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DO EXPORT PRICE ELASTICITIES SUPPORT TENSIONS IN CURRENCY MARKETS? EVIDENCE FROM CHINA AND SIX OECD COUNTRIES

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Do export price elasticities support tensions in currency markets? Evidence from China and six OECD countries•

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Abstract The empirical literature on trade imbalances does not make currency tensions easy to understand, because tensions across traders originate from the assumption that export-price elasticity is high. This paper provides new evidence by analysing the export-behaviour of China, France, Germany, Italy, Japan, UK, and the USA from 1990 to 2012. Estimates of export-price elasticities have been made using panel data techniques for non-stationary data. Long run relationships are stable to any structural break and indicate that exports are heavily dependent on world income, with long run income elasticity significantly higher than unity in many cases (China, Japan, Germany, UK and USA). Conversely, exports are price inelastic for most of the countries in the sample, in both the long and short runs. The exception is France, whose exports in the long run would increase by 2 percent if the country experienced a 1 percent depreciation of its real exchange rate.

Keywords: export elasticity, competitive Devaluation, currency wars, panel data
JEL classification: C23, F10, F17, F37, P33

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Abstract The empirical literature on trade imbalances does not make currency tensions easy to understand, because tensions across traders originate from the assumption that export-price elasticity is high. This paper provides new evidence by analysing the export-behaviour of China, France, Germany, Italy, Japan, UK, and the USA from 1990 to 2012. Estimates of export-price elasticities have been made using panel data techniques for non-stationary data. Long run relationships are stable to any structural break and indicate that exports are heavily dependent on world income, with long run income elasticity significantly higher than unity in many cases (China, Japan, Germany, UK and USA). Conversely, exports are price inelastic for most of the countries in the sample, in both the long and short runs. The exception is France, whose exports in the long run would increase by 2 percent if the country experienced a 1 percent depreciation of its real exchange rate.

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1. Introduction

The analysis of trade flows reveals many cases of national current account imbalances. The USA was a net exporter until 1975, when its trade surplus accounted for 1.07% of GDP; it then experienced rapidly growing trade deficits and since the 1990s it is the world's largest debtor nation. In 2000 Germany was a net importer with a trade deficit of 1.83% of GDP; then in just a few years it became an export-oriented economy and by 2013 had a trade surplus of 7.58% of GDP. China ran a current account surplus averaging 4.24% of GDP from 1998 until 2013, peaking at over 10% in 2007.

Sizable and persistent national trade surpluses in large economies generate global imbalances and tensions in world markets: there is serious concern over exporters managing their national currency to gain from competitive devaluation strategies, such as quantitative easing, currency intervention and capital controls. Disputes between national interests can turn into currency wars, when trading partners accuse each other of unfair practices in manipulating their exchange rates in order to boost exports and curb imports.

Although the most prominent recent case is China - much criticised by the USA - Germany, Japan and the UK have also been accused of manipulating their real exchange rates. Japan and the UK used quantitative easing in order to counter the current recession (Gagnon, 2013; Joyce et al.

2011), and, according to the U.S. Treasury, Germany's low level of investment and high savings rate contributed substantially to the Eurozone crisis, which is characterised by increasing trade troubles for EU peripheral countries and huge surpluses for Germany. Therefore, in order to rebalance the global economy, Germany should promote domestic investments and demand. Paul Krugman agrees with the U.S. Treasury and argues that Germany's insistence on an export-driven economic model:

> "has… been a bad thing for the rest of the world. It's simply arithmetic. Since southern Europe has been forced to end its deficits while Germany hasn't reduced its surplus, Europe as a whole is running large trade surpluses, helping to keep the world economy depressed." (www.forbes.com, "If Paul Krugman Supports The U.S. Treasury's Currency Wars, Then Treasury Must Be Confused" 14/11/2013)

After reading the literature on national current account imbalances and currency tensions, one would expect that controlling exchange rates is a feasible policy to improve trade balance: tensions on currency markets are understandable if devaluations lead to substantial increases in exports. Differently phrased, exports are expected to be price elastic. This expectation, however, is not supported by the evidence emerging from the related literature: empirical studies present a wide range of results, many of which estimate long and short run export price elasticities less than unity.

While this heterogeneity in results poses some questions about the impact of real devaluation on exports, it also suggests that price competitiveness remains one of the most difficult, controversial and intriguing issues in international trade. The literature refers to contributions made some time ago, for example by Orcutt (1950), by Houthakker and Magee (1969) and by Kravis and Lipsey (1978). There are surveys of initial papers in Stern et al. (1976), that cite 130 articles from the period 1960-1975, in Goldstein and Khan (1985) and in Sawyer and Sprinkle (1996), who reviewed approximately 50 articles. While these review papers demonstrate the wide range of values of price elasticities estimated by various scholars over four decades, it is noteworthy that the picture does not change with more recent studies. Limiting attention to the price elasticities of aggregate trade-flows, a number of authors show that exports are price inelastic. For instance, Anaraki (2014) uses a Keynesian model and quarterly data over the 2001–2010 period and finds that a 10 percent Euro devaluation against the major currencies (yuan, dollar and yen) would increase the Eurozone's exports to China by 3.4 percent, to the USA by 2.4 percent and to Japan by 1.9 percent. Algieri (2011) reports that the price elasticities of the exports of France, Italy, Japan, Netherland, Spain, UK and USA are, over the period 1978-2009, rather small (in the range -0.3/-0.8). France's total exports are also found to be price inelastic in Magnani et al. (2013), who simulate the macroeconomic effect of a 10 percent Euro devaluation and found that French total exports would increase by 3.2 percent. Similarly the price elasticity of the total exports of the Eurozone countries was found to be low by Bayoumi et al. (2011) and in Chen et al. (2012), at -0.6 over the 1980-2009 period and -0.46 over the 1990-2009 period, respectively. Ketenci and Uz (2011) looked at EU bilateral trade flows over the 1980-2007 period and found an export price elasticity in the range -0.08 to -0.64. In Thorbecke and Kato (2012) the price elasticity of Germany's overall exports was found to be -0.6, with higher values for consumption than for capital goods and for Eurozone than for extra-Eurozone exports. Thorbecke and Kato (2012) focus on total Japanese exports to 17 trading partners over the period 1988-2009 and find that those exports are price inelastic, although when they restrict the analysis to consumption products they obtain a unitary long run export price elasticity. Crane et al. (2007), using quarterly data over the 1981–2006 period, show that the price elasticity of Japanese exports is -0.34. Yao et al. (2013) looked at total Chinese exports from 1992 to 2006 and, even after controlling for the increase of product-variety, they find a short run price elasticity of -0.65.

The conclusion that can be drawn from this discussion is that total exports are found to be price inelastic in several studies. This outcome seems to hold whichever country and time period are examined and whichever methods are used in estimating the export equation (these range from OLS to more sophisticated time-series tools). Thus the results from macro-analyses do not make currency tensions easy to understand, because they originate from the controversial assumption of a high exports sensitivity to price competitiveness. Indeed, if the estimates conducted at macro-level are reliable, then competitive devaluations will not lead to increased current account surpluses in the 'aggressive' countries and, therefore, will not penalize trading partners. If this holds, tension between trading actors is not justified (on economic grounds alone, anyway), because the changes of exchange rate do not make much difference to exports.

