New Evidence on Export Price Elasticity from China and Six OECD Countries

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Abstract

This paper provides new evidence on export price elasticities by analyzing the cases of China, France, Germany, Italy, Japan, UK and the USA over the period 1990–2012. Estimates have been made using panel data techniques for non-stationary data. After demonstrating that long-run relationships are stable to any structural break, it is found that exports are significantly determined by foreign demand, with long-run income elasticity significantly higher than unity for China, Japan, Germany, the UK and the USA. Conversely, exports are price inelastic for most of the countries in the sample, in both the long run and the short run. The exception is France, whose export price elasticity is lower (higher) than unity in the short run (long run).

Key words: competitive devaluation, currency wars, export price elasticity, panel data
JEL codes: C23, F10, F17, F37, P33

I. Introduction

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 percent of its GDP; it then experienced rapidly growing trade deficits and since the 1990s it has been the world's largest debtor. In 2000 Germany had a trade deficit of 1.83 percent of its GDP, but became a net-exporter by 2013, with a trade surplus of 7.58 percent of its GDP. China ran a trade surplus averaging 4.24 percent of its GDP from 1998 to 2013, peaking at approximately 10 percent 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

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their currency to gain from competitive devaluations. Disputes over 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 that of China, Germany, Japan and the UK have also manipulated their real exchange rates. Japan and the UK used quantitative easing to counter the current recession (Joyce *et al.*, 2011; Gagnon, 2013), and, according to the US Treasury, Germany's low level of investment and high savings rate contributed to the eurozone crisis, which is characterized by increasing trade troubles for the EU periphery and huge surpluses for Germany. Following this, one would expect that controlling exchange rates is a feasible policy to improve trade balances. In other words, tensions in currency markets are expected if devaluations lead to substantial increases in exports. In short, exports are expected to be price elastic. This expectation, however, is not empirically supported, with price elasticity found in many studies to be less than unity.

While this heterogeneity of results casts doubt on the effect of a real devaluation, it also indicates that price competitiveness remains a controversial and intriguing issue in international trade (see e.g. Orcutt, 1950; Houthakker and Magee, 1969; Kravis and Lipsey, 1978). There are surveys of initial papers in Stern *et al.* (1976), Goldstein and Khan (1985) and Sawyer and Sprinkle (1996). Although these review papers demonstrate the wide range of price elasticities, it is noteworthy that the picture does not change in more recent studies. Limiting attention to price elasticities of aggregate trade flows, several authors show that exports are price inelastic. 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 (the yuan, the dollar and the 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, the Netherlands, Spain, the UK and the USA are rather small (in the range of 0.3/0.8) over the period 1978 – 2009. Similarly, the export price elasticity of eurozone countries is found to be low in Bayoumi *et al.* (2011) and in Chen *et al.* (2012), at 0.6 and 0.46, respectively. Ketenci and Uz (2011) study the EU bilateral trade flows over 1980–2007 and find an export price elasticity ranging in the 0.08/0.64 interval. The price elasticity of Germany 's exports is reported as 0.6 in Thorbecke and Kato (2012). Thorbecke and Kato (2012) focus on Japanese exports to 17 partners over the period 1988–2009 and find that exports are price inelastic, although a unitary long-run elasticity is found for consumption products. Crane *et al.* (2007) find that during the 1981–2006 period the price elasticity is low for Italy (0.7), Japan (0.34) and the USA (0.6). Yao *et al.* (2013) look at total Chinese exports from 1992 to 2006 and, even after controlling for an increase in product variety, they find a short-run price elasticity of 0.65. Dezeure and Teixeira (2014) argue that in spite of depreciation of the pound, the weak growth of British exports in the 2000s is due to the virtually zero elasticity between exports and the exchange rate.

The conclusion drawn from this discussion is that exports are price inelastic, whichever country and time period are examined and whatever methods are used. Thus, macro-analyses do not make currency tensions easy to understand, because they originate from the controversial assumption of high export price sensitivity. Indeed, if the macro-level estimates are reliable, then competitive devaluations will not lead to increased trade surpluses in the "aggressive" countries and, therefore, will not penalize trading partners.

This paper contributes to the debate in three ways. First, it proposes an updated analysis of the export behavior of six OECD countries (France, Germany, Italy, Japan, the UK and the 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 WTO in 2001. Total exports are examined from 1990 to 2012, a period during which there were a number of changes in world trade structures. Second, the research is enriched by testing for structural breaks. Structural breaks can affect model parameters, thereby inducing different policy implications. Third, estimates of export price elasticities are based on panel data techniques for non-stationary data. This represents an important novelty because these methods are rarely used to estimate trade elasticities, although they were developed in the 1990s. Within this analytical framework, we use the export equation derived from the imperfect substitutes model proposed by Goldstein and Khan (1985). After checking for non-stationarity, stability and cointegration of time-series, the analysis is carried out by applying the pooled mean group (PMG) estimator developed by Pesaran *et al.* (1999) and the mean group (MG) estimator of Pesaran *et al.* (1996), thus allowing for full country heterogeneity of short-run price elasticity. Long-run elasticity is assumed to be common across countries for the PMG estimator and country-specific for the MG estimator. Finally, we also use a vector error correction model (VECM) to check the robustness of panel data results.

The paper proceeds as follows: Section II presents the empirical setting and describes the data sample; Section III presents the results; finally, Section IV concludes.

II. Empirical Setting and Data

The empirical setting relates to the imperfect substitutes model proposed by Goldstein and Khan (1985), whose major assumption is that neither imports nor exports are perfect substitutes for domestic goods. This implies that if domestic and foreign goods were perfect substitutes, then one should observe either of the goods having a market share of unity, and each country acts as an importer or exporter of a traded good but not both (Goldstein and Khan, 1985). Again, the coexistence of trade flows and domestic production makes the hypothesis of perfect substitutes unrealistic. Following the literature, the econometric log-linear specification of the export demand is: [1]

$$\ln X_{it} = \alpha_i + \beta_1 \ln REX_{it} + \beta_2 \ln Y_{it}^{w} + u_{it}\,, \tag{1}$$

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 export elasticity to the real exchange rate and β_2 is the export elasticity to foreign income. Generally speaking, it is expected that β_1 is negative, as currency depreciations determine an increase in exports. The parameter β_2 is expected to be positive, indicating that exports rise with world income. 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. [2] Data are from DataStream and are expressed on a quarterly basis covering the period 1990:Q1 – 2012:Q1. They are in real terms (2005 is the base year) and are seasonally adjusted.

The sample comprises China and the 6-OECDs (France, Germany, Italy, Japan, the UK and the USA). While the OECDs have always been important traders, China is the subject of interest in the current debate on exchange rate misalignments because of its growing role as an exporter. The sample of countries provide much of the world exports; their total export shares comprised approximately 45 percent of world exports for the 2-year period from 2001 to 2002 and approximately 40 percent for 2010 –2012. The data also highlight the impressive pattern of Chinese export shares, which increased by approximately 7 percentage points, moving from 4.3 percent in 2001 –2002 to 11.3 percent in 2012. Interestingly, market shares have decreased for the other exporters (e.g. the USA 's market share was 8.6 percent in 2012, but 11.9 percent in 2001), except for Germany, whose market share was 8 percent in 2012. What the data clearly highlight is that China has become an important exporter in the past few years.

