument sketched above workable. Specifically, we will estimate equations for the trade balance (total, and with the EU) of the Southern European members of the euro area (Greece, Italy, Portugal and Spain).

In addition, we will also address the issue of the disconnection between exchange rates and trade in recent times by testing for the existence of structural change in the estimated equations, which is complemented with an estimation by subperiods. In this way, we will be able to assess the extent to which the orthodox argument might hold – at least its first step, since even if the real exchange rate significantly influences the trade balance, the latter is only a share of total aggregate demand. In Section 2, the underlying theoretical framework of the paper is discussed, Section 3 presents the empirical results, and Section 4 collects the main conclusions and policy implications.

2 Theoretical framework

We will develop in this section a simple version of the imperfect substitutes model of international trade (Goldstein and Khan, 1985). This approach can be summarized in the following five equations:

$$M(Y, p_m) = X^*(p_m^*) \quad (1)$$

$$X(p_x) = M^*(Y^*, p_x^*) \quad (2)$$

$$p_m = \frac{p_m^*}{q} \quad (3)$$

$$p_x = \frac{p_x^*}{q} \quad (4)$$

$$B = p_x X - p_m M \quad (5)$$

where M and X are the domestic demand for imports and supply of exports; X^* and M^* the foreign supply of exports and demand for imports; p_m and p_x the price of the goods imported at home and the price of domestic exports, relative to the domestic price level; p_m^* and p_x^* the price of foreign exports and the price of the goods

imported abroad, relative to the foreign price level; Y and Y^* the levels of domestic and foreign real income; q the real exchange rate, defined as the price of domestic goods relative to foreign goods; and B the domestic trade balance.

In the above model (see Dornbusch, 1975 for a similar framework in nominal terms), equations (1) and (2) represent the equilibrium conditions in the markets for domestic imports and domestic exports, respectively; Equations (3) and (4) relate the domestic and foreign currency relative prices of goods through the real exchange rate; and Equation (5) defines the home country's trade balance in real terms.

By solving Equations (1) to (4) for the levels of domestic imports and exports and the relative prices p_x and p_m^* , as functions of Y , Y^* and q , the trade balance can be written as a function:

$$B = B(Y, Y^*, q) \quad (6)$$

where:

$$\frac{\partial B}{\partial Y} < 0, \frac{\partial B}{\partial Y^*} > 0, \frac{\partial B}{\partial q} \geq 0$$

Notice that, unlike the cases of changes in domestic and foreign income, the trade balance effect of changes in the real exchange rate is ambiguous. In particular, a depreciation of the real exchange rate, while increasing the value of exports, could lead to either an increase or a decrease in the value of imports, depending on whether its effect on prices or quantities prevails. A depreciation of the real exchange rate will improve the trade balance if the Bickerdike–Robinson–Metzler condition is satisfied, i.e., when:

$$\frac{\partial B}{\partial q} = - \left[\frac{p_x}{q} X(1 + \varepsilon) \frac{\eta^*}{\varepsilon + \eta^*} - \frac{p_m}{q} M(1 - \eta) \frac{\varepsilon^*}{\varepsilon^* + \eta} \right] < 0 \quad (7)$$

or, in terms of the elasticity of the trade balance to the real exchange rate:

$$\frac{\partial B}{\partial q} q = - \left[p_x \frac{X}{B} (1 + \varepsilon) \frac{\eta^*}{\varepsilon + \eta^*} - p_m \frac{M}{B} (1 - \eta) \frac{\varepsilon^*}{\varepsilon^* + \eta} \right] < 0 \quad (7')$$

where ε and ε^* denote the (absolute values of the) domestic and foreign price elasticities of the supply of exports; and η and η^* the (absolute values of the) domestic and foreign price elasticities of the demand for imports. In other words, a depreciation of the real exchange rate will improve the trade balance if the term in square brackets in the above expressions is positive (recall that, given our definition of the real exchange rate, a decrease in q means a depreciation, and an increase in q means an appreciation). In the particular case of infinite supply elasticities and a trade balance initially in equilibrium, the Bickerdike–Robinson–Metzler condition becomes the well-known Marshall–Lerner condition:¹

$$\eta^* + \eta > 1$$

The first empirical studies estimating export and import functions obtained price elasticities that were quantitatively small, although their sum was greater than one in absolute value; in other words, the Marshall–Lerner condition was satisfied so that a depreciation would improve the trade balance (Houthakker and Magee, 1969). Subsequent evidence still found that a depreciation of the real exchange rate improved the trade balance, but with the effect operating with a considerable lag, measured in terms of years (Dornbusch and Krugman, 1976). These lags in the effect of exchange rates on prices and trade volumes means that it is costly to adjust trade flows quickly, which could be explained by the tendency of those firms engaged in international trade to commit themselves to particular suppliers for extended periods of time (Krugman and Baldwin, 1987).