This paper contributes to the discussion on trade elasticities in two ways. Firstly, it proposes an updated analysis of the export behaviour of six OECD countries (France, Germany, Italy, Japan, UK and USA, henceforth the 6-OECDs) and China. The 6-OECDs have played a dominant role in international trade for some time, while China has become a big player since it joined the World Trade Organization in 2001. Total exports from these countries are analyzed from 1990 to 2012, a period which saw a number of tensions in the global currency market.

Secondly, estimates of export price elasticities have been made by using panel data techniques for non-stationary data. While advances in the econometrics of these methods began in the early 1990s and most empirical research is from the last decade, few researchers have used them to estimate trade elasticities (Kubota 2009; Jovanovic 2012). Within this analytical framework, we use an export equation derived from the imperfect substitutes model proposed by Goldstein and Khan (1985). After detecting for non-stationarity and co-integration of time-series, the empirical analysis is carried out by applying the Pooled Mean Group estimator (PMG) developed by Pesaran, Shin and Smith (1999) and the Mean Group estimator (MG) of Pesaran, Smith and Im (1996). The export equations are specified with an error correcting mechanism and allow for full country heterogeneity of short run price elasticity and of the dynamics towards the long run equilibrium, which is assumed to be common across countries in the PMG procedure and different for each country in the MG method. The analysis is enriched by testing for structural breaks occurring during the period analyzed.

Results indicate that exports are heavily dependent on world income, with long run income elasticity significantly higher than unity in many cases (China, Japan, Germany, UK and USA). Conversely, exports are price inelastic for most of the countries in the sample, both in the long and in the short run. The exception is France, whose exports, in the long run would increase by 2 percent if the country experienced a 1 percent depreciation of its real exchange rate.

The remainder of the paper is structured in 5 sections. Section 2 presents the empirical setting and describes the data. Section 3 presents and discusses the results, while section 4 draws the conclusions.

2. Empirical setting and data

The empirical setting of this study refers to the imperfect substitutes model proposed by Goldstein and Khan (1985). The major assumption of the model is that neither imports nor exports are perfect substitutes for domestic goods. In the vein of much research on this subject, we proceed using aggregate data. In this respect, according to a great deal of literature, the export demand is specified as a function of the real exchange rate and of the income in the 'Rest of the World' (RoW):[2]

$$\log X_{it} = \alpha_i + \beta_1 \log REX_{it} + \beta_2 \log Y^{w}_{it} + u_{it} \qquad (1)$$

[2] See for instance, Hamori and Yin (2011), Ketenci and Uz (2011), Shigeyuki and Yoichi (2009), Caporale and Chui (1999), Senhadji and Montenegro (1999), Bahmani-Oskooee and Niroomand (1998), Sawyer and Sprinkle (1996), Thorbecke (2011).

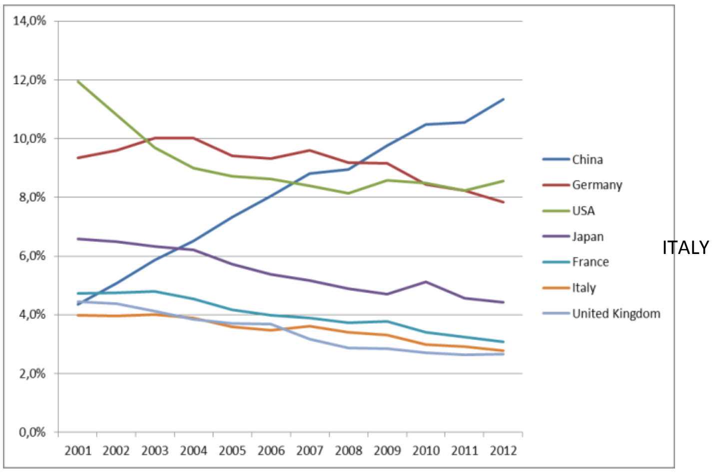
where X_{it} refers to the total national exports of country *i* at time *t*, REX_{it} is the relative price variable gauged by the real exchange rate of country *i* at time t, and foreign demand is measured by world income Y_t^w. Given the log-linear form of equation (1), β_1 is the exports elasticity to the real exchange rate and β_2 is the exports elasticity to foreign income. Based on the theory, it is expected that β_1 is negative, implying an increase in demand of exported goods when national currency depreciates (an increase in *REX* stands out for real appreciation, i.e., loss of price competitiveness). The parameter β_2 is expected to be positive, indicating that exports rise as world income increases. For each exporter, *REX* is constructed as a weighted average of the real exchange rates against each trade partner and is based on the Consumer Price Index.[3] Data are from Datastream and are expressed on a quarter basis covering the years between 1990:Q1 and 2012:Q1. All the time series are in real terms (2005 is the base year) and are seasonally adjusted.

The sample of countries comprises China and the 6-OECDs, which are France, Germany, Italy, Japan, the UK, and the USA. While the six OECDs are important traders and have been for a long time, China represents the matter of interest in the current debate on devaluation and exchange rate misalignments because of its growing role as a production hub and exporter. As can be observed from figure 1, the sample of countries absorbs much of the world's exports market, as their total export shares are around 45 per cent in the 2-year period from 2001 to 2002 and about 40 per cent in 2010-2012. Data also highlight the impressive pattern of Chinese export shares, which increased by about seven percentage points, moving from 4.3 per cent between 2001 and 2002 to 11.3 per cent in 2012. It is interesting to note that market shares are decreased for the other exporters (e.g. the USA's market share is 8.6 per cent in 2012, but 11.9 per cent in 2001), while Germany's market share amounts to 8 per cent at the end of the period. What the data clearly highlight is that China became, in just a few years, a first world exporter.

Figure 2 displays the plot of the time series of exports and real exchange rates over the whole time period from 1990 to 2012, analysed using econometric modelling. Although a strong positive increase is revealed for the value of exports in each country, the highest increase belongs to China, followed by the UK and the USA. Another common result across countries is the drop of exports at the time of the 2008 financial crisis. Exports reduced much more in Italy and Japan than in other countries. All countries observed a recovery of exports after 2008. From a statistical perspective, Figure 2 clearly highlights that exports in the time series exhibit a non-stationary pattern. The same does not appear for the real exchange rate, which is a fact that deserves more statistical attention (see § 3). In the case of REX, each time series highlights much more instability along the trend than a strict trend pace itself. It becomes very interesting to evaluate the effects of this variability on export behaviour. It is an issue that will be addressed in the following paragraphs when measuring the short run relationship between exports and exchange rates.