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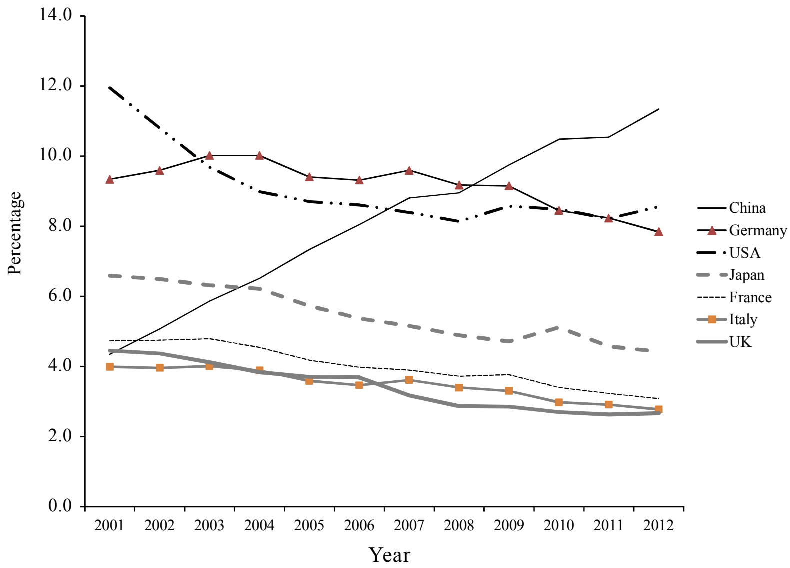
[1]As the economic model from which the foreign demand originates is well-known, the system of eight equations proposed by Goldstein and Khan (1985) is not displayed in the present study. In this, we follow the choice made by Hamori and Yin (2011), Ketenci and Uz (2011), Hamori and Matsubayashi (2009), Caporale and Chui (1999), Senhadji and Montenegro (1999), Bahmani-Oskooee and Niroomand (1998), Sawyer and Sprinkle (1996) and Thorbecke (2011).

[2]The real exchange rate is $REX_{i,t} = \frac{CPI_{it}}{CPI_{RoW,t}} \times E_{it}$, where CPI_{it} is the consumer price index of domestic goods and services in country i at time t and $CPI_{RoW,t}$ is the corresponding index for the rest of the world. The nominal exchange rate E_{it} is the domestic currency price of one unit of foreign currency.

Figure 1. Dynamics of World Export Market Shares (by Country, 2001 –2012)



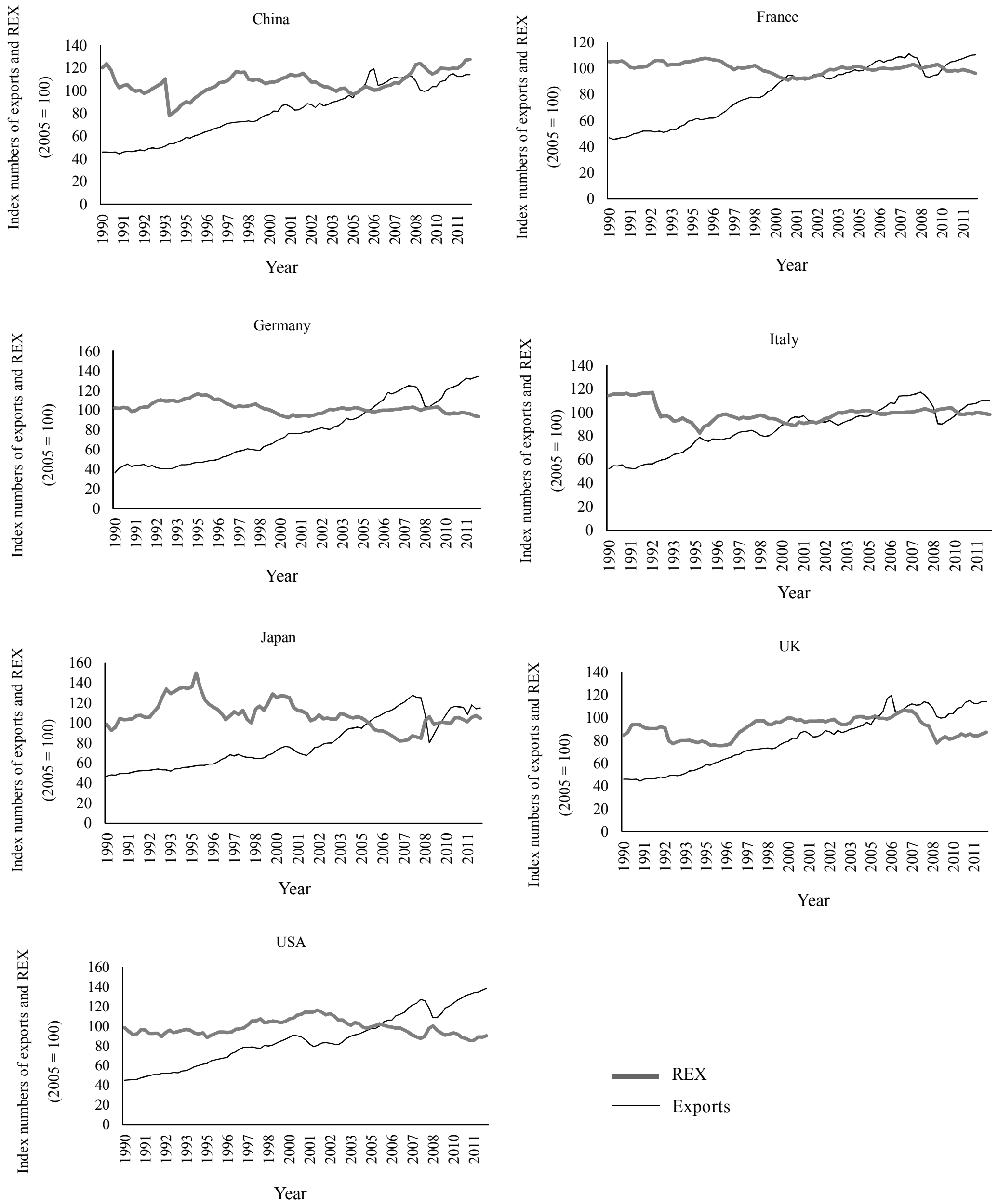
Source: Data elaborated from DataStream.

Figure 2 plots the time series of exports and real exchange rates over the 1990 –2012 period. Although a strong positive increase is revealed for the exports in each country, China has the highest increase, followed by the UK and the USA. Another common result is the drop in exports at the time of the 2008 financial crisis. Exports reduced much more in Italy and Japan than in the other countries. All countries observed a recovery of exports after 2008. Moreover, Figure 2 clearly highlights that exports in the time series exhibit a non-stationary pattern. The same does not appear to be true for the real exchange rate, a fact that deserves more attention (see Section III). In the case of the real exchange rate (REX), there is much more instability along the trend than a strict trend pace itself. Hence, it becomes interesting to evaluate the effects of this variability on export behavior. It is an issue that will be addressed in the following paragraphs when measuring the short-run export-price elasticity.

