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Real effective exchange rate and the constant elasticity of substitution assumption

Antonio Spilimbergo*, Athanasios Vamvakidis

International Monetary Fund, 700 19th Street, N.W., Washington, DC 20431, USA

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Abstract

The real effective exchange rate (REER) is an aggregation of several bilateral real exchange rates assuming constant elasticity of substitution (CES) between goods from different countries. We investigate the validity of the CES assumption by estimating manufacturing export equations for 56 countries over 26 years. Under the CES assumption, splitting the REER into two components should not increase the fit in an export equation and the coefficients on the two REERs should be equal. We reject both these implications and find that the export equations with two REERs—vs. OECD and vs. nonOECD countries—perform better than the traditional ones.

Keywords: Foreign exchange; Trade forecasting and simulation; Panel unit root tests; Panel cointegration

JEL classification: F31; F17

1. Introduction

A commonly used measure of competitiveness is the real effective exchange rate (REER), calculated as a weighted average of the relative prices of a country with its main trade competitors. The trade share of a country in each industry and the importance of each industry in the total trade of each country determine the

*Corresponding author. Tel.: +1-202-623-6346; fax: +1-202-623-4661.
E-mail address: aspilimbergo@imf.org (A. Spilimbergo).

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weights in these calculations. This methodology takes into account both domestic and third-market effects. The Information Notice System (INS) of the IMF provides such weights for almost all countries. INS

The current methodology for determining the weights in the REER estimations assumes that the elasticity of substitution is constant for products coming from different countries. For example, the level of development of the country of origin for each product is not relevant (the elasticity of substitution between US products and Japanese products is equal to the elasticity of substitution between US products and any other country's products, either developed or developing). This assumption becomes even stronger in practice since the REER is calculated based on total trade weights, and, therefore, the elasticity of substitution is assumed to be constant even across different products.

If this assumption does not hold, then the commonly used aggregated REERs are not accurate, since each country's weights would depend on the elasticity of substitution with respect to each other country. For example, a country with a large trade share in an industry but with a relatively small elasticity of substitution for the exports of all other countries in this industry should have a relatively small weight in the REER calculation. In this case, the weights used in the existing literature to calculate the REER would result in an inaccurate picture of which countries are the main competitors of a given country.

The debate on a possible devaluation in China during the East Asian financial crisis of 1997–1998 illustrates this point. A concern often expressed during that time was that if the Chinese authorities devalued their currency, they would fuel a new wave of competitive devaluations in the area. However, based on the INS weights, this concern was not justified. According to the INS weights, China's weight is less than 3% for all East Asian economies, except for Hong Kong POC. If we consider the list of the 10 major competitors for each East Asian economy, defined as the 10 economies with the largest weights, China makes it to the list only for Hong Kong POC (Hong Kong POC in turn makes it to the list only for Korea and Singapore, but with very small weights, 2.5% and 3%, respectively). However, most economists and politicians involved in the East Asian financial crisis would disagree that devaluation in China during this period would have had no impact on other East Asian economies. If they were right, then the current REER weights have to be wrong. Indeed, if China's exports have a relatively high elasticity of substitution for the exports of the other East Asian economies, then China should be assigned a higher weight than is currently assigned in the standard REER calculation for these economies.

This paper tests the validity of the assumption that the elasticity of substitution is constant between products coming from different countries. The test determines whether the assumption holds, and whether it matters when it does not, by estimating manufacturing export equations for 56 countries, for a period of 26 years. The equations include either the standard aggregated REER or two REERs disaggregated into two components—one REER with respect to OECD countries and one with respect to nonOECD countries—as well as other alternative disaggregations. As explained below, if the assumption of the constant elasticity of substitution holds, any disaggregation of the REER should worsen the fit of the export equation, while the estimated coefficients on the disaggregated REERs should be equal.

The data reject the hypothesis of constant elasticity of substitution because the estimated coefficients on the two REERs are significantly different, and export equations with two REERs perform better than export equations with only one aggregated REER. These results are robust to panel and time series regressions, to regressions in levels and in differences, as well as to several estimation techniques.

The disaggregation of the standard REER in the OECD REER and the nonOECD REER is arbitrary, but justified by the idea that products from similar countries may have a constant elasticity of substitution. We assess the plausibility of this idea by performing a simulation exercise in which we show that the OECD vs. nonOECD disaggregation increases the fit with respect to randomly chosen groups of countries in most cases. We also split the REER into one in relation to neighboring countries and one in relation to non-neighboring countries, which also performs better than the aggregated REER, but worse than a grouping of OECD versus nonOECD countries.

