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THE CHALLENGING ESTIMATION OF TRADE ELASTICITIES:

TACKLING THE INCONCLUSIVE EUROZONE EVIDENCE

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Abstract

This paper sheds light on serious methodological difficulties of employing the empiric export equation in order to derive long-run trade elasticities. The unreliable estimated price coefficient (Kaldor Paradox) and the potential presence of cointegration are identified as the most relevant points. It can be shown that difficulties are in part due to methodological issues. New empirical evidence, encompassing eleven Euro area countries and the timespan 1995–2019, has been obtained from different cointegration techniques. In seven out of eleven cases a robust long-run relationship can be detected and price elasticity was consistently found being significant and negative.

Keywords: International trade, Competitiveness, Kaldor Paradox, Export equation, ARDL

JEL codes: F14, C13

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1. INTRODUCTION

Differences in international competitiveness are considered an important determinant for trade imbalances. In modern trade economics price competitiveness is distinguished from non-price competitiveness. Both are understood to affect the trade performance of single countries when compared to the rest of the world. In this context, empirical evidence is commonly obtained from the estimation of the standard equation of export demand.[1] According to this standard income-and-price-elasticity framework, in its most basic version, exports are explained by the dynamics of foreign demand and of relative prices or, in many cases, of relative costs. Unfortunately, the utilisation of this procedure does not seem to produce consistent results throughout the empiric literature. In particular, the role of prices and costs[2] is discussed extensively. The interpretation of an empirical analysis by Cambridge economist Nicholas Kaldor (1978) is often seen as the starting point of a long-lasting price competitiveness debate. Since Kaldor couldn't detect the expected negative impact of relative prices on export market shares, he concluded that prices are a result, rather than a determinant, of the trade process. Several studies (e.g. Storm & Naastepad (2015) for Germany, Breuer & Klose (2015) for Italy) confirmed this unexpected finding of insignificant or even positive cost coefficients by using a version of the empirical export equation. However, having used a similar research design, comparable studies (e.g. Horn & Watt (2017) for Germany, Paternesi-Meloni (2018) for Italy) were able to detect negative and significant cost coefficients. In general, a common country pattern regarding the trade elasticities is hard to obtain. Thus, various scholars have focused on specification issues of the underlying empirical approaches testing the validity of the so-called Kaldor paradox. Since cointegration analysis appears to be applicable only with certain limitations (Bairam (1993), Bagnai (2010)), a consensus for employing the export equation in its first difference form was established. However, Boggio & Barbieri (2017) highlight the differences in outcomes between employing the level and the first difference of relative prices and costs. According to the authors, the level version is the preferable choice in empirical applications. The relationship employing the cost level, in contrast to the use of the first-difference, is significant, robust and the coefficient shows the expected negative sign. Building on related literature, two central difficulties can be identified, which significantly affect the estimation outcome: the first-difference or levels specification issue and the problematic estimation of long-term coefficients.

[1] An important strand of literature utilising the export equation exists around the estimation and the existence of a balance-of-payments constraint for national growth. The original concept by Thirlwall (1979) builds on the standard export equation and focuses on the income elasticity, whereas the role of prices is neglected for various reasons.

[2] The expressions of relative prices and costs are used as synonyms and refer to the usually employed real effective exchange rate (REER, the nominal exchange rate corrected by nominal unit labour costs). See section 3 for theoretical substantiation.

This paper approaches these issues in a more systematic way and aims at giving conclusive evidence in order to answer pending questions. More precisely, the aim is to clarify whether standard cointegration techniques employing the basic export equation can yield robust long-run evidence and whether the impact of costs on exports is negative and significant. This will be done by a systematic replication of established approaches in empirical international economics.

In section two I explain the export demand equation - the income-and-price elasticity framework - and analyse its standard interpretation from a theoretical point of view. The necessity for employing long-run information (using variables in levels instead of first-differences) is substantiated. As a next step, some recent and inconclusive empirical literature on the Euro area trade imbalances as an emblematic case is presented. Section three takes a closer look at technical difficulties coming along with the stationarity of relative prices and it discusses implications for further cointegration analysis. Building on the insights gathered, new empirical evidence is presented in section four. The export equation is tested for eleven member countries of the Euro area in the period from 1996Q2 to 2019Q2. Utilising different cointegration techniques (in particular, the estimation of an ARDL model) helps assessing whether the export equation produces robust and reliable long-run results or spurious evidence.

2. THE EMPIRICAL EXPORT EQUATION

2.1 The standard export equation and the determinants of exports

Trade flows are commonly treated as being determined by the size of the world market and by domestic prices vis-à-vis the prices of foreign competitors. This causality is assumed to hold true both in a static and a dynamic sense. Respective empirical equations have become an integral part of economics:

> *"Trade equations are one of the older, and more rewarding, parts of empirical economics. Numerous standard trade equations [...] have been estimated over the last 20 years or so with notable empirical success, so much that they are now an accepted part of most policy and applied academic work in international economics."* (Bayoumi, 1999, p.3)

Within the empirical standard export equation (which is primarily an import equation, explaining the foreign importers buying behaviour), the dynamics of exports x_t result as a function of foreign income y_t and of relative prices p_t:

$$x_t = \alpha + \beta_1 y_t + \beta_2 p_t + e_t. \tag{1}$$

According to the examples of this standard specification[3], some measure of real export flows is used as dependent variable. Either foreign real income or real expenditure is used to proxy foreign demand. The choice of the adequate relative price proxy is a more difficult issue, since the price of domestic production and of substitutes produced abroad should be put in relation to each other. Employing a version of the real effective exchange rate (REER) represents the standard for this purpose. That is the nominal exchange rate adjusted by a weighted scheme of relative export prices, wholesale prices, consumer prices, the implicit GDP deflator or unit labour costs (ULC).

Once this relationship is empirically tested, the estimated coefficients are understood as representations of the trade elasticities. The export's elasticity with respect to foreign income $\beta_1$is expected to be positive. However, that of prices $\beta_2$is assumed to be negative due to a falling demand curve in competitive markets[4]. When interpreting the results, it should be taken into consideration that the estimated elasticities not only reflect an identity, but rather a behavioural relation, with an economic mechanism behind it (McCombie, 1997). Hence, the export equation might explain the behaviour and the determinants of the buying decisions of final consumers as well as of suppliers of an importing foreign country. At first sight, econometric handling and economic interpretation appears to be straightforward, but it becomes more complex when inspected more closely. Indeed, the history of the export equation's application is full of debates regarding its reliability:

> *"The success of the standard trade model does not mean, however, that there are no controversies. Rather, such disagreements have centred more upon the size and importance of specific parameters and variables than on the underlying empirical approach. For example, differing views about the size of the long-run relative price elasticities has generated extensive discussions on the effectiveness of the exchange rate in altering nominal trade balances in the long-run..."* (Bayoumi, 1999, p.3)

An example of such a controversy is the strand of empirical literature trying to explain the different trade dynamics of the Euro area's member countries mentioned above. The interpretation of trade elasticities is often controversial and the results appear to be highly sensitive to the specification (i.e. levels or first-differences) of the export function.

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[3] See for specification issues chapter 20 of the Handbook of International Economics (Goldstein & Khan, 1985) and for notes and history of the export equation Sawyer & Sprinkle (1997).

[4] This article focusses on the determinants of the trade of manufactured goods which represent the bulk of trade. Since competition in the market of manufactured goods is imperfect, a falling demand curve can be assumed.

