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A SMOOTH TRANSITION REGRESSION MODEL FOR SIX EURO AREA COUNTRIES

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Abstract

This paper argues that, under certain conditions, firms consider export activity as a substitute of serving domestic demand. 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 non-linear smooth transition regression model. We find that domestic demand developments are relevant for the short-run dynamics of exports in particular during more extreme stages of the business cycle. A strong substitutive relationship between domestic and foreign sales can most clearly be found for Spain, Portugal and Italy providing evidence of the importance of sunk costs and hysteresis in international trade.

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

Keywords: domestic demand, exports, smooth transition models, hysteresis, sunk costs

Non-technical summary

Several euro area countries – we consider Spain, Portugal, Italy, France, Ireland and Greece – exhibited large current account deficits in the pre-crisis period. Over recent years, these countries have experienced a significant correction of their external imbalances, in particular in the trade balance. Although driven to a large extent by falling imports, a significant part of the correction has resulted from rising exports. Traditional export specifications, which explain export performance by foreign demand and price competitiveness indicators, are not able to describe the recent developments in the countries under consideration. Shrinking unit labour costs, falling real effective exchange rates and thus an increase in price competitiveness explain only a part of the gains in export market shares.

In our paper, we suggest the fall of domestic demand and low capacity utilisation as an additional explanation of increasing exports, resulting in a substitutive relationship between domestic demand and exports. When production capacities are highly utilised amid strong domestic demand, an increase in exports is often not possible in the short-term. Conversely, during a domestic recession, capacities will become free and firms will be able to export more; to compensate for low domestic sales, they might increase their efforts towards the export market. This consideration is of particular relevance during the current weak economic situation in the countries under consideration.

We argue that shifting sales towards the export market or even entering the export market altogether comes with costs. Firms are only willing to pay these costs once their capacities are so lowly utilised that paying these costs is more effective than producing at very low capacities. The opposite might also be true for periods of very high capacity utilisation when firms are reluctant to shift their sales back to the domestic market. This generates a so called 'band of inaction' during periods of relatively normal capacity utilisation when a substitution effect between domestic and foreign sales does not take place. Therefore, these underlying considerations give rise to a possibly non-linear relationship between exports and domestic demand: the substitution effect may arise only during periods of strong economic stress and boom. We also consider the possibility that the export response to domestic demand developments is sharper in a recession than during an economic expansion.

The non-linear relationship is manifested at the firm level: only after passing a certain firm-specific threshold of high or low capacity utilisation will firms leave the 'band of inaction' and shift their sales from one market to the other. As we are interested in economy-wide outcomes and rely on country-level data, we employ a non-linear estimation technique that permits specifying our model even without knowing each firm's exact thresholds. The so called 'smooth transition regression model' does not specify one abrupt switch between being above or below threshold values, but rather allows for a smooth and continuous change from a substitutive to an non-substitutive relationship between exports and domestic demand.

Our empirical analysis builds on quarterly data for six euro area countries during the time period 1980-2012. The results support our hypotheses most clearly for Spain, Portugal and Italy. For these countries, we find a substitutive relationship between domestic demand and exports for low and high levels of capacity utilisation, with particular strong effects during periods of economic stress. For

France, exports seem to be hardly influenced by domestic demand developments. This may be related to the lower openness of the French economy, but also due to lower fluctuations in the business cycle in the time frame under consideration. For Ireland and Greece, we find weak evidence for a substitutive relationship between domestic sales and exports during periods of low capacity utilisation. The weak domestic demand-export relation in Ireland can 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. By contrast, results for Greece are consistent with the lacklustre export performance of the Greek economy during the crisis which can be explained by the lack of adjustment capacity, institutional weaknesses and structural rigidities in labour and product markets.

Our findings provide one possible explanation for the rising exports in the countries under consideration over recent years: in the current situation of cyclical weakness with low rates of capacity utilisation and a strong decline in domestic demand, many firms have tried to compensate for weak domestic sales by increasing their effort of selling on foreign markets or even entering the export market in the first place. Our results suggest that the negative relationship between domestic demand and exports is a short-run phenomenon linked to current economic conditions; in the long-run, a lot of the gains in export market shares of vulnerable euro area countries could be lost. 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. However, several reasons could make the gains in export market performance last in the long-run: domestic firms have paid sunk costs for shifting sales to the export market and a reversal could be unlikely within the 'band of inaction', once domestic demand returns to more normal levels; production has been adjusted for foreign consumers and investment has been shifted to export-oriented projects; lastly, an overall efficiency improvement could lead to higher exports even when domestic sales and exports behave as complements during average levels of domestic demand and capacity utilisation. It can therefore be expected that a substantial part of the gains in export market shares may indeed be structural.

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, in particular the trade balance, over recent years. Although driven to a large extent by falling imports, a significant part of the correction has resulted from rising exports (see ECB 2013). The standard approach to model exports appears unable to exactly trace for the significant and continuous increase of exports market shares in many of these countries since 2009. Shrinking unit labour 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. Specifically, the emerging residuals can be potentially matched by the parallel dramatic fall of domestic demand In fact, as shown in this paper, the relationship between domestic demand and exports could be particularly important in the current economic scenario of significant macroeconomic adjustment needs and 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 utilisation is high and when production is sold mainly 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. Recently, Esteves and Rua (2013) present a survey of the literature covering the theoretical reasons and the empirical evidence concerning the relation between domestic demand and exports, while addressing the Portuguese case. Our study goes beyond the above 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 non-linear econometric procedures not applied in this context up to now.

Building on hysteresis models of international trade, we explicitly test for a non-linear relationship between domestic demand and foreign sales in the short-run. A particular asymmetric effect was already considered in Esteves and Rua (2013). In this paper, as an innovation to the literature, we implement a smooth transition regression model which allows us to specify aggregated non-linearities with a high degree of flexibility. The basic idea is that the strength of the relation between domestic demand and exports depends on the stage of the business cycle. Hysteresis considerations suggest that firms substitute between domestic and foreign sales only during extreme stages of the business cycle. Sunk costs prevent a sharp export reaction to domestic demand developments during average business cycle periods, measured by capacity utilisation. The substitution effect between domestic sales and exports may arise during periods of economic stress, when it is worth for firms to pay sunk costs of entering or shifting sales to the export market, or during economic booms, when capacity constraints provide a lower incentive to export. The export response to domestic demand

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

developments might even be sharper in a recession than during an economic expansion. In an environment of weak domestic demand and low capacity utilisation, exporting firms increase their efforts to shift sales from domestic to export markets or 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, non-exporting firms might be more willing to pay sunk costs of export market entry given the reduced prospects for domestic sales.

The paper proceeds as follows. In section 2, we present different theoretical approaches which help to explain the relationship between domestic demand and exports. We consider a simple sunk cost-based model which serves to capture the non-linear hysteresis-type dynamics inherent in the relation between capacity utilisation 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 non-linearities imposed by our theoretical considerations. In section 4, we perform the smooth transition regression model (STR) suggested by Teräsvirta (1994). We also present several robustness tests. Section 5 concludes.

2. Theoretical motivation

The export response to a domestic demand shock is not straightforward. A recent survey of literature is presented in Esteves and Rua (2013). A positive link between domestic and foreign sales may be due to learning by doing effects which raise overall efficiency (Belke et al. 2013) or due to a "liquidity channel" if there is a liquidity constraint and the cash flow generated by exports is used to finance domestic operations (Berman et al. 2011). By contrast, recent theoretical and empirical research *at the firm level* argues that, in the short-run, exporting firms substitute sales between their domestic and export markets due to capacity constraints or increasing marginal costs (see, e.g., Ilmakunnas and Nurmi 2007, Máñez et al. 2008, Berman et al. 2011, Blum et al. 2011, Vannoorenberghe 2012 or Ahn and McQuoid 2013).

The main lesson from the literature is that any exercise of modelling 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 analysed in a linear framework. The relationship between domestic demand and export performance may however vary with economic conditions and thus be of a non-linear nature. This could be due to irreversible costs firms need to pay to enter a foreign market or to shift more sales towards that 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 export market, quite in contrast, as soon as the firm leaves the export market, the significance of this knowledge diminishes rapidly (Belke et al. 2013).

Substitutability implies that firms try to shift more sales to the export market following a negative domestic demand shock. With sunk costs for entering or shifting to the export market, this might not be considered worthy as long as capacity is relatively highly utilised. The investment might pay off, however, once capacity utilisation falls below a certain threshold. At this point, the cost of running excess capacity may outweigh the additional costs and effort of selling in the foreign market. Shifting sales to foreign markets and increasing overall exports could then be considered as "survival-driven" rather than primarily being due to an increase in competitiveness.[3] The opposite could also hold for a positive domestic demand shock. Firms might prefer selling to the domestic market instead of exporting if highly utilised capacities do not allow them to satisfy both markets. The sunk costs of shifting sales between the markets or risking paying entry market costs again in the future means that only once a certain threshold of high capacity utilisation has been reached, firms would consider this shift. All in all, this suggests that firms consider shifting their sales activities following a domestic demand shock only once certain thresholds of low and high capacity utilisation have been reached; within these thresholds, firms operate in a "band of inaction" in which capacities exist for serving export markets, but sunk entry costs deter firms from entering or shifting their sales (Belke et al. 2013).[4] In a theoretical model, these arguments can be analysed based on the Dixit-type "investment under uncertainty" model (Dixit and Pindyck 1994) or, as a modern variant, based on Impullitti et al. (2013). 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 this context, Esteves and Rua (2013) using macroeconomic data tested for an asymmetric effect of domestic demand on exports. We argue that the strength of the relation between domestic demand and exports depends on capacity constraints and the business cycle in general in a more flexible formulation.

