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Exports and capacity constraints – A smooth transition regression model for six euro area countries

Abstract

Traditional specifications of export equations incorporate foreign demand as a demand pull factor and the real exchange rate as a relative price variable. However, such standard export equations have failed to explain the export performance of euro area countries during the crisis period. In particular, the significant gains in export market shares in a number of vulnerable euro area crisis countries did not coincide with an appropriate improvement in price competitiveness. This paper argues that, under certain conditions, firms consider export activity as a substitute of serving domestic demand. The strength of the link between domestic demand and exports is dependent on capacity constraints. Our econometric model for six euro area countries suggests domestic demand pressure and capacity constraint restrictions as additional variables of a properly specified export equation. As an innovation to the literature, we assess the empirical significance through the logistic and the exponential variant of the nonlinear smooth transition regression model. In the first case, we differentiate between positive and negative changes in capacity utilization and in the second case between small and large changes of the same transition variable. We find that domestic demand developments are relevant for the short-run dynamics of exports when capacity utilization is low. For some countries, we also find evidence that the substitution effect of domestic demand on exports turns out to be stronger the larger is the deviation of capacity utilization from its average value over the cycle.

JEL Codes: F14, C22, C50, C51, F10

Keywords: domestic demand pressure, error-correction models, asymmetry, play-hysteresis; modeling techniques; switching/spline regression; smooth transition models; exports; sunk costs

Preliminary Draft

December 2013

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1. Introduction

A number of euro area countries which recorded large current account deficits in the pre-crisis period have seen a significant correction of their external imbalances over recent years. Although driven to a large extent by falling imports, a significant part of the correction has also resulted from rising exports (see ECB 2013). Interestingly, the standard approach to model exports appears unable to exactly trace for the export performance since 2009. The recent significant and continuous increase of exports market shares cannot be explained by changes in the usual price competitiveness indicators as positive developments such as shrinking unit labor costs and falling real effective exchange rates are able to explain only a part of the gains in export market shares. This suggests that non-price related factors have been important in explaining export performance of euro area countries. The emerging residuals can, however, be potentially matched by the parallel dramatic fall of domestic demand, as shown by Esteves and Rua (2013) for the case of Portugal. In fact, the relationship between domestic demand and exports could be particularly important in the current economic scenario of cyclical weakness. It may have a bearing beyond the Portuguese case and may well extend to other euro area member countries facing significant macroeconomic adjustment needs and thus a strong decline in domestic demand.

While studies on the effects of domestic demand pressure on the inclination and/or capacity to export are not numerous, they have their roots already in the 1960s.[2] Generally, it is argued that increases in export demand cannot be satisfied in the short-run when capacity utilization is high and when production is mainly sold on the domestic market. Conversely, during a domestic recession, firms will be able to shift more resources to export activities. In these periods, firms strive to compensate for the decline in domestic sales through increased efforts to export in order to stay in or enter the export market. The studies overall identified a significant negative effect of domestic demand pressure on exports for several countries, among them the United Kingdom, the United States, Germany, Spain, Israel, Turkey, Morocco and India. Our study goes beyond this country sample by focusing on six euro area countries with significant current account deficits in the pre-crisis period (Spain, Portugal, Italy, France, Ireland and Greece), using an adequate set of nonlinear econometric procedures not applied up to now in this context. Building on hysteresis models of international trade, we explicitly test for a nonlinear relationship between domestic demand and foreign sales in the short-run. The basic idea is that non-exporting firms are more willing to pay sunk costs of export market entry in an environment of weak domestic demand and low capacity utilization, while exporting firms strive to stay in the foreign market and accept lower or even negative profits in order to avoid exit costs and costs of re-entry.

Moreover, we try to put the empirical analysis on a more sound and coherent theoretical footing by explicitly incorporating the trade hysteresis approach. The latter integrates the micro approach (which tends to dominate the preceding studies on the topic) and the macro approach in a unified approach allowing for a so-called "symmetric nonlinearity". By this, we are enlarging the set of testable hypotheses on the impact of domestic demand pressure on the export equation.

The paper proceeds as follows. In section 2, we present different theoretical approaches which help to explain a negative relationship between domestic demand and exports. We consider a simple sunk

[2] See, for instance, Ball et al. (1966), Smyth (1968), Artus (1970, 1973), Zilberfarb (1980), Faini (1994) and Sharma (2003).

cost-based model which serves to capture the nonlinear hysteresis-type dynamics inherent in the relation between the degree of capacity utilization and exports as the most promising one. Taking this model as a starting point, we conduct some pre-testing in terms of unit roots and cointegration in section 3. This enables us to model an error-correction export equation and to incorporate nonlinearities imposed by our theoretical considerations. In section 4, we explicitly refer to two different kinds of nonlinearities: exports might react sharper in a recession than during an economic expansion, or might react to a negligibly low extent to a small change in economic conditions, but the effect may strongly increase for larger changes in the business cycle. Whereas the previous literature investigates only the former hypothesis (see e.g. Berman et al. 2011, Esteves and Rua 2013,), we also focus – as an innovation – on the latter. In doing so, we introduce capacity utilization as the so-called transition variable to capture business cycle effects. We structure our proceedings in accordance with the modeling cycle for the smooth transition regression model suggested by Teräsvirta (1994) which consists of three stages: specification, estimation and evaluation. In the first stage, we perform linearity tests for our basic linear model, and then select between a logistic and an exponential STR model. In the second stage, we estimate the parameter values by multivariate nonlinear least squares, and in a last stage evaluate and test our model. Section 5 finally concludes.

2. Theoretical motivation

The export response to a domestic demand shock is not straightforward. A standard hypothesis in international trade has been that firms face constant marginal costs and maximize profits on the domestic and export markets independently of each other. Das et al. (2007) argue for instance that "shocks that shift the domestic demand schedule do not affect the optimal level of exports". Other theoretical considerations suggest a positive link between domestic and foreign sales, i.e. *complementarity* between the two, at least in the long-run. This may be due to learning by doing effects emerging from domestic sales to export activities and in opposite direction which in turn raises overall efficiency in the long-run (Belke et al. 2013, Esteves and Rua 2013). A positive and complementary correlation may also emerge in the short-run if there is a liquidity constraint and the cash flow generated by exports is used to finance domestic operations (Berman et al. 2011; referred to in the following as the short-run "liquidity channel").

More recently, however, much theoretical and empirical research *at the firm level* has been conducted which allows for a deeper foundation of the relationship between domestic demand and exports (Berman et al. 2011, Blum et al. 2011, Vannoorenberghe 2012). These studies generally argue that, in the short-run, exporting firms substitute sales between their domestic and export markets. Vannoorenberghe (2012) shows theoretically and empirically that a higher than average sales growth in one market is associated with a lower than average growth in the other. Máñez et al. (2008) find that foreign markets became a relevant alternative in periods of low domestic demand, and that the probability of exporting increases in these periods. In turn, Ahn and McQuoid (2013) and Ilmakunnas and Nurmi (2007) conclude that positive domestic demand shocks may exert a downward pressure on exports.

The arguments put forward to motivate a short-run *substitutive* relationship between domestic demand and exports are twofold: A first possible reason is related to the *demand side* of exports. With growing domestic demand, inflationary pressure increases which in turn should diminish price competitiveness of exports and therefore reduce export demand. This effect is usually taken into account by means of the real exchange rate in empirical export demand equations (Esteves and Rua 2013). Alternatively, one could argue that prices are relatively rigid in the short-run, especially in the downward direction. Hence, they may not react adequately to changes in domestic demand pressure (Zilberfarb 1980). In this case, domestic demand would exert an impact on exports (via competitiveness and export demand) only after some time has elapsed and/or if business cycle fluctuations are pronounced.

A second and more direct impact of domestic demand pressure on exports refers to the *supply side* of exports. In their excellent survey, Ahn and McQuoid (2013) deal in detail with the sources of export-domestic sales trade-offs and assess the growing literature that traces back a negative correlation between domestic and export sales to capacity constraints or increasing marginal costs.[3] Using a standard Cobb-Douglas production function, the assumption of increasing marginal costs is motivated by production factors which are difficult (or costly) to adjust in the short-run, as evidenced by lengthy hiring procedures or overtime pay for labor. When a firm experiences a demand increase in one market and thus raises its sales in that market, the firm's marginal costs will increase. Because of the higher marginal costs, it would be optimal to then reduce the sales in the other market and vice versa. With marginal costs increasing in the short run, firms therefore face a trade-off between serving the domestic and foreign market.

The same pattern can be explained using a Melitz (2003) type of model of international trade with demand uncertainty in which firms face market-specific shocks and short-run convex costs of production. In these frameworks, the optimal output for the domestic and the foreign market are not independent of each other. Firms react to a shock in the domestic market by adjusting their sales in the foreign market. Faced with a negative domestic demand shock, firms would sell relatively less to the domestic market and target their sales more towards foreign markets as the costs of excess capacity may outweigh the additional costs and effort of selling in the foreign market. By contrast, firms will prefer selling to the domestic market in detriment of export sales if domestic demand increases.

Overall, the main lesson from the available empirical literature is that any exercise of modeling export performance should take into account not only the factors driving external demand (and thus impact export activity from the demand side), but also those influencing domestic demand (which affect export activity mostly through the supply side). Moreover, the studies underline the necessity of clearly differentiating between the short and the long-run.

One potential limitation of the previous literature is that the "complementarity" versus "substitutability" property of domestic demand and export activity has typically been analyzed in a linear framework. The relationship between domestic demand and export performance may however vary with economic conditions and thus be of a nonlinear nature. This could be due to irreversible costs firms need to pay to enter a foreign market, which are sunk ex post (Baldwin and Krugman 1989). Activity in export markets and building a global network for exports requires considerable set up costs such as market research costs, marketing, finding suitable foreign suppliers and setting up networks for distribution. Most of these costs cannot be reversed on leaving the

[3] Supporting empirical evidence is delivered by Blum et al. (2011) for Chilean, Soderbery (2011) for Thai and Ahn and McQuoid (2013) for Indonesian firms.

export market; quite the contrary, these costs mainly refer to knowledge and information that needs to be gathered to set up a global export network. As soon as the firm leaves the export market, the significance of this knowledge diminishes rapidly (Belke et al. 2013). These sunk costs imply that firms not yet participating in export markets consider export market entry only during certain conditions: as long as domestic demand is strong and capacity is highly utilized, there is no reason and even no capacity to export. With average capacity utilization, capacities exist for serving export markets, but sunk entry costs might deter firms from entering. Only when domestic demand is very low and expected to stay low and there is much capacity for exports, firms consider exports as a substitute for domestic sales. In return, firms already participating in exports markets would tend to exit these markets only when domestic demand becomes very strong and due to capacity constraints not both domestic and foreign markets can be served at the same time.