Other studies that have disaggregated the REER in a similar way have also found that the explanatory power of trade equations improves significantly. Giorgianni and Milesi-Ferretti (1997) split the exchange rate for industrialized versus non-industrialized countries to explain export demand for Korea. Faini et al. (1992) used such disaggregation to determine the benefits of devaluation for developing economies following an export-led strategy, when other developing economies were following similar policies. These studies in part motivate the disaggregation chosen in our paper.

This paper focuses only on the REER relevant for international trade. There are many different methodologies to calculate the REER, depending on its use. Maciejewski (1983) provides a review of the different real exchange rates. The calculation based on trading weights is more relevant for trade purposes, which is why it is commonly used in export equations.

This paper focuses only on manufacturing exports. Weights for other sectors such as agriculture and tourism are available from INS, but a disaggregation of the REER to OECD and nonOECD countries is not clearly justified for these sectors. Furthermore, it may be more difficult for the CES assumption to hold for the manufacturing sector, since many manufacturing products are in oligopolistic markets.[1]

As noted above, since the weights in the calculation of the REER in the existing

[1] See, for example, Berry et al. (1995) for an estimation of demand parameters in the oligopolistic differentiated US automobile industry, and Blonigen and Wilson (1999) for an estimation of Armington elasticities between US domestic and foreign goods across 100 industrial sectors.

literature are based on aggregate trade shares, the CES assumption is made even for different products within manufacturing. This aggregation makes the CES assumption considerably stronger, since, even if the elasticity of substitution is constant for the same manufacturing product across countries, it may not be constant across different products. In this case, if OECD and nonOECD countries export different products, then a disaggregation of the REER with respect to these two groups should improve the explanatory power of trade equations.

The rest of the paper is organized in three main sections: Section 2 gives a brief description of the theoretical foundations for the empirical part of the paper; Section 3 presents the methodology, the empirical results and robustness tests; and Section 4 concludes the paper. Appendix A has the data sources and description.

2. Theoretical foundations

This section briefly discusses the form of the export demand equation in the general case, in which the elasticity of substitution of products coming from different countries varies, and the CES case, in which the elasticity of substitution is constant. Since such models have been discussed extensively in the literature, this section focuses on the main results, while giving references for their derivation.

2.1. The general case

The export demand equation when the elasticity of substitution of products coming from different countries is not the same has the following form:

$$\ln D_j = b + \ln D + \sum_{l \neq j}^{m} \sigma_{lj} W_{lj} \ln \left(\frac{P_l}{P_j} \right) \tag{1}$$

where b is a constant term, $D_j = P_j X_j$ is the demand for country j's exports, P_j is country j's export price index, X_j is country j's real exports, D is total world demand, σ_{lj} is the elasticity of substitution between products l and j, $W_{jl} = \Sigma_{\psi=1}^{m} T_j^{\psi} S_l^{\psi}$ is the weight of country j on country l, $S_l^{\psi} = (P_l X_l^{\psi} / \Sigma_{k=1}^{m} P_k X_k^{\psi})$, is the share of exports of country l in country ψ over the total external demand in country ψ, $T_j^{\psi} = (P_j X_j^{\psi} / \Sigma_{\psi=1}^{m} P_j X_j^{\psi})$ is the share of country's j exports in country ψ over the total exports of country j, and there are m countries. Therefore, the weight of country j on country l depends on two other weights: firstly, a weight for the importance of country l in each market, and secondly, a weight for the importance of each market in country l's exports. The last term of Eq. (1) is the real effective exchange rate, but after multiplying each price ratio with the respective elasticity of substitution.

This export equation is based on the derivation described in Artus and McGuirk

(1981) and McGuirk (1987). Similar models that assume that the CES assumption does not hold have been discussed by Johansen (1969), Barten (1977), and Caplin and Nalebuff (1991), among many other papers.

2.2. The CES case

If we assume that the elasticity of substitution of products coming from different countries is constant, then Eq. (1) has the following form:

$$\ln D_j = b + \ln D + \sigma \sum_{l \neq j}^{m} W_{lj} \ln \left(\frac{P_l}{P_j} \right) \tag{2}$$

where we have assumed that $\sigma_{lj} = \sigma$. The last term in Eq. (2) is the real effective exchange rate, multiplied by the constant elasticity of substitution. This export equation is derived in Armington (1969), and is discussed in Artus and McGuirk (1981) and McGuirk (1987). The CES case is also discussed in Dixit and Stiglitz (1977), in Helpman and Krugman (1985), and in Caplin and Nalebuff (1991), among many other papers.

