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Abstract

We argue that, under certain conditions described by a sunk cost hysteresis model, firms consider exports as a substitute for domestic demand. This is valid also on the macroeconomic level where the switch from the domestic market to the export market and vice versa takes place in a smooth manner. Areas of weak reaction of exports to changes in domestic demand are widened by uncertainty. Our econometric model for six euro area countries suggests domestic demand and capacity constraints as additional variables for export equations. We apply the exponential and logistic variant of a smooth transition regression model and find that domestic demand developments and uncertainty are relevant for short-run export dynamics particularly during more extreme stages of the business cycle. A substitutive relationship between domestic and foreign sales can most clearly be found for France, Greece and Ireland (ESTR model) and France, Portugal and Italy (LSTAR model), providing evidence of the importance of sunk costs and hysteresis in international trade in these EMU member countries. What is more, our empirical results are robust to the inclusion of a variable measuring European policy uncertainty. In some cases (Italy, Greece and Portugal) the results underscore the empirical validity of the export hysteresis under uncertainty model.

JEL-Codes: F140, C220, C500, C510, F100.

Keywords: domestic demand pressure, exports, error-correction models, hysteresis, modelling techniques, smooth-transition models, sunk costs, uncertainty.

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1. Introduction

A number of euro area countries which recorded large current account deficits in the period prior to the European debt and banking crisis starting in 2010 have seen a significant correction of their external imbalances, in particular the trade balance, over recent years. Falling imports have been an important part of this correction due to low domestic demand. However, at the same time, exports and export market shares have been continuously increasing in most of these countries since 2009. Shrinking unit labour costs and falling real effective exchange rates are able to explain only part of the gains in export market shares. Christodoulopoulou and Tkacevs (2014) find that only 60 to 70% of variation in exports can be explained by standard export equations. It thus seems likely that non-price related factors have been important in explaining export performance. The residuals from a standard approach to model exports are potentially consistent with the parallel dramatic fall of domestic demand. A possible relationship between domestic demand and exports could be particularly important in the current economic situation of substantial macroeconomic adjustment needs and very low domestic demand.

The relation between domestic demand and exports is not straightforward and could be either negative (substitutive) or positive (complementary). A recent survey of literature on this topic is presented in Esteves and Rua (2013). Theoretical reasons for a *positive* link between domestic demand and exports may be due to increased efficiency from learning by doing effects (Belke, Goecke and Guenther (2013)) or due to liquidity generated by cash flow from exports which can help overcome liquidity constraints for domestic operations (Berman et al. (2011)). Theory has identified a *negative* relationship between domestic demand and exports mostly at the firm level. Several studies have been concerned with the effects of domestic demand pressure on the inclination and capacity to export. These studies are not numerous, but go back several decades.[1]

The main argument is that - in the short-run - exporting firms face capacity constraints or increasing marginal costs and thus have to substitute sales between their domestic and foreign markets. An increase in demand for exports cannot be satisfied in the short-run as long as capacity is highly utilised and most of production is sold on the domestic market. Conversely, with low domestic demand, for instance during a domestic recession, firms will be able to shift more resources to export activities; to compensate for the decline in domestic sales, firms will increase their efforts to export. Besides pull factors (e.g. foreign demand), export performance can thus also be determined by push factors (such as low capacity utilisation). Besides the studies mentioned above, more recent empirical literature (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)) generally identifies a significant negative effect of domestic demand pressure on exports for several

[1] Examples are Ball et al. (1966), Smyth (1968), Artus (1970, 1973), Dunlevy (1980), Zilberfarb (1980), Faini (1994) and Sharma (2003).

countries, among them the United Kingdom, the United States, Germany, Spain, Israel, Turkey, Morocco and India.

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 exports has often been analyzed in a linear framework. The relationship between domestic demand and export performance may however vary with economic conditions and thus be of a non-linear nature.

Assuming a substitutive relationship between domestic demand and exports, following a domestic demand shock, firms will try to shift sales between the two markets. However, entering the export market or shifting more sales towards it usually implies sunk costs. These are costs firms need to pay that are irreversible ex post (Baldwin and Krugman (1989)) and the significance of this knowledge diminishes rapidly after leaving the export market (Belke, Goecke and Guenther (2013)).

In that respect, we can distinguish two cases. First, with a negative domestic demand shock and sunk costs for entering or shifting to the export market, firms will therefore be reluctant to pay these costs as long as capacity is still relatively highly utilised. Once capacity utilisation falls below a certain threshold, firms might be more willing to pay sunk investment costs as these costs and the effort of selling in the foreign market might be lower than the cost of running excess capacity.[2] Exports in this case can be considered as "survival-driven". Second, following a positive domestic demand shock, firms might not be able to serve both domestic and foreign markets due to highly utilised capacities. If they prefer producing for the domestic market, firms would consider shifting sales to that market once a certain high capacity utilisation threshold has been crossed. With sunk costs, leaving the export market or shifting sales away from it implies that these costs would have to be paid again upon trying to re-enter the export market or re-shifting sales towards it in the future.

Overall, these arguments suggest that only if certain low or high capacity utilisation thresholds have been crossed, firms will change their export behaviour. Only if a domestic demand shock is accompanied by extreme changes in capacity utilisation will firms shift their sales to another market. As long as capacity is utilised to a more normal degree and operates within these lower and upper thresholds, firms are working in a "band of inaction" where sunk costs hinder firms from changing their export behaviour even though

[2] Alternatively, some firms might be constrained by technical limitations that allow production at a certain capacity utilisation rate only; facing a certain low capacity utilisation threshold they might face the decision to either not produce at all or shift their production to serving foreign markets.

capacities might exist for those firms that are not yet very active in foreign markets.[3] This also implies that, once capacity utilisation thresholds have been crossed on either end and firms have shifted sales among markets, they will be reluctant to shift again once capacity returns back to more normal levels. There is thus strong persistence in export behaviour which can be traced back to the theory of hysteresis (Baldwin and Krugman (1989). Export hysteresis is the tendency of a temporary change in export behaviour to become permanent. It is particularly important in the current weak economic situation of several euro area member states; firms increase efforts to shift sales to the export market given weak domestic demand and this might not be a cyclical change but rather a persistent improvement as firms will often decide to stay in the foreign market even once domestic demand picks up again as they are trying to avoid repaying sunk costs.

