Are Agri-food Exports any Special? Exchange Rate Nonlinearities in European Exports to the US

Sind Agrar- und Lebensmittelexporte besonders? Wechselkurs-nichtlinearitäten bei europäischen Exporten in den USA

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Abstract

Using aggregated EMU exports to the US as an example, VERHEYEN (2013a) showed, that in the long run exports react to exchange rate changes in a nonlinear way. This paper tests whether this holds true for agri-food exports as well. To address this question a partial sum decomposition approach and the NARDL framework of SHIN et al. (2013) is applied to the aggregated agri-food exports of eleven European countries to the US, which is currently the major partner of the EU in agricultural trade. The outcomes suggest, that the exchange rate nonlinearities are even more pronounced in agri-food than in total exports. European exporters seem to benefit more from Euro depreciation, than its appreciation harms them. These findings might be interpreted as a sign of pricing strategies application (e.g., pricing-to-market) to the European agri-food exports.

Key Words

agri-food exports; asymmetry; exchange rate nonlinearity; export demand; NARDL

Zusammenfassung

Am Beispiel aggregierter Exporte aus dem Euroraum in die USA zeigte VERHEYEN (2013a), dass Exporte in der langen Frist in nichtlinearer Weise auf Veränderungen des Wechselkurses reagieren. In diesem Artikel wird getestet, ob dies auch für Agrar- und Lebensmittelexporte gilt. Um diese Fragestellung zu untersuchen, wird der Teilsummenzerlegungsansatz und das NARDL-Modell von SHIN et al. (2013) auf die aggregierten Agrar- und Lebensmittelexporte elf europäischer Länder in die USA, die derzeit der wichtigste Partner der EU im Agraraußenhandel sind, angewendet. Die Ergebnisse deuten darauf hin, dass Wechselkursnichtlinearitäten für Agrarexporte sogar noch ausgeprägter sind als für die gesamten Exporte. Europäischen Exporteuren scheinen Abwertungen des Euro in stärkerem Maße zu nutzen, als ihnen Aufwertungen schaden. Dies lässt sich als Anzeichen für die Anwendung von Preissetzungsstrategien (z. B. Pricing-to-Market) bei europäischen Agrar- und Lebensmittelexporten interpretieren.

Schlüsselwörter

Agrar- und Lebensmittelexporte, Asymmetrie, Wechselkursnichtlinearitäten, Exportnachfrage, NARDL

1 Introduction

First studies on trade elasticities were published about seventy years ago (e.g. ADLER, 1945). Since then, a vast empirical literature on determinants of trade flows has emerged. Besides the geographic focus and difference in the estimation techniques, those studies typically varied in their way of including the exchange rate into the model. Some of them focused on income and relative prices as the main drivers of trade (e.g. MARQUEZ, 1990; CAPORALE and CHUI, 1999; NARAYAN and NARAYAN, 2005) and ignored the influence of the exchange rate. ORCUTT (1950) was the first to show that trade flows respond differently to changes in relative prices and exchange rates. Later BAHMANI-OKOOEE (1986), BAHMANI-OSKOOEE and KARA (2003 and 2008), RAO and SINGH (2005), KUMAR (2009) and VERHEYEN (2013a and 2013b) among the others argued, that exchange rates belong to major determinants of trade and hence cannot be excluded from the model.

An alternative framework to study trade determinants is the gravity equation. The first empirical gravity model was introduced by TINBERGEN (1962). Later, the gravity equation was derived also theoretically from different models of international trade (e.g. ANDERSON, 1979; BERGSTRAND, 1990; DEARDORFF, 1998) and became a widely accepted tool of trade analysis. Empirical gravity studies focused mostly on standard determinants of bilateral trade flows, including the distance, income, openness and socio-cultural factors. Exchange rates traditionally were not considered in a standard gravity model due to lack of theoretical rationale for their inclusion (e.g. ANDERSON and VAN WINCOOP, 2003). And even though some empirical studies included exchange rates in their analysis (e.g. MARTINEZ-ZARZOSO and NOWAK-LEHMANN, 2003), it was ANDERSON et al. (2013) who first showed, that the exchange rate needs to be modelled within a gravity framework, once the exchange rate pass-through is (expected to be) incomplete.

Although the investigation of trade determinants and trade elasticities has been playing an important role in international economics for many decades now, the question of possible nonlinearities in international trade stayed unaddressed till the end of 80s, when the sunk costs and hysteresis literature emerged (e.g., BALDWIN, 1990). According to the hysteresis literature, nonlinearities in the export demand might be driven by strategic behaviour of the exporters, who, if once invested an amount of sunk costs into entering the market, are willing to protect their market shares in the destination country. That might lead to an incomplete pass-through of cost- and other types of shocks (as e.g. exchange rate changes) to prices, paid in local currencies in strategic destination markets. Then small exchange rate changes are expected to have no or only weak effect on exports, while large changes might lead to larger-scaled reactions of export volumes (BELKE et al., 2013).

Empirical literature often addressed the uncertainty of the exchange rate development and the effect of exchange rate volatility on trade (e.g. CHO et al., 2002; BONROY et al., 2007; LONGJIANG, 2011; SHELDON et al., 2013 for the case of agricultural trade). Studies of exchange rate nonlinearities in determining trade volumes are scarce. The few empirical studies which address nonlinearities and asymmetries of the trade reaction on the exchange rates of different sign or magnitude either ignore time-series properties of the underlying variables or capture solely short-run effects (e.g. KANNEBLEY, 2008, and BELKE et al., 2013).

The only exception the author is aware of is the study of VERHEYEN (2013a), in which he addressed exchange rate nonlinearities and asymmetries in the export demand function both in the long and in the short run. VERHEYEN adopted the nonlinear autoregressive distributed lag approach (NARDL) of SHIN et al. (2011) and applied it to the data on total exports of twelve European countries to the US. Asymmetric reactions of export demand to Euro appreciations and depreciations were found for many countries, especially in the long run, but no clear conclusions could have been made with regard to hysteresis. Although the study did not explain why the exchange rates were considered to be exogenous to the value of total exports, it showed that ignoring the nonlinear effect of the exchange rate on the export demand is too restrictive.

As there has been no study concerning this issue conducted for agri-food trade, this paper fills this gap and analyses exports of “Food and live animals” as defined by the Standard International Trade Classification (SITC) of eleven European countries to the US over the last 25 years.

