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Bilateral Exports from Euro Zone Countries to the US – Does Exchange Rate Variability Play a Role?

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Does Exchange Rate Variability Play a Role?

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Abstract

The financial crisis and the debt crisis in Europe lead to pronounced swings of the $/€-exchange rate. The influence of this exchange rate uncertainty on exports is neither theoretically nor empirically unambiguous. Therefore, this investigation tries to find out what effect exchange rate volatility has got on exports from eleven euro zone countries to the US. Our results suggest that if exchange rate volatility exerts a significant influence on exports, it is typically negative. Furthermore, exports of SITC categories 6 and 7 seem to be affected negatively most often.

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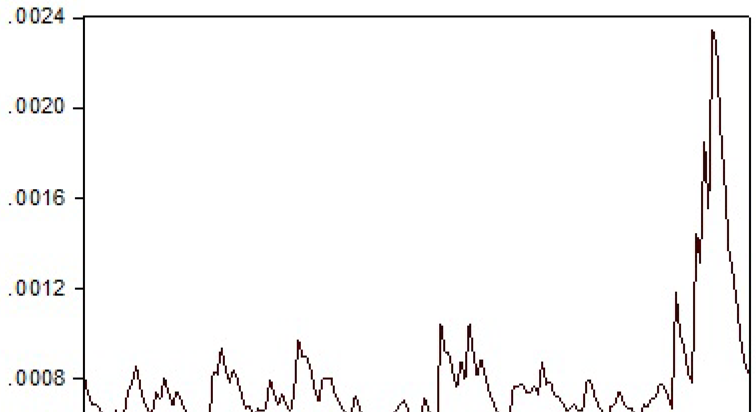
1. Introduction

The financial crisis and the debt crisis in Europe lead to pronounced swings of the \$/€-exchange rate. After reaching values of 1.60 US-\$ per € in mid-2008, the euro devaluated sharply until the end of 2008. In 2009 the trend was reversed and the euro recovered to 1.50 \$/€ but as the European debt crisis intensified, the euro dropped against the dollar once more. At the end of March 2011 the exchange rate notifies at about 1.40 \$/€ which is still somewhat above its purchasing power parity level of about 1.15 \$/€. These developments on international financial markets together with the Chinese exchange rate policy lead IMF chief Dominique Strauss-Kahn to voice fears of a currency war. Additionally, World Bank chief Robert Zoellick brought up a debate about a new gold standard.

The recent events highlight the fact, that exporters in the euro zone face an undeniable uncertainty with respect to the exchange rate, which seems to be high especially during the financial crisis. A look at the volatility measures used in this paper confirms this view. For instance, figure 1 displays the bilateral exchange rate volatility of the euro and the US-\$ for Germany, when using a GARCH(1,1) model.[2] This exchange rate variability might well affect bilateral exports from countries of the euro zone to the United States. Severe differences with regard to the effect of exchange rate volatility on exports could provoke tensions about the future role of foreign exchange policy in the EMU.

Fig. 1: Exchange rate volatility of the \$/€ exchange rate for Germany



[2]See equations (3) and (4) for details.

Therefore, this paper takes a closer look at the role of exchange rate variability on bilateral exports from eleven euro zone countries to the United States.[3] In order to do this, the next section gives a brief overview of the existing literature on the relationship between exchange rate variability and export performance. Section 3 describes the ARDL bounds testing approach of Pesaran and Shin (1999) and Pesaran et al. (2001), which is applied in this paper. In Section 4, we turn to the results, which indicate that there seems to be mostly a negative effect of exchange rate volatility on exports to the US. The last section concludes.

2. Literature Review

There is a vast literature dealing with exchange rate uncertainty and exports. For a comprehensive overview see McKenzie (1999) or Bahmani-Oskooee and Hegerty (2007). In general, as economic theory gives both reasons for a positive or negative impact of exchange rate volatility on exports, no consensus whether exchange rate variability stimulates or depresses exports has emerged. Early work of Ethier (1973) states a negative relationship between exchange rate uncertainty and a firm's exports, when the firm does not know how its revenue depends on the future exchange rate. But this uncertainty can be reduced by engaging in forward contracts. Other research claims that higher volatility might increase trade. For example, Franke (1991) derives conditions under which higher exchange rate volatility leads to more exports. Especially disadvantaged firms can benefit from volatility as it gives them leeway to set prices more freely.

Because of this ambiguity, the question how exchange rate volatility affects trade volumes, has been investigated empirically in various studies, especially for developing countries.[4] Fewer papers deal with trade of industrialized economies. This fact might be due to the notion that financial markets in developing countries might not deliver the appropriate tools to hedge the exchange rate risk (Caglayan and Di 2008). This opinion is supported by Grier and Smallwood (2007) and Ćorić and Pugh (2010).

To the best of our knowledge, there is no study which deals with the issue of the effect of exchange rate variability on exports from euro zone countries to the US. Caglayan and Di (2008) provide a comparative study of trade relationships between 13 industrial and developing countries and the US but their focus is somewhat different than ours. Although they use the same disaggregation as we do, their focus is on both exchange rate variability and

[3] The corresponding economies are Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal and Spain. They were selected due to data availability.

[4] To give some examples, Grobar (1993) or Arize et al. (2000).

income variability. Furthermore, they do not employ cointegration techniques as they estimate their model in first differences

One could argue that for euro zone countries the overall role of exchange rate uncertainty is reduced as many exports are within the currency union, where per definition no exchange rate risk exists. Nevertheless, the US are a main trading partner for many countries of the euro zone. Therefore, we believe that our comparative study adds additional insights to the literature. If exchange rate volatility does influence exports to the US differently in some countries of a currency union, this might result in a struggle whether the political authorities should focus on the exchange rate or not.