We thus essentially present a story of dynamic investment in the presence of high fixed cost and capacity constraint. This story is consistent with firms switching from selling in the domestic market to the foreign market as soon as the level of domestic demand falls short of a given trigger threshold. But what about the heterogeneity element that induces switching only by some firms, but not in all firms? Here we refer to Belke and Goecke (2005) who, starting from the idea of a "band of inaction" focus on the issue of aggregation. They are able to derive an aggregation process, considering heterogeneity of sunk exit/entry costs and/or the extent of uncertainty of the future market situation and/or the elasticity of demand. This is resulting in different triggers for different firms. This (realistic) consideration of heterogeneity alters the hysteresis characteristics when aggregating from the microeconomic to the macroeconomic level. Due to heterogeneity in firm characteristics such as the magnitude of sunk costs or the productivity level firms exit (and entry) sequentially and not all in a time from (into the) the market and the resulting aggregated hysteresis loop thus shows no discontinuities. This is rather important in our context because, absent this feature, all firms would switch if there is a large negative domestic demand shock. This would contradict the abundant micro-evidence in the trade literature that actually the most engaged exporters are also faring best on the domestic market.[4]

Notably, in this model of export hysteresis, the band of inaction is widened by uncertainty (Belke and Goecke, 2005). This is because a forward-looking firm considers future effects of a present sunk cost 'investment'. If the exogenous variable demand is stochastic, a real option approach applies (Belke and Goecke (2001), Dixit (1989), Pindyck (1998, 1991)). An inactive firm deciding on a present entry or to stay passive, will include the option to enter later as a potential alternative. Demand which is presently contributing to cover costs, may in a stochastic situation decrease in the future. By staying passive the firm can avoid future losses if this

[3] In the European case and the countries under consideration, potential for shifting production to foreign markets seems to exist. As an example, Esteves and Rua (2013) specify that in 2010, only one third of Portuguese manufacturing firms was exporting and for them the exports to sales ratio was on average around 30 per cent.

[4] The aggregation procedure of firm heterogeneity under consideration is explained in detail in Belke and Goecke (2005), pp. 196-201.

situation will realize. Moreover, an instantaneous entry kills the option to enter later and to “wait-and-see” if the future demand movement will turn out to be (un)favorable. Thus, in a stochastic situation, the sunk costs and, additionally, an option value of waiting have to be covered in order to trigger an entry. Therefore, uncertainty implies an upward shift of the entry trigger demand. The same is valid for the exit trigger demand which will shrink in a situation with uncertainty. Belke and Goecke (2005) show that this line of reasoning is valid also at the macroeconomic aggregated level. Thus, uncertainty leads to a widening of the band of inaction also at the macroeconomic level, aggravating the hysteresis property of the firm's export behavior.[5]

Our paper builds on this sunk-cost hysteresis model and explicitly tests for a short-run non-linear relationship between domestic demand and exports from a macroeconomic perspective. A particular asymmetric effect was already considered in Esteves and Rua (2013) for the case of Portugal. Belke, Oeking and Setzer (2015) consider the relation of domestic demand and export of goods in several euro area countries. Our analysis goes beyond these studies by investigating six euro area countries with significant current account deficits in the pre-crisis period (Spain, Portugal, Italy, France, Ireland and Greece) employing the export of both goods and services.

The focus upon these former current account deficit countries reflects our intent to analyze countries in which firms were experiencing a fall in domestic demand. After the start of EMU, a real appreciation set in for peripheral Euro area member countries, especially in those experiencing a massive housing boom, namely Ireland and Spain. As a consequence, their competitiveness went down further. At the same time nominal long-term interest rates converged among EMU member countries, inducing low real interest rates in the peripheral countries with higher inflation. This in turn stifled spending and inflation even further, leading to growing current account deficits (see, for instance, Krugman, Obstfeld and Melitz (2015), pp. 687ff.). Since currency devaluation was no option for these countries, it became clear that the necessary real exchange rate adjustment implied a period of low inflation or even deflation in combination with significant unemployment and protracted recession including weakness of domestic demand.[6] The “doom loop” among banks and governments contributed significantly to this development.

Moreover, we go beyond the papers mentioned above by thoroughly conducting tests for structural breaks common to the countries under investigation and integrating an uncertainty variable in our estimations.

Following Belke, Oeking and Setzer (2015), we implement a smooth transition regression model such that we can specify aggregated non-linearities with a high degree of flexibility. We argue that the strength of the

[5] The aggregation procedure under consideration of firm heterogeneity and uncertainty is explained in detail in Belke and Goecke (2005), pp. 189-192.

[6] See, for instance, Belke and Gros (2017). But, for example, in Ghironi and Melitz (2005), a negative current account emerges also in times of a positive productivity shock or a reduction of entry barriers. In this case, the home economy experiences the most attractive conditions and becomes nevertheless a net borrower on international markets to finance the creation of new firms that a positive productivity shock or the reduction of competitive barriers warrant. However, this was not the scenario the six EMU member countries in our sample were faced with.

relation between domestic demand and exports depends on capacity constraints and more generally the business cycle. Besides the possibility that substitutability will increase after reaching either the upper or lower threshold (i.e. giving rise to symmetry around the band of inaction), we also allow for the possibility that exports react sharper in a recession than during an economic expansion (giving rise to asymmetry around the band of inaction). This is achieved by relying on either an exponential or logistic variant of smooth transition specification. The aggregation at the macro level allows us to draw results on net effects of capacity utilisation on the economies as a whole. This is of special importance in the discussion of macroeconomic adjustment and the reduction of current account imbalances in the euro area.

The paper proceeds as follows. Taking the simple sunk cost-based hysteresis model as a starting point, we carry out some pre-testing in terms of unit roots and cointegration in section 2. Based on the cointegration results, we set up an error-correction export equation and incorporate non-linearities as suggested by our theoretical considerations. These smooth transition regression models (STR), including several robustness tests among them the incorporation of an uncertainty variable, are estimated in section 3. Section 4 finally concludes and conveys an outlook on further research avenues.

2. Empirical strategy

Data

Our data stems from different sources (cf. Table A1): Data on real exports (x_t) and real domestic demand (dd_t) comes from the national statistical offices (either obtained from Eurostat or Oxford Economics). 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 has been obtained from Eurostat and is an index deflated by consumer price indices with a country's 15 main trading partners (r_t). Alternatively, the same source provides an index deflated by unit labour costs with a country's 24 main trading partners (r_t^{ULC}). Data on foreign demand (y_t^*) from the ECB is based on trade-weighted imports for a country's 15 main trading partners. Capacity utilisation data in the manufacturing industry (z_t) comes from the Business and Consumer Surveys by the European Commission, available from Eurostat or Insee in the case of France. For Ireland, data on capacity utilisation is not available. Instead, we use the output gap (interpolated data from AMECO). As an uncertainty variable for our robustness checks we employ the economic policy uncertainty index relevant for the EU as a whole because the respective index was not available for the individual Euro area member countries for such a long sample period like ours (http://www.policyuncertainty.com/europe_monthly.html). The final data set is quarterly and mostly available from 1980:Q1 to 2012:Q4.