The existence of sunk costs thus suggests that if there is substitutability between serving domestic and export demand, it will only reflect a short-term phenomenon during certain stages of the business cycle, i.e. if the deviation of capacity utilisation from its normal level is either highly positive (upper threshold) or highly negative (lower threshold). In the context of this paper, we therefore analyse the relationship between domestic demand and export activity in a non-linear framework. Based on the micro foundation of sunk cost induced hysteresis in export market participation, we rely on an aggregation approach which appears to be adequate to fit a macro data set as used in this contribution. Most importantly, because thresholds for shifting activities or entering 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. The aggregation at the macro level allows us to draw results on net effects of capacity utilisation on the economies as a whole.

[3] Likewise, the decision to shift activity to the export market could be driven by technical limitations. Firms such as refineries or steel producers might only be able to produce at a certain capacity utilisation rate or otherwise have to shut down their production completely once a certain threshold has been reached. For these firms, it might pay off to shift production to the export market instead of not producing at all.

[4] There appears to be ample scope for relocation in terms of market destination from the home to the foreign market in the countries under consideration. In 2010, for instance, only one third of the firms in the Portuguese manufacturing sector was exporting and for them the exports to sales ratio was on average around 30 per cent (Esteves and Rua 2013).

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 by finding a long-run cointegration relation between these variables.[5] Our analysis focuses, however, on the second step, in which we estimate an error-correction model including the short-run adjustment to our long-run equilibrium. As explained in Section 2, we apply a non-linear framework to capture any non-linear impact regarding the state of the economies. We consider each country's economic conditions by looking at deviations of its capacity utilisation from its mean.

Data

Our data stems from different sources (cf. table A1): Data on real exports (x_t, x_t^{goods}) (both goods & services or goods only) and real domestic demand (dd_t) comes from the national statistical offices (either obtained from Eurostat or Oxford Economics). Value added exports (x_t^{va}) have been constructed by data from the World Input-Output Database (wiod.org); the annual data were converted to quarterly data by applying cubic spline interpolation. The real effective exchange rate is either an index deflated by consumer price indices with a country's 15 main trading partners available at Eurostat (r_t) or an index deflated by unit labour costs with a country's 24 main trading partners also available at Eurostat (r_t^{ULC}). The series on foreign demand (y_t^*) is based on trade-weighted imports for 15 main trading partners and comes from the ECB. Finally, data on capacity utilisation in the manufacturing industry (z_t) stems from the Business and Consumer Surveys by the European Commission, available from Eurostat. For France, this data comes from Insee. In the case of Ireland, data on capacity utilisation is not available. For this country, we use 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.

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. For this purpose, we apply the augmented Dickey-Fuller test (ADF-test) with an intercept in the auxiliary regressions for the real effective exchange rate series and both an intercept and a time trend for all other series. To account for possible structural breaks in the series, we also apply the LM unit root testing procedure based on Lee and Strazicich (2003) to the levels of the series to test for the correctness of the ADF test results.[6] The results for both the ADF test and the Lee-Strazicich test can be found in table 1 and let us conclude that the series are all I(1).

– Table 1 about here –

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

[6] The LM test by Lee and Strazicich is 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.

Testing for cointegration

As the variables are non-stationary, we use the Engle-Granger approach to consider cointegration and estimate the following long-run equilibrium relationship:

$$x_t = b_1 + b_2 y_t^* + b_3 r_t + b_4 d + b_5 d \cdot y_t^* + b_6 d \cdot r_t + e_t \qquad (1)$$

with log of exports x_t, log of foreign demand y_t^*, log of the real effective exchange rate r_t and a dummy d and the respective interaction terms to capture a structural break. With time series data for the countries in question, the introduction of the euro and the time leading up to it might cause a break in the long run relationship. The break point for each country is found by a multiple structural change analysis as described in Bai and Perron (2003)[7] 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. It is defined as $d = 1$ if $t \geq break\ point$, otherwise $d = 0$.

We estimate equation (1) by fully modified least squares and compute an Engle-Granger test for cointegration ; test results with the respective critical values from MacKinnon (1991) can be found in table 2. 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 from the FMOLS estimation can be found in table 3.

– Table 2 about here –

– Table 3 about here –

The results need to be interpreted with caution, as further structural breaks or omitted variables can have an influence on the outcomes (cf. e.g. Esteves and Rua 2013).[8] Our results are broadly in line with other studies, both in terms of sign and size of the coefficients (cf. e.g. European Commission 2011). We refrain however from a more detailed analysis, given that this paper's focus is on the short-run relation.

Types of non-linearity

As a next step, we look at short-run adjustments and in particular at the short-run relation between exports and domestic demand, taking into account the long-run equilibrium estimated above. For this purpose, we apply an error-correction model. As already mentioned in section 2, we take into account the possibility of a non-linear adjustment process to a linear long-run equilibrium relationship depending on the state of the economy. For an economy's export performance where

[7] 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.

[8] As robustness checks, we also included further variables such as trade openness in the long-run relation or restricted the coefficient for foreign demand to unity. Certainly, other non-price competitiveness variables could have an influence on exports as well. Since our focus is on the short-run results and slightly different long-run specifications did not change the final short-run non-linear estimation results in a noteworthy way, we do not report these results here.

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 on an aggregated level. The error-correction model with non-linear short-run adjustment in STR form then looks like this:

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

$$\hat{\varepsilon}_{t-1} = x_{t-1} - \hat{b}_1 - \hat{b}_2 y_{t-1}^{*} - \hat{b}_3 r_{t-1} - \hat{b}_4 d - \hat{b}_5 d \cdot y_{t-1}^{*} - \hat{b}_6 d \cdot r_{t-1} \tag{3}$$

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] Contrary to the long-run estimation, we do not include a structural break in the short-run specification because our short-run specification already includes non-linearities by applying the smooth transition regression model. Furthermore, 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 set of brackets of the regression model (3) 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 non-linear part of the model. F is called the transition function of the smooth transition model. This is a smooth and continuous function which is 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:

$$F(z_{t-j}, \gamma, c) = \left[1 + \exp(-\tfrac{\gamma}{\sigma_z}\left(z_{t-j} - c\right))\right]^{-1} \quad \text{with } \gamma > 0. \tag{4}$$

[9] We also included domestic demand in the long-run cointegration relation, but it neither turned out to be statistically significant nor did it help to constitute a better long-run relation.

Here, z is the transition variable, i.e. the variable that distinguishes different regimes in our non-linear approach. In our case z is operationalised by the degree of capacity utilisation 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 utilisation over our sample time period. γ represents the smoothness parameter which determines the speed and strength of the 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 in the extreme and a smooth transition between these two: (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.

In our setting, this implies testing the hypothesis that domestic sales are substituted by foreign sales once capacity utilisation falls below a certain threshold. Further reductions in capacity reduction strengthen the substitution of domestic demand by exports. Note that there is no threshold for the opposite case of high capacity utilisation. In other words, the band of inaction is only constrained on one side (for negative but not for positive deviations of capacity utilisation from its average values).

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

$$F(z_{t-j},\gamma,c)=1-\exp[-\frac{\gamma}{\sigma_z}\left(z_{t-j}-\mathrm{c}\right)^2] \quad \text{with } \gamma>0. \tag{5}$$

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 non-linear part disappears in the latter extreme.

In our case, this refers to the hypothesis of symmetric hysteresis in exports, i.e. both positive and negative deviations of capacity utilisation from its average value c matter. This implies that as long as the deviation of the transitional variable capacity utilisation from c is small, there would be no or only small substitution effects from domestic demand to exports (band of inaction). However, if the capacity utilisation variable is either significantly above or below its average value, we would expect substitution effects.

Thus, the two forms of non-linear 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.

4. The modelling cycle and empirical results

The modelling 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 non-linear 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 non-linear behaviour 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 of the following form:

$$\Delta x_t = \varphi_0 + \varphi_1 W_t + \varphi_2 W_t z_{t-j} + \varphi_3 W_t {z_{t-j}}^2 + \varphi_4 W_t {z_{t-j}}^3 + \epsilon_t \qquad (6)$$

where $W_t = (\Delta dd_t, \Delta dd_{t-1}, \dots, \Delta dd_{t-p}, \Delta y_t^*, \dots, \Delta y_{t-p}^*, \Delta r_t, \dots, \Delta r_{t-p}, \Delta x_{t-1}, \dots, \Delta x_{t-p}, \hat{\varepsilon}_{t-1})$ and $\varphi_i = (\varphi_{i1}, \dots, \varphi_{iq})'$ 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 non-linear 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 non-linearity 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; when 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 $j = 1, \dots, 6$.[10] 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.[11] A non-linear model therefore seems to be suitable for the countries in our sample.