The ARDL bounds testing approach of Pesaran and Shin (1999) and Pesaran et al. (2001) used in this study has been employed by various other studies dealing with exchange rate volatility and export performance. De Vita and Abbott (2004) examine whether exchange rate volatility impacts on UK exports to the EU. They find that short-term exchange rate volatility has no effect on exports whereas long-term volatility exhibits a significant negative effect. Todani and Munyama (2005) deal with the same question for South Africa. They find that the results depend on the measure of exchange rate volatility. Furthermore, they either find no significant relationship or a positive one. Bustaman and Jayanthakumaran (2006) investigate whether Indonesia's exports to the US are sensitive to exchange rate volatility. They find positive as well as negative coefficients for exchange rate volatility, but overall their results support the view that higher volatility depresses exports. Hosseini and Moghaddasi (2010) investigate this issue for Iran. They find mixed results of cointegration, depending on the volatility measure used. Their results suggest that if there is any effect of exchange rate uncertainty, it is positive. Jiranyakul (2010) inspects whether Thailand's exports to the US and Japan are influenced by exchange rate uncertainty. He concludes that exchange rate volatility affects exports to Japan negatively while no significant effects for exports to the US occur.

To conclude our literature survey, Ćorić and Pugh (2010) perform a meta-regression analysis for a total of 49 studies on exchange rate variability and international trade with 686 observations. Their results suggest that there is a negative link between exchange rate volatility and trade. This is especially true for less developed countries where forward markets are less effective compared to industrial countries. Additionally, they find that studies that use cointegration and error correction techniques are more likely to find a negative link between exchange rate variability and trade.

3. Empirical methodology

As mentioned by Bahmani-Oskooee and Hegerty (2007), various studies on exchange rate volatility and trade flows use the bounds testing approach of Pesaran and Shin (1999) and Pesaran et al. (2001), to investigate, whether a volatile exchange rate depresses or supports trade flows.[5] This approach is quite suitable for this issue for several reasons. The time series approximating exchange rate volatility might be stationary, while others (exports or price levels) might be integrated of order one (I(1)). In this case, when there is a mixture of stationary and non-stationary time series or if there is uncertainty about the order of integration, other cointegration techniques such as the Engle/Granger (1987) method or the Johansen (1988, 1991) approach, cannot be employed. Instead, for this scenario, the ARDL bounds testing approach of cointegration is appropriate (Narayan and Smyth 2004), because it does not require a precise pretesting for unit roots of the time series under consideration. However, it is important to check whether no time series is integrated of order two as the bounds testing approach gives critical values for the two border cases where all time series are I(0) and where all time series are I(1). Therefore, it covers the case, when there is a mixture of time series integrated of order one and zero.

We start with an export demand function of the following form (Arize et al. 2000):

$$x_t = \pi_0 + \pi_1 r_t + \pi_2 y_t + \pi_3 v_t + \mu_t. \quad (1)$$

Thereby x denotes bilateral, seasonally adjusted real exports from one euro zone country to the United States in €. The export data is available at monthly frequency from January 1995 onwards. Our sample ends in August 2010.[6] Furthermore, the data is classified by the Standard International Trade Classification (SITC), which gives us the opportunity to avoid the so called “aggregation bias”, as we can split exports by product groups and countries (Bini-Smaghi 1991). The presence of the aggregation bias is confirmed by Ćorić and Pugh (2010). When using disaggregated data, it is more likely to find negative coefficients for exchange rate uncertainty. In our empirical analysis, we will use the main categories (namely product categories 0 to 9) as well as total exports, to investigate whether exchange rate volatility does influence exports to the US.[7] To get real exports, nominal exports are deseasonalized applying the Census-X12 procedure and then deflated with the consumer price

[5] See section 2 for some examples.

[6] Due to the calculation of the moving standard deviation and the lagged variables in the models, our effective sample is 1996M02-2009M10.

[7] The SITC main categories are: 0: food and live animals; 1: beverages and tobacco; 2: crude materials, inedible, except fuels; 3: mineral fuels, lubricants and related materials; 4: animal and vegetable oils, fats and waxes; 5: chemicals and related products; 6: manufactured goods; 7: machinery and transport equipment; 8: miscellaneous manufactured articles; 9: commodities and transactions not classified elsewhere.

index, to get exports in real terms.[8] π_0 is a constant and r represents the bilateral real exchange rate, measured as ratio of national CPI and US CPI multiplied with the bilateral exchange rate in \$ per €.[9] y stands for foreign, therefore US demand, which is approximated by US already seasonally adjusted industrial production. We use industrial production as proxy for foreign demand, because this allows the use of monthly data. Finally, v is our measure of exchange rate volatility. Each variable enters equation (1) in logarithms. μ is an i.i.d. error term.

We employ various definitions of exchange rate variability found in the literature. Furthermore, Ćorić and Pugh (2010) conclude that new innovative measures of exchange rate volatility do not affect the main results. Our first measure, the moving standard deviation of the changes in the nominal exchange rate e, is defined as follows:[10]

$$v_t = \sqrt{\frac{1}{s}\sum_{i=1}^{s}(e_{t+i-1} - e_{t+i-2})^2} \quad (2)$$

We calculate this moving standard deviation for various horizons s (namely 3, 6 and 12 months) to check for robustness and to avoid an arbitrary choice of s (named MSTDEV3, MSTDEV6, MSTDEV12).[11]

Alternatively, we employ an autoregressive conditional heteroskedasticity model (denoted as GARCH(1,1)).[12]

$$\Delta e_t = \tau_0 + \tau_1 \Delta e_{t-1} + u_t \quad (3)$$

$$v_t = \varphi_0 + \varphi_1 u_{t-1}^2 + \vartheta_t \quad (4)$$

Our expectations with respect to the coefficients of equation (1) are as follows: an increase in the real exchange rate should depress exports. Thus π_1 is expected to be negative. Rising demand from the US should stimulate exports. Therefore, π_2 ought to be positive. The sign of π_3 is ambiguous as mentioned in section 2.

To test, whether equation (1) represents a cointegrating relationship, we start with the following autoregressive distributed lags model.