In a theoretical model, if there is uncertainty about returns, the decision to switch on or off export activity can be analyzed based on the Dixit-type "investment under uncertainty" model (Dixit and Pindyck 1994) or, as a modern variant, based on Impullitti et al. (2013). They derive export market entry and exit decisions in a general equilibrium context with heterogeneous firms andshow that sunk costs induce hysteresis, i.e. history-dependency when it comes to export markets participation. Empirical studies with firm level data, among them Roberts and Tybout (1997), Bernard and Wagner (2001), Bernard and Jensen (2004) and Campa (2004) confirm these findings.

In these micro models of hysteresis in export market participation, a band of inaction due to switching costs for firms between serving the domestic and foreign market typically emerges. The existence of sunk costs thus suggests that if there is substitutability among serving domestic and export demand, it will only be reached if the deviation of capacity utilization from its normal level is either highly positive ("strong") or highly negative ("weak"). We call this pattern "symmetric nonlinearity". It will require a significant negative domestic demand shock for firms to reach a threshold where they pay the sunk entry costs and switch to export activity. In the same vein, in order to avoid paying exit costs and repaying the entry costs, active exporters may only leave the export market if domestic demand pressure increases strongly and capacity constraint considerations become pressing (Belke and Goecke 2005, Esteves and Rua 2013).

Besides this "symmetric nonlinearity", we also take into account the possibility of an "asymmetric nonlinearity". Here, the effect of low capacity utilization on exports has the effect just explained, i.e. sunk costs will discourage export market entry until very weak capacity utilization levels due to very low domestic demand are reached. The opposite case, i.e. very strong capacity utilization leading to export market exits, might, however, not be of importance. Capacity constraints could be less binding in some countries due to flexible prices or flexible labor markets and immigration.

In the context of this paper, we will therefore analyze the relationship between domestic demand and export activity in a nonlinear framework. Based on the depicted micro foundation, we develop an aggregation approach which appears to be adequate to fit a macro data set as used in this contribution. Most importantly, because thresholds for entering or exiting export markets are firm- and sector-specific, we apply a so-called "smooth transition" model that makes specifying an explicit threshold on the macro level unnecessary, but rather allows for a smooth change between regimes. In the following empirical analysis, we will test for the relevance of the hysteresis channel by applying an exponential smooth transition regression model to capture the "symmetric nonlinearity" and a logistic smooth transition regression to model the "asymmetric nonlinearity". The aggregation at the macro level allows us to draw results on net effects of capacity utilization on the economies as a whole.

3. Estimation design and pre-testing

Standard international trade models predict that the volume of exports of a country is in the long-run a function of its foreign demand and its relative price level vis-à-vis its main trading partners. As a first step, we therefore estimate an export equation which relates real exports of goods and services x_t to real foreign demand y_t^* and the real effective exchange rate r_t. We consider the (non-) stationarity of our series and then apply the Engle-Granger cointegration technique to find a long-run relation between exports, foreign demand and the real effective exchange rate.[4] As a second step, we estimate an error-correction model which includes the short-run adjustment to our long-run equilibrium. As explained in section 2, it is rather straightforward from theory that domestic demand d_t may exert an important short-run effect on exports and that the strength and direction of this effect depends on the business cycle stance. Deviating from Esteves and Rua (2013) and the literature cited therein, we do not only take into account the possibility that downturns often have a sharper impact on export activities of a country than recoveries and that this effect is particularly strong for large changes in economic conditions. Instead, we also allow for the possibility that export activity reacts only to a negligibly low extent to a small change in economic conditions (as measured by the degree of capacity utilization), but the effect strongly increases for larger changes in conditions. We therefore apply a nonlinear framework to capture any nonlinear impact regarding the state of the economies. We consider each country's economic conditions by looking at deviations of its capacity utilization from its mean.

Data

Our data stems from different sources: Data on real exports (x_t) and real domestic demand (dd_t) comes from the national statistical offices (either obtained from Eurostat or Oxford Economics). These data are adjusted for price by relying on prices of a reference year. The real effective exchange rate (r_t) is an index deflated by consumer price indices with a country's 15 main trading partners available at Eurostat.[5] The series on foreign demand (y_t^*) is based on trade-weighted imports for 15 main trading partners and comes from the ECB. Finally, the data on capacity utilization in the manufacturing industry (z_t) stems from the Business and Consumer Surveys by the European Commission, available from Eurostat. For France and Ireland, this data was not available (or only for a very short time period). For these countries, we used the output gap instead (interpolated data from AMECO). The series are all available as quarterly data, for most variables in the time period 1980:Q1 to 2012:Q4.[6]

[4] Such a "standard" export demand equation has also been estimated by many others, for instance by the European Commission (2011).

[5] We also used REERs for the main 24 trading partners and deflated by unit labor costs; results did not change with these different measures.

[6] For the exact definitions of variables cf. Table A1.

Unit root tests

As is commonly done, we take each series in (natural) logarithms. In a first step, we check whether the variables in our model are stationary or not, i.e. whether they are integrated of order zero, I(0), or of a higher order, e.g. I(1). For this purpose, we apply the augmented Dickey-Fuller test (ADF-test) with different auxiliary regressions: for the real effective exchange rate series, the regression includes an intercept, but no deterministic time trend; all other series show a time-dependent mean which is then incorporated into the auxiliary regressions via both an intercept and a time trend.

To account for possible structural breaks in the series, we also apply the LM unit root testing procedure based on Lee and Strazicich (2003). If there were structural breaks in the series, the ADF test would have very low power and would be biased towards non-rejection. Thus we apply another test for those times when the null hypothesis of the ADF test cannot be rejected, i.e. to the levels of the series to test for the correctness of the ADF test results.[7]

The results for both the ADF test and the Lee-Strazicich test can be found in Table 1. For the series in levels, we cannot reject the null hypothesis of a unit root for both the ADF test and the Lee-Strazicich test. At the same time, the null hypothesis can be rejected for the series in first differences. Thus, we conclude that the series are all I(1).

– Table 1 about here –

Testing for cointegration

As the variables are non-stationary, we cannot estimate an export equation in a straightforward fashion, but first need to consider cointegration. This will be done by the Engle-Granger approach. The Engle-Granger approach estimates the following long-run equilibrium relationship:

$$x_t = b_1 + b_2 y_t^* + b_3 r_t + e_t \tag{1}$$

with log of exports x_t, log of foreign demand y_t^*, and log of the real effective exchange rate r_t. With time series data for the countries in question, there might be the issue of structural breaks in their long-run relationship, mostly due to the introduction of the euro and the time leading up to it. For this purpose, we allow for a structural break (d) in this relation. The break point for each country is found by a multiple structural change analysis as described in Bai and Perron (2003)[8] and by a Gregory-Hansen cointegration test (Gregory and Hansen 1996a, 1996b) which allows for one break in the cointegration regression. The identified break points all lie in the time period between the European Exchange Rate Mechanism (ERM) crisis of 1992/1993 and the introduction of the euro in 1999. For Spain and France, the break point occurs in 1993, the time of the ERM crisis. For Italy – which left the ERM during its crisis, so can be assumed to have been affected differently than the former mentioned countries – the introduction of the euro in 1999 constitutes the break point. For

[7] The LM test by Lee and Strazicich will be applied to each series with both one break and two breaks (each break representing a shift in levels), where the structural break is allowed to occur at an endogenously set date.

[8] The maximum number of breaks allowed was two, but due to the relatively short time series at hand we concentrate on one break for estimation of the cointegration relation. Otherwise, events such as the global crisis in 2008 would have been considered as another break (which, however, would have included only a short number of time periods after the break).

Ireland and Portugal, the structural breaks were identified in 1995, around the start of convergence to the euro. The break for Greece is 1998 when it joined the ERM.

The dummy d is defined as $d = 1$ if $t \geq break\ point$, otherwise $d = 0$. The dummy and interaction terms with the regressors are included in the equation, so that we have:

$$x_t = b_1 + b_2 y_t^* + b_3 r_t + b_4 d + b_5 d \cdot y_t^* + b_6 d \cdot r_t + e_t \qquad (2)$$

If there was a long-run linear relation between these series, the residuals $\widehat{e_t}$ from this regression had to be stationary. In this case, the OLS results would yield super-consistent estimates for the cointegrating parameters. We estimate equation (2) by fully modified least squares (which corrects the OLS estimator for endogeneity and serial correlation) and compute an Engle-Granger test for cointegration using the residuals $\widehat{e_t}$ from this first-stage regression. The null hypothesis for this test is that there is no cointegration (i.e. that the residual series has a unit root). The test results with the respective critical values from MacKinnon (1991) can be found in Table 2.

– Table 2 about here –

For each country, we find that $\widehat{e_t} \sim I(0)$ and therefore conclude that the variables are cointegrated. The resulting long-run relationship comes from the results of the FMOLS estimation and can be found in Table 3.

– Table 3 about here –

Based on theory, the expected outcome for the long-run relationship is a positive relation between x_t and y_t^*, i.e. when foreign demand increases, so do exports. For x_t and r_t we expect a negative relation, as the REER is a measure of the change in competitiveness of a country. A rise in the index of the respective REER means a loss of competitiveness, i.e. exports should decline. This is exactly what the results show: a positive sign for β_2 and $(\beta_2 + \beta_5)$ and a negative sign for β_3 and $(\beta_3 + \beta_6)$. Also, the size of the coefficients is overall plausible. They are generally not too much different from one for the income elasticity and broadly in line with other studies for the price elasticity (see e.g. European Commission 2011).