The combination of a short-run deterioration with a long-run improvement in the trade balance (or, in terms of Equation (7), of a positive short-run with a negative long-run derivative), following a depreciation of the real exchange rate, is known in the literature as a J-curve, with the graph of the response over time of the trade balance to a real depreciation resembling a “J” sloping to the right. The first use of the term “J-curve” is normally attributed to Magee (1973), who noticed that, while exchange rates adapt instantly, there is a delay on the side of consumers and producers to adapt to the changes in relative prices; that is, trade balances may not improve in the short run because the exchange rate short-term elasticities are

¹ Notice that the theoretical model given by Equations (1) to (5) is a partial equilibrium one, since the explanatory variables Y , Y^* and q are taken as exogenous. In contrast, in a general equilibrium setting, the three variables would be endogenous, and the Bickerdike–Robinson–Metzler condition would be no longer valid.

smaller than the long-term ones. In other words, recalling that an exchange rate variation has two effects on trade – i.e., price and volume – the price effect would dominate in the short run and the volume effect in the long run, once the amounts of exports and imports begin to respond to the change in exchange rates.

Following the influential paper of Rose and Yellen (1989), a large number of empirical papers have attempted to estimate equations for the trade balance, as in (6), allowing for a dynamic response of the different explanatory variables and, in particular, the real exchange rate, so that a J-curve may be detected. Rose and Yellen’s initial results were not favourable to the existence of a J-curve for the bilateral trade of the US with Japan, Canada, the UK, France, Germany and Italy between 1960 and 1985, and these results were confirmed by Rose (1991) for the total trade of five countries (the UK, Canada, Germany, Japan and the US) between 1974 and 1986. The evidence, however, is inconclusive depending on the countries and periods analyzed; a survey of this empirical literature is provided in Bahmani-Oskooee and Ratha (2004).

In the next section, we will present estimates of dynamic equations based on Equation (6), for the cases of Greece, Italy, Portugal and Spain.

3 Empirical results

In order to show the effects of the different explanatory variables of the trade balance, we will estimate Equation (6) using the method of Dynamic Ordinary Least Squares (DOLS) of Stock and Watson (1993), with the methodology of Shin (1994). The advantage of this method is that provides a robust correction to the possible presence of both endogeneity in the explanatory variables, and serial correlation in the error terms of the OLS estimation. Hence, we will estimate a long-run dynamic equation that includes leads and lags of the (first difference of the) explanatory variables in Equation (6):

$$\begin{aligned} \Delta LTB_t = & \text{constant} + \theta_1 LY_t + \theta_2 LY_t^* + \theta_3 LREER_t \\ & + \sum_{i=-q}^q \beta_i \Delta LY_{t-1-i} + \sum_{i=-q}^q \gamma_i \Delta LY_{t-1-i}^* + \sum_{i=-q}^q \delta_i \Delta LREER_{t-1-i} + v_t \end{aligned} \quad (8)$$

where LTB , LY , LY^* and $LREER$ denote, for each country, the logarithms of the trade balance, domestic real income, foreign real income and the real effective exchange rate, respectively; Δ is the first difference operator; and v_t is an error term. Next, Shin's (1994) test will be performed by computing C_{μ} , an LM statistic from the DOLS residuals that tests for deterministic cointegration (i.e., when no trend is present in the regression).²

Equation (8) has been estimated using quarterly data for Greece, Italy, Portugal and Spain. Regarding definition of the variables, the trade balance is defined as the ratio of real exports to real imports; in addition to total trade (i.e., that made with the whole world), we will also consider the case of trade with the EU (see below). We take as proxies for domestic and foreign real incomes the GDP of each country, and the GDP of either the OECD or the EU, in the equations for total trade or trade with the EU, respectively, all of them in real terms. Finally, three different REERs have been employed in the estimations – computed alternatively using, as deflators, CPIs, export prices or ULCs – and in all cases vis-à-vis 37 industrialized countries and the EU, which will be used in the equations for total trade and trade with the EU, respectively. The data sources are OECD.Stat for total exports and imports and all GDPs, and Eurostat for the REERs and exports and imports to and from the EU. All the data are seasonally adjusted, and the time period is 1994:1–2014:4 for total trade (1995:1–2014:4 in the cases of Greece and Portugal), and, due to problems of data availability, 1999:1–2014:4 for trade with the EU.

We will now test for the order of integration of the variables. To that end, we will first make use of the Phillips–Perron test (Phillips and Perron, 1988), which, under the null hypothesis of a unit root, corrects non-parametrically for the autocorrelation that might appear in the standard Dickey–Fuller test. The results are shown in Table 1, in panels A and B for the variables used in the estimations for total trade and trade with the EU, respectively. According to the test, all variables can be assumed to be integrated of order one, i.e., stationary in first differences. In addition, given the small power of the Phillips–Perron test under certain stochastic properties of the series, we have also performed the KPSS test

² We have also tested for the presence of stochastic cointegration, by including a trend as an additional regressor in the estimated equations, but this trend never proved to be significant (results available from the authors upon request).