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[8] Unfortunately, we were not able to gather comparable export price indices for all countries, so that we had to use the CPI. This is done by Todani and Munyama (2005) or Grier and Smallwood (2007) as well.

[9] For the time before 1999M01, the exchange rate is converted back in \$ per €.

[10] There is a vigorous debate whether one should use nominal or real exchange rates. McKenzie (1999) concludes that this does not affect the results significantly. Moreover, over short horizons nominal and real exchange rates move closely together.

[11] The choice of s or generally of the measure of exchange rate volatility does not significantly affect the results. Therefore, we will only show the results for the GARCH(1,1) model. Results for the other measures are available upon request.

[12] For example, Todani and Munyama (2005) or Bustaman and Jayanthakumaran (2006) use a GARCH(1,1) volatility measure as well.

$$\Delta x_t = \alpha_0 + \alpha_1 x_{t-1} + \alpha_2 r_{t-1} + \alpha_3 y_{t-1} + \alpha_4 v_{t-1} + \sum_{i=1}^{m} \beta_i \Delta x_{t-i} + \sum_{j=0}^{n} \gamma_j \Delta r_{t-j} + \sum_{k=0}^{p} \delta_k \Delta y_{t-k} + \sum_{l=0}^{q} \gamma_l \Delta v_{t-l} + \varepsilon_t \quad (5)$$

To test for cointegration, one has to test the null hypothesis $\alpha_1 = \alpha_2 = \alpha_3 = \alpha_4 = 0$. For this test, Pesaran et al. (2001) tabulate critical value bounds for the two border cases where all time series are I(0) or otherwise all of them are I(1). Thus, if the test statistic exceeds the upper critical value, we conclude that there exists a long-run relationship among the variables. If it falls in between the critical bounds, the test is inconclusive and if it falls below the lower critical value, there is no evidence for cointegration.

As the test might likely depend upon the lag length (Bahmani-Oskooee and Brooks (1999)), we estimate equation (5) for each country and SITC group with one to twelve lags and perform the F-test.[13] When there is evidence for cointegration, the long-run coefficients are calculated from equation (5) as $\pi_0 = -\frac{\alpha_0}{\alpha_1}$, $\pi_1 = -\frac{\alpha_2}{\alpha_1}$, $\pi_2 = -\frac{\alpha_3}{\alpha_1}$ and $\pi_3 = -\frac{\alpha_4}{\alpha_1}$. Thus we normalize on exports.

Finally, we reformulate equation (5) in error-correction form by replacing the lagged levels in equation (5), with the lagged residuals μ_{t-1} from the export demand equation (1). For the error-correction term, we expect to find a significant negative coefficient, which would support our evidence for cointegration.

4. Results

As we consider eleven countries of the European Monetary Union we start with some general findings, which apply to all countries. For brevity, we resign to present all these results in detail here.[14]

To check if the bounds testing approach is suitable for our investigation as proposed in section 3, unit root tests for the time series under consideration have been performed. The tests strictly rule out the possibility of I(2) evidence. ADF- and PP-tests do reject the null hypothesis of a unit root in the first differenced series without exception. The real exchange rate is found to be basically I(1) and many export series are I(1) as well. For the volatility measures we find mixed evidence of I(0) and I(1). US industrial production is classified as I(0). Generally, the PP-test does reject the null of a unit root more often than the ADF-test. Therefore, as there is evidence for I(1) as well as I(0) time series, the bounds testing approach is appropriate for our concern (Narayan and Smyth (2004)).

[13] We do not have data for all ten SITC main categories for all eleven countries. See tables 2-12 for more details.
[14] The results are available from the author upon request.

Turning to the results, table 1 indicates that the export demand equation (1) represents a cointegrating relationship for the majority of bilateral export relationships. As the bounds testing procedure is sensitive with respect to the lag length, one to twelve lags have been considered in equation 5. Similar to Bahmani-Oskooee and Goswami (2004), cointegration cannot be rejected for at least one lag length at the 10% level of significance for most cases. The SITC categories where cointegration is not supported are dropped from the forthcoming analysis. Furthermore, there are cases where we find cointegration for almost all lag lengths. A closer look at the lag lengths for which cointegration is found, suggests that cointegration is more likely when fewer lags are considered. These findings are quite robust with regard to the choice of the volatility measure. These results are in line with the findings of Ćorić and Pugh (2010). Their meta-regression analysis indicates that the choice of the volatility measure rarely affects the results.

Tab. 1: Results of the bounds testing approach

	AUT	BEL	ESP	FIN	FRA	GER	GRE	IRE	ITA	NED	POR	Total
GARCH(1,1)	5/9	5/10	10/11	8/8	8/11	10/11	9/9	7/9	9/11	9/11	6/8	86/108
MSTDEV3	4/9	6/10	9/11	7/8	8/11	10/11	9/9	6/9	8/11	9/11	5/8	81/108
MSTDEV6	5/9	5/10	10/11	8/8	8/11	10/11	9/9	7/9	9/11	9/11	5/8	85/108
MSTDEV12	6/9	7/10	8/11	7/8	9/11	10/11	9/9	7/9	8/11	10/11	6/8	87/108

Number of cointegrating relationships found for each country and volatility measure

The results of the bounds testing procedure are summarized in table 1. While the first column displays the measure of exchange rate volatility, the following columns show how many cointegating relationships can be found for each country. For example, total exports and eight SITC categories have been analyzed for Austria, which give a total of nine export relationships. According to the chosen volatility measure, we detect four to six cases with evidence of cointegration at least at the 10% level of significance.

Inspecting the other countries, an interesting result is that evidence of cointegration for almost each SITC category is found for Finland, Germany or Greece. On the other hand, for Austria and Belgium there is less evidence of cointegration. Overall, the results are fairly robust with regard to the chosen exchange rate volatility measure. To sum up, we find cointegration in about 75-80% of cases.