European countries hold the second position among the world top agri-food exporters. The US is the largest export market of the EU (EUROPEAN COMMISSION, 2013). As most of all agri-food EU exports are final goods (e.g. dairy products, cereal, fruit and vegetable preparations and confectionery) it is expected to find more pronounced evidence of exchange rate nonlinearities in agri-food than it was recorded for total exports by VERHEYEN (2013a). This hypothesis is mainly driven by application of pricing-to-market strategies, which were often documented in empirical studies, suggesting that European exporters tend to hinder the pass-through of the Euro changes to the domestic US prices in order to protect their market shares in the US market (e.g. KNETTER, 1989; GLAUBEN and LOY, 2003). In total exports this might be less visible, as they include exports of raw materials and other homogeneous goods, for which world prices (and hence no room for pricing-to-market) are typically assumed.

To allow for nonlinearities in the short and in the long run and to address the time-series properties of the data (including possible hidden cointegration as in GRANGER and YOON, 2002), a partial sum decomposition and the NARDL approach of SHIN et al. (2013) combined with the bounds testing approach by PESARAN et al. (2001) are applied here.

The remainder of the paper is structured as follows: Section 2 describes the methodology in more detail, Section 3 introduces the data, Section 4 presents the results, Section 5 discusses the outcomes and the last section concludes.

2 Modeling Approach

As a starting point it is assumed that exports can be described by a reduced-form demand function[1] similar to one employed in BAHMANI-OSKOOEE (1986), BAHMANI-OSKOOEE and KARA (2003) and VERHEYEN (2013a and 2013b):

$$X_t = A * R_t^{\alpha} * Y_t^{\beta} \qquad (1),$$

where X_t are the European exports to the US at the time t, which are determined by some constant parameter A, the US demand Y, and the real exchange rate R. Alternatively, the effects of the nominal exchange rate (E) and relative prices (P) can be separated by substituting the real exchange rate as $R_t^{\alpha} = E_t^{\gamma} * P_t^{\delta}$.

Taking logs of Equation (1) results in Equation (2), which represents the long-run relationship between exports and its determinants (lower case letters *x, r, y* denote logs of variables):

$$x_t = a + \alpha r_t + \beta y_t \qquad (2).$$

To model potential asymmetries in the reaction of the US export demand to a change in the exchange rate, a partial sum decomposition and the NARDL framework by SHIN et al. (2013) are applied. This approach allows to test, whether exchange rate changes of different signs and magnitudes have a similar impact on the export demand, both in the short and the long run. Besides separating the effects of exchange rate appreciations and depreciations, this study aims to account for hysteresis (nonlinear reaction to small and large exchange rate changes), which is expected due to "wait-and-see" strategy of exporters, who neglect minor changes in exchange rate until some "pain threshold" is passed (e.g., BALDWIN, 1990; BELKE et al., 2013). To address these issues the exchange rate decomposition takes the following form:

$$r_t = r_0 + r_t^- + r_t^{\pm} + r_t^+ \qquad (3),$$

where

$$r_t^- = \sum_{j=1}^{t} \Delta r_j^- = \sum_{j=1}^{t} \Delta r_j I\{\Delta r_j \leq -STD\} \qquad (4);$$

$$r_t^{\pm} = \sum_{j=1}^{t} \Delta r_j^{\pm} = \sum_{j=1}^{t} \Delta r_j I\{-STD < \Delta r_j < +STD\} \qquad (5)$$

$$r_t^+ = \sum_{j=1}^{t} \Delta r_j^+ = \sum_{j=1}^{t} \Delta r_j I\{+STD \leq \Delta r_j\} \qquad (6),$$

and r_0 is the value of the exchange rate at the time t_0.

[1] This reduced-form specification of export demand is used in order to compare the outcomes for agri-food exports with the outcomes of the VERHEYEN (2013a) for total exports, which is the only benchmark study available.

Unlike SHIN et al. (2013) and VERHEYEN (2013a) here the logarithm of the exchange rate, not the original series of exchange rates, is decomposed. This allows us to avoid the problem of taking a logarithm of a negative number (exchange rate changes related to depreciations). Furthermore, instead of using various quantiles, the thresholds are fixed at the level of one positive and negative standard deviation (STD) as proposed by BUSSIÈRE (2013). This allows us to test how the export reaction changes within the range of standard fluctuations of exchange rates and outside of it.[2]

The NARDL specification for the Equation (2) is:

$$\Delta x_t = a_0 + a_1\left(x_{t-1} - a_2 r_{t-1}^- - a_3 r_{t-1}^{\pm} - a_4 r_{t-1}^+ - a_5 y_{t-1}\right) + \sum_{\tau=0} \eta_\tau \Delta r_{t-\tau}^- + \sum_{\tau=0} \theta_\tau \Delta r_{t-\tau}^{\pm} + \sum_{\tau=0} \iota_\tau \Delta r_{t-\tau}^+ + \sum_{\tau=0} \kappa_\tau \Delta y_{t-\tau} + \sum_{\omega=1} \lambda_\omega \Delta x_{t-\omega} + u_t \qquad (7).$$

As European agri-food exports are only a small part of total exports, and even a smaller fraction of these is shipped to the US, endogeneity between the export value and the exchange rate is not an issue for this specification.

The appropriate lag structure is chosen according to the Schwarz criterion (a maximum lag length of 12 is considered as monthly data is used). When autocorrelation is still present in the chosen specification lags of the first difference of the dependent variable are added in order to overcome the problem.

As the estimation of NARDL with OLS delivers only the product of the exchange rate estimates and the coefficient of the lagged export demand variable, (approximated values of) long-run elasticities need to be recalculated as follows:

$$rer^- = -\frac{a_2}{a_1};\ rer^{\pm} = -\frac{a_3}{a_1};\ rer^+ = -\frac{a_4}{a_1}. \qquad (8).$$

Standard errors and significance levels of the recalculated coefficients can be obtained using the Delta method. To test for the long-run relationship between variables, the Bounds testing[3] by PESARAN et al. (2001) is applied. The symmetry is tested by means of a Wald test.

[2] Alternatively, one could use diverse quantiles to test how the outcomes change at the different levels of thresholds as done by VERHEYEN (2013a).

[3] Bounds testing can be applied to variables of I(0), I(1) or mixed order of integration. None of variables used in the empirical specification is of order I(2) or higher. Results of the unit-root pretesting are available upon a request.

Positive values for the estimates of the foreign demand (y) are expected, as an increase in the US income is supposed to raise American imports. As for the exchange rates, it is assumed that a Euro depreciation should stimulate European exports (negative long run coefficient rer^-), while a Euro appreciation should have an opposite effect and decrease the export demand (negative coefficient rer^+). If European agri-food exporters apply pricing strategies to mitigate negative impacts of a strong Euro, asymmetry of the export demand reactions to Euro appreciations and depreciations is awaited. For the inner regime ($rer^{\pm}$) coefficients, which are smaller in absolute terms than those of the other regimes are expected, if hysteresis is present in exports (similar to VERHEYEN, 2013a).