Both of these cases, the general and the CES case, assume that products coming from different countries are not perfect substitutes (imperfect substitutes model). However, the second case adds the CES assumption. Because of this assumption, while a different elasticity of substitution multiplies each price term in Eq. (1), the same elasticity of substitution multiplies all price terms in Eq. (2).

There are many reasons to believe that the CES assumption may not hold and that the general case is true. For example, a product produced and exported by a developed country is often very different from a product produced and exported by a developing country, even when the two products are in the same industry. Products coming from developed countries are often more advanced technologically and more expensive than products coming from developing economies, and they often target different consumer groups.

If we assume that the elasticity of substitution is the same for products coming from similar countries, then a version of Eq. (1) holds. To simplify the analysis, we assume that the elasticity of substitution is constant within the groups of developing and developed countries, but varies for countries from different groups. In this case, Eq. (1) has the following form:

$$\ln D_j = b + \ln D + \sigma^{\beta} \sum_{l \neq j}^{n} W_{lj} \ln \left(\frac{P_l}{P_j} \right) + \sigma^{\gamma} \sum_{k \neq j}^{s} W_{kj} \ln \left(\frac{P_k}{P_j} \right) \tag{3}$$

where σ^{β} is the elasticity of substitution between products coming from developed countries, σ^{γ} is the elasticity of substitution between products coming from developing countries.

3. Empirical analysis

According to the CES assumption, the elasticity of substitution between products coming from different countries is constant. The purpose of this section is to test the validity of this assumption.

An export demand equation is a natural test of the CES assumption. If the CES assumption is valid, splitting the real exchange rate into two or more components should not increase the predictive power of the export demand equation. If it does increase its predictive power, then the CES assumption does not hold. This empirical experiment is conducted for a set of 56 countries during a period of 26 years—1970 through 1995.

Ideally, we should estimate a system comprising export demand and export supply for each country (see Goldstein and Khan, 1985). However, uncertainty about the supply model specification and the lack of proper supply data for many countries makes this task impossible. There is a trade-off between estimating the demand equation with potential simultaneity bias, and estimating a system of two equations with potential misspecification of the supply curve. The second problem becomes more relevant when working with panel data from many different countries, because the uncertainty about the supply equations concerns all countries in the sample.[2] Moreover, estimating only a demand equation does not give rise to simultaneity bias when the supply curve is perfectly elastic to price. Evidence in support of a flat supply curve has been found by Muscatelli et al. (1995), Duttagupta and Spilimbergo (2000), and Giorgianni and Milesi-Ferretti (1997), which justifies focusing only on the demand side of exports.[3]

The ideal specification of an export demand equation should split the real exchange rate into bilateral components—potentially using all the bilateral real exchange rates. Obviously, lack of degrees of freedom makes this estimation impossible. An intermediate solution, as explained in Section 2, is to use two exchange rates: for example, one with respect to OECD countries and the other with respect to nonOECD countries. This split assumes that the elasticity of substitution is the same within the groups of the OECD and nonOECD countries, but different between the two groups. Other splits could be used, as discussed in more detail below. The demand equation to estimate is similar to Eq. (3):

$$\ln(D_{jt}) = \alpha_j + \beta_j \ln\left(\frac{P_{l1t}}{P_{jt}}\right) + \gamma_j \ln\left(\frac{P_{l2t}}{P_{jt}}\right) + \theta_j \ln(TRMI_{jt}) + \varepsilon_{jt} \tag{4}$$

where D_{jt} is an index of manufacturing export volume; P_{l1t}/P_{jt} is an index of

[2] This is probably why many authors focus only on the demand side when working with relatively large panel data (see Reinhart, 1995).

[3] However, Riedel (1989) argued that specification of export demand and supply can drive the results on elasticity. In particular, he argues that export supply is more upward sloping than what is found in other studies.

relative prices with respect to one group of countries (for instance, OECD countries); P_{l2t}/P_{jt} is an index of relative prices with respect to all other countries (for instance, nonOECD countries); and these price ratios are weighted by the INS weights for manufacturing exports; $TRMI_{jt}$ is the weighted average of the main trading partners' real manufacturing imports; and all the export prices are in US dollar terms.[4] Under the null hypothesis that the CES assumption holds, we should find that $\beta_j = \gamma_j$ and that the fit of the export equation increases if the two exchange rates are collapsed into one. Under the null hypothesis, Eq. (4) collapses into a standard demand equation:

$$\ln(D_{jt}) = \alpha_j + \beta_j \ln\left(\frac{P_{lt}}{P_{jt}}\right) + \theta_j \ln(TRMI_{jt}) + \varepsilon_{jt}. \quad (5)$$

The existing REER weights assume equal weights for all domestic and foreign producers in a third market (domestic producers have a weight of 0.5, while all foreign producers combined have a weight of 0.5). This assumption is strong and arbitrary, but there are no easy alternatives, since there are no comprehensive input–output tables for most countries (Wickham, 1987). Therefore, this assumption is taken as a given, and is left to future research to test its validity and importance.