Non-stationarity and cointegration tests

By focusing on the volume of exports for a specific country as our main purpose of this paper it is necessary to specify a function, which depends on foreign demand and the difference in price levels concerning trading partners in the long run. For this purpose, an equation

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

is specified, where x_t is the logarithm of exports, y_t^* the log of foreign demand, r_t the log of the real effective exchange rate and d is a dummy. By taking logarithms on each side of the equation we apply a non-linear framework to short-run-effects. Before applying the Engle-Granger approach (1987) to test for cointegration we need to check whether the variables introduced above are non-stationary. In an augmented Dickey-Fuller-test (ADF-test) we include an intercept for the real effective exchange rate series and an intercept plus a time trend for all other series and specify the auxiliary regression accordingly. In addition, using the two-breaks minimum LM unit root test (Lee and Strazicich (2003)) trend stationarity is established if the null hypothesis is rejected. The ADF-test is complemented by LM unit root tests and corroborates our ADF-results. The results are displayed in Table A2 in the annex.

We adopt a methodology developed by Bai and Perron (1998, 2003) to account for possible instabilities in the long-run coefficients of equation (1). Their basic idea is to choose breakpoints such that the sum of squared residuals for all observations is minimized. The estimated breakpoints by definition represent the linear combination of these segments which achieve a minimum of the sum of squared residuals (Bai and Perron, 2003). Table 1 shows the two most important break points for the six countries analyzed, accompanied by the 95% upper and lower confidence intervals. In case of Spain the first breakpoint occurs in 1993Q4 with 1993Q3 and 1994Q1 providing the 95% confidence intervals.

The results of Table 1 are also useful in the context of unifying our testing and estimation approach. One may ask, for instance, to what extent the differences in the cointegrating test and cointegrating equation estimation results across countries we usually gain for the export equations of the six EMU member countries analyzed here (available on request) are driven by the fact that the specific breakpoints are different across countries?[7] To check this, we follow the option to see what happens when imposing a common break point for all countries. Do the data strongly reject such an assumption? In order to account for this issue, we have implemented one common breakpoint for all countries in all estimations contained in this paper with an eye on the results displayed in Table 1, i.e. by a (permanent) dummy denoting the most common break in 1993:04 which may proxy the fallout of the 1992/93 crisis of the European Monetary System (EMS).[8]

[7] We owe this point to an anonymous referee.

[8] Estimation results for a common dummy denoting the introduction of the Euro are available on request.

Table 1: Breakpoints with lower and upper 95 %

Country	*Break-point*	*Lower 95%*	*Upper 95%*
Spain	1993 Q4	1993 Q3	1994 Q1
	2004 Q1	2004 Q1	2004 Q4
Portugal	1986 Q3	1986 Q1	1987 Q1
	1993 Q4	1993 Q2	1994 Q3
Italy	1993 Q4	1993 Q3	1994 Q4
	1997 Q4	1997 Q3	1998 Q1
France	1983 Q2	1983 Q1	1984 Q1
	1993 Q4	1993 Q3	1994 Q1
Ireland	1993 Q4	1993 Q3	1994 Q1
	2001 Q2	2001 Q1	2001 Q3
Greece	1985 Q3	1985 Q2	1986 Q1
	1996 Q4	1996 Q3	1997 Q1

Notes: The table provides the two most important breakpoints according to the Bai and Perron (1998) methodology for all countries under investigation.

The respective findings (long-term relation and Engle-Granger test for cointegration) based on a common break in the fourth quarter of 1993 are provided in Table 2. Alternative specifications which country-specific break points were contained in the previous version of this paper and are available upon request. To be more concrete, we focus on estimating the long-run equilibrium of equation (1) by FMOLS (Fully Modified Least Squares). To test for cointegration, we apply the Engle-Granger test for cointegration. The results are displayed in the last column with the respective critical values from MacKinnon (1991). Because $\hat{e}_t \sim I(0)$, we can conclude that for each country the error-terms are stationary and a cointegration relationship between the variables is thus present. It should be mentioned that the findings are not greatly affected by the specification of the break points. The error terms from estimations based on common and individual breakpoints turn out to be highly correlated and the short-term findings provided in the following are not affected by these findings.

Table 2: Long-run relationships and Engle-Granger test for cointegration based on a specification with common breaks.

Country	*Long-run relationship*	*Engle-Granger Test*
Spain	$x_t = 4{,}444^{***} - 0{,}189d - 1{,}023^{***}r_0 - 0{,}867^{***}r_1 + 1{,}161^{***}y_0 + 1{,}164^{***}y_1$ (147,62) (-0,56) (-9,80) (-4,17) (9,37) (20,84)	-4.5136***
Portugal	$x_t = 3{,}749^{***} - 0{,}765d - 1{,}075^{***}r_0 - 0{,}280r_1 + 1{,}246^{***}y_0 + 0{,}828^{***}y_1$ (88,38) (-0,92) (-20,52) (-0,59) (21,71) (10,68)	-2.6114*
Italy	$x_t = 4{,}751^{***} + 0{,}347d - 0{,}724^{***}r_0 - 0{,}609^{**}r_1 + 0{,}851^{***}y_0 + 0{,}555^{***}y_1$ (154,77) (0,88) (-17,10) (-2,55) (17,38) (9,93)	-3.9215*
France	$x_t = 4{,}772^{***} + 1{,}492^{***}d - 0{,}523^{***}r_0 - 1{,}230^{***}r_1 + 0{,}604^{***}y_0 + 0{,}629^{***}y_1$ (230,07) (5,13) (-25,93) (-8,74) (26,50) (33,22)	-3.8773*
Ireland	$x_t = 3{,}951^{***} - 1{,}649^{***}d - 1{,}405^{***}r_0 - 0{,}596^{**}r_1 + 1{,}585^{***}y_0 + 1{,}733^{***}y_1$ (69,19) (-3,92) (-19,89) (-2,18) (19,78) (17,69)	-4.0386*
Greece	$x_t = 3{,}501^{***} + 2{,}683^{**}d - 0{,}331^{***}r_0 - 2{,}417^{***}r_1 + 0{,}370^{***}y_0 + 1{,}287^{***}y_1$ (71,39) (2,53) (-6,45) (-3,92) (6,66) (10,55)	-4.5760***

Notes: The final column 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 taken from MacKinnon (1991) for the respective number of variables. */**/*** statistical significance at the 10%/5%/1% level. The common breakpoint is located in 1993:04. d denotes a dummy which takes a value of 1 from 1994:01. For the regressors r and y, 0 denotes the first and 1 the second subperiod. The number of observations is 131 (Portugal, France, Ireland), 128 (Italy), 111 (Spain) and 105 (Greece).