3 Data

The sample includes monthly agri-food bilateral exports of eleven European countries to the US. These countries are Austria (AT), Belgium (BE), Germany (DE), Spain (ES), Finland (FI), France (FR), Greece (GR), Ireland (IE), Italy (IT), the Netherlands (NL), and Portugal (PT). Nominal export data are taken from Eurostat and are measured in Euro. Agri-food exports are defined as those included into the SITC group 0 "Food and live animals". Exports are deflated by the corresponding consumer prices for food, taken from the OECD Main Economic Indicators (MEI). Presumably, using export price indices or a GDP deflator would be a better solution to adjust exports. But as the export price indices are not acquirable for all the countries and the GDP deflator is only available quarterly, this study follows GRIER and SMALLWOOD (2007) and uses consumer prices.

Figures 1 and 2 show the development of nominal agri-food exports to the US and its share in total exports to the US over time.

Most of the considered exporting countries adopted the Euro in 1999 with the exception of Greece, which introduced the Euro one year later. Nominal exchange rates are measured as units of American Dollar (USD) per 1 Euro and are taken from Eurostat. Hence, an increase of the exchange rate corresponds to a Euro appreciation. Official conversion rates are used to obtain bilateral exchange rate series for the period before the Euro introduction. In order to calculate real exchange rates, nominal exchange rates are multiplied by relative prices. Those are measured as consumer price indexes (CPI) of the corresponding European country divided by the American CPI. The US demand is approximated by the index of industrial production (IIP), as it is available on a monthly basis, as opposed to GDP, which is only available quarterly. Both CPI and IIP are taken from the OECD MEI database.

Exports, relative prices and industrial production are deseasonalised using the Census-12 procedure to exclude the influence of seasonal effects. For most of the countries the analysis covers the period from January 1988 to December 2013. For Austria and Finland the export data is available only from 1995. The data on consumer prices for food is available from 1991 for Belgium and 1993 for Spain. For these countries the time span used for the analysis is shorter. All variables enter estimations in logarithms.

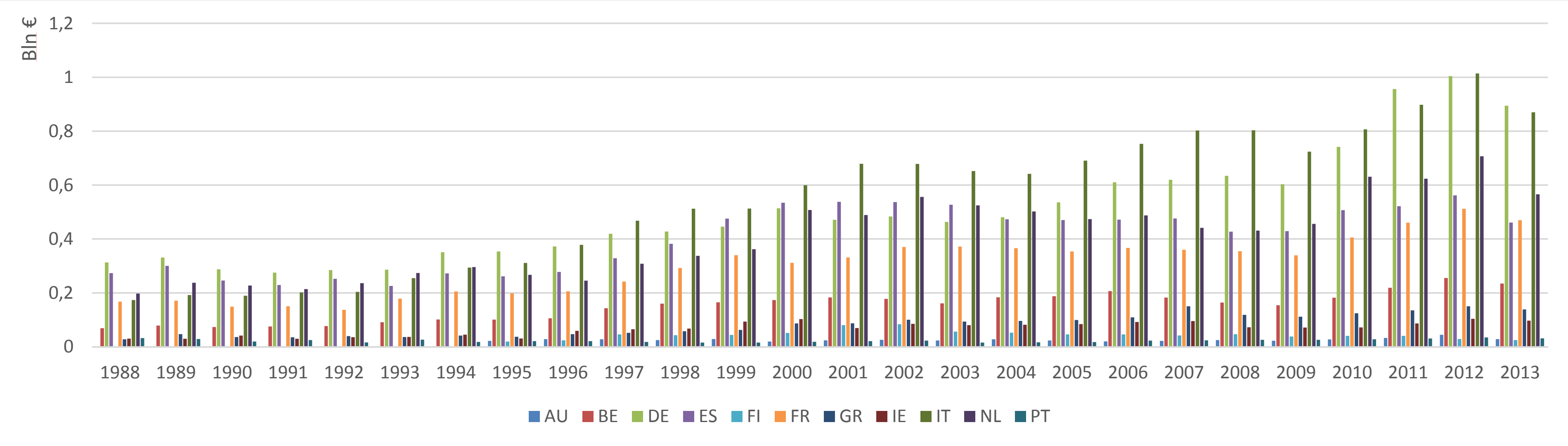
Descriptive statistics on export volumes and real exchange rates are reported in Table 1. Annex 1 provides some additional descriptive statistics on exchange rates (logs, in first differences).

4 Results and Discussion

In this section the general outcomes of the estimated models are first presented and then put into the context of existing studies. Furthermore, some reasons which might be driving the obtained results are provided. Detailed outcomes of the estimated NARDL models are reported in Annex 2. The adjusted coefficients of determination vary from 0.28 (Greece) to 0.47 (France). On average, included variables are able to explain about 39 percent of the variation of exports' changes. All the models pass diagnostic tests, including the Breusch-Godfrey serial correlation LM test, the Ramsey RESET test, and the CUSUM test, according to which most of the models are stable over time. Bounds testing rejects the hypothesis of no long-run relationship for level variables for all the models. Additionally, the IIP, which is a measure of the US income, enters the equations with an expected positive sign in eight out of eleven cases. In the other three models, the estimated income parameter does not statistically differ from zero at any conventional levels of significance.