III. Econometric Evidence

Equation (1) is estimated using panel data techniques. The analysis starts by performing the panel unit root test proposed by Levin *et al.* (2002) and the panel cointegration test of Westerlund (2007). In addition, we use the Gregory and Hansen (1996) test for checking the structural stability of each time series. After performing these tests, we proceed by using

Figure 2. Dynamics of Total Exports and REX (by Country, 1990 –2012)



Source: Elaboration of data from DataStream.
Note: REX, real exchange rate.

the MG estimator of Pesaran *et al.* (1996) and the PMG method proposed by Pesaran *et al.* (1999).

1. Testing Stationarity and Cointegration

In order to detect the stochastic properties of the time series as a whole, we use the Levin *et al.* (2002) panel unit root test. This test performs well in case of homogeneous panel and assumes that each individual unit shares the same AR(1) coefficient, but allows for individual and time effects. Lags of the dependent variable are introduced to allow for serial correlation. The test is a pooled Dickey–Fuller test, or an augmented Dickey–Fuller (ADF) test when lags are included, with the null hypothesis of non-stationarity. The *t*-statistic converges to the standard normal distribution. Table 1 shows the results.

Regarding exports, the estimated coefficient of the one-period lagged variable is –0.03 and the Levin Lin Chu test supports the hypothesis of non-stationarity with a high level of significance (the *p*-value is approximately 0.82). This corroborates what we have deduced from Figure 2: exports are not stationary. The same applies for the exchange rate, as the coefficient of the one-period lagged variable is –0.07 (the *p*-value is 0.19).[3]

After non-stationarity has been ascertained, the next step is to verify the existence of any cointegrating process. This is done by implementing the Westerlund (2007) test. Rejection of H_0 should be taken as rejection of cointegration for the entire panel. The

Table 1. Levin Lin Chu Test for Exports and Real Exchange Rate Time-series

Levin–Lin–Chu test for exports		
Pooled ADF test (1 lag)	N,T = (7,89)	Observation = 609
	Coefficient	–0.031
	p-value	0.8167
Levin–Lin–Chu test for real exchange rate		
Pooled ADF test (1 lag)	N,T = (7,89)	Observation = 609
	Coefficient	–0.069
	p-value	0.186

Source: Elaboration of data from DataStream.
Note: ADF, Augmented Dickey Fuller.

[3] World income (Y^w) is also non-stationary. This comes from the augmented Dickey–Fuller test (1981). The statistic test used is the tau test (τ), as tabulated by MacKinnon (2010). The estimated coefficient of Y^w is –3.03 with a *p*-value = 0.12. Furthermore, the evidence of Table 1 overlaps that obtained when performing the ADF *t*-test for a heterogeneous panel as proposed by Im *et al.* (2003; results are available upon request). It is also important to emphasize that our panel is composed of a sectional dimension of seven exporters. This issue belongs to the ongoing discussion comparing large and small panel data (Eberhardt, 2011). It can be addressed by performing robustness analyses as in, for example, Roud *et al.* (2007). In our case, the large T dimension ensures the reliability of panel data results; we also find that panel data estimations overlap a lot of those obtained from individual country studies (see footnote 11 and Table 9).

Table 2. Westerlund ECM Panel Cointegration Test

Results for H_0 = No cointegration with seven series and two covariates		
Test for cointegration between exports and (*REX* & Y^w) – lags(1):		
Statistic value	*Z*-value	*p*-value
−7.353	−3.668	0

Source: Elaboration of data from DataStream.
Notes: REX, real exchange rate; ECM, error correction model.

underlying idea is to test for the absence of cointegration 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 evidence shows that the H_0 of no cointegration is rejected, implying that there is a significant cointegrating relationship between exports and Y^w and REX (Table 2).

2. Testing for Structural Stability

Previous studies disregard possible structural breaks in the cointegration relationship between exports and exchange rates. A break may be the result of governmental policies, institutional reforms and other country-specific factors. If the break is significant, it alters the cointegration parameters. Thereby, after accepting the hypothesis of stability we learn more about the links between exports and exchange rates, in the sense that the long-run relationship that will be estimated through panel data will be seen as reliable.

The existence of structural breaks was detected using the Gregory and Hansen (1996) test, who consider cointegration processes allowing intercepts and/or slope coefficients to break at an unknown time point. We have the following formulas:

$$\ln X_t = \mu_1 + \mu_2 \varphi_{t\tau} + \beta \ln REX_t + u_t \tag{2}$$

$$\ln X_t = \mu_1 + \mu_2 \varphi_{t\tau} + \delta T + \beta \ln REX_t + u_t \tag{3}$$

$$\ln X_t = \mu_1 + \mu_2 \varphi_{t\tau} + \delta T + \beta \ln REX_t + \beta \ln REX_t \varphi_{t\tau} + u_t, \tag{4}$$

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

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

The parameter $\tau \in (0,1)$ denotes the timing of the break point (the regime shift) and $[n\tau]$ is the integer part, where *n* is the number of periods in the analysis. In Equation (2), the break is modeled as a change in the intercept. If a break occurs at time *t*, the intercept is μ_1 before *t* and $\mu_1 + \mu_2$ after *t*. As it allows for a level shift in the long-run relationship, it is known as a "level shift model." For Equation (3) a time trend is added to Equation (2), yielding a "level shift with trend model." Finally, the "regime shift model" allows for breaks of slopes (Equation 4).

The Gregory and Hansen test is used to identify potential breaks in the long-run relationship between exports and exchange rates. The null hypothesis is the absence of change in the long-run relationship. Under the alternative hypothesis there is a pace towards a new long-run equilibrium. The test is an extension of the ADF, Z_t and Z_a test statistics for cointegration and, therefore, allows us to detect the stability of cointegration in the presence of structural change.[4]

Table 3 shows the results of Equation (3). Several breakpoints are identified, but no significant changes are identified in the elasticities before and after the break. These findings hold for every country, as the three statistics tests (ADF, Z_t and Z_a) converge to the same decision. The Gregory and Hansen test is also highly informative about the time of the break. A break is detected for China at the 52nd period, which corresponds to the first quarter of 2003 when the test is carried out by using the ADF approach. The break is identified at the 46th period (third quarter of 2001) if the test is implemented through Z_a and Z_t. The 2001 accession to the WTO and Chinese Taipei 's accession in 2002 may be the reasons for these breakpoints (Kerr and Hobbs, 2001). This shock, however, was not strong enough to affect the long-run elasticity. Interestingly, in testing for changes in the constant, a break is identified for Italy in 1993 (13th period with Z-statistics and 15th with ADF): the Gregory and Hansen test captures some shocks arising from the 1992 devaluation of the national currency adopted to stimulate exports (Macis and Schivardi, 2012). Even in this case the long-run path of Italian exports is robust to the break.

Results from the other two models (Equations 3 and 4) are qualitatively similar to those obtained with the first test (Tables 4 and 5). In particular, the findings show that the long-run elasticity does not change before and after the structural breaks. The calculated ADF, Z_a and Z_t statistics are always lower than the asymptotic critical values at any convenient level of significance (1, 5 and 10 percent).

Regarding the breakpoint time, the evidence indicates that Germany faced a break in

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[4]The starting point to calculate Z_a and Z_t statistics is to estimate the first-order serial correlation coefficient of OLS residuals, $\hat{\rho}^*$. The difference between Z_a and Z_t lies in the fact that Z_t considers also a transformation of the long-run variance $\hat{s}^2$ of 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 OLS residuals (in first differences) against their lags and the lagged first differences. The ADF statistics, 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 they 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 $ADF = \inf_{\tau \in \mathrm{T}} ADF(\tau)$.