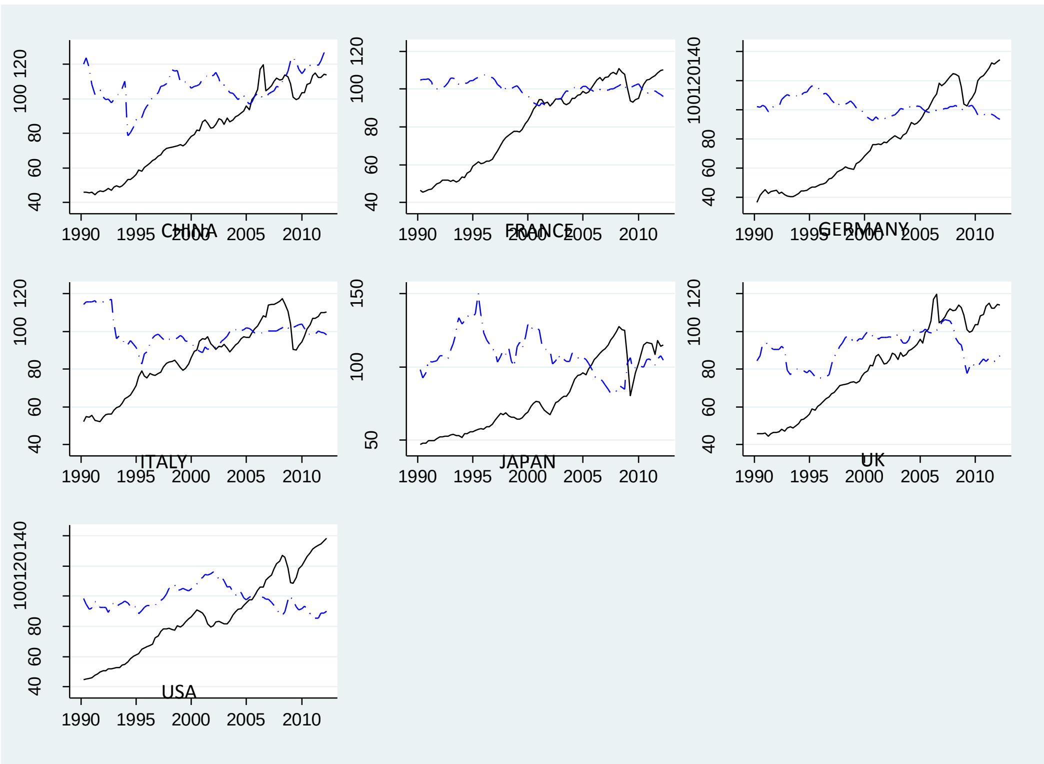
[3] The real exchange rate is given as $REX_{i,t} = \frac{CPI_{it}}{CPI_{RoW,t}} \times E_{it}$ where CPI_{it} is the Consumer Price Index of domestic goods and services in country i at time t and CPI_{RoWt} is the Consumer Price Index in the RoW at time t. The nominal exchange rate E_{it} is the domestic currency price of one unit of foreign currency. Details on the use of different indicators of international price competitiveness are in Durand *et al*. (1998) and Jovanovic (2012).

Figure 1. Dynamics of export market shares country (2001-2012)



Source: calculation on data from Datastream.

Figure 2. Dynamics of total exports and REX by country from 1990 to 2012



Source: our elaborations on data from Datastream.
Legend: —— Export; -·- REX.

3 Econometric evidence

3.1 Testing Stationarity and Co-integration

Equation (1) is estimated using panel-data techniques. The analysis starts by performing the panel unit root test proposed by Levin, Lin and Chu (2002) and the panel co-integration test developed by Westerlund (2007). After performing these two tests, we proceed by using the Mean Group (MG) estimator of Pesaran, Smith and Im (1996) and the Pooled Mean Group (PMG) method proposed by Pesaran et al. (1999).

In order to detect the stochastic properties of time-series, we use the homogeneous Levin, Lin and Chu (2002) panel unit root test (LLC). This test assumes that each individual unit in the panel shares the same AR(1) coefficient, but allows for individual and time effects and considers as additional explanatory variable, a time trend. Lags of the dependent variable are introduced to allow for errors' serial correlation. The test is a pooled Dickey-Fuller test, or an ADF test when lags are included, with the null hypothesis of non-stationarity [*I(1)* behavior]. The *t*-statistic converges to the standard normal distribution. Table 1 shows the results:

Table 1 Levin Lin Chu test for exports and REX time series

Levin-Lin-*Chu* test for **exports**		
Pooled ADF test (1 lag)	N,T = (7,89)	Obs = 609
	Coefficient	-0,0307
	p-value	**0,8166**
Levin-Lin-Chu test for **exchange rate**		
Pooled ADF test (1 lag)	N,T = (7,89)	Obs = 609
	Coefficient	-0,0691
	p-value	**0,1857**

Source: our elaborations on data from Datastream.

With regards to exports, the estimated coefficient of the one-year lagged variable is -0,0307 and the LLC test allows to accept the hypothesis of non-stationarity with a high level of statistical significance (the p-value is about 0.82). This corroborates what we have deducted when looking at figure 2: exports are not stationary. The test allows to draw the same conclusion for the exchange rate. In this case, the estimated coefficient of the one-year lagged variable is -0,0691, with a p-value around 0.19.[4]

After non-stationarity has been ascertained, the next step is to verify the existence of any co-integrating process. This is done by implementing the Westerlund (2007) test. Rejection of H_0 should be taken as rejection of co-integration for the panel as a whole. The underlying idea is to test for the absence of co-integration by determining whether the individual time-series follow an error correction model. The test is very flexible and allows for an almost completely heterogeneous specification of both the long run equilibrium and the short run dynamics. The results for our sample of countries are reported in table 2. The evidence shows that the H_0 of no co-integration is

[4] The world income (Y^w) is also non-stationary. This comes from the augmented Dickey-Fuller test (1981) on the presence of a unit root. The statistic-test is the tau-test (τ), as tabulated by MacKinnon (2010). The estimated coefficient of Y^w is -3,0317 with a p-value around 0.1234.

rejected, implying that a co-integrating relationship between exports and its fundamentals (Y^w and REX) exists and it is highly significant.

Table 2 Westerlund ECM panel co-integration test

Results for H_0 = no co-integration	with 7 series and 2 covariates	
Test for co-integration between exports and (REX & Y^w) - lags(1):		
Statistic value	Z-value	p-value
-7,3530	-3,668	0,000

Source: see table 1

3.2 Panel estimations of exports-price elasticities

Having found that the co-integrating relationship exists, we proceed by using panel methods for non-stationary and co-integrated time series. In this respect and after introducing dynamics and an error correction mechanism, the estimation of equation (1) has been made by performing the Pooled Mean Group estimator (PMG) proposed by Pesaran Shin and Smith (1999) and the Mean Group estimator (MG) of Pesaran, Smith and Im (1996). Both approaches address the non-stationarity of time series for heterogeneous panels.

