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Exports and Capacity Constraints A smooth transition regression model for six euro-area countries

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Abstract

The significant gains in export market shares made in a number of vulnerable euro-area crisis countries have not been accompanied by an appropriate improvement in price competitiveness. This paper argues that, under certain conditions, firms consider export activity as a substitute for 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 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 pressure; error-correction models; hysteresis; modelling techniques; smooth transition models; exports; sunk costs

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1. Introduction

A number of euro area countries that 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 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 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. The emerging residuals can, however, be potentially matched by the parallel dramatic fall of domestic demand, as shown by Esteves & 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 there have not been many studies on the effects of domestic demand pressure on the inclination and/or capacity to export, their roots date back to the 1960s.[1] 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

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

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 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. The basic idea is that the strength of the relationship between domestic demand and exports depends on the stage of the business cycle. Specifically, two non-linearities can be distinguished: First, the export response to domestic demand developments might 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. Second, hysteresis considerations suggest that firms substitute between domestic and foreign sales only during extreme stages of the business cycles. Sunk costs prevent then a sharp export reaction to domestic demand developments during periods of average capacity utilisation, while the substitution effect may increase during periods of economic stress and booms. Empirically, a substitutive relationship should then be found both for very low and very high levels of capacity utilisation. Whereas the previous literature investigates only the first type of non-linearity (see e.g. Berman et al., 2011; Esteves & Rua, 2013), we also focus – as an innovation – on the latter.

The paper proceeds as follows. In section 2, we present different theoretical approaches that 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 relationship 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 use the logistic and the exponential STR model to account for the two different kinds of non-linearities described above. We also present several robustness tests. 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 maximise 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 the opposite direction, which raises overall efficiency in the long-run (Belke et al., 2013; Esteves & 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, much theoretical and empirical research *at the firm level* has been conducted that allows for a deeper foundation of the relationship between domestic demand and exports (Berman et al., 2011; Blum et al., 2011 and 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 & McQuoid (2013) and Ilmakunnas & 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 two-fold: 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 & Rua, 2013).[2]

A second and more direct impact of domestic demand pressure on exports refers to the *supply side* of exports, assuming that foreign sales can be increased infinitely. In their survey, Ahn & McQuoid (2013) deal with the sources of export-domestic sales trade-offs and trace 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 labour. When a firm experiences a demand increase in one market and increases its sales in that market, the firm's marginal costs will increase. Due to higher marginal costs, it would then be optimal to 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.

Overall, the main lesson from the available empirical literature is that any exercise in 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 due to sunk costs. We argue that the strength of the relationship between domestic demand and exports depends on capacity constraints and the business cycle in general. The substitutability argument above demonstrates that firms will try to export more following a

[2] 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.

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

negative domestic demand shock. With sunk costs for shifting between domestic and export markets, making this shift might not be considered worthy as long as capacity is still 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 therefore increasing overall exports could then be considered as 'survival-driven', rather than simply being due to an increase in competitiveness. Likewise, the decision to shift activity to the export market could also 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. Finally, 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.

Non-linearities in the relationship between domestic and foreign sales are further strengthened by sunk costs of export market entry. In principle, 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 & Rua, 2013). For firms selling only on the domestic market, the capacity utilisation threshold can be considered an important factor in determining export market entry. This could be due to irreversible costs that firms need to pay to enter a foreign market, which are sunk ex post (Baldwin & Krugman, 1989). Moving into export markets and building a global network for exports require 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; on 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 under certain conditions: as long as domestic demand is strong and capacity is highly utilised, there is no strong reason and often even no capacity to export. With average capacity utilisation, capacities exist for serving export markets, but sunk entry costs might deter firms from entering. Only following a negative domestic demand shock resulting in ample available capacity for exporting, do 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 both domestic and foreign markets cannot be served at the same time due to capacity constraints.