Table 3. Gregory–Hansen Test for Cointegration

Test statistic		Breakpoint date		Asymptotic critical values		
				1%	5%	10%
China						
ADF	−3.20	52	2003: Q1	−5.13	−4.61	−4.34
Z_t	−3.13	46	2001: Q3	−5.13	−4.61	−4.34
Z_a	−13.26	46	2001: Q3	−50.07	−40.48	−36.19
France						
ADF	−3.19	55	2003: Q4	−5.13	−4.61	−4.34
Z_t	−3.40	54	2003: Q3	−5.13	−4.61	−4.34
Z_a	−17.25	54	2003: Q3	−50.07	−40.48	−36.19
Germany						
ADF	−3.28	58	2004: Q3	−5.13	−4.61	−4.34
Z_t	−3.66	57	2004: Q2	−5.13	−4.61	−4.34
Z_a	−17.63	57	2004: Q2	−50.07	−40.48	−36.19
Italy						
ADF	−3.23	15	1993: Q4	−5.13	−4.61	−4.34
Z_t	−3.41	13	1993: Q2	−5.13	−4.61	−4.34
Z_a	−18.09	13	1993: Q2	−50.07	−40.48	−36.19
Japan						
ADF	−3.54	59	2004: Q4	−5.13	−4.61	−4.34
Z_t	−3.70	53	2003: Q2	−5.13	−4.61	−4.34
Z_a	−18.75	53	2003: Q2	−50.07	−40.48	−36.19
UK						
ADF	−3.46	70	2007: Q3	−5.13	−4.61	−4.34
Z_t	−3.05	66	2006: Q3	−5.13	−4.61	−4.34
Z_a	−13.74	66	2006: Q3	−50.07	−40.48	−36.19
USA						
ADF	−3.58	29	1997: Q2	−5.13	−4.61	−4.34
Z_t	−3.86	29	1997: Q2	−5.13	−4.61	−4.34
Z_a	−26.25	29	1997: Q2	−50.07	−40.48	−36.19

Source: Elaboration of data from DataStream.

Notes: ADF, augmented Dickey–Fuller. Number of observations = 89; lags = 2 chosen by Akaike criterion; maximum lags = 5.

2004. This might be related to the labor market reforms introduced in Germany in 2003 –2005 (Jacobi and Kluve, 2006; Bodegan *et al.*, 2010). Furthermore, a break is detected for China, France, Italy and the UK in 2008 (Table 5); that is, when shocks were triggered by the US sub-prime loans and propagated worldwide (Grigor ’ev and Salikhov, 2009). The Gregory and Hansen test fails to capture any remarkable circumstances in the USA in 2008. Conversely, the USA exhibited a structural change during the last quarter of 2001 (Table 5, third test), surely due to the World Trade Center terrorist attack and to the dot.com crisis (Abadie and Gardeazabal, 2003). Thus, we can argue that the long-run path of US exports is not significantly affected by the break detected in 2001, which, according to the results, is more important than the expected effect of the 2008 crisis.

3. Panel Estimations of Export Price Elasticities

Having found that there is a cointegrating relationship and that it is stable over time, we proceed by using panel methods for non-stationary and cointegrated time series. In this

Table 4. Gregory–Hansen Test for Cointegration

Test statistic		Breakpoint date		Asymptotic critical values 1%	5%	10%
China						
ADF	−3.20	52	2003: Q1	−5.47	−4.95	−4.68
Z_t	−3.13	46	2001: Q3	−5.47	−4.95	−4.68
Z_a	−13.26	46	2001: Q3	−57.17	−47.04	−41.85
France						
ADF	−3.19	59	2004: Q4	−5.47	−4.95	−4.68
Z_t	−3.42	36	1999: Q1	−5.47	−4.95	−4.68
Z_a	−16.45	36	1999: Q1	−57.17	−47.04	−41.85
Germany						
ADF	−3.28	58	2004: Q3	−5.47	−4.95	−4.68
Z_t	−3.66	53	2003: Q2	−5.47	−4.95	−4.68
Z_a	−17.63	53	2003: Q2	−57.17	−47.04	−41.85
Italy						
ADF	−4.22	32	1998: Q1	−5.47	−4.95	−4.68
Z_t	−4.26	35	1998: Q4	−5.47	−4.95	−4.68
Z_a	−19.49	35	1998: Q4	−57.17	−47.04	−41.85
Japan						
ADF	−3.54	31	1997: Q4	−5.47	−4.95	−4.68
Z_t	−3.70	33	1998: Q2	−5.47	−4.95	−4.68
Z_a	−18.75	33	1998: Q2	−57.17	−47.04	−41.85
UK						
ADF	−3.46	71	2007: Q4	−5.47	−4.95	−4.68
Z_t	−3.05	74	2008: Q3	−5.47	−4.95	−4.68
Z_a	−13.74	74	2008: Q3	−57.17	−47.04	−41.85
USA						
ADF	−3.58	57	2004: Q2	−5.47	−4.95	−4.68
Z_t	−3.86	58	2004: Q3	−5.47	−4.95	−4.68
Z_a	−26.25	58	2004: Q3	−57.17	−47.04	−41.85

Source: Elaboration of data from DataStream.

Notes: ADF, augmented Dickey–Fuller. Number of observations = 89, lags = 2 chosen by Akaike criterion; maximum lags = 5.

respect and after introducing dynamics and an error correction mechanism, the estimation of Equation (1) was made using the PMG estimator proposed by Pesaran *et al.* (1999) and the MG estimator of Pesaran *et al.* (1996). Both approaches address the non-stationarity of time series for heterogeneous panels.

Generally speaking, an econometric specification of export 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-run and short-run coefficients across countries can be estimated consistently using the unweighted average of any individual coefficient estimated at country level. This is made possible 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 these two extremes there is

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

Table 5. Gregory–Hansen Test for Cointegration

Test statistic		Breakpoint date		Asymptotic critical values 1%	5%	10%
China						
ADF	−3.20	74	2008: Q3	−5.45	−4.99	−4.72
Z_t	−3.13	74	2008: Q3	−5.45	−4.99	−4.72
Z_a	−13.26	74	2008: Q3	−57.28	−47.96	−43.22
France						
ADF	−3.19	73	2008: Q2	−5.45	−4.99	−4.72
Z_t	−3.42	74	2008: Q3	−5.45	−4.99	−4.72
Z_a	−16.45	74	2008: Q3	−57.28	−47.96	−43.22
Germany						
ADF	−3.28	58	2004: Q3	−5.45	−4.99	−4.72
Z_t	−3.66	59	2004: Q4	−5.45	−4.99	−4.72
Z_a	−17.63	59	2004: Q4	−57.28	−47.96	−43.22
Italy						
ADF	−4.27	73	2008: Q2	−5.45	−4.99	−4.72
Z_t	−4.48	73	2008: Q2	−5.45	−4.99	−4.72
Z_a	−23.43	73	2008: Q2	−57.28	−47.96	−43.22
Japan						
ADF	−3.54	59	2004: Q4	−5.45	−4.99	−4.72
Z_t	−3.70	57	2004: Q2	−5.45	−4.99	−4.72
Z_a	−18.75	57	2004: Q2	−57.28	−47.96	−43.22
UK						
ADF	−3.46	74	2008: Q3	−5.45	−4.99	−4.72
Z_t	−3.05	75	2008: Q4	−5.45	−4.99	−4.72
Z_a	−13.74	75	2008: Q4	−57.28	−47.96	−43.22
USA						
ADF	−3.58	48	2002: Q1	−5.45	−4.99	−4.72
Z_t	−3.86	47	2001: Q4	−5.45	−4.99	−4.72
Z_a	−26.25	47	2001: Q4	−57.28	−47.96	−43.22

Source: Elaboration of data from DataStream.