Generally speaking, an econometric specification of exports demand allows for different degrees of parameter heterogeneity across countries. At one extreme, the full heterogeneity imposes no cross-country parameter restrictions. As our span-period of each time series is large enough, the mean of long and short run coefficients across countries can be estimated consistently by the un-weighted average of any individual country coefficient. This is made by the MG method. At the other extreme, the fully homogeneous coefficient model requires that all slopes and intercepts be equal across countries.[5] This is the simple “pooled” estimator. In ‘between two extremes’ there is the PMG method, which restricts the long run coefficients to be the same across countries, but allows the short run coefficients and the speed of adjustment to be country-specific. The PMG also generates consistent estimates of the mean of short run coefficients across countries by taking the unweighted average of the individual country coefficients (given that the cross-sectional dimension is large). In *I(1)* panels this estimator “allows for mix of co-integration and no co-integration” (Eberhardt, 2011).[6] The MG yields parameters as averages of the N individual group regressions and assumes heterogeneity across countries also for the long run coefficients. The econometric specification of equation (1) to be estimated is as follows:

$$\Delta\log X_{it} = \alpha_i + \beta_{1i}\,\Delta\log REX_{it} + \beta_{2i}\,\Delta\log Y^{w}_{t} - \lambda_i\,(X_{it-1} - \theta_{1i}\,REX_{it-1} - \theta_{2i}\,Y^{w}_{t-1}) + \mathrm{v}_{it} \qquad (2)$$

with $\mathrm{v}_{it} \sim \mathrm{iidN}(0, \sigma_i^2)$ and con (i=1,..,7; t= 1,..,89). PMG estimator differs from MG one because the long run parameters are constant (the subscript *i* is omitted). The spam period is 1990_{Q1}-2012_{Q1}.

[5] They are basically the traditional pooled estimators such as the fixed and random effects estimators, where the intercepts are allowed to differ across groups while all the other coefficients and error variances are constrained to be the same (Pesaran et al. 1996).

[6] Both MG and PMG offers the best available compromise in the search for consistency and efficiency. Indeed, the PMG is particularly useful if countries share the determinants of steady-state, whereas the short-run adjustment are related to country characteristics, such as the financial development and the relative price flexibility. In other words, the PMG predicts not only a common long run equilibrium relationship but also short run dynamics of each single country. In short, it is possible to show that MG always yields consistent estimates, while PMG results are consistent and efficient only if the hypothesis of common long run elasticity is empirically accepted (Pesaran et al. 1996; Pesaran et al. 1999).

In order to control for non-stationarity, the variables in equation (2) are in first differences, as they are non-stationary in level.[7] The coefficients β_i are short run parameters which, like σ_i^2, differ across countries. The error-correcting speed of adjustment term λ_i also differs across *i*. The long run parameters θ_{i1} and θ_{i2} differ country-by-country for MG.

The PMG specification of equation (2) differ from the MG only for what concerns the long run parameters θ_1 and θ_2, which, in PMG, are constant across the groups. In other words, the subscript *i* is omitted in θ_1 and θ_2, consistently with the hypothesis of common long run equilibrium.[8]

As said, short run country heterogeneity is allowed in both estimators, while long run elasticities differ country-by-country in the MG framework and are common across countries in the PMG. However, in using the MG it is also possible to collapse short and long run elasticities to their average values. The same applies in the PMG for what concerns the short run dynamics. Table 3 reports these results, while tables 4 and 5 display the full estimates at individual country level.

From table 3 it emerges that all the elasticities have the expected sign and are highly significant. The main results are two-fold. On the one hand, exports are income elastic in the long run. Indeed, the income elasticity is higher than 1 either when using the PMG or the MG model, even though the magnitude of the effect differs: exports are more income-elastic when considering the MG instead of the PMG approach, since a shock of 1% in world demand would determine an increase of exports of 1.08% under PMG and 1.39% under MG. However, it is meaningful to point out that 1.08 is not statistically different than 1 and thus it is possible to argue that, under PMG, exports have a unitary income elasticity. Differently, the averaged long run income elasticity in MG is statistically different than 1.[9] Our estimates reveal that the income sensitivity of exports is even higher in the short run, being 3.8 the average of the elasticities in PMG as well as in the MG model. A world income shock of 1% induces an increase of 3.8% of exports in the short run.

Turning to price elasticity, table 3 indicates that the demand of exports of all countries, as a whole, is price inelastic, whatever the model. Long run price elasticity is -0.89 in PMG and -0.86 as far as the MG estimator is concerned (the value from MG is the average of the elasticities predicted country-by-country). In both the cases, exports are inelastic, even though the estimated elasticities are not significantly less than unity.[10] The low price sensitivity becomes even more noticeable in the short run: the elasticity ranges from -0.11 in the case of MG model to -0.17 under PMG. Based on these results we can argue that if countries adopted competitive devaluation policies the effect would be an increase in their total national exports, but not so large to be considered aggressive in the world market equilibrium. The evidence demonstrates that a real devaluation of 10% (as averaged across all countries in the sample) would induce an increase of exports of 8.6% in the long run and of, at best 1.7%, in the short run.

However, as already mentioned, the PMG restricts the long run coefficients to be the same across countries, but allows for short run coefficients heterogeneity (including the speed of adjustment). Elasticities differ country-by-country either in the long or in the short run (tables 4 and 5). It is interesting to note that the short run elasticities and the adjustment terms do not differ when comparing PM and PMG results. This seems to be an indirect proof that both models run pretty well

[7] The MG offers the opportunity to get only one short run and long run elasticities simply by averaging the estimations of each individual country. This advantage is due to use panel data instead of time series.

[8] The estimator PMG is quite appealing when studying small sets of arguably 'similar' countries rather than large diverse macro panels (Eberhardt, 2011). The main requirements for the validity of both these methods are such that: *(i)* there exists a long run relationship among the variables of interest and, *(ii)* the dynamic specification of the model be augmented such that the regressors are exogenous and the resulting residual is not serially correlated.

[9] For PMG, we do not reject the null hypothesis of unitary elasticity (the value of the test-statistic is equal to 1.58 with p-value of 0.2082), while for MG estimations, we reject the null hypothesis given that the test-statistic is equal to 8.50 (p-value = 0.0035).

[10] For PMG, the test-statistic is equal to 0.66 (p-value = 0.4175), while for MG, the test-statistic is equal to 0.22 (p-value = 0.6357).

in the short run and they just differ with respect to the hypothesis regarding long run behavior (which is a hypothesis confirmed by our data). Phrased differently, this result implies that if the interest was only to short run dynamics, then it would be indifferent to use PM or PMG estimations.

Nevertheless, long run elasticities vary at country level (table 5) and, thus, it becomes important to verify which is the best performing model between PM and PMG. To this end we run a LR test. The two models are *nested* in each other: the PMG is the unrestricted model, while the PM is without restrictions. The long run elasticities are common across countries under the H_0 hypothesis, while the alternative is that they differ from one country to another (as assumed by the MG estimator).