In a theoretical model, if there is uncertainty about returns, the decision to shift sales from domestic to export markets or to switch on or off export activity can be analysed, based on the Dixit-type 'investment under uncertainty' model (Dixit & 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 and show 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 & Tybout (1997), Bernard & Wagner

(2001), Bernard & Jensen (2004) and Campa (2004) confirm these findings. In these micro models of hysteresis in export market participation, a band of inaction emerges due to switching costs for firms between serving the domestic and foreign market. 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 utilisation from its normal level is either highly positive (upper threshold) or highly negative (lower threshold). Small changes in capacity utilisation will not induce these effects. It will require a significant negative domestic demand shock for firms to reach the lower threshold where they pay either the sunk costs for shifting sales between the markets or the sunk entry costs to switch to export activity. In the same vein, in order to avoid paying the shifting or exit costs and repaying the entry costs, active exporters may only shift back sales to the domestic market or leave the export market altogether if domestic demand pressure increases strongly and capacity-constraint considerations become pressing at the upper threshold (Belke & Goecke, 2005; Esteves & Rua, 2013).

In the context of this paper, we will therefore analyse the relationship between domestic demand and export activity in a non-linear framework. Based on the depicted micro foundation, 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. 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 that 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 relationship 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 previous literature, 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 utilisation), but the effect strongly increases for larger changes in conditions. We therefore 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.

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

Data

Our data stem from different sources (see Table A1 in the Appendix): Data on real exports (x_t, x_t^{goods}) (both goods and services or goods only) 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. Data on 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) stem from the Business and Consumer Surveys by the European Commission, available from Eurostat. For France, these data come from Insee. In the case of Ireland, no data on capacity utilisation are available. For this country, 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.

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.[5]

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).

[5] 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.

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
	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

Notes: ADF test: lag length is chosen by minimising 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. See Lee & Strazicich (2004 and 2003).

*/**/*** statistical significance at the 10%/5%/1% level.

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.[6] With time series data for the countries in question, there might be the issue of structural breaks in their long-run relationship, e.g. 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 & Perron (2003)[7] and by a Gregory-Hansen cointegration test (Gregory & Hansen, 1996a and 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 1992-93 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 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 which results in the following long-run equilibrium relationship:

$$x_t = b_1 + b_2 y_t^* + b_3 r_t + b_4 d + b_5 d \cdot y_t^* + b_6 d \cdot r_t + e_t \tag{2}$$

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.

[6] As robustness checks, we also included further variables such as trade openness in the long-run relationship or dropped the long-run relationship altogether. Since this did not change our final short-run non-linear estimation results in a noteworthy way, we do not report these results here.

[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 break (which, however, would have included only a short number of time periods after the break).

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

Notes: 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.

For each country, we find that $\hat{e}_t \sim I(0)$ 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. 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)	1998Q1	0.951

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

Based on theory, the expected outcome for the long-run relationship is a positive relationship 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 relationship, 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 the one for the income elasticity and broadly in line with other studies for the price elasticity (see e.g. European Commission, 2011).

Types of non-linearity

As a next step, we look at short-run adjustments and in particular at the short-run relationship 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 non-linearities. This allows us to investigate a non-linear adjustment process to a linear long-run equilibrium relationship depending on the state of the economy. A variable might e.g. react more sharply 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 non-linear short-run adjustment in STR form then looks like this:

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

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

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.[8] This is a finding also supported e.g. by Esteves & Rua (2013) who argue that it is unclear in which way domestic demand should theoretically enter the long-run export demand equation: periods of strong domestic demand could lead to neglect of export possibilities, but periods of weak domestic demand could stimulate investment towards exports. Contrary to the long-run estimation, we do not include a structural break in the short-run specification. This is because our short-run specification already includes non-linearities by applying the smooth transition regression

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

model. Furthermore, a break in the long-run relationship 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 $F(z_{t-j}, \gamma, c)$ 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 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:

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

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 the time period in our sample in each country. 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 normalise it by σ_z in order to be scale-free (see 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: $\lim_{z_{t-j} \to -\infty} F(z_{t-j}, \gamma, c) = 0$, i.e. the model collapses to just the linear part, and ii) positive deviations of the transition variable from its threshold value: $\lim_{z_{t-j} \to +\infty} F(z_{t-j}, \gamma, c) = 1$. 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 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:

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

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: $\lim_{z_{t-j} \to \pm\infty} F(z_{t-j}, \gamma, c) = 1$ and ii) small deviations of the transition variable from the threshold: $\lim_{z_t - j \to c} F(z_{t-j}, \gamma, c) = 0$, i.e. the non-linear part disappears in the latter extreme.