– Table 4 about here –

Based on equation (6), we also approach the choice between an ESTR and an LSTR model (cf. Teräsvirta 1994, 1998). After the null hypothesis H_{01} has been rejected (i.e. the model is regarded as non-linear), 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 favour 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.

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

[11] France is an exception; here, null hypothesis cannot be rejected for higher lag lengths.

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 the Teräsvirta test in particular in small samples is that if the true model is an ESTR model which behaves closely to an LSTR model, the 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 argue that using equation (6) does not capture all important features and suggest a second-order Taylor approximation yielding the following auxiliary regression:

$$\Delta x_t = \varphi_0 + \varphi_1 W_t + \varphi_2 W_t z_{t-j} + \varphi_3 W_t {z_{t-j}}^2 + \varphi_4 W_t {z_{t-j}}^3 + \varphi_5 W_t {z_{t-j}}^4 + \epsilon_t \qquad (7)$$

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). This explains why some of the estimated specifications do not match the original proposal by the above tests: the suggested specifications either showed strong autocorrelation, did not mathematically converge or – as is the case with France – linearity was not very strongly rejected such that the two models form very close substitutes to an almost linear result.

Estimation and Evaluation

The second stage of the modelling cycle consists of estimating our parameter values. We estimate equation (2) in combination with either (4) or (5) as the transition function $\mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c})$ with non-linear 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 modelling cycle consists of evaluation. The estimation results are examined by simple judgment concerning the plausibility of the parameter values, the convergence of the models, goodness of fit and by inspecting the regimes which the models imply. Our results are also subjected to the misspecification test of no residual autocorrelation by applying a special case of the Breusch-Godfrey Lagrange Multiplier (BG) test suitable for non-linear 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 model, which explains why a few results do not match the original test results for ESTR vs. LSTR. Final results for β can be found in table 7.[12]

A substitution effect from domestic demand to exports should result in a negative coefficient for β. The two extreme regimes in our non-linear estimation are coefficient β_{1i} for $\mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c}) = 0$ (i.e. the linear model) and $\beta_{1i}+\beta_{2i}$ 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 $\beta_i = \beta_{1i}+\beta_{2i} \cdot \mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c})$.

– 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. As evident from figure 1, which is based on an ESTR model for Spain, β_0 displays negative values for low and high levels of past capacity utilisation. This suggests a substitutive relationship between domestic and foreign sales when the economy is close to peak or trough. When capacity utilisation is very low, firms react to a fall in domestic demand by increasing their efforts to export. Conversely, if the economy operates at high capacity utilisation, 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 utilisation (coefficient $\beta_{10} + \beta_{20}$ in table 7), a one percentage point fall in domestic demand generates close to a one percentage point increase in exports; the 95% confidence interval mostly confirms this negative relation, while reaching small positive parameter values as well. For peak times, this elasticity is slightly lower. By contrast, a positive link is identified between domestic demand and exports during normal economic conditions (coefficient β_{10}). 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). 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.

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 from zero for Portugal (both the substitution effects during peak and trough and the positive links during normal times), this is not the case for Italy. Here, the small contemporaneous substitution effect during trough and peak ($\beta_{10} + \beta_{20}$) is found not to be different from zero contrary to the statistically significant positive coefficient for normal times; this also holds for higher lag

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

lengths of the coefficient on domestic demand. Overall, the results suggest that the net effect is a substitutive relation. This indicates 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 tend to sell relatively less to the domestic market and just switch to foreign markets or new firms would enter the export market. During normal economic times, the relationship is 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 found that exports and domestic sales are not substitutive but complementary for a panel of French firms. Our results also show that this complementary relationship holds over the entire values of the transition variable and is not as strong as it is for other countries; we find an elasticity of around 0.5. In addition, for France, the test on non-linearity did not reject linearity as strongly as it did for the other countries. Rather, linearity was only rejected for low lag lengths of the transition variable. The estimation results for France pointed to an ESTR specification while the specification tests suggested an LSTR model; this could also be due to the fact that non-linearity is not as strong as it is for other countries. Figure 4 confirms the notion that non-linearity might not play an important role for the French data. One possible explanation could be that the French business cycle in the years under consideration did not vary as much as that of the other countries. The overall finding of no substitutive relationship may also 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.

Looking at Ireland and Greece, the two countries for which we estimate an LSTR model, we equally find at least weak evidence for a negative link between domestic and foreign sales during periods of low capacity utilisation (figures 5 and 6; coefficients β_{1i} in table 7). This effect, however, is statistically insignificant for both countries and economically only of very modest size. For Greece, the coefficient β_{1i} also shows different qualitative results for higher lag lengths. After passing a critical threshold, exports and domestic demand become complements with an increasing degree of capacity utilisation (coefficient $\beta_{10} + \beta_{20}$; again, for Greece the higher lag lengths for coefficients $\beta_{1i} + \beta_{2i}$ yield different outcomes). A further threshold, for positive domestic demand shocks and high capacity utilisation is not reached. For both countries, therefore, the band of inaction is only restricted to one side. In the case of Ireland, the finding that only economic recessions but not periods of booms might 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. At the same time, the overall small coefficients around zero (both positive and negative) might be due to the large number of multinational corporations in Ireland, which are presumably less tied to the domestic situation and should therefore react less to domestic demand shocks than firms with a strong domestic focus. For Greece, the estimated model somewhat resembles a simple two-regime threshold model where marginal changes of capacity utilisation around its average have strong effects on the relation between domestic demand and exports. Further strong changes, however, do not have any additional effects. Also, at least during the time period under consideration, Greece has never displayed a capacity utilisation rate of more than 80 percent and its average degree of utilisation is much lower than that of the other countries. This could explain why the band of inaction for Greece seems to be restricted only to the side of low capacity utilisation. The interlinkages between exports and domestic demand changes under high capacity utilisation rates remain unknown. Results seem to be less robust during longer lag lengths. Finally, it needs to be noted that the weak substitutive relation could be due to the fact that there is no strong tradable sector in Greece and that the lack of adjustment capacity, institutional weaknesses and structural rigidities in labour and product markets prevented a further rise in Greek exports during the period of the massive fall in domestic demand (Böwer et al. 2014).

Overall, our empirical results strongly suggest that the relationship between domestic sales and exports depends on capacity utilisation and the business cycle. A substitutive relationship between domestic and foreign sales is evident during economic downturns when capacities are only weakly utilised; we obtain a negative coefficient for β_0 in all countries except France.[13] 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.

Robustness Checks

In the following, we are performing some robustness checks to our estimations. We begin with employing two different export variables. First, we take a look at exported goods only. While exported services seem to play an important role for the countries under consideration – for instance in the field of travel and tourism –, for exported goods, capacity constraints should be even more binding. Second, we consider value added exports rather than gross exports. By disregarding imported intermediate goods, we obtain a measure which is more closely related to capacity constraints. Due to data availability reasons, the sample had to be restricted to the period until 2011.

Moreover, we also consider a different type of real effective exchange rate to measure competitiveness. While the results above are based on the REER deflated by consumer price indices with respect to a country's 15 main trading partners, we are here also using the REER deflated by unit labour costs for a country's 24 main trading partners, i.e. capturing cost competitiveness rather than price competitiveness. Also, influencing the values of the respective transition functions, we employ the median instead of the arithmetic mean as the threshold value for our transition variable in order to limit the impact of outliers. Last, as the long-run relation can be sensitive to further structural breaks or omitted variables, we drop the long-run coefficient altogether.

The results of our robustness tests can be found in figures 7 – 11 and tables A3 –A7. Overall, the findings confirm the results we presented above with slight refinements. For Spain and Portugal, the results for the different estimations strongly resemble the original estimations, even though the size of the coefficients decreases considerably when employing value added exports (the same holds for the other countries' results). For Italy, the main finding – namely a substitutive relation between

[13] 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}.

domestic demand and exports during low capacity utilisation – is confirmed in all of the robustness estimations, even though the specification changed from an ESTR to an LSTR model in some cases. The upper threshold for the band of inaction thus seems to be less robust. The original estimation for France showed that non-linearity was less important; it also found a slightly positive relation between domestic demand and exports throughout different values of capacity utilisation. This result is confirmed by most of the robustness estimations, with even smaller coefficients around zero. Only in the case of exported goods do we find a slightly negative coefficient for the domestic demand and export relation for low capacity utilisation values. For Ireland, we also find only weak non-linearities and coefficients around zero in all of our robustness estimations. This again strongly resembles our original findings, reflecting the high flexibility of the Irish economy. Lastly, for Greece we confirm the finding of a coefficient around zero for low capacity utilisation levels. We find a positive coefficient for higher capacity utilisation levels similar to the original findings for our estimations with the ULC deflated REER and median threshold value. When using export goods or value added exports, this positive relation disappears.