Table 3. Estimation of the export function of China and 6-OECDs.
PMG and MG averaged estimations over the period 1990-2012

PMG Estimations

	Coef.	Std. Err.	Z	P>|z|	[95% Conf. Interval]	
Long Run						
log(REX)	-0.8906	0.1350	-6.6	0	-1.15511	-0.6260
log(Y^w)	1.0813	0.0646	16.74	0	0.95470	1.2079
Short Run						
Error correction term	-0.0703	0.0189	-3.73	0	-0.1073	-0.0333
Δlog(REX)	-0.1734	0.0589	-2.94	0.003	-0.2889	-0.0580
Δlog(Y^w)	3.8339	0.5836	6.57	0	2.6900	4.9777
Intercept	0.2422	0.0662	3.66	0	0.1124	0.3720

MG Estimations

	Coef.	Std. Err.	Z	P>|z|	[95% Conf. Interval]	
Long Run						
log(REX)	-0.8663	0.2822	-3.07	0.002	-1.4194	-0.3133
log(Y^w)	1.3935	0.1349	10.33	0	1.1290	1.6579
Short Run						
Error correction term	-0.1467	0.0374	-3.93	0	-0.2199	-0.0735
Δlog(REX)	-0.1136	0.0677	-1.68	0.093	-0.2463	0.0191
Δlog(Y^w)	3.8236	0.5565	6.87	0	2.7329	4.9143
Intercept	0.0848	0.1620	0.52	0.601	-0.2327	0.4024

Obs = 616; Number of Groups = 7; Obs per Group = 88
Source: see table 1

Table 4. Estimation of the export function of China and 6-OECDs.
Results from Pooled Mean Group Estimator (1990:Q1-2012:Q1)

	Coef.	Std. Err.	Z	P>|z|	[95% Conf. Interval]	
Long Run						
log(REX)	-0,8906	0,1350	-6,6	0	-1,1551	-0,6260
$\log(Y^w)$	1,0813	0,0646	16,74	0	0,9547	1,2079
China – Short Run						
Error correction term	-0,0345	0,0161	-2,14	0,032	-0,0661	-0,0029
Δlog(REX)	*0,0371*	0,0608	0,61	0,542	-0,0820	0,1561
$\Delta\log(Y^w)$	2,9605	0,5728	5,17	0	1,8378	4,0832
Intercept	0,1176	0,0605	1,94	0,052	-0,0010	0,2363
France - Short Run						
Error correction term	-0,0648	0,0175	-3,71	0	-0,0990	-0,0305
Δlog(REX)	-0,3225	0,1258	-2,56	0,01	-0,5690	-0,0759
$\Delta\log(Y^w)$	3,0207	0,3279	9,21	0	2,3780	3,6634
Intercept	0,2251	0,0760	2,96	0,003	0,0763	0,3740
Germany – Short Run						
Error correction term	-0,0280	0,0153	-1,83	0,067	-0,0579	0,0019
Δlog(REX)	*-0,1888*	0,1795	-1,05	0,293	-0,5406	0,1630
$\Delta\log(Y^w)$	3,2094	0,5918	5,42	0	2,0495	4,3692
Intercept	0,0935	0,0579	1,62	0,106	-0,0200	0,2069
Italy – Short Run						
Error correction term	-0,1297	0,0335	-3,87	0	-0,1954	-0,0641
Δlog(REX)	-0,3261	0,0878	-3,71	0	-0,4982	-0,1539
$\Delta\log(Y^w)$	3,9644	0,3951	10,03	0	3,1900	4,7388
Intercept	0,4606	0,1533	3,01	0,003	0,1602	0,7609
Japan - Short Run						
Error correction term	-0,1516	0,0344	-4,4	0	-0,2191	-0,0841
Δlog(REX)	*0,0482*	0,0732	0,66	0,511	-0,0953	0,1916
$\Delta\log(Y^w)$	7,1225	0,7090	10,05	0	5,7327	8,5122
Intercept	0,5184	0,1372	3,78	0	0,2494	0,7874
UK – Short Run						
Error correction term	-0,0365	0,0146	-2,5	0,012	-0,0652	-0,0079
Δlog(REX)	-0,2337	0,0988	-2,36	0,018	-0,4274	-0,0400
$\Delta\log(Y^w)$	4,0029	0,6411	6,24	0	2,7464	5,2595
Intercept	0,1130	0,0525	2,15	0,031	0,0101	0,2158
USA – Short Run						
Error correction term	-0,0469	0,0131	-3,59	0	-0,0725	-0,0213
Δlog(REX)	-0,2282	0,0750	-3,04	0,002	-0,3753	-0,0811
$\Delta\log(Y^w)$	2,5566	0,3928	6,51	0	1,7867	3,3266
Intercept	0,1672	0,0516	3,24	0,001	0,0660	0,2684

Obs = 616; Number of Groups = 7; Obs per Group = 88 Log Likelihood = 1512.67

Table 5. Estimation of the export function of China and 6-OECDs.
Results from Mean Group Estimator (1990:Q1-2012:Q1)

	Coef.	Std. Err.	Z	P>|z|	[95% Conf. Interval]	
China – LR						
log(REX)	*-0,2207*	0,3009	-0,73	0,463	-0,8104	0,3690
log(Y^w)	1,5546	0,1527	10,18	0	1,2554	1,8538
China – SR						
Error correction term	-0,1175	0,0455	-2,58	0,01	-0,2067	-0,0284
Δlog(REX)	*0,0430*	0,0623	0,69	0,49	-0,0791	0,1650
Δlog(Y^w)	3,1020	0,6107	5,08	0	1,9050	4,2989
Intercept	*-0,1951*	0,1897	-1,03	0,304	-0,5669	0,1768
France –LR						
log(REX)	-2,0405	0,5828	-3,5	0	-3,1828	-0,8982
log(Y^w)	1,0052	0,1682	5,98	0	0,6754	1,3349
France –SR						
Error correction term	-0,0764	0,0248	-3,08	0,002	-0,1251	-0,0277
Δlog(REX)	-0,2626	0,1334	-1,97	0,049	-0,5241	-0,0012
Δlog(Y^w)	3,0248	0,3332	9,08	0	2,3716	3,6779
Intercept	0,6982	0,2514	2,78	0,005	0,2055	1,1910
Germany – LR						
log(REX)	-0,6702	0,1759	-3,81	0	-1,0150	-0,3254
log(Y^w)	2,0309	0,0534	38,03	0	1,9263	2,1356
Germany – SR						
Error correction term	-0,3287	0,0677	-4,86	0	-0,4613	-0,1961
Δlog(REX)	*0,1100*	0,1775	0,62	0,536	-0,2380	0,4579
Δlog(Y^w)	3,0716	0,5455	5,63	0	2,0023	4,1408
Intercept	*-0,5654*	0,3704	-1,53	0,127	-1,2914	0,1605
Italy – LR						
log(REX)	-0,7249	0,2217	-3,27	0,001	-1,1594	-0,2905
log(Y^w)	0,9768	0,0947	10,32	0	0,7913	1,1624
Italy – SR						
Error correction term	-0,1218	0,0344	-3,54	0	-0,1893	-0,0544
Δlog(REX)	-0,3283	0,0899	-3,65	0	-0,5045	-0,1520
Δlog(Y^w)	4,0579	0,4101	9,89	0	3,2541	4,8617
Intercept	0,3950	0,1802	2,19	0,028	0,0417	0,7482
Japan – LR						
log(REX)	-0,5254	0,1469	-3,58	0	-0,8133	-0,2375
log(Y^w)	1,3637	0,0975	13,98	0	1,1726	1,5549