There are two major econometric issues involved in the estimation of Eq. (4). Firstly, it is necessary to establish the dynamic properties of the data in order to use the proper estimation techniques; secondly, a test should determine whether it is possible to pool the data and estimate the system as a panel, or to estimate a separate equation for each country. The section proceeds by addressing these two issues.

3.1. Dynamic properties of the data

At a theoretical level, it is not clear whether we should estimate an export demand equation in differences or in levels, and, therefore, the issue should be resolved empirically.[5]

A first pass is to test the data for the presence of unit roots by applying standard single equation methodologies—the augmented Dickey–Fuller (ADF) and Phillips–Perron tests. The ADF tests reject a unit root at the 5% significance level only in eight countries for the real exchange rate, in six countries for manufacturing

[4]The weights for trading partners' imports are also based on INS. We experimented with several other scale variables (e.g. world GDP, world trade). While the subsequent results do not change, we have used trading partners' imports because it is time- and country-specific.

[5]Senhadji and Montenegro (1999) derive an export equation from an intertemporal maximization model and show that estimation of both in levels and in differences are theoretically justified, depending on the nature of the income innovations.

exports, and in 11 countries for the trading partners' demand; the evidence is similar for Phillips–Perron tests.[6]

However, given the low power of these tests in relatively small samples, it is difficult to reject the null hypothesis of unit roots.[7] Therefore, we investigate the issue of the stationarity in the data by using two tests for unit roots in the context of panel data.[8] We exclude from the panel test the countries for which the presence of a unit root is already rejected at the 5% significance level in the single equation approach. Im et al. (1995) propose a test based on separate Dickey–Fuller unit root tests for each of the N cross-section units. They show that the average of the t-statistics associated with the unit root tests is distributed according to a normal variable with mean μ and variance σ^2 that depends on N and T. The same study also suggests that this test can be extended to the case of panels with unobserved common time-specific components by taking out cross-sectional averages; however, they warn that this procedure is not robust to misspecification of time trends, if the effect of the common component varies across groups.

Maddala and Wu (1999) proposed an alternative test based on the P-values of the separate Dickey–Fuller unit root tests for each of the N cross-section units. Under the assumption that the tests are independent, the sum of the (log) significance levels P_i is distributed as a χ^2 with $2N$ degrees of freedom. Unfortunately, the P-values associated with the augmented Dickey–Fuller test are not tabulated. We constructed our P-values with a bootstrapping of 10,000 draws. The results of both tests are reported in Table 1, where we consider three groupings for the REER: the aggregated REER, splitting the REER with respect to OECD and nonOECD countries, and splitting the REER with respect to neighboring and non-neighboring countries (defined as countries in the same or different continent). Both tests are more general than the Levin and Lin (1993) test, since each panel group is allowed to have a different process under the alternative hypothesis.

Both tests fail to reject the null hypothesis of a unit root for all the variables in the majority of countries.[9] Given these results, it is appropriate to test whether

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[6]The countries for which a unit root in the real exchange rate is rejected at the 5% level are: Australia, Canada, Finland, Ireland, Mauritius, Morocco, New Zealand, and Thailand.

[7]For instance, through a Monte Carlo simulation, the percentage of rejection of a unit root at the 5% confidence interval is below 6% when the alternative of a $\rho = 0.96$ is true. We thank Nelson Mark for sharing his simulations on this point.

[8]Levin and Lin (1993) propose such a test, in which the null hypothesis is the presence of unit root ($\rho_i = 1$ for all i), while the alternative is that none of the panels has a unit root and all follow the same dynamic process ($\rho_i = \rho < 1$ for all i). We do not use such a test here because the alternative hypothesis is too restrictive.

[9]In recent years, there has been a considerable literature on the existence of unit roots for real exchange rates. O'Connell (1998) finds that panel data evidence in favor of stationarity of the REER disappears if the testing procedure controls for cross-sectional dependence among the error terms; however, Higgins and Zakrajšek (1999) use SUR techniques correcting for upward bias to argue for REER stationarity. Our results are not directly comparable, given that we considered an exchange rate based only on manufacturing products.