When cointegration is evident for more than one lag length, we make use of information criteria to determine the appropriate model. To be exact, we use the Akaike, Schwarz and Hannan-Quinn information criterion to select the lag length. It points out, that all criteria usually favour quite parsimonious models with usually no more than three lags. Thereby, the

Akaike criterion normally prefers models with more lags than the other two criteria. In many cases, the optimal model according to the information criteria corresponds to the model where evidence for cointegration is strongest. In the following, when we turn to the results for each country, we will use the results with the GARCH(1,1) volatility measure as proxy for exchange rate uncertainty. First, because the results are more or less similar regardless of the choice of the volatility measure and second, Grier and Smallwood (2007) give several reasons that volatility proxied by a GARCH model is superior to other measures of exchange rate volatility such as moving standard deviations.

Tab. 2: Austria: Long-run coefficient estimates with diagnostic tests

SITC	constant	r	y	v	ec	Adj. R^2	LM	ARCH	RESET	CUSUM
Total					no cointegration					
0 (1)	19.7 (-6.15)	0.20 (0.98)	-1.58 (-3.75)	-0.32 (-2.07)	-0.82 (-8.01)	0.43	0.41	0.94	0.39	stable
1					no cointegration					
2					no cointegration					
3 (1)	6.88 (0.71)	0.10 (0.10)	2.66 (1.46)	1.56 (2.10)	-0.75 (-6.90)	0.46	0.16	0.95	0.87	stable
4					no data					
5 (1)	-0.94 (-0.24)	-1.80 (-3.90)	4.87 (3.93)	0.73 (2.60)	-0.33 (-5.66)	0.30	0.14	0.11	0.07	stable
6 (1)	-2.12 (-0.88)	-0.01 (-0.06)	3.56 (3.76)	-0.36 (-2.00)	-0.23 (-4.29)	0.33	0.01	0.96	0.19	stable
7					no cointegration					
8 (1)	3.04 (1.23)	0.18 (0.75)	3.03 (4.39)	0.07 (0.40)	-0.51 (-5.58)	0.38	0.00	1.00	0.00	stable
9					no data					

Numbers in brackets under the coefficient estimates are t-values; LM: Breusch-Godfrey test for serial correlation, p-values; ARCH: test of ARCH effects, p-values; RESET: Ramsey's test for functional mis-specification, p-values; CUSUM: test of parameter instability

Tables 2-12 show the long-run coefficients which can be interpreted as elasticities for all export demand models and countries (Arize et al. 2000). The coefficient estimates are usually in line with our a priori expectations. The real exchange rate coefficient turns out to be negative in most cases, while the coefficient of US industrial production is positive in the majority of cases. For the coefficient of exchange rate volatility positive, negative as well as insignificant coefficients can be observed. The error correction term *ec* is significantly negative in all cases, which supports the results of the bounds testing procedure for cointegration. Moreover, the magnitude is comparatively large. This is plausible as exports usually follow no smooth evolution but are rather volatile. Thus, a rapid return to the long-run equilibrium seems reasonable. The adjusted coefficient of determination shows that the fit is of satisfying magnitude as we estimate our models in first differences. The diagnostic tests mainly do not show severe mis-specification.

Tab. 3: Belgium: Long-run coefficient estimates with diagnostic tests

SITC	constant	r	y	v	ec	Adj. R^2	LM	ARCH	RESET	CUSUM
Total					no cointegration					
0 (1)	10.05 (5.86)	-0.55 (-4.01)	1.09 (4.02)	-0.14 (-1.81)	-0.71 (-7.32)	0.41	0.21	0.34	0.01	stable
1					no cointegration					
2 (1)	5.05 (2.10)	-1.37 (-4.57)	1.20 (2.53)	-0.62 (-3.37)	-0.85 (-7.26)	0.51	0.16	0.94	0.52	stable
3 (1)	-17.29 (-3.26)	0.36 (0.74)	6.20 (4.49)	-0.66 (-1.87)	-0.57 (-6.96)	0.46	0.31	1.00	0.05	instable
4					no data					
5					no cointegration					
6					no cointegration					
7 (1)	12.86 (4.38)	-0.97 (-3.81)	0.96 (2.39)	-0.24 (-2.02)	-0.43 (-5.01)	0.29	0.00	0.89	0.95	stable
8 (3)	5.99 (2.33)	-0.69 (-3.10)	2.32 (4.15)	-0.07 (-0.49)	-0.43 (-4.68)	0.52	0.86	0.77	0.01	stable
9					no cointegration					

See tab. 2 for explanations.

Turning now to the country specific results, table 2 displays the long-run elasticities and some diagnostic tests of the export demand equations for Austria. The first column indicates the corresponding SITC trade category. The number in brackets refers to the lag length of the estimated model. As mentioned earlier, we list the export demand equation for total exports to show that aggregation might lead to misleading results as is the case for Austria. While there is no cointegration when using total exports, disaggregation gives five cointegrating relationships. Turning to the coefficient estimates (t-values in brackets underneath), for our volatility measure *v* we find significant negative coefficients for SITC categories 0 and 6, while there is a significant positive coefficient for category 3 and 5. The real exchange rate *r* does not seem to be a major factor for Austrian exports to the US. Only for chemicals and related products (SITC 5) we get a significant coefficient. In contrast, US demand *y* is of more importance. Rising US demand stimulates Austrian exports except food and live animals (SITC 0). The long-run elasticity is always greater than unity which indicates that Austrian exports to the US are income elastic.

The picture for Belgium is somewhat different. Exchange rate volatility exerts a negative impact on all export categories which is significant in four of five cases. But its elasticity is small compared to the coefficients of *r* and *y*. Furthermore, a rise of the real exchange rate depresses exports in four of five cases. US demand is a major factor for Belgian exports as well as it significantly increases Belgium's exports more than proportional for each SITC category except number 7 (machinery and transport equipment).