Notes: ADF, augmented Dickey–Fuller. Number of observations = 89, lags = 2 chosen by Akaike criterion; maximum lags = 5.

the PMG method, which restricts the long-run coefficients to being 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, 23).[6]

The econometric specification of Equation (1) aligned to the PMG framework is as follows:

$$\Delta\log X_{i,t} = \delta_i + \beta_{1i}\Delta\log REX_{i,t} + \lambda_{1i}(\theta_1 REX_{i,t-1} - X_{i,t-1}) + \beta_{2i}\Delta\log Y^w_{i,t} + \lambda_{2i}(\theta_2 Y^w_{i,t-1} - X_{i,t-1}) + v_{i,t} \quad (5)$$

[6]Both MG and PMG estimations offer a good compromise between consistency and efficiency. The PMG is useful if countries share the determinants of the steady state, whereas the short-run adjustments are related to country characteristics. In other words, the PMG predicts a common long-run equilibrium relationship and short-run dynamics of each country. In brief, the MG framework 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).

with $v_{it} \sim \text{iid}N(0, \sigma_i^2)$ and ($i = 1, .., 7$; $t = 1, .., 89$). The MG specification differs from the PMG only for what concerns the long-run parameters θ_1 and θ_2, which, in the MG method, vary across countries. In other words, the subscript i is inserted in θ_1 and θ_2, consistent with the hypothesis of country-specific long-run equilibrium; that is: [7]

$$\Delta\log X_{i,t} = \delta_i + \beta_{1i}\Delta\log REX_{i,t} + \lambda_{1i}(\theta_{1i}REX_{i,t-1} - X_{i,t-1}) + \beta_{2i}\Delta\log Y_{i,t}^{w} + \lambda_{2i}(\theta_{2i}Y_{i,t-1}^{w} - X_{i,t-1}) + v_{i,t}\,. \quad (6)$$

To control for non-stationarity, the variables in Equations (5) and (6) are in first differences, as they are non-stationary in level. [8] The coefficients are short-run parameters, which, like, differ across countries. The error-correcting speed of adjustment term also differs across i. The long-run parameters θ_{i1} and θ_{i2} differ country-by-country for the MG estimator.

As mentioned above, short-run country heterogeneity is allowed for both estimators, while long-run elasticities differ country-by-country in the MG framework and are common across countries in the PMG framework. However, in using the MG estimator it is also

Table 6. Estimation of the Export Function of China and 6-OECDs
(PMG and MG Averaged Estimations over the Period 1990–2012)

PMG estimations	Coefficient	Standard error	Z	$P > \|z\|$	[95% confidence	interval]
Long- run						
ln(*REX*)	−0.8906	0.1350	−6.6	0	−1.15511	−0.6260
ln(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
Δln(*REX*)	−0.1734	0.0589	−2.94	0.003	−0.2889	−0.0580
Δln(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	**Coefficient**	**Standard error**	**Z**	**$P > \|z\|$**	**[95% confidence**	**interval]**
Long-run						
ln(*REX*)	−0.8663	0.2822	−3.07	0.002	−1.4194	−0.3133
ln(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
Δln(*REX*)	−0.1136	0.0677	−1.68	0.093	−0.2463	0.0191
Δln(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

Source: Elaboration of data from Datastream.

Notes: Observations = 616; number of Groups = 7; observations per group = 88. MG, mean group; PMG, pooled mean group; *REX*, real exchange rate; Y^w, world income.

[7] The PMG estimator is quite appealing when studying small sets of arguably "similar" countries rather than heterogeneous panels (Costantini and Destefanis, 2009; Eberhardt, 2011). The requirements for the validity of both these methods are such that: (i) there is a long-run relationship among the variables of interest; and (ii) the dynamic specification can be augmented such that the regressors are exogenous and the residuals are serially uncorrelated.

[8] The MG estimator offers the opportunity to obtain only one short-run and long-run elasticity value simply by averaging the estimations of each individual country.

Table 7. Estimation of the Export Function of China and 6-OECDs
(Results from Pooled Mean Group Estimator, 1990: Q1–2012: Q1)

	Coefficient	Standard error	Z	$P>\|z\|$	[95% confidence	interval]
Long-run						
ln(*REX*)	−0.8906	0.1350	−6.6	0	−1.1551	−0.6260
ln(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
Δln(*REX*)	0.0371	0.0608	0.61	0.542	−0.0820	0.1561
Δln(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
Δln(*REX*)	−0.3225	0.1258	−2.56	0.01	−0.5690	−0.0759
Δln(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
Δln(*REX*)	−0.1888	0.1795	−1.05	0.293	−0.5406	0.1630
Δln(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
Δln(*REX*)	−0.3261	0.0878	−3.71	0	−0.4982	−0.1539
Δln(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
Δln(*REX*)	0.0482	0.0732	0.66	0.511	−0.0953	0.1916
Δln(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
Δln(*REX*)	−0.2337	0.0988	−2.36	0.018	−0.4274	−0.0400
Δln(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
Δln(*REX*)	−0.2282	0.0750	−3.04	0.002	−0.3753	−0.0811
Δln(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

Source: Elaboration of data from Datastream.

Notes: Observations = 616; number of groups = 7; observations per group = 88; log likelihood = 1512.67. *REX*, real exchange rate; Y^w, world income.

possible to collapse short-run and long-run elasticities to their average values. The same applies for the PMG estimator regarding the short-run dynamics. Table 6 reports these results, while Tables 7 and 8 display the full estimates at country level.

All the elasticities have the expected sign and are highly significant. The results are twofold. On the one hand, exports are income elastic in the long run. Indeed, the income elasticity is higher than 1 both when using the PMG and the MG model, even though the magnitude of the effect differs: exports are more income elastic when considering the MG estimator instead of the PMG approach. A shock of 1 percent in world demand would result in an increase of exports of 1.08 percent under the PMG framework and 1.39 percent under the

Table 8. Estimation of the Export Function of China and 6-OECDs
(Results from Mean Group Estimator, 1990: Q1–2012: Q1)