Table 1: Phillips–Perron unit root tests

Total trade	A) Total trade											
	Greece			Italy			Portugal			Spain		
	$Z(t_{\hat{\alpha}})$	$Z(t_{\alpha^*})$	$Z(t_{\hat{\alpha}})$	$Z(t_{\hat{\alpha}})$	$Z(t_{\alpha^*})$	$Z(t_{\hat{\alpha}})$	$Z(t_{\hat{\alpha}})$	$Z(t_{\alpha^*})$	$Z(t_{\hat{\alpha}})$	$Z(t_{\hat{\alpha}})$	$Z(t_{\alpha^*})$	$Z(t_{\hat{\alpha}})$
<i>LTB</i>	-1.593	-0.618	-0.506	-0.450	-1.581	-1.497	-1.181	-0.524	-0.763	-0.887	-0.960	-0.739
Δ <i>LTB</i>	-13.40 ^a	-12.17 ^a	-12.15 ^a	-8.641 ^a	-8.422 ^a	-8.475 ^a	-9.419 ^a	-9.197 ^a	-9.184 ^a	-9.035 ^a	-8.990 ^a	-9.012 ^a
<i>LY</i>	-0.121	-1.744	0.659	-0.927	-2.552	1.122	-1.504	-3.592	1.670	0.125	-2.617 ^c	3.176
Δ <i>LY</i>	-7.644 ^a	-6.354 ^a	-6.283 ^a	-5.009 ^a	-4.455 ^a	-4.391 ^a	-6.959 ^a	-5.867 ^a	-5.395 ^a	-3.117 ^c	-2.308	-1.579 ^c
<i>LY*</i>	-1.552	-2.129	5.065	-1.395	-2.085	5.371	-1.552	-2.129	5.065	-1.395	-2.085	5.371
Δ <i>LY*</i>	-3.977 ^b	-3.869 ^a	-2.493 ^b	-4.082 ^a	-3.908 ^a	-2.557 ^b	-3.977 ^b	-3.869 ^a	-2.493 ^b	-4.082 ^a	-3.908 ^a	-2.557 ^b
<i>LREER_CPI</i>	-1.669	-1.543	0.232	-2.869	-2.007	0.493	-1.642	-1.221	0.354	-1.610	-1.131	0.982
Δ <i>LREER_CPI</i>	-6.523 ^a	-6.551 ^a	-6.600 ^a	-7.606 ^a	-7.651 ^a	-7.688 ^a	-7.069 ^a	-7.109 ^a	-7.139 ^a	-6.431 ^a	-6.461 ^a	-6.432 ^a
<i>LREER_EXP</i>	-2.004	-1.457	1.170	-2.389	-1.890	1.057	-2.070	-1.473	0.215	-1.215	-1.966	1.737
Δ <i>LREER_EXP</i>	-7.772 ^a	-7.793 ^a	-7.748 ^a	-8.499 ^a	-8.524 ^a	-8.498 ^a	-6.877 ^a	-6.915 ^a	-6.961 ^a	-7.411 ^a	-7.206 ^a	-7.041 ^a
<i>LREER_ULC</i>	-0.737	-1.670	0.127	-2.551	-1.195	0.631	-0.967	-1.882	0.367	-0.848	-1.293	0.212
Δ <i>LREER_ULC</i>	-5.136 ^a	-4.988 ^a	-5.025 ^a	-6.584 ^a	-6.601 ^a	-6.616 ^a	-5.740 ^a	-5.423 ^a	-5.469 ^a	-5.162 ^a	-5.101 ^a	-5.132 ^a

	B) Trade with the EU											
	Greece			Italy			Portugal			Spain		
	$Z(t_{\hat{\alpha}})$	$Z(t_{\alpha^*})$	$Z(t_{\hat{\alpha}})$	$Z(t_{\hat{\alpha}})$	$Z(t_{\alpha^*})$	$Z(t_{\hat{\alpha}})$	$Z(t_{\hat{\alpha}})$	$Z(t_{\alpha^*})$	$Z(t_{\hat{\alpha}})$	$Z(t_{\hat{\alpha}})$	$Z(t_{\alpha^*})$	$Z(t_{\hat{\alpha}})$
<i>LTB</i>	-2.033	-0.160	-1.157	-2.016	-1.998	-1.632 ^c	-1.633	-0.959	-0.832	-1.752	-0.534	-0.941
Δ <i>LTB</i>	-7.831 ^a	-7.774 ^a	-7.642 ^a	-8.911 ^a	-8.812 ^a	-8.879 ^a	-8.795 ^a	-8.740 ^a	-8.741 ^a	-9.061 ^a	-8.996 ^a	-8.941 ^a
<i>LY</i>	-0.893	-1.304	0.139	-1.916	-2.195	0.359	-1.785	-2.381	0.610	-1.334	-3.475 ^b	2.151
Δ <i>LY</i>	-6.891 ^a	-5.815 ^a	-5.845 ^a	-3.819 ^b	-3.476 ^b	-3.501 ^a	-6.019 ^a	-5.622 ^a	-5.609 ^a	-2.178	-2.161	-2.003 ^b
<i>LY*</i>	-1.882	-2.441	2.563	-1.882	-2.441	2.563	-1.882	-2.441	2.563	-1.882	-2.441	2.563
Δ <i>LY*</i>	-3.264 ^c	-3.105 ^b	-2.821 ^a	-3.264 ^c	-3.105 ^b	-2.821 ^a	-3.264 ^c	-3.105 ^b	-2.821 ^a	-3.264 ^c	-3.105 ^b	-2.821 ^a
<i>LREER_CPI</i>	-1.353	-1.121	0.039	-2.660	-1.241	0.626	-0.797	-1.664	0.807	0.373	-1.632	2.059
Δ <i>LREER_CPI</i>	-4.514 ^a	-4.539 ^a	-4.582 ^a	-6.048 ^a	-6.097 ^a	-6.074 ^a	-7.948 ^a	-7.519 ^a	-7.476 ^a	-6.493 ^a	-6.021 ^a	-5.590 ^a
<i>LREER_EXP</i>	-2.089	-0.846	1.186	-1.845	-1.273	1.494	-1.510	-1.706	0.596	1.134	-2.563	2.440
Δ <i>LREER_EXP</i>	-5.123 ^a	-5.154 ^a	-4.899 ^a	-5.579 ^a	-5.567 ^a	-5.352 ^a	-5.837 ^a	-5.724 ^a	-5.719 ^a	-6.844 ^a	-5.686 ^a	-5.090 ^a
<i>LREER_ULC</i>	-0.720	-1.208	-0.445	-2.098	-0.807	1.297	-1.372	-0.649	-0.420	0.026	-1.161	0.053
Δ <i>LREER_ULC</i>	-2.978	-2.922 ^b	-2.918 ^a	-4.448 ^a	-4.549 ^a	-4.374 ^a	-4.187 ^a	-3.725 ^a	-3.741 ^a	-3.638 ^b	-3.132 ^b	-3.154 ^a