Tab. 4: Spain: Long-run coefficient estimates with diagnostic tests

SITC	constant	r	y	v	ec	Adj. R^2	LM	ARCH	RESET	CUSUM
Total (2)	10.18 (3.97)	-0.22 (-2.20)	1.75 (4.33)	-0.15 (-2.44)	-0.56 (-5.19)	0.48	0.27	0.52	0.63	stable
0 (1)	12.41 (5.91)	-1.05 (-5.97)	1.07 (4.08)	0.03 (0.50)	-0.63 (-6.85)	0.37	0.05	0.84	0.83	stable
1 (1)	3.84 (3.25)	-0.48 (-4.23)	2.78 (6.45)	0.15 (2.30)	-0.79 (-7.57)	0.41	0.12	0.89	0.06	stable
2					no cointegration					
3 (1)	-8.41 (-1.38)	1.16 (2.07)	5.02 (3.72)	-0.15 (-0.40)	-0.70 (-7.57)	0.43	0.37	1.00	0.02	stable
4 (1)	5.04 (1.27)	0.85 (2.26)	1.98 (2.50)	-0.10 (-0.43)	-0.39 (-5.01)	0.42	0.04	0.09	0.19	instable
5 (1)	-5.38 (-1.55)	0.60 (1.82)	4.75 (3.73)	0.03 (0.12)	-0.30 (-4.33)	0.27	0.00	0.02	0.00	stable
6 (1)	13.66 (3.62)	-0.78 (-3.13)	0.33 (0.88)	-0.43 (-3.25)	-0.36 (-4.80)	0.32	0.00	0.01	0.00	stable
7 (6)	4.85 (1.34)	0.02 (0.16)	2.20 (3.51)	-0.41 (-2.66)	-0.68 (-4.94)	0.58	0.77	0.77	0.89	stable
8 (2)	18.35 (3.92)	-1.47 (-3.74)	-0.56 (-1.74)	-0.30 (-2.22)	-0.32 (-4.04)	0.46	0.51	0.08	0.08	stable
9 (1)	19.07 (6.28)	-1.89 (-6.38)	-1.19 (-2.94)	-0.34 (-2.67)	-1.01 (-8.30)	0.56	0.18	0.73	0.56	stable

See tab. 2 for explanations.

Table 4 shows the results for Spain. The first impression is that we find more cointegrating relationships than for Austria and Belgium. When the coefficient of exchange rate volatility is significant, it is negative in five of six cases. The influence of the real exchange rate varies across categories but US demand mostly increases Spanish exports.

For Finland we find cointegration in each case. As for Spain, exchange rate uncertainty depresses exports significantly in five trade categories. Just for miscellaneous manufactured articles (SITC 8), we find a significant positive coefficient. The real exchange rate influences exports negatively in every case but it is insignificant in some cases. US demand stimulates exports when there is a significant coefficient but the elasticities are rather low compared with other countries. Thus, Finland might not benefit that much from rising US demand like other countries of the euro zone.

Tab. 5: Finland: Long-run coefficient estimates with diagnostic tests

SITC	constant	r	y	v	ec	Adj. R^2	LM	ARCH	RESET	CUSUM
Total (1)	14.44 (6.39)	-0.63 (-3.18)	0.78 (2.44)	-0.17 (-1.35)	-1.14 (-9.11)	0.51	0.91	0.96	0.01	stable
0 (6)	14.99 (3.82)	-1.87 (-4.14)	0.20 (0.40)	0.11 (0.35)	-0.82 (-4.91)	0.51	0.88	0.10	0.56	stable
1 (1)	19.96 (3.69)	-1.95 (-3.46)	-0.98 (-1.10)	0.23 (0.67)	-0.70 (-7.36)	0.39	0.19	1.00	0.02	stable
2 (1)	-18.43 (-2.18)	-0.74 (-0.84)	2.59 (1.55)	-2.91 (-4.22)	-0.34 (-5.01)	0.29	0.13	0.78	0.60	stable
3					no data					
4					no data					
5 (1)	16.13 (4.20)	-1.96 (-4.12)	-0.48 (-1.04)	-0.35 (-1.93)	-0.40 (-5.39)	0.41	0.33	0.83	0.12	stable
6 (1)	14.38 (5.36)	-1.05 (-4.32)	0.22 (0.64)	-0.36 (-2.46)	-0.54 (-6.57)	0.24	0.31	0.04	0.18	stable
7 (1)	19.32 (5.95)	-1.24 (-3.86)	-0.78 (-1.61)	-0.43 (-2.22)	-1.17 (-9.40)	0.51	0.85	0.87	0.15	stable
8 (1)	6.08 (2.47)	-0.37 (-1.49)	2.97 (-4.64)	0.55 (-3.03)	-0.43 (-5.42)	0.41	0.26	0.09	0.13	stable
9					no data					

See tab. 2 for explanations.

Tab. 6: France: Long-run coefficient estimates with diagnostic tests

SITC	constant	r	y	v	ec	Adj. R^2	LM	ARCH	RESET	CUSUM
Total (0)	14.17 (4.60)	-1.04 (-4.31)	1.23 (3.86)	-0.17 (-2.68)	-0.44 (-5.09)	0.45	0.00	0.59	0.74	stable
0 (2)	11.74 (3.96)	-0.48 (-2.57)	1.18 (2.68)	0.06 (0.55)	-0.39 (-4.24)	0.40	0.30	0.82	0.02	stable
1 (1)	9.91 (4.44)	-0.90 (-4.26)	1.73 (3.79)	-0.04 (-0.42)	-0.41 (-5.22)	0.35	0.08	1.00	0.34	stable
2 (1)	4.30 (2.73)	-0.39 (-2.58)	1.83 (5.08)	-0.45 (-4.58)	-0.65 (-7.19)	0.38	0.79	0.96	0.00	stable
3 (1)	-22.61 (-3.02)	1.66 (2.42)	7.71 (4.24)	-0.36 (-0.83)	-0.50 (-5.86)	0.39	0.58	0.99	0.00	stable
4 (1)	7.42 (1.31)	-0.98 (-1.71)	0.21 (0.22)	-0.67 (-1.91)	-0.52 (-5.72)	0.42	0.11	0.00	0.00	instable
5 (2)	7.11 (2.77)	-0.55 (-2.36)	2.44 (3.59)	-0.09 (-0.66)	-0.31 (-4.71)	0.48	0.04	0.25	0.17	stable
6					no cointegration					
7					no cointegration					
8 (2)	10.02 (3.64)	-0.78 (-3.55)	1.74 (3.67)	-0.09 (-1.16)	-0.39 (-4.09)	0.46	0.13	0.38	0.12	stable
9					no cointegration					

See tab. 2 for explanations.