2.2 Trade elasticities in recent literature

A particular empirical issue derives from the different export dynamics of Germany, France and Italy, as biggest manufacturing countries in the Euro area. The scientific debate on the driving forces of the German export success and the relative weakness of Italy and France are emblematic cases regarding the difficult handling of the export equation. While the huge German trade surplus represents the debate`s starting point, the divergence among Euro member countries in terms of real export of goods and export market shares is the more interesting fact. Over the last two decades, Germany has experienced a higher export growth path (Figure 1) compared to other European industries, though before, single country dynamics had been similar up to the mid-90s.

[FIGURE 1 HERE]

In the last two decades France and Italy have lost 50% and 30% of their export world market share respectively, while Germany's share has fallen by only 15%. Ultimately, there are some signs of stabilisation, although the accumulated divergence persists. The export dynamics, expressed in growth rates (Table 1), confirm this pattern. The German export grew much faster, particularly in the period from 2001 to 2010. Since 2011 the performance of France and Italy has been much closer to the German benchmark than previously.

[TABLE 1 HERE]

In search of the drivers of the dynamics described above, Breuer & Klose (2015) represent an often cited starting point for recently estimated long-run trade elasticities (timespan 1995Q1 to 2012Q4) of the Euro area countries. The estimated elasticity of German exports with respect to foreign demand is 1.75, and -0.82 with respect to the level of the ULC deflated real effective exchange rate. For Italy, the estimation yields an income effect of 0.9 and an insignificant price effect, yet with the expected sign. France is subject to a higher price elasticity of -1.73.

When comparing estimates for the German exports only, Horn & Watt (2017) estimated an income elasticity of 1.1 and a price elasticity (export deflator based REER) of -0.5. Baccaro & Benassi (2017) find significant short-run price coefficients for German manufacturing exports. According to their rolling regression model the price sensitiveness has even increased in the period from 1971 to 2014. The average coefficients from the first differences equation range from -0.4 for ULC based REER to -0.8 for export prices in relation to import prices. The general outcome of higher income coefficients and exports being less sensitive to prices represents an empirical regularity. The majority of studies on the German case find significant price or cost elasticities with the expected sign.[5] The research of Storm and Naastepad (2015), however, represents an often-mentioned 'against the current' example in the ongoing

[5] For a literature survey on the German trade elasticities, see Heinze (2018) and on the specific topic of German price elasticities, see Baccaro and Benassi (2017).

debate. Using growth rates of income and relative ULC, they detect an insignificant coefficient for relative costs, even with a positive sign. Yet, the estimated effect size of 2.79 for the foreign income is an outstanding result. The outcome is interpreted as a consequence of Germany's extraordinary strength in terms of its corporative industrial framework and technological level, which renders export demand for its product range insensitive to changes in the industrial costs and price structure.

Due to the nature of the commodities produced, Italy's industry is often seen as being exposed to a more intense price competition than Germany (European Commission; 2010). However, as mentioned above, Breuer & Klose couldn't find a significant price effect for Italy. Defying this result, the analysis of Paternesi Meloni (2018) provides empirical evidence for high price and cost sensitiveness of Italian exports.[6] Using level variables, the author found a remarkable long-run importance of price and cost competitiveness (with a magnitude of around -0.8 for the ULC deflated REER with respect to real exports) for Italian total exports from 1994Q1 to 2014Q1. The estimated price coefficient for various price specifications and geographic samples of trade partners often shows an even greater magnitude than the income coefficient. This holds true in particular for the export of manufacturing goods. However, in the case of the bilateral export performance a higher cost sensitiveness of Germany (-0.8) in relation to Italy (-0.3) was detected by Baccaro and Tober (2017). The sample ranges from 1999 to 2014 and uses the first-difference export equation. Giordano and Zollino (2016) employed a first-difference version in the search for determinants of the trade performance of the most important Euro area countries, too. The reported difference in unit cost sensitiveness among Germany, Italy and France for the period 1993-2012 is very small. The coefficients range from -0.3 to -0.4. Although the relevance of the price effect is confirmed in most applications of the export equation, exceptions still keep the debate on the price paradox alive.

Although the cited research is based on a version of the export equation and focuses on similar periods, it appears difficult to derive conclusive evidence. As described above, the price elasticity of German exports can be found to be higher, as well as similar, in comparison to that of Italy and France, or even to be statistically insignificant. Apparently there is a high sensitiveness of the results to the individual specification and the comparability of the cited work appears to be very low. A general country pattern of the results is hard to reproduce. This raises the need for a more systematic approach, taking the known difficulties of the export equation's econometric handling and some theoretical issues as a potential source of error more seriously into account.

[6] European Commission (2010), through a cross-country analysis found significant and negative price elasticity for Italy , too.

2.3 Interpretative remarks

The question whether the trade performance depends more on quality or on prices must be answered by utilising the appropriate statistical methods. First of all, it seems useful to step back from empirics and to pose the question: What do the estimated coefficients tell us on a theoretical level? According to Romero & McCombie (2018), the determinants of the elasticity are not fully understood. Income elasticity differences are often explained by different national compositions of traded goods (McGregor & Swales, 1985) and by different productive structures (Thirlwall & Dixon, 1979). High quality commodities are assumed to be only hardly substitutable by competing goods, and their higher apparent income elasticity in combination with a low price elasticity might tell that they, therefore, do not compete on prices (e.g. Krugman, 1989). This might explain differences in income elasticities since countries specialise in the production of different types of commodities. Irrespective of how the coefficient is seen precisely, the interpretations mentioned suggest that the income elasticity catches aspects not connected to the relative price: non-price competitiveness. In the words of the European Commission non-price competitiveness is "*viewed as the sum of all factors other than prices and cost that impact on trade performance*" (European Commission, 2009, p.21). This includes product quality, design, product differentiation, liability, practicality, renovation, efficiency of sales network and all kinds of services connected to the product.[7] A more macroeconomic interpretation is, for instance, that of firm size, taxation, and access to finance (Altomonte & Osbat, 2013). Commonly used macroeconomic metrics are, i.e., indices of complexity or total factor productivity. Hence, the apparent income elasticity reflects aspects of the preference structure of the importing country, but mainly aspects of the productive structure of the exporting country.

The price elasticity magnitude, however, is assumed to be a result of the products substitutability and, therefore, reflects product properties, too. According to this substitutability interpretation, a country specialised in the production of qualitatively outstanding superior goods should experience a lower price elasticity of demand for its exports. As foreign competitors catch up on the technological level, price elasticity for domestically produced goods will rise. If, however, the price effect is low or even insignificant, the standard interpretation is as follows: the offered products do not compete on prices, because of their particular properties and consequently their low substitutability. On the contrary, a high price elasticity suggests that the product properties and the product range of an economy are highly substitutable by foreign goods and, therefore, the products have to be sold at a competitive price.[8]

[7] According to the European Commission (2017), product quality can be considered the most important feature of non-price competition.

[8] Krugman's (1989, p. 1031) interpretation is emblematic of the substitutability view.

In general, the determinants of the trade elasticities are assumed to be rather exogenous and to be a result of given fundamentals.[9] The effects of price changes reflected in the price elasticity estimates are often treated as being independent from the mentioned determinants of the income effect. Low or insignificant price elasticity coefficients often lead to the view that goods were sold only because of their non-price characteristics.[10]

Apart from this intuitive interpretation, one important logical link regarding the importance of the price level should be kept in mind. Within the standard model of a competitive market, the price level never loses its significance. Unless the demand curve is vertical, the magnitude of both income and price elasticity depends always on the current price level. In this sense caution is required when interpreting the empiric results. Even for the most outstanding superior commodity, the price still counts. A high utility of a good justifies a higher price only up to a certain level. An even higher price would foster the sale of competitor goods, although having much less utility. From a customer's perspective this is not only due to budgetary reasons, but also due to the important price-to-quality ratio. This implies that both effects, resulting from observable past changes of prices or foreign demand, can never be disconnected from the price level. Therefore, an appropriate empirical investigation shouldn't neglect information of the prevailing relative price level.