	Coefficient	Standard error	Z	$P > \|z\|$	[95% confidence	interval]
China, long-run						
ln(*REX*)	−0.2207	0.3009	−0.73	0.463	−0.8104	0.3690
ln(Y^w)	1.5546	0.1527	10.18	0	1.2554	1.8538
China, short-run						
Error correction term	−0.1175	0.0455	−2.58	0.01	−0.2067	−0.0284
Δln(*REX*)	0.0430	0.0623	0.69	0.49	−0.0791	0.1650
Δln(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, long-run						
ln(*REX*)	−2.0405	0.5828	−3.5	0	−3.1828	−0.8982
ln(Y^w)	1.0052	0.1682	5.98	0	0.6754	1.3349
France, short-run						
Error correction term	−0.0764	0.0248	−3.08	0.002	−0.1251	−0.0277
Δln(*REX*)	−0.2626	0.1334	−1.97	0.049	−0.5241	−0.0012
Δln(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, long-run						
ln(*REX*)	−0.6702	0.1759	−3.81	0	−1.0150	−0.3254
ln(Y^w)	2.0309	0.0534	38.03	0	1.9263	2.1356
Germany, short-run						
Error correction term	−0.3287	0.0677	−4.86	0	−0.4613	−0.1961
Δln(*REX*)	0.1100	0.1775	0.62	0.536	−0.2380	0.4579
Δln(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, long-run						
ln(*REX*)	−0.7249	0.2217	−3.27	0.001	−1.1594	−0.2905
ln(Y^w)	0.9768	0.0947	10.32	0	0.7913	1.1624
Italy, short-run						
Error correction term	−0.1218	0.0344	−3.54	0	−0.1893	−0.0544
Δln(*REX*)	−0.3283	0.0899	−3.65	0	−0.5045	−0.1520
Δln(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, long-run						
ln(*REX*)	−0.5254	0.1469	−3.58	0	−0.8133	−0.2375
ln(Y^w)	1.3637	0.0975	13.98	0	1.1726	1.5549
Japan, short-run						
Error correction term	−0.2331	0.0501	−4.66	0	−0.3313	−0.1350
Δln(*REX*)	0.0619	0.0743	0.83	0.405	−0.0837	0.2075
Δln(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, long-run						
ln(*REX*)	−0.1159	0.3412	−0.34	0.734	−0.7846	0.5529
ln(Y^w)	1.4688	0.1706	8.61	0	1.1345	1.8031
UK, short-run						
Error correction term	−0.0990	0.0472	−2.1	0.036	−0.1915	−0.0065
Δln(*REX*)	−0.2270	0.1020	−2.23	0.026	−0.4270	−0.0271
Δln(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, long-run						
ln(*REX*)	−1.7666	1.1816	−1.5	0.135	−4.0825	0.5494
ln(Y^w)	1.3541	0.2893	4.68	0	0.7870	1.9212
USA, short-run						
Error correction term	−0.0502	0.0305	−1.65	0.1	−0.1100	0.0096
Δln(*REX*)	−0.1921	0.0810	−2.37	0.018	−0.3508	−0.0333
Δln(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

Source: Elaboration of data from Datastream.

Notes: Observations = 616; number of groups = 7; observations per group = 88. *REX*, real exchange rate; Y^w, world income.

Table 9. Long-run Elasticities: MG versus VECM Estimations

	MG		VECM	
	Price elasticity	Income elasticities	Price elasticity	Income elasticities
China	–0.22 [a]	1.55	–0.22 [a]	1.45
France	–2.04	1.01	–1,41	1.00
Germany	–0.67	2.03	–0.02 [a]	2.23
Italy	–0.72	0.98	–0.72	1.01
Japan	–0.52	1.36	–0.55	1.34
UK	–0.11 [a]	1.47	–0.83	1.60
USA	–1.77	1.35	–2.42	1.20

Source: Data elaborated from Datastream.

Notes: [a]not significant; MG, mean group; VECM, vector error correction model.

MG framework (Table 6). However, 1.08 is not statistically different from 1 and, thus, it could be argued that, under PMG estimation, exports have a unitary income elasticity. In contrast, the average long-run income elasticity for the MG estimation is statistically different from 1. [9] Our estimates reveal that the income sensitivity of exports is even higher in the short run, 3.8 being the average of the elasticities in PMG as well as in the MG model. A world income shock of 1 percent induces an increase of 3.8 percent in exports in the short run.

Turning to price elasticity, Table 6 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 for the PMG estimator and –0.86 for the MG estimator (the value from the MG estimator is the average of the elasticities predicted country-by-country). In both 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 the MG model to –0.17 under the PMG model. Based on these results, it is clear that the increase in national exports as a result of competitive devaluation is not so large to be considered aggressive in the world market. The evidence demonstrates that a real devaluation of 10 percent (as averaged across all countries in the sample) would induce an increase in exports of 8.6 percent in the long run and of, at best, 1.7 percent in the short run.

However, as already mentioned, the PMG framework 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 both in the long run and in the short run (Tables 7 and 8). It is interesting to note that the short-run elasticities and the adjustment terms do not differ when comparing MG and PMG results. This is

[9]For the PMG estimator we accept the null hypothesis of unitary elasticity (the test statistic is 1.58 with a p-value of 0.21), while for MG estimations we reject the null hypothesis as the test statistic is 8.50 (p-value = 0.0035).

[10]For the PMG estimator, the test statistic is 0.66 (p-value = 0.42), while for the MG estimator it is 0.22 (p-value = 0.64).

indirect proof that both models run pretty well in the short run and just differ with respect to the hypothesis regarding long-run behavior, thereby implying that it would not make a difference whether PM or PMG estimations were used for short run analyses.

Nevertheless, long-run elasticities vary at country level (Table 8) and, thus, it becomes important to verify which performs best, the MG or the PMG model. To this end we ran a likelihood-ratio test (LR). The two models are nested in each other: the PMG is the restricted model, while the MG estimator 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).

According to the LR results, we reject the null hypothesis: the LR yields a $\chi^2(12)=44.0$ with a *p*-value = 0. This means that the assumption that countries share the same equilibrium is unrealistic and not supported by the data. In contrast, we find that each country converges to its own long-run equilibrium. Based on this finding, our discussion then focuses only on the price and income elasticities estimated through the MG method. [11]

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 the MG results, we already know that the average long-run income elasticity is equal to 1.39 (Table 6). However, this value disregards high country heterogeneity. Indeed, foreign income results are very effective for Germany (the estimated elasticity is 2.03), China (1.55), the UK (1.46), Japan (1.36) and the 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 percent in world income occurred, then exports would increase, in the short run, by 6.94 percent in Japan, 4.06 percent in Italy, 3.9 percent in the UK, approximately 3 percent in China, France and Germany, and by 2.6 percent in the USA.

Furthermore, we reveal significant differences in the values of export price elasticity.

[11]To check the robustness of panel data estimations, we replicated the analysis at single-country level by using a VECM model. In this we follow Roudet *et al.* (2007), who investigate the long-run equilibrium paths of the real effective exchange rates of a sample of seven African developing countries. Our results regarding long-run elasticities are summarized in Table 9. Is it noteworthy that the sign and the statistical significance of each parameter do not vary when moving from MG to VECM models. Interestingly, even the magnitude of export elasticities is very similar. This contrasts with the evidence of Roudet *et al.* (2007), whose estimations are very sensitive to the estimation methods. In our case, the similarity in results is in favor of panel data estimations over individual time series as the former have the advantage of coming from a common analytical framework, thereby assuring a faithful comparability across countries.

This holds true in the long run and in the short run. In the long run, the analyzed countries have, as expected, a statistically significant negative coefficient with respect to 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. The result regarding China and the UK, whose exports are independent of price in the long run, is negative but not significant. The USA's exports exhibit a high (–1.77 percent) long-run price elasticity, although the statistical significance is just 13 percent. In brief, 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 percent in the presence of a real depreciation of 1 percent. For the other countries, real devaluation would induce an increase in exports but less than the relative change in national currency. Export 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), the UK (–0.23) and the USA (–0.19). Aggregate exports from China, Germany and Japan exhibit a wrong sign, but not significant, short-run price elasticity.