– Figures 7 to 11 about here –

5. Conclusions

In this paper, we have analysed the relation between domestic demand and export activity for six euro area countries using non-linear estimations. The results of our macro-econometric smooth transition regression approach indicate that domestic demand behaviour is relevant for the short-run dynamics of several euro area member countries' exports. In particular, the estimation results suggest that on an aggregated level, contemporary and lagged domestic demand developments can affect a country's export performance significantly. In the cases of Spain, Portugal and Italy, the symmetric non-linearity of the relation expresses itself in a substitutive relationship between domestic demand and export activity if deviations from average capacity utilisation are large, independent of their sign. In other words, the substitution effect from domestic demand to exports turns out to be stronger and more significant during more extreme stages of the business cycle. For periods with average levels of capacity utilisation, we find a band of inaction in which the relation between domestic and foreign sales is complementary. On a micro level, theoretical reasons for these findings can be found in the sunk costs hysteresis approach. Only after reaching an upper or lower threshold of capacity utilisation, firms are willing to pay sunk costs to shift activities to another market.

In the cases of Ireland and Greece, we find that the non-linear relationship between domestic demand and exports is asymmetric. Domestic demand and exports are slightly substitutive during a business cycle trough and complements during normal times and in a boom. The sign of the deviation of capacity utilisation from its normal level matters, suggesting that the liquidity channel plays an important role in these countries. For France, the evidence for non-linearity is weaker. We find evidence for mostly complementary relationships.

Overall, we can therefore confirm the short-run non-linear relationship between domestic demand and foreign sales depending on capacity constraints for most countries in our sample. A strong substitutive relationship for times of low capacity utilisation can most clearly be found for Spain, Portugal and Italy. However, we believe there are valid reasons for the different findings in the other countries (such as the high number of multinational corporations in Ireland, the lower openness of the French economy or the small Greek tradable sector).

In recent years, the six countries under consideration have been able to correct their external imbalances, partly by increasing their exports. Our findings provide one possible explanation for the rising exports. The countries are currently in a situation of cyclical weakness with generally low rates of capacity utilisation and a strong decline in domestic demand. We argue that many firms have tried to compensate for weak domestic sales by increasing their effort of selling on foreign markets or even entering the export market in the first place. Our results point to the fact that the observed increase in export market shares accompanying the reduction of the current account deficits might have been due to non-price related factors, such as "survival-driven" exports instead of an increase in price competitiveness as expected by sustainable structural reforms.

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 shifting sales or 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 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, would render 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 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 $(c = 0.780)$

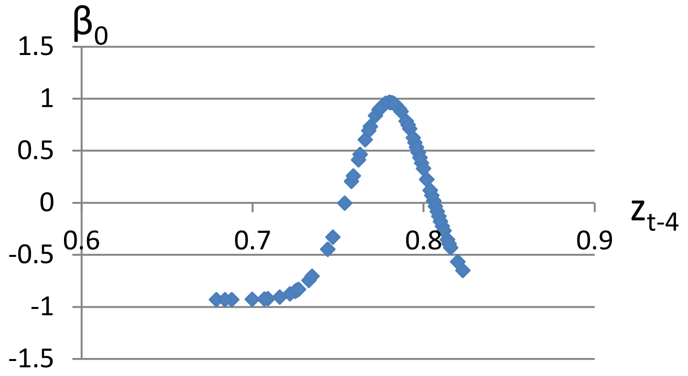


Figure 2: Estimation Results for Portugal $(c = 0.793)$

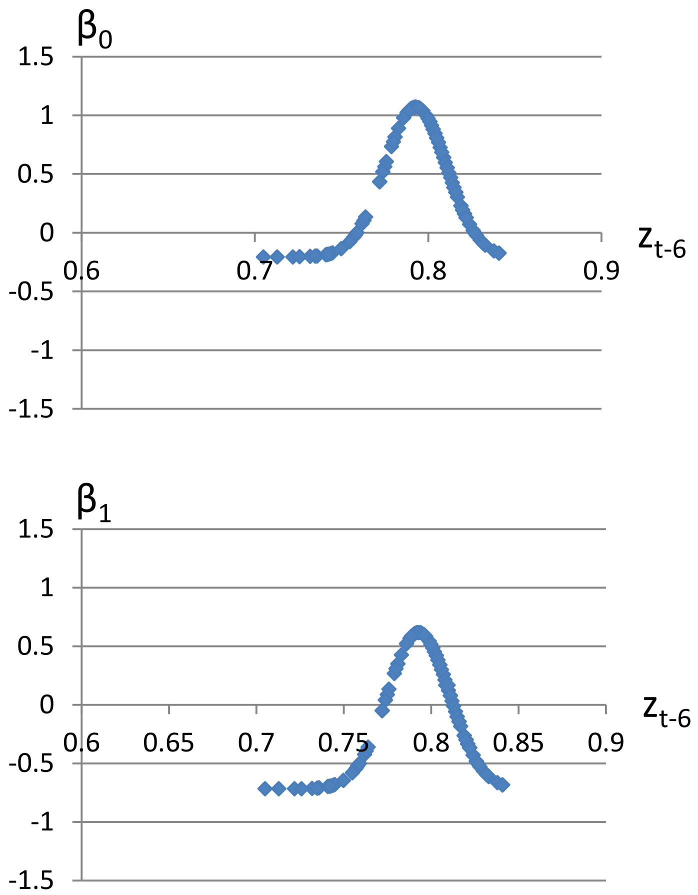


Figure 3: Estimation Results for Italy ($c = 0.751$)

