Econometric analysis of the exchange rate pass through into import prices for the Euro Area. Some new results

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Abstract

Exchange rate pass through (ERPT) has been extensively explored in the empirical literature. In this paper, we focus on the specific case of the Euro Area to analyze pass-through into import prices from outside the Eurozone. More sophisticated econometric techniques are applied to revised previous results. In particular, by making use of a wide variety of testing procedures for two different, but alternative and competing, models incorporating a nonlinear and asymmetric effect of the nominal exchange rate on the magnitude and statistical significance of the ERPT, we seek for the existence of empirical evidence supporting such an effect for the considered series. However, globally considered, the empirical analysis provides strongly support for the no existence of such a significant effect for the series and time period used as well as evidence for a complete and time-invariant ERPT.

Keywords: exchange rates, import prices, pass-through, euro area, nonlinear cointegration, threshold effects

JEL Classification: *F31, F41, C22, C12*

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1. Introduction

Exchange rate pass-through (ERPT) can be defined as the rate at which exchange rate changes are transmitted into import prices expressed in the currency of the importing country. This analysis has attracted a great deal of attention in the empirical literature where the debate focuses on the prevalence of the producer-currency pricing (PCP) versus local-currency pricing (LCP). When the PCP holds, the ERPT is complete since the producers (exporters) absorb the exchange rate variation, and consequently the mark-up does not change. On the other hand, if the LCP holds, producers may decide not to vary prices in the importing country's currency and assume exchange rate fluctuations within the mark-up. Many papers have analyses ERPT by considering different countries, sample period and methodology (See for instance; Goldberg and Knetter, 1997; Gosh and Rajan, 2007 or Aaron et al, 2014)

With the nominal exchange rate being eternally fixed at one rate and a single authority being in charge of the monetary policy for the entire area; certain mechanisms and instruments are not available to the euro area members to deal with an exchange rate shock to facilitate individual preferences (Langwasser, 2009). Thus, the analysis of the ERPT is particularly relevant for the case of the Eurozone since it is a key factor for external adjustment. For instance, incomplete pass-through could delay or diminish the response of the trade balance to shocks and produce a certain degree of exchange rate disconnect (Devereux and Engel, 2002). Moreover, if differences in the ERPT between the Eurozone countries exist, a common external shock may have different impact on the euro-area economies.

Faruqee (2004) analyses the pattern of ERPT in the euro area along the pricing chain by using a VAR approach to identify the effects of an exogenous exchange rate shock. Results of the analysis suggest incomplete pass-through. Campa and Gonzalez-Mínguez (2006) explore differences on short and long-run ERPT into the prices of extra-Euro area imports. Again, the authors find evidence of incomplete pass-through as well as differences in the ERPT across the euro area countries. More recently, María-Dolores (2010) studies the ERPT in the new member states and candidate countries of the European Union obtaining that neither the hypothesis of LCP nor the hypothesis of PCP can be rejected. Furthermore, differences in the ERPT across countries are found. Saiki (2011) examines how monetary integration affects the exchange rate pass through, by testing whether monetary policy convergence in the euro area led to a convergence in terms of exchange rate pass-through.

As an alternative to the more traditional modeling based on a linear, and possibly stable, long-run relationship among the system variables, there could be some theoretical arguments indicating the possible existence of a nonlinear adjustment mechanism determining more accurate measures of the ERPT. Thus, the main objective of the paper is to empirically evaluate this evidence on the basis of two different, but alternative, specifications incorporating such nonlinear behavior, namely: (i) the so-called threshold cointegration model by Balke and Fomby (1997), where this effect is incorporate through the disequilibrium error term following a self-exiting threshold autoregression (SETAR) process, (ii) and a functional-coefficient cointegrating regression model with a time-varying cointegration vector with values changing from one regime to another according to the magnitude of a predetermined threshold variable proposed by Gonzalo and Pitarakis (2006a). With this, the paper is organized as follows. Section 2 presents a brief revision of the theoretical basis for the determination of the ERPT, while data description is presented in Section 3. The main body of the paper is section 4 devoted to the econometric analysis of the series involved and the statistical significance of two different but competing representations of the nonlinear and asymmetric adjustment process in a single-equation cointegrating regression model. Finally, some conclusions are drawn in section 5.

2. Theoretical framework

The ERPT can be defined as the transmission of exchange rate changes into import prices in the destination market currency price of goods. Therefore, it can be presented as follows:

$$MP_t = ER_t \cdot XP_t$$

Taking logs we obtain:

$$mp_t = er_t + xp_t \qquad [1]$$

where *MP* is the import price denominated in local currency, *ER* is the bilateral exchange rate expressed in units of local currency per unit of foreign currency, while *XP* is the price of exports, denominated in foreign currency, which consists of the exporters marginal cost (*FMC*) and a mark-up (*FMKUP*), that is

$$XP_t = FMC_t \cdot FMKUP_t .$$

Being in logs

$$xp_t = fmc_t + fmkup_t \qquad [2]$$

Consequently, considering [1] and [2], ERPT can be defined as:

$$ERPT_t = \partial mp_t / \partial er_t = 1 + \partial fmc_t / \partial er_t + \partial fmkup_t / \partial er_t , \qquad [3]$$

so that, the three main determinants of ERPT into import prices are: (i) changes in the nominal exchange rate, (ii) changes in the producers' marginal costs and (iii) the mark-up response to exchange rate movements. In particular, producers may decide to absorb part of the exchange rate fluctuation within the mark-up instead of complete passing them through the price expressed in the importing country local currency. In that case, there is partial or incomplete ERPT. Moreover, the two following extreme situations can be found. Firstly, exporters will opt to adjust the mark-up absorbing all the fluctuation of the nominal exchange rate within it. In that case, there is zero ERPT and the LCP held. Secondly, ERPT is complete when producers opt not to adjust the mark-up when exchange rate fluctuates, passing all the variation into prices of imports expressed in local currency. In that case, the PCP held[1].

Also, taking into account that the mark-up has two components, a specific constant component and a reaction to exchange rate movements, and that the marginal cost, which is a function of demand conditioning the imports, also depends on a set of components determining the world price of the product, *fp*, then [1] can be rewritten as[2]:

$$mp_t = \alpha + \beta_1 er_t + \beta_2 fp_t + u_t \qquad [4]$$

where β_1 captures the pass-through elasticity and $\beta_2 fp_t$ is considered as independent of the exchange rate. In practice, a simple framework to estimate the ERPT is to estimate the following equation, where all the variables are expressed in first differences

[1] With our definition of nominal exchange rates (in terms of foreign currency per unit of domestic currency), PCP corresponds in our paper to an ERPT of −1.

[2] Particularly, we have that $\beta_2 fp_t = \boldsymbol{\delta}'\mathbf{w}_t$, $\boldsymbol{\delta} = (\delta_0, \delta_1, \delta_2)'$, $\mathbf{w}_t = (y_t, fw_t, fcp_t)'$, where y_t is the income in the importing country, fw_t the wage and fcp_t the commodity price index in foreign currency.

$$\Delta mp_t = \Delta d_t + \sum_{k=0}^{m} \beta_k \Delta er_{t-k} + \sum_{k=0}^{m} \gamma_k \Delta fp_{t-k} + v_t \qquad [5]$$

where we can use a quite general representation for the deterministic component in [4] as a linear trend component, that is, $d_t = \alpha_0 + \alpha_1 t$, so that $\Delta d_t = 0$ when $\alpha_1 = 0$, and $\Delta d_t = \alpha_1$ when the full specification in levels is considered. This distributed lag regression model, first proposed by Campa and Gonzalez-Mínguez (2006) and also considered by María-Dolores (2010), allows to measure the short-run ERPT through the coefficient β_0, while $\sum_{k=0}^{m} \beta_k$ is interpreted as the long-run ERPT. In our empirical analysis of section 4, this model will serve us as a benchmark to compare the results obtained by other more sophisticate techniques.

3. Data description and some basic evidence

In this paper, we are considering ERPT into import prices from outside the Eurozone for the initial eleven countries that adopt the euro in 1999 (Austria, Belgium, Finland, France, Germany, Ireland, Italy, Luxembourg, Netherlands, Spain and Portugal) and Greece who adopts the euro two years later. Extra-Euro Area trade is considered since is the one still exposed to exchange rate fluctuations after the euro was created. We use time series data on import unit values for total imports. In particular, monthly unit value index of imports from Extra-Eurozone (17) expressed in euro is used as a proxy for import prices.

Regarding marginal costs, or foreign price index, there are two possibilities. On the one hand, there is possible integration in the world market, and hence there exists only a single international market for the product, regardless of product origin, destination market or currency denomination. In this case measuring the world price should be the same when expressed in a common currency. In particular, the world price will be expressed in US dollars, although any other currency could be used as *numeraire.* On the other hand, there is possible highly segmented market with a significant degree of price discrimination for the product depending on the countries of origin and destination of imports. Under this view of the world, an appropriate measure for the bilateral exchange rate and the foreign price index should be made contingent, for a given destination country, on the countries in which these imports originate. Campa and Gonzalez-Mínguez (2006) analyze which market structure hold for the euro area trade originated from outside the Eurozone and obtain that the best option is the integrated market structure. Therefore, monthly series of unit value indices of imports expressed in US$ is used as foreign price index.