The sign of the estimated coefficients (negative for our exchange rate variable and positive for the foreign demand variable) overall correspond with our priors from theory. The effects are in line with theory. As an example, the effect of y (exp) is always positive. Our estimation results for the long-run relations largely match those of other studies; both in terms of sign and size of the coefficients (see, for instance, European Commission (2011)). We do not come up with a more detailed analysis here as our main focus is on the short-run relation and slightly different long-run specifications did not change the following results in a noteworthy way.[9]

[9] As robustness checks, we also included additional variables in the long-run relation, e.g. trade openness, or restricted the coefficient for foreign demand to unity. Other non-price competitiveness variables could also have an influence on exports. As Esteves and Rua (2013) point out, the long-run results need to be interpreted with caution, as further structural breaks or these potential omitted variables could have an influence on the outcomes. Since our focus is on the

Empirical model

As explained above, we apply a non-linear framework to capture any short-run non-linear impact in the relation between domestic demand and exports regarding the state of the economy. We consider each country's economic condition by looking at deviations of its capacity utilisation from its mean. Looking at short-run adjustments and in particular at the short-run relation between exports and domestic demand, we take into account the long-run equilibrium estimated above. For this purpose, we apply an error-correction model. As already mentioned in the introduction, 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. Based on an economy's export performance where individual firm level decisions are aggregated, it may not seem adequate 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 when the timing of the regime switch varies on an aggregated level.

According to Engle and Granger (1987), for every (long-run) cointegration model an error-correction model describes the short-run dynamics of the system. Our main interest is in parameter β, the short-run elasticity of exports to a change in domestic demand concerning the state of economy, looking at the capacity utilizations and especially its deviations from its mean (z_t). The long-run-equilibrium (1) takes the possibility into account that a non-linear-adjustment process leads, depending on z_t, to the long-run equilibrium. The error-correction model (see eq. 2 below) derived from eq. (1) can best be modelled as a smooth-transition regression (STR). We will therefore estimate the following error-correction model with non-linear short-run adjustment in STR form:

$$\Delta x_t = \left[\alpha_1 + \sum_{i=0}^{n-1} \beta_{1i}\Delta dd_{t-i} + \sum_{i=0}^{n-1} \theta_{1i}\Delta y^*_{t-i} + \sum_{i=0}^{n-1} \mu_{1i}\Delta r_{t-i} + \sum_{i=1}^{n-1} \eta_{1i}\Delta x_{t-i} + \delta_1 \hat{\varepsilon}_{t-1}\right] +$$

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

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

as a non-linear short-run STR-model which includes gradual regime changes when the timing of the regime switch varies on an aggregated level. Δx_t represents a function of lagged 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 in 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

short-run results and the short-run non-linear estimation appear to be relatively insensitive to slightly different long-run specifications.

parameters $\alpha, \beta, \theta, \mu, \eta$ are the short-run effects. Our main parameter of interest is β, the short-run 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 given in the introduction above, domestic demand should enter our estimations in the short-run only.[10] In contrast to the long-run estimation, we do not include a structural break in the short-run estimations of equation (2) because this specification already includes the smooth transition non-linearities. Furthermore, a break in the long-run relation does not imply that short-run dynamics change as well; by excluding these breaks we are also able to reduce our model's complexity.

The first set of brackets in equation (2) is a standard linear error-correction model. Non-linearity is introduced via the second set of brackets which includes the same regressors, but is multiplied with the transition function $F(z_{t-j}, \gamma, c)$. The transition function in a STR model is a smooth, continuous and bounded function between 0 and 1. We consider two popular forms of smooth transition models based on the transition function. These are the *logistic* STR model (LSTR) and *exponential* STR model (ESTR). The LSTR model uses a *logistic* transition function of the following form:

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

with the transition variable z distinguishing different regimes in our non-linear estimation. 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.[11] Smoothness parameter γ determines strength and speed 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 follow Teräsvirta (1998) and normalize it by σ_z in order to be scale-free).

The logistic transition function increases monotonically from 0 to 1 as the value of transition variable z increases. We can therefore distinguish two different regimes in the extreme and a gradual transition between these two: (i) negative deviations of the transition variable from its threshold: $\lim_{z_{t-j}\to-\infty} F(z_{t-j}, \gamma, c) = 0$, when the model collapses to just the first set of brackets in equation (2), i.e. the linear part, and (ii) positive deviations of the transition variable from its threshold: $\lim_{z_{t-j}\to+\infty} F(z_{t-j}, \gamma, c) = 1$. The coefficients $\alpha, \beta, \theta, \mu, \eta, \delta$ gradually change between these two extreme values with changing z_{t-j}.

[10] As a robustness test, we also included domestic demand in the above long-run cointegration relation. Its coefficient did neither turn out to be statistically significant nor did it help to constitute a better long-run relation.

[11] As a robustness check, we also apply the same estimations by looking at deviations of z from its mean value. Final results remain similar. Results are available from the authors upon request.

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 mean).

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

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

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

In our case, the ESTR model represents the hypothesis of symmetric hysteresis in exports. Here, both positive and negative deviations of the threshold variable capacity utilisation from its average value c matter. This implies that as long as the deviation of the transitional variable 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.

The main difference between these two forms of non-linear error-correction model refers to different deviations of the transition variable from its threshold value (its mean): the LSTR case distinguishes positive vs. negative deviations and the ESTR model large vs. small deviations from equilibrium. The former will therefore yield asymmetric results around the threshold, and the latter symmetric deviations above or below the threshold.