One example of application is 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:

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

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; 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 $j = 1,2,3,4,5,6$.[9] 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.[10] A non-linear model therefore seems to be suitable for the countries in our sample.

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

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

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*
	372.18	178.31	85.41	920.17	118.78	111.00	
Spain	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	4
	[0.58]	[0.51]	[0.53]	[0.60]	[0.56]	[0.58]	
	34.50	33.48	108.94	121.89	251.97	1270.97	
Portugal	(0.001)	(0.001)	(0.000)	(0.000)	(0.000)	(0.000)	6
	[0.34]	[0.38]	[0.37]	[0.33]	[0.41]	[0.45]	
	105.25	137.53	55.13	79.38	116.32	113.27	
Italy	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	6
	[0.46]	[0.46]	[0.42]	[0.50]	[0.51]	[0.59]	
	35.016	23.955	20.509	14.832	15.798	7.532	
France	(0.002)	(0.014)	(0.042)	(0.192)	(0.111)	(0.755)	1
	[0.39]	[0.41]	[0.38]	[0.39]	[0.39]	[0.39]	
	188.90	249.53	182.05	204.51	100.73	89.36	
Ireland	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	4
	[0.65]	[0.64]	[0.65]	[0.68]	[0.64]	[0.60]	
	1764.02	1619.83	146.17	97.69	137.47	180.74	
Greece	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	2
	[0.51]	[0.58]	[0.49]	[0.49]	[0.51]	[0.47]	

Notes: 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 $\bar{R}^2$.

Based on equation (7), we also approach the choice between an ESTR and an LSTR model (see Teräsvirta, 1994 and 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.

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. 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	12.200 (0.032)	11.759 (0.038)	5.526 (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

Note: χ^2 test statistic realisations are displayed with p-values in parentheses.

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 (see 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 & Jordá (1999). They argue that using equation (7) does not capture all important features and suggest a second-order Taylor approximation yielding the following auxiliary regression:

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

The hypotheses tested here are H_{0E}: $\varphi_3 = \varphi_5 = 0$ and H_{0L}: $\varphi_2 = \varphi_4 = 0$. Escribano and Jordá suggest choosing 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. 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	14.684 (0.066)	15.212 (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

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

In general, it can be argued that once linearity has been rejected, the LSTR and ESTR models 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.

Estimation and evaluation

The second stage of the modelling 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 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, correctness of fit and by inspecting the regimes the models imply. Our results are also subjected to the mis-specification 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 models. Final results for β can be found in Table 7.[11]

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_{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. 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*** [-0.08; 2.57] (0.22)	1.072*** [-0.006; 3.027] (0.13)	0.950** [-0.482; 2.496] (0.46)	0.535** [-0.147; 1.916] (0.23)	-0.086 [-0.670; 0.442] (0.23)	-0.226 [-1.117; 0.416] (0.20)
β_{20}	-1.897*** [-8.461; -0.085] (0.22)	-1.278*** [-4.962; 0.318] (0.14)	-1.214*** [-10.851; 1.341] (0.38)	-0.135 [-1.967; 0.913] (0.356)	0.538* [-0.275; 1.544] (0.31)	1.569*** [0.761; 3.297] (0.31)
$\beta_{10} + \beta_{20}$	-0.933*** [-7.478; 0.117]	-0.206** [-4.325; 0.426]	-0.264 [-9.693; 1.387]	0.399** [-0.432; 1.195]	0.452*** [0.045; 0.940]	1.343*** [0.894; 2.361]

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

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

Notes: 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_{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 $n = 1000$.

β_{j0} $(j = 1,2)$ is the coefficient for domestic demand in the non-linear error correction model. The two extreme regimes are $F(z_{t-j}, \gamma, c) = 0$ 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 $F(z_{t-j}, \gamma, c) = 1$ 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).