Over the 1990–2012 period, we find that China's exports are price insensitive both in the long run and in the short run. The same applies for the UK in the long run. The long-run price elasticity of the USA's exports is high, but not strongly significant. In the remaining cases, exports are price inelastic. The only exception is France, whose exports are price elastic in the long run and price inelastic in the short run. However, the finding that France performs differently from other exporters is not a novelty in this strand of the literature. For instance, in Crane *et al.* (2007) the price elasticity of France is 2.9, which is a high value compared to the values estimated in that work for Italy (0.7) and the USA (0.6). In Borey and Quille (2013), France also registers the highest value (1.1) of price elasticity (for the UK and Germany the values are 0.5 and 0.1, respectively). Evidently this mixed evidence reflects differences in the countries' export structure.

Discerning the causes of "high or low" price elasticity deserves further research based on highly disaggregated trade flows to capture the sectorial and geographic positioning and the quality ranges of each exporter. Nevertheless, some country peculiarities can be detected by using macro-level data. Table 10 displays some trade statistics for each exporter (we maximize the data availability of each source, whose time coverage differs from each other). The following facts may be highlighted.

The first regards the capital goods that have fewer close substitutes than other products and, therefore, are less sensitive to price: a low proportion of exports of capital goods is expected to be associated with high price elasticity. The results satisfy the expectations, as the correlation between the estimated long-run elasticity and the share of capital goods is 0.64. Importantly, among our sample countries, France registers the highest long-run price elasticity and the lowest (11 percent) proportion of capital goods exported. Conversely, the

Table 10. Export Price Elasticity and Countries' Export Structure

	Long-run exports price elasticity (MG)	Exports of capital goods as % of total exports (1995–2012) [a]	Proportion of services in exports [b]		Proportion of food in exports [b, c]		Exports to high-income economies (% of total exports) [b]	
Country			2005	2012	2012	1990–2012	1990–1991	2011–2012
China	−0.22 [d]	22	12	9	3	6	86	75
France	−2,04	11	21	27	13	13	84	80
Germany	−0,67	17	15	15	5	5	87	81
Italy	−0,72	15	20	17	8	7	85	78
Japan	−0,52	19	15	15	1	1	79	59
UK	−0,11 [d]	15	35	38	6	6	89	83
USA	−1.77*	16	29	30	10	9	75	61

Sources: [a]Data are from Comtrade (2-digit code "41" of BEC classification); [b]Data are from the World DataBank (World Development Indicators 2105).

Notes: [c]Food and live animals, beverages and tobacco, animal and vegetable oils and fats; [d]not-significant; **p*-value = 0.135. MG, mean group.

highest proportion (22 percent) of capital goods is found in China, whose exports are price inelastic. Similarly, services are more differentiated than goods. Then a high proportion of services in exports should be associated with low price elasticity. The contrary holds for food, which tends to be more homogeneous than other goods: hence, the higher the food content in exports the higher the price elasticity. In our case, price elasticity is wrongly correlated with services in exports, while the correlation between price elasticity and food in exports is high and, as expected, positive (0.75). Interestingly, the peak (13 percent) of the proportion of food in exports indicates France as the country with the highest estimated value of export price elasticity. Finally, we find that market destination matters in understanding cross-country differences in export elasticity: in our case the correlation with the share of exports to high-income countries is positive (although is not high and decreases from 0.23 in 1990–1991 to 0.19 in 2011–2012). These facts highlight the role of capital goods and food in exports and, at the same time, suggest that the explanation for heterogeneity in price elasticity requires more detailed study of the country specialization than the discussion we present here, and an extensive analysis of export quality.

IV. Concluding Remarks

This paper investigates the relationship between the real exchange rate and the export demand of seven exporting countries (China, France, Germany, Italy, Japan, the UK and the USA) over the period 1990–2012. The analysis is based on the model proposed by Goldstein and Khan (1985), while the econometric specification is adapted to non-stationary panel data and conducted using the PMG and the MG estimators. Importantly, 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. Because MG allows for full country heterogeneity of the relationships between exports and income and price, we may draw four general conclusions. First, 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). Second, each country accordingly converges towards its own long-run equilibrium with a country-specific speed of adjustment. Third, the differences in the short-run income and price elasticities underscore that the starting point of the transition path towards the final equilibrium varies country-by-country. Finally, the analysis allows us to glean the relevant finding that there is a stable long-run relationship between exports and the real exchange rate.

From an economic perspective, we find that the aggregate exports are highly income elastic in both the long run and the short run, implying that increases in world aggregate demand positively affect 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 price elasticity is –0.89, meaning that exports would increase by 8.9 percent after a 10-percent depreciation of the real exchange rate. In other words, total exports do increase in cases of competitive devaluation policies, but far less than the expansions one expects after having observed how intense and crude the tensions in currency markets are. The low export price sensitivity holds true when focusing on individual countries. Surprisingly, the nexus of export–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 signed wrongly (although, again, not significantly). 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 elasticity of –2, while its exports are price inelastic in the short run. A similar high long-run price elasticity (–1.77) is found for the USA, albeit it is weakly significant. Noticeably, these outcomes are robust over time, as there is no significant change in the long-run cointegrated 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 imbalances, which assumes that exports are highly price elastic. On the contrary, our findings suggest that devaluation gains are less than expected, because aggregate exports are price inelastic. This particularly holds true for China, as we find that the demand of importing countries rather than the price of exported goods plays a crucial role in boosting Chinese exports. We also report some evidence that this country is changing its export structure, from food and low-technology products to trade in mid-technology goods. If this process to gain position in the global value chain is stable over time, then the advantages that China would gain from updating the technological contents in exports will be less dependent on prices than in the past. Under these circumstances, a yuan appreciation would have a limited effect on Chinese exports owing to both the low level of real exchange rate elasticity and the productivity gains related to the structural changes under way in the country; the international pressure forcing Chinese authorities to appreciate the yuan becomes pointless.

A number of extensions to this paper could yield further insights into the nexus between exchange rates and exports. These might include the use of disaggregated trade flows, which would help to verify whether and to what extent export price elasticity is robust to market destination and product variety. Furthermore, future work could examine countries ' export structure by extending the estimation framework to allow for cross-country common factors, thereby controlling for unobserved time-varying omitted common variables.

References

Abadie, A. and J. Gardeazabal, 2003, "The economic cost of conflict: A case study of the Basque Country," *American Economic Review*, Vol. 93, No. 1, pp. 113–32.

Algieri, B., 2011, "Modelling export equations using an unobserved component model: The case of the euro area and its competitors," *Empirical Economics*, Vol. 41, No. 3, pp. 593–637.

Anaraki, N. K., 2014, "Effects of EURO devaluation on eurozone exports," *International Journal of Economics and Finance*, Vol. 6, No. 2, pp. 19–24.

Bahmani-Oskooee, M. and F. Niroomand, 1998, "Long-run elasticities and the Marshall–Lerner condition revisited," *Economic Letters*, Vol. 61, No. 1, pp. 101–9.