Notes: $Z(t_{\hat{\alpha}})$, $Z(t_{\alpha^*})$ and $Z(t_{\hat{\alpha}})$ are the Phillips-Perron statistics with drift and trend, with drift, and without drift, respectively.

^a, ^b and ^c denote significance at the 1%, 5% and 10% levels, respectively. The critical values are taken from MacKinnon (1996).

(Kwiatkowski et al., 1992) under the null hypothesis of stationarity, contrary to the standard Dickey–Fuller-type tests. The results from this test appear in panels A and B of Table 2 and allow us to reject the null hypothesis of stationarity, with the sole exception of *LTB* for Italy in the case of trade with the EU (even though the statistic η_{μ} is very close to significance at the 10% level: 0.118 versus 0.119).

The results of the econometric estimation of Equation (8) for the total trade of Greece, Italy, Portugal and Spain, in terms of the long-run coefficients and the statistic C_{μ} , appear in Table 3. For each country, three columns appear, according to the measure of the REER employed in the estimation, with CPIs, export prices or ULCs (first, second and third column, respectively) as deflators. The number of leads and lags for the first-differentiated variables in the right-hand side of Equation (8) has been selected as $INT(T^{1/3})$ (with T being the number of observations), as proposed in Stock and Watson (1993).³

As can be seen in the table, the null of deterministic cointegration is not rejected in all cases at the 1% level of significance, suggesting the existence of a long-run relationship between the trade balance and their explanatory variables.⁴ The estimated coefficients on domestic and foreign real incomes always have the expected signs and are clearly significant, except for foreign income in the case of Italy, which is not significant. The results for the REER, however, are not so clear-cut. The coefficients are estimated with the expected sign (i.e., negative) and are significant only for Portugal and Spain (except for the REER deflated with export prices in the case of Spain); even so, the estimated elasticities are clearly lower than those of domestic and foreign income. In turn, the estimated coefficients are never significant for Italy, unlike the case of Greece, where they are always significant but with a sign opposite to that expected (i.e., positive). Finally, Table 4 shows the estimated coefficients on the lags of the REER, which can give us some information about the existence of a J-curve effect; as we can see, some weak evidence of a J-curve appears only for Italy and Portugal, when using the REER deflated with export prices, and with CPIs and export prices, respectively.

³ The full results, including leads and lags of all the explanatory variables, are available from the authors upon request.

⁴ The critical values for the C_{μ} statistic are 0.271 and 0.159, at the 1% and 5% significance levels, respectively; and are taken from Shin (1994), Table 1, for $m = 3$.

Table 2: KPSS stationarity tests

A) Total trade								
	Greece		Italy		Portugal		Spain	
	η_{μ}	η_{τ}	η_{μ}	η_{τ}	η_{μ}	η_{τ}	η_{μ}	η_{τ}
<i>LTB</i>	0.295 ^a	0.458 ^c	0.241 ^a	0.657 ^b	0.229 ^a	0.410 ^c	0.240 ^a	0.239
<i>LY</i>	0.286 ^a	0.574 ^b	0.311 ^a	0.706 ^b	0.288 ^a	0.799 ^a	0.271 ^a	1.023 ^a
<i>LY*</i>	0.271 ^a	1.205 ^a	0.287 ^a	1.121 ^a	0.271 ^a	1.205 ^a	0.287 ^a	1.121 ^a
<i>LREER_CPI</i>	0.138 ^c	0.745 ^a	0.079	0.916 ^a	0.157 ^b	0.827 ^a	0.148 ^b	0.920 ^a
<i>LREER_EXP</i>	0.096	1.106 ^a	0.174 ^b	1.064 ^a	0.160 ^b	0.919 ^a	0.162 ^b	1.060 ^a
<i>LREER_ULC</i>	0.168 ^b	0.709 ^b	0.100	0.927 ^a	0.279 ^a	0.483 ^b	0.159 ^b	0.666 ^b

B) Trade with the EU								
	Greece		Italy		Portugal		Spain	
	η_{μ}	η_{τ}	η_{μ}	η_{τ}	η_{μ}	η_{τ}	η_{μ}	η_{τ}
<i>LTB</i>	0.215 ^b	0.770 ^a	0.118	0.136	0.176 ^b	0.380 ^c	0.226 ^a	0.631 ^b
<i>LY</i>	0.252 ^a	0.254	0.236 ^a	0.242	0.229 ^a	0.319	0.254 ^a	0.770 ^a
<i>LY*</i>	0.221 ^a	0.914 ^a	0.221 ^a	0.914 ^a	0.221 ^a	0.914 ^a	0.221 ^a	0.914 ^a
<i>LREER_CPI</i>	0.129 ^c	0.759 ^a	0.075	0.832 ^a	0.233 ^a	0.622 ^b	0.226 ^a	0.920 ^a
<i>LREER_EXP</i>	0.090	0.917 ^a	0.156 ^b	0.857 ^a	0.213 ^b	0.824 ^a	0.250 ^a	0.926 ^a
<i>LREER_ULC</i>	0.185 ^b	0.294	0.147 ^b	0.901 ^a	0.244 ^a	0.381 ^c	0.222 ^a	0.435 ^c

Notes: η_{μ} and η_{τ} are the KPSS statistics with trend, and without trend, respectively.