Figure 1. Estimation results for Spain $(c = 0.780)$

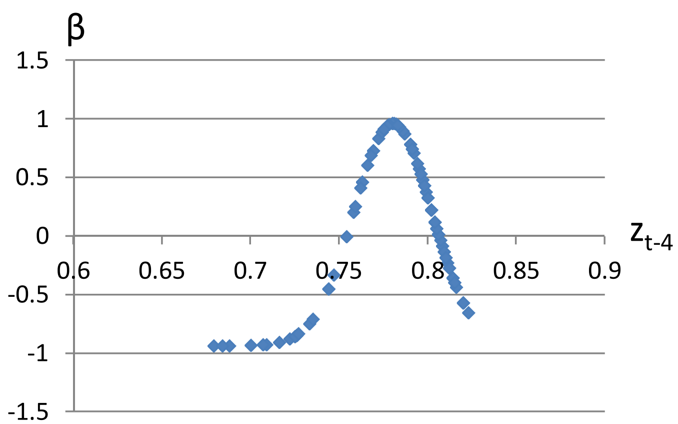


Figure 2. Estimation results for Portugal $(c = 0.793)$

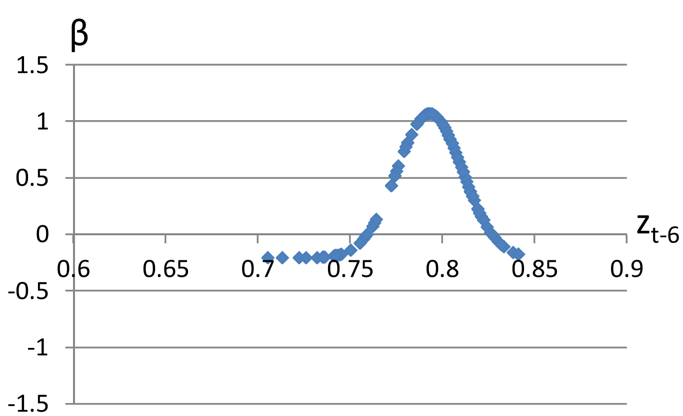


Figure 3. Estimation results for Italy ($c = 0.751$)

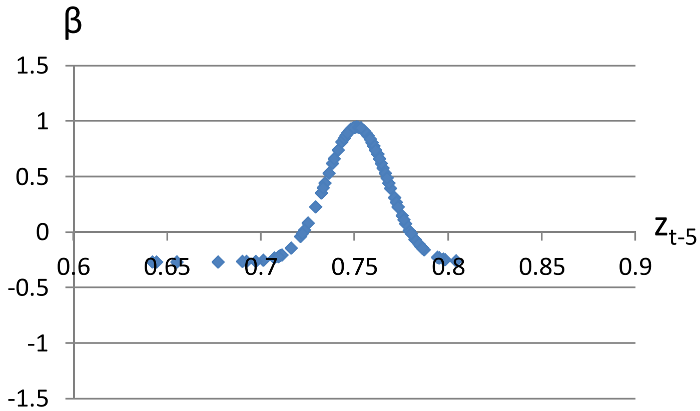


Figure 4. Estimation results for France ($c = 0.847$)

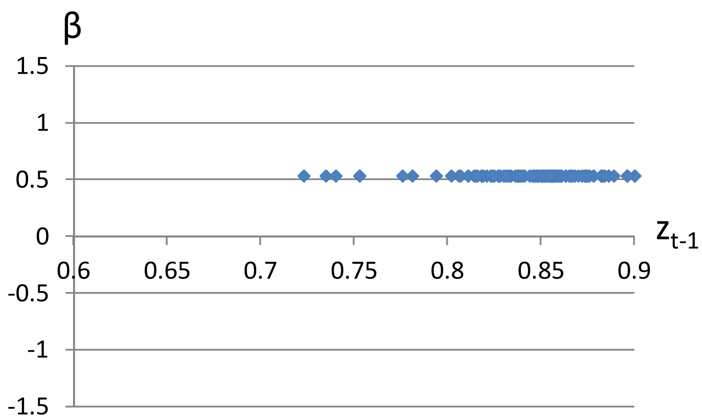


Figure 5. Estimation results for Ireland ($c = -0.330$)