3. Empirical results

Specification tests

The modeling cycle for smooth transition models suggested by Teräsvirta (1994) starts with a test for nonlinearity. The null hypothesis of linearity can be expressed as either H_0: $\gamma = 0$, or H_0: $\beta_1 = \beta_2$. However, both γ and β_2 are unidentified under the null hypothesis. Consequently, standard asymptotics cannot be applied because of the existence of nuisance parameters (Van Dijk, Teräsvirta and Franses, 2002). To overcome this, Teräsvirta (1994) suggests an approximation of the transition function by a third-order Taylor approximation. Thus, the corresponding Lagrange multiplier (LM) test for linearity introduced by Luukkonen et al. (1988) can be expressed as:[12]

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

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).

The null hypothesis, which refers to the linear model being adequate, is tested as H_0: $\varphi_i = 0$ with $i = 2,3,4$ against the alternative H_1 where at least one $\varphi_i \neq 0$, implying that the higher order terms are significant (Teräsvirta, 1998). The test statistic has an χ^2 distribution with three degrees of freedom.[13] This procedure also enables the choice of an adequate transition variable. In the case of the linearity hypothesis being rejected, a method for choosing the latter lies in computing the test statistic for several transition functions, i.e. different values of the lag order j, and selecting the configuration for which its value is maximized (van Dijk et al., 2002). Teräsvirta (1994, 1998) has proven that this procedure is adequate.

According to Granger and Teräsvirta (1993), Teräsvirta (1994, 1998) as well as van Dijk, Teräsvirta and Franses (2002), the LM testing procedure described above can also be applied to *distinguish between an exponential and a logistic transition function* and thus, to choose the appropriate specification. If the linearity null has been rejected, equation (6) is used to test the following hypotheses successively

$$\begin{aligned} &\boldsymbol{H_{04}}\text{: } \boldsymbol{\varphi_4 = 0}, \\ &\boldsymbol{H_{03}}\text{: } \boldsymbol{\varphi_3 = 0 \mid \varphi_4 = 0}, \\ &\boldsymbol{H_{02}}\text{: } \boldsymbol{\varphi_2 = 0 \mid \varphi_3 = \varphi_4 = 0}. \end{aligned} \qquad (7)$$

The decision rule to select the most adequate transition function introduced by Teräsvirta (1994) is as follows. If the rejection of $\boldsymbol{H_{03}}$ is the strongest one in terms of lowest p-value or largest test statistic, respectively, the ESTR model should be chosen, otherwise the LSTR model should be preferred.[14] Table 3 displays the empirical realizations of the nonlinearity test statistics, while Table 4 provides the test statistic to distinguish between both configurations.

The common procedure behind selecting the lag length of the transition variable in the Teräsvirta testing and modelling cycle intuitively seems to pick the lags for which the chance to observe non-linearity is strongest (Belke, Oeking and Setzer, 2015). However, this would seem to artificially favor our prior which is to find non-

[12] In the case of small samples in combination with a large number of explanatory variables, F-versions of the LM test statistics are preferable, as they have better size properties (Granger and Teräsvirta, 1993, Teräsvirta, 1998, and van Dijk, Teräsvirta and Franses, 2002).

[13] The number of degrees of freedom $3p$ refers to the number of regressors p. Furthermore, the test assumes that all regressors, as well as the transition variable z_t, are stationary and uncorrelated with the error in eq. (4) u_{t+k} (Teräsvirta, 1998).

[14] See Granger and Teräsvirta (1993) or Teräsvirta (1994) for details.

linearity in the data.[15] In order to react fully as possible to this important caveat we provide findings from Table 3 on where a common lag order is just imposed for all countries to make the results more comparable.

Table 3: Teräsvirta test for non-linearity

	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
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]
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]
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]
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]
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]
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]

Notes: The 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. The tests are conducted based on specifications with a common breakpoint in 1993, fourth quarter. j denotes the lag length.

Table 4: Test for the appropriate specification

Country	lags	(ii)	(iii)	(iv)
Spain	4	48.32 (0.000)	47.97 (0.000)	43.52 (0.000)
Portugal	6	47.66 (0.000)	5.89 (0.435)	18.02 (0.006)
Italy	6	47.11 (0.000)	28.36 (0.001)	8.29 (0.405)
France	1	12.20 (0.032)	11.76 (0.038)	5.53 (0.355)
Ireland	4	50.42 (0.000)	16.70 (0.054)	32.79 (0.000)
Greece	2	72.42 (0.000)	54.98 (0.000)	70.47 (0.000)

Notes: For the Teräsvirta test, χ^2 test statistic realizations are displayed with p-values in parentheses. The test is based on a long-run specification with a common breakpoint. The choice is made for an individual lag length for each country. As mentioned in the text, however, the lag length is unified by us for the following estimations and both LSTR and ESTR are estimated. (ii), (iii) and (iv) refer to H_{02}, H_{03} and H_{04} in equation (7) respectively.

[15] We owe this caveat to an anonymous referee. The results for potentially different country-specific lag lengths are contained in the previous version of this paper and are available on request.

The findings in Table 3 show that nonlinearity is essentially never rejected for all lag orders. Table 4 shows that a distinction between both model configurations turns out to be difficult. Teräsvirta (1998) suggests estimating different models and choosing between the different specifications and different lag lengths only during evaluation of the estimation results. LSTR and ESTR models generally form very close substitutes. Tests as the ones above should thus be seen as a starting point for estimation instead of providing clear-cut outcome at this early stage of analysis. Taking the ambiguous findings into account, we therefore estimate both LSTR and ESTR models for all countries. To allow for a direct comparison of our findings*, we always use a unified lag length of 2 for our transition function* in our study.

Estimation

To evaluate our parameters, we estimate equation (2) with non-linear least squares (NLS). Our main coefficient of interest β depends on the transition function $\mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c})$ as depicted in either equation (4) or (5). To choose the final specifications, we examine our estimation results by simple judgment regarding the plausibility of the parameter values and the regimes which the models imply, the models' convergence properties, goodness of fit measures and a test of no residual autocorrelation. For this misspecification test we apply a variant of the Breusch-Godfrey Lagrange Multiplier (BG) test suitable for non-linear estimation as suggested in Teräsvirta (1998). The test's null hypothesis is that there is no p^{th} order serial correlation in our residuals u_t. The test regresses $\tilde{u}_t$ (the estimated residuals) on $\tilde{u}_{t-1}, \dots, \tilde{u}_{t-p}$ and the partial derivatives of the regression function with respect to γ.