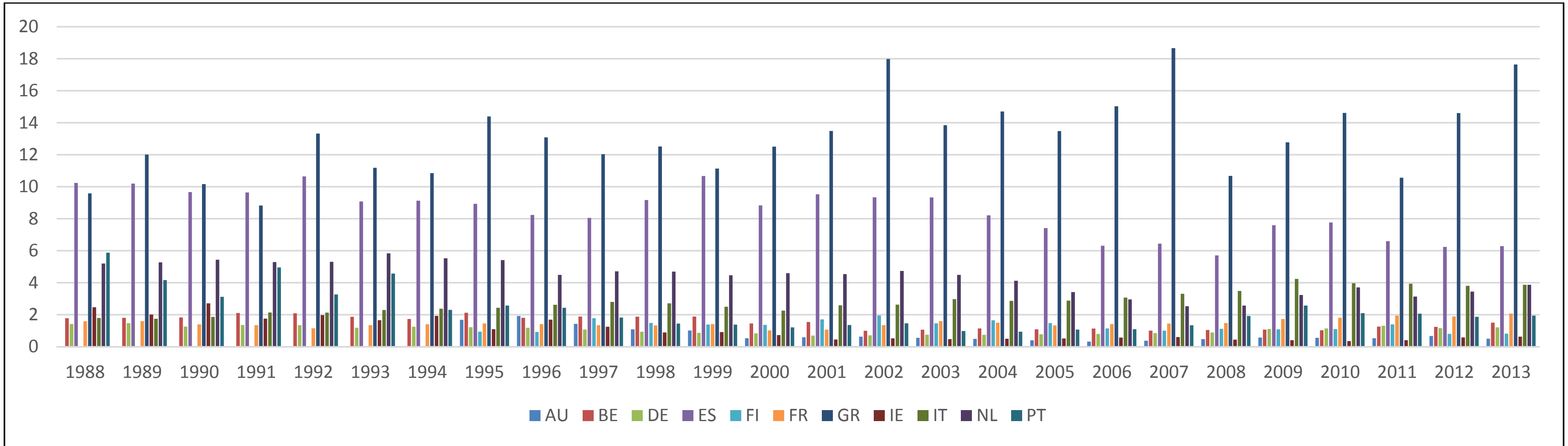
Table 2 shows the approximated long-run exchange rate elasticities and the significance levels which have been calculated using the Delta method. Few observations can be made based on reported results. First, in ten out of eleven cases the effect of the Euro depreciation is much higher in absolute terms than that of appreciations. This implies that the Euro,

Figure 1. Agri-food exports to the US (nominal, bln Euro, 1988-2013



Source: own presentation with the data from Eurostat

Figure 2. Share of agri-food exports (SITC 0) in total exports to the US, %



Source: own presentation with the data from Eurostat

Table 1. Descriptive statistics

Country	AT	BE	DE	ES	FI	FR	GR	IE	IT	NL	PT
Real exports in Euro (Nominal exports, consumer prices (food) and seasonally adjusted)											
Mean	2478114	15745175	46570596	43783203	4156766	28662713	9061192	6278776	54004150	37659402	2342672
Median	2334871	16133576	43616951	42848341	3687365	30538223	8874861	6229161	58963691	37040678	2055536
Maximum	5675545	28055587	89484022	81268009	16523282	51091056	27831168	15614785	89339645	65524356	6941385
Minimum	568688	6147281	19976290	24032391	228452	9840877	1944665	1306652	20135384	12634434	809802
Std. Dev.	797079	4074074	15315703	9942620	2373264	8156079	3083728	2384116	18457111	11512615	1097167
Real exchange rate (Nominal exchange rate, relative prices adjusted)											
Mean	1.250	1.257	1.316	1.177	1.270	1.330	0.949	1.227	1.170	1.272	1.143
Median	1.281	1.293	1.329	1.196	1.295	1.336	0.966	1.242	1.202	1.291	1.170
Maximum	1.558	1.592	1.672	1.570	1.569	1.729	1.503	1.687	1.619	1.553	1.588
Minimum	0.889	0.878	0.918	0.823	0.928	0.901	0.220	0.856	0.748	0.888	0.647
Std. Dev.	0.155	0.157	0.155	0.179	0.147	0.178	0.348	0.167	0.200	0.147	0.214
Observations	228	276	312	252	228	312	312	312	312	312	312

Source: own computations

Table 2. Long-run elasticities of agri-food exports to the US with respect to exchange rates

	AT	BE	DE	ES	FI	FR	GR	IE	IT	NL	PT
rer^{-}	-0.240	-0.891 ***	-0.522 *	-1.262 ***	-0.939 *	-1.045 ***	-1.678 ***	-0.913 ***	-0.641 ***	-1.156 ***	-1.704 **
	(0.573)	(0.205)	(0.309)	(0.155)	(0.493)	(0.271)	(0.296)	(0.272)	(0.090)	(0.265)	(0.716)
$rer^{\pm}$	0.779	-0.271	0.460	0.935 ***	-3.201 ***	-0.388	-0.925 **	-0.618	-0.128	-1.412 ***	-0.322
	(0.836)	(0.301)	(0.575)	(0.278)	(0.626)	(0.310)	(0.383)	(0.508)	(0.116)	(0.374)	(1.340)
rer^{+}	-0.147	-0.762 ***	-0.008	-1.158 ***	-1.230 ***	-0.801 ***	-0.853 ***	-0.877 ***	-0.314 ***	-0.612 ***	-0.804
	(0.538)	(0.153)	(0.275)	(0.114)	(0.467)	(0.253)	(0.224)	(0.230)	(0.073)	(0.199)	(0.559)

Notes: Delta method standard errors are in parentheses. ***, ** and * refer to significance at the 1, 5, 10 % level.
Source: own computations

Table 3. Symmetry testing summary

	AT	BE	DE	ES	FI	FR	GR	IE	IT	NL	PT
$a_2 = a_3 = a_4$	0.505	0.146	0.011	0.071	0.000	0.003	0.000	0.688	0.000	0.014	0.043
$a_2 = a_3$	0.362	0.103	0.160	0.193	0.002	0.070	0.068	0.529	0.000	0.387	0.345
$a_3 = a_4$	0.421	0.178	0.493	0.431	0.014	0.311	0.879	0.646	0.162	0.035	0.742
$a_2 = a_4$	0.562	0.129	0.004	0.167	0.116	0.024	0.000	0.829	0.000	0.004	0.015
Short run (contemporaneous)	0.652	0.915	0.943	0.978	0.027	0.777	0.498	0.726	0.974	0.174	0.144
Short run (lag1)			0.390		0.768						0.626

Notes: Wald test results (p-values) are reported.
Source: own computations

appreciations do not harm the exported values to the same extent as exporters profit from the depreciating Euro. The only exception from this pattern is Finland, whose exports benefit less from the weak Euro than they suffer from a strong Euro (similar outcomes were obtained by VERHEYEN (2013a) for total exports).

The second observation refers to hysteresis, which can be defined as a weaker reaction (in absolute terms) of the exports to smaller changes of the exchange rate (inner regime) as opposed to large changes. In my sample, the hypothesis of hysteresis cannot be rejected for most of the countries. For the cases of Belgium, Spain, France, Ireland, Italy and Portugal the absolute value of the coefficient of the inner regime is much smaller than the value of both coefficients referring to large appreciations and depreciations. And in the case of Austria, Germany and Portugal the inner regime coefficients are not statistically different from zero, so there is, once again, an evidence in favour of hysteresis.