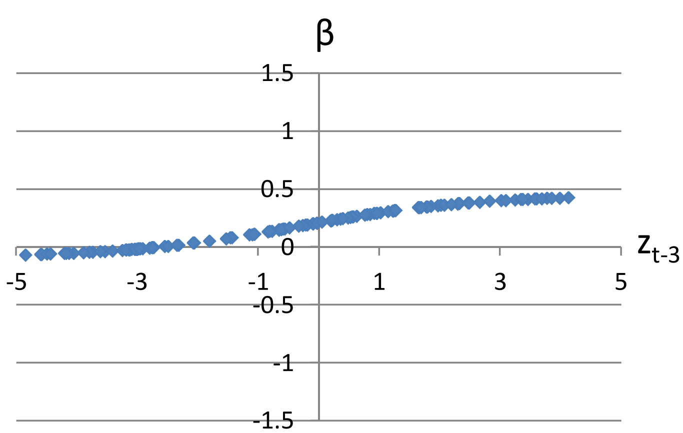
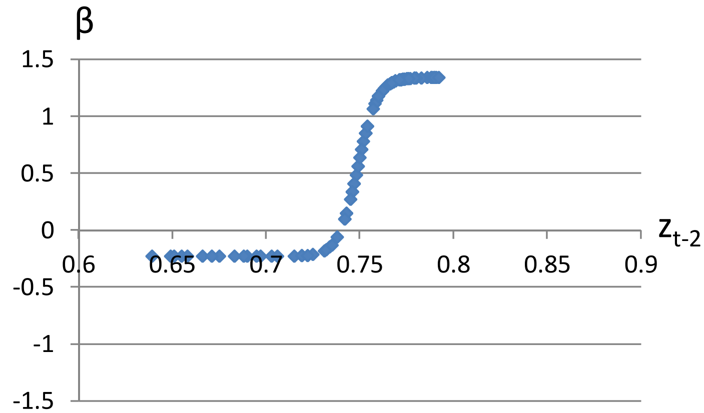


Figure 6. Estimation results for Greece ($c = 0.748$)



Estimation results

Let us first turn to the countries for which the econometric specification warrants an ESTR model. As is 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 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 (although somewhat less strong in economic terms) are found for Portugal and Italy, as shown in Figures 2 and 3. Whereas the estimated coefficients for domestic demand are statistically significant from zero 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 ($\beta_{10} + \beta_{20}$) is found not to be different from zero, contrary to the statistically significant positive coefficient for normal times. The 95% confidence interval for the joint coefficient $\beta_{10} + \beta_{20}$ also includes not only the expected negative values, but positive values as well. Overall, the results suggest that the net effect is a substitutive relationship. This indicates that, as a reaction to a negative domestic demand shock, firms that 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 & 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. The 95% confidence interval includes negative elasticities as well. 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. In addition, 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 characterised 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 (see Figures 5 and 6; coefficient β_{10} in Table 7). This effect, however, is statistically insignificant for both countries and economically only of very modest size; the confidence intervals also include small positive values. After passing a critical threshold, exports and domestic demand become complements with an increasing degree of capacity utilisation (coefficient $\beta_{10} + \beta_{20}$), supported by positive confidence interval values. 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 relationship 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% 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 and rates remain unknown. Finally, it needs to be noted that the weak substitutive relationship could be due to the fact that there is no strong tradable sector in Greece.

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 in all countries except France.[12] 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 this section, we perform some robustness checks to our estimations. We begin by 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 that is more closely related to capacity constraints. Due to data availability reasons, the sample had to be restricted to the period until 2011.