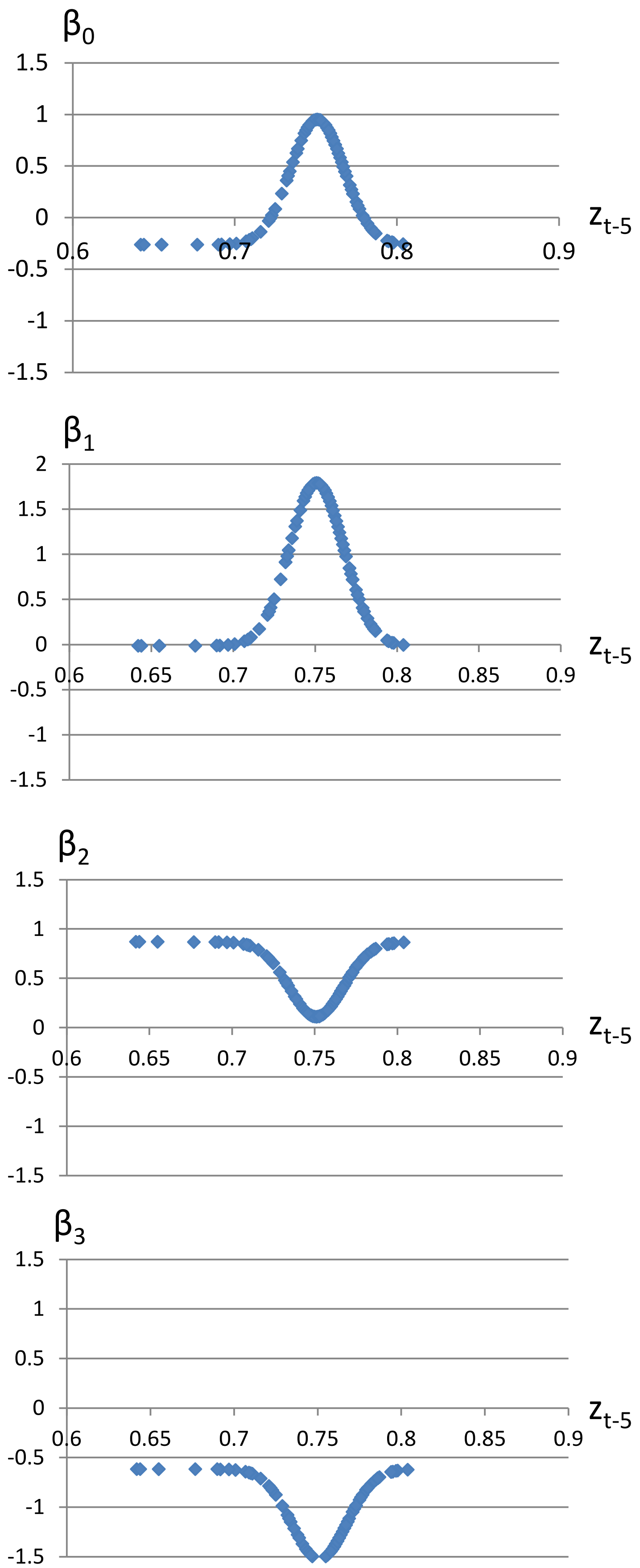


Figure 4: Estimation Results for France $(c = 0.847)$

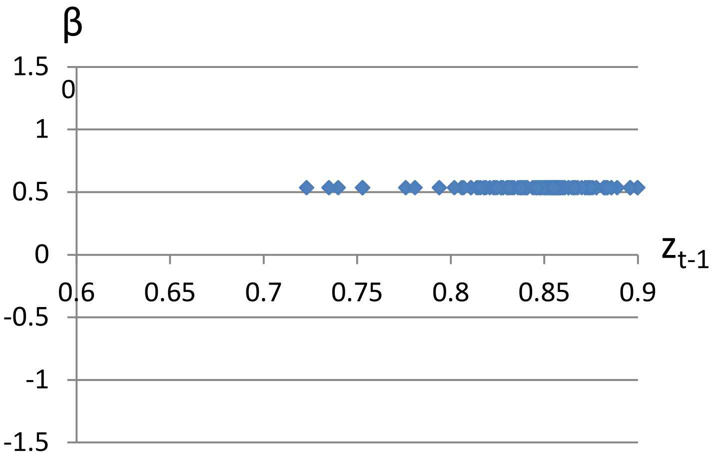


Figure 5: Estimation Results for Ireland $(c = -0.330)$

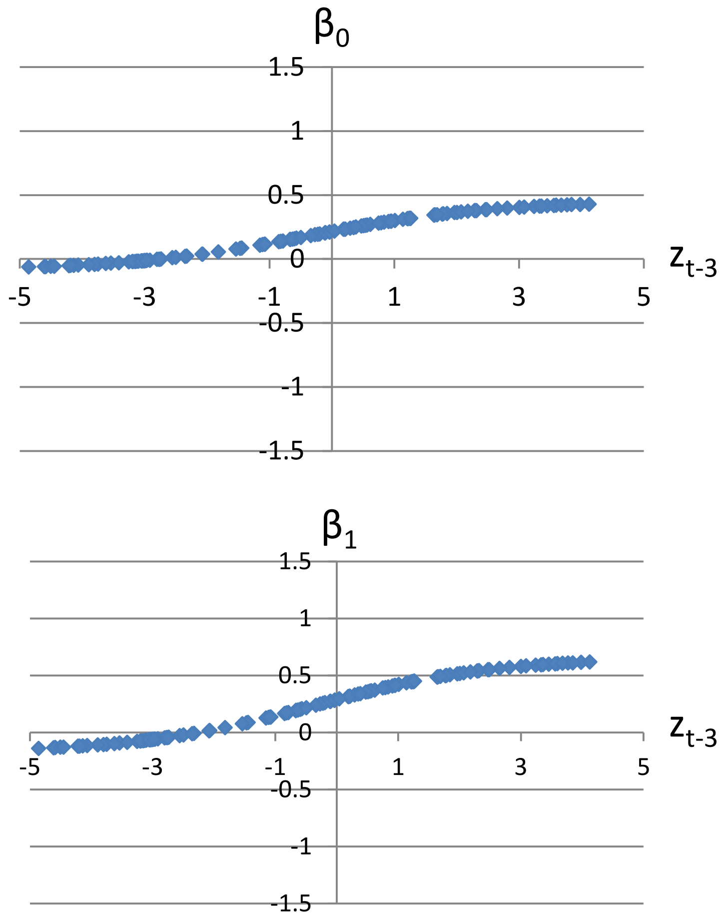


Figure 6: Estimation Results for Greece $(c = 0.748)$

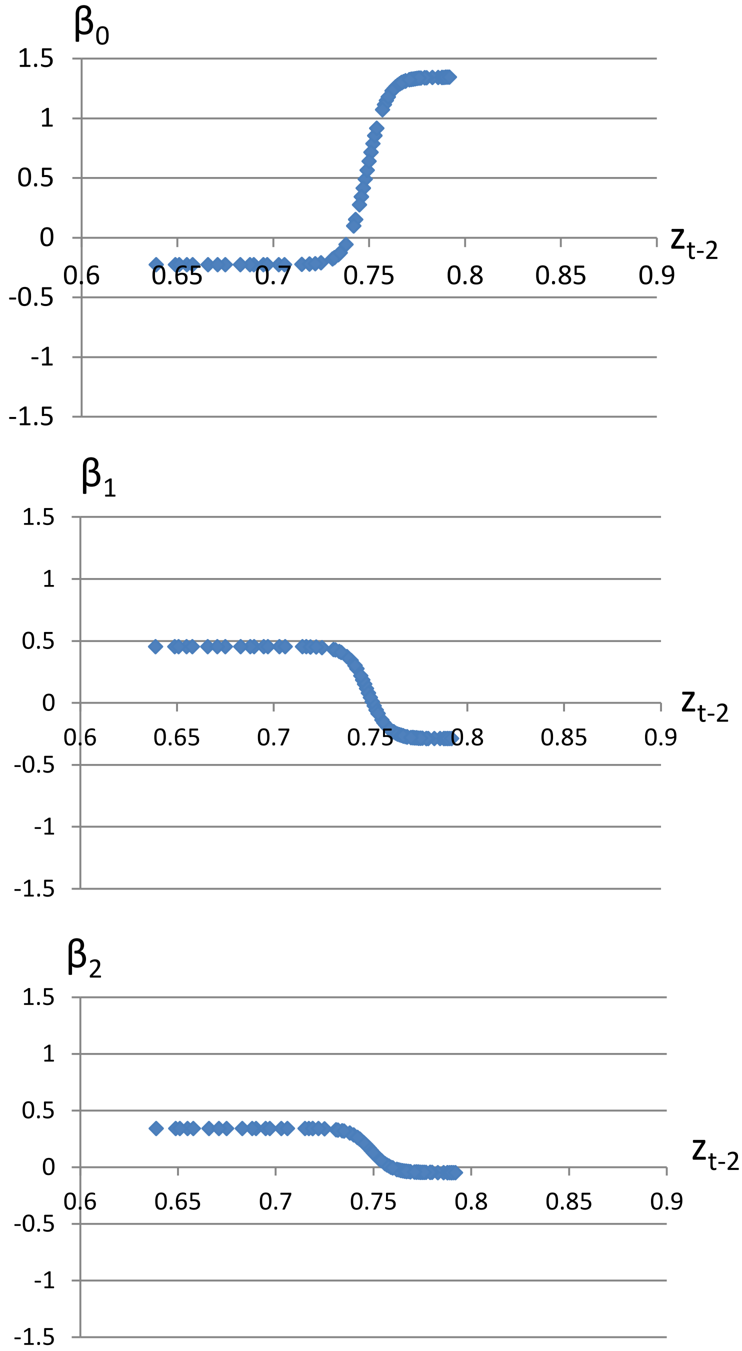
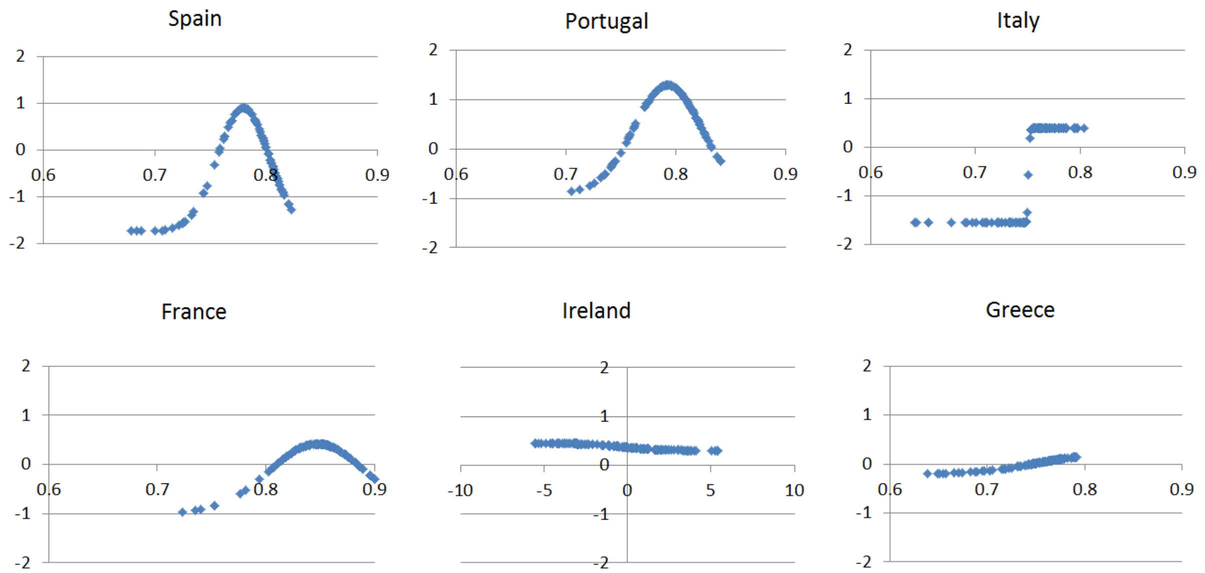
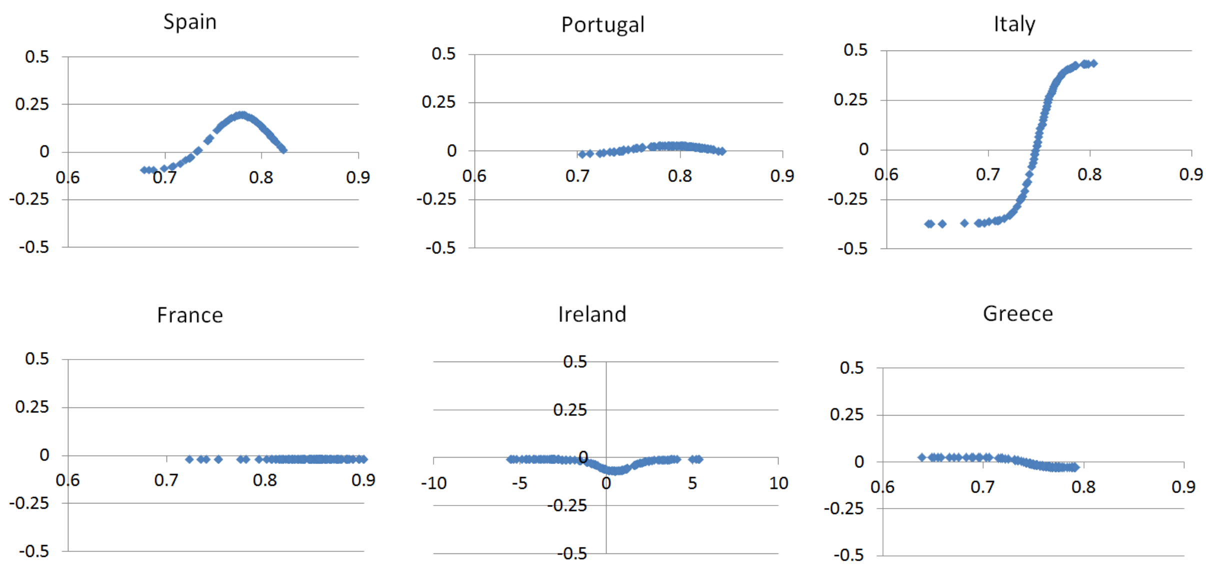