Table 3 reports results of the symmetry testing for the agri-food export demand. The equality of all the long-run exchange rate coefficients is rejected in eight out of eleven cases. Hence, appreciations and depreciations seem to have a different impact on exports in these eight countries. As for Austria, Belgium and Ireland large appreciations and depreciations seem to affect exports symmetrically. Furthermore, asymmetry between the appreciations and depreciations is more pronounced, than between those and the inner regime. Short-run dynamics do not seem to play an important role in shaping the European exports. The short-run coefficients are mostly of a minor statistical significance and the F-test could not reject symmetry of short-run coefficients for any country, but Finland. The outcomes for Finland might be partially due to a shorter data sample. As the export data for Austria and Finland are only available from 1995, these two models lack 84 observations compared to countries with data available from 1988.

Moreover, the results obtained for agri-food exports differ somewhat between countries. This might be due to various factors, including (but not limiting to) the composition of exported goods, different market niches, which these products target on the US market, the intensity of competition exporters face on the US market, the sensitivity of the US demand (preferences) or pricing strategies of exporters.

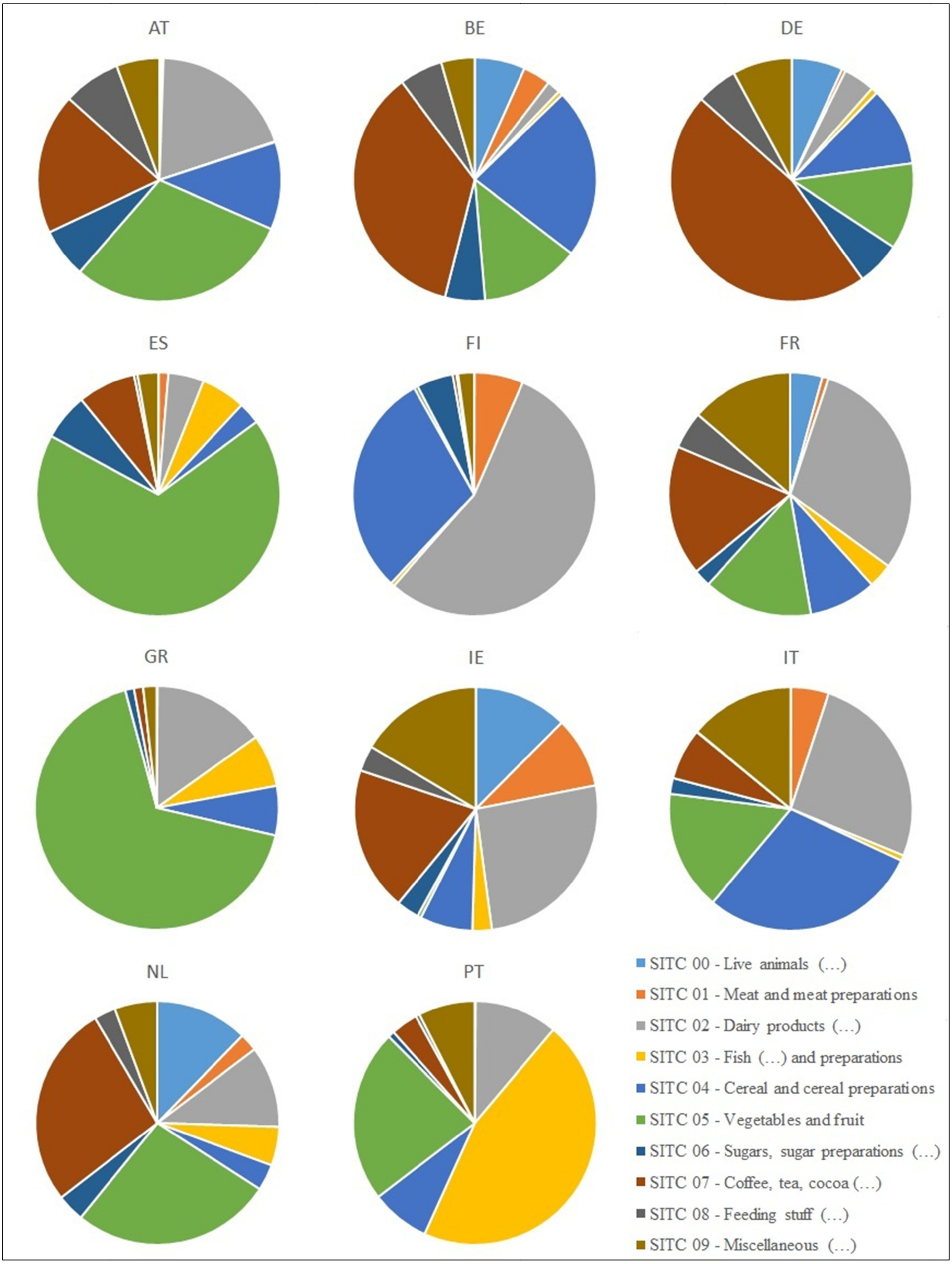
Figure 3 depicts the structure of the aggregated agri-food exports (the SITC code 0) on a 2-digit level.

Products of the SITC group 02 - Dairy products and birds' eggs, 05 - Vegetables and fruits, 07 - Coffee, tea, cocoa, spices, and manufactures thereof and 04 - Cereals and cereal preparations prevail in exports of the countries of our sample. Products from the group 07 were among the three main exported product groups for Austria (19 %, on average from 1988-2013), Belgium (36 %), Germany (46 %), France (18 %), Ireland (19 %), and the Netherlands (27 %). Dairy products exports are crucially important for Austria (19 %), Finland (55 %), France (29 %), Greece (15 %), Ireland (26 %), and Italy (26 %). Cereals and preparations are an important source of export revenues for Belgium (23 %), Germany (10 %), Finland (30 %), and Italy (29 %). Also, nearly all countries export vegetables and fruits and their preparations: Austria (29 %), Belgium (13 %), Spain (68 %), France (14 %), Greece (67 %), Italy (16 %), the Netherlands (27 %) and Portugal (23 %). Some deviations from the standard pattern shows, for example, Portugal, which mainly exports fish products (46 %). The composition of the exported goods, even on such an aggregated level varies considerably between the countries, which might, at least partially, explain the cross-country difference in the obtained coefficients. Another reason for cross-country discrepancies related to the composition of exports might be the subsidization policy of the European agricultural sector by the government, as support levels differ a lot across products. [1]

European countries specialize in the exports of final goods, which according to EUROPEAN COMMISSION (2013) account for two thirds of total agricultural exports. As this paper focuses on the products from the SITC 0 group, which does not include e.g. fats and oils (SITC 4) and targets exclusively exports to the US, the share of final processed goods in exports is even higher in my sample. Presumably, some of exported items might have gained reputation on the American market, so that the US consumers do not switch away from European goods as their local price in US Dollars rises, and consume more, once the Dollar price falls. It is also plausible that the European food exporters, who perceive the US market as strategically important and invested an amount of sunk cost in order to enter the market, use some pricing strategies (e.g., pricing-to-market). This might be done by partial offsetting of exchange rate changes in order to smooth fluctuations in shipped quantities by adjusting the markup that exporters set on marginal costs. Strategic pricing might be a plausible explanation behind

[1] If (and when yes, how) these subsidies might lead to asymmetric adjustments in exports remains a question for future research.