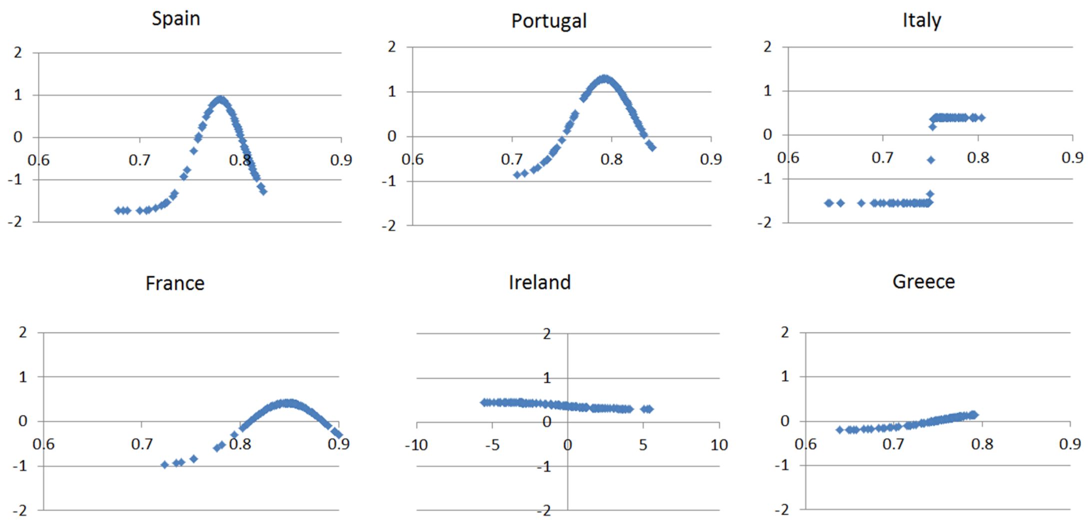
Moreover, we also consider a different type of real effective exchange rate to measure competitiveness. While the results above are based on the REER deflated by consumer price indices with respect to a country's 15 main trading partners, we are here also using the REER deflated by unit labour costs for a country's 24 main trading partners, i.e. capturing cost competitiveness rather than price competitiveness. Lastly, 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.

The results of our robustness tests can be found in Figures 7-10 and Tables A3-A6. 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 relationship between 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 seems to be less important, however. The original estimation for France showed that non-linearity was less important; it also found a slightly positive relationship 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 relationship 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

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

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 relationship disappears.

Figure 7. Estimation with export goods

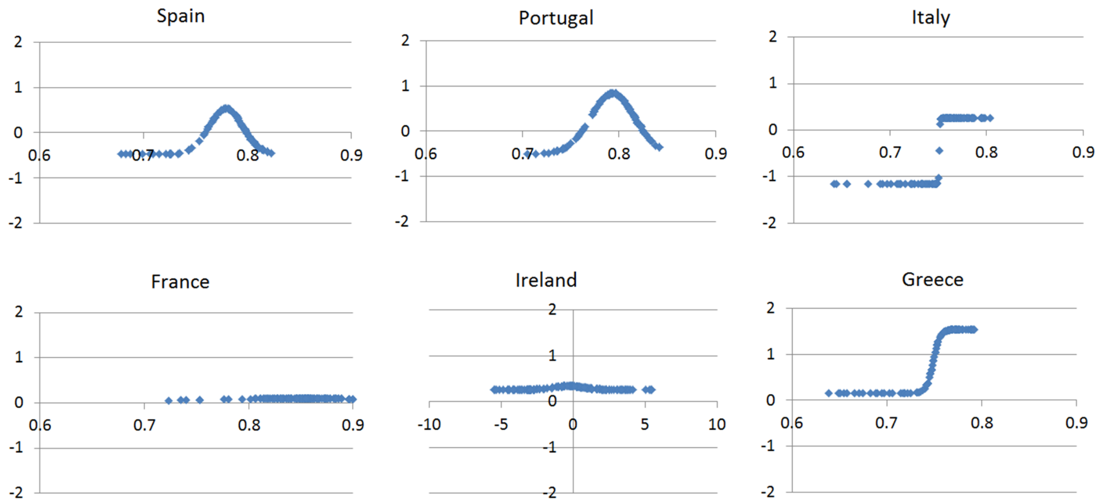


Notes: The figures refer to coefficient β which is depicted on the vertical axis; β is defined as $\beta = \beta_{10} + \beta_{20} \cdot \mathrm{F}(z_{t-j}, \gamma, c)$. The transition variable z_{t-j} is displayed on the horizontal axis.