(continue)

(Table 5 continue)

Japan – SR						
Error correction term	*-0,2331*	*0,0501*	*-4,66*	*0*	*-0,3313*	*-0,1350*
Δlog(REX)	*0,0619*	0,0743	0,83	0,405	-0,0837	0,2075
Δlog(Y^w)	6,9404	0,7197	9,64	0	5,5299	8,3510
Intercept	*0,1251*	0,2245	0,56	0,577	-0,3149	0,5652
UK – LR						
log(REX)	*-0,1159*	0,3412	-0,34	0,734	-0,7846	0,5529
log(Y^w)	1,4688	0,1706	8,61	0	1,1345	1,8031
UK – SR						
Error correction term	-0,0990	0,0472	-2,1	0,036	-0,1915	-0,0065
Δlog(REX)	-0,2270	0,1020	-2,23	0,026	-0,4270	-0,0271
Δlog(Y^w)	3,9665	0,6597	6,01	0	2,6735	5,2594
Intercept	*-0,1837*	0,2268	-0,81	0,418	-0,6283	0,2609
USA – LR						
log(REX)	-1,7666	1,1816	-1,5	0,135	-4,0825	0,5494
log(Y^w)	1,3541	0,2893	4,68	0	0,7870	1,9212
USA – SR						
Error correction term	-0,0502	0,0305	-1,65	0,1	-0,1100	0,0096
Δlog(REX)	-0,1921	0,0810	-2,37	0,018	-0,3508	-0,0333
Δlog(Y^w)	2,6022	0,4052	6,42	0	1,8081	3,3964
Intercept	0,3195	0,1563	2,04	0,041	0,0132	0,6258

Obs = 616; Number of Groups = 7; Obs per Group = 88

Source: see table 1

According to LR results, we reject the null hypothesis: the LR yields a chi2(12)=44.0 with a p-value=0. This means that the assumption that countries share the same equilibrium is unrealistic and not supported by data. On the contrary, we find that each country converges to its own long run equilibrium. Based on this, our discussion then regards only the price and income elasticities estimated through the MG method.

Before concentrating on price elasticity, it is fruitful to point out that the aggregate export function is, as expected, foreign income (Y^w) elastic both in the long run and in the short run. From MG results, we already know that the average long run income elasticity is equal to 1.39 (table 3). However this value disregards high country heterogeneity. Indeed, if on the one hand the estimated parameters indicate how important the foreign demand is for each country's exports, on the other hand we find that income is very effective for Germany (the estimated elasticity is 2.03), China (1.55), UK (1.46), Japan (1.36) and USA (1.35). France and Italy exhibit a unitary income elasticity of exports. Income is even more important in the short run, as the elasticity is extremely high. According to our estimates, if a positive shock of 1% in world income occurred, then exports would increase, in the short run, of 6.94% in Japan, 4.06% in Italy, 3.9% in the UK, about 3% in China, France and Germany and of 2.6% in the USA.

As for the scope of this paper, we reveal significant differences in the values of export price elasticity. This holds true in the long and in the short run. In the long run, the analyzed countries have, as expected, a statistically significant negative coefficient with respect to the real exchange rate (*REX*). Estimates vary from -0.52 (Japan) to -2.04 (France). Between these two values we find that the export price elasticity is -0.72 for Italy and-0.67 for Germany. Negative, but not significant is the result regarding China and UK, whose exports are independent on price in the long run. USA's exports exhibit a high (-1.77%) long run price elasticity, although the statistical significance is just 13%. In short, we find that exports from six out of seven countries of the sample are price inelastic, with the exception of France, whose exports would increase by 2% in the presence of a real depreciation of 1%. For the other countries, real devaluation would induce an increase of exports but less than the relative chance of national currency. Exports insensitivity to prices is even more apparent in the short run, as we find a significant relationship between exports and REX only for Italy (-0.33), France (-0.25), UK (-0.23) and USA (-0.19). Aggregate exports from China, Germany and Japan exhibit a wrongly signed, but not significant short run price elasticity.

Results can be synthesized in a few lines. Over the period 1990-2012, the panel-data estimations indicate that China's exports are price insensitive either in the long and in the short run. The same applies for UK exports in the long run. High, but not strongly significant, is the long run price elasticity of USA's exports. In the remaining cases, exports are price inelastic (in the short run, Germany and Japan's exports are even insensitive to changes in price). The only exception is France, whose exports are price elastic in the long run.[11]

3.3 Testing for structural breaks: the Gregory-Hansen test

Previous results disregard possible structural breaks in the co-integration relationship between exports and exchange rates during the 1990:Q1-2012:Q1 period. When structural breaks occur, there could be significant changes of the co-integration parameters or even a change regarding the existence of co-integration relationships itself. Structural breaks may occur for different reasons such as governmental policies, institutional reforms and other country-specific factors. If this is the case, then testing for structural breaks becomes essential because we learn more about the relationships among exports and the real exchange rate, and verify if the estimates so far discussed are reliable.

The existence of structural breaks has been tested by implementing the procedure proposed by Gregory and Hansen (1996), who consider co-integration processes allowing intercepts and/or slope coefficients to break at an unknown point over the period under scrutiny. In formulas we have

$$\log X_t = \mu_1 + \mu_2\, \varphi_{t\tau} + \beta \log REX_t + \mathrm{u}_t \qquad (3)$$

$$\log X_t = \mu_1 + \mu_2\, \varphi_{t\tau} + \delta \mathrm{T} + \beta \log REX_t + \mathrm{u}_t \qquad (4)$$

$$\log X_t = \mu_1 + \mu_2\, \varphi_{t\tau} + \delta \mathrm{T} + \beta \log REX_t + \beta \log REX_t\, \varphi_{t\tau} + \mathrm{u}_t \qquad (5)$$

where $\varphi_{t\tau}$ is the dummy variable

$$\varphi_{t\tau} = \begin{cases} 0 \ \ if \ \ t \leq [n\tau] \\ 1 \ \ if \ \ t > [n\tau] \end{cases}$$

[11] Interesting insights come from the dynamics towards the long run equilibrium. The error correction speed of adjustment is high in the case of Japan and Germany (-0.23 and -0.33 respectively) meaning that they reach their long run equilibrium faster than the other countries of the panel. At the extreme side, the speed of adjustment is very low (-0.05) for USA. Italy and China converge towards their equilibria at the same speedy (-0.11), as do France and the UK, although at slower adjustment (-0-074 for France and -0009 for UK).

The unknown parameter $\tau \in (0,1)$ denotes the timing of the break point (the so-called regime shift) and $[n\tau]$ denotes the integer part, where *n* is the number of periods in the analysis.

In equation (3), the break is modelled as a change in the intercept, while the slopes are constant. If a break occurs at time t, the intercept is μ_1 before t and $\mu_1 + \mu_2$ after t. In that it allows for a level shift in the long run relationship and is known as *level shift model*. equation (4) refers to the *level shift with trend model*, where a time trend is added to the level shift model. Another possible structural break allows the slope vector to also change. This permits the equilibrium relation to rotate and allows a shift as well. This is the third test applied, based on the *regime shift model* (equation 5).