Finally, monthly nominal exchange rate (period average) defined as units of foreign currency ($) per unit of domestic currency (€) is needed. Data on import prices expressed in both domestic and foreign currency were extracted from Eurostat-Comext dataset from January 2000 to December 2011 for the twelve countries that initially joined to the Eurozone. Data on nominal exchange rate were obtained from the International Financial Statistics database by the International Monetary Fund for the same sample period.

Different degree of ERPT among the Eurozone countries can be cause by distinct degrees of openness to Extra-Eurozone imports as well as by heterogeneity in the structure of imports. Figure 1 presents the share of each country total imports as a percentage of its GDP and imports from outside the Euro Area (17) as a percentage of total imports. Important divergences in the openness rates as well as share of imports from Extra-Euro Area on total imports are observed. On the one hand, Belgium, Luxembourg and Netherlands present the highest share of imports on GDP, being nearly

90% in the case of Luxembourg. On the other hand, Spain, France, Italy and Greece present a percentage below 30%. Regarding the share of imports from outside de Eurozone over the total imports, Ireland, Finland and Netherland present the highest percentage, above 60%, while Luxembourg, Austria a Portugal present the lowest shares.

[Figure1, here]

The composition of imports from outside the Euro Area by type of product is presented in Figure 2. Again, this distribution varies across the Eurozone countries. In general, machinery and transport equipment, mineral fuels, and manufactured goods are the categories that present a larger share on the composition of total Extra-Euro Area imports, while chemical products, crude materials, and food and live animals present a the lowest. However, the shares of the different categories in non-euro area are largely heterogeneous between the Eurozone countries. Therefore, divergences in both openness rates and import composition across countries may explain the existence of different degree of ERPT in the Euro Area[3].

[Figure 2, here]

4. Econometric analysis

The starting point of the subsequent analysis is based on the general representation of the relation between the logs of the import series and the nominal exchange rate, given in [4], as a standard linear single-equation regression model of the form

$$y_t = \alpha_0 + \mathbf{x}_t'\boldsymbol{\beta} + u_t \quad t = 1,\ldots,n \qquad [6]$$

where $\mathbf{x}_t$ the $k\times1$ vector of regressors, with k = 2 in our case, that is $\mathbf{x}_t = (er_t, fp_t)'$. Given the usual concern about the nonstationarity of the model variables, this equation is interpreted as a cointegrating regression model, although there could be some reasons questioning the linearity of the relation. As it is well known in the econometric literature on time series analysis, the pioneering work of Granger (1981) and Engle and Granger (1987) states that a linear cointegrating relationship implies a continuous and symmetric adjustment to the long run equilibrium. However, there could be some situations where a non-linear adjustment mechanism could be more appropriate (see for instance the recent papers by Camacho, 2005; Gonzalo and Pitarakis, 2006a, b; Li and Lee, 2010; and Hansen, 2011; for a review of the relevant references on this issue, theoretical justifications, and some different proposals).

Basically, the idea behind these modifications of a standard cointegrated system consists on generalizing the generating mechanism of the equilibrium error sequence $u_t = y_t - \mathbf{x}_t'\boldsymbol{\beta} = \boldsymbol{\kappa}_1'\mathbf{z}_t$, with $\mathbf{z}_t = (y_t, \mathbf{x}_t')'$ and $\boldsymbol{\kappa}_1 = (1, -\boldsymbol{\beta}')'$ being the $(k+1)\times1$ cointegrating vector when u_t is stationary, to allow for a particular form of a nonlinear adjustment processes by usually specifying a random coefficient-type, or time-varying, autoregressive process, such as $u_t = \alpha_t u_{t-1} + \upsilon_t$, with α_t governing the stochastic properties of u_t, generally in terms of an usually observable transition variable.

In some cases, there could be some theoretical arguments supporting this type of

[3] According to Campa and González-Mínguez (2006), differences in the degree of ERPT among Euro Area countries are consequence of the different degrees of openness of each country to the non-eurozone rather than the heterogeneity in the structure of imports.

behaviour or, at least, some guidance on how to obtain such specification. Thus, following Al-Abri and Goodwin (2009), given the assumption that small changes in the nominal exchange rate will not imply any significant adjustment in the mark-up due to menu costs, but large ones will force to some changes in the producer's mark-up, the ERPT can be written, when $\partial xmc_t / \partial er_t = 0$, as $ERPT_t = 1 + \varphi_t (1 - I_{t-d}(\gamma))$ with $\varphi_t = \partial xmkup_t / \partial er_t$, and $I_{t-d}(\gamma)$ being an indicator function defined as $I_{t-d}(\gamma) = I(| q_{t-d} | \leq \gamma)$, with $1 - I_{t-d}(\gamma) = J_{t-d}(\gamma) = I(| q_{t-d} | > \gamma)$. In this formulation, q_{t-d} is known as the threshold, or transition, variable with respect to the threshold parameter γ, with $d \geq 0$ the time-delay parameter reflecting the possibility that economic agents or controllers may react to deviations from the equilibrium with a certain lag. It is particularly convenient that the threshold parameter γ take values on a closed and bounded subset of the sample space of the variable q_{t-d}. In this framework, an obvious candidate for the transition variable is the first difference of the nominal exchange rate, that is, $q_{t-d} = \Delta er_{t-d}$, with $d = 0$, as a measure of the changes in the nominal exchange rates, but that can be generalized to any other factors that introduce adjustment costs to changes in the import prices. Hence, with the aim to incorporate such potential source of nonlinear relationship between import prices and nominal exchange rates, these same authors choose to follow the threshold cointegration methodology first developed by Balke and Fomby (1997), as an alternative to specifying a nonlinear version of the cointegrating regression able to directly capture this asymmetric response.

Following Balke and Fomby (1997), under the assumption that, among the set of k+1 variables in $\mathbf{z}_t = (y_t, \mathbf{x}_t')'$ there is at most a single cointegration relationship such that $u_t = \boldsymbol{\kappa}_1' \mathbf{z}_t$ is stationary, the simplest choice for a generating mechanism allowing for eventual departures from the equilibrium is a two-regime threshold AR(1) (TAR) model such as $\alpha_t = I(| q_{t-d} | \leq \gamma) + \alpha I(| q_{t-d} | > \gamma) = \alpha + (1 - \alpha) I(| q_{t-d} | \leq \gamma)$, where $|\alpha| < 1$. Therefore, cointegration only occurs in the outer regime given by large values of the threshold variable, so that for small values this one corresponds to a long-run disequilibrium relation (that is, no cointegration). The role of the threshold variable is played by the regression errors, i.e., $q_t = u_t$ with $d = 1$, which implies that, despite the assertion made by these authors that while "locally" u_t may have a unit root (when $| u_{t-1} | \leq \gamma$), "globally" it is stationary, the results in Petruccelli and Woolford (1984) and Chan, et.al. (1985), determine that a two-regime self exciting threshold AR(1) process (SETAR), such as $\alpha_t = \alpha_1 I(| q_{t-d} | \leq \gamma) + \alpha_2 I(| q_{t-d} | > \gamma)$, is ergodic and second-order stationary only if α_1, $\alpha_2 < 1$; $\alpha_1 \cdot \alpha_2 < 1$ and the error sequence driving the SETAR process has finite second moment. In any case, this simple formulation can be extended to more general structures as, a three-regime SETAR model given by $\alpha_t = \alpha_1 I(u_{t-d} \leq \gamma_1) + \alpha_2 I(\gamma_1 < u_{t-d} \leq \gamma_2) + \alpha_3 I(u_{t-d} > \gamma_2)$, with two thresholds, or

$$u_t = \alpha_{0t} + \alpha_t(L) u_{t-1} + \upsilon_t \qquad [7]$$

With multiple thresholds, where $\alpha_{0t} = \sum_{i=1}^{m+1} \alpha_{0i} I_{t-d}(\gamma_{i-1}, \gamma_i)$, and time-varying polynomial $\alpha_t(L) = \sum_{i=1}^{m+1} \alpha_i(L) I_{t-d}(\gamma_{i-1}, \gamma_i)$, with $I_{t-d}(\gamma_{i-1}, \gamma_i) = I(\gamma_{i-1} < u_{t-d} < \gamma_i)$, $\gamma_0 = -\infty$, and $\gamma_{m+1} = +\infty$. Generalizing the dynamics of the equilibrium error by allowing nonzero intercepts that can act as additional driving forces to push the series back towards the equilibrium bands and asymmetric thresholds, we obtain a nonlinear generating process that could satisfy the stationary conditions in Chan et.al. (1985).[4] In any case, the usual way to examine the existence and effects of this driving mechanism is through the corresponding threshold error correction model representation associated to this generalized version of a cointegrating system. By defining the non-singular matrix $\boldsymbol{\kappa} = (\boldsymbol{\kappa}_1, \boldsymbol{\kappa}_2)$, with $\boldsymbol{\kappa}_2$ a $(k+1)\times k$ matrix, with $\boldsymbol{\kappa}_1$ and $\boldsymbol{\kappa}_2$ linearly independent, i.e. $\boldsymbol{\kappa}_1'\boldsymbol{\kappa}_2 = \mathbf{0}_k'$, we have