Figure 7: Estimation with export goods



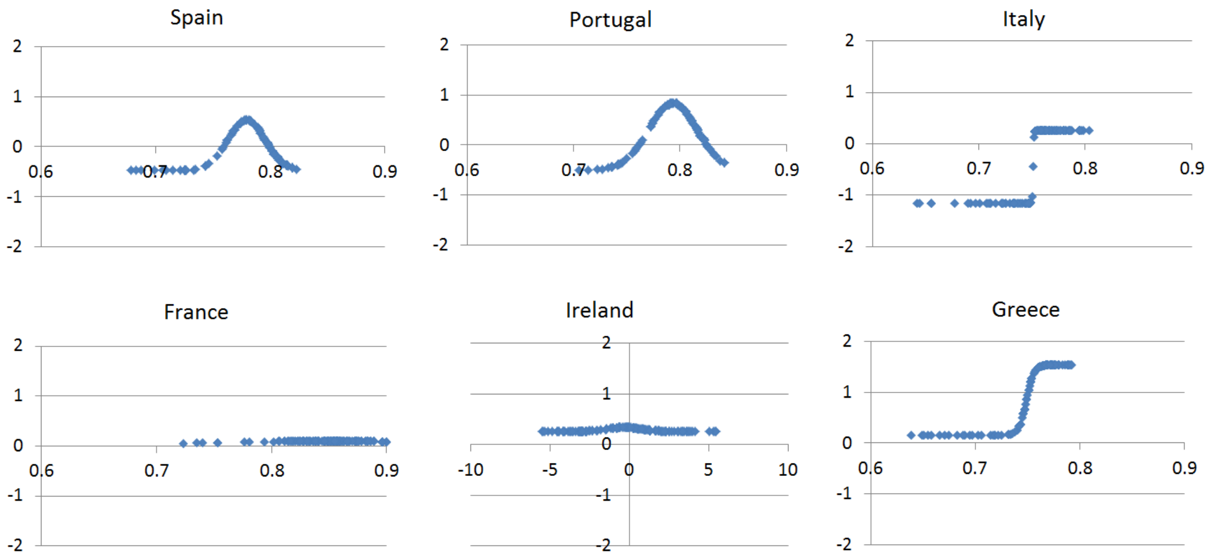
The figures refer to coefficient β_0 which is depicted on the vertical axis; β_0 is defined as $\beta_0 = \beta_{10}+\beta_{20} \cdot \mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c})$. The transition variable z_{t-j} is displayed on the horizontal axis.

Figure 8: Estimation with value added exports



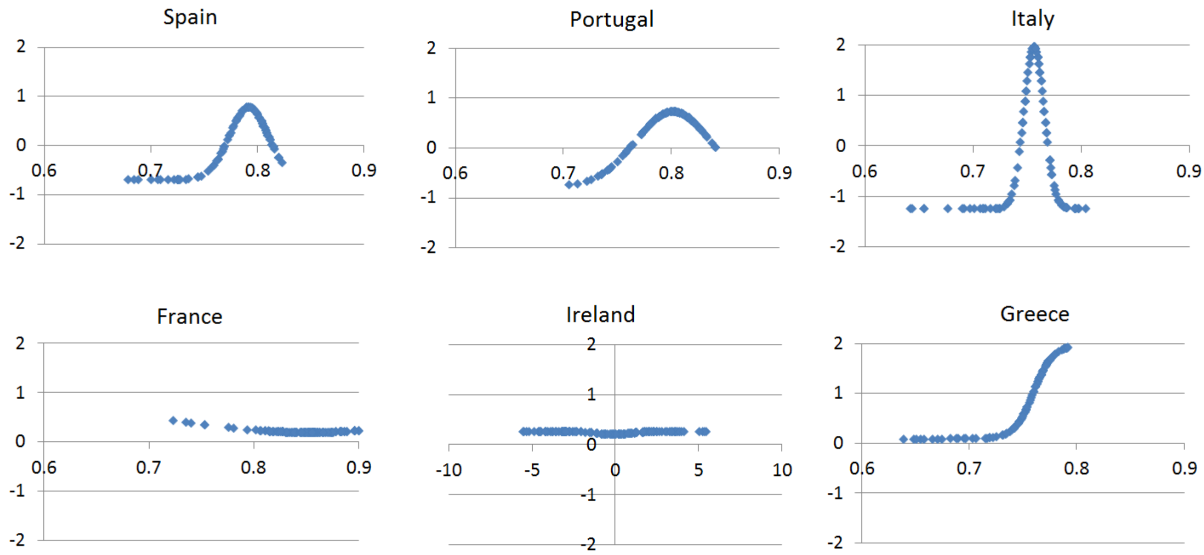
The figures refer to coefficient β_0 which is depicted on the vertical axis; β_0 is defined as $\beta_0 = \beta_{10}+\beta_{20} \cdot \mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c})$. The transition variable z_{t-j} is displayed on the horizontal axis.

Figure 9: Estimation with ULC deflated REER



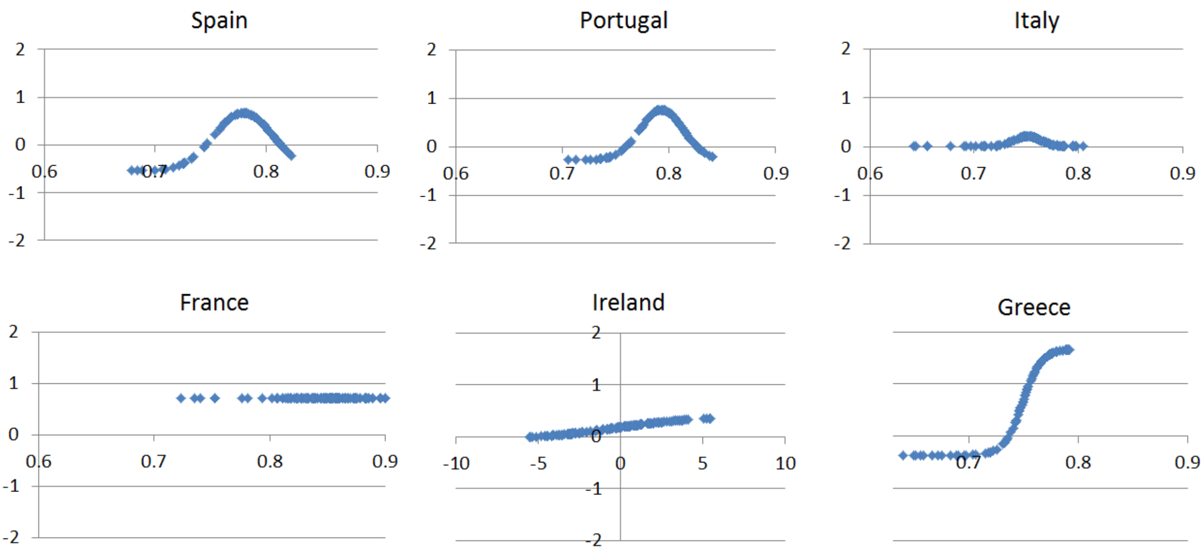
The figures refer to coefficient β_0 which is depicted on the vertical axis; β_0 is defined as $\beta_0 = \beta_{10} + \beta_{20} \cdot \mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c})$. The transition variable z_{t-j} is displayed on the horizontal axis.

Figure 10: Estimation with median as threshold value



The figures refer to coefficient β_0 which is depicted on the vertical axis; β_0 is defined as $\beta_0 = \beta_{10} + \beta_{20} \cdot \mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c})$. The transition variable z_{t-j} is displayed on the horizontal axis.

Figure 11: Estimation without long-run adjustment coefficient



The figures refer to coefficient β_0 which is depicted on the vertical axis; β_0 is defined as $\beta_0 = \beta_{10} + \beta_{20} \cdot \mathrm{F}(\mathrm{z_{t-j}}, \gamma, \mathrm{c})$. The transition variable z_{t-j} is displayed on the horizontal axis.