^a, ^b and ^c denote significance at the 1%, 5% and 10% levels, respectively. The critical values are taken from Kwiatkowski et al. (1992, Table 1).

Table 3: Estimation of long-run relationships for total trade. Stock-Watson-Shin cointegration tests

	Greece			Italy			Portugal			Spain		
constant	-4.044 ^a (-4.465)	-1.088 (-1.008)	-1.330 (-1.149)	8.430 ^a (13.47)	7.674 ^a (11.11)	8.548 ^a (12.28)	3.115 ^a (5.439)	4.398 ^a (9.382)	1.796 ^b (2.242)	-1.044 (-1.057)	-1.996 ^b (-2.074)	-0.126 (-0.122)
LY_t	-1.725 ^a (-12.87)	-1.661 ^a (-12.27)	-2.109 ^a (-10.31)	-1.255 ^a (-6.048)	-1.242 ^a (-6.296)	-1.259 ^a (-5.789)	-2.538 ^a (-14.80)	-2.173 ^a (-18.73)	-1.878 ^a (-8.326)	-1.838 ^a (-7.476)	-1.969 ^a (-6.817)	-1.551 ^a (-5.696)
LY_t^*	1.361 ^a (7.272)	0.872 ^a (3.187)	1.230 ^a (4.622)	-0.054 (-0.483)	0.057 (0.572)	-0.073 (-0.585)	1.513 ^a (9.611)	1.380 ^a (10.42)	1.207 ^a (7.274)	1.709 ^a (5.655)	1.935 ^a (6.208)	1.351 ^a (4.087)
$LREER_t$	1.334 ^a (4.496)	1.531 ^a (3.398)	1.534 ^a (4.311)	-0.105 (-0.556)	-0.189 (-0.906)	-0.084 (-0.724)	-0.570 ^b (-2.640)	-1.682 ^a (-4.600)	-0.528 ^b (-2.407)	-0.405 ^b (-2.233)	-0.388 (-1.288)	-0.384 ^a (-3.062)
R^2	0.998	0.998	0.998	0.924	0.949	0.916	0.997	0.998	0.996	0.993	0.993	0.994
C_{μ}	0.058	0.086	0.070	0.066	0.061	0.060	0.064	0.079	0.044	0.043	0.048	0.044

Notes: t-statistics in parentheses. ^a, ^b and ^c denote significance at the 1%, 5% and 10% levels, respectively.

Table 4: Short-run coefficients on the real exchange rate. Total trade

	Greece			Italy			Portugal			Spain		
$\Delta LREER_t$	-0.654 (-0.945)	-0.970 (-1.288)	-0.793 (-1.032)	0.347 (0.938)	0.804 ^b (2.529)	0.158 (0.522)	1.465 ^b (2.134)	2.045 ^a (3.215)	0.980 (1.241)	0.867 (1.561)	0.707 (1.201)	0.654 (1.306)
$\Delta LREER_{t-1}$	0.126 (0.186)	-0.240 (-0.292)	0.312 (0.423)	0.200 (0.542)	0.620 ^c (1.962)	0.046 (0.138)	1.178 ^c (1.893)	0.801 (1.362)	0.415 (0.567)	0.066 (0.119)	0.161 (0.255)	0.066 (0.126)
$\Delta LREER_{t-2}$	0.257 (0.388)	0.763 (0.968)	0.520 (0.734)	0.386 (0.804)	0.355 (0.895)	0.195 (0.490)	0.797 (1.076)	0.065 (0.100)	0.462 (0.557)	-0.047 (-0.077)	0.371 (0.543)	0.216 (0.388)
$\Delta LREER_{t-3}$	0.581 (0.849)	1.013 (1.280)	0.667 (0.930)	0.595 (1.239)	0.293 (0.708)	0.498 (1.251)	0.209 (0.297)	-0.417 (-0.668)	0.626 (0.780)	0.164 (0.287)	0.470 (0.695)	0.155 (0.295)
$\Delta LREER_{t-4}$	0.687 (1.053)	1.257 (1.618)	0.798 (1.167)	0.038 (0.083)	-0.075 (-0.191)	0.055 (0.148)	0.535 (0.778)	-0.126 (-0.206)	0.829 (1.128)	0.006 (0.010)	0.190 (0.285)	-0.343 (-0.677)

Notes: See Table 3.