Table 1
Panel unit root tests

	Number of countries	Im–Pesaran–Shin	Im–Pesaran–Shin (controlling for time specific effects)	Maddala–Wu
REER	48	−1.20 (0.12)	−1.76 (0.04)	99.03 (0.40)
REER vs. OECD countries	47	−1.24 (0.11)	−1.65 (0.05)	97.26 (0.39)
REER vs. nonOECD countries	44	−2.26 (0.01)	−0.25 (0.40)	119.87 (0.07)
REER vs. neighbors	49	−0.35 (0.36)	−0.72 (0.23)	91.56 (0.66)
REER vs. non-neighbors	51	−2.37 (0.01)	−2.54 (0.01)	122.99 (0.08)
Log (trading partners' demand)	45	−3.42 (0.00)	−2.46 (0.01)	132.61 (0.00)
Log (manufacturing)	50	0.04 (0.52)	0.66 (0.74)	88.98 (0.78)

The *P*-values are in parentheses. The tests are for the null hypothesis that all countries have a unit root; the alternative hypothesis is that some (possibly all) countries have no unit root. The IPS statistic is distributed according to a Normal(0,1) and the Maddala–Wu statistics is distributed according to a χ^2 with $2N$ degrees of freedom. As explained in the text we exclude from the panel all countries for which the standard augmented Dickey–Fuller test on an individual country rejects a unit root at the 5% level.

there is a cointegrating relationship between (log) manufacturing export, (log) trading partners' demand, and various definitions of (log) real exchange rate.

As for the unit root tests, we perform a cointegration test for each country taken singularly and in a panel framework. The standard single-equation residual-based tests, as suggested by Engle and Yoo (1987), reject the null hypothesis of no cointegration at the 10% significance level in 14 countries, and at the 5% significance level in five countries.[10] If a trend is introduced, the null hypothesis of no cointegration is rejected in three cases at the 10% significance level, and in one case at the 5% significance level.

However, the power of tests based on single groups is limited, and therefore, panel cointegration tests proposed by Pedroni (1999), involving the estimation of seven different statistics, may be more appropriate. Under the null hypothesis of no cointegration, all the statistics are distributed as a normal (0, 1). Under the alternative hypothesis, the first statistic diverges to positive infinity, while the other six statistics diverge to negative infinity. Accordingly, the right tail of the normal distribution is used to reject the null hypothesis in the first statistic, while the left tail is used for the other six statistics.

Table 2 reports the results for the panel cointegration tests (the corresponding P-values are reported in parentheses). Five out of seven tests strongly reject the null of no cointegration, and two out of seven do, if heterogeneous trends are allowed.[11] Given these results, we estimate Eqs. (4) and (5) in levels, imposing a

Table 2
Panel cointegration tests among the variables: ln (D_{jt}), ln (P_{l1}/P_{jt}), ln ($TGDP_t$)

	With common trend	With heterogeneous trend
Panel v-stat	2.89 (0.00)	1.11 (0.13)
Panel rho-stat	−0.54 (0.29)	1.82 (0.97)
Panel pp-stat	−2.46 (0.01)	−0.83 (0.20)
Panel adf-stat	−4.30 (0.00)	−3.04 (0.00)
Group rho-stat	1.61 (0.95)	3.80 (1.00)
Group pp-stat	−1.75 (0.04)	−0.12 (0.45)
Group adf-stat	−7.03 (0.00)	−4.71 (0.00)

Null hypothesis is no cointegration in all countries. Under the null tests all the statistics are distributed as a normal (0, 1). See text for explanations.[17]

[10] The sample includes only countries for which the presence of a unit root was not rejected at the 5% level in the real exchange rate. The critical levels for the cointegration tests are based on an augmented Dickey–Fuller test for the residuals, using MacKinnon's (1991) estimated response surface regressions.

[11] We thank Peter Pedroni for useful clarification on these tests. The panel consists of 47 countries for which the single tests fail to reject the hypothesis of no cointegration at 5%. We also tried the same exercise excluding the four countries for which the single equation approach rejects the null of no cointegration at the 10% level, finding similar results.

[17] The statistics shown in Table 2 are calculated using a program kindly provided by Peter Pedroni.

cointegrating relationship. We restrict our sample to the 47 countries for which we reject the hypothesis of no cointegration at 5%.

Our empirical strategy is in line with several other studies that find and estimate a cointegrating relationship in the export equation (for instance, Reinhart (1995) and Bayoumi (1999)).

3.2. Should the data be pooled together?

A Chow test on the joint restrictions ($\beta_j = \beta \cup \theta_j = \theta$) in Eq. (4) can determine whether the data can be pooled together in a panel. This test is applied separately to OECD and nonOECD countries, to see whether countries within each group behave differently. Given that the presence of a cointegrating relationship has not been rejected, we use dynamic ordinary least square (DOLS).