Table 1: Unit Root Tests

Country	Series	*ADF test* Level t-stat. [lags]	*ADF test* 1st Diff. t-stat. [lags]	*Lee-Strazicich test* 1 break t-stat.	*Lee-Strazicich test* 2 breaks 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
	x_t^{goods}	-1.875 [0]	-12.457*** [0]	-2.4443	-2.9754
	x_t^{va}	-2.407 [8]	-2.093** [10]	-0.7349	-0.7597
	y_t^*	-3.418* [1]	-4.569*** [0]	-1.9472	-2.0878
	r_t	-1.250 [1]	-8.763*** [0]	-1.8106	-1.9323
	r_t^{ULC}	-1.373 [1]	-7.905*** [0]	-1.0327	-1.0664
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
	x_t^{goods}	-1.967 [4]	-3.257*** [3]	-2.6350	-2.9542
	x_t^{va}	-0.750 [8]	-1.843* [3]	-1.1552	-1.1895
	y_t^*	-2.742 [1]	-4.400*** [0]	-1.6444	-1.7162
	r_t	-1.353 [1]	-8.784*** [0]	-2.4693	-2.5850
	r_t^{ULC}	-0.917 [1]	-6.849*** [0]	-1.0068	-1.0402
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
	x_t^{goods}	-3.906** [2]	-8.076*** [0]	-2.5597	-2.9079
	x_t^{va}	-3.251* [7]	-2.585** [7]	-1.4249	-1.4481
	y_t^*	-2.944 [2]	-4.750*** [1]	-2.0089	-2.1816
	r_t	-2.501 [1]	-8.336*** [0]	-1.8317	-1.9321
	r_t^{ULC}	-2.279 [1]	-7.685*** [0]	-1.6470	-1.7732
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
	x_t^{goods}	-2.297 [1]	-7.339*** [0]	-1.2483	-1.3156
	x_t^{va}	-1.509 [8]	-1.842* [7]	-0.7760	-0.8076
	y_t^*	-3.268* [1]	-4.703*** [0]	-2.0007	-2.0854
	r_t	-1.921 [0]	-10.654*** [0]	-2.6688	-2.7981
	r_t^{ULC}	-3.129* [1]	-8.750*** [0]	-1.5954	-1.6572
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
	x_t^{goods}	-1.273 [4]	-4.099*** [3]	-1.3362	-1.4306
	x_t^{va}	-2.308 [8]	-2.059** [7]	-0.5018	-0.5126
	y_t^*	-2.580 [2]	-5.141*** [1]	-1.8182	-1.9890
	r_t	-1.837 [0]	-9.162*** [0]	-1.8346	-1.9568
	r_t^{ULC}	-1.896 [1]	-7.549*** [0]	-1.2778	-1.3429
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
	x_t^{goods}	-3.015 [4]	-5.130*** [3]	-4.1321**	-4.8821***
	x_t^{va}	-1.232 [8]	-1.271 [6]	-0.8985	-0.9393
	y_t^*	-3.646** [1]	-4.249*** [0]	-1.8027	-1.9790
	r_t	-0.810 [0]	-12.329*** [0]	-3.5230*	-3.8786**
	r_t^{ULC}	-2.029 [1]	-9.804*** [0]	-1.9257	-2.0192

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

Tests the null hypothesis that there is no cointegration (i.e. that the residual series has a unit root). The (approximate) critical values for the t-test are from MacKinnon (1991) for the respective number of variables. */**/*** statistical significance at the 10%/5%/1% level.

Table 3: Long-Run Relationship

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

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

Table 4: Teräsvirta test for non-linearity 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	35.016 (0.002) [0.39]	23.955 (0.014) [0.41]	20.509 (0.042) [0.38]	14.832 (0.192) [0.39]	15.798 (0.111) [0.39]	7.532 (0.755) [0.39]	1
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	1	12.20 (0.032)	11.76 (0.038)	5.53 (0.355)	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	1	14.68 (0.066)	15.21 (0.033)	LSTR

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	1	3	2
β_{10}	0.964*** (0.22)	1.072*** (0.13)	0.950** (0.46)	0.535** (0.23)	-0.086 (0.23)	-0.226 (0.20)
β_{20}	-1.897*** (0.22)	-1.278*** (0.14)	-1.214*** (0.38)	-0.135 (0.356)	0.538* (0.31)	1.569*** (0.31)
$\beta_{10}+\beta_{20}$	-0.933*** [0.00]	-0.206** [0.01]	-0.264 [0.54]	0.399** [0.02]	0.452*** [0.00]	1.343*** [0.00]
β_{11}	-	0.617*** (0.15)	1.791*** (0.61)	-	-0.174 (0.29)	0.454*** (0.17)
β_{21}	-	-1.336*** (0.18)	-1.806* (1.00)	-	0.827* (0.45)	-0.743*** (0.27)
$\beta_{11}+\beta_{21}$	-	-0.719*** [0.00]	-0.015 [0.97]	-	0.653*** [0.00]	-0.289** [0.04]
β_{12}	-	-	0.110 (0.49)	-	-	0.341 (0.22)
β_{22}	-	-	0.758** (0.33)	-	-	-0.390* (0.23)
$\beta_{12}+\beta_{22}$	-	-	0.868 [0.12]	-	-	-0.049 [0.86]
β_{13}	-	-	-1.526*** (0.56)	-	-	-
β_{23}	-	-	0.907 (0.64)	-	-	-
$\beta_{13}+\beta_{23}$	-	-	-0.619*** [0.00]	-	-	-
γ	35.566* (18.61)	49.762*** (19.27)	59.061*** (20.89)	1.6381** (0.684)	1.872** (0.86)	6.662*** (2.29)
R^2	*0.773*	*0.566*	*0.603*	*0.568*	*0.683*	*0.686*
p-value BG test	*0.506*	*0.687*	*0.741*	*0.110*	*0.079*	*0.714*

Coefficients estimated by NLS; 95% confidence intervals in brackets; Newey-West standard errors in parentheses. */**/*** statistical significance at the 10%/5%/1% level. For the joint significance of β_{1i} and β_{2i}, the linear restriction $\beta_{1i}+\beta_{2i}=0$ has been tested with Chi-squared test statistics; p-value in brackets. The Breusch-Godfrey Lagrange Multiplier (BG) test is based on the null hypothesis of no serial correlation of the residuals of order $p=4$. Bootstrapped confidence intervals with $n=1000$.
β_{ji} $(j=1,2)$ is the coefficient for domestic demand in the non-linear error correction model. The two extreme regimes are $\mathrm{F}(z_{t-j},\gamma,c)=0$ given by β_{1i} (i.e. for the ESTR model around the threshold value, for the LSTR model for large negative deviations from the threshold) and $\mathrm{F}(z_{t-j},\gamma,c)=1$ given by $\beta_{1i}+\beta_{2i}$ (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 (in prices of reference year)	1980Q1 – 2012Q4; IT: 1981Q1 – 2012Q4
Exports (Goods)	National Statistical Offices	real exports of goods (in prices of reference year)	1980Q1 – 2012Q4; IT: 1981Q1 – 2012Q4
Exports (Value Added)	World Input-Output Database (interpolated)	value added exports (converted to prices of reference year)	1995Q1 – 2011Q1
Domestic Demand	National Statistical Offices	real domestic demand (in prices of reference year)	1980Q1 – 2012Q4; IT: 1981Q1 – 2012Q4
Real Effective Exchange Rate (CPI)	Eurostat	index deflated by consumer price indices with a country's 15 main trading partners	1980Q1 – 2012Q4
Real Effective Exchange Rate (ULC)	Eurostat	index deflated by unit labour costs with a country's 24 main trading partners	1980Q1 – 2012Q4
Foreign Demand	ECB	trade-weighted imports for 15 main trading partners	1980Q1 – 2012Q4
Capacity Utilisation	Eurostat	current level of capacity utilisation in manufacturing industry based on business surveys	PT: 1987Q1 – 2012Q4; IT, GR: 1985Q1 – 2012Q4; ES: 1987Q2 – 2012Q4
Capacity Utilisation	Insee	capacity utilisation rate based on quarterly business survey	FR: 1980Q1 – 2012Q4
Output Gap	AMECO (interpolated)	gap between actual GDP and potential GDP as percentage of potential GDP	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	1	3	2
α_1	0.0307*** (0.01)	0.007*** (0.00)	0.002 (0.00)	-0.010*** (0.002)	0.012*** (0.00)	0.011 (0.01)
β_{10}	0.964*** (0.22)	1.072*** (0.13)	0.950** (0.46)	0.535** (0.23)	-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)	0.514*** (0.11)	-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)	-0.023 (0.11)	-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)	0.448*** (0.15)	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)	-0.173*** (0.05)	0.065*** (0.03)	-0.374*** (0.07)
α_2	0.039 (0.03)	-0.013*** (0.00)	-0.007* (0.00)	0.017*** (0.00)	-0.005 (0.00)	-0.009 (0.01)
β_{20}	-1.897*** (0.22)	-1.278*** (0.14)	-1.214*** (0.38)	-0.135 (0.356)	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)	-0.233** (0.09)	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)	-0.534*** (0.20)	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)	-0.373*** (0.14)	-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)	-0.168*** (0.06)	-0.175*** (0.05)	0.025 (0.13)
γ	35.566* (18.61)	49.762*** (19.27)	59.061*** (20.89)	1.638** (0.68)	1.872** (0.86)	6.662*** (2.29)
R^2	*0.773*	*0.566*	*0.603*	*0.568*	*0.683*	*0.686*
p-value BG test	*0.506*	*0.687*	*0.741*	*0.110*	*0.079*	*0.714*