$$\boldsymbol{\kappa}'\mathbf{z}_t = \begin{pmatrix} u_t \\ \boldsymbol{\xi}_t \end{pmatrix},$$

with $\boldsymbol{\xi}_t = \boldsymbol{\kappa}_2'\mathbf{z}_t$ the k-vector of stochastic trends, such that $\Delta\boldsymbol{\xi}_t = \boldsymbol{\upsilon}_t$ with $\boldsymbol{\upsilon}_t$ a stationary sequence given by $\boldsymbol{\upsilon}_t = \boldsymbol{\mu}_k + \mathbf{B}(L)\boldsymbol{\upsilon}_{t-1} + \mathbf{v}_t$, where $\mathbf{v}_t$ is a zero-mean iid (or martingale difference) sequence with finite second moments. Given that Δu_t can be written as $\Delta u_t = \alpha_{0t} - [1-\alpha_t(L)]u_{t-1} + \upsilon_t$ or, alternatively, as

$$\Delta u_t = \alpha_{0t} - (1-\alpha_t(1))u_{t-1} - \tilde{\alpha}_t(L)\Delta u_{t-1} + \upsilon_t$$

The above formulation implies the following time-varying error correction model representation

$$\Delta\mathbf{z}_t = \begin{pmatrix} \boldsymbol{\kappa}_1' \\ \boldsymbol{\kappa}_2' \end{pmatrix}^{-1} \begin{pmatrix} \Delta u_t \\ \boldsymbol{\upsilon}_{k,t} \end{pmatrix} = \boldsymbol{\mu}_t + \boldsymbol{\alpha}_t u_{t-1} + \mathbf{C}_t(L)\Delta\mathbf{z}_{t-1} + \mathbf{u}_t \qquad [8]$$

with parameters defined as $\boldsymbol{\mu}_t = \boldsymbol{\kappa}'^{-1}(\alpha_{0t}, \boldsymbol{\mu}_k')'$, $\boldsymbol{\alpha}_t = \boldsymbol{\kappa}'^{-1}(-(1-\alpha_t(1)), \mathbf{0}_k')'$, and

$$\mathbf{C}_t(L) = \begin{pmatrix} \boldsymbol{\kappa}_1' \\ \boldsymbol{\kappa}_2' \end{pmatrix}^{-1} \begin{pmatrix} -\tilde{\alpha}_t(L) & \mathbf{0}_k' \\ \mathbf{0}_k & \mathbf{B}(L) \end{pmatrix} \begin{pmatrix} \boldsymbol{\kappa}_1' \\ \boldsymbol{\kappa}_2' \end{pmatrix}$$

with $(k+1)\times 1$ error vector $\mathbf{u}_t = \boldsymbol{\kappa}'^{-1}(\upsilon_t, \mathbf{v}_t')'$, where the source of variability of the functional parameters in [8] is associated to the magnitude of the regression error u_t through the indicator functions introduced in the SETAR model [7] with a time-varying drift component. For a review of the specification and estimation of several generalized forms of the standard error correction model incorporating nonlinear adjustments based on the magnitude of disequilibrium error, see the recent paper by Seo (2011).

An alternative to the above formulation, arising from direct incorporation of the asymmetric response of the ERPT to unit changes in the nominal exchange rate into the systematic component of the regression model [6], is the so-called cointegrating regression model with threshold effects, first proposed by Gonzalo and Pitarakis (2006a), which is given by

$$y_t = \alpha_0 + \mathbf{x}_t'\boldsymbol{\beta} + (\kappa_0 + \mathbf{x}_t'\boldsymbol{\lambda})J_{t-d}(\gamma) + u_t \quad t = 1,\ldots,n \qquad [9]$$

with $J_{t-d}(\gamma) = I(q_{t-d} > \gamma)$, so that the long-run equilibrium relationship may change according to the magnitude of a threshold variable q_{t-d}, which it is assumed to be stationary and predetermined, and for technical reasons when dealing with OLS estimation the regression error must be restricted to be serially uncorrelated.

Once described these two competing and alternative models to accommodate the asymmetric response of the ERPT to changes in the nominal exchange rate, the rest of this section deals with the empirical analysis based on a wide set of testing procedures. First of all, we explore the existence of unit root behaviour in the individual series employing a number of robust semi parametric test statistics both for testing the null hypothesis of stationarity in levels against difference stationarity and for the reversed

[4] For a more detailed discussion on this question see, e.g., Balke and Fomby (1997)

hypothesis. Secondly, given the evidence of a stochastic trend component in the series analyzed, we explore the existence of a stable long-run relationship among these series in the framework of the static linear cointegrating regression [6], and also consider possible misspecifications due to structural instabilities in the long-run relationship and the existence of threshold effects, as described in equation [9]. Finally, we compute some recently proposed Wald-type test statistics to test the adequacy of the threshold cointegration assumption in two modified versions of the nonlinear error correction model described in [8].

4.1. Univariate analysis of stationarity and autoregressive unit root

All the testing procedures used in this stage are based on examining the behaviour of the OLS-detrended observations for the most commonly used representations of the deterministic component underlying the generating mechanism of each individual series as $z_t = d_t + \eta_t$, with $d_t = \boldsymbol{\alpha}'\boldsymbol{\tau}_t$ and $\boldsymbol{\tau}_t = (1, t, ..., t^p)'$ for p = 0, 1, and $\eta_t = \rho\eta_{t-1} + \varepsilon_t$. As indicated in note (a) of Table B, we employ a set of semi parametric statistics robust to serially correlated errors ε_t to test the null of a trend stationary behaviour, when $|\rho| < 1$, against the alternative of a unit root ($\rho = 1$), and for the reversed hypothesis. As presented in Table B, all these testing procedures reject the corresponding null hypothesis for large values of the estimated test statistic, except in the case of the variance-ratio statistic by Breitung (2002) that rejects the null of a fixed unit root for values smaller than the critical value. The statistic labelled as $\hat{H}_{n,p}$ is a modified version of the one proposed in McCabe et.al. (2006) to test the null of stationary cointegration against a heteroskedastic alternative and is given by

$$\hat{H}_{n,p} = \frac{1}{n\sqrt{12n\hat{\omega}_n^2(a_t)}} \sum_{t=1}^{n} t(\hat{\xi}_t^2 - \hat{\sigma}_{\xi,n}^2)$$

where $\hat{\xi}_t = \Delta z_t - \hat{\boldsymbol{\alpha}}_n' \Delta d_t$ are the OLS residuals from correcting for the deterministic component, and $\hat{\omega}_n^2(a_t)$ is the usual kernel estimator of the long-run variance of the sequence $a_t = \hat{\xi}_t^2 - \hat{\sigma}_{\xi,n}^2$, that is, $\hat{\omega}_n^2(a_t) = n^{-1}\sum_{-(n-1)}^{n-1} w(jq_n^{-1})\sum_{t=j+1}^{n} a_t a_{t-j}$. The limiting null distribution of all these test statistics is no standard and involves different functional of Brownian processes, except that of $\hat{H}_{n,p}$ which is standard normal, that is $\hat{H}_{n,p} \Rightarrow N(0,1)$ under the null hypothesis of a fixed unit root.

For the usual specifications of the deterministic component, that is only a constant term or a constant term and a linear trend component and even for the raw series, in many cases the general evidence support the existence of a unit root in each of the series considered. These results justify the subsequent treatment of the regression model relating nonstationary variables in the context of the cointegration analysis.