Figure 3. The structure of European agri-food exports to the US (SITC, 2 digits, average 1988-2013)



Source: own presentation with the data from Eurostat the nonlinearity of the export volumes' reactions towards Euro appreciations and depreciations, as empirical literature often found evidence of a pricing-to-market policy of European exporters, especially in their trade with the US (e.g., KNETTER 1989; FALK and FALK, 2000; GLAUBEN and LOY, 2003; STAHN, 2007). Few studies suggest that pricing-to-market strategies might be nonlinear and asymmetric as well, implying that the exporters act strategically, when adjusting their markups and decide on passing-through of exchange rate changes to prices in local currencies. Such an observation was made e.g. by FEDOSEEVA (2013), who investigated German exports of sugar confectionery and found evidence in favour of both asymmetry and hysteresis of pricing-to-market. According to this study, the exporters tend to partially offset the Euro appreciations on the most important markets, while depreciations are often fully passed-through. If this is true for other agri-food products as well, the imports of the European goods by the US do not change much, as Euro appreciates, which results in a less pronounced reaction of exports towards the Euro appreciation and a larger positive effect of the depreciations. This would lead exactly to the pattern of export reactions to the exchange rates, which is observed for most countries of this study.

As empirical literature concerning asymmetries and hysteresis of the exchange rate impact on export demand is scarce, the outcomes for agri-food exports are compared with those of VERHEYEN for total exports. One should keep in mind, that the considered export goods, data span, thresholds and the way of inclusion of the exchange rate in these two studies are not identical. [1]

VERHEYEN (2013a: 73) found indications of nonlinearity of the exchange rate in about 50 % of his estimations. He compared the outcomes of a one-threshold model and a two-threshold model and concluded that if nonlinearities are detected, exports tend to respond more to currency depreciations than to appreciations. Regarding hysteresis, the evidence was more mixed. A number of coefficients were of unreasonable magnitude and only for few countries an evidence of hysteresis was found.

In agri-food exports nonlinearities seem to be more pronounced. This holds for both, asymmetries and hysteresis. The effect of depreciations in absolute terms is much higher than that of appreciations for all countries, but Finland. An evidence in favour of hysteresis, accepting the definition of VERHEYEN (2013a), is recorded for all but three countries (Finland, Greece and the Netherlands). The reason for such discrepancies between total and agri-food exports might be due to aggregation issues. As total exports include raw materials or other homogenous goods, which are typically assumed to be priced at a world price level, asymmetric or nonlinear reactions of individual product export groups might be well hindered.

5 Conclusion

This study returns to the estimation of trade elasticities in an attempt to re-address the role of the exchange rate in determining exports. Traditionally, empirical investigations concentrated either on exchange rate volatility or limited themselves to the inclusion of the exchange rate into the model in a linear way. Though exchange rates proved to be an important determinant of trade, the question of possible nonlinearities of the exchange rate stayed undeservedly neglected. This study fills the gap in the literature by allowing the export demand to react differently to exchange rate changes of different magnitudes and directions. These nonlinearities and asymmetries of the export reaction are modelled using a partial sum decomposition and the NARDL approach of SHIN et al. (2013). The application of such a framework allowed me to assess how exports react to various types of exchange rate changes both in the long and the short run. The empirical test is conducted using the data on aggregated agri-food exports (SITC 0) of eleven European countries to the US, which is the main trade partner of these countries outside of the EU. The data span includes 25 years of monthly observations, from 1988 to 2013, for most countries.

The results support the importance of exchange rates in shaping exports, as indicated by previous studies (e.g. BAHMANI-OSKOOEE and KARA, 2008). Still, allowing for asymmetries and nonlinearities provides some more insights on the pattern of exports' reactions to exchange rate changes of different nature, compared to earlier literature. The outcomes of this study might contribute to the lively debates, which

[1] In an earlier version of this paper (FEDOSEEVA, 2014) export demand equations for total exports are also estimated. The outcomes are very close to those of VERHEYEN (2013a). The estimation of the model, in which the real exchange rate is substituted by the nominal exchange rate and relative prices results in somewhat different coefficients, but does not alter the outcomes regarding asymmetry and hysteresis of the export reaction towards exchange rate changes.

discuss the hampering effect of a strong Euro on European trade.

The estimation results suggest that the strong Euro does not harm agri-food exports as much, as exporters are able to benefit from the weak Euro. Such outcomes are found for ten out of eleven countries in the sample, as the reaction to depreciations is larger in absolute terms, than that to appreciations. Furthermore, the evidence in favour of hysteresis is found for most of the exporting countries, which implies that slender exchange rate changes have a smaller effect on exports. This finding is in line with BELKE et al. (2013), who suggests that there is a pain threshold, passing of which leads to a greater reaction of exports to the exchange rate shifts. Although BELKE et al. (2013) made an effort to quantify this effect, they did not consider the possibility of an asymmetric impact of appreciations and depreciations.

Moreover, the outcomes are generally in line with those of VERHEYEN (2013a), who applied a similar methodology to the data on total exports. Still, the estimates for agri-food exports indicate both a more pronounced asymmetry and hysteresis in reactions of export demand to exchange rate changes. Those reactions also differ between countries. Cross-country discrepancies might be well due to a different structure of exported agri-food goods, to a divergent degree of competition these products face in the US market, to various pricing strategies of exporters, or some other factors. Regardless the specific reason for each concrete exporting country, it seems like on average the European agri-food exporters have found a way to cope with the negative impacts of the currency appreciations.[2]

As European countries export a lot of final goods to the US, the European exporters might apply pricing-to-market strategies to hinder the pass-through of the exchange rate appreciations in order to stay competitive on the US market and to protect their market shares. Euro depreciations might then be used to expand exports. Numerous empirical pricing-to-market studies provide evidence in favour of such markup (price) adjustments made by European exporters for the case of agri-food exports, chemical products and manufactured goods.

Still, testing those hypothesis empirically would require employing a more detailed dataset and application of some alternative models (e.g. residual demand elasticity, gravity or pricing-to-market models for individual products/exporters), which is left for future research.