Coefficients estimated by NLS; Newey-West standard errors in parentheses. */**/*** statistical significance at the 10%/5%/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: Estimation with export goods

	Spain	**Portugal**	**Italy**	**France**	**Ireland**	**Greece**
specification	ESTR	ESTR	LSTR	ESTR	LSTR	LSTR
lag length	5	5	6	3	3	2
β_{10}	0.893*** [-0.38; 2.36] (0.18)	1.291*** [-0.09; 3.33] (0.25)	-1.556*** [-2.91; 0.02] (0.58)	0.420 [-12.19; 22.04] (0.55)	0.452* [-0.52; 1.50] (0.24)	-0.204 [-1.19; 0.84] (0.28)
β_{20}	-2.633*** [-44.38; -0.10] (0.38)	-2.182*** [-7.82; 0.48] (0.58)	1.944*** [0.14; 4.23] (0.46)	-1.409** [-24.64; 12.32] (0.65)	-0.148 [-1.79; 1.36] (0.24)	0.427 [-1.33; 2.10] (0.32)
$\beta_{10} + \beta_{20}$	-1.739*** [-43.53; -0.21]	-0.891** [-6.04; 0.76]	0.387 [-0.82; 2.01]	-0.990*** [-3.02; 1.23]	0.305*** [-0.47; 1.07]	0.223** [-0.69; 1.12]
γ	31.952*** (3.68)	16.717*** (3.37)	67.460** (27.61)	8.211*** (1.13)	2.519* (1.33)	1.192*** (0.40)
R^2	*0.861*	*0.502*	*0.720*	*0.225*	*0.547*	*0.616*
p-value BG test	*0.445*	*0.002*	*0.861*	*0.055*	*0.434*	*0.695*

Coefficients estimated by NLS; 95% confidence intervals in brackets; Newey-West standard errors in parentheses. */**/*** statistical significance at the 10%/5%/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 $p = 4$. Bootstrapped confidence intervals with $n = 1000$.

Table A4: Estimation with value added exports

	Spain	Portugal	Italy	France	Ireland	Greece
specification	ESTR	ESTR	LSTR	ESTR	ESTR	LSTR
lag length	6	6	6	3	2	2
β_{10}	0.193 [0.09; 0.35] (0.01)	0.027 [-0.41; 0.48] (0.02)	-0.372 [-2.78; 1.70] (0.19)	-0.023 [-2.02; 3.06] (0.08)	-0.068 [-1.37; 0.65] (0.00)	0.024 [-0.05; 0.14] (0.01)
β_{20}	-0.293 [-1.47; 0.22] (0.05)	-0.046 [-3.34; 3.39] (0.03)	0.807 [-1.72; 3.43] (0.38)	0.459 [-2.87; 2.50] (0.16)	0.061 [-0.66; 1.36] (0.00)	-0.055 [-0.23; 0.06] (0.01)
$\beta_{10} + \beta_{20}$	-0.100 [-1.26; 0.34]	-0.019 [-2.98; 3.26]	0.435 [-0.08; 0.95]	0.436 [-0.45; 1.23]	-0.006 [-0.03; 0.03]	-0.031 [-0.12; 0.03]
γ	17.346 (4.35)	11.611 (4.15)	3.455 (1.31)	0.687 (0.13)	1.409 (0.11)	3.456 (0.42)
R^2	*0.996*	*0.948*	*0.937*	*0.919*	*0.999*	*0.996*
p-value BG test	*0.437*	*0.006*	*0.000*	*0.000*	*0.036*	*0.113*

Coefficients estimated by NLS; 95% confidence intervals in brackets; Newey-West standard errors in parentheses. */**/*** statistical significance at the 10%/5%/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 $p = 4$. Bootstrapped confidence intervals with $n = 1000$.

Table A5: Estimation with ULC deflated REER

	Spain	Portugal	Italy	France	Ireland	Greece
specification	ESTR	ESTR	LSTR	ESTR	ESTR	LSTR
lag length	5	6	6	3	2	2
β_{10}	0.540*** [-0.51; 1.47] (0.08)	0.846*** [-0.16; 2.20] (0.09)	-1.161*** [-2.25; 0.27] (0.32)	0.083 [-0.81; 1.22] (0.59)	0.353*** [-0.27; 1.08] (0.05)	0.159 [-1.16; 0.94] (0.11)
β_{20}	-1.003*** [-4.38; 0.42] (0.13)	-1.335*** [-5.75; 0.32] (0.18)	1.418*** [-0.41; 3.40] (0.24)	-0.960 [-0.98; 1.93] (0.68)	-0.087 [-1.00; 0.77] (0.11)	1.380*** [0.39; 3.78] (0.29)
$\beta_{10} + \beta_{20}$	-0.463*** [-3.61; 0.23]	-0.489*** [-5.35; 0.40]	0.257 [-0.79; 1.61]	-0.877*** [-0.09; 1.44]	0.266* [-0.15; 0.70]	1.539*** [1.02; 2.78]
γ	65.930*** (6.59)	30.800** (13.03)	72.346** (30.92)	10.665*** (1.37)	1.378*** (0.48)	9.688 (7.32)
R^2	*0.840*	*0.569*	*0.724*	*0.064*	*0.678*	*0.652*
p-value BG test	*0.810*	*0.738*	*0.372*	*0.069*	*0.159*	*0.957*

Coefficients estimated by NLS; 95% confidence intervals in brackets; Newey-West standard errors in parentheses. */**/*** statistical significance at the 10%/5%/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 $p = 4$. Bootstrapped confidence intervals with $n = 1000$.

Table A6: Estimation with median as threshold value

	Spain	Portugal	Italy	France	Ireland	Greece
specification	ESTR	ESTR	ESTR	ESTR	ESTR	LSTR
lag length	6	6	5	2	3	2
β_{10}	0.768*** [-0.26; 1.67] (0.11)	0.731*** [-0.18; 1.65] (0.73)	1.970*** [0.11; 4.07] (0.22)	0.191 [-3.05; 4.42] (0.19)	0.206*** [-0.55; 0.87] (0.07)	0.091 [-2.03; 0.73] (0.18)
β_{20}	-1.477*** [-5.29; 0.15] (0.20)	-1.506*** [-8.19; 1.52] (0.39)	-3.214*** [-79.96; -0.18] (0.66)	0.518 [-4.16; 4.28] (0.42)	0.066 [-0.86; 1.06] (0.21)	1.881*** [0.17; 3.86] (0.30)
$\beta_{10} + \beta_{20}$	-0.709*** [-4.74; 0.43]	-0.775 [-7.44; 1.25]	-1.244*** [-78.48; 0.70]	0.710** [-0.54; 1.94]	0.272 [-0.18; 0.72]	1.971*** [1.12; 3.07]
γ	51.476*** (9.79)	12.917 (11.78)	181.427*** (25.62)	1.154** (0.46)	1.140** (0.47)	3.869*** (0.65)
R^2	*0.811*	*0.492*	*0.689*	*0.520*	*0.681*	*0.805*
p-value BG test	0.599	0.771	0.380	0.258	0.110	0.766

Coefficients estimated by NLS; 95% confidence intervals in brackets; Newey-West standard errors in parentheses. */**/*** statistical significance at the 10%/5%/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 $p = 4$. Bootstrapped confidence intervals with $n = 1000$.

Table A7: Estimation without long-run adjustment coefficient

	Spain	Portugal	Italy	France	Ireland	Greece
specification	ESTR	ESTR	ESTR	ESTR	LSTR	LSTR
lag length	4	6	5	1	3	2
β_{10}	0.671*** (0.19) [-0.20; 1.58]	0.762*** (0.12) [-2.27; 4.30]	0.213 (0.68) [-1.55; 1.80]	0.714*** (0.26) [-0.82; 1.95]	-0.062 (0.24) [-0.45; 0.66]	-0.365* (0.20) [-1.62; 0.39]
β_{20}	-1.205*** (0.18) [-7.67; 0.73]	-1.034*** (0.17) [-6.92; 5.29]	-0.209 (0.55) [-4.94; 2.74]	0.039 (0.54) [-1.82; 2.11]	0.464 (0.31) [-0.60; 1.14]	2.036*** (0.34) [1.04; 4.58]
$\beta_{10} + \beta_{20}$	-0.534* [-6.94; 0.78]	-0.271*** [-4.47; 3.13]	0.004 [-4.41; 1.75]	0.754** [-0.13; 1.63]	0.403*** [-0.08; 0.80]	1.671*** [1.12; 3.10]
γ	24.507 (23.95)	38.141** (17.69)	96.197*** (13.60)	1.362* (0.72)	1.032*** (0.00)	4.064*** (1.52)
R^2	*0.751*	*0.547*	*0.504*	*0.358*	*0.667*	*0.626*
p-value BG test	0.638	0.735	0.164	0.101	0.058	0.934

Coefficients estimated by NLS; 95% confidence intervals in brackets; Newey-West standard errors in parentheses. */**/*** statistical significance at the 10%/5%/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 $p = 4$. Bootstrapped confidence intervals with $n = 1000$.

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