The results in Table 3 show that the null hypothesis is always rejected.[12] Rejecting the hypothesis that the coefficients are the same across the panel does not imply that estimating the pooled data is less efficient. If the final aim is to obtain more precise estimates of the parameters, one may prefer a biased estimator with low variance to an unbiased estimator with high variance (see Baltagi, 2001). What follows presents both panel estimates and estimates based on single equations.

3.3. Results

Table 4 shows the average P-value associated with an F-test for all coefficients equal to zero, adjusted R^2 and coefficients obtained by estimating Eqs. (4) and (5) for each country. We exclude from the sample the countries for which the cointegration is rejected at 5%, so that the number of countries is 47 (see Section 3.1). The equation for each country is estimated with DOLS, allowing for one lag and one lead of the right-hand side variables.

In the context of cointegration, the P-value attached to an F-test on the hypothesis that all coefficients in the regression are equal to zero is a good

Table 3
Test for poolability. F-tests for the restrictions ($\beta_j = \beta \cup \theta_j = \theta$)

	F-test	P-values
All countries	7.45	0.00
OECD	10.11	0.00
NonOECD	10.80	0.00

We use DOLS estimators, allowing for two lags and two leads in the right-hand side variables.

[12] These results are confirmed if we allow country-specific trends.

Table 4
Average coefficients (Eqs. (4) and (5))

Variable	Mean	S.D.	Minimum	Maximum
P-value (5)	0.03	0.14	0.00	0.84
P-value (4)	0.01	0.06	0.00	0.41
Adjusted R^2 (5)	0.82	0.26	−0.18	0.99
Adjusted R^2 (4)	0.85	0.22	0.03	0.99
β_j (4)	−0.26	2.47	−8.38	7.25
t-stat (β_j) (4)	0.00	2.77	−5.18	11.21
γ_j (4)	8.48	14.33	−14.72	73.31
t-stat (γ_j) (4)	1.74	1.80	−1.75	5.19
Test ($\gamma_j = \beta_j$) (4)	5.96	7.82	0.02	32.56
P-value (4)	0.31	0.35	0.00	0.99

Each equation is estimated with DOLS. Number of countries: 47. See text for explanations.

measure of fit. Using this criterion, specification (4) dominates specification (5) on average. Note that a comparison of adjusted R^2s may not be proper in the context of cointegration, given that asymptotically all R^2s converge to zero. With this caveat, the adjusted R^2s are higher for Eq. (4) than for Eq. (5).

As noted above, if the existing REER weights are correct, the two coefficients on the exchange rate in Eq. (4) should be the same. However, the average γ_j is equal to 8.48, while the average β_j is equal to -0.26. Unfortunately, these coefficients are estimated quite imprecisely for some countries, and this makes it difficult to test the hypothesis that $\gamma_j = \beta_j$ for all j. However, the null hypothesis that $\gamma_j = \beta_j$ can be rejected at the 10% significance level based on a t-test for each equation in 24 out of 47 countries, and in 18 cases at the 5% significance level. For the remaining countries we fail to reject the hypothesis that $\gamma_j = \beta_j$, primarily because the estimated parameters are not precise and the power of the t-test is very low in the single equation setting.

In order to investigate further whether $\gamma_j = \beta_j$ and to avoid the problem of lack of power in the single equation approach, three approaches are followed. The first applies the procedure suggested by Fisher (1970) to test the significance of the results from N independent tests of hypothesis as explained in Maddala and Wu (1999): '... if the test statistics are continuous, the significance levels P_i ($I = 1, 2, \ldots, N$) are independent uniform (0, 1) variables, and $-2 \ln (P_i)$ has a χ^2 distribution with two degrees of freedom. Using the additive property of χ^2 variables, we get $\lambda = -2 \ln (P_i)$ has a χ^2 distribution with $2N$ degrees of freedom...'. In our case, λ is 270.93 and the corresponding P-value is 0.00 so that we can reject the hypothesis that $\gamma_j = \beta_j$ for every j.

The second approach is to estimate a panel. While this will lead to biased estimates, as noted above, it will gain more power for the test. Given the cointegrating relationships, we use panel DOLS, which includes one lead and lag of all explanatory variables. Table 5 reports the results.