4.2. Cointegration analysis in a single-equation cointegrating regression. The linear case

Given the specification as in equation [5], that is

$$y_t = \alpha_0 + \mathbf{x}_t'\boldsymbol{\beta} + v_t \quad t = 1, ..., n \qquad [10]$$

with nonstationary I(1) regressors, we have that the normalized OLS estimation errors of α_0 and $\boldsymbol{\beta}$ are given by

$$n^{\kappa}(\hat{\alpha}_{0,n} - \alpha_0) = n^{-(1-\kappa)}\sum_{t=1}^{n} v_t - n^{-3/2}\sum_{t=1}^{n}\mathbf{x}_t'[n^{1/2+\kappa}(\hat{\boldsymbol{\beta}}_{k,n} - \boldsymbol{\beta}_k)]$$

and

$$n^{1/2+\kappa}(\hat{\boldsymbol{\beta}}_{k,n}-\boldsymbol{\beta}_k)=\left(n^{-1}\sum_{t=1}^{n}\tilde{\mathbf{x}}_t\tilde{\mathbf{x}}_t'\right)^{-1}n^{-(1-\kappa)}\sum_{t=1}^{n}\tilde{\mathbf{x}}_t v_t$$

respectively, with $\tilde{\mathbf{x}}_t=\mathbf{x}_t-n^{-1}\sum_{s=1}^{n}\mathbf{x}_s$ the OLS demeaned observations of the integrated regressors, where the power κ takes values ±1/2 depending on the stochastic properties of the error sequence v_t. It is a well known result that with stationary regression errors, with κ = 1/2, the OLS estimators of these model parameters are consistent at the rates $n^{1/2}$ and n, respectively, but their limiting distributions contains two sources of bias in finite samples induced by the endogeneity of the regressors and the possible serial correlation in the regression errors, that invalidates its use to built valid pivotal statistics. Among the existing alternative estimation methods to OLS, the semiparametric correction called Fully Modified OLS estimator (FM-OLS), proposed by Phillips and Hansen (1990), is a convenient alternative that allows to asymptotically remove these bias components. Recently, Vogelsang and Wagner (2014) propose a very simple transformation of a modified specification of the cointegrating regression that does not depend on any nuisance parameter and will produce asymptotically almost efficient estimates as compared with FM-OLS.[5] In order to test the null hypothesis of cointegration we employ the OLS and FM-OLS versions of the KPSS statistic for stationarity proposed by Shin (1994) (labeled as $\hat{C}_{n,p}$ and $\hat{C}_{n,p}^{+}$ in Table C), and the semiparametric statistics $S\hat{C}_{n,p}$ and $H\hat{C}_{n,p}$ proposed by McCabe, et.al. (2006), based on another modification of the OLS estimator called Asymptotically Instrumental Variable (AIV) estimator, that allows to testing two related null hypothesis of a generalized definition of stationary cointegration. Finally, to complete this basic analysis we also present the estimated values of the Z statistics proposed by Phillips (1987) to test the null hypothesis of no cointegration against stationary cointegration. Again, for the most commonly used specifications of the deterministic component, we find a general empirical consensus supporting the evidence of a stable linear long-run relationship among the model variables.

However, to empirically evaluate the potential source of asymmetric adjustment discussed above, we first consider the results of the Hansen's (1992) LM-type test for the null of structural stability in a linear cointegrating regression against the alternative of a martingale process describing the behavior of the regression parameter vector, that is $y_t=\mathbf{x}_t'\mathbf{A}_t+u_t$ with $\mathbf{A}_t=\mathbf{A}_{t-1}+\mathbf{v}_t$, where $\mathbf{v}_t$ a zero-mean martingale difference sequence with finite and time-varying covariance matrix, $E[\mathbf{v}_t\mathbf{v}_t']=\delta^2\mathbf{G}_t$, for different forms of the matrix $\mathbf{G}_t$, with $\delta^2=0$ representing the null hypothesis of time-invariant coefficients. To see the connection between these two seemly unrelated concepts we may rewrite equation [9] as

$$y_t=\alpha_0+\kappa_0 J_{t-d}(\gamma)+\mathbf{x}_t'(\boldsymbol{\beta}+\boldsymbol{\lambda}J_{t-d}(\gamma))+u_t=\theta_t(\gamma)+\mathbf{x}_t'\boldsymbol{\theta}_t(\gamma)+u_t \qquad [11]$$

which takes the general form of a functional-coefficient cointegrating regression equation as defined recently by Pitarakis (2012). On the other hand, combining [6], [7], and the martingale structure for the complete set of the model variables in $\mathbf{z}_t=(y_t,\mathbf{x}_t')'$, $\mathbf{z}_t=\mathbf{z}_{t-1}+(\varepsilon_{0,t},\boldsymbol{\varepsilon}_t')'$, then we get

[5] For more details on this new estimation method of cointegrating regressions, called Integrated Modified OLS (IM-OLS) estimator, see the cited paper and its application to the cointegrating regression with threshold effects in subsection 4.3.1 below.

$$y_t = \alpha_{0t} + \mathbf{x}_t'\boldsymbol{\theta}_t + v_t \quad t = 1,...,n \qquad [12]$$

with $\boldsymbol{\theta}_t = (1-\alpha_t(L))\boldsymbol{\beta}$, and $v_t = u_t + \alpha_t(L)(Y_{t-1} + \boldsymbol{\varepsilon}_t\boldsymbol{\beta})$, Again, we expect to find some type of instability in the analysis of [6] when omitting the nonlinear adjustment process. However, theoretical and numerical results not reported here indicate that Hansen's LM test behaves different in these two cases, being unable to significantly identify the time-varying structure of the functional coefficients in [11], with nontrivial power against the specification in [12].

4.3. Nonlinear extensions of the standard cointegrating regression model

This final section deals with the analysis of the two competing nonlinear extensions of the basic model considered above.

4.3.1. Cointegration with threshold effects

Following Gonzalo and Pitarakis (2006a), and for a general deterministic component in the cointegrating regression, we may define the two sets of regressors in [9] as $\mathbf{m}_t = (\boldsymbol{\tau}_t', \mathbf{x}_t')'$, and $\mathbf{m}_t(\gamma) = \mathbf{m}_t J_{t-d}(\gamma)$, with $J_{t-d}(\gamma) = I(q_{t-d} > \gamma)$, so that the nonlinear cointegrating regression with threshold effects can be written as

$$y_t = \boldsymbol{\theta}_0'\mathbf{m}_t + \boldsymbol{\theta}_1'\mathbf{m}_t(\gamma) + u_t = \boldsymbol{\theta}'\mathbf{A}_t(\gamma) + u_t \quad t = 1,...,n \qquad [13]$$

where the null hypothesis of no threshold effects (that is, linear cointegration) can be tested by the LM-type test statistic defined as

$$L\hat{M}_n(\gamma) = \frac{1}{\hat{\sigma}_n^2}\hat{\mathbf{F}}_n'(\gamma)\mathbf{M}_n^{-1}(1-\gamma)\hat{\mathbf{F}}_n(\gamma) \qquad [14]$$

where $\hat{\mathbf{F}}_n(\gamma) = \sum_{t=1}^n \mathbf{m}_t(\gamma)\hat{v}_t$, with $\hat{v}_t = Y_t - \hat{\boldsymbol{\theta}}_{0,n}'\mathbf{m}_t$ the OLS residuals from estimation under the null, and

$$\mathbf{M}_n(1-\gamma) = \sum_{t=1}^n \mathbf{m}_t(\gamma)(\mathbf{I}_{m+k} - \mathbf{m}_t'\mathbf{Q}_n^{-1}\mathbf{m}_t)\mathbf{m}_t'(\gamma)$$

with $\mathbf{Q}_n = \sum_{t=1}^n \mathbf{m}_t\mathbf{m}_t'$. Note that the LM statistic can also be computed as $L\hat{M}_n(\gamma) = n(1 - \sum_{t=1}^n \hat{\xi}_t^2(\gamma)/\sum_{t=1}^n \hat{v}_t^2)$, with $\hat{\xi}_t(\gamma)$ the OLS residuals from the auxiliary regression $\hat{v}_t = \hat{\boldsymbol{\delta}}_0\mathbf{m}_t + \boldsymbol{\theta}_1'\mathbf{m}_t(\gamma) + \xi_t$, for any given value of the threshold parameter. For any admissible value of γ, with any additional assumptions and the results obtained in Caner and Hansen (2001), the limiting null distribution of the LM test is nonstandard, but coincides with that of Andrews (1993) when testing for parameter instability in a linear regression with stationary regressors. Thus, given a grid of values for the threshold parameter, the global statistic is given by $\sup L\hat{M}_n(\gamma) = \max_{\gamma\in[\gamma_L,\gamma_U]} L\hat{M}_n(\gamma)$. Also, as a byproduct of this testing procedure, we get an estimate of the threshold parameter as $\hat{\gamma}_n = \arg\max_{\gamma\in[\gamma_L,\gamma_U]} L\hat{M}_n(\gamma)$, where it can be proved that this estimate coincides with $\hat{\gamma}_n = \arg\min_{\gamma\in[\gamma_L,\gamma_U]} SC_n(\hat{\boldsymbol{\theta}}_n(\gamma),\gamma)$, where $SC_n(\hat{\boldsymbol{\theta}}_n(\gamma),\gamma)$ is the sum of squared residuals of the profiled OLS estimation of the complete regression with threshold components, that is, $SC_n(\hat{\boldsymbol{\theta}}_n(\gamma),\gamma) = \sum_{t=1}^n (y_t - \hat{\boldsymbol{\theta}}_n'(\gamma)\mathbf{A}_t(\gamma))$, with

$$\hat{\boldsymbol{\theta}}_n(\gamma) = \left(\sum_{t=1}^n \mathbf{A}_t(\gamma)\mathbf{A}_t'(\gamma)\right)^{-1}\sum_{t=1}^n \mathbf{A}_t(\gamma)y_t$$

the OLS estimator in [13]. Given the factorization $\mathbf{m}_t = \mathbf{W}_n\mathbf{m}_{nt}$, with weighting matrix $\mathbf{W}_n = \mathrm{diag}(\boldsymbol{\Gamma}_n^{-1}, \sqrt{n}\mathbf{I}_{k,k})$, where $\boldsymbol{\Gamma}_n$ such that $\boldsymbol{\tau}_{nt} = \boldsymbol{\Gamma}_n\boldsymbol{\tau}_t \to \boldsymbol{\tau}(r)$, then we get

$$n^{\kappa}\begin{pmatrix}\mathbf{W}_n(\hat{\boldsymbol\theta}_{0,n}(\gamma)-\boldsymbol\theta_0)\\ \mathbf{W}_n(\hat{\boldsymbol\theta}_{1,n}(\gamma)-\boldsymbol\theta_1)\end{pmatrix}=\begin{pmatrix}\mathbf{Q}_n^{-1}(\gamma)n^{-(1-\kappa)}\sum_{t=1}^{n}\mathbf{m}_{nt}I_{t-d}(\gamma)u_t\\ \mathbf{M}_n^{-1}(1-\gamma)n^{-(1-\kappa)}\sum_{t=1}^{n}\mathbf{m}_{nt}J_{t-d}(\gamma)u_t-\mathbf{Q}_n^{-1}(\gamma)n^{-(1-\kappa)}\sum_{t=1}^{n}\mathbf{m}_{nt}u_t\end{pmatrix}$$

where $\mathbf{Q}_n(\gamma)=n^{-1}\sum_{t=1}^{n}\mathbf{m}_{nt}\mathbf{m}'_{nt}I_{t-d}(\gamma)$, which implies that, under the same set of basic assumption as in Gonzalo and Pitarakis (2006a) relating the stochastic properties of the stationary threshold variable and the regression error sequence, we have that under cointegration and the true value of the threshold parameter, $\gamma=\gamma_0$,

$$n^{1/2}\mathbf{W}_n(\hat{\boldsymbol\theta}_{1,n}(\gamma_0)-\boldsymbol\theta_1)\Rightarrow-\frac{1}{F(\gamma_0)(1-F(\gamma_0))}\mathbf{M}^{-1}\int_0^1\mathbf{m}(s)dK_u(s,F(\gamma_0))$$

where $\mathbf{M}=\int_0^1\mathbf{m}(s)\mathbf{m}'(s)ds>0$, with $\mathbf{m}(s)=(\boldsymbol\tau'(s),\mathbf{B}'(s))'$, and $K_u(s,F(\gamma_0))$ the zero-mean Gaussian process $K_u(s,F(\gamma_0))=B_u(s,F(\gamma_0))-F(\gamma_0)B_u(s,1)$, called a Kiefer process, with $B_u(s,F(\gamma_0))$ a two-parameter Brownian (or Brownian sheet) process, where $F(\gamma)$ can be called the threshold quantile with $F(\cdot)$ the cumulate distribution function of the stationary threshold variable. The main relevant characteristic of this limiting result is that the Kiefer process is statistically independent of $\mathbf{m}(s)$ which means that this limiting distribution is mixed Gaussian. However, the same does not apply to the limiting distribution of the normalized estimation error of $\boldsymbol\theta_0$. To avoid these complications, we propose to adapt the IM-OLS estimator formulated by Vogelsang and Wagner (2014) which consists in the OLS estimation of the integrated modified version of the cointegrating regression given by

$$S_t=\boldsymbol\theta_0'\mathbf{g}_t+\boldsymbol\theta_1'\mathbf{g}_t(\gamma)+\boldsymbol\pi'\mathbf{x}_t+Z_t=\boldsymbol\theta'\mathbf{B}_t(\gamma)+\boldsymbol\pi'\mathbf{x}_t+Z_t\quad t=1,\dots,n \qquad [15]$$

where $S_t=\sum_{j=1}^{t}y_j$, $\mathbf{g}_t=\sum_{j=1}^{t}\mathbf{m}_j$, and $Z_t=U_t-\boldsymbol\pi'\mathbf{x}_t$, with $U_t=\sum_{j=1}^{t}u_j$. Taking into account that the normalized error sequence in [15] can be written as $n^{-(1-\kappa)}Z_t=n^{-(1-\kappa)}U_t-n^{-(1/2-\kappa)}\boldsymbol\pi'n^{-1/2}\mathbf{x}_t$, with the value of κ chosen as before so that the limiting distribution is well defined both under cointegration and no cointegration, then we have

$$\begin{pmatrix}n^{\kappa}\mathbf{W}_n^*(\tilde{\boldsymbol\theta}_n(\gamma)-\boldsymbol\theta)\\ n^{-1/2+\kappa}(\tilde{\boldsymbol\pi}_n(\gamma)-\boldsymbol\pi)\end{pmatrix}=\left(n^{-1}\sum_{t=1}^{n}\begin{pmatrix}\mathbf{B}_{nt}(\gamma)\\ \mathbf{x}_{nt}\end{pmatrix}(\mathbf{B}'_{nt}(\gamma),\mathbf{x}'_{nt})\right)n^{-1}\sum_{t=1}^{n}\begin{pmatrix}\mathbf{B}_{nt}(\gamma)\\ \mathbf{x}_{nt}\end{pmatrix}n^{-(1-\kappa)}Z_t$$

with $\mathbf{x}_{nt}=n^{-1/2}\mathbf{x}_t$, and $\mathbf{W}_n^*=\mathrm{diag}(\mathbf{W}_n,\mathbf{W}_n)$. Then, given the true value of the threshold parameter $\gamma=\gamma_0$ and under cointegration, we obtain the following limiting distribution[6]

$$\begin{pmatrix}n^{1/2}\mathbf{W}_n^*(\tilde{\boldsymbol\theta}_n(\gamma)-\boldsymbol\theta)\\ \tilde{\boldsymbol\pi}_n(\gamma)-\boldsymbol\pi\end{pmatrix}\Rightarrow\left(\int_0^1\mathbf{g}(s,\gamma_0)\mathbf{g}'(s,\gamma_0)ds\right)^{-1}\int_0^1\mathbf{g}(s,\gamma_0)B_{u.x}(s)ds$$

which is zero mean mixed Gaussian (given that $E[\mathbf{g}(s,\gamma_0)B_{u.x}(s)]=\mathbf{0}$), with

$$\mathbf{g}(s,\gamma_0)=\begin{pmatrix}\mathbf{B}(s,\gamma_0)\\ \mathbf{B}(s)\end{pmatrix},\ \mathbf{B}(s,\gamma_0)=\begin{pmatrix}\mathbf{g}(s)\\ (1-F(\gamma_0))\mathbf{g}(s)\end{pmatrix}$$

where $\mathbf{g}(s)=\int_0^s\mathbf{m}(a)da$ is the weak limit of $\mathbf{g}_{nt}=n^{-1}\sum_{j=1}^{t}\mathbf{m}_{nj}$. The main advantage of this result is that, in general, it allows to built asymptotically pivotal Wald-type test statistics. However, in our application of this limiting result we are not going to develop

[6] The formal proof of this result is not presented here, can be requested from the authors, although it only requires the application of standard partial results on weak convergence for each component and the Continuous Mapping Theorem.

this possibility, but we will only use the superconsistent point estimates of the slope parameters determining the ERPT. These point estimates are shown at bottom of Table C, together with the OLS estimates of the cointegrating regression with threshold effects (labeled as IM-OLS threshold and OLS threshold, respectively), and the OLS estimates of the benchmark simple model in equation [5] and the linear cointegrating regression in equation [6].

In both the short and long-run, we find evidence that PCP hypothesis held in the twelve Eurozone countries considered in the analysis. That is, ERPT is complete since producers do not to adjust the mark-up when exchange rate fluctuates, passing all the variation into prices of imports expressed in local currency. Although, Campa and González-Mínguez (2006) also found a relatively high degree of ERPT in the Euro Area, their results suggest partial ERPT.

4.3.2. Threshold cointegration

Finally, we consider the search for empirical evidence supporting the assumption that there exists an asymmetric adjustment mechanism drive by the disequilibrium error sequence, as has been described above. Since the work of Balke and Fomby (1997), there have been a number of testing procedures to test for the existence of threshold cointegration mainly by making use of the nonlinear extension of the error correction model (see, e.g., Hansen and Seo (2002), Seo (2006), and Gonzalo and Pitarakis (2006b)), while that Enders and Siklos (2001) employ a threshold version of the ADF regression to obtain a pseudo-T ratio statistic in the same spirit as the Engle and Granger (1987) methodology. Recently, Li and Lee (2010) proposed new testing procedures for threshold cointegration using a single-equation autoregressive distributed lag (ADL) regression model, with a not pre specified cointegrating vector, by generalizing the ADL-based cointegration tests in the linear case proposed by Boswijk (1994) (BO test) and Banerjee, et.al. (1998) (BDM test). Before introducing the particular expressions of the ADL-type representation of the error correction models used to implementing these testing procedures, Li and Lee (2010) propose to consider, following the same idea as in Enders and Granger (1998) in the context of the univariate analysis of a threshold unit root process, two different types of indicator functions based on the estimated disequilibrium errors arising from the OLS estimation of the linear cointegrating regression, $\hat{v}_t = y_t - \hat{d}_{t,n} - \hat{\boldsymbol{\beta}}_n' \mathbf{x}_t$, with $\hat{d}_{t,n}$ the estimated deterministic component. Particularly, these authors propose to consider the *Indicator A* defined as

$$I_{t-d}(\gamma) = I(\hat{v}_{t-d} < Q_{n,t-d}(\gamma))$$

where the threshold value, $Q_{n,t-d}(\gamma)$, is given by the γth percentile of the empirical distribution of $\hat{v}_{t-d}$ or, alternatively, the *Indicator B* function (also called the "momentum" indicator function), given by

$$I_{t-d}(\gamma) = I(\Delta\hat{v}_{t-d} < q_{n,t-d}(\gamma))$$

where $q_{n,t-d}(\gamma)$ denotes the γth percentile element of the empirical distribution of the stationary series $\Delta\hat{v}_{t-d}$. For a more detailed discussion on these indicators and the asymptotics they determine, see Li and Lee (2010). Thus, given the nonlinear specification of the associated error correction model incorporating these threshold effects, these authors propose to examine their statistical significance in the context of the augmented version of the error correction model as an ADL regression model. They provide two versions of the resulting testing procedure called, the threshold BO test that examines the significance of the coefficients of the lagged regressand as well as the lagged conditioning variables in the ADL regression model

$$\Delta y_t = \boldsymbol{\alpha}_p' \boldsymbol{\tau}_{p,t} + \mathbf{B}_1' \mathbf{z}_{t-1} I_{t-d}(\gamma) + \mathbf{B}_2' \mathbf{z}_{t-1} J_{t-d}(\gamma) + \sum_{i=1}^{l} \phi_i \Delta y_{t-i} + \sum_{i=0}^{l} \boldsymbol{\Phi}_i' \Delta \mathbf{x}_{t-i} + u_{1,t}$$

with $J_{t-d}(\gamma) = 1 - I_{t-d}(\gamma)$, where the null hypothesis of no threshold cointegration is given by $\mathbf{B}_1 = \mathbf{B}_2 = \mathbf{0}_{k+1}$, and the threshold BDM test, that involves testing only the significance of the coefficients of the lagged regressand and requires adding leads of differences of conditioning variables when strict exogeneity fails. In this case, the ADL regression model is written as

$$\Delta y_t = \boldsymbol{\alpha}_p' \boldsymbol{\tau}_{p,t} + (\kappa_{11} y_{t-1} + \boldsymbol{\beta}_1' \mathbf{x}_{t-1}) I_{t-d}(\gamma) + (\kappa_{12} y_{t-1} + \boldsymbol{\beta}_2' \mathbf{x}_{t-1}) J_{t-d}(\gamma)$$
$$+ \sum_{i=1}^{l_2} \phi_i \Delta y_{t-i} + \sum_{i=-l_1}^{l_2} \boldsymbol{\Phi}_i' \Delta \mathbf{x}_{t-i} + v_t$$

where the null hypothesis of no threshold cointegration is given by $\kappa_{11} = \kappa_{12} = 0$. In both cases, the Wald-type statistics are, for a given threshold value, equal to the pseudo-F ratio test statistic computed from the OLS estimation of the corresponding restricted and unrestricted versions of the ADL regression model.

Table D in the appendix shows the results for the two test statistics for the two versions of the indicator function. Globally, there appears to be little statistical support to the assumption of an asymmetric transmission of the changes in the nominal exchange rate on the magnitude of the ERPT.

Again, evidence of complete ERPT across the Euro Area countries is found. Consequently, the degree of ERPT seems to be homogeneous, i.e. PCP hypothesis held in all of them, regardless the testing procedure used. These results contrast to the ones obtained by Campa and Gonzalez-Mínguez (2006) where differences in the degree that a common exchange rate movement gets transmitted into consumer prices exist across Euro Area countries. Conversely, they are similar to the results obtained by Saiki (2011) where evidence for convergence of exchange rate pass-through for the euro area economies was found.

5. Conclusions

The purpose of this study is to provide new evidences on the ERPT into imports prices on the Euro Area. As an alternative to the more traditional modeling based on a linear, and possibly stable, long-run relationship among the system variables, there could be some theoretical arguments indicating the possible existence of a nonlinear adjustment mechanism determining more accurate measures of the ERPT. Thus, the main objective of the paper is to empirically evaluate this evidence on the basis of two different, but alternative, specifications incorporating such nonlinear behavior, namely: (i) the so-called threshold cointegration model by Balke and Fomby (1997), where this effect is incorporate through the disequilibrium error term following a self-exiting threshold autoregression (SETAR) process, (ii) and a functional-coefficient cointegrating regression model with a time-varying cointegration vector with values changing from one regime to another according to the magnitude of a predetermined threshold variable proposed by Gonzalo and Pitarakis (2006a). Globally, the empirical analysis provides strongly support for the no existence of such a significant effect for the series and time period used as well as evidence for a complete and time-invariant ERPT. Therefore, the degree of ERPT seems to be homogeneous, i.e. PCP hypothesis held in all of twelve Euro Area countries considered in the analysis.

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Appendixes

A. Figures

Figure 1. *Imports from outside the Euro Area*

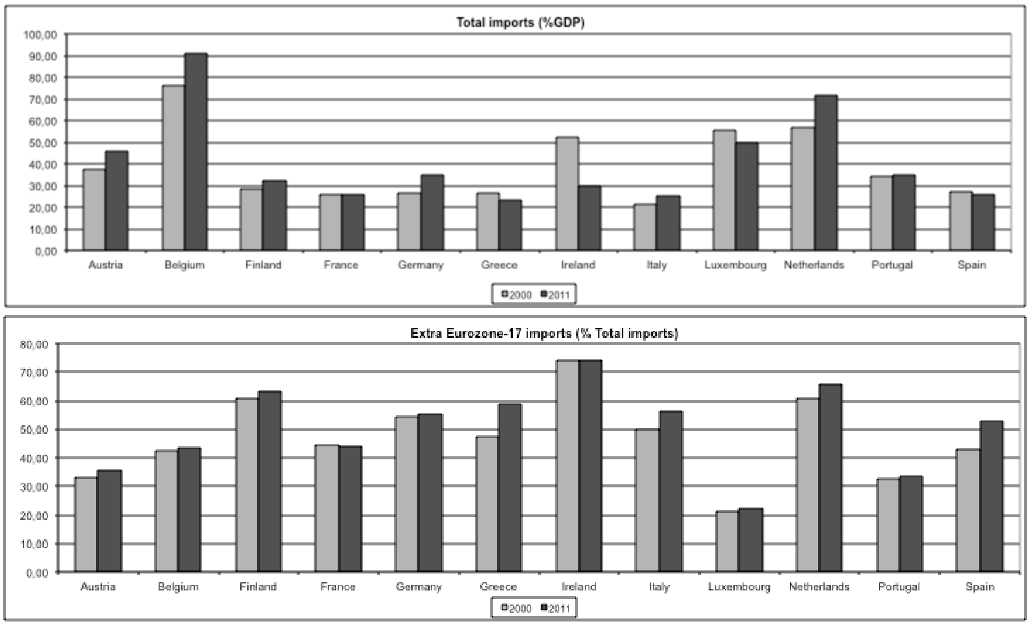
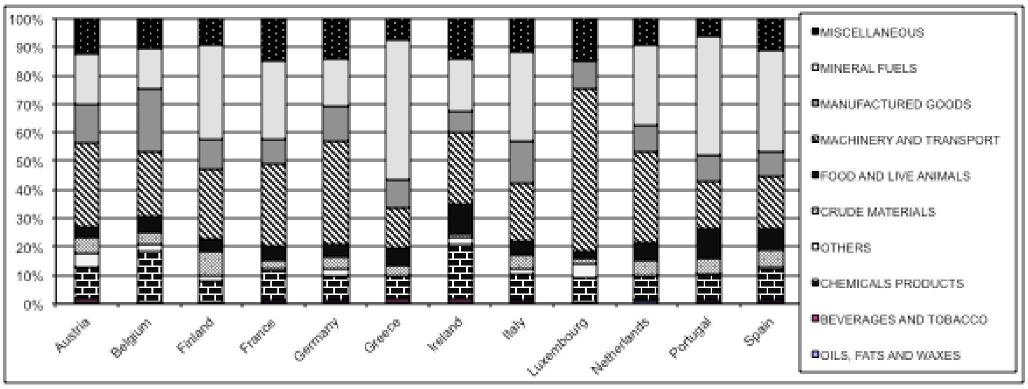