When Equation (8) is estimated taking as dependent variable the trade balance with the EU, the results are shown in Table 5. Recall that in this case the sample period is notably shorter than for total trade (i.e., 16 years instead of 20–21), so the results should be taken with care. Again, the null of deterministic cointegration is not rejected at the 1% level of significance (5% for Greece and Portugal with the REER deflated with export prices), suggesting the existence of a long-run relationship between the trade balance and their explanatory variables. The long-run effect of both domestic and foreign real income is always significant, with the expected signs. In turn, the coefficient on the REER is negative and significant for Italy in all cases, for Portugal with the REER based on CPIs and ULCs, and for Spain with the REER based on ULCs; and non-significant in the rest of the cases. Lastly, from Table 6 we can see that evidence of a J-curve only appears for Portugal in the equations using the REER deflated with export prices and ULCs.

As already mentioned in the introduction, due to the expansion of global value chains, a debate has arisen regarding the disconnection between exchange rates and trade that might have emerged in recent years. In the remainder of this section, we will address this issue by analyzing the possible existence of structural changes in the above estimated equations. Specifically, we will make use of the approach of Kejriwal and Perron (2008, 2010) that tests for multiple structural changes in cointegrated regression models. These authors develop a sequential procedure that makes it possible to identify the presence of parameter instability in cointegration regression models, in addition to allowing for consistency in the number of breaks detected. Three types of test statistics are presented: (i) a sup-Wald test of the null hypothesis of no structural break versus the alternative of a fixed (arbitrary) number of breaks k ; (ii) a test of the null hypothesis of no structural break versus the alternative of an unknown number of breaks, given some upper bound; and (iii) a sequential test of the null hypothesis of k breaks versus the alternative of $k + 1$ breaks.

In Table 7, we report the number of breaks selected according to the Kejriwal–Perron tests for multiple structural changes in cointegrated regression models. To save space, we only show the results for the cases in which at least one structural change was detected; due to the small length of our sample period, we have allowed up to two breaks under the alternative hypothesis. The number of breaks has been selected using three alternative procedures – a sequential procedure, the

Table 5: Estimation of long-run relationships for trade with the EU. Stock-Watson-Shin cointegration tests

	Greece			Italy			Portugal			Spain		
constant	-16.01 ^a (-11.60)	-16.23 ^a (-3.936)	-14.58 ^a (-6.505)	4.507 ^a (6.067)	1.014 (1.308)	-0.858 (-0.973)	2.271 ^a (4.510)	2.978 ^a (3.845)	-1.122 ^c (-2.027)	-7.785 ^a (-4.186)	-6.972 (-1.414)	-4.941 ^b (-2.253)
LY_t	-1.759 ^a (-9.529)	-1.586 ^a (-8.933)	-1.101 ^a (-2.840)	-0.778 ^a (-3.477)	-0.754 ^a (-3.760)	-0.723 ^a (-3.000)	-0.711 ^a (-3.906)	-0.931 ^a (-3.815)	-0.608 ^a (-4.077)	-3.217 ^a (-4.427)	-2.761 ^a (-3.846)	-2.211 ^b (-2.575)
LY_t^*	3.222 ^a (7.951)	3.280 ^b (2.483)	3.045 ^a (5.937)	0.721 ^a (3.191)	0.978 ^a (4.263)	1.107 ^a (3.463)	0.641 ^a (4.288)	0.704 ^c (1.707)	0.815 ^a (4.924)	3.708 ^a (6.490)	3.254 ^b (2.638)	2.828 ^a (3.101)
$LREER_t$	1.050 (0.922)	0.481 (0.207)	-0.818 (-0.987)	-2.407 ^a (-5.697)	-1.649 ^a (-4.576)	-1.271 ^a (-4.376)	-1.511 ^a (-4.206)	-1.497 (-0.714)	-0.716 ^b (-3.660)	0.401 (0.267)	0.236 (0.089)	-0.929 ^b (-2.101)
R^2	0.999	0.999	0.999	0.890	0.873	0.875	0.940	0.887	0.951	0.994	0.992	0.996
C_μ	0.088	0.139	0.064	0.085	0.084	0.083	0.113	0.127	0.104	0.070	0.074	0.062

Notes: See Table 3.

Table 6: Short-run coefficients on the real exchange rate. Trade with the EU

	Greece			Italy			Portugal			Spain		
$\Delta LREER_t$	1.364 (0.833)	2.925 (0.994)	-0.014 (-0.007)	0.676 (0.928)	0.977 (1.147)	0.407 (0.590)	0.599 (1.040)	2.637 ^b (2.101)	0.222 (0.472)	1.567 (1.099)	0.540 (0.234)	2.156 ^c (1.934)
$\Delta LREER_{t-1}$	-0.036 (-0.023)	-4.376 (-1.666)	0.453 (0.239)	-1.748 ^a (-2.796)	-1.061 (-1.395)	-1.034 (-1.690)	-1.279 ^b (-2.073)	0.575 (0.399)	-0.816 (-1.645)	1.097 (0.695)	-0.396 (-0.194)	0.301 (0.315)
$\Delta LREER_{t-2}$	1.826 (1.009)	-1.493 (-0.417)	0.312 (0.180)	-2.156 ^a (-3.224)	-1.351 (-1.628)	-1.103 (-1.638)	-1.238 ^c (-1.940)	0.097 (0.075)	-0.677 (-1.330)	1.453 (1.158)	-0.702 (-0.371)	0.298 (0.347)
$\Delta LREER_{t-3}$	1.702 (0.876)	1.791 (0.495)	2.019 (1.104)	-0.267 (-0.370)	0.604 (0.662)	0.349 (0.478)	1.475 ^c (2.020)	2.682 ^b (2.238)	1.476 ^b (2.574)	0.453 (0.351)	-0.369 (-0.223)	0.266 (0.289)
$\Delta LREER_{t-4}$	1.264 (0.822)	4.101 (1.378)	-2.013 (-1.153)	-0.455 (-0.652)	0.266 (0.310)	-0.368 (-0.511)	1.014 (1.406)	1.305 (1.084)	0.900 (1.561)	-0.314 (-0.197)	0.122 (0.068)	-0.248 (-0.238)

Notes: See Table 3.