Table 5
Eqs. (4) and (5) estimated as a panel with fixed effects. Dependent variable ln (manufacturing)

	All countries		OECD		NonOECD	
Equations	5	4	5	4	5	4
REER	0.93		0.42		1.12	
	(0.00)		(0.00)		(0.00)	
REER with OECD (β_j)		0.58		0.34		0.73
		(0.00)		(0.00)		(0.00)
REER with nonOECD (γ_j)		2.76		1.25		2.54
		(0.00)		(0.03)		(0.00)
Ln (TRMI)	0.70	0.77	0.67	0.70	0.75	0.85
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Observations	1081	1081	391	391	690	690
Number of countries	47	47	17	17	30	30
Log-likelihood	−99.58	−80.51	318.76	324.22	−387.25	−379.11
t-test on $\gamma_j = \beta_j$		29.79		1.89		12.35
		(0.00)		(0.17)		(0.00)

The six equations are estimated with a panel DOLS. P-value of t-statistics in parentheses. We allow for country-specific heteroskedasticity so that the panel is estimated as a GLS. Number of countries: 47.

The first two columns of Table 5 report the results for the panel DOLS estimation with fixed effects of Eqs. (4) and (5) for all countries. All the coefficients have the expected sign and are significant. The two coefficients of the REER with OECD and nonOECD countries have quite different values, and we reject that they are equal at the 1% significance level.

Between the two extreme approaches of estimating equation by equation and estimating a panel, there is the intermediate approach, which consists of grouping countries into broad groups with similar characteristics. This can be done by splitting the countries in the sample into 17 OECD countries and 30 nonOECD countries, as shown in the last four columns of Table 5. Splitting the sample in this way reveals an interesting difference: while we reject the null hypothesis of equal coefficients for the two REERs in nonOECD countries, we fail to reject the null for OECD countries. Finally, the coefficients of the partners' manufacturing imports are higher in nonOECD countries than in OECD countries, consistent with the idea that nonOECD exports are more homogeneous and thus less price sensitive.

The third approach consists of an F-test based on group mean fully modified OLS (FMOLS) as described in Pedroni (2000, 2001).[13] We estimate the following two panel regressions:

$$\ln (D_{jt}) = \alpha_j + \beta_j \ln \left(\frac{P_{l1t}}{P_{jt}}\right) + \theta_j \ln (TRMI_t) + \varepsilon_{jt} \tag{6a}$$

[13]We thank an anonymous referee for suggesting this procedure.

and

$$\ln(D_{jt}) = \alpha_j + \gamma_j \ln\left(\frac{P_{l2t}}{P_{jt}}\right) + \theta_j \ln(TRMI_t) + \varepsilon_{jt}. \tag{6b}$$

Using the group mean panel FMOLS and allowing for heterogeneous trends, we estimate the group means $\bar{\beta}$ and $\bar{\gamma}$. An F-test of the null that $\bar{\beta} = \bar{\gamma}$ gives a value of 3.85 with a P-value of 0.00; therefore, this non-nested test also clearly rejects the null hypothesis.[14]

While the majority of the 47 equations perform reasonably, some equations do not behave very well (for instance, the maximum value for γ_j is due to an outlier). In an unreported exercise, we re-estimated our panel excluding one country at a time. We found that our panel results do not depend on the inclusion of specific countries and are robust to the exclusion of outliers.[15]

3.4. Is OECD versus nonOECD countries grouping efficient?

In the previous sections, we have compared the standard REER with two REERs, one with respect to the OECD and one with respect to the nonOECD countries. As already noted above, this grouping is arbitrary, but justified by the idea that products from similar countries may have a constant elasticity of substitution.

Another plausible grouping is based on the proximity of the trading partner, so that one exchange rate is vis-à-vis countries in the same continent and the other is vis-à-vis countries in all other continents (neighboring vs. non-neighboring countries). In order to evaluate the predictive power of these two alternative groupings, we compare the F-statistics using this splitting for each country. In 33 out of 56 countries, the OECD vs. nonOECD is a superior splitting of the sample. Interestingly, many countries for which the regional grouping works better are in Europe, suggesting a peculiarity of trade within the European Union.

An alternative approach is to see how the OECD vs. nonOECD split compares with random splits. For this exercise, we select random pairs of country groups, and calculate two REERs vis-à-vis these two groups for each country.[16] We then

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[14] In order to perform this test we use a modified version of a program provided by Peter Pedroni and described in Pedroni (2000).

[15] The previous estimations have imposed a cointegrating relationship between exports, real exchange rate, and foreign demand. While the data do not reject this assumption, the power of the unit root and cointegration tests is limited. Furthermore, there is also a theoretical motivation to check the robustness without imposing the cointegrating relationship. In fact, the debate on the time series properties of the REER is not yet settled (see O'Connell, 1998, and Higgins and Zakrajšek, 1999). Accordingly, we estimate Eqs. (4) and (5) in differences obtaining similar results.