3. THE UNRELIABLE PRICE COEFFICIENT

3.1 Historical remarks

Since the famous work of Houthakker and Magee (1969) on trade elasticities, the inconsistency of the estimated price elasticity coefficients has been a central issue. Their estimations yielded insignificant or in some cases even positive coefficients – a result which provoked questions. Morgan (1970) replied to this study, arguing that already controlling for serial correlation renders results more plausible and the price effect turns negative and significant in many cases. He concluded that "*the fact that the price elasticities are, to say the least, unreliable throws doubt on the validity of the whole exercise*" (Ibid, p.304) and that this should be a reason "*for treating the results of almost any econometric analysis of international trade with great caution*" (Ibid, p.305). His criticism was the first to point at the reliability and international comparability of the price measures available at the time. Taking aggregated values and, thus, taking many prices not important for the actual manufacturing trade into

[9] Exceptions are, for instance, the advancements of the balance-of-payments-constrained-growth model with endogenous elasticities. For further insights see Missio et al. (2015).

[10] Kaldor (1978) interprets his findings as prices being endogenous and trade performance depending on non-price "*factors non susceptible to measurement*" (Ibid, p. 104).

account, would bias their estimated coefficient downward, as manufacturing prices show only small changes and low variance through time. Unfortunately, there are more difficulties in this context. Hence, this section will subsequently focus on three important issues connected to the specification of the price variable and its econometric consequences.

3.2 First differences or levels

Surveying the empirical literature further confirms the guess that the price coefficient is the most problematic and most contended component of the export equation. In Post-Keynesian economics, the problem regarding its sign and significance attained a lot of attention. The income effect, on the contrary, never raised similar doubts. The debate's starting point was an empirical investigation on the determinants of changes in international export market shares, done by Nicholas Kaldor (1978). Relative prices and market shares of most countries under investigation increased simultaneously. The conventional wisdom about the role of prices had been questioned and the discussion about the so-called Kaldor-Paradox was born. Since then, a whole strand of literature has emerged debating the validity of the paradox. Romero & McCombie (2018) sum up the debate, stating that the majority of studies employing the standard export equation is able to find positive and significant income elasticities. In contrast, price elasticities are not significant sometimes. These statements are a response to an earlier debate, in which Razmi (2015) argued that the cost level (in this case the REER level) possesses a notably higher statistical significance explaining trade and growth performance than a periodical REER change, that is usually employed. The insignificance of prices is, therefore, purely a methodological issue. Razmi proposes to use both price level and price change in the equation. Boggio & Barbieri (2017) present further support for the relevance of the cost effect. Regressing the export growth rates on the unit cost level on a cross-country basis yields evidence in favour of the cost competitiveness hypothesis. Although the authors did a market share analysis, results can be compared with some limitation. While the unit cost level is found to be significant in explaining different export performances, its first difference is not.

Besides these purely empirical arguments in favour of the use of the level variables another interpretative reason is intuitive as well. Small but persistent growth rates of relative prices can lead to a significant gap in price or cost levels among countries. Let's assume a country having a cost level 10% above that of its competitor countries producing close substitutes. The accumulated gap may have a stronger weight on the current buying decision of potential importers than a small periodical reversal of this gap. Estimating growth rate coefficients only will not deliver a comprehensive picture. The relative price level accounts for the sum of the past periodical price movements and gives a more comprehensive picture concerning the buying criteria of potential buyers. The "short-term" first difference specification therefore doesn't take this crucial information into consideration as the price level variable is omitted. Using certain lag structures won't make up for this loss of information, since the accumulation of the gap probably does not follow a constant time pattern. Unfortunately, the econometric handling of the export equation in levels form is not as straightforward as required. According to Bairam (1993), it appears difficult to detect a cointegrating relationship among the level variables of the export equation. The "*variables in levels form are not cointegrated*" and the "*elasticity coefficients do not have any desirable properties*". He concluded that "*statistical significance of the estimated elasticities can only be tested from the dynamic* [first differences; note from the author] *specifications*" (Ibid, p.740).

In accordance with the estimation results, the author concludes that the magnitude and the significance of the estimated income effect do not change much when switching from a static to a dynamic specification. Since Bairam is interested in the size of the income effect only, which, again, seems insensitive to the specification, the abandonment of the level equation shouldn't pose a problem for the relevance of this exercise. Possible consequences regarding the price coefficient weren't taken into consideration any further since the balance-of-payments-constrained growth literature, from where Bairam's work originates, assumes prices to have at most a very limited impact. Bagnai (2010), however, questions the preference to the first differences equation because it may contain a loss of long-run information. Switching simply from the long-term to a sort of short-term equation, without keeping the long-run information, might render the estimates biased and inconsistent. Furthermore, Bagnai argues that the differencing of variables used in a spurious level regression reduces the goodness of the fit and generates spurious significance of the effects, too. According to Bagnai, the commonly undertaken pre-differencing is not justified.

3.3 Choice of price measure

Another potential source of error is represented by the specification of the variables, since a certain margin for choosing a reasonable proxy exists. As regards the price variable, the choice should be well-considered. The most common approaches[11] are based on terms of trade, i.e. export prices, or on relative cost structures, i.e. nominal unit labour costs, both measured in a common currency. Terms of trade denote the relation of export prices to import prices. However, it can be questioned, whether this relation, really represents the basis of decision-making for potential importers. Thus, the use of terms-of-trade is subject to criticism:

> *"The role of terms-of-trade in these equations is to capture the effect that export price have on export demand. When choosing to consume a country's exports, foreigners compare between the price of goods from that country compared to the prices of the same goods from competing exporting countries. These prices are not the same as the import prices faced by the domestic country, yet these prices are used in conventional terms-of-trade. Similarly, domestic consumers choose between purchasing*

[11] See Table 2 in Neumann (2019) for a survey.

imports at those prices and domestic substitutes at domestic prices." (Duvall-Pelham, 2019, p.5)

It should be added that the export price indices only account for the price of the commodity sold, not for the price of the *relevant* competitor good not being sold, which, again, renders the measure less appropriate. The difficulty of choosing data that matches the corresponding requirements raises the need for a second-best solution. The volumes of nominal unit labour costs appear to be more reliable in this context. Unit costs is a widely used and acknowledged measure, but truly subject to some serious concerns, too. So why use unit costs as a measure to obtain information on the price sensitivity of exports? The cost per unit of output relates labour costs to the real output. An advantage of the cost approach is that, by employing real output, it accounts for productivity. In this sense, to some degree, it controls for industrial efficiency, which represents important supply side information.

When focusing on the analysis of manufacturing trade flows, which, again, represents the bulk of trade, the market is characterised by monopolistic competition.[12] This is a situation of increasing returns to scale, characterising the supply side, where firms offer their products via mark-up on costs. In this sense, the cost per unit of output is seen as the most important basis of pricing behaviour and therefore a good proxy for the price level of domestic production.[13] Thus, it represents the basis of pricing for domestic and foreign producers, regardless whether their products have been exported or not. Furthermore, the indicator catches the whole cost structure of the national production system, which is not subject to short-term price adjustments due to reasons of competitive pricing policy. This characteristic highlights the need for employing the cost level, too.