The tests work as follows: when considering the long run relationship between exports and exchange rates, the procedure allows us to identify possible breaks; when this occurs, it tests the null hypothesis of absence of change in the long run relationship. Under the alternative hypothesis, there is movement towards a new long run equilibrium (Gregory and Hansen, 1996). The test is based on an extension of the ADF, Z_t and Z_a test-statistics for co-integration and allows us to detect the stability of co-integration over time in the presence of structural change, if any.[12]

Table 6 shows the results for the test constructed to search for a change in constant (equation 3). What data suggest, is that the tests identify several break points, but none of them determines significant changes in the elasticities before and after the break. These findings are true for all the countries. As can be seen, all the three statistics tests (ADF, Z_t and Z_a) converge to the same result. Another meaningful aspect concerns the breakpoints identified by the Gregory-Hansen (GH) test. For example, in table 6, a break is detected for China at the 52nd period that corresponds to the first quarter of 2003 when the test is run with ADF. The break is identified at the 46th period (third quarter of 2001) if the test is implemented through Z_a and Z_t. This shock, however, was not strong enough to impact the long run elasticity. The accession of China to the World Trade Organization in 2001 and that of the Chinese Taipei in 2002 maybe are the reasons of these breakpoints (Kerr and Hobbs, 2001). In testing to search for a change in constant, it is worth noticing that for Italy a break is identified in 1993 (13th period with Z-statistics and 15th with ADF) that means that the GH test can catch some shocks arising from competitive devaluation of the national currency that the country adopted in the previous year to stimulate its exports (Macis and Schivardi, 2012). However, also in this case the long run relationships remain the same before and after the break.

The results achieved when considering equations (4) and (5) allowing for changes in constant and in both constant and slope in models with time trend (tables 7 and 8) are qualitatively similar to those obtained with the first test. In particular, the findings show that the long run elasticity does not change before and after the structural breaks. The calculated statistics are in all cases lower than the asymptotic critical values at 1%, 5% and 10%.

From the side of the breakpoint time, it should be noted that the three tests provide different results for the same countries, except for Germany that seems to be affected by some events in

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[12] The starting point to calculate Z_a and Z_t statistics is to estimate the first-order serial correlation coefficient, $\hat{\rho}$, for the OLS residuals of equations (3), (4) and (5). Then, a bias-correction version, $\hat{\rho}^*$, is used, where the correction involves the estimate of a weighted sum of auto-covariances. The difference between Z_a and Z_t consists in the fact that Z_t consider also a transformation of the long run variance $\hat{s}^2$ of the OLS residuals (in formulas: $Z_a(\tau) = n(\hat{\rho}_\tau^* - 1)$ and $Z_t(\tau) = (\hat{\rho}_\tau^* - 1)/\hat{s}$). The $ADF(\tau)$ statistic is calculated by regressing the first difference of OLS residuals upon the lagged residuals and their lagged first differences for some suitably chosen lag truncation. The three statistics ADF, Z_a and Z_t are calculated across all estimated values of the regime shifts $\tau \in \mathrm{T}$. Then, the GH test is performed by taking the smallest values of each statistics, as these values constitute evidence against the null hypothesis. The test-statistics become $Z_a = \inf_{\tau \in \mathrm{T}} Z_a(\tau)$, $Z_t = \inf_{\tau \in \mathrm{T}} Z_t(\tau)$ and $ADF = \inf_{\tau \in \mathrm{T}} ADF(\tau)$. For details, see Gregory and Hansen (1996).

2004. In the case of this country, all the tests identify a break in 2004, even if in different quarters. This might be related to the fact that Germany implemented a comprehensive set of labour market reforms during the period 2003-2005, the so-called Hartz act (Bodegan et al., 2010; Jacobi and Kluve, 2006). Furthermore, table 8 indicates that a break is detected for China, France, Italy and UK in 2008, that is when some shocks due to the financial crisis started with the US sub-prime loans and propagated in many other countries (Grigor'ev and Salikhov, 2009).

Obviously, we expected the same for the USA, but the GH test fails to capture any remarkable circumstance in the country in 2008. Conversely, the USA exhibited a structural change during the last quarter of 2001 (table 8, third test), surely due to the World Trade Center terrorist attack and to the *dot.com* crisis (Abadie and Gardeazabal, 2003). Taking into account how the GH test works, we can argue that the revealed 2001 break is more important than that related to the expected effect of 2008 crisis, however is not so important to significantly affect the long run path of US exports.

Table 6 Gregory-Hansen Test for Co-integration: searching for Change in Constant

Number of obs = 89 Lags= 2 chosen by Akaike criterion
Maximum Lags = 5

Test Statistic	Breakpoint date		Asymptotic Critical Values 1%	5%	10%
Country: China					
ADF -3,20	52	2003: Q1	-5.13	-4.61	-4.34
Zt -3,13	46	2001: Q3	-5.13	-4.61	-4.34
Za -13,26	46	2001: Q3	-50,07	-40,48	-36,19
Country: France					
ADF -3,19	55	2003: Q4	-5.13	-4.61	-4.34
Zt -3,40	54	2003: Q3	-5.13	-4.61	-4.34
Za -17,25	54	2003: Q3	-50,07	-40,48	-36,19
Country: Germany					
ADF -3,28	58	2004: Q3	-5.13	-4.61	-4.34
Zt -3,66	57	2004: Q2	-5.13	-4.61	-4.34
Za -17,63	57	2004: Q2	-50,07	-40,48	-36,19
Country: Italy					
ADF -3,23	15	1993:Q4	-5.13	-4.61	-4.34
Zt -3,41	13	1993:Q2	-5.13	-4.61	-4.34
Za -18,09	13	1993:Q2	-50,07	-40,48	-36,19
Country: Japan					
ADF -3,54	59	2004: Q4	-5.13	-4.61	-4.34
Zt -3,70	53	2003: Q2	-5.13	-4.61	-4.34
Za -18,75	53	2003: Q2	-50,07	-40,48	-36,19
Country: UK					
ADF -3,46	70	2007: Q3	-5.13	-4.61	-4.34
Zt -3,05	66	2006: Q3	-5.13	-4.61	-4.34
Za -13,74	66	2006: Q3	-50,07	-40,48	-36,19
Country: USA					
ADF -3,58	29	1997: Q2	-5.13	-4.61	-4.34
Zt -3,86	29	1997: Q2	-5.13	-4.61	-4.34
Za -26,25	29	1997: Q2	-50,07	-40,48	-36,19

Source: see table 1

Table 7 Gregory-Hansen Test for Co-integration: searching for Change in Constant in model with time trend