Estimation results are found in Table 5 for countries with an ESTR specification and in Table 6 for countries with an LSTR specification. Our theoretical priors suggest a negative coefficient for the coefficient β, i.e. a substitution effect from domestic demand to exports during times of low or high capacity utilisation. When estimating the ESTR model, coefficient β_{1i} for $\mathrm{F}\big(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c}\big) = 0$ (i.e. the linear model) shows us results for capacity utilisation levels around the threshold level. The joint coefficient $\beta_{1i}+\beta_{2i}$ for the case when $\mathrm{F}\big(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c}\big) = 1$ yields the results for positive and negative deviations from our threshold. In the LSTR case, β_{1i} represents low levels of capacity utilisation and $\beta_{1i}+\beta_{2i}$ high values of capacity utilisation.

Estimation Results

Let us first turn to the econometric specification based on an ESTR model (Table 5). For France, Greece and Ireland, the effects for one or two lags display negative values for extreme levels of past capacity utilisation while negative contemporaneous effects are not identified except for the case of Italy. The contemporaneous coefficient is positive for Ireland; for Italy the results are ambiguous. 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. A positive coefficient may imply that 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 (Belke, Oeking and Setzer (2015), and Berman, Berthou, and Héricourt (2011)). As argued above, also 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.

Table 5: Estimation results for domestic demand effects on exports based on ESTR specification

		Spain	Portugal	Italy	France	Greece	Ireland
Contemporaneous coefficients							
Domestic demand 1st regime (β_{10})		0.7377 (1.6703)	-0.1910 (-1.3809)	1.0906*** (4.5453)	-0.3952 (-1.2791)	-0.0258 (-0.3100)	0.4505*** (3.9030)
Domestic demand 2nd regime (β_{20})		-0.4091 (-0.3879)	0.3087** (2.6784)	-0.9441*** (-3.1921)	0.9451 (1.5624)	-0.2483 (-1.0574)	3.6635 (0.3619)
Sum	$\beta_{10} + \beta_{20}$	0.3286	0.1177	0.1465	0.5499	-0.2741	4.1140
Lagged coefficients with 1 lag							
Domestic demand 1st regime (β_{11})		-0.7260 (-1.4029)	-0.5166*** (-5.4342)	-1.8717** (-2.9751)	2.4935** (2.0281)	0.5064** (2.6547)	-0.0039 (-0.0767)
Domestic demand 2nd regime (β_{21})		1.3413** (2.9055)	1.0025*** (8.1215)	2.4579*** (4.4976)	-2.2968 (1.5967)	-0.7399** (-2.6968)	-0.8538** (-3.0011)
Sum	$\beta_{11} + \beta_{21}$	0.6153	1.5191	0.586	0.197	-0.2034	-0.8577
Lagged coefficients with 2 lags							
Domestic demand 1st regime (β_{12})		0.0018 (0.0051)	0.0819 (1.2300)	0.4710 (1.1663)	1.4331** (2.1650)	0.6935*** (6.7361)	-0.0359 (-0.3502)
Domestic demand 2nd regime (β_{22})		-0.0572 (-0.5623)	0.4357*** (5.1060)	-0.7331** (-2.1199)	-1.7396** (-3.0831)	0.0641 (0.1176)	0.6747** (2.5383)
Sum	$\beta_{12} + \beta_{22}$	-0.0554	0.5176	-0.2621	-0.3065	0.7576	0.6388
	γ cont	1.0319 (1.5611)	2.8357** (2.0973)	5.0390** (2.6401)	2.3023** (2.0015)	2.5524*** (4.8027)	0.0205 (0.3239)
	γ 1 lag	3.2174 (0.9334)	2.8703** (2.1083)	13.9106*** (3.9557)	36.0609 (0.2189)	2.5483*** (5.2896)	0.1179 (1.4186)
	γ 2 lags	-0.4824 (-1.5965)	2.4393** (2.2700)	4.8830 (1.2392)	2.8426*** (3.7361)	3.6794*** (4.9411)	8.4277*** (4.3715)

Notes: Coefficients estimated by NLS; 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 $p = 4$.
β_{ji} $(j = 1{,}2)$ is the coefficient for domestic demand in the non-linear error correction model. The two extreme regimes are $\mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c}) = 0$ given by β_{1i} (i.e. for the ESTR model around the threshold value) and $\mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c}) = 1$ given by $\beta_{1i}+\beta_{2i}$ (i.e. for the ESTR model for large deviations from threshold). Numbers of observations: Italy (125), Spain (100), Greece (95), Portugal (104), France (124), Ireland (62).

We now turn to our findings based on the LSTR specification (Table 6). The contemporaneous substitution coefficient is positive for France and Ireland but insignificant for the other countries in the first regime (beta 0, business cycle trough). For Portugal and France, the substitution coefficient becomes negative in a boom (which is reflected by the sum of both coefficients). While there is hardly any significance of the coefficients for a lag of two quarters, we find a positive coefficient for Spain and Portugal in case of negative capacity utilisation (trough) and a negative one for Italy. However, the sum of both coefficients becomes positive for Italy, France and Greece in case of positive capacity utilisation (boom).

Table 6: Estimation results for domestic demand effects on exports based on LSTR specification

	Spain	Portugal	Italy	France	Greece	Ireland
Contemporaneous coefficients						
Domestic demand 1st regime (β_{10})	0.5842	-0.0493	0.1413	0.5018***	-0.2646	0.5059**
	(1.0028)	(-0.3718)	(0.5148)	(4.0670)	(-1.1375)	(2.6860)
Domestic demand 2nd regime (β_{20})	-0.2537	-0.4210***	1.1692**	-0.8151**	0.3259	0.0707
	(-0.3407)	(-3.2759)	(2.4850)	(-1.9890)	(0.4689)	(0.1721)
Sum $\beta_{10}+\beta_{20}$	0.3305	-0.4703	1.3105	-0.3133	0.0613	0.5766
Lagged coefficients with 1 lag						
Domestic demand 1st regime (β_{11})	0.6569***	0.1893**	-1.7207***	-0.2950**	-0.1988	
	(4.4806)	(2.0073)	(-4.3676)	(-2.0709)	(-1.7014)	
Domestic demand 2nd regime (β_{21})	-0.3124**	-0.2388	3.9681***	1.2857***	0.9990**	
	(-1.9651)	(-1.3240)	(4.4281)	(3.7642)	(2.4368)	
Sum $\beta_{11}+\beta_{21}$	0.3445	-0.0495	2.2474	0.9907	0.8002	
Lagged coefficients with 2 lags						
Domestic demand 1st regime (β_{12})	-0.4857	0.0143	0.0291	0.8431	0.8653*	
	(-1.6104)	(0.0786)	(0.1011)	(1.5604)	(1.8661)	
Domestic demand 2nd regime (β_{22})	1.2434***	0.5524**	0.7532	-1.1847*	-0.1405	
	(4.5419)	(2.3334)	(0.8090)	(-1.9363)	(-0.2531)	
Sum $\beta_{12}+\beta_{22}$	0.7577	0.5667	0.7823	-0.3416	0.7248	
γ cont	2.2329	4.1112**	4.0047**	36.0609	2.0200***	0.5575
	(1.7211)	(2.5128)	(2.9547)	(0.2189)	(3.3607)	(1.1220)
γ 1 lag	9.6231	4.1054**	0.8425***	2.2374	2.1150***	
	(1.6827)	(2.6452)	(3.2729)	(0.9153)	(8.2651)	
γ 2 lags	5.4792**	2.4577	1.7537***	7.2044	11.1546***	
	(2.2027)	(1.3123)	(6.0526)	(0.6369)	(6.4563)	