3.4 Price stationarity and cointegration techniques

In a market economy under free trade, the price series put in relation to each other cannot diverge indefinitely due to economic reasons. The underlying question is, whether purchasing power parity (PPP) applies, since its presence implies stationarity of relative prices. As most of the recent literature suggests, the time series is mean reverting, at least in the long run (see Blecker, 2016). Prices, in a common currency, may diverge up to a certain extent for a certain period of time, but the divergence does not last. This should hold true even more in the case of a common market within a currency union without relevant trade barriers. To demonstrate this point more clearly, Figure 2 shows the real effective exchange rates of Germany, France and Italy.

[12] For instance, in the trade and growth models of Kaldor (1970) and Krugman (1979), world markets are assumed to be imperfectly competitive. The production of tradable goods would be characterised by economies of scale, which render competition monopolistic.

[13] For instance, the Deutsche Bundesbank (2016) reports a higher reliability of the GDP deflator and ULC in comparison to consumer price indices explaining trade flows.

[FIGURE 2 HERE]

Despite periods characterized by REER divergence are significant, the extent of divergence appears to be limited in the short sample as well as in the longer one. Looking closer at the period following the fixation of exchange rates and the onset of the euro in 1998/1999, relative prices have diverged steadily for 10 years.[14] Afterwards the time series partially revert. It can be observed that the single time series aren't characterised by any clear trends.

The fact that the time series of relative prices is not trending has important implications for the econometric treatment of the export equation. The most popular testing methods for long-run relationships among level variables require the involved time series to be trending, namely to be integrated of order one. For instance the Engle Granger test for cointegration or the Johansen approach[15] demand regressors to be purely I(1). While in the majority of cases the time series of exports and foreign income are clearly I(1), it appears not feasible to assume relative prices to be purely trending, either. Only the choice of rather exceptionary time spans where prices diverge, or other particular constructions of the variable, may yield a price time series which exhibits a trending behaviour.[16] It is more likely that relative prices are fractionally integrated; that is neither stationary nor non-stationary.[17] Such time series can be characterised as a long memory and slow mean-reverting process (see Habermeier & Mesquita, 1999).

In order to run a cointegration analysis the order of integration of the variables involved needs to be established. However, unit root tests, like the often used Augmented Dickey Fuller methodology (ADF), suffer from low power, when the time series is a close to, but not exactly a unit root process. Elliot (1998), for instance, reports that the ADF test is only hardly capable to distinguish an I(0.5) process from a unit root process. It is likely that a series is reported to be I(1) when in fact it is not. In order to demonstrate this point more clearly, two different unit root tests are carried out on the ULC deflated real effective exchange rate. First, an ordinary Augmented Dickey Fuller test is employed, to test whether the variable is non-stationary or not. As described above, the Dickey Fuller methodology has low power in the case of partially integrated time series. Therefore, a KPSS test (Kwiatkowski, Phillips, Schmidt & Shin, 1992) is carried out to give complementary information on relative prices and whether they are stationary or not. The two methods differ in terms of their respective null hypothesis.

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[14] This period rather appears as an exception. In the words of the European Commission: "*The current level of divergence in competitiveness does not appear extremely large by historical standards but its persistence does*" (European Commission, 2009, p.19).

[15] This holds true in the bivariate case.

[16] For instance, Bagnai (2010) utilised the ratio of the GDP deflator and the import deflator, which yields an I(1) time series.

[17] This characterisation of time series behavior was established by Granger & Joyeux (1980) and Hosking (1981)

While the ADF method tests against a null hypothesis of non-stationarity I(1), the KPSS test has a H0 of an I(0) stationary process. The sample contains the annual (1960-2018) and quarterly (1995Q1-2019Q2) REER (ULC deflated) data of eleven founder states of the Euro area.[18] The annual sample is added to the analysis for the sake of comparison only and is assumed more likely to be stationary. The quarterly sample is the same as the one that will be used later in the empirical analysis. Table 2 presents the rejection share of the respective null hypothesis at 10% confidence level. Analysing the annual single country results, both tests suggest three cases out of eleven to be clearly stationary. Other five are reported to be non-stationary. In the other three cases the tests yield contradictory results, namely ADF non-stationarity and KPSS stationarity. Hence, the long-term sample suggests that clear non-stationarity for relative prices cannot be assumed. However, the results of the short quarterly sample are even more confusing. The tests report two cases to be clearly non-stationary and one to be clearly stationary. In eight cases, the order of integration cannot be detected unambiguously. The results tend to confirm the conjecture that the relative price time series is fractionally integrated.

[TABLE 2 HERE]

Detecting time series spuriously as I(1) can generate spurious evidence of cointegration methods, too. The Engle Granger approach, as the most basic method testing for a long-run relationship, relies strongly on the assumption of I(1) variables. This holds true for a bivariate Johansen cointegration test as well. Considering the contradictory results obtained above, there is no surprise that researchers struggle to find a long-run relationship of the export equation variables. Since the assumption of I(1) processes is not satisfied, the problem of serial correlation of the error term is likely to occur using the cointegrating equations mentioned. Furthermore, spurious cointegration can be another possible pitfall (Gonzalo & Lee, 2000). Thus, an appropriate empirical exercise should take the points mentioned into account.

[18] For theoretical reason deterministic trends of the time series are not assumed. Nonetheless, where statistically significant, deterministic trends were added to the single unit root test (mainly in the quarterly sample). The general outcome was roughly unaffected.

4. INFERENTIAL EVIDENCE

4.1 Specification

Regarding the precise model specification, the work of Breuer & Klose (2015) serves as point of reference. The export variable X_t^i is proxied by real export of goods (data source: EUROSTAT quarterly national accounts), accounted in purchasing power parity US-Dollar. The countries considered ($i = 1, \ldots, 11$) are Austria, Belgium, Germany, Finland, France, Ireland, Italy, Luxembourg, the Netherlands, Portugal and Spain which all belong to the currency union of the euro.[19] Foreign demand FD_t^i is measured in purchasing power parity dollar (data source: OECD quarterly national accounts) of real gross domestic product of each 44 partner countries. The sample contains the countries of Australia, Argentina, Austria, Belgium, Brazil, Bulgaria, Canada, Chile, Costa Rica, Czech Republic, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Iceland, India, Indonesia, Ireland, Israel, Italy, Japan, Korea, Latvia, Lithuania, Luxembourg, Mexico, the Netherlands, New Zealand, Norway, Poland, Portugal, Romania, Russia, Slovak Republic, Spain, South Africa, Sweden, Switzerland, Turkey, the United Kingdom, the United States and ranges from 1996Q2 to 2019Q2. The sample ought to catch the potential demand a single exporter country faces. Since the subtraction of the single-country demand from world demand has a rather negligible effect and the data is not weighted, the potential foreign demand is similar for each country.

As described in section three, the real effective exchange rate deflated by nominal unit labour costs (data source: European Commission Competitiveness Database) is chosen to proxy relative prices ($REER_t^i$). This includes each domestic ULC in relation to 36 partner country average ULC (double export weights), both accounted in a common currency. The sample consists of the countries of the EU28 plus Australia, Canada, the United States, Japan, Norway, New Zealand, Mexico, Switzerland, and Turkey. In difference to the foreign demand series, the sample should primarily account for the main industrial competitor nations of each country under consideration. Since there is a need for interpreting the estimated coefficients as trade elasticities, all data of the sample is expressed in logs.

[19] The geographic sample choice was made according to reasons of data availability.

4.2 Methodology

The methodical response to the potential problems described in section 3 is to use different methods in conjunction. This setting should help avoiding a pitfall due to the choice of inappropriate single techniques, which are not capable of producing reliable results. As a first step, the necessary unit root test on the variables involved will be performed. A particular test procedure has already been chosen for the case of relative prices and the time series characteristics discussed in the section 3.3. An Augmented Dickey Fuller test is carried out on the time series of exports and foreign income. The test only rejects the null hypothesis of non-stationarity of the real export variable in the cases of Italy, the Netherlands, Austria and Finland. This is an unexpected result which may lead to further difficulties. However, the variable of foreign demand is found to be I(1) in all eleven cases.[20] Since the method of this paper is to replicate the standard approach, single country equations will be preferred over a panel data approach. This allows for different elasticities reflecting the relevant differences in productive structures.

Surveying the empirical literature utilising the export equation, no common additional controls have been established for a variety of reasons. For instance, many studies try to control for quality or the variety of the product range by employing proxies like total factor productivity or economic complexity. The usefulness of such additional control variables can be doubted in many cases as their use gives rise to further problems like endogeneity or multicollinearity, both in statistical and in theoretical terms.[21] The particularity of the methodological setting used in this study is to use different estimation approaches and cointegration tests in conjunction.[22] As a first step the Engle-Granger approach is carried out based on an OLS estimation of the long-run coefficients. This method is expected to be particularly problematic since the time series involved do not meet the method's assumptions. Two additional techniques, that are expected to suit better, are chosen. First, a maximum likelihood estimation of the coefficients, within a vector error correction model, is applied. Secondly, an autoregressive distributed lag model is employed since it represents a remedy in case of spurious regression results that are due to non-trending time series of the regressors.[23] This strategy is a reaction to the stationarity issue described above and might help obtaining the long-run evidence required in the context of international trade.

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[20] Detailed results are reported in appendix tables A1 and A2.

[21] For instance, Paternesi Meloni (2018) mentions the endogeneity issue using nominal unit labour costs and total factor productivity, which both account for national productivity dynamics.

[22] Gonzalo and Lee (1998) recommend using various techniques in order to avoid empirical pitfalls in the case of fractional integration of the time series.

[23] See Ghouse et. al (2018) for further insights on the ARDL method in the context of spurious regression.

Furthermore, it sheds light on different time series characteristics and their potential impact on the econometric outcome.

4.3 The general setting

For purposes of comparison the first task is simply to check for the short-run relationship with differenced variables.[24] The outcome might help to assess if short-run elasticities differ from long-run elasticities. In doing so, the already discussed hypothesis of Bairam (1993) will be tested, estimating whether the export equation in first differences is equivalent to its levels form. The model specification is

$$\Delta X_t^i = \theta + \beta_1 \Delta FD_t^i + \beta_2 \Delta REER_t^i + \epsilon_t. \tag{2}$$

Since the time series of exports and foreign income are non-stationary and the pre-differencing of the variables may result in a misspecification due to a potential omission of the long-run relationship, there is a need for employing cointegration techniques. Cointegration exists when non-stationary variables are tied together by a stationary relationship in the long-run. The general empirical approach can be expressed by the subsequent version of the export equation:

$$X_t^i = \delta_0 + \delta_1 FD_t^i + \delta_2 REER_t^i + EC_t^i, \tag{3}$$

where EC denotes the error term, which represents the deviation from the assumed long-run equilibrium. Under cointegration, the whole relationship can be expressed by the following general error correction model:

$$\Delta X_t^i = \theta + \alpha EC_{t-1}^i + \sum_{n=1}^{L} \beta_1 \Delta X_{t-n}^i + \sum_{n=1}^{L} \beta_2 \Delta FD_{t-n}^i + \sum_{n=1}^{L} \beta_3 \Delta REER_{t-n}^i + \epsilon_t, \tag{4}$$

where L represents the chosen number of lags of the respective variable. This is basically equation (2) controlling for the long-run information. Since this analysis is focused on the existence of a long-run equilibrium, short-run dynamics out of the error correction model won't be taken further into consideration. The corresponding cointegrating equation is expressed by

$$EC_{t-1}^i = X_{t-1}^i - \delta_1 FD_{t-1}^i - \delta_2 REER_{t-1}^i - \delta_0, \tag{5}$$

The potential relationship will be tested using the following econometric techniques:

i. The short-run dynamics within equation (2) are estimated by means of ordinary least squares. In line with Storm & Naastepad (2015), the Prais-Winsten AR(1) estimation method was chosen. This technique is capable handling the serial correlation problems.

ii. The Engle-Granger (EG) two-step procedure (Engle & Granger, 1987) is probably the most widespread method testing the empirical export equation. The first step is to estimate equation (3) by OLS. The second step is to conduct a unit root test on the residuals of step one. This hypothesis test uses critical values of MacKinnon (2010). One of the most critical assumptions of this method is that it demands all employed variables to be non-stationary. Lag length selection is done by the minimisation of the Akaike criterion (AIC).

iii. The Johansen approach (Johansen, 1988) works differently and is expected to eliminate some shortcomings of the residual based Engle-Granger method. In principle, the critical assumption of non-stationarity of the variables is not relaxed. However, in case of testing the presence of a cointegrating relationship in a multivariate setting among more than two variables the Johansen procedure requires only two series to be I(1), whereas the other variables can differ (Hansen & Juselius, 1995). In general, the method treats all chosen variables of the export equation as endogenous in a VAR system and allows for multiple cointegrating relationships. The long-run information is part of a short-term vector error-correction model. The cointegrating vectors are estimated by the means of maximum likelihood method. Based on Johansen's trace and maximal-eigenvalue tests the rank of the estimated long-run coefficient matrix can be specified. The rank is defined as the number of cointegrating vectors respective relationships. In the case of the export equation, only one cointegrating relationship is expected. As before, AIC was used to specify the number of lags.

iv. The residual based cointegration test within an autoregressive distributed lag model (ARDL) approach by Pesaran, Shin and Smith (2001) appears to be the preferable choice for the difficult time series properties described above. In a combination of an autoregressive and a distributed lag model, an I(1) variable can be regressed on variables, whether they are I(0) or (1). The condition is that no variable is I(2) and that only one cointegrating relationship can exist. In this setting the chosen dependent variable is regressed on its own lags and contemporary and lagged independent variables via OLS. The so-called bounds test, which basically is an F test, is carried out for testing the presence of a long-run equilibrium using critical values of Kripfganz and Schneider (2018). Again, AIC was used to select the appropriate lag length.

4.4 Results

Estimating the export equation in its pre-differenced version by means of OLS (i) does not provide any surprises. The model explains roughly 60% of the dependent variable's variance. The short-run income elasticity magnitudes, where significant on a 90% confidence level, range from 2.5 to 4 and are higher than recent empirical literature has shown. Short-run price elasticity, however, varies from -0.4 to -0.7. Since the income effect is estimated to be at least five times the price effect, it seems reasonable to consider exports primarily driven by changes in foreign income and to highlight the importance of non-price competitiveness. Since no long-run information is included, some suspicion against the reliability of the results remains. If cointegration among the levels can be detected, the results of (i) can be regarded as spurious evidence.