Types of nonlinearity

As a next step, we look at short-run adjustments and in particular at the short-run relation between exports and domestic demand, but take into account the long-run equilibrium we have estimated above. For this purpose, we apply an error-correction model. As already mentioned in section 2, in this context we are also taking into account the possibility of nonlinearities. This allows us to investigate a nonlinear adjustment process to a linear long-run equilibrium relationship depending on the state of the economy. A variable might e.g. react sharper in a recession than during an economic expansion, or might hardly react to a small change in economic conditions, but the effect strongly increases for larger changes in conditions. This could be estimated in the context of a simple threshold model. However, for some processes such as an economy's export performance where individual firm level decisions are aggregated, it may not seem reasonable to assume that this threshold is a sudden and abrupt change which is identical for all firms and which is commonly known; the smooth-transition regression (STR) model thus allows for gradual regime change or for a change when the exact timing of the regime switch is not known with certainty. The error-correction model with nonlinear short-run adjustment in STR form then looks like this:

$$\Delta x_t = \left[\alpha_1 + \sum_{i=0}^{n-1} \beta_{1i} \Delta dd_{t-i} + \sum_{i=0}^{n-1} \theta_{1i} \Delta y^*_{t-i} + \sum_{i=0}^{n-1} \mu_{1i} \Delta r_{t-i} + \sum_{i=1}^{n-1} \eta_{1i} \Delta x_{t-i} + \delta_1 \hat{\varepsilon}_{t-1}\right] +$$
$$\left[\alpha_2 + \sum_{i=0}^{n-1} \beta_{2i} \Delta dd_{t-i} + \sum_{i=0}^{n-1} \theta_{2i} \Delta y^*_{t-i} + \sum_{i=0}^{n-1} \mu_{2i} \Delta r_{t-i} + \sum_{i=1}^{n-1} \eta_{2i} \Delta x_{t-i} + \delta_2 \hat{\varepsilon}_{t-1}\right] F\left(z_{t-j}, \gamma, c\right) + u_t \,, \tag{3}$$

$$\hat{\varepsilon}_{t-1} = x_{t-1} - \hat{b}_1 - \hat{b}_2 y^*_{t-1} - \hat{b}_3 r_{t-1} - \hat{b}_4 d - \hat{b}_5 d \cdot y^*_{t-1} - \hat{b}_6 d \cdot r_{t-1} \tag{4}$$

such that the change of x_t is a function of past equilibrium errors (the error-correction term $\delta_1 \hat{\varepsilon}_{t-1}$, where $\hat{\varepsilon}_t$ refers to the error term of the long-run cointegration relation between x_t, y^*_t and r_t determined in the previous step), changes of the variables domestic demand dd_t, foreign demand y^*_t, the real effective exchange rate r_t and past changes of its own value. The parameter δ is referred to as the adjustment effect which gives information about the speed of adjustment when there is disequilibrium and parameters $\alpha, \beta, \theta, \mu, \eta$ are the short-run effects. The parameter β is the parameter we are most interested in, namely the elasticity of exports to a change in domestic demand.

The main difference between our short and long-run specification is the inclusion of the domestic demand variable. Based on the theoretical arguments in section 2 above, domestic demand should enter our estimations in the short-run only.[9] This is a finding also supported e.g. by Esteves and Rua (2013) who argue that it is unclear in which way domestic demand should theoretically enter the long-run export demand equation. Contrary to the long-run estimation, we also do not include a structural break in the short-run specification. Most importantly, our short-run specification already includes nonlinearities by applying the smooth transition regression model. Besides, a break in the long-run relation does not imply that short-run dynamics change as well; by excluding breaks we also reduce the complexity of our model.

The first brackets of the regression model (3) in the is a standard linear error-correction model. The second set of brackets picks up the same regressors, but this part is multiplied with function $F\left(z_{t-j}, \gamma, c\right)$ and constitutes the nonlinear part of the model. F is called the transition function of the smooth transition model. This is a smooth and continuous function which is always bounded and lies between 0 and 1. Here, we consider two different forms of smooth transition models, depending on the specification of the transition function. These are the LSTR model (*logistic* STR model) and ESTR (*exponential* STR model).

The LSTR model relies on a *logistic* transition function of the following form:

$$F(z_{t-j}, \gamma, c) = \left[1 + \exp(-\tfrac{\gamma}{\sigma_z}\left(z_{t-j} - c\right))\right]^{-1} \quad \text{with } \gamma > 0. \tag{5}$$

[9] We also included domestic demand in the long-run cointegration relation, but it did not turn out to be significant.

Here, z is the transition variable, i.e. the variable that distinguishes different regimes in our nonlinear approach. In our case we employ capacity utilization. We employ capacity utilization to capture business cycle effects in particular in the manufacturing industry. We look at deviations of z from a threshold value c which we set as the average value of capacity utilization over the time period in our sample in each country. γ represents the smoothness parameter which determines the speed of transition and σ_z is the standard deviation of the transition variable. As the smoothness parameter γ depends on the scaling of the transition variable, we normalize it by σ_z in order to be scale-free (cf. Teräsvirta (1998)).

The logistic function increases monotonically from 0 to 1 when the value of the transition variable z increases. The threshold thus separates two different regimes:[10] (i) negative deviations of the transition variable from its threshold value: $\lim_{z_{t-j}\to-\infty} F\left(z_{t-j},\gamma,c\right)=0$, i.e. the model collapses to just the linear part, and (ii) positive deviations of the transition variable from its threshold value: $\lim_{z_{t-j}\to+\infty} F\left(z_{t-j},\gamma,c\right)=1$. The coefficients $\alpha,\beta,\theta,\mu,\eta,\delta$ smoothly change between these two extreme values as the value of z_{t-j} changes.

One example of application is the hypothesis that domestic demand is substituted by exports when the degree of capacity utilization is below its normal level (the more strongly it is below its normal level the more significantly the substitution between domestic demand and exports) whereas domestic demand substitutes exports to a lesser extent (or, both even become complements) if, in turn, capacity utilization is higher than normal.

The ESTR model uses an *exponential* transition function of the following functional form:

$$F(z_{t-j},\gamma,c)=1-\exp[-\frac{\gamma}{\sigma_z}\left(z_{t-j}-\mathrm{c}\right)^2] \quad \text{with } \gamma>0. \tag{6}$$

Due to the quadratic term, this transition function is symmetric (U-shaped) around $z_{t-j}=c$ so that the two different regimes to distinguish between are: (i) large deviations of the transition variable from the threshold: $\lim_{z_{t-j}\to\pm\infty} F\left(z_{t-j},\gamma,c\right)=1$ and (ii) small deviations of the transition variable from the threshold: $\lim_{z_{t-j}\to c} F\left(z_{t-j},\gamma,c\right)=0$, i.e. the nonlinear part disappears in the latter extreme.

One example of application is the hypothesis of symmetric hysteresis in exports referred to in section 2. This implies that there are only small effects from domestic demand on exports if the deviation of the transitional variable capacity utilization from c is small ("band of inaction") and large effects if the deviation of the capacity utilization variable is large.

Thus, the two forms of nonlinear error-correction mentioned here refer to different deviations of the transition variable from its threshold value: positive vs. negative deviations in the case of LSTR or large vs. small deviations from equilibrium (but symmetric deviations above or below the threshold) in the case of ESTR.

[10] There are two different ways of interpreting a smooth transition model. On the one hand, it can be regarded as a model with two regimes (the two extremes when the transition function takes on the values 0 or 1) and a smooth transition between these two regimes. On the other hand, it may be interpreted as a continuous number of regimes between the two extremes.

4. The modeling cycle and empirical results

The modeling cycle for the STR model as suggested by Teräsvirta (1994) consists of three stages: specification, estimation and evaluation. In the first stage, we perform linearity tests for the linear model, and then propose either an LSTR or ESTR model. In the second stage, we estimate the parameter values by multivariate nonlinear least squares, and in a last stage evaluate and test our model.

Specification

To test for the presence of an STR model, Teräsvirta (1994) developed the following framework which tests both for the presence of nonlinear behavior and for an LSTR vs. ESTR process. The basis for this test is a Taylor series expansion of the STR model in which the transition function is approximated by a third-order Taylor expansion. The approximated model has the following form:

$$\Delta x_t = \varphi_0 + \varphi_1 W_t + \varphi_2 W_t z_{t-j} + \varphi_3 W_t z_{t-j}{}^2 + \varphi_4 W_t z_{t-j}{}^3 + \epsilon_t \qquad (7)$$

where $W_t = (\Delta dd_t, \Delta dd_{t-1}, \ldots, \Delta dd_{t-p}, \Delta y_t^*, \ldots, \Delta y_{t-p}^*, \Delta r_t, \ldots, \Delta r_{t-p}, \Delta x_{t-1}, \ldots, \Delta x_{t-p}, \hat{\varepsilon}_{t-1})$ and $\varphi_i = (\varphi_{i1}, \ldots, \varphi_{iq})'$ with q equal to the number of regressors (i.e. the number of elements in W_t). To get a first idea of how many regressors and how many lags of each variable to include in W_t, we first estimate the linear part of the VECM model with all different combinations of lags (up to $p = 4$) and choose the number of lags based on the Schwarz information criterion.

Testing for linearity means testing the joint restriction that every nonlinear term in this expression is zero. The alternative hypothesis is that of a STR model. Formally, this is H_{01}: $\varphi_i = 0$ for $i = 2,3,4$ against the alternative H_{11}: $\varphi_i \neq 0$ for at least one of $i = 2,3,4$, implying nonlinearity due to significant higher order terms (Teräsvirta 1998). The test assumes that all regressors and the transition variable are stationary, i.e. OLS is valid. We apply the test for different lag lengths j of the transition variable and select the value of j that results in the smallest p-value, as this is believed to provide the best estimate of j; where the p-values are the same, we also consider the values of $\bar{R}^2$ of the particular regression model. Plausible values for the lag length for quarterly data are here assumed to be $j = 1,2,3,4,5,6$.[11] The results of the test in Table 4 show that the null hypothesis of linearity can be clearly rejected for each country and every lag length.[12] A nonlinear model therefore seems to be suitable for the countries in our sample.

– Table 4 about here –

Based on equation (7), we also approach the choice between an ESTR and an LSTR model (cf. Teräsvirta 1994, 1998). After the first null hypothesis H_{01} has been rejected (i.e. the model is regarded as nonlinear), we test the null hypothesis H_{02}: $\varphi_4 = 0$ against H_{12}: $\varphi_4 \neq 0$. A rejection of this null hypothesis can be seen as a rejection of the ESTR model. Next, we test the hypothesis H_{03}: $\varphi_3 = 0 \mid \varphi_4 = 0$ against H_{13}: $\varphi_3 \neq 0 \mid \varphi_4 = 0$. Not rejecting H_{03} can be seen as evidence in favor of an LSTR model. Lastly, one can test the hypothesis H_{04}: $\varphi_2 = 0 \mid \varphi_3 = \varphi_4 = 0$ against H_{14}: $\varphi_2 \neq 0 \mid \varphi_3 = \varphi_4 = 0$. If H_{04} is rejected, this again points to the LSTR model.

[11] Longer lag lengths (up to j=8) were carried out as robustness checks, but turned out to be less suitable.

[12] However, for France with lag length $j = 1,2,3$ the null hypothesis can only be rejected at the 5% level.

In short, the specification tests point to an LSTR model if H_{02} is rejected and if H_{04} is rejected after H_{03} could not be rejected and to an ESTR model if H_{02} cannot be rejected, or if H_{04} was not rejected after rejecting H_{03}. As Teräsvirta (1994) argues, however, this way, an LSTR model could be erroneously selected and he suggests to compare the relative strengths of the rejections instead, i.e. the p-values. For an LSTR model, H_{02} and H_{04} are usually more strongly rejected than H_{03} and the opposite is expected for an ESTR model. Results for the test are shown in Table 5 including the model tentatively proposed for each country.