Figure 2. *Composition by product of Extra-Eurozone imports (2011)*



B. Univariate semiparametric tests for stationarity/unit root

Table B. *Univariate analysis of nonstationarity based on semiparametric tests*

			Spain		France		Netherlands	
		er	*mp*	*fp*	*mp*	*fp*	*mp*	*fp*
Stationarity tests								
$\hat{K}_{n,p}$	No det.	1.538[b]	4.873	4.771	4.886	4.786	4.916	4.820
	$p = 0$	1.254[a]	1.320	1.436	1.344	1.447	1.022	1.416
	$p = 1$	0.249[a]	0.122	0.165	0.149	0.171	0.183	0.141
$C\hat{S}_{n,p}$	No det.	1.104	0.038	0.079	0.032	0.072	0.019	0.059
	$p = 0$	1.631[a]	1.740	1.715	1.728	1.706	1.546	1.693
	$p = 1$	0.885[b]	0.672	0.903	0.708	0.920	0.759	0.851
$V\hat{S}_{n,p}$	No det.	0.977[a]	1.254	1.261	1.254	1.261	1.253	1.261
	$p = 0$	0.301[a]	0.289	0.307	0.287	0.303	0.272	0.303
	$p = 1$	0.249[a]	0.122	0.165	0.149	0.171	0.183	0.141
$J\hat{C}_{n,p}$	No det.	3.889	4.208	4.002	4.255	3.999	4.305	4.005
	$p = 0$	2.313[a]	2.113	2.382	2.156	2.367	1.972	2.328
	$p = 1$	1.122	1.035	1.351	1.051	1.352	1.266	1.218
Unit root tests								
$\bar{\rho}_{n,p}$	No det.	0.102[c]	0.330	0.324	0.331	0.325	0.333	0.327
	$p = 0$	0.081[c]	0.082	0.093	0.083	0.093	0.061	0.091
	$p = 1$	0.013[c]	0.007	0.009	0.008	0.009	0.009	0.007
$\hat{H}_{n,p}$	No det./$p = 0$	0.034	0.035	0.107	0.066	0.113	0.114	0.137
	$p = 1$	0.035	0.019	0.092	0.048	0.100	0.108	0.125

		Germany		Italy		Ireland	
		mp	*Fp*	*mp*	*fp*	*mp*	*fp*
Stationarity tests							
$\hat{K}_{n,p}$	No det.	4.910	4.814	4.851	4.746	4.944	4.853
	$p = 0$	1.297	1.430	1.394	1.469	0.614	1.389
	$p = 1$	0.138	0.216	0.175	0.149	0.201	0.175
$C\hat{S}_{n,p}$	No det.	0.022	0.062	0.046	0.088	0.008	0.046
	$p = 0$	1.748	1.707	1.757	1.727	1.175	1.658
	$p = 1$	0.711	0.935	0.801	0.824	0.843	0.890
$V\hat{S}_{n,p}$	No det.	1.256	1.263	1.252	1.259	1.255	1.262
	$p = 0$	0.288	0.309	0.306	0.309	1.226	0.296
	$p = 1$	0.138	0.216	0.174	0.149	0.201	0.175
$J\hat{C}_{n,p}$	No det.	4.218	3.958	4.222	4.015	4.366	3.984
	$p = 0$	2.049	2.389	2.224	2.361	1.686	2.366
	$p = 1$	1.134	1.371	1.411	1.157	1.349	1.293
Unit root tests							
$\bar{\rho}_{n,p}$	No det.	0.333	0.327	0.329	0.322	0.335	0.329
	$p = 0$	0.080	0.093	0.087	0.095	0.033	0.089
	$p = 1$	0.008	0.012	0.009	0.007	0.011	0.009
$\hat{H}_{n,p}$	No det./$p = 0$	–0.017	0.021	0.091	0.130	–0.052	0.068
	$p = 1$	–0.036	0.019	0.086	0.123	–0.056	0.065

Table B. *Univariate analysis of nonstationarity based on semiparametric tests (continuation)*

		Greece		Portugal		Belgium	
		mp	*fp*	*mp*	*fp*	*mp*	*fp*
Stationarity tests							
$\hat{K}_{n,p}$	No det.	4.845	4.738	4.864	4.761	4.909	4.813
	$p = 0$	1.386	1.442	1.319	1.434	1.237	1.440
	$p = 1$	0.093	0.175	0.125	0.153	0.178	0.179
$C\hat{S}_{n,p}$	No det.	0.049	0.092	0.042	0.083	0.022	0.062
	$p = 0$	1.772	1.722	1.739	1.718	1.674	1.703
	$p = 1$	0.541	0.968	0.675	0.886	0.809	0.851
$V\hat{S}_{n,p}$	No det.	1.254	1.262	1.253	1.261	1.254	1.262
	$p = 0$	0.283	0.301	0.293	0.308	0.285	0.306
	$p = 1$	0.093	0.175	0.125	0.153	0.178	0.179
$J\hat{C}_{n,p}$	No det.	4.199	4.009	4.181	4.005	4.288	3.989
	$p = 0$	2.099	2.357	2.141	2.386	2.099	2.384
	$p = 1$	2.115	1.435	0.989	1.416	1.329	1.069
Unit root tests							
$\bar{\rho}_{n,p}$	No det.	0.329	0.321	0.329	0.323	0.333	0.327
	$p = 0$	0.085	0.093	0.081	0.092	0.076	0.094
	$p = 1$	0.004	0.009	0.006	0.008	0.009	0.009
$\hat{H}_{n,p}$	No det./$p = 0$	−0.087	−0.087	0.089	0.121	0.030	0.100
	$p = 1$	−0.088	−0.083	0.079	0.107	0.019	0.089

		Luxembourg		Finland		Austria	
		mp	*fp*	*mp*	*fp*	*mp*	*fp*
Stationarity tests							
$\hat{K}_{n,p}$	No det.	5.038	4.959	4.879	4.779	4.896	4.798
	$p = 0$	1.205	0.210	1.297	1.413	1.347	1.454
	$p = 1$	0.071	0.183	0.117	0.171	0.208	0.160
$C\hat{S}_{n,p}$	No det.	0.033	0.013	0.035	0.077	0.026	0.067
	$p = 0$	1.610	0.983	1.719	1.708	1.667	1.687
	$p = 1$	0.746	0.839	0.672	0.880	0.856	0.827
$V\hat{S}_{n,p}$	No det.	1.258	1.266	1.255	1.262	1.253	1.261
	$p = 0$	0.275	0.182	0.286	0.305	0.289	0.301
	$p = 1$	0.071	0.183	0.117	0.171	0.208	0.160
$J\hat{C}_{n,p}$	No det.	4.129	3.839	4.206	3.997	4.319	4.008
	$p = 0$	2.163	1.588	1.999	2.386	2.212	2.376
	$p = 1$	1.903	1.586	1.134	1.088	1.266	1.042
Unit root tests							
$\bar{\rho}_{n,p}$	No det.	0.342	0.337	0.331	0.324	0.332	0.325
	$p = 0$	0.068	0.010	0.081	0.092	0.083	0.094
	$p = 1$	0.003	0.009	0.006	0.009	0.012	0.008
$\hat{H}_{n,p}$	No det./$p = 0$	0.195	0.213	0.024	0.088	0.160	0.000
	$p = 1$	0.195	0.213	0.011	0.078	0.143	−0.004

Notes. (a) The semiparametric statistics for testing the null of stationary against a fixed unit root are the KPSS statistic ($\hat{K}_{n,p}$) by Kwiatkowski, et.al. (1992), the maximum absolute fluctuation of the scaled partial sum of OLS residuals proposed by Xiao (2001) ($C\hat{S}_{n,p}$), the rescaled variance-ratio statistic by Giraitis, et.al. (2003) ($V\hat{S}_{n,p}$), and $J\hat{C}_{n,p}$, the statistic proposed by Xiao and Lima (2007) to testing the null of weak, or second-order, stationarity. On the other hand, the semiparametric statistics for testing the reverse hypothesis, are the one proposed by Breitung (2002) ($\bar{\rho}_{n,p}$), and $\hat{H}_{n,p}$ the statistic proposed by McCabe, et.al. (2006) against a heteroskedastic integrated alternative. **(b)** [a,b,c] indicate statistically significant results at 1, 5, and 10% level. **(c)** All these statistics, except the variance-ratio $\bar{\rho}_{n,p}$ are computed with a fixed bandwidth $m_n = [8(n/100)^{1/3}]$ and the Bartlett kernel.