[TABLE 3 HERE]

Testing the level equation (3) via OLS (ii) yields smaller foreign income coefficients. The magnitude varies from 1.0 to 2.1. The price effect is found to be negative and significant in ten cases and ranges from -0.2 to -1.9. Unfortunately, the evidence obtained is spurious, as an cointegration relationship is found only for three countries just on a 10% confidence level. Furthermore, in all eleven cases the reported Durbin Watson statistic reveals problems of first order autocorrelation of the residuals. The DW statistic is lower than the R^2, hence the spurious-regression-rule-of-thumb is violated, too. This does not represent a surprise considering that the necessary time-series characteristics for the estimation are not given here. Proceeding to the maximum-likelihood estimation of the long-run coefficients within a vector-error-correction model (iii) yields more promising results. At least one cointegrating vector was found in six cases, at a critical value of even 1%. Another three-country time-series (Italy, the Netherlands, Portugal) possess a cointegrating relationship on a 90% confidence level, but results appear unstable, once further robustness tests are being applied.[25] It is well known that the Johansen approach is very sensitive to changes in the lag structure. This holds true in this case, too. However, where cointegration is present, the income effect ranges from 0.8 to 1.6. The price effect is reported to be negative, but only in five out of nine cointegrated cases it is statistically significant (Table A4). The effect size lays between -1.1 and -0.2. Turning to the ARDL bounds approach (iv) almost the same cases as before are found to possess a long-run relationship. Where the null hypothesis of no cointegration of a single country equation is rejected by both the Johansen and the ARDL approach on a significance level of at least 95%, the estimated magnitudes are strikingly similar. However, no cointegration is confirmed for Finland, Ireland, Italy and the Netherlands. It is notable that, when employing the ARDL model, most price coefficients turn significant.

[25] Robustness tests were carried out altering the lag length choice, period and geographical sample of foreign demand and the price variable. Results are available upon request.

Drawing conclusions from the models (iii) and (iv), robust cointegration evidence was found for Austria, Belgium, Germany, France, Luxembourg, Portugal and Spain. The case of Finland yields mixed results. No cointegration, however, was found for the time series of Ireland, Italy and the Netherlands. At this point it should be recalled, that in the cases of Finland, Italy and the Netherlands the simple unit root test was not able to rule out the possibility of stationarity of the dependent export variable. Considering that the dependent variable is not trending in the cases mentioned, cointegrating techniques become superfluous and inappropriate. This finding suggests that the result of no cointegration in the case of single countries can be simply due to pure statistical and methodical reasons apart from the general validity of the export equation. The estimated effect magnitudes differ significantly among countries, but the pattern obtained does not fit into some of the established narratives mentioned in section two. The sensitivity of exports with respect to foreign demand is above unity in every cointegrating case. Germany shows the highest coefficient with an elasticity of 1.972. Austria, Belgium, Luxembourg and Spain show an effect size of around 1.5, however, France has the lowest value (1.0). The cost sensitiveness, however, of exports varies more. The lowest elasticities above -1.0 can be observed for Spain, Luxembourg and France. Austrian, Belgian and German exports are more price elastic with effect magnitudes below -1.0, confirming that their manufacturing export is significantly driven by favourable cost levels, too. For instance, Germany is characterised by a relatively high cost sensitiveness (-1.0) of its exports. Spain's exports, in comparison, are less elastic to the relative cost level (-0.3). This, in turn, confirms that German exports are very sensitive with respect to the REER, too.

The differences in terms of effect size, significance and cointegration among the different methods are highly significant. This finding highlights the importance of assessing the time series characteristics carefully before deriving an appropriate estimation model. Since relative prices in most cases are neither clearly I(0) nor I(1), an econometric pitfall is very likely to occur. No reliable result was produced by employing the Engle Granger two-step procedure, where the problem of spurious regression occurs in all eleven cases. The Johansen approach, however is able to find cointegration in most cases. Where cointegration was detected only on a significance level below 95%, the results suffer from a high instability regarding the coefficient magnitudes and the presence of cointegration when the specification is altered. The empiric example shows that the ARDL model is most likely being capable of producing robust results and detecting cointegration among the variables of the export equation in its simplest version. At this point it should be highlighted, that the differences between short-run and long-run coefficients are notable, too. Different income movements appear to be strongly driving trade flows in the short-run. In the long-run, however, the effect size moderates and that of cost increases significantly in most cases.

5. CONCLUSION

In this analysis, I identified three central difficulties regarding the estimation of trade elasticities: (1) the pre-differencing of the export equation, (2) the fractionally integrated time series of the real effective exchange rate and (3) inappropriate cointegration techniques. Several studies estimating the impact of foreign income and relative prices on export dynamics haven't succeeded in detecting cointegration among the variables. Thus, using pre-differenced variables is commonly considered a remedy for the cointegration pitfall. I substantiate why pre-differencing of the export equation is disputable for methodological and theoretical reasons. The resulting need for employing the equation in its levels version generates difficulties for cointegration analysis in terms of the regressor's time series properties. In case purchasing power parity is valid, the relative price time series cannot expose a clear trending behaviour. Testing the time series (by ADF and KPSS tests) of the real effective exchange rate of eleven Euro area countries confirms the theoretical conjecture. Relative unit costs can be characterised as being fractionally integrated since they trend over a certain period of time before reverting back. This characteristic violates the assumptions of commonly applied long-run estimation techniques leading to problems of autocorrelation.

By utilising three different econometric techniques, which all work differently regarding the data characteristics, additional evidence is provided. The main findings can be summarised as follows: As expected, the Engle Granger two step procedure produces spurious evidence, since a cointegrating relationship cannot be detected. By having employed the Johansen approach (based on a vector error correction model) at least one cointegrating vector was found in six out of eleven cases. The ARDL model output shows an equilibrium relationship for seven countries. In these cases, price coefficients are found to be significant and negative. The ARDL model appears as the preferable choice since it yields the most consistent and robust outcome. Assessing the results of both methods, robust cointegration is found for Austria, Belgium, Germany, France, Luxembourg, Portugal and Spain. The export elasticity to foreign income varies from 1.0 to 1.9. The price effect ranges from -0.3 to -1.5. However, the five cases that are not cointegrated possess a dependent variable not clearly trending which renders common long-run estimation technique superfluous. Furthermore, there is neither empirical nor theoretical evidence to neglect the importance of relative prices and cost structures for exports as has been argued earlier in literature.

Provided that the methodological assumptions are met and that the model is specified with diligence, in most cases the use of the differenced export equation is neither necessary nor yields robust empiric evidence. This finding confirms that the pre-differencing of variables can yield spurious results and that controlling for the long-run relationship is indispensable.

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FIGURES

Figure 1: Export of goods, world market share (left) and volume 2010Q1=100 (right)

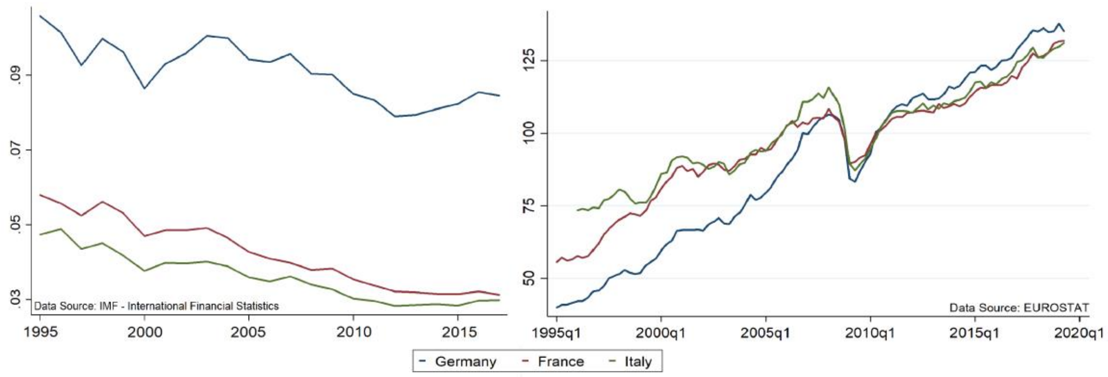
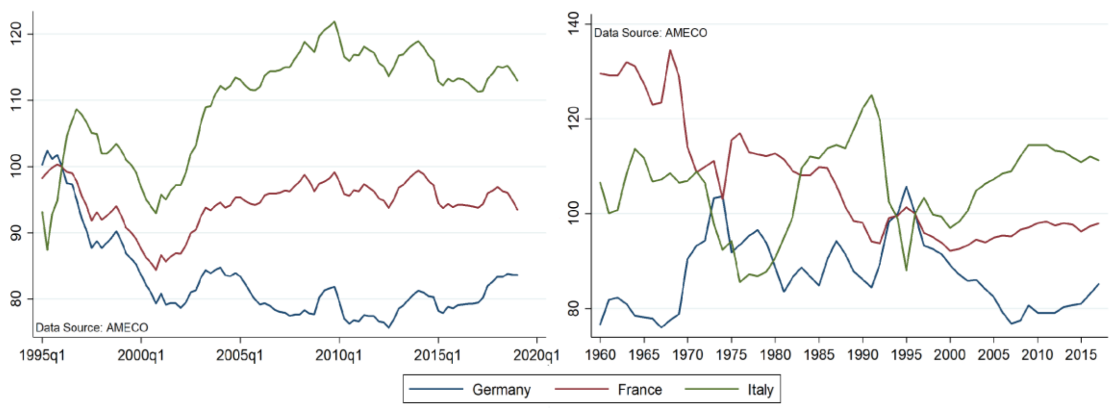