– Table 5 about here –

One problem with this test, however, is the fact that in particular in small samples, if the true model is an ESTR model which behaves closely to an LSTR model, the Teräsvirta test often erroneously chooses an LSTR model (cf. Teräsvirta 1994). Because the test does not give clear cut results for the selection of the transition function, we also apply another procedure, proposed by Escribano and Jordá (1999). They claim that using equation (7) above does not capture all important features and suggest a second-order Taylor approximation yielding the following auxiliary regression:

$$\Delta x_t = \varphi_0 + \varphi_1 W_t + \varphi_2 W_t z_{t-j} + \varphi_3 W_t {z_{t-j}}^2 + \varphi_4 W_t {z_{t-j}}^3 + \varphi_5 W_t {z_{t-j}}^4 + \epsilon_t \qquad (8)$$

The hypotheses tested here are $H_{0\mathrm{E}}: \varphi_3 = \varphi_5 = 0$ and $H_{0L}: \varphi_2 = \varphi_4 = 0$. Escribano and Jordá suggest to choose an LSTR model if the lowest p-value is obtained for H_{0L} and an ESTR model if the lowest p-value is obtained for H_{0E}. Results for this test can be found in Table 6.

– Table 6 about here –

In general, it can be argued that once linearity has been rejected, the LSTR and ESTR model form very close substitutes. The decision rules might not be fully important, but can rather be seen as a starting point for estimation. As Teräsvirta (1998) argues, it might make sense to estimate different models and choose between them only during the next stages, i.e. during the estimation and evaluation of the estimation results (the same holds for the choice of the lag length).

Estimation and Evaluation

The second stage of the modeling cycle consists of estimating our parameter values. We estimate equation (3) in combination with either (5) or (6) as the transition function $\mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c})$ with nonlinear least squares (NLS). The results for our main coefficient of interest β are thus made dependent on the state of the economy. The third and last stage of the modeling cycle consists of evaluation. The estimation results are examined by simple judgment concerning the convergence of the models, goodness of fit and by inspecting the regimes the models imply. Our results are also subjected to the misspecification test of no residual autocorrelation. To test for this, we apply a special case of the Breusch-Godfrey Lagrange Multiplier (BG) test suitable for nonlinear estimation (Teräsvirta 1998). The null hypothesis for the test is that there is no p^{th} order serial correlation in our residuals u_t. The test regresses our estimated residuals $\tilde{u}_t$ on lagged residuals $\tilde{u}_{t-1}, \dots, \tilde{u}_{t-p}$ and the partial derivatives of the regression function with respect to γ. Where necessary, we then re-specify our estimated models. Final results for β can be found in Table 7.[13]

A substitution effect between exports and domestic demand should result in a negative coefficient for β. The two extreme regimes in our nonlinear estimation are coefficient β_{10} for $\mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c}) = 0$ (i.e. the linear model) and $\beta_{10}+\beta_{20}$ for the case when $\mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c}) = 1$. To show how β evolves between these two extremes (and thus through all stages of the business cycle), β is drawn in combination with the transition variable $\mathrm{z}_{\mathrm{t-j}}$ in Figures 1 to 6. In these figures, β is defined as $\beta = \beta_{10}+\beta_{20} \cdot \mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c})$.

– Table 7 about here –

– Figures 1 to 6 about here –

Estimation Results

Let us first turn to the countries for which the econometric specification warrants an ESTR model and what we called “symmetric nonlinearity”. As evident from Figure 1, which is based on an ESTR model for Spain, β displays negative values for low and high levels of past capacity utilization. This suggests a substitutive relationship between domestic and foreign sales when the economy is close to peak or trough. When capacity utilization is very low, firms react to a fall in domestic demand by increasing their efforts to export or enter the export market. Similarly, if the economy operates at high capacity utilization, capacity constraints imply that an increase in domestic demand triggers a reallocation of resources from external to domestic clients. The estimation for Spain yields statistically significant results and the economic significance is also meaningful. For very low capacity utilization, a one percentage point fall in domestic demand generates close to a one percentage point increase in exports (cf. Table 7); for peaks, this elasticity is slightly lower. By contrast, a positive link is identified between domestic demand and exports during normal economic conditions. As argued above, this general pattern is in line with the prevalence of hysteresis and the band of inaction due to switching costs for suppliers between serving the domestic and foreign market. It is likely that during this interval, the short-run liquidity channel dominates, whereby the cash flow generated by exports is used to finance domestic operations and the existence of increasing returns dominates the capacity constraints channel (Berman et al. 2011).

Similar results (though somewhat less strong in economic terms) are found for Portugal and Italy as evident in Figures 2 and 3. Whereas the estimated coefficients for domestic demand are statistically significant for Portugal (both the substitution effect during peak and trough and the positive link during normal times), this is not the case for Italy. Here, the small substitution effect during trough and peak is found not to be different from zero contrary to the statistically significant positive coefficient for normal times. Overall, the results indicate that, as a reaction to a negative domestic demand shock, firms which are already in the export market and have thus already incurred market entry costs would sell relatively less to the domestic market and just switch to foreign markets. Especially in the Portuguese case, there appears to be ample scope for relocation in terms of market destination from the home to the foreign market. In 2010, only one third of the firms in the

[13] Complete estimation results are shown in Table A2 along with R^2 values and p-values for the test of no autocorrelation.

Portuguese manufacturing sector was exporting and for them the exports to sales ratio was on average around 30 per cent (Esteves and Rua 2013). During normal economic times, the relationship is strongly complementary for both countries. As former entry costs can be considered to be sunk, one could argue that in order to avoid exiting the markets and paying entry costs anew in the future (Belke and Goecke 2005), firms try to serve both domestic and foreign markets.

The results for France (Figure 4) do not correspond with our theoretical priors but with the results by Berman et al. (2011) who suggest that exports and domestic sales are not substitutive but complementary for a panel of French firms. Our results also show that this complementary relationship is not as strong as it is for other countries; we find an elasticity of around 0.3. Contrary to most other countries' results, the coefficients on domestic demand also do not turn out to be jointly statistically significant. These findings may be related to the lower openness of the French economy and potentially the lower foreign demand elasticity of French exports. Generally, the effect of increases in marginal costs gains importance with foreign demand elasticity, which makes a substitutive relationship between domestic demand and exports more likely in small open economies characterized by highly elastic foreign demand.[14]

Looking at Ireland and Greece, the two countries for which we estimate an LSTR model, we equally find evidence for a negative link between domestic and foreign sales during periods of low capacity utilization (Figures 5 and 6). This effect, however, is statistically insignificant for both countries and economically only of very modest size. After passing a critical threshold, exports and domestic demand become complements with an increasing degree of capacity utilization. The two countries therefore show the "asymmetric nonlinearity" case explained above. In the case of Ireland, the finding that only economic recessions but not periods of booms lead to a substitutive relationship between domestic and export sales may be explained by the higher flexibility of the Irish economy compared to its Southern European counterparts. Flexible prices and immigration may have made capacity constraints less binding. For Greece, the estimated model somewhat resembles a simple two-regime threshold model where marginal changes of capacity utilization around its average have strong effects on the relation between domestic demand and exports. Further strong changes, however, do not have any further effects. Also, at least during the time period under consideration, Greece has never displayed a capacity utilization rate of more than 80 percent and its average degree of utilization is therefore lower compared to the other countries. Whether the relation between exports and domestic demand changes under high capacity utilization rates (such that the asymmetric nonlinearity switches to symmetric linearity) remains unknown.

[14] The coefficient β_{10}, i.e. the coefficient marking the regime when the transition variable is situated around its mean, turns out to be strongly negative for France. This suggests a substitutive relation around average capacity utilization. However, this effect seems to be negligible for three reasons. First, the proxy "output gap" used for capacity utilization does not contain many observations around the average which make statements about this interval rather vague. Second, even though data on capacity utilization for France are available for a short time period only, we perform the same regression with capacity utilization instead of the proxy. The results (not shown here, but available on request) suggest that the coefficient around average utilization is much smaller and not significantly different from zero. Third, relying on the output gap as a proxy or the short time series on capacity utilization, the rejection of linearity is not as strong for France as it is for the other countries; nonlinearity might therefore not be very strong which leads us to assuming a mostly complementary relationship.

Overall, our empirical results strongly suggest that the relationship between exports and domestic sales depends on capacity utilization and the business cycle. A substitutive relationship between domestic and foreign sales is evident during economic downturns when capacities are only weakly utilized; we obtain a negative coefficient for β in all countries except France.[15] This is in line with the gain in export market shares in several euro area crisis countries during the current recession. There is more diversity across countries during other stages of the business cycle suggesting that capacity constraints and the liquidity channel play a different role across countries and/or partly cancel each other out.

Adjustment Effects

Besides the effect of domestic demand on exports, we are also interested in the adjustment coefficient δ, which shows how much of the long-run disequilibrium between exports and its explanatory factors is being corrected in each period. In particular, the coefficient tells us the extent to which disequilibrium prevailing in the previous period has an impact on export adjustments.

If there was a negative shock and exports in the previous period (x_{t-1}) were below its long-run equilibrium path, the value of $\hat{\varepsilon}_{t-1}$ from equation (4) would turn negative. Since we add δ with a positive sign in the error-correction model of equation (3), we expect a negative adjustment coefficient in order for Δx_t to return to the long-run equilibrium. The opposite holds for a positive shock to exports. The speed with which exports return to equilibrium depends on the size of δ. In our specification, the adjustment coefficient depends on the transition function (and therefore the transition variable). To show the adjustment effect for the respective countries over the business cycle (i.e. as the transition variable changes), the trajectory of the estimated coefficient $\delta = \delta_1 + \delta_2 \cdot \mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c})$ is displayed in Figure 7 for the different countries.

– Figure 7 about here –

For Spain and Italy, the adjustment coefficient turns out to be somewhere between -0.1 and -0.3 depending on the state of the economy, i.e. 10 to 30 per cent of the adjustment from disequilibrium takes place in one quarter. The adjustment for France and Greece does not vary substantially over the business cycle and ranges from 30 to about 37 per cent. For Portugal and Ireland, the case is different: the maximum correction of the disequilibrium between exports and its explanatory factors is one fifth (Portugal) and one tenth (Ireland), while no adjustment takes place during strong economic downturns.

Robustness Checks

In the following, we are performing some robustness checks to our estimations by e.g. changing some of the parameters or splitting our sample. As one of these robustness checks, we modify the lag length of the transition variable. In Table 4, we proposed suitable lag lengths based on tests for nonlinearity and chose the one with the lowest p-value (given that it passed the estimation and evaluation stage). Now, we also estimate our nonlinear error-correction model with the second

[15] In case of the ESTR model (for Spain, Portugal, Italy and France) the coefficient of interest for strong economic downturns is $\beta_{10} + \beta_{20}$, for the LSTR model (Ireland and Greece) it is β_{10}.

lowest p-value's lag length. This also reduces the problem of the relatively long lag length for some countries, notably Portugal, Italy and France. As can be seen in Table A3 and in Figure 8, our results quantitatively do not change for most countries if we choose a different lag length of the transition variable. Exceptions are Italy and France, those two countries that did not yield significant results for the estimation of joint β in the original estimation. For France, there is also the problem of very few observations around the threshold value of its transition variable. Most of the observations of the output gap as proxy for capacity utilization are found in the left and right tail, and for these parts the estimation results stay roughly the same.

In addition, we vary the initial values for our smoothness parameter γ when our estimations are iterated by nonlinear least squares. We originally started with a small value of $\gamma=2$, but now also use smaller and larger values. When $\gamma \to \infty$, our model resembles a simple two-regime threshold model. When $\gamma \to 0$, the transition function becomes a constant (0.5 for the logistic version and 0 for the exponential version) and our model collapses into a linear error-correction model. For a large range of starting values for our iteration, the estimations converge to the original results and our original γ. Only for some values very close to zero, the models do not converge anymore. We take this as further evidence for nonlinear estimation.

Last, we estimate our results for a shorter time period to judge how our results could have been influenced by the recent crisis period. We split our sample right before the financial crisis (taking the Lehman Brothers' bankruptcy in September 2008 as break point), i.e. between 2008Q2 and 2008Q3. Because the time span after and including the financial crisis is very short and estimations of this time period would include a very limited number of observations, we concentrate on estimating the first time period, i.e. up to and including 2008Q2. Results can be found in Table A4 and Figure 9.[16] Compared to the original estimation, results are quantitatively robust for Spain, Italy, France, Ireland and Greece. For these countries, inclusion of the crisis time period thus does not seem to strongly impact our findings. Results change, however, for Portugal. Whereas we originally found a small substitution effect between domestic demand and exports when capacity utilization was particularly high or low, this effect now disappears. The recent years thus seem to have influenced Portugal's relation of exports and domestic demand. The most important reason for this might be the fact that very low capacity utilization in our sample only occurred during the last few years, so substitution effects simply might not have been an issue before. Overall, the results from our robustness checks confirm the results we presented above with slight refinements.

5. Conclusions

The results of our macro-econometric smooth transition regression approach indicate that domestic demand behavior is relevant for modeling the short-run dynamics of several euro area member countries' exports. In particular, the estimation results suggest that contemporary and lagged domestic demand developments affect a country's export performance significantly and negatively. In the cases of Spain, Portugal and Italy, the symmetric nonlinearity of the relation expresses itself in a substitutive relationship between domestic demand and export activity if deviations from average capacity utilization are large, independent of their sign. To be more concrete: if our data indicate that

[16] The average value of the transition variable changed as well during the shorter sample and can be found in Figure 9.

the ESTR model has to be applied, the substitution effect from domestic demand on exports turns out to be stronger and more significant the larger the deviation of capacity utilization from its average value over the cycle is. On the contrary, in the cases of Ireland and Greece where the LSTR model turns out to be the better modeling choice, we find that the nonlinear relationship between domestic demand and exports is asymmetric. Domestic demand and exports are substitutes during a business cycle trough and complements in a boom. In other words, positive versus negative deviations of capacity utilization from its normal level matter. For France, we find evidence for mostly complementary relationships instead.

What are the implications of these results for the discussion of macroeconomic adjustment and the reduction of euro area current account imbalances? Prima facie, our results suggest that the negative link between domestic demand and exports is a short-run phenomenon linked to current economic conditions. In the long-run, export performance is closely related to price developments. This would imply that a lot of the gains in export market shares of vulnerable euro area countries are cyclical and could be lost in the long-run. Analyses of cyclically adjusted current account balances, as done in the context of the macroeconomic imbalance procedure or the macroeconomic adjustment programs, could then possibly overestimate the structural adjustment of the current account to the extent that weak domestic economic conditions exert an impact not only on the import side of the net trade equation, but also on the export side.

On the other hand, at least three factors give rise to the hope that the gains in export market performance may be of a more long-run nature. First, if domestic producers have paid sunk costs for export market entry and adapted their production to meet the requirements of foreign clients, attraction by foreign markets should remain high even in an economic upswing. There seems to be no strong reason to leave the export market again as long as variable costs are covered (Belke et al. 2013) and as long as there are capacities for serving both foreign and domestic market. After all, hysteresis refers to history or path dependency; once a certain state has been reached, e.g. participation in export markets, we do not expect it to be reversed anytime soon, at least not as long as a firm is within its “band of inaction”. Second, the effect may also be more long-run to the extent that the current economic crisis leads to a change in investment activities: With an eye on the depressed domestic demand conditions, firms in vulnerable euro area countries may increasingly consider export-oriented foreign direct investment into distribution networks and other hedging activities (Belke et al. 2013). This, in turn, renders the hypothesized negative relationship between domestic demand and exports more long-run. Third, as argued above, a positive correlation between domestic sales and exports might emerge in the long-run due to general efficiency improvements induced by learning-by-doing effects. Overall, it can therefore be expected that a substantial part of the gains in export market shares may indeed be structural. This is supported by the ECB (2013), arguing that policies have lately taken place that are aiming for a rebalancing of the respective economies towards the tradable sector. These policies imply a more structural and sustainable current account adjustment.