[16] One group contains 22 countries, and the other one the remaining 34 countries, in order to have the same size as the OECD versus nonOECD classification.

estimate the export equation specified in (4) for each of the 56 countries in the sample, and compute the proportion of countries for which the F-test that all coefficients are equal to zero is bigger than the analogous test in the equation that uses the OECD vs. nonOECD splitting. We have repeated this exercise 1000 times and we have found that only 13.5% of the cases using random splitting perform better, i.e. have a lower P-value associated with the F-test, than the OECD vs. the nonOECD sample. Additionally, we calculate the test for $\beta_j = \gamma_j$ for each of the simulations with random splitting. We find that the tests reject the null hypothesis at the 10% confidence level in 46% of the cases.

This simulation exercise provides two conclusions. Firstly, from a theoretical point of view, it shows that the elasticities of substitution are generally asymmetric across countries. From a practical point of view, we show that there is little to be gained by using two exchange rates randomly and that the partition of the OECD versus nonOECD countries can improve considerably the fit of the export demand equation for most countries.

4. Conclusions

The real effective exchange rate of a country is an aggregation of several bilateral real exchange rates with respect to other countries. The aggregation is usually done under the assumption of constant elasticity of substitution between products coming from different countries. We have investigated the validity of this assumption, by estimating manufactures export equations.

Using both individual equation and panel approaches, we have selected 47 countries for which the hypothesis of no cointegration cannot be rejected. Based on this finding, we have performed several tests to check whether exports react similarly to change in the exchange rates vis-à-vis OECD and nonOECD countries.

The results show that the specifications that do not impose the restriction of constant elasticity of substitution perform better than the specifications that do impose such a restriction. The hypothesis of constant elasticity of substitution is rejected and the export equations that contain two REERs perform on average considerably better than the traditional ones. These results are robust to several estimation techniques, suggesting that it may be useful to use disaggregated REERs when trying to reach conclusions on competitiveness.

Our results also have important implications for the imperfect competition trade literature. This literature is based on the product variety model with constant elasticity of substitution (Helpman and Krugman, 1985). However, we have shown that the data reject the hypothesis of constant elasticity of substitution and that there is a pattern for the elasticity of substitution between OECD and nonOECD countries. We do not know whether the main results of models based on product variety hold when taking into account our results, and we think that examination of this question should be undertaken.

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Appendix A. Data sources and description

The weights in the REER calculations are export weights for manufacturing derived in Zanello and Desruelle (1997). The methodology is explained in their paper, and is the same methodology used for the IMF INS trade weights. All other data are from the World Development Indicators of the World Bank (WDI). WDI provides data for manufacturing exports. The export price index is used as the price index for all REER calculations.

The 22 OECD economies are: Australia, Austria, Belgium–Luxembourg, Canada, Denmark, Finland, France, Germany, Federal Republic, Greece, Iceland, Ireland, Italy, Japan, Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, United Kingdom, United States. The 34 nonOECD economies in the sample are: Algeria, Argentina, Brazil, Chile, China, Colombia, Costa Rica, Ecuador, Egypt, El Salvador, Guatemala, Honduras, Hong Kong POC, Hungary, India, Indonesia, Israel, Jamaica, South Korea, Malaysia, Mauritius, Mexico, Morocco, Nicaragua, Pakistan, Paraguay, Peru, Philippines, Singapore, Thailand, Trinidad and Tobago, Tunisia, Turkey, Uruguay.

The summary statistics are:

Variable	Observations	Mean	S.D.	Minimum	Maximum
All economies					
Ln (manufacturing exports)	1456	21.76	2.47	15.68	26.70
REER	1456	−0.08	0.22	−1.18	1.15
REER-OECD	1456	−0.09	0.18	−1.04	0.77
REER-nonOECD	1456	0.01	0.05	−0.17	0.38
Ln (world GDP)	1456	3.68	0.96	1.03	6.35

OECD economies					
Ln (manufacturing exports)	572	23.59	1.93	16.64	26.69
REER	572	−0.01	0.11	−0.41	0.38
REER-OECD	572	−0.02	0.10	−0.37	0.32
REER-nonOECD	572	0.01	0.02	−0.04	0.09
Ln (wgdp)	572	4.31	0.66	2.62	6.35
NonOECD economies					
Ln (manufacturing exports)	884	20.57	2.01	15.68	25.57
REER	884	−0.12	0.26	−1.18	1.15
REER-OECD	884	−0.13	0.21	−1.04	0.77
REER-nonOECD	884	0.01	0.06	−0.17	0.38
Ln (wgdp)	884	3.28	0.91	1.03	5.61

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