Figure 2: Real effective exchange rate, ULC deflated (double export weights)[26], quarterly data 1995q1-2019q2, 1996q1=100 (left) and annual data 1960-2018, 1996=100 (right)



[26] Double export weights are based on the competitor's share in total exports.

TABLES

Table 1: Export of goods, annual growth rates, 1996-2018 and subperiods

Period	Germany	Italy	France	EZ11
1996-2000	0.026	-0.018	0.019	0.017
2001-2005	0.129	0.080	0.092	0.105
2006-2010	0.086	0.054	0.045	0.068
2011-2015	-0.018	-0.021	-0.022	-0.014
2016-2018	0.059	0.046	0.057	0.060
1996-2018	0.056	0.026	0.036	0.046

Data source: Own calculation based on Eurostat database, EZ11: AU, BE, FI, FR, GE, IR, IT, LU, NL, PT, SP.

Table 2: Cases of H0 rejection at 10% confidence level

Method	H0	1960 - 2018	1995q1 - 2019q1
ADF	I(1)	4 / 11	6 / 11
KPSS	I(0)	5 / 11	5 / 11

More detailed results are reported in the appendix in table A1.

Table 3: Export of goods, results survey of effects size of income (ß1/δ1) and prices (ß2/δ2).

	(i) Δ		(ii) EG		(iii) VECM		(iv) ARDL	
	ß1	ß2	δ1	δ2	δ1	δ2	δ1	δ2
Austria	**3.275**	**-0.829**	**1.687**	**-1.922**	**1.552**	**-1.110**	**1.599**	**-1.490**
Belgium	**3.399**	-0.121	1.399	-0.857	**1.261**	**-0.177**	**1.463**	**-1.059**
Finland	**5.431**	-0.335	1.277	-1.433	0.777	-0.168	0.762	-9.780
France	**2.731**	**-0.431**	1.103	-1.033	**0.999**	**-0.685**	**1.019**	**-0.877**
Germany	**3.616**	**-0.595**	1.653	-1.025	**1.598**	**-1.091**	**1.972**	**-1.028**
Ireland	0.328	**-0.969**	2.106	-0.570	2.247	-0.792	1.123	-1.582
Italy	**3.522**	**-0.588**	**0.966**	**-0.391**	**0.895**	**-0.178**	1.075	-0.575
Luxembourg	**3.097**	-0.057	**2.147**	**-1.123**	**1.445**	**-0.336**	**1.696**	**-0.649**
The Netherlands	**2.570**	**-0.461**	1.623	-0.220	**1.518**	**-0.038**	1.433	-0.253
Portugal	**3.031**	0.265	1.456	-0.115	**1.541**	**-0.136**	**1.446**	**-0.517**
Spain	**3.477**	**-0.399**	1.527	-0.272	**1.519**	**-0.278**	**1.458**	**-0.309**

Bold values in case i signal significance on 10% confidence level. Bold values in the cases ii, iii, iv signal the presence of cointegration of the underlying equation and the coefficients significance, both at 10% critical value of respective test approach. Caution is needed in case of (ii) since the OLS estimator might not have standard distribution. ß1: short-run income elasticity, ß2: short-run price elasticity, δ1: long-run income elasticity, δ2: long-run price elasticity. The detailed results are presented in appendix tables A3-A6.

APPENDIX

Table A 1: Unit root tests of relative prices / Real effective exchange rate deflated by ULC

	REER_ULC37 (q)				REER_ULC37 (a)			
	ADF	l, t	KPSS	l	ADF	l, t	KPSS	l, t
Austria	-3.808**	1, t	0.303	6	**-1.734**	1	**0.705****	5
Belgium	-3.654**	1, t	0.505**	6	-3.386**	1	0.0995	5, t
Finland	-3.300*	1, t	0.450*	6	**-1.977**	2	**0.588****	5
France	-2.108	2	0.319	6	**-1.795**	2	**0.932*****	5
Germany	-3.706***	1	0.840***	6	-2.503*	1	0.182	5
Ireland	**-1.303**	1, t	**0.422***	6	-1.017	1	0.231	5
Italy	-3.623***	1	1.02***	6	-2.461*	1	0.170	5
Luxembourg	-3.371*	1, t	1.360***	6	**-0.357**	0	**0.589****	4
The Netherl.	**-2.548**	1, t	**0.636****	6	-2.988**	1	0.261	5
Portugal	-1.726	1	0.324	6	-1.604	2	0.150	5
Spain	**-1.100**	1	**0.473****	6	**-1.927**	4	**0.926*****	5

Quarterly data (q) 1995q1-2019q1, annual data (a) 1960-2018. ADF: Augmented Dickey-Fuller test with constant, H0: I(1) non stationarity. KPSS: Kwiatkowski–Phillips–Schmidt–Shin test, H0: I(0) stationarity. Lag length choice (l) according to AIC minimisation. The presence of a deterministic trend is denoted by (t). Bold value indicate consistent non-stationarity of the single country time series. Critical values: 1% ***; 5% **; 10% *.

Table A 2: Unit root test (ADF) of real export of goods, foreign demand and relative prices.

	EXPORTG		FD37		REER_ULC37	
	ADF	l, t	ADF	l, t	ADF	l, t
Austria	-3.430*	1; t	-2.603	2, t	-3.808**	1, t
Belgium	-2.809	2, t	-2.608	2, t	-3.654**	1, t
Finland	-2.897**	3	-2.608	2, t	-3.300*	1, t
France	-2.303	1, t	-2.612	2, t	-2.108	2
Germany	-2.458	1, t	-2.482	2, t	-3.706***	1
Ireland	-0.745	1	-2.596	2, t	-1.303	1, t
Italy	-3.462**	1, t	-2.609	2, t	-3.623***	1
Luxembourg	-2.608	0, t	-2.603	2, t	-3.371*	1, t
The Netherl.	-3.198*	1, t	-2.601	2, t	-2.548	1, t
Portugal	-3.124	0, t	-2.605	2, t	-1.726	1
Spain	-3.009	2, t	-2.711	2, t	-1.100	1

Quarterly data 1995q1-2019q1, ADF: Augmented Dickey-Fuller test with constant, H0: I(1) non stationarity. Lag length choice (l) according to AIC minimisation. The presence of a deterministic trend is denoted by (t). Critical values: 1% ***; 5% **; 10% *.

Table A 3: Export of goods, short-run estimates (OLS Prais-Winsten AR(1)), equation (2).