Figure 1: Estimation Results for Spain $(c = 0.780)$

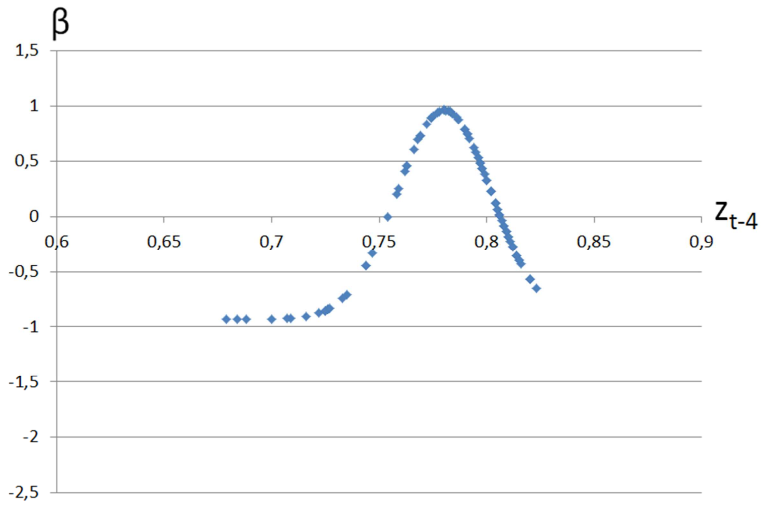


Figure 2: Estimation Results for Portugal $(c = 0.793)$

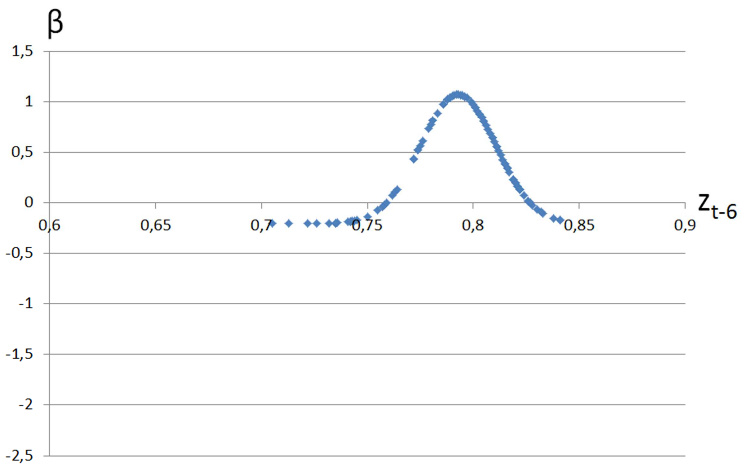


Figure 3: Estimation Results for Italy $(c = 0.751)$

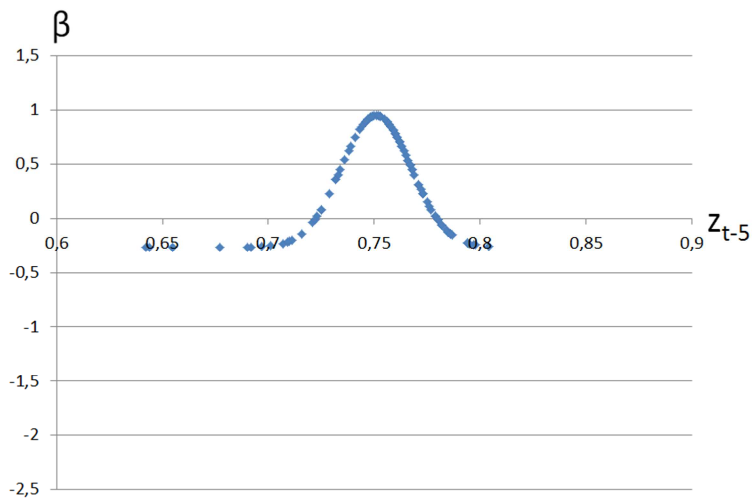


Figure 4: Estimation Results for France $(c = 0.847)$

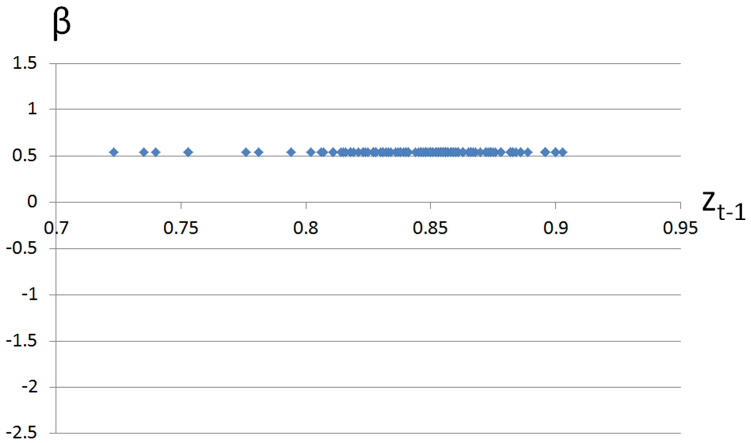


Figure 5: Estimation Results for Ireland $(c = -0.330)$

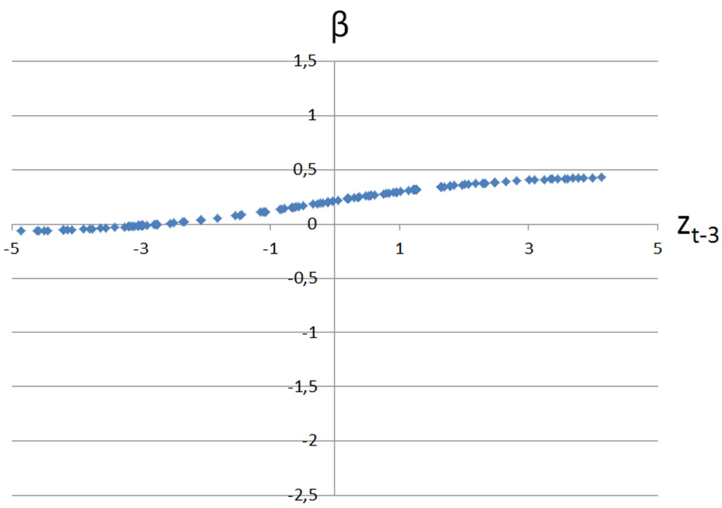


Figure 6: Estimation Results for Greece $(c = 0.748)$

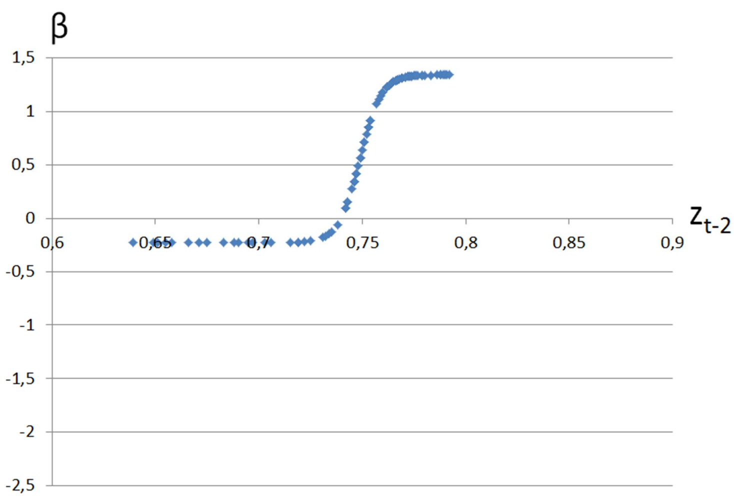
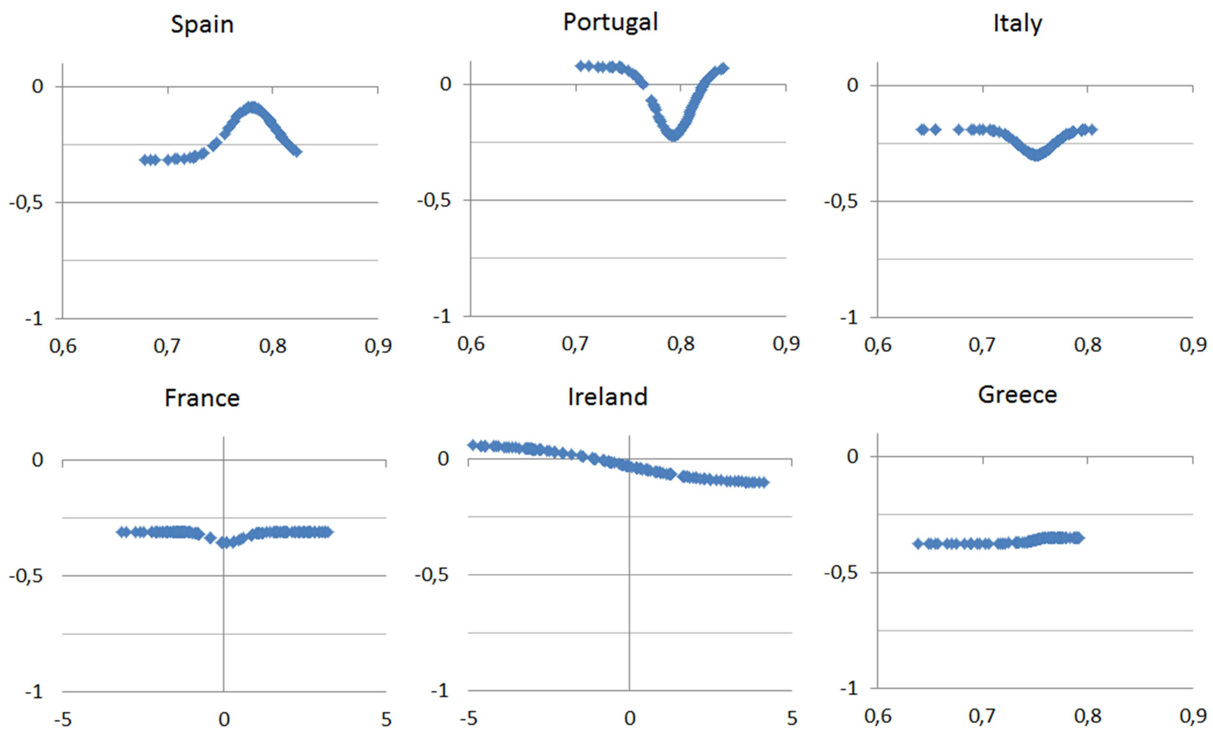
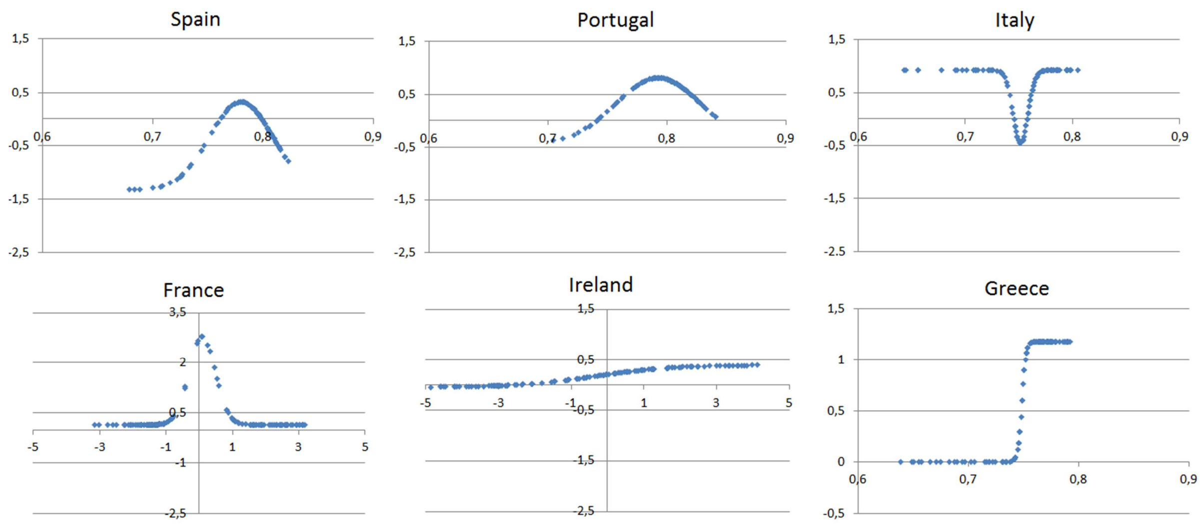


Figure 7: Adjustment coefficients



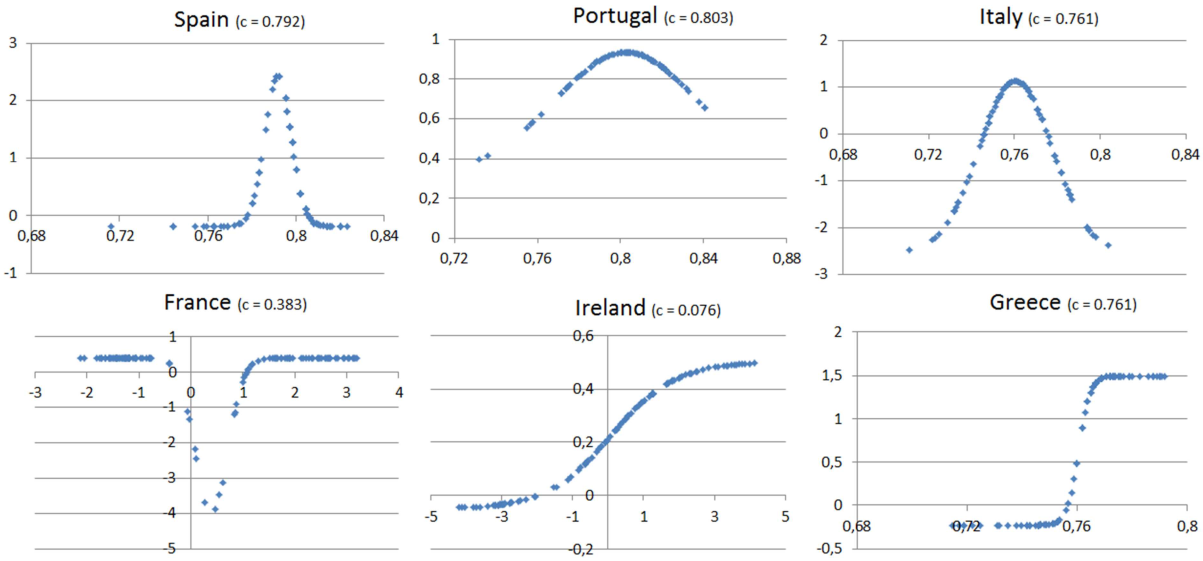
The figures refer to the adjustment coefficient δ which is depicted on the vertical axis; δ is defined as $\delta = \delta_1 + \delta_2 \cdot \mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c})$. The transition variable z_{t-j} is displayed on the horizontal axis.

Figure 8: Robustness: estimation results for different lag lengths



The figures refer to coefficient β which is depicted on the vertical axis; β is defined as $\beta = \beta_{10} + \beta_{20} \cdot \mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c})$. The transition variable z_{t-j} is displayed on the horizontal axis.

Figure 9: Robustness: estimation results for time period 1980Q1 – 2008Q2



The figures refer to coefficient β which is depicted on the vertical axis; β is defined as $\beta = \beta_{10} + \beta_{20} \cdot \mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c})$. The transition variable z_{t-j} is displayed on the horizontal axis.

Figure 10: Robustness: estimation results for export goods only

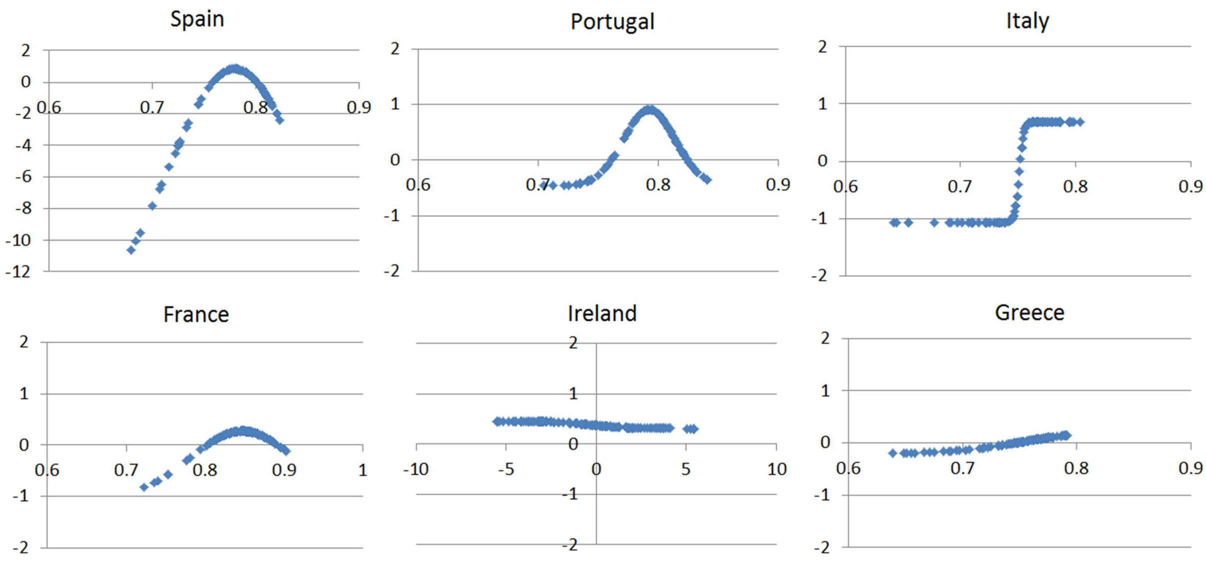


Table 1: Unit Root Tests

		ADF test		*Lee-Strazicich test*	
		Level	*1st Diff.*	*1 break*	*2 breaks*
Country	*Series*	*t-stat. [lags]*	*t-stat. [lags]*	*t-stat.*	*t-stat.*
Spain	dd_t	-1.054 [3]	-2.111** [2]	-0.6281	-0.6370
	x_t	-1.275 [0]	-10.565*** [0]	-1.7927	-2.0560
	y_t^*	-3.418* [1]	-4.569*** [0]	-1.9472	-2.0878
	r_t	-1.250 [1]	-8.763*** [0]	-1.8106	-1.9323
Portugal	dd_t	-0.199 [3]	-3.017*** [2]	-0.5972	-0.6117
	x_t	-0.731 [0]	-7.321*** [0]	-1.4594	-1.5466
	y_t^*	-2.742 [1]	-4.400*** [0]	-1.6444	-1.7162
	r_t	-1.353 [1]	-8.784*** [0]	-2.4693	-2.5850
Italy	dd_t	-0.153 [2]	-3.637*** [1]	-0.7875	-0.8090
	x_t	-1.318 [0]	-5.907*** [1]	-2.0700	-2.3491
	y_t^*	-2.944 [2]	-4.750*** [1]	-2.0089	-2.1816
	r_t	-2.501 [1]	-8.336*** [0]	-1.8317	-1.9321
France	dd_t	-1.692 [2]	-2.659***[1]	-0.9772	-1.0018
	x_t	-1.160 [1]	-4.640*** [1]	-1.0702	-1.1443
	y_t^*	-3.268* [1]	-4.703*** [0]	-2.0007	-2.0854
	r_t	-1.921 [0]	-10.654*** [0]	-2.6688	-2.7981
Ireland	dd_t	-1.650 [3]	-2.805*** [2]	-0.6024	-0.6188
	x_t	-0.764 [4]	-1.401 [6]	-1.1048	-1.1648
	y_t^*	-2.580 [2]	-5.141*** [1]	-1.8182	-1.9890
	r_t	-1.837 [0]	-9.162*** [0]	-1.8346	-1.9568
Greece	dd_t	-0.109 [5]	-2.906*** [4]	-1.1719	-1.2182
	x_t	-1.734 [4]	-5.125*** [3]	-2.4917	-2.8454
	y_t^*	-3.646** [1]	-4.249*** [0]	-1.8027	-1.9790
	r_t	-0.810 [0]	-12.329*** [0]	-3.5230*	-3.8786**

ADF test: lag length is chosen by minimizing the Schwarz Information Criterion with a prior defined maximum lag length of 12. Critical values for an intercept: 1%: -3.43, 5%: -2.86, 10%: -2.57. Critical values for both an intercept and a time trend: 1%: -3.96, 5%: -3.41, 10%: -3.13. Critical values without deterministic trends (for first differences): 1%: -2.56, 5%: -1.94, 10%: -1.62.
Lee-Strazicich test: critical values with one break: 1%: -4.239, 5%: -3.566, 10%: -3.211. Critical values with two breaks: 1%: -4.545, 5%: -3.842, 10%: -3.504. Cf. Lee and Strazicich (2004) and Lee and Strazicich (2003).

Table 2: Engle-Granger Test for Cointegration

Country	*Lags*	*Test Statistic*	*Critical value 1%*	*Critical value 5%*	*Critical value 10%*
Spain	0	-5.88026***	-5.44302	-4.83614	-4.52609
Portugal	2	-4.45270*	-5.13257	-4.52552	-4.21549
Italy	2	-4.63834**	-5.13676	-4.52809	-4.21747
France	3	-5.50043***	-5.44784	-4.83923	-4.52847
Ireland	1	4.67103**	-5.13121	-4.52468	-4.21486
Greece	0	-5.75130***	-5.44302	-4.83614	-4.52609

The (approximate) critical values for the t-test are from MacKinnon (1991) for the respective number of variables.
* statistical significance at the 10% level, ** statistical significance at the 5% level, *** statistical significance at the 1% level.

Table 3: Long-Run Relationship

Country	*Long-run relationship*	*Break-point*	*R^2*
Spain	$x_t = 0.901y_t^* - 0.307r_t + 3.748d + 0.360y_t^*d - 1.055r_td + 7.712$ (20.42) (-2.42) (3.93) (5.53) (-4.20) (16.51)	1993Q4	0.996
Portugal	$x_t = 1.233y_t^* - 0.318r_t + 1.741d - 0.404y_t^*d + 5.306$ (29.72) (-2.02) (8.90) (-8.44) (8.35)	1995Q3	0.988
Italy	$x_t = 0.983y_t^* - 0.961r_t + 11.660d - 2.540r_td + 11.576$ (41.19) (-9.97) (9.32) (-9.39) (26.82)	1999Q1	0.983
France	$x_t = 0.570y_t^* - 0.668r_t + 8.405d - 0.045y_t^*d - 1.716r_td + 11.786$ (31.64) (-3.76) (7.15) (-1.93) (-7.12) (13.63)	1993Q4	0.996
Ireland	$x_t = 1.551y_t^* - 1.654r_t - 6.508d + 1.530r_td + 10.398$ (27.27) (-4.69) (-3.21) (3.47) (6.35)	1995Q1	0.990
Greece	$x_t = 0.493y_t^* + 0.191r_t + 8.646d + 0.433y_t^*d - 2.204r_td + 5.628$ (9.87) (0.93) (3.58) (2.58) (-3.35) (6.57)	1998Q1	0.951

Estimated by FMOLS. T-values in parentheses. The structural break dummy d is defined as $\mathrm{d} = 1$ if t $\geq$ break point, otherwise $\mathrm{d} = 0$.

Table 4: Teräsvirta test for nonlinearity and choice of lag length of transition variable

	test statistic for j=1	test statistic for j=2	test statistic for j=3	test statistic for j=4	test statistic for j=5	test statistic for j=6	*Proposed lag length*
Spain	372.18 (0.000) [0.58]	178.31 (0.000) [0.51]	85.41 (0.000) [0.53]	920.17 (0.000) [0.60]	118.78 (0.000) [0.56]	111.00 (0.000) [0.58]	4
Portugal	34.50 (0.001) [0.34]	33.48 (0.001) [0.38]	108.94 (0.000) [0.37]	121.89 (0.000) [0.33]	251.97 (0.000) [0.41]	1270.97 (0.000) [0.45]	6
Italy	105.25 (0.000) [0.46]	137.53 (0.000) [0.46]	55.13 (0.000) [0.42]	79.38 (0.000) [0.50]	116.32 (0.000) [0.51]	113.27 (0.000) [0.59]	6
France	25.51 (0.044) [0.44]	28.48 (0.019) [0.43]	25.71 (0.041) [0.42]	35.50 (0.002) [0.43]	65.76 (0.000) [0.44]	42.18 (0.000) [0.44]	5 or 6
Ireland	188.90 (0.000) [0.65]	249.53 (0.000) [0.64]	182.05 (0.000) [0.65]	204.51 (0.000) [0.68]	100.73 (0.000) [0.64]	89.36 (0.000) [0.60]	4
Greece	1764.02 (0.000) [0.51]	1619.83 (0.000) [0.58]	146.17 (0.000) [0.49]	97.69 (0.000) [0.49]	137.47 (0.000) [0.51]	180.74 (0.000) [0.47]	2

Test statistic has asymptotic χ^2-distribution with 3m degrees of freedom under the null hypothesis (m = number of regressors). The table shows the values of the test statistic and p-values in parentheses and $\bar{R}^2$ in brackets.
Lag length of the transition variable is chosen based on the lowest p-value and – if p-values are the same – based on the goodness of fit measure $\bar{R}^2$.

Table 5: Teräsvirta test for the appropriate specification

Country	lags	H_{02}	H_{03}	H_{04}	*Proposed specification*
Spain	4	48.32 (0.000)	47.97 (0.000)	43.52 (0.000)	ESTR/LSTR
Portugal	6	47.66 (0.000)	5.89 (0.435)	18.02 (0.006)	LSTR
Italy	6	47.11 (0.000)	28.36 (0.001)	8.29 (0.405)	ESTR/LSTR
France	5	14.80 (0.011)	5.88 (0.318)	8.72 (0.121)	LSTR
France	6	3.17 (0.673)	4.55 (0.474)	14.14 (0.015)	ESTR/LSTR
Ireland	4	50.42 (0.000)	16.70 (0.054)	32.79 (0.000)	LSTR
Greece	2	72.42 (0.000)	54.98 (0.000)	70.47 (0.000)	ESTR/LSTR

χ^2 test statistic realizations are displayed with p-values in parentheses.

Table 6: Escribano Jordá test for the appropriate specification

Country	*lags*	H_{0E}	H_{0L}	*Proposed specification*
Spain	4	37.06 (0.000)	46.80 (0.000)	ESTR/LSTR
Portugal	6	6.56 (0.584)	3.57 (0.827)	ESTR
Italy	6	32.05 (0.000)	19.80 (0.031)	ESTR
France	5	9.29 (0.505)	10.35 (0.410)	LSTR
France	6	8.92 (0.540)	6.93 (0.732)	ESTR
Ireland	4	113.20 (0.000)	96.53 (0.000)	ESTR/LSTR
Greece	2	158.03 (0.000)	15.50 (0.050)	ESTR

LM test statistic with asymptotic χ^2 distribution given with p-values in parentheses. Degrees of freedom: 4(p+1).