For France the coefficient of exchange rate volatility is significant in only three cases where always a negative impact on exports to the US arises. The real exchange rate depresses exports in every category but SITC 3. US industrial production raises exports where the elasticity is usually above unity. Comparing the coefficients of US demand across SITC categories and countries, one can see that rising US demand stimulates especially exports of category 3 (mineral fuels, lubricants and related materials). This applies to Belgium, Spain, France, Germany, Italy and the Netherlands.

The influence of exchange rate uncertainty on German exports is negative in most significant cases. Furthermore, a real appreciation of the euro against the greenback will depress exports to the US. In case of significance, the relevant coefficient is always negative. On the other hand, rising US demand will increase German exports in every SITC group. Therefore, Germany might benefit from a booming US economy more than other countries of the euro zone.

Tab. 7: Germany: Long-run coefficient estimates with diagnostic tests

SITC	constant	r	y	v	ec	Adj. R^2	LM	ARCH	RESET	CUSUM
Total (2)	9.83 (4.01)	-0.60 (-4.02)	2.30 (4.16)	-0.18 (-2.58)	-0.44 (-4.88)	0.50	0.63	0.10	0.44	stable
0 (1)	12.64 (5.99)	0.05 (0.48)	1.11 (4.67)	0.09 (1.18)	-0.67 (-6.98)	0.39	0.30	0.95	0.40	stable
1 (1)	5.01 (3.25)	0.16 (1.07)	2.59 (5.37)	0.10 (0.94)	-0.72 (-6.85)	0.50	0.06	0.57	0.00	stable
2					no cointegration					
3 (1)	-32.41 (-5.26)	-0.07 (-0.14)	10.29 (6.65)	0.02 (0.05)	-0.87 (-8.32)	0.51	0.13	1.00	0.00	stable
4 (1)	-1.55 (-0.33)	-0.22 (-0.45)	2.94 (2.88)	-0.11 (-0.31)	-0.53 (-6.00)	0.31	0.53	0.50	0.23	stable
5 (2)	9.93 (4.88)	-0.32 (-2.18)	2.60 (4.81)	0.32 (2.78)	-0.63 (-5.71)	0.46	0.45	1.00	0.33	stable
6 (1)	6.84 (3.19)	-0.16 (-1.30)	2.31 (4.34)	-0.25 (-2.75)	-0.36 (-5.08)	0.40	0.11	0.75	0.63	instable
7 (2)	9.21 (4.11)	-0.80 (-4.30)	2.17 (4.24)	-0.29 (-3.40)	-0.48 (-5.14)	0.51	0.17	0.44	0.07	stable
8 (2)	9.75 (4.53)	-0.39 (-3.86)	2.15 (4.66)	0.06 (0.91)	-0.54 (-5.02)	0.42	0.04	0.92	0.27	stable
9 (1)	-2.83 (-0.35)	0.46 (0.61)	1.94 (1.46)	-1.56 (-2.85)	-0.17 (-4.48)	0.19	0.04	0.19	0.00	stable

See tab. 2 for explanations.

For Greece, the evidence of cointegration is strong as well. An increase in the real exchange rate will depress exports in all cases, when there is a significant coefficient and US demand stimulates exports in the majority of cases. We find three significant coefficients for exchange rate uncertainty, two positive and one negative.

The results for Ireland as shown in table 9 generally fit theoretical suggestions. The real exchange rate coefficient is negative when significant and US industrial production is positive when significant. The coefficient estimates for exchange rate volatility are positive in three cases of significance. This picture is somewhat different from the other countries where we found at least some cases with a negative relationship.

Table 10 delivers the results for Italy. The coefficient estimates for *r* and *y* confirm theory. Just like Germany, Italy might benefit from growing US demand more than other euro zone countries because *y* influences each export category significantly positive and even more than proportional. The coefficient of *v* is negative in all but one case. Moreover, it is significant in five cases.

The coefficient estimates for the Netherlands mostly correspond to our expectations. Exchange rate volatility seems not to be a major factor for Dutch exports to the US. There is only one significant coefficient which is negative.

Tab. 8: Greece: Long-run coefficient estimates with diagnostic tests

SITC	constant	r	y	v	ec	Adj. R^2	LM	ARCH	RESET	CUSUM
Total (6)	9.90 (4.07)	-0.24 (-1.94)	1.46 (4.16)	-0.07 (-0.65)	-1.23 (-5.86)	0.60	0.28	0.54	0.53	stable
0 (1)	5.10 (2.42)	-0.12 (-0.65)	2.05 (3.48)	-0.06 (-0.46)	-0.41 (-5.20)	0.28	0.51	0.96	0.00	instable
1 (1)	19.31 (3.39)	-1.06 (-2.21)	-1.35 (-1.34)	-0.28 (-0.82)	-0.95 (-8.78)	0.47	0.07	0.02	0.71	stable
2 (1)	2.64 (0.37)	-0.61 (-0.91)	1.61 (1.13)	-0.48 (-1.01)	-0.58 (-5.84)	0.48	0.40	0.02	0.03	stable
3	no data									
4 (12)	-0.76 (-0.22)	0.25 (1.06)	1.27 (2.15)	-1.10 (-2.80)	-0.71 (-4.06)	0.59	0.05	0.97	0.62	stable
5 (1)	3.15 (1.23)	-0.63 (-2.56)	2.25 (3.71)	0.00 (0.02)	-0.79 (-7.71)	0.40	0.35	0.01	0.01	stable
6 (1)	8.64 (2.09)	-0.43 (-1.25)	0.77 (1.06)	-0.54 (-2.20)	-0.58 (-5.94)	0.47	0.34	0.00	0.76	instable
7 (1)	7.18 (1.77)	0.17 (0.49)	2.80 (3.48)	0.82 (3.06)	-0.87 (-8.29)	0.40	0.10	0.64	0.91	stable
8 (1)	25.61 (4.20)	-2.61 (-4.40)	-2.14 (-2.52)	-0.09 (-0.38)	-0.49 (-5.62)	0.30	0.01	1.00	0.00	stable
9	no data									