	δ1	δ2	α	DW	R2
Austria	3.275***	-0.829***	-0.009***	1.997	0.687
	(0.287)	(0.187)	(0.002)		
Belgium	3.399***	-0.121	-0.124	1.988	0.465
	(0.399)	(0.200)	(0.003)		
Finland	5.431***	-0.335	-0.027***	2.135	0.413
	(0.848)	(0.360)	(0.006)		
France	2.731***	-0.431***	-0.009***	1.982	0.584
	(0.265)	(0.129)	(0.002)		
Germany	3.616***	-0.595***	-0.013***	2.021	0.698
	(0.324)	(0.128)	(0.002)		
Ireland	0.328	-0.969***	0 .014	1.985	0.233
	(0.894)	(0.187)	(0.007)		
Italy	3.522***	-0.588***	-0.017***	2.019	0.665
	(0.302)	(0.130)	(0.002)		
Luxembourg	3.097***	-0.057	-0.009	2.011	0.111
	(1.021)	(0.547)	(0.008)		
The Netherl.	2.570***	-0.461***	-0.005***	1.974	0.604
	(0.263)	(0.150)	(0.002)		
Portugal	3.031***	0.265	-0.009**	1.984	0.317
	(0.474)	(0.277)	(0.003)		
Spain	3.477***	-0.399*	-0.011***	1.987	0.457
	(0.414)	(0.208)	(0.003)		

n=91. Prais-Winsten AR(1) regressions. δ1: Short-run income elasticity. δ2: Short-run price elasticity. Critical values: 1% ***; 5% **; 10% *. Standard errors in brackets.

Table A 4: Export of goods, long-run estimates (OLS) from model ii.

	δ 1	δ2	α	DF	DW	R2	l
Austria	1.688***	-1.923***	-0.258***	-3.932**	0.377	0.991	2
	(0.017)	(0.097)	(0.075)				
Belgium	1.399***	-0.858***	-0.244***	-2.770	0.289	0.967	2
	(0.032)	(0.161)	(0.067)				
Finland	1.278***	-1.433***	-0.099**	-1.551	0.213	0.687	1
	(0.029)	(0.334)	(0.048)				
France	1.103***	-1.033***	-0.202***	-2.385	0.142	0.943	3
	(0.031)	(0.144)	(0.045)				
Germany	1.653***	-1.026***	-0.303***	-3.250	0.364	0.990	2
	(0.022)	(0.074)	(0.076)				
Ireland	2.106***	-0.571***	-0.105***	-2.147	0.220	0.909	1
	(0.082)	(0.086)	(0.053)				
Italy	0.967***	-0.391***	-0.179***	-3.615*	0.219	0.923	2
	(0.036)	(0.095)	(0.050)				
Luxembourg	2.148***	-1.124***	-0.282***	-3.652*	0.563	0.934	1
	(0.121)	(0.206)	(0.090)				
The Netherl.	1.623***	-0.221***	-0.152***	-2.672	0.124	0.981	1
	(0.029)	(0.134)	(0.050)				
Portugal	1.457***	-0.116	-0.295***	-2.697	0.402	0.966	1
	(0.029)	(0.106)	(0.083)				
Spain	1.527***	-0.273***	-0.343***	-3.383	0.263	0.970	1
	(0.029)	(0.077)	(0.078)				

n=92-l. Engle-Granger two-step procedure for cointegration test. Null hypothesis: No long-term relationship exists. Critical values: -4,456 - 1%***; -3,836 - 5%**; -3,521 - 10%*. Long term coefficients δ1 (foreign demand) and δ2 (REER) from cointegrating equation. Adjustment coefficient α from error correction model. Model lag length (l) according to AIC minimisation. Standard errors in brackets.

Table A 5: Export of goods, long-run estimates (MLM) of VECM, model iii.

	δ1	δ2	α	Trace	l
Austria	1,552***	-1.110***	-0.225***	40.321***	2
	(0.046)	(0.255)	(0.042)		
Belgium	1.261***	-0.177	-0.253***	41.444***	2
	(0.063)	(0.311)	(0.055)		
Finland	0.777***	-0.168	-0.125***	19.697	1
	0.268	0.941	(0.043)		
France	0.999***	-0.685***	-0.214***	37.782***	3
	(0.046)	(0.202)	(0.042)		
Germany	1.598***	-1.091***	-0.302***	37.185***	2
	(0.046)	(0.152)	(0.065)		
Ireland	2.247***	-0 .792***	-0.145***	23.366	1
	(0.228)	(0.241)	(0.045)		
Italy	0.895***	-0.178	-0.169***	27.548*	2
	(0.073)	(0.185)	(0.042)		
Luxembourg	1.445***	-0.336	-0.270***	55.709***	1
	(0.226)	(0.384)	(0.060)		
The Netherl.	1.518***	-0.038	-0.146***	27.672*	1
	(0.07)	(0.319)	(0.047)		
Portugal	1.541***	-0.136	-0.158**	33.304**	1
	(0.051)	(0.103)	(0.068)		
Spain	1.519***	-0.278***	-0.231***	47.402***	1
	(0.040)	(0.103)	(0.060)		

n=92-l. Johansen trace statistic for cointegration test. Null hypothesis: No long-term relationship exists. Critical values: 35,65 - 1%***; 29,68 - 5%**; 26,79 - 10%*. Long term coefficients δ1 (foreign demand), δ2 (REER) from vector error correction model. Adjustment coefficient α obtained from export equation (Exports as dependent variable) within VECM. Model lag length choice (l) according to AIC minimisation. Standard errors in brackets.

Table A 6: Export of goods, long-run estimates (OLS) of ARDL model iv.

	δ1	δ2	α	F	t	l
Austria	1.599***	-1.490***	-0.224***	6.049**	-3.798**	3,3,4
	(0.046)	(0.269)	(0.059)			
Belgium	1.463***	-1.059*	-0.153***	4.229**	-3.421*	1,1,0
	(0.119)	(0.562)	(0.044)			
Finland	0.762	-9.780	-0.023	1.441	-0.494	3,4,3
	(2.057)	(19.904)	(0.046)			
France	1.019***	-0.877***	-0.224***	15.89***	-6.67***	4,4,0
	(0.0426)	(0.193)	(0.033)			
Germany	1.972***	-1.028***	-0.294***	9.361***	-4.204***	3,3,1
	(0.491)	(0.153)	(0.070)			
Ireland	1.123	-1.582*	-0.074	2.555	-1.45	2,0,2
	(0.831)	(0.873)	(0.051)			
Italy	1.075***	-0.575	-0.088**	1.839	-2.194	1,2,1
	(0.170)	(0.404)	(0.040)			
Luxembourg	1.696***	-0.649	-0.282***	5.271**	-3.805**	1,2,0
	(0.441)	(0.708)	(0.074)			
The Netherl.	1.433***	-0.253	-0.102**	3.468	-2.299	4,4,4
	(0.118)	(0.476)	(0.044)			
Portugal	1.446***	-0.517**	-0.296***	6.632**	-3.893**	4,4,2
	(0.056)	(0.233)	(0.076)			
Spain	1.458***	-0.309***	-0.305***	6.651**	-4.233***	2,4,2
	(0.045)	(0.109)	(0.072)			

n=92-l. Bounds procedure (F and t Test) for cointegration test. Null hypothesis: No long-term relationship exists. Critical values depend on lag structure choice: 1% ***; 5% **; 10% *. Adjustment coefficient α and long term coefficients δ1 (foreign demand), δ2 (REER) from ARDL equation. Individual lag length choice (l) according to AIC minimisation. Standard errors in brackets.