Table 7: Number of breaks selected. Kejriwal-Perron tests for structural change

Total trade		Trade with the EU
Greece	Spain	Spain
1	1	1
(2004:4)	(2008:1)	(2009:1)

Note: Date of the estimated structural change in parentheses.

Bayesian information criterion, and the modified Schwarz criterion proposed by Liu et al. (1997)⁵ – all of which led to the same result. As shown in the table, we have detected one structural change in just three cases (in all three cases, independently of the measure of the REER used in the estimations):

- For total trade, in the cases of Greece and Spain, dated at 2004:4 and 2008:1, respectively.
- For trade with the EU, in the case of Spain, dated at 2009:1.

Finally, our trade balance equations for these three cases have been re-estimated by subperiods, i.e., before and after the break dates detected by the Kejriwal–Perron tests. The results of the estimations are shown, respectively, in Table 8 for the total trade of Greece and Spain, and in Table 9 for the EU trade of Spain. In all cases, the null hypothesis of deterministic cointegration is not rejected at the 1% level of significance. It should be noticed, however, that the results in these tables must be taken more carefully than usual given the very small number of observations available for the estimations. This problem is particularly serious in the two second-period estimations for Spain, which used just 28 and 24 observations (i.e., 7 and 6 years), respectively. Regarding the estimation results for total trade, in the case of Greece only the REER computed using ULCs shows statistically significant coefficients, but with a (correct) negative sign in the first subperiod and a positive sign in the second one. In turn, in the case of Spain, the three REERs appear with a negative and significant coefficient in the first subperiod; unlike the second subperiod, where this result only keeps for the REER computed using export prices. A similar conclusion shows for the Spanish trade with the EU, even though the coefficient on the REER computed using ULCs is now negative and significant for the second subperiod too.

⁵ The complete results of the tests are available from the authors upon request.

Table 8: Estimation of long-run relationships for total trade, by subperiods. Stock-Watson-Shin cointegration tests

	Greece						Spain					
	1995:1-2004:3		2004:4-2014:4		1994:1-2007:4		2008:1-2014:4					
constant	18.94 ^a (5.133)	6.820 (0.620)	20.51 ^a (4.242)	-15.50 ^a (-7.982)	-13.34 ^a (-4.652)	-11.05 ^a (-11.46)	-17.04 ^b (-2.126)	-26.93 ^b (-2.248)	-25.96 ^a (-5.120)	29.24 ^a (3.128)	53.81 ^a (9.743)	17.98 ^c (1.919)
LY_t	1.026 ^c (1.804)	-1.192 (-0.573)	1.324 (1.454)	-1.181 ^a (-9.147)	-1.272 ^a (-6.289)	-1.513 ^a (-10.98)	-5.421 ^a (-2.909)	-7.308 ^a (-2.796)	-7.501 ^a (-6.411)	-4.760 ^a (-5.886)	-4.883 ^a (-10.61)	-5.833 ^a (-7.177)
LY_t^*	-3.256 ^a (-3.882)	-0.512 (-0.192)	-3.305 ^b (-2.650)	-2.659 ^a (-6.809)	2.360 ^a (3.729)	2.244 ^a (18.39)	6.827 ^b (2.636)	9.784 ^b (2.579)	9.678 ^a (5.921)	-0.063 (-0.126)	-1.540 ^a (-5.286)	1.927 ^b (2.210)
$LREER_t$	-0.195 (-0.832)	1.603 (1.443)	-1.695 ^a (-6.469)	0.661 (0.556)	0.963 (0.499)	0.925 ^a (3.264)	-0.816 ^a (-3.186)	-1.299 ^b (-2.352)	-0.808 ^a (-5.724)	0.102 (0.132)	-6.126 ^a (-10.70)	1.439 ^b (2.688)
R^2	0.999	0.999	0.999	0.999	0.999	0.999	0.999	0.995	0.997	0.997	0.999	0.998
C_μ	0.099	0.100	0.097	0.044	0.046	0.047	0.038	0.039	0.042	0.036	0.036	0.036

Notes: See Table 3.

Table 9: Estimation of long-run relationships for trade with the EU, by subperiods. Stock-Watson-Shin cointegration tests

	Spain					
	1999:1–2008:4			2009:1–2014:4		
	constant	6.696 ^b (2.707)	1.589 (0.267)	3.595 (1.140)	-24.93 ^a (-11.29)	28.90 ^a (3.318)
LY_t	1.514 ^b (2.744)	2.677 (1.062)	1.170 (1.571)	-5.539 ^a (-37.79)	-4.264 ^a (-17.01)	9.201 ^a (9.276)
LY_t^*	-1.359 ^c (-1.789)	-0.543 (-0.249)	-0.804 (-0.788)	4.136 ^a (23.91)	2.777 ^a (6.390)	-2.981 ^a (-4.214)
$LREER_t$	-3.146 ^a (-11.52)	-7.036 ^b (-2.312)	-2.522 ^a (-9.764)	14.37 ^a (17.19)	-11.30 ^a (-5.219)	-6.455 ^a (-11.99)
R^2	0.999	0.998	0.999	0.994	0.994	0.995
C_μ	0.052	0.045	0.057	0.036	0.037	0.038

Notes: See Table 3.