Figure 8. Estimation with value-added exports

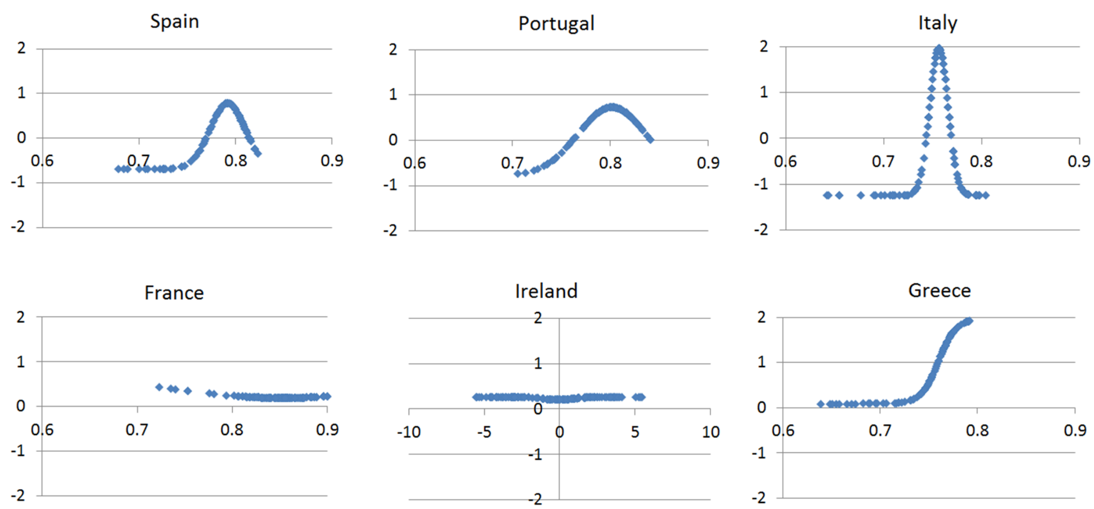
Notes: The figures refer to coefficient β which is depicted on the vertical axis; β is defined as $\beta = \beta_{10} + \beta_{20} \cdot \mathrm{F}(z_{t-j}, \gamma, c)$. The transition variable z_{t-j} is displayed on the horizontal axis.

Figure 9. Estimation with ULC-deflated REER



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

Figure 10. Estimation with median as threshold value



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

5. Conclusions

In this paper, we have analysed the net relationship between domestic demand and export activity for six euro-area countries using non-linear estimations. The results of our macroeconometric 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 relationship 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 relationship 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 are firms 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 efforts at 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 triggered by 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 programmes, 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 markets. 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 any time 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 hypothesised 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 been adopted with the aim of rebalancing the respective economies towards the tradable sector. These policies imply a more structural and sustainable current-account adjustment.

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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**			-0.390*
			(0.33)			(0.23)
β_{23}			0.907			
			(0.64)			
θ_{20}	1.013***	1.027***	0.301	-0.233**	0.776**	-0.200
	(0.38)	(0.23)	(0.25)	(0.09)	(0.33)	(0.40)
θ_{21}		1.026***				
		(0.18)				
μ_{20}	-0.480*	-0.326	-0.391	-0.534***	0.232	-1.807***
	(0.27)	(0.23)	(0.25)	(0.20)	(0.20)	(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	-0.460***	0.638***	-0.373***	-0.242**	0.318*
	(0.12)	(0.16)	(0.10)	(0.14)	(0.10)	(0.17)
η_{22}	-0.079				0.215***	0.156**
	(0.09)				(0.07)	(0.07)
η_{23}					-0.279***	0.292***
					(0.10)	(0.05)
η_{24}					-0.424***	-0.170
					(0.09)	(0.19)
δ_2	-0.224	0.298***	0.110	-0.168***	-0.175***	0.025
	(0.16)	(0.06)	(0.08)	(0.06)	(0.05)	(0.13)
γ	35.566*	49.762***	59.061***	1.638**	1.872**	6.662***
	(18.61)	(19.27)	(20.89)	(0.68)	(0.86)	(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*

Notes: 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*

Notes: 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*

Notes: 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*

Notes: 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

Notes: 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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