[2] Of course, it might be not the case for each individual exporter, but as such disaggregated data is not available, this statement cannot be assessed at this stage.

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References

ADLER, H.J. (1945): United States import demand during the interwar period. In: The American Economic Review 35 (3): 418-430.

ANDERSON, J.E. (1979): A theoretical foundation for the gravity equation. In: The American Economic Review 69 (1): 106-116.

ANDERSON, J.E. and E. VAN WINCOOP (2003): Gravity with gravitas: A solution to the border puzzle. In: The American Economic Review 93 (1): 170-192.

ANDERSON, J.E., M. VESSELOVSKY and Y.V. YOTOV (2013): Gravity, scale and exchange rates. NBER Working Paper, No. 18807. NBER, Cambridge, USA.

BAHMANI-OSKOOEE, M. (1986): Determinants of international trade flows: The case of developing countries. In: Journal of Development Economics 20 (1): 107-123.

BAHMANI-OSKOOEE, M. and O. KARA (2003): Relative responsiveness of trade flows to a change in prices and exchange rate. In: International Review of Applied Economics 17 (3): 293-308.

– (2008): Relative responsiveness of trade flows to a change in prices and exchange rate in developing countries. In: Journal of Economic Development 33 (1): 147-163.

BALDWIN, R. (1990): Hysteresis in trade. In: Empirical Economics 15 (2): 127-142.

BELKE, A., M. GÖCKE and M. GÜNTHER (2013): Exchange rate bands of inaction and play-hysteresis in German exports – sectoral evidence for some OECD destinations. In: Metroeconomica 64 (1): 152-179.

BERGSTRAND, J.H. (1990): The Heckschner-Ohlin-Samuelson model, the Linder hypothesis and the determinants of bilateral intra-industry trade. In: The Economic Journal 100 (403): 1216-1229.

BONROY, O., J.-P. GERVAIS and B. LARUE (2007): Are exports a monotonic function of exchange rate volatility?

Evidence from disaggregated pork exports. In: Canadian Journal of economics 40 (1): 127-154.

BUSSIÈRE, M. (2013): Exchange rate pass-through to trade prices: The role of non-linearities and asymmetries. In: Oxford Bulletin of Economics and Statistics 75 (5): 731-58.

CAPORALE, G.M. and M.K.F. CHUI (1999): Estimating income and price elasticities of trade in a cointegration framework. In: Review of International Economics 7 (2): 254-264.

CHO, G., I. SHELDON and S. MCCORRISTON (2002): Exchange rate uncertainty and agricultural trade. In: American Journal of Agricultural Economics 84 (4): 931-942.

DEARDORFF, A.V. (1998): Determinants of bilateral trade: Does gravity work in a neoclassical world? In: Frankel, J.A. (ed.): The Regionalization of the World Economy. University of Chicago Press, Chicago, London: 7-32.

EUROPEAN COMISSION (2013): Monitoring agri-trade policy. Agricultural trade in 2012: A good story to tell in a difficult year? In: http://ec.europa.eu/agriculture/trade-analysis/map/index_en.htm. Call: 12.02.2014.

FALK, M. and R. FALK (2000): Pricing to market of German exporters: Evidence from panel data. In: Empirica 27 (1): 21-46.

FEDOSEEVA, S. (2013): (A)symmetry, (non)linearity and hysteresis of pricing-to-market: Evidence from German sugar confectionery exports. In: Journal of Agricultural and Food Industrial organization 11 (1): 69-85.

– (2014): Is there a difference? Exchange rate nonlinearities in European agri-food (versus total) exports to the US. Discussion Paper No. 66. Center for International Development and Environmental Research, University of Giessen.

GLAUBEN, T. and J.-P. LOY (2003): Pricing-to-market versus residual demand elasticity analysis of imperfect competition in food exports: Evidence from Germany. In: Journal of Agricultural and Food Industrial Organization 1 (1): article 3.

GRANGER, C.W.J. and G. YOON (2002): Hidden cointegration. Economics Working Paper No. 2002-02. University of California, San Diego.

GRIER, K.B. and A.D. SMALLWOOD (2007): Uncertainty and export performance: Evidence from 18 countries. In: Journal of Money, Credit and Banking 39 (4): 965-979.

KANNEBLEY, JR. S. (2008): Tests for the hysteresis hypothesis in Brazilian industrialized exports: A threshold cointegration analysis. In: Economic Modelling 25 (2): 171-190.

KNETTER, M.M. (1989): Price discrimination by U.S. and German exporters. In: The American Economic Review 79 (1): 198-210.

KUMAR, S. (2009): An empirical evaluation of export demand in China. In: Journal of Chinese Economic and Foreign Trade Studies 2 (2): 100-109.

LONGJIANG, C. (2011): The effect on China's RMB exchange rate movement on its agricultural exports: A case study of export to Japan. In: China Agricultural Economic Review 3 (1): 26-41.

MARQUEZ, J. (1990): Bilateral trade elasticities. In: The Review of Economics and Statistics 72 (1): 70-77.

MARTINEZ-ZARZOSO, I. and F. NOWAK-LEHMANN (2003): Augmented gravity model: An empirical application to Mercosur-European union trade flows. In: Journal of Applied Economics 6 (2): 291-316.

NARAYAN, P.K. and S. NARAYAN (2005): Estimating income and price elasticities of imports for Fiji in a cointegration framework. In: Economic Modelling 22 (3): 423-438.

ORCUTT, G.H. (1950): Measurement of price elasticities in international trade. In: Review of Economics and Statistics 32 (2): 117-132.

PESARAN, M.H, Y. SHIN and R.J. SMITH (2001): Bounds testing approaches to the analysis of level relationships. In: Journal of Applied Econometrics 16 (3): 289-326.

RAO, B.B. and R. SINGH (2007): Estimating export equations. In: Applied Economic Letters 14 (11): 799-802.

SHELDON, I., K.S. MISHRA, D. PICK and S.R. THOMPSON (2013): Exchange rate uncertainty and US bilateral fresh fruit and fresh vegetable trade: An application of the gravity model. In: Applied Economics 45 (15): 2067-2082.

SHIN, Y., B. YU and M.J. GREENWOOD-NIMMO (2011): Modelling asymmetric cointegration and dynamic multipliers in a nonlinear ARDL framework. In: http://papers.ssrn.com/sol3/papers.cfm?abstract_id=1807745.

SHIN, Y., B. YU and M.J. GREENWOOD-NIMMO (2013): Modelling asymmetric cointegration and dynamic multipliers in a nonlinear ARDL framework. In: Horrace, W.C. and R.C. Sickles (eds.): Festschrift in Honor of Peter Schmidt. Springer Science & Business Media, New York (forthcoming).