C. Regression results. Cointegration analysis based on the cointegrating regression

Table C. *Testing for cointegration rank, single-equation tests for cointegration and parameter instability, and ERPT estimates*

		Country	Spain	France	Netherlands	Germany
Testing for the cointegration rank						
$\hat{\Lambda}_{n,p}(q)$	$p = 0$	$q = 1$	10.74	10.64	10.93	10.58
		$q = 2$	92.70	84.38	84.22	92.29
	$p = 1$	$q = 1$	72.65	69.90	67.73	73.57
		$q = 2$	267.27	299.76	289.93	246.88
Single-equation tests for cointegration (case: $p = 0$)						
$\hat{Z}_{n,p}(\alpha)$			–164.65	–155.86	–131.96	–113.77
$\hat{Z}_{n,p}(\tau)$			–41.72	–85.76	–51.53	467.47
$\hat{C}_{n,p}$			0.069	0.044	0.047	0.039
$\hat{C}^{+}_{n,p}$			0.043	0.045	0.046	0.035
$S\hat{C}_{n,p}$			1.006 (0.843)	–0.001 (0.499)	–0.600 (0.274)	–0.955(0.169)
$H\hat{C}_{n,p}$			–0.022 (0.412)	–0.207 (0.418)	–0.093 (0.463)	–0.217(0.414)
Tests for parameter instability (case: $p = 0$)						
$\hat{L}_{n,p}$			0.233 (0.25)	0.189 (0.33)	0.148 (0.41)	0.123(0.469)
$\hat{L}^{+}_{n,p}$			0.210 (0.29)	0.299 (0.16)	0.143 (0.42)	0.128(0.457)
$\sup L\hat{M}_{n,p}$			15.37	6.58	5.89	12.89
$F(\hat{\gamma}_n)$			0.22	0.00	1.00	0.57
Estimates of the ERPT and tests						
DL regression	S/R		–0.998 (0.59)	–0.995 (1.15)	–1.003 (–0.81)	–1.001 (–0.11)
	L/R		–1.001 (–0.13)	–0.998 (0.28)	–1.000 (–0.06)	–0.996 (–0.49)
OLS linear			–0.999	–1.001	–1.002	–0.999
OLS threshold			$-1.002-0.002\,J_{t-1}$	–1.001	–1.002	$-0.998-0.003J_{t-1}$
IM-OLS threshold			$-1.000-0.0002J_{t-1}$	–1.000	–1.003	$-0.999-0.002J_{t-1}$

Table C. *Testing for cointegration rank, single-equation tests for cointegration and parameter instability, and ERPT estimates (continuation)*

		Country	Italy	Ireland	Greece	Portugal
Testing for the cointegration rank						
$\hat{\Lambda}_{n,p}(q)$	$p = 0$	$q = 1$	10.48	11.14	10.80	10.87
		$q = 2$	79.13	88.97	98.22	90.78
	$p = 1$	$q = 1$	67.69	68.49	72.91	71.21
		$q = 2$	257.55	335.24	328.24	282.11
Single-equation tests for cointegration (case: $p = 0$)						
$\hat{Z}_{n,p}(\alpha)$			−83.76	−139.05	−158.99	−158.07
$\hat{Z}_{n,p}(\tau)$			307.62	−0.48	−225.68	106.24
$\hat{C}_{n,p}$			0.064	0.062	0.086	0.051
$\hat{C}^{+}_{n,p}$			0.057	0.071	0.092	0.057
$S\hat{C}_{n,p}$			−0.725 (0.23)	−0.439 (0.33)	0.333 (0.63)	−0.641 (0.26)
$H\hat{C}_{n,p}$			−0.241 (0.41)	−0.076 (0.47)	−0.255 (0.39)	−0.211 (0.42)
Tests for parameter instability (case: $p = 0$)						
$\hat{L}_{n,p}$			0.209 (0.29)	0.439 (0.06)	0.401 (0.8)	0.491 (0.05)
$\hat{L}^{+}_{n,p}$			0.274 (0.19)	0.341 (0.13)	0.312 (0.15)	0.666 (0.02)
$\sup L\hat{M}_{n,p}$			7.59	5.81	10.08	3.179
$F(\hat{\gamma}_n)$			0.12	0.03	0.90	0.26
Estimates of the ERPT and tests						
DL regression	S/R		−1.002 (−0.57)	−1.000 (−0.03)	−1.004 (−1.42)	−0.998 (0.38)
	L/R		−1.004 (−0.77)	−1.001 (−0.11)	−1.004 (−0.69)	−1.001 (−0.11)
OLS linear			−1.001	−0.999	−1.000	−1.000
OLS threshold			$-1.001+0.0007J_{t-1}$	$-0.989+0.011J_{t-1}$	$-1.001+0.003J_{t-1}$	$-1.000-0.001J_{t-1}$
IM-OLS threshold			$-0.994-0.006J_{t-1}$	$-0.947-0.052J_{t-1}$	$-1.001+0.013J_{t-1}$	$-1.006+0.006J_{t-1}$

		Country	Belgium	Luxembourg	Finland	Austria
Testing for the cointegration rank						
$\hat{\Lambda}_{n,p}(q)$	$p = 0$	$q = 1$	10.52	11.66	10.89	10.45
		$q = 2$	84.28	133.38	93.09	75.41
	$p = 1$	$q = 1$	70.74	72.02	72.13	65.82
		$q = 2$	261.25	464.06	257.01	302.01
Single-equation tests for cointegration (case: $p = 0$)						
$\hat{Z}_{n,p}(\alpha)$			−137.81	−131.02	−134.41	−116.12
$\hat{Z}_{n,p}(\tau)$			330.84	458.72	120.67	409.09
$\hat{C}_{n,p}$			0.053	0.179	0.056	0.049
$\hat{C}^{+}_{n,p}$			0.044	0.075	0.062	0.048
$S\hat{C}_{n,p}$			0.419 (0.66)	0.826 (0.79)	0.358 (0.64)	−0.808 (0.21)
$H\hat{C}_{n,p}$			−0.276 (0.39)	0.022 (0.51)	−0.193 (0.42)	−0.205 (0.42)
Tests for parameter instability (case: $p = 0$)						
$\hat{L}_{n,p}$			0.214 (0.28)	0.459 (0.06)	0.188 (0.33)	0.249 (0.23)
$\hat{L}^{+}_{n,p}$			0.209 (0.29)	1.594 (-----)	0.247 (0.23)	0.258 (0.22)
$\sup L\hat{M}_{n,p}$			10.98	4.043	4.39	5.54
$F(\hat{\gamma}_n)$			0.59	0.47	0.67	1.00
Estimates of the ERPT and tests						
DL regression	S/R		−0.992 (1.49)	−0.997 (1.68)	−0.998 (0.59)	−1.001 (−0.15)
	L/R		−0.997 (0.30)	−0.999 (0.01)	−0.995 (0.74)	−1.004 (−0.43)
OLS linear			−0.998	−1.000	−1.002	−1.000
OLS threshold			$-0.997-0.001J_{t-1}$	$-0.999-0.001J_{t-1}$	$-1.001-0.0004J_{t-1}$	−1.001
IM-OLS threshold			$-0.998-0.001J_{t-1}$	$-1.000-0.001J_{t-1}$	$-1.001-0.004J_{t-1}$	−1.001

D. Testing for threshold cointegration

Table D. *Single-equation ADL tests for threshold cointegration*

Country	Threshold BO Test					
	Indicator A			Indicator B		
	No det.	$p = 0$	$p = 1$	No det.	$p = 0$	$p = 1$
Spain	4.33	19.13	114.11[a]	18.99[c]	7.83	72.01[a]
France	4.33	15.96	96.57[a]	7.49	7.46	45.93[a]
Netherlands	6.99	18.93	63.89[a]	12.76	5.82	33.11[a]
Germany	8.45	14.66	62.64[a]	17.79	10.08	35.49[a]
Italy	6.51	35.82[a]	135.85[a]	2.99	14.35	71.33[a]
Ireland	14.00	28.18[b]	40.52[a]	7.58	6.55	18.98
Greece	20.28	23.23[c]	152.22[a]	13.22	24.17[b]	94.52[a]
Portugal	7.13	11.31	101.78[a]	6.69	11.98	64.94[a]
Belgium	11.89	20.73	94.79[a]	8.86	3.93	38.90[a]
Luxembourg	13.77	31.09[a]	160.94[a]	6.63	16.53	119.15[a]
Finland	7.69	22.95[c]	109.16[a]	11.76	9.85	67.61[a]
Austria	11.09	29.99[b]	134.74[a]	12.13	7.04	44.58[a]

Country	Threshold BDM Test					
	Indicator A			Indicator B		
	No det.	$p = 0$	$p = 1$	No det.	$p = 0$	$p = 1$
Spain	2.54	2.68	5.84	17.77[b]	3.99	7.25
France	2.75	4.28	2.74	2.68	5.77	5.17
Netherlands	3.38	8.49	8.43	7.39	6.45	7.35
Germany	6.72	6.64	4.81	8.74	2.05	2.19
Italy	3.75	2.54	3.59	1.99	6.09	7.71
Ireland	12.30	3.42	3.35	2.19	5.70	7.47
Greece	16.61[b]	11.48	20.40[c]	6.88	18.46[b]	25.87[a]
Portugal	3.68	1.68	5.05	3.53	9.98	9.67
Belgium	3.11	3.23	5.49	3.37	3.41	5.47
Luxembourg	7.71	1.49	8.27	4.19	2.76	5.39
Finland	2.79	8.59	9.19	2.31	8.59	8.83
Austria	4.37	1.75	2.71	8.97	2.32	2.65

Notes. [a,b,c] indicate statistically significant results at 1, 5, and 10% levels, respectively.