See tab. 2 for explanations.

Tab. 9: Ireland: Long-run coefficient estimates with diagnostic tests

SITC	constant	r	y	v	ec	Adj. R^2	LM	ARCH	RESET	CUSUM
Total (1)	6.13 (2.23)	-1.05 (-3.69)	3.77 (4.08)	0.50 (3.44)	-0.35 (-4.86)	0.36	0.01	0.27	0.00	stable
0 (1)	10.08 (2.92)	-0.74 (-2.70)	1.12 (1.90)	-0.01 (-0.09)	-0.57 (-6.33)	0.40	0.08	0.36	0.43	stable
1 (1)	9.71 (2.19)	-1.22 (-3.06)	1.07 (1.31)	-0.21 (-1.07)	-0.42 (-5.06)	0.31	0.16	0.91	0.00	stable
2 (1)	-22.76 (-2.42)	-1.69 (-2.33)	7.01 (3.28)	-0.54 (-1.48)	-0.33 (-4.41)	0.33	0.11	0.04	0.45	stable
3	no data									
4	no data									
5 (1)	-2.43 (-0.79)	-1.15 (-3.81)	5.59 (4.98)	0.62 (3.70)	-0.46 (-5.50)	0.38	0.09	0.97	0.03	stable
6	no cointegration									
7 (1)	12.68 (3.18)	-2.11 (-4.03)	1.41 (2.25)	0.03 (0.23)	-0.33 (-4.32)	0.34	0.03	0.06	0.73	instable
8	no cointegration									
9 (1)	31.13 (3.16)	0.28 (0.51)	-1.09 (-0.83)	1.24 (2.84)	-0.24 (-3.90)	0.26	0.59	0.29	0.08	stable

See tab. 2 for explanations.

Tab. 10: Italy: Long-run coefficient estimates with diagnostic tests

SITC	constant	r	y	v	ec	Adj. R^2	LM	ARCH	RESET	CUSUM
Total (1)	13.26 (7.56)	-0.86 (-7.19)	1.38 (6.87)	-0.18 (-6.09)	-0.88 (-8.15)	0.47	0.80	0.87	0.18	stable
0 (1)	10.45 (5.25)	-0.19 (-1.78)	1.39 (4.13)	-0.05 (-1.10)	-0.47 (-5.93)	0.40	0.06	0.02	0.34	stable
1	no cointegration									
2 (1)	2.13 (1.33)	-0.88 (-4.50)	2.41 (5.21)	-0.31 (-3.17)	-0.76 (-7.16)	0.45	0.29	0.89	0.36	stable
3 (1)	-41.73 (-3.63)	0.06 (0.06)	11.45 (4.42)	-0.48 (-0.97)	-0.52 (-6.02)	0.38	0.00	0.04	0.16	stable
4 (1)	6.21 (1.52)	0.07 (0.21)	1.93 (2.71)	-0.19 (-1.11)	-0.23 (-4.32)	0.20	0.37	0.03	0.65	stable
5 (1)	9.58 (6.71)	-1.09 (-7.04)	1.89 (6.81)	0.00 (-0.08)	-1.00 (-8.97)	0.48	0.53	1.00	0.00	stable
6 (1)	8.03 (3.65)	-0.94 (-4.02)	1.80 (3.90)	-0.37 (-4.42)	-0.34 (-5.07)	0.28	0.01	0.15	0.99	stable
7 (1)	12.52 (8.73)	-0.59 (-6.79)	1.41 (7.48)	-0.11 (-3.50)	-1.12 (-9.79)	0.58	0.79	0.02	0.01	stable
8	no cointegration									
9 (1)	-14.16 (-2.74)	-0.53 (-1.14)	4.41 (4.06)	-0.73 (-2.76)	-0.74 (-6.85)	0.41	0.02	0.08	0.00	stable

See tab. 2 for explanations.

Tab. 11: Netherlands: Long-run coefficient estimates with diagnostic tests

SITC	constant	r	y	v	ec	Adj. R^2	LM	ARCH	RESET	CUSUM
Total (1)	5.19 (2.31)	0.02 (0.12)	3.13 (3.68)	-0.02 (-0.13)	-0.26 (-4.55)	0.27	0.00	0.01	0.72	stable
0 (2)	10.28 (3.49)	-1.06 (-3.70)	1.69 (3.24)	0.16 (1.03)	-0.36 (-4.29)	0.29	0.12	0.41	0.21	instable
1 (1)	4.30 (1.24)	-0.54 (-1.45)	1.99 (1.97)	-0.56 (-2.24)	-0.19 (-4.41)	0.30	0.18	0.01	0.02	stable
2 (3)	8.55 (3.55)	-0.63 (-3.25)	1.78 (3.67)	0.01 (0.09)	-0.49 (-5.16)	0.38	0.13	0.77	0.29	stable
3 (1)	-45.88 (-3.09)	3.01 (1.94)	13.99 (4.03)	0.68 (0.67)	-0.41 (-5.14)	0.27	0.00	0.00	0.06	instable
4 (1)	12.85 (1.60)	0.34 (0.40)	0.50 (0.31)	0.18 (0.30)	-0.30 (-4.43)	0.21	0.02	0.01	0.00	instable
5	no cointegration									
6 (2)	7.53 (2.50)	0.15 (0.70)	1.78 (3.42)	-0.24 (-1.39)	-0.44 (-4.35)	0.54	0.37	0.40	0.53	stable
7 (2)	4.46 (2.02)	-0.46 (-2.21)	3.18 (3.83)	0.06 (0.47)	-0.36 (-4.05)	0.46	0.10	0.47	0.64	stable
8	no cointegration									
9 (1)	-28.06 (-2.73)	-4.83 (-3.62)	9.29 (3.51)	0.00 (0.00)	-0.30 (-4.66)	0.19	0.02	0.47	0.00	instable

See tab. 2 for explanations.