*Notes: Coefficients estimated by NLS; 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 $p=4$.*
β_{ji} $(j=1,2)$ 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_{1i} (i.e. for the ESTR model around the threshold value) and $F(z_{t-j},\gamma,c)=1$ given by $\beta_{1i}+\beta_{2i}$ (i.e. for the ESTR model for large deviations from threshold). Numbers of observations: Italy (125), Spain (100), Greece (95), Portugal (104), France (124), Ireland (64).

Overall, our empirical results presented in Tables 5 and 6 suggest that the relationship between domestic sales and exports depends on capacity utilisation and the business cycle. Evidence of a substitutive relationship between domestic and foreign sales varies among countries and with different lag lengths.[16] The findings are broadly in line with the gain in export market shares in several euro area (debt and banking) crisis countries during the subsequent 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.

Seen on the whole, thus, more analytical rigour by imposing a common breakpoint in all country-specific estimations and tests and common lags for the transition variable comes at a “cost” in terms of less adaptation of country specifics and thus economically plausible results.[17]

Robustness check: taking stock of political uncertainty

In the following, we are reporting the results of an important robustness check of our estimations.[18] As a final check, we included *economic policy uncertainty* in our analysis, a variable playing a prominent role in explaining band of inactions in the reaction of exports to its main drivers (see, for instance, Belke and Kronen, 2016). More precisely, we rely on the change of European policy uncertainty in period t-1 (i.e. lagged one quarter) as a transition variable in equation (3). As derived in section 1, the empirical results of the hysteretic export equation may turn out to be more pronounced because the band of inaction gets larger with increasing uncertainty. The findings for the LSTR model are given in Table 7.

[17] More evidence in favour of „substitutability“ for at least four countries in our sample (Spain, Portugal, Ireland and Greece) is gained if one admits different breakpoints for different countries. The results are available on request.

[18] Belke, Oeking and Setzer (2015) present results in a framework comparable to ours but using export goods (a measure which is more closely related to capacity utilization but is unfortunately not available for the sample period used by us) only and yield similar results. We also performed additional robustness tests by using different types of real effective exchange rates (deflated by unit labor costs and deflated by consumer price indices) and using the median instead of the mean value as threshold. The results are available upon request.

Table 7: Estimation results for domestic demand effects on exports with uncertainty – LSTR specification

Country	**Spain**	**Portugal**	**Italy**	**France**	**Ireland**	**Greece**
Lagged coefficients with 1 lag						
Domestic demand 1st regime (β_0)	0.1170	0.4589*	0.9710***	1.0858		-0.0183
	(0.5968)	(2.4077)	(5.8344)	(1.9293)		(-0.3000)
Domestic demand 2nd regime (β_{01})	0.4703*	-0.2768	-1.5012***	-0.6302		-0.4811***
	(1.8622)	(-0.5282)	(-3.9177)	(-1.0603)		(-5.2344)
Lagged coefficients with 2 lags						
Domestic demand 1st regime (β_0)	0.8381***	0.7809	0.0475	0.9144**	0.0063	0.8299**
	(8.0184)	(7.2030)	(0.1425)	(2.616)	(0.0324)	(4.6914)
Domestic demand 2nd regime (β_{01})	0.0809	-0.61220**	0.6130	0.5287	0.0345	0.0091
	(0.5704)	(-4.0477)	(1.5312)	(0.9899)	(0.1491)	(0.0556)

Notes: Coefficients estimated by NLS; 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 $p = 4$.
β_{ji} $(j = 1{,}2)$ is the coefficient for domestic demand in the non-linear error-correction model. The two extreme regimes are $\mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c}) = 0$ given by β_{1i} (i.e. for the ESTR model around the threshold value) and $\mathrm{F}(\mathrm{z}_{\mathrm{t-j}}, \gamma, \mathrm{c}) = 1$ given by $\beta_{1i}+\beta_{2i}$ (i.e. for the ESTR model for large deviations from threshold). *Numbers of observations: Italy (61), Spain (61), Greece (61), Portugal (61), France (62), Ireland (64).*

For the following interpretation of the estimation results, we have to keep in mind that the first coefficient holds if economic policy uncertainty decreases while the sum of both coefficients is relevant for an increase in economic policy uncertainty. This interpretation is based on equation (3) for both values of the transition function. The evolution of the transition function reflects the band of inaction. The transition function increases from 0 to 1 if economic policy uncertainty increases. The second coefficient in Table 7 always provides the additional effect once the transition function increases from 0 to 1. Hence, the overall effect for the highest increase in uncertainty is given by the sum of both coefficients which reflects the case where the transition function is 1. We focus on two potential effects in the following: the effect on the substitution coefficient and the one on the global demand coefficient. In each case, we analyze the impact with a delay of the regressor of one or two lags. Due to the rich number of models estimated, several coefficients are as usual insignificant. Nevertheless, a few important results are worth mentioning.

The original effect of domestic sales on exports is always either positive or insignificant in the regime with a decrease in uncertainty. However, the effect for an increase in uncertainty (measured by the sum of both coefficients) becomes negative for Italy (1 lag) and Greece (1 lag) and is strongly reduced for Portugal (2 lags).