Table 7: Estimation results for domestic demand

	Spain	**Portugal**	**Italy**	**France**	**Ireland**	**Greece**
specification	ESTR	ESTR	ESTR	ESTR	LSTR	LSTR
lag length	4	6	5	6	3	2
β_{10}	0.964*** (0.22)	1.072*** (0.13)	0.950** (0.46)	-2.399*** (0.65)	-0.086 (0.23)	-0.226 (0.20)
β_{20}	-1.897*** (0.22)	-1.278*** (0.14)	-1.214*** (0.38)	2.707*** (0.64)	0.538* (0.31)	1.569*** (0.31)
$\beta_{10} + \beta_{20}$	-0.933***	-0.206**	-0.264	0.308	0.452***	1.343***
γ	35.566* (18.61)	49.762*** (19.27)	59.061*** (20.89)	4.095*** (0.64)	1.872** (0.86)	6.662*** (2.29)
R^2	*0.773*	*0.566*	*0.603*	*0.517*	*0.683*	*0.686*
p-value BG test	*0.506*	*0.687*	*0.741*	*0.159*	*0.079*	*0.714*

Coefficients estimated by NLS; Newey-West standard errors in parentheses. * statistical significance at the 10% level, ** statistical significance at the 5% level, *** statistical significance at the 1% level. For the joint significance of the coefficients β_{10} and β_{20}, the linear restriction $\beta_{10} + \beta_{20} = 0$ has been tested with Chi-squared test statistics. β_{j0} $(j = 1,2)$ is the coefficient for domestic demand in the nonlinear error correction model. The two extreme regimes are $\mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c}) = 0$ given by β_{10} (i.e. for the ESTR model around the threshold value, for the LSTR model for large negative deviations from the threshold) and $\mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c}) = 1$ given by β_{10}+β_{20} (i.e. for the ESTR model for large deviations from threshold, for LSTR for large positive deviations from threshold).

Appendix

Table A1: Data Sources

Series	*Source*	*Definition*	*time periods available*
Exports	National Statistical Offices	real exports of goods and services	1980Q1 – 2012Q4; IT: 1981Q1 – 2012Q4
Domestic Demand	National Statistical Offices	real domestic demand	1980Q1 – 2012Q4; IT: 1981Q1 – 2012Q4
Real Effective Exchange Rate	Eurostat	index deflated by consumer price indices with a country's 15 main trading partners	1980Q1 – 2012Q4
Foreign Demand	ECB	trade-weighted imports for 15 main trading partners	1980Q1 – 2012Q4
Capacity Utilization	Eurostat	current level of capacity utilization in manufacturing industry based on business surveys	PT: 1987Q1 – 2012Q4; IT, GR: 1985Q1 – 2012Q4; ES: 1987Q2 – 2012Q4
Output Gap	AMECO (interpolated)	gap between actual GDP and potential GDP as percentage of potential GDP	FR, IE: 1980Q1 – 2012Q4

Table A2: Estimation results

	Spain	**Portugal**	**Italy**	**France**	**Ireland**	**Greece**
specification	ESTR	ESTR	ESTR	ESTR	LSTR	LSTR
lag length	4	6	5	6	3	2
α_1	0.0307*** (0.01)	0.007*** (0.00)	0.002 (0.00)	0.055*** (0.01)	0.012*** (0.00)	0.011 (0.01)
β_{10}	0.964*** (0.22)	1.072*** (0.13)	0.950** (0.46)	-2.399*** (0.65)	-0.086 (0.23)	-0.226 (0.20)
β_{11}		0.617*** (0.15)	1.791*** (0.61)		-0.174 (0.29)	0.454*** (0.17)
β_{12}			0.110 (0.49)			0.341 (0.22)
β_{13}			-1.526*** (0.56)			
θ_{10}	0.403** (0.18)	0.336*** (0.11)	0.598*** (0.17)	0.629*** (0.12)	-0.247** (0.14)	0.593 (0.42)
θ_{11}		-0.843*** (0.16)				
μ_{10}	0.020 (0.11)	0.219** (0.10)	-0.232** (0.09)	0.386** (0.19)	-0.468*** (0.15)	-0.111 (0.22)

μ_{11}	-0.686*** (0.17)					
μ_{12}	-0.417*** (0.13)					
μ_{13}	0.265** (0.11)					
μ_{14}	-0.446*** (0.13)					
η_{11}	-0.070 (0.11)	0.225*** (0.05)	-0.364*** (0.06)	-0.993*** (0.11)	0.141*** (0.05	-0.205 (0.15)
η_{12}	-0.205*** (0.06)				-0.325*** (0.03)	-0.027 (0.08)
η_{13}					0.134** (0.06)	-0.089** (0.04)
η_{14}					0.720*** (0.09)	0.403*** (0.06)
δ_1	-0.090*** (0.03)	-0.222** (0.09)	-0.300*** (0.04)	-0.358*** (0.08)	0.065*** (0.03)	-0.374*** (0.07)
α_2	0.039 (0.03)	-0.013*** (0.00)	-0.007* (0.00)	-0.055*** (0.01)	-0.005 (0.00)	-0.009 (0.01)
β_{20}	-1.897*** (0.22)	-1.278*** (0.14)	-1.214*** (0.38)	2.707*** (0.64)	0.538* (0.31)	1.569*** (0.31)
β_{21}		-1.336*** (0.18)	-1.806* (1.00)		0.827* (0.45)	-0.743*** (0.27)
β_{22}			0.758** (0.33)			-0.390* (0.23)
β_{23}			0.907 (0.64)			
θ_{20}	1.013*** (0.38)	1.027*** (0.23)	0.301 (0.25)	-0.290 (0.23)	0.776** (0.33)	-0.200 (0.40)
θ_{21}		1.026*** (0.18)				
μ_{20}	-0.480* (0.27)	-0.326 (0.23)	-0.391 (0.25)	-0.673*** (0.17)	0.232 (0.20)	-1.807*** (0.45)
μ_{21}	0.843*** (0.25)					
μ_{22}	-0.553*** (0.18)					
μ_{23}	-1.217*** (0.22)					
μ_{24}	0.221*** (0.14)					
η_{21}	0.035 (0.12)	-0.460*** (0.16)	0.638*** (0.10)	1.231*** (0.20)	-0.242** (0.10)	0.318* (0.17)
η_{22}	-0.079 (0.09)				0.215*** (0.07)	0.156** (0.07)
η_{23}					-0.279*** (0.10)	0.292*** (0.05)
η_{24}					-0.424*** (0.09)	-0.170 (0.19)
δ_2	-0.224 (0.16)	0.298*** (0.06)	0.110 (0.08)	0.047 (0.12)	-0.175*** (0.05)	0.025 (0.13)

γ	35.566* (18.61)	49.762*** (19.27)	59.061*** (20.89)	4.095*** (0.64)	1.872** (0.86)	6.662*** (2.29)
R^2	*0.773*	*0.566*	*0.603*	*0.517*	*0.683*	*0.686*
p-value BG test	*0.506*	*0.687*	*0.741*	*0.159*	*0.079*	*0.714*

Coefficients estimated by NLS; Newey-West standard errors in parentheses. * statistical significance at the 10% level, ** statistical significance at the 5% level, *** statistical significance at the 1% level. The Breusch-Godfrey Lagrange Multiplier (BG) test is based on the null hypothesis of no serial correlation of the residuals of order p. Due to quarterly data, we report the results for this test for $p = 4$.

Table A3: Robustness: estimation results for different lag lengths

	Spain	**Portugal**	**Italy**	**France**	**Ireland**	**Greece**
specification	ESTR	ESTR	ESTR	ESTR	LSTR	LSTR
lag length	1	5	4	4	2	1
β_{10}	0.318*** (0.07)	0.810*** (0.24)	-0.458 (0.74)	2.792*** (0.26)	-0.046 (0.19)	-0.001 (0.15)
β_{20}	-1.652*** (0.24)	-1.253*** (0.32)	1.383* (0.83)	-2.642*** (0.22)	0.445* (0.23)	1.177*** (0.31)
$\beta_{10} + \beta_{20}$	-1.334***	-0.444	0.925***	0.159	0.399***	1.179***
γ	20.863*** (3.01)	12.286 (13.14)	342.207*** (91.19)	5.777** (2.43)	2.522 (1.10)	19.505 (13.53)
R^2	0.809	0.488	0.645	0.548	0.682	0.653
p-value BG test	0.408	0.593	0.132	0.053	0.043	0.944

Coefficients estimated by NLS; Newey-West standard errors in parentheses. * statistical significance at the 10% level, ** statistical significance at the 5% level, *** statistical significance at the 1% level. For the joint significance of the coefficients β_{10} and β_{20}, the linear restriction $\beta_{10} + \beta_{20} = 0$ has been tested with Chi-squared test statistics. The Breusch-Godfrey Lagrange Multiplier (BG) test is based on the null hypothesis of no serial correlation of the residuals of order p. Due to quarterly data, we report the results for this test for $p = 4$.

Table A4: Robustness: estimation results for time period 1980Q1 – 2008Q2

	Spain	Portugal	Italy	France	Ireland	Greece
specification	ESTR	ESTR	ESTR	ESTR	LSTR	LSTR
lag length	4	6	5	6	3	2
β_{10}	2.418*** (0.16)	0. .938*** (0.10)	1.143*** (0.41)	-3.980*** (1.29)	-0.050 (0.29)	-0.227*** (0.08)
β_{20}	-2.607*** (0.17)	-0. 610*** (0.15)	-3.680*** (0.72)	4.365*** (1.36)	0.554 (0.36)	1.730*** (0.29)
$\beta_{10} + \beta_{20}$	-0.190**	0.328***	-2.537***	0.385***	0.504***	1.502***
γ	299.112*** (34.11)	9.705** (4.79)	33.175*** (9.62)	8.683*** (1.70)	2.883** (1.25)	8.466*** (1.19)
R^2	0.823	0.430	0.524	0.405	0.720	0.753
p-value BG test	0.790	0.916	0.961	0.232	0.041	0.867

Coefficients estimated by NLS; Newey-West standard errors in parentheses. * statistical significance at the 10% level, ** statistical significance at the 5% level, *** statistical significance at the 1% level. For the joint significance of the coefficients β_{10} and β_{20}, the linear restriction $\beta_{10} + \beta_{20} = 0$ has been tested with Chi-squared test statistics. The Breusch-Godfrey Lagrange Multiplier (BG) test is based on the null hypothesis of no serial correlation of the residuals of order p. Due to quarterly data, we report the results for this test for $p = 4$.

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