4 Conclusions

In this paper, we have estimated trade balance equations for the Southern European members of the euro area (Greece, Italy, Portugal and Spain). The equations have been estimated both for total trade and for the trade performed with the EU, and taking three alternative measures of the REER – i.e., computed using, as deflators, CPIs, export prices and ULCs, respectively. The estimation period was 1994:1–2014:4 for total trade (1995:1–2014:4 in the cases of Greece and Portugal) and 1999:1–2014:4 for the trade with the EU; notice that in the latter case the results should be taken with care, since the sample period is notably shorter. Our ultimate aim was to assess whether changes in the real exchange rates, as a proxy of external competitiveness, affect the trade balance, thus validating the conventional policy recommendation for these countries.

The trade balance equations have been estimated using the DOLS methodology. The null hypothesis of deterministic cointegration was not rejected in all cases, suggesting the existence of a long-run relationship between the trade balance and their explanatory variables. While the estimated coefficients on domestic and foreign real incomes always had the expected signs and were clearly significant (with the sole exception of foreign income in the case of Italy), this was

not generally true for the REER. Specifically, regarding the long-run effect of the REER:

- In the case of Greece, it was always significant for total trade, but with the sign of the estimated coefficient opposed to that expected; and it was non-significant for trade with the EU
- In the case of Italy, it was always significant for trade with the EU, and non-significant for total trade
- In the case of Portugal, it was always significant, except when using the REER based on export prices in the equations for trade with the EU; and the estimated coefficients were lower than those on domestic and foreign incomes in the equations for total trade
- In the case of Spain, it was significant when using the REER based on CPIs and ULCs in the equations for total trade, and with the REER based on ULCs in the equations for trade with the EU; and the estimated coefficients were clearly lower than those on domestic and foreign incomes

In addition, some weak evidence of a J-curve appeared for total trade only for Italy and Portugal, when using the REER deflated with export prices, and with CPIs and export prices, respectively; and, in the case of trade with the EU, for Portugal in the equations using the REER deflated with export prices and ULCs.

Next, we tested for the existence of structural change in the estimated equations by means of the Kejriwal–Perron test and found one structural change in just three cases, independently of the measure of the REER: for total trade in the cases of Greece and Spain (dated at 2004:4 and 2008:1, respectively), and for trade with the EU in the case of Spain (dated at 2009:1). Finally, when re-estimating, by subperiods (i.e., before and after the break dates detected), our trade balance equations for these three cases, and despite the presence of cointegration, the results proved to be rather unclear and difficult to interpret, which might be related to the small number of observations available.

Several competing explanations have been recently suggested to account for European current account imbalances; see, e.g., Constantine et al. (2016). However, our main result (i.e., that the trade balance is not clearly related to changes in the REER) should not come as a surprise. Almost 40 years ago, Kaldor (1978) showed that those countries with the greatest increase in their market shares

during the post-war period were also the ones that experienced the highest increases in ULCs, and hence the highest decline in their price competitiveness: this result is known as Kaldor's paradox. In subsequent years, the so-called "new trade theories" were extensively developed. Based on imperfect competition, where specialization is allowed to be driven by increasing returns, and firms differentiate their products (in terms of design, quality, and the like), these models predict a lower influence of prices on the trade balance. In such a context, Krugman (1989) claims that those countries experiencing higher growth rates "expand their share of world markets, not by reducing the relative prices of their goods, but by expanding the range of goods that they produce as their economies grow" (Krugman, 1989, p. 1039).

Hence, following this line of reasoning, in order to improve a country's competitiveness, attention should be addressed to the four critical channels highlighted by Ottaviano et al. (2007): market accessibility, market size, Ricardian technological advantage, and the institutional and political framework; see also Di Mauro and Forster (2008).

On the other hand, especially in Spain and Italy, the bulk of exports are made by large firms with high levels of productivity. For instance, Spanish exports have shown rather good performance in recent years, despite a large fall in price competitiveness, which might be explained either by the singularity and quality of Spanish products or the concentration of exports in large firms enjoying productivity levels above the average of the economy (Myro, 2013). In particular, there are some other factors that can offset a worsening in price competitiveness, e.g., supplying products tailored to world demand and oriented towards expanding markets (Myro, 2015). Finally, these conclusions would reinforce the idea that peripheral economies of the euro area are described by a wage-led regime, in the sense of Bhaduri and Marglin (1990) – i.e., when an increase in the wage share of total income has a positive impact on aggregate demand. Accordingly, our results are in line with those of a recent empirical contribution by Onaran and Obst (2016), which shows that most European economies, and in particular Spain, Italy, Portugal and Greece, are wage-led.

To conclude, it seems that demand appears to be more relevant than relative prices when explaining the evolution of trade flows. Accordingly, in order to improve the trade balance, one should have greater trust in an increase in external demand rather than lowering relative prices; at the same time, an improvement of

domestic demand might jeopardise any favourable developments of the trade balance.

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