STAHN, K. (2007): Has the export behavior of German enterprises changed? Empirical evidence from German sectoral prices. In: Journal of Economics and Statistics 227 (3): 295-329.

TINBERGEN, J. (ed.) (1962): Shaping the world economy. Suggestions for an international economic policy. Twentieth Century Fund, New York.

VERHEYEN, F. (2013a): Exchange rate nonlinearities in EMU exports to the US. In: Economic Modelling 32 (C): 66-76.

– (2013b): The stability of German export demand equations – have German exports suffered from the strength of the Euro? In: International Economics and Economic Policy (forthcoming).

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Annex 1. Descriptive statistics for real exchange rates (Δ ln)

	AT	BE	DE	ES	FI	FR	GR	IE	IT	NL	PT
Mean	0.000	-0.000	-0.000	0.002	-0.000	-0.001	0.006	0.000	0.001	-0.000	0.002
Median	0.001	-0.001	-0.000	0.001	0.000	0.000	0.005	0.001	0.002	-0.000	0.002
Maximum	0.066	0.063	0.066	0.080	0.065	0.065	0.076	0.068	0.140	0.068	0.069
Minimum	-0.069	-0.077	-0.075	-0.070	-0.069	-0.075	-0.070	-0.075	-0.089	-0.075	-0.077
Std. Dev.	0.023	0.024	0.024	0.024	0.023	0.024	0.027	0.026	0.027	0.024	0.025
Skewness	0.100	-0.130	-0.127	0.221	0.090	-0.131	0.179	-0.098	0.240	-0.117	0.041
Kurtosis	3.103	3.329	3.160	3.191	3.101	3.093	2.816	2.925	4.721	3.205	3.072
Jarque-Bera	0.481	2.006	1.163	2.433	0.402	1.007	2.095	0.567	41.356	1.255	0.155
Probability	0.786	0.367	0.559	0.296	0.818	0.604	0.351	0.753	0.000	0.534	0.925
Sum	0.040	-0.123	-0.152	0.393	-0.026	-0.203	1.733	0.025	0.428	-0.108	0.656
Sum Sq. Dev	0.120	0.159	0.179	0.142	0.122	0.185	0.225	0.206	0.233	0.180	0.200
Observations	227	275	311	251	227	311	311	311	311	311	311

Source: own computations

Annex 2. Outcomes of the estimated NARDL models

	AT	BE	DE	ES	FI	FR	GR	IE	IT	NL	PT
Const.	8.15 ***	5.08 ***	3.76 ***	7.14 ***	8.86 ***	4.59 ***	4.50 ***	4.65 ***	7.17 ***	5.68 ***	5.73 ***
x_{t-1}	-0.44 ***	-0.49 ***	-0.23 ***	-0.58 ***	-0.70 ***	-0.34 ***	-0.43 ***	-0.55 ***	-0.69 ***	-0.34 ***	-0.26 ***
r^{-}_{t-1}	-0.11	-0.44 ***	-0.12	-0.73 ***	-0.66 *	-0.35 ***	-0.72 ***	-0.51 ***	-0.44 ***	-0.39 ***	-0.44 *
$r^{\pm}_{t-1}$	0.34	-0.13	0.11	-0.54 ***	-2.24 ***	-0.13	-0.40 **	-0.34	-0.09	-0.48 ***	-0.08
r^{+}_{t-1}	-0.07	-0.37 ***	-0.00	-0.67 ***	-0.86 **	-0.27 ***	-0.37 ***	-0.49 ***	-0.22 ***	-0.21 ***	-0.21
y_{t-1}	-0.36	0.64 ***	0.05	0.70 ***	0.37	0.24	0.58 ***	0.87 **	1.07 ***	-0.01	-0.45
Δr^{-}_{t}	-1.37	-0.93	-0.15	-0.41	-2.04	0.48	-0.97	-0.88	0.01	-0.76	1.49
$\Delta r^{\pm}_{t}$	0.68	-1.38 *	0.05	-0.57	4.09 *	0.61	-1.38	-0.37	0.07	-0.14	-0.48
Δr^{+}_{t}	-0.32	-1.17 **	-0.32	-0.54	-3.86 **	0.03	0.05	-1.82	-0.04	0.67	-1.79
Δy_{t}	1.64	3.28 **	0.21	0.50	-5.03	2.94 **	1.60	-0.80	1.58	4.05 ***	0.49
Δr^{-}_{t-1}			-1.00 **		-0.74						-3.01 ***
$\Delta r^{\pm}_{t-1}$			-1.75 **		1.80						-1.30
Δr^{+}_{t-1}			-0.36		0.05						-2.16 **
Δy_{t-1}			-0.46		1.43						-2.90
Δx_{t-1}	-0.39 ***	-0.31 ***	-0.54 ***	-0.10	-0.13 **	-0.60 ***	-0.19 *	-0.26 ***	-0.20 ***	-0.37 ***	-0.58 ***
Δx_{t-2}	-0.28 ***	-0.21 **	-0.32 ***	0.04		-0.51 ***	-0.04	-0.10		-0.38 ***	-0.38 ***
Δx_{t-3}		-0.06	-0.13 *	0.14 **		-0.29 ***	-0.11			-0.09	-0.20 **
Δx_{t-4}		-0.13 **				-0.24 ***	-0.09			-0.14 **	-0.16 **
Δx_{t-5}						-0.11 *	-0.02			0.02	
Δx_{t-6}							-0.00				
Δx_{t-7}							-0.09				
Δx_{t-8}							0.00				
Adj. R^2	0.42	0.39	0.36	0.32	0.41	0.47	0.28	0.39	0.43	0.37	0.43
LM-Corr. (p-val.)	0.99	0.09	0.16	0.05	0.06	0.21	0.07	0.05	0.12	0.16	0.10
R-Reset (p-val.)	0.55	0.06	0.36	0.78	0.59	0.32	0.65	0.05	0.30	0.10	0.41
Cusum	+	+	-	+	+	+	-	+	+	+	+
Bounds t. (F-stat.)	6.14[a]	5.37[a]	3.28[c]	9.69[a]	12.44[a]	4.81[a]	6.20[a]	8.16[a]	11.87[a]	3.28[c]	3.12[c]

Notes: ***, ** and * denote significance at the 1, 5 and 10 percent error term (White robust standard errors).
[a], [b], [c] denote significance at the 1, 5 and 10 percent level, respectively, and refer to the outcomes of the bounds testing according to PESARAN et al. (2001: 300, Table CI(ii), k=4).

Source: own computations