Tab. 12: Portugal: Long-run coefficient estimates with diagnostic tests

SITC	constant	r	y	v	ec	Adj. R^2	LM	ARCH	RESET	CUSUM
Total (1)	6.88 (2.51)	-0.66 (-2.56)	1.93 (2.70)	-0.31 (-1.74)	-0.31 (-4.15)	0.23	0.01	0.17	0.96	stable
0 (1)	14.72 (4.54)	0.01 (0.05)	-0.14 (-0.31)	0.00 (-0.01)	-0.67 (-7.93)	0.40	0.03	0.23	0.10	stable
1 (1)	11.03 (7.28)	-0.66 (-5.07)	0.60 (2.60)	-0.15 (-2.16)	-1.00 (-9.03)	0.49	0.03	0.36	0.80	stable
2 (1)	5.10 (0.81)	1.59 (2.31)	1.49 (1.23)	-0.20 (-0.53)	-0.27 (-4.24)	0.24	0.01	0.83	0.89	stable
3	no data									
4	no data									
5 (3)	17.65 (4.17)	-2.07 (-4.14)	0.64 (1.00)	0.76 (3.17)	-0.62 (-5.23)	0.44	0.09	1.00	0.84	stable
6 (1)	12.88 (3.48)	-1.02 (-3.06)	0.36 (0.81)	-0.39 (-2.66)	-0.27 (-3.93)	0.34	0.01	0.45	0.01	stable
7	no cointegration									
8	no cointegration									
9	no data									

See tab. 2 for explanations.

Lastly, we have got the results for Portugal in table 12. Once again, estimates for *r* and *y* match theory. Exchange rate volatility affects exports negatively more often than positively.

Finally, table 13 summarizes our results with respect to the influence of exchange rate variability on exports from euro zone countries to the United States. We have estimated a total of 86 cointegrating relationships. Altogether we find 56 negative and 30 positive coefficients for exchange rate volatility, while 33 respectively 10 are significant. To sum up, for the eleven countries of the euro zone, exchange rate uncertainty does more often harm exports to the US than it supports them. But in 50% of cases, there is no significant influence.

Tab. 13: Coefficients of exchange rate volatility

	positive	negative	Σ
significant	10	33	43
insignificant	20	23	43
Σ	30	56	86

Taking a closer look for which SITC categories exchange rate volatility exerts a significant impact on exports, we find that it is especially categories 6 (manufactured goods) and 7 (machinery and transport equipment) which are affected negatively. These two export categories account for more than two thirds of total exports. Thus, the finding of significant negative coefficients for these categories might be of special relevance for the political authorities.

Altogether, our results suggest a negative effect of exchange rate volatility on trade, but the influence is rather small compared to the elasticities of the real exchange rate or US demand. The estimated long-run elasticities are usually far below unity while exports from the countries of the euro zone to the US respond to US demand more than proportional. Nonetheless, neglecting the influence of exchange rate risk might cause problems, especially in a currency union, when the influence of exchange rate volatility differs across individual countries and there can per definition be only one foreign exchange policy. For example, we find that exchange rate volatility does not affect Dutch exports to the US while Italian exports are depressed in many cases. Contrary, Irish exports even seem to benefit from increasing volatility.

5. Conclusion

We have investigated whether exchange rate volatility proxied by measures of moving standard deviations or a GARCH(1,1) model is crucial for exports to the United States for eleven countries of the European Monetary Union. Referring to economic theory, there is no consensus whether there should be a positive or a negative connection. Furthermore, our literature review shows that there are cases where a positive, a negative or no significant relationship is found. But no study investigates this question for the countries of the euro zone.

Applying the ARDL bounds testing approach for cointegration of Pesaran and Shin (1999) and Pesaran et al. (2001), we investigate this issue for disaggregated SITC export categories. Using a simple export demand model, we find evidence of cointegration in more than 75% of cases. The models are generally robust to the choice of the volatility measure and the coefficient estimates usually support a priori expectations. The real exchange rate reduces exports while US demand stimulates them. Furthermore, exports react more than proportional to rising US demand. Thereby, countries like Germany or Italy might benefit from rising US demand more strongly than others.

When taking a look at the elasticities of exchange rate volatility, we find mixed results. There are positive, negative and insignificant coefficients. Nonetheless, our results indicate that it is most likely that exchange rate variability depresses exports. This finding is in line with Ćorić and Pugh (2010) who draw the same conclusion. Especially the main SITC trading categories 6 and 7 are affected negatively. With regard to the magnitude of the exchange rate volatility coefficients it becomes obvious that these are generally rather small compared with the other long-run elasticities of the real exchange rate or US industrial production. Therefore, analogously one might suppose that the eleven countries under study could have benefitted from accession to EMU as it has erased any exchange rate risk between these countries. Additionally, as there are countries for which exchange rate volatility is of more importance, this can cause different opinions and thereby disagreement about the extent to which the political authorities should take care of the exchange rate. On the national level, the same reasoning applies as well as there are sectors that are affected positively and others that suffer from exchange rate uncertainty.

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