This points to a substitution effect as a result of higher uncertainty. Interestingly, the sum of both coefficient becomes positive for Spain. These results broadly confirm robustness of the results gained before with respect to the inclusion of a formerly omitted variable "policy uncertainty". In other words, there is no omitted variable bias in our case. On the contrary, our model of export hysteresis presented in section 1 is corroborated for some EMU member countries such as Italy, Greece and, a bit less so, also for Portugal. The "non-case" of Ireland 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 (see, for example, Belke, Oeking and Setzer, 2015).

For information only: differences between the two regimes are also observed with regard to the effect of world demand on exports. The effect of world demand on exports of France and Greece is negative in case of an increase in policy uncertainty. The opposite is observed for Portugal and Spain where the effects on exports increase in case of higher uncertainty. While no effect is observed for Italy, the specific effect on Greece depends on the lag order but the sum of coefficients for higher uncertainty is always significant. A useful extension for further research will be the consideration of country-specific economic policy uncertainty indices. However, these are not available over the full sample (see www.policyuncertainty.com).

4. Conclusions

In this paper, we have analyzed the relation between domestic demand and exports for six euro area countries using non-linear smooth transition estimations faced with a *strong a priori restriction of common breakpoints and common lags across individual country specifications*. In order to illustrate the results gained in this paper, it seems worthwhile to contrast them with those identified by us without these a priori restrictions.

Our empirical results based on individual and potentially different breakpoint specifications which have been gained in a *previous* version of this paper (available on request) clearly indicated that domestic demand behaviour is relevant for the short-run dynamics of several euro area member countries' exports. The estimation results suggested that on an aggregated level, contemporary and lagged domestic demand developments can affect a country's export performance significantly.

We found that in the cases of Spain, Portugal and Italy, the symmetric non-linearity of the relation manifests itself in a contemporary substitutive relationship between domestic demand and export activity if deviations from average capacity utilisation are large. This is somewhat independent of their sign, but we found stronger evidence for notably low levels of capacity utilisation. 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. During periods of more average levels of capacity utilisation, our empirical evidence pointed to 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. For France, the evidence for non-linearity was weaker. We found evidence of mostly complementary relationships. In the cases of Ireland and Greece, we detected an asymmetric non-linear relationship between domestic demand and exports. Domestic demand and exports are slightly substitutive during a business cycle trough and complements during normal times and in a boom.

Overall, our results mostly confirmed the short-run non-linear relationship between domestic and foreign sales depending on capacity constraints. A substitutive relationship with low capacity utilisation shows up most clearly for Spain, Portugal and Italy.

We also provided first ideas for why 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). However, deriving more detailed explanations for these heterogeneous results for some countries in our sample provide an interesting area for future research. A further interesting avenue could lead to a more disaggregated, sectoral analysis to understand the underlying firm behaviour in more detail (Esteves and Prades (2017)).

A final interesting avenue was taken in this paper: we conducted all necessary estimations and tests based on a common breakpoint implementation in order not to bias the results into the direction of the empirical model with the highest degree of non-linearity and on common lags for all countries in order to avoid the impression that the country-specific regression models were over-fitted till significance. The pattern of the results changed as follows. A substitutive relationship between domestic and foreign sales can now most clearly be found for France, Greece and Ireland (ESTR model) and France, Portugal and Italy (LSTAR model), providing evidence of the importance of sunk costs and hysteresis in international trade in these EMU member countries.

What is more, our empirical results are robust to the inclusion of a variable measuring European policy uncertainty. In some cases (Italy, Greece and Portugal) the results underscore the empirical validity of the export hysteresis under uncertainty model. While we do not feel legitimised to go more deeply into economic interpretations of the country-specific results due to the pronounced ex ante restrictions such as the imposition of common breakpoints and of common lags across all country-specific empirical models, we would like to stress the finding of a general non-linear pattern of export activity of the Euro area member countries with a remarkable goodness-of-fit.

Seen on the whole, the macroeconomic perspective is able to offer insights on overall adjustment effects for euro area countries with previous imbalances. In recent years, the six countries under consideration which recorded large current account deficits before the European debt and banking crisis starting in 2010 have seen a significant correction of their external imbalances. This holds in particular for their trade balances, and exports have been a key adjustment factor. Our results provide one explanation for the rising exports besides standard competitiveness arguments; the observed increase in export market shares accompanying the reduction of the current account deficits could be due to non-price related factors, such a low domestic demand leading to survival-driven exports, instead of an increase in price competitiveness as expected by sustainable structural reforms. This argument appears to be especially relevant in the current period for the countries under consideration in which their capacities have been utilised only to a low degree and domestic demand has fallen strongly. Low domestic demand then did not only affect imports, but at the same time exports and has thus strongly contributed to the external adjustment.

Regarding policy implications, our findings provide important insights for the discussion of macroeconomic adjustment and the reduction of imbalances in the euro area. Prima facie, our results for specific countries could suggest that domestic demand and exports are negatively related only in the short-run, triggered by current economic conditions. To the extent that the closure of the output gap is driven by a pick-up in domestic demand, a lot of the gains in export market shares of vulnerable euro area countries could be lost in the long-run. In such a scenario, analyses of cyclically adjusted current account balances could possibly overestimate the structural adjustment to the degree that weak domestic economic conditions impact not only the import side of the net trade equation, but also 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, domestic demand conditions in peripheral economies are likely to remain depressed as long as the debt burden of both private and public sector remains high. An extended period of deleveraging pressure increases the chances that the reallocation of resources to the export sector will also be more permanent, possibly also fostering increased export-oriented foreign direct investment into distribution networks and other hedging activities (Belke, Goecke and Guenther (2013)). This would make the hypothesized substitutive relationship between domestic demand and exports more long-run. Second, our sunk-cost hysteresis model suggests that once domestic producers have paid sunk costs for shifting production to exports, they remain in a band of inaction even as the business cycle improves. Reversing export market participation should not be expected as long as there are capacities for serving both domestic and foreign market. Third, with increasing exports today and a pick-up in domestic demand in the future, a complementary relation between domestic sales and exports might develop in the long-run due to improvements in efficiency encouraged by learning-by-doing effects. In conclusion, the export increase could therefore be lasting and a substantial part of the gains in export market shares may not only a cyclical phenomenon, but indeed be of a more structural nature.

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 (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	Federal Reserve board	gap between actual GDP and potential GDP as percentage of potential GDP	IE: 1980Q1 – 2012Q4 FR: 1980Q1 – 2012Q4
Policy Uncertainty	www.policyuncertainty.com	Newspaper based uncertainty index	1987Q1 – 2012Q4

Table A2: 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: the 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).

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