Number of obs = 89 Lags= 2 chosen by Akaike criterion

Maximum Lags = 5

Test Statistic	Breakpoint date		Asymptotic Critical Values 1%	5%	10%
Country: **China**					
ADF -3,20	52	2003: Q1	-5.47	-4.95	-4.68
Zt -3,13	46	2001: Q3	-5.47	-4.95	-4.68
Za -13,26	46	2001: Q3	-57,17	-47,04	-41,85
Country: **France**					
ADF -3,19	59	2004: Q4	-5.47	-4.95	-4.68
Zt -3,42	36	1999: Q1	-5.47	-4.95	-4.68
Za -16,45	36	1999: Q1	-57,17	-47,04	-41,85
Country: **Germany**					
ADF -3,28	58	2004:Q3	-5.47	-4.95	-4.68
Zt -3,66	53	2003:Q2	-5.47	-4.95	-4.68
Za -17,63	53	2003:Q2	-57,17	-47,04	-41,85
Country: **Italy**					
ADF -4,22	32	1998:Q1	-5.47	-4.95	-4.68
Zt -4,26	35	1998:Q4	-5.47	-4.95	-4.68
Za -19,49	35	1998:Q4	-57,17	-47,04	-41,85
Country: **Japan**					
ADF -3,54	31	1997: Q4	-5.47	-4.95	-4.68
Zt -3,70	33	1998:Q2	-5.47	-4.95	-4.68
Za -18,75	33	1998:Q2	-57,17	-47,04	-41,85
Country: **UK**					
ADF -3,46	71	2007: Q4	-5.47	-4.95	-4.68
Zt -3,05	74	2008:Q3	-5.47	-4.95	-4.68
Za -13,74	74	2008:Q3	-57,17	-47,04	-41,85
Country: **USA**					
ADF -3,58	57	2004:Q2	-5.47	-4.95	-4.68
Zt -3,86	58	2004:Q3	-5.47	-4.95	-4.68
Za -26,25	58	2004:Q3	-57,17	-47,04	-41,85

Source: see table 1

Table 8 Gregory-Hansen Test for Co-integration: searching for Change in Constant and Slopes in model with time trend

Number of obs = 89 Lags= 2 chosen by Akaike criterion

Maximum Lags = 5

Test Statistic	Breakpoint	date	Asymptotic Critical Values		
			1%	5%	10%
Country: **China**					
ADF -3,20	74	2008: Q3	-5.45	-4.99	-4.72
Zt -3,13	74	2008: Q3	-5.45	-4.99	-4.72
Za -13,26	74	2008: Q3	-57.28	-47.96	-43.22
Country: **France**					
ADF -3,19	73	2008: Q2	-5.45	-4.99	-4.72
Zt -3,42	74	2008: Q3	-5.45	-4.99	-4.72
Za -16,45	74	2008: Q3	-57.28	-47.96	-43.22
Country: **Germany**					
ADF -3,28	58	2004: Q3	-5.45	-4.99	-4.72
Zt -3,66	59	2004: Q4	-5.45	-4.99	-4.72
Za -17,63	59	2004: Q4	-57.28	-47.96	-43.22
Country: **Italy**					
ADF -4,27	73	2008: Q2	-5.45	-4.99	-4.72
Zt -4,48	73	2008: Q2	-5.45	-4.99	-4.72
Za -23,43	73	2008: Q2	-57.28	-47.96	-43.22
Country: **Japan**					
ADF -3,54	59	2004:Q4	-5.45	-4.99	-4.72
Zt -3,70	57	2004:Q2	-5.45	-4.99	-4.72
Za -18,75	57	2004:Q2	-57.28	-47.96	-43.22
Country: **UK**					
ADF -3,46	74	2008: Q3	-5.45	-4.99	-4.72
Zt -3,05	75	2008: Q4	-5.45	-4.99	-4.72
Za -13,74	75	2008: Q4	-57.28	-47.96	-43.22
Country: **USA**					
ADF -3,58	48	2002:Q1	-5.45	-4.99	-4.72
Zt -3,86	47	2001: Q4	-5.45	-4.99	-4.72
Za -26,25	47	2001: Q4	-57.28	-47.96	-43.22

Source: see table 1

4 Concluding remarks

This paper investigates the relationship between the real exchange rate and export demand of seven exporting countries (China, France, Germany, Italy, Japan, UK, and the USA) over the period 1990-2012. The analysis is based on the economic model proposed by Goldstein and Khan (1985), whilst the econometric specification is adopted to non-stationary panel data and conducted using the Pooled Mean Group and the Mean Group estimators developed by Pesaran, Shin, and Smith (1999) and Pesaran, Smith, and Im (1996), respectively. These methods allow for country heterogeneity in the long run equilibrium as well as for short run dynamics. The evidence shows that the MG model better fits the data, as supported by the LR post-estimation test. Since MG allows for full country-heterogeneity of the relationships between exports and their fundamentals (income and price competitiveness), the LR result helps to draw three general conclusions. Firstly, the hypothesis of common long run equilibrium across countries is not supported by the data (the result evidently reflects the absence of homogeneity within the sample). Secondly, each country accordingly converges towards its own long run equilibrium with a country-specific speed of adjustment. Thirdly, the differences in the short run income and price elasticities underscore that the starting point of the transition path towards the final equilibrium varies by country.

From an economic perspective, we find that the aggregate exports are highly income elastic in both the long and short runs, implying that global economic growth induces increases in world aggregate demand and thus impacts positively and significantly on the total exports of the countries considered in the study. This result is consistent with the expectations and the evidence provided by others. Furthermore, exports are, on average, price inelastic. As far as the seven countries are concerned, long run export price elasticity is -0.89, meaning that exports would increase by 8.9 percent if a 10 percent of real exchange rate depreciation occurs. In other words, total exports do increase in cases of aggressively competitive devaluation policies, but far less than the expansions one expects after having observed how intense and crude the tensions on currency markets are. The low export price sensitivity holds true when focusing on individual countries. Surprisingly, the *nexus* exports-price competitiveness is difficult to interpret in the case of China, whose long run price elasticity is low (-0.22) and not significant and in the short run is also signed wrongly (although again not significant). Similarly, the long run level of exports appears to be unrelated to the real exchange rate for the UK (whose elasticity is -0.11 but not significant). When results are significant, the long run price elasticity is -0.52 for Japan, -0.67 for Germany, and -0.72 for Italy. The exception is France, whose exports exhibit a long run exchange rate elasticity of -2. The similar high price elasticity (-1.77) of the USA's exports is not immediately interpretable because it is significant, but only weakly so. Noticeably, all the findings are robust over time, as no significant change exists in the long run co-integrated path of exports and real exchange rates, even after having identified some structural country-level breaks at specific points of time.

This mixed evidence supports the pessimistic view that exchange rate policies may not be fully successful in promoting export growth: if a competitive devaluation is carried out by aggressive countries, total exports will in fact increase, but only moderately. This is puzzling in light of the debate on currency devaluation, which assumes that exports are highly price elastic. On the contrary, our findings suggest that the gains in exports are less than expected, because aggregate exports are price inelastic, as previously documented.

There are many promising avenues for further research in this area. For instance, it would be valuable to investigate whether disaggregation could yield greater insights into the counterintuitive evidence on export price sensitivity. Along this line of reasoning, it would be fruitful to deepen the analysis by considering sectoral bilateral trade flows and different measures of price competitiveness.

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