Bayoumi, T., Harmsen R. T. and J. Turunen, 2011, "Euro-area export performance and competitiveness," *IMF Working Papers*, IMF, Washington, DC.

Bodegan, C., B. Kraemer and M. Maucher, 2010, "The changing labour market in Germany in time of crisis," *WP19*, Institute of Economic and Social Research, Hans–Böckler Foundation, Germany, Dü sseldorf.

Borey, G. and B. Quille, 2013, "How to explain the recent shift in balance-of-trade trends in Europe?" *Special Analysis*, July 2013, National Institute for Statistics and Economic Studies, Paris.

Caporale, G. M. and M. K. F. Chui, 1999, "Estimating income and price elasticities of trade in a co-integration framework," *Review of International Economics*, Vol. 7, No. 2, pp. 192–202.

Chen, R., G. M. Milesi-Ferretti and T. Tressel, 2012, "External imbalances in the euro area," *IMF Working Paper* 12/236, IMF, Washington, DC.

Costantini, M. and S. Destefanis, 2009, "Cointegration analysis for cross-sectionally dependent panels:

The case of regional production functions," *Economic Modelling*, Vol. 26, No. 2, pp. 320–27.

Crane, L., M. A. Crowley and S. Quayyum, 2007, "Understanding the evolution of trade deficits: Trade elasticities of industrialized countries," *Economic Perspectives*, Vol. 31, No. 4, pp. 2–17.

Dezeure, N. and B. Teixeira, 2014, "Why are British exports so weak?" *Flash Economics*, Vol. 31, No. 236, pp. 1–11.

Dickey, D. A. and W. A. Fuller, 1981, "Likelihood ratio statistics for autoregressive time series with a unit root," *Econometrica: Journal of the Econometric Society*, Vol. 49, No. 4, pp. 1057–72.

Eberhardt, M., 2011, "Panel time-series modelling: New tools for analyzing xt data," *United Kingdom Stata Users' Group Meetings* 2011, Stata, London.

Gagnon, J. E., 2013, "Currency wars," *The Milken Institute Review*, Vol. 15, No. 1, pp. 47–50.

Goldstein, M. and M. S. Khan, 1985, "Income and price effects in foreign trade," in R. W. Jones and P. B. Kenen, eds, *Handbook of International Economics*, edn 1, Vol. 2, Netherlands: Elsevier, pp. 1041–105.

Gregory, A.W. and B. E. Hansen, 1996, "Residual-based tests for co-integration in models with regime shifts," *Journal of Econometrics*, Vol. 70, No. 1, pp. 99–126.

Grigor'ev, L. and M. Salikhov, 2009, "Financial crisis 2008. Entering global recession," *Problems of Economic Transition*, Vol. 51, No. 10, pp. 35–62.

Hamori, S. and Y. Matsubayashi, 2009, "Empirical analysis of export demand behavior of LDCs: Panel co-integration approach," *Economics Bulletin*, Vol. 29, No. 3, pp. 1990–99.

Hamori, S. and F. Yin, 2011, "Estimating the import demand function in the autoregressive distributed lag framework: The case of China," *Economics Bulletin*, Vol. 31, No. 2, pp. 1576–91.

Houthakker, H. S. and S. P. Magee, 1969, "Income and price elasticities in world trade," *The Review of Economics and Statistics*, Vol. 51, No. 2, pp. 111–25.

Im, K. S., M. H. Pesaran and Y. Shin, 2003, "Testing for unit roots in heterogeneous panels," *Journal of Econometrics*, Vol. 115, No. 1, pp. 53–74.

Jacobi, L. and J. Kluve, 2006, "Before and after the Hartz reforms: The performance of active labour market policy in Germany," *WP No.* 2100, IZA Institute, Germany, Bonn.

Joyce, M., A. Lasaosa, I. Stevens and M. Tong, 2011, "The financial market impact of quantitative easing in the United Kingdom," *International Journal of Central Banking*, Vol. 7, No. 3, pp. 113–61.

Kerr, W. A. and A. L. Hobbs, 2001, "Taming the dragon: The WTO after the accession of China," *Estey Centre Journal of International Law and Trade Policy*, Vol. 2, No. 1, pp. 1–9.

Ketenci, N. and I. Uz, 2011, "Bilateral and regional trade elasticities of the EU," *Empirical Economics*, Vol. 40, No. 3, pp. 839–54.

Kravis, I. and R. Lipsey, 1978, "Price behaviour in the light of balance of payments theories," *Journal of International Economics*, Vol. 8, No. 2, pp. 193–246.

Levin, A., C. F. Lin and J. C. S. Chu, 2002, "Unit root tests in panel data: Asymptotic and finite-sample properties," *Journal of Econometrics*, Vol. 108, No. 1, pp. 1–24.

Macis, M. and F. Schivardi, 2012, "Exports and wages: Rent sharing, workforce composition or returns to skills?" *Discussion Paper* No. 6466, IZA Institute, Bonn, Germany.

MacKinnon, J. G., 2010, "Critical values for co-integration tests," *Queen's Economics Department*

Working Paper No. 1227, Queen's University, Kingston, Ontario.

Orcutt, G. H., 1950, "Measurement of price elasticities in international trade," *The Review of Economics and Statistics*, Vol. 32, No. 2, pp. 117–32.

Pesaran, M. H., Y. Shin and R. P. Smith, 1999, "Pooled mean group estimation of dynamic heterogeneous panel," *Journal of the American Statistical Association*, Vol. 94, No. 446, pp. 621–34.

Pesaran, M. H., Y. Smith and K. S. Im, 1996, "Dynamic linear models for heterogeneous panels," in L. Mátyás and P. Sevestre, eds, *The Econometrics of Panel Data. A Handbook of the Theory with Applications*, Part I, Springer, Netherlands, pp. 145–95.

Roudet, S., M. Saxeggard and C. G. Tsangarides, 2007, "Estimation of equilibrium exchange rates in the WAEMU: A robustness analysis," IMF-WP-07/194, IMF, Washington, DC.

Sawyer, W. C. and R. L. Sprinkle, 1996, "The demand for imports and exports in the U.S.: A survey," *Journal of Economics and Finance*, Vol. 20, No. 1, pp. 147–78.

Enhadji, A. S. and C. E. Montenegro, 1999, "Time-series analysis of export demand equations: A cross-country analysis," *IMF Staff Papers*, IMF, Washington, DC.

Stern, R. M., J. Francis and B. Schumacher, 1976, *Price Elasticities in International Trade: An Annotated Bibliography*, London: Macmillan for the Trade Policy Research Centre, Macmillan, London.

Thorbecke, W., 2011, "Investigating the effect of exchange rate on China's processed exports," *Journal of the Japanese and International Economics*, Vol. 25, No. 2, pp. 33–46.

Thorbecke, W. and A. Kato, 2012, "The effect of exchange rate changes on Japanese consumption exports," *Japan and the World Economy*, Vol. 24, No. 1, pp. 64–71.

Westerlund, J., 2007, "Testing for error correction in panel data," *Oxford Bulletin of Economics and Statistics*, Vol. 69, No. 6, pp. 709–48.

Yao, Z., F. Tian and Q. Su, 2013, "Income and price elasticities of China's exports," *China & World Economy*, Vol. 21, No. 1, pp. 91–106.

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