Al Collegio dei Docenti della Scuola di Dottorato

in Scienze Economiche e Aziendali

Università della Calabria

Valutazione della Tesi di Dottorato *“Long-run export elasticities for industrial countries, 1990-2012"*, dottoranda Alessia Via.

Alessia Via presenta una tesi di Dottorato dal titolo *“Long-run export elasticities for industrial countries, 1990-2012"*.
La stima delle elasticità degli scambi internazionali rispetto ai prezzi relativi è uno dei temi più importanti, controversi e affascinanti dell'economia internazionale.
Il tema é di grande importanza e attualità, poichè dai valori di queste elasticità dipende la performance di diversi regimi di tassi di cambio; in particolare, la stessa sopravvivenza dell'area euro dipende dall'entità degli aggiustamenti in termini di prezzi relativi, e quindi di costi del lavoro, necessari per ridurre gli squilibri degli scambi con l'estero all'interno dell'area euro, che sono probabilmente la causa di fondo della crisi dell'euro. Il tema é controverso, poiché le stime ottenute per queste elasticità sono fortemente discordanti, e spesso in contrasto con le esperienze concrete di tanti paesi. Addirittura, fra gli anni 20 e gli anni settanta del secolo scorso emerse fra gli studiosi di economia internazionale una frattura fra "*elasticity pessimists*" ed "*elasticity optimists*", che ebbe il suo momento di massima contrapposizione con la proposizione negli anni settanta di un "approccio monetario all'analisi della bilancia dei pagamenti", che, rifiutando in blocco i risultati di tante stime econometriche, si basava sul postulato di valori infinitamente grandi delle elasticità degli scambi internazionali rispetto ai prezzi relativi (generalizzazione dell'ipotesi del "piccolo paese"). Il fatto che gli impegni richiesti ai paesi che hanno aderito all'Unione monetaria europea riguardino i saldi di finanza pubblica, ma non la dinamica del costo del lavoro, potrebbe forse essere interpretata come una implicita accettazione dell'approccio monetario alle bilance dei pagamenti.
La tesi di Dottorato di Alessia Via fornisce un contributo interessante su questo argomento. Nel primo capitolo viene presentato il *framework* teorico in cui il lavoro si colloca. Il secondo capitolo presenta alcuni dei modelli econometrici più utilizzati per la stime delle elasticità-prezzo degli scambi internazionali. Il terzo capitolo illustra i risultati delle stime econometriche ottenute per Italia, Germania, Francia, Stati Uniti, Giappone, Regno Unito e Cina, con riferimento al periodo 1990-2012. La tesi, accanto alle stime prodotte utilizzando il VECM, contiene anche un'applicazione della metodologia dei panel non-stazionari, ovviando così al problema di avere stime puntuali per singolo paese che non permettevano di fare un confronto significativo tra tutti i paesi sotto indagine e di capirne le dinamiche temporali; sono state inoltre prese in considerazione diverse variabili di controllo.
La valutazione complessiva sulla tesi é positiva.

Università della Calabria, 08 ottobre 2013

(Prof. Antonio Aquino)

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CICLO XXV

Long-run Export Elasticities for Industrial Countries, 1990-2012

Settore Scientifico Disciplinare SECS – P/01

Direttore: Ch.mo Prof. Patrizia Ordine

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CHAPTER 3: ESTIMATING EXPORT ELASTICITIES USING A VECTOR ERROR CORRECTION MODEL

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However, my greatest debt is with my children, my husband, my parents and my whole family without whose encouragement, understanding, patience and absolutely unselfish support, this thesis could not have been written. It is to them that I dedicate my research.

LONG-RUN EXPORT ELASTICITIES FOR INDUSTRIAL COUNTRIES

1990-2012

Abstract

External imbalances are a threat for the global economy and disorderly adjustments as well as errors in forecasting the effects of policies can yield strongly negative outcomes. Focusing on export price elasticities, my main purpose is to provide an overall view of the previous research carried out on trade elasticity issues and to analyze the implications of global current account imbalances. Export price elasticities estimated in the previous literature feature a high variability with values ranging from -0.14 to -3.13. Some of these results can be considered controversial with respect to one side of the current debate and cause complexity in their interpretation. I have first applied a cointegration model in an error correction framework to estimate export elasticities covering the period from 1990 to 2012 for countries that represent both surplus and deficit sides of the current debate: Italy, Germany, France, USA, UK, Japan and China. Furthermore, I have used a non-stationary panel technique to take into account both inter-country differences and dynamic variations. Using these estimates, in combination with the prevalent macroeconomic forecasts related to the issue, I have illustrated how variations in exchange rates and incomes can produce effects on exports.

Sintesi

Gli squilibri nei pagamenti internazionali rappresentano una minaccia per l'economia globale. L'obiettivo principale di questo lavoro è fornire una visione complessiva delle ricerche svolte sulle problematiche riguardanti le elasticità del commercio internazionale, ed analizzane le implicazioni per gli squilibri delle bilance commerciali. Per ciò che riguarda le esportazioni, le elasticità dei prezzi stimate nella letteratura presentano una forte variabilità con valori che variano da -0.14 a -3.13. Alcune di queste stime, in particolare, possono essere considerate controverse e sono complesse nella loro interpretazione. Per le stime è stato utilizzato in questa tesi un modello di cointegrazione nell'ambito del Meccanismo di Correzione dell'Errore per stimare le elasticità delle esportazioni per il periodo che va dal 1990 al 2012 per Paesi sia in surplus che in deficit di bilancia dei pagamenti: l'Italia, la Francia, la Germania, gli USA il Regno Unito, il Giappone e la Cina. Successivamente, per arricchire l'analisi con le variazioni nel tempo e nello spazio, è stato implementato un panel non-stazionario. Le stime ottenute sono state utilizzate per illustrare l'entità dell'impatto sulle esportazioni dei diversi paesi delle variazioni dei tassi di cambio.

CHAPTER 1

INTRODUCTION

1. PREFACE

One of the most important issues in applied International Economics is the effect on trade flows of changes in income and relative prices. The increasing interdependence among countries and their efforts to maximize benefits from international trade makes the import and export demand equation specifications essential not only for forecasting, planning and policy formulation but also for the quantification of welfare gains from trade (Hamori and Yin, 2011). The estimation of income and price elasticities of trade is consequently the main object of several studies on the determinants of imports and exports. Price elasticities are particularly important for estimating the effects of changes of relative prices on trade flows and for determining to which degree they adjust to these changes.

The "elasticities" approach of the econometric specifications has, in fact, always been used in international economics to determine the causes of trade just for its capacity both to explain the past and to forecast and, consequently, plan the future. The main elements of this model are the elasticities of exports and imports with respect to economic activity and to relative prices, and the influence of other factors, including global supply and increased variety and interdependence. Export elasticities, in particular, are often used to show the relative flexibility of certain exporters when facing a loss of competitiveness while the price elasticity of imports reflects consumers' fidelity to domestic or foreign goods.

All these reasons can only partially explain why the role played by trade elasticities is considered fundamental in translating economic analysis into policy-making.

Given the importance of the issue, economists are interested in understanding how it will evolve in the future and, above all, how empirical models and techniques can improve with respect to the existing literature.

2. OBJECTIVES AND CONTRIBUTIONS

Export elasticity estimation is one of the most important, controversial and intriguing topics in International trade as well as the oldest empirical efforts in economics. According to Goldstein and Khan (1985), there were 42 books and articles by 1957 and Stern et al. (1976) cite 130 articles from the period 1960-1975, which estimate the trade elasticities. Sprinkle and Sawyer (1996) pick up in 1976 and survey approximately 50 articles, which estimate the trade elasticities. Most econometric estimations indicate that price elasticities fall in a range of 0 to –4.0, while income elasticities fall between 0.17 and 4.5. The high variability of trade elasticities estimates suggests that there are still gaps in this research area: in fact, since the values of price elasticities vary considerably, the recent literature questions the effectiveness of real devaluation in affecting exports. The importance and the interest of the issue lies on the fact that performance of the different kind of exchange rate policies and systems depends on the results of these estimations.

In spite of more than 50 years of analyses, the estimation of price and income trade elasticities in the international scenario is still an open and highly significant empirical subject; perhaps, this interest can be addressed, among other factors, to questions that do not achieve a total concurrence of results:

- do exports actually expand after depreciations? If so, by how much?
- can exchange rates alone represent a feasible policy to improve the trade balance?

The topic is controversial because the estimated price elasticities of the last decades are extremely contrasting not only between one another but also with the concrete experience of many countries like Germany. It is true that countries like the USA and Germany both push other countries (i.e. China) to appreciate their currencies but, while the USA are facing deficit issues, Germany cannot say as well: its export market share has increased in recent years, especially towards Asia[1].

Something in this context does not figure: or the complaints of the major exporters are without basis, meaning that any sort of exchange rate manipulation is nearly worthless (i.e., exports are not so sensitive to movements in exchange rates at least at an aggregate level) or the elasticities reported by the literature are, for some reason, inexact.

The motivation for this research is, therefore, explained by the extremely important consequences of trade elasticity estimates; my interest arose after reading several articles (and the related issues and studies) on the global imbalances and on how changes in exports are habitually seen as the key for realignments in the international scenarios. This represents the starting point and inspiration of the study: the estimation of trade elasticities. During the course of the study, an

[1] Deutsche Bundesbank, OECD, National Accounts database, www.oecd.org.

increasing focus on export functions occurred and, in particular, on long-run export price elasticities. From an economic point of view, the constant complaints about devaluation policies could been seen, indeed, as a signal and could disclose further outcomes useful for the study of international interdependencies and trade patterns; from an empirical point of view, both research articles and reference texts provide alternative methods for deciding on the model structure and this was also a challenging aspect: the over fifty years of econometric development in time series analyses offer different milestones for the building of an appropriate model. Indeed, despite the immense literature, and perhaps, as a consequence of this, there are different approaches and a lack of uniformity not only in the models applied but also in the results of the estimations. This leads to another important consideration: due to the highly relevant use of trade elasticities, if the results of the estimation vary through sample periods, methods and models, other empirical analyses could be useful.

In an attempt to address all the above mentioned issues, the objectives of this research are to:

- review trade elasticity literature both from a theoretical and an empirical perspective;
- identify contradictions and/or discrepancies in the estimates of export price elasticities in the published literature in a comparative framework and for chosen countries;
- describe and apply appropriate models within a time series and non-stationary panel cointegration framework approach;
- summarize, interpret and discuss the elicited results and the techniques used for the estimation in order to address the question of the nexus between weak currencies and improved trade balances.

More specifically, my goal is to assess whether policymakers can exploit the relationship (if any) between exports and policy to weaken the currency and promote growth through competitive devaluations. In this context, after providing an overall view of the previous research carried out on this topic and illustrating the main issues related to it, this thesis complements the existing literature by implementing both time series and panel data techniques for non-stationary series using aggregate trade data.

The novelty of this research is the use of frontier panel techniques to estimate the coefficients of the long-run export price elasticities. This study can be considered relevant because, by estimating long-run export price elasticities and their dynamics over time, it could allow to understand more about the consequences of changes in relative prices, so that policymakers could assess better their interventions and attempt to implement correct adjustments.

To sum up, the research contributes to the body of knowledge by:

- summarizing the literature on export elasticities by describing the most established approaches ;
- estimating export price elasticities for seven countries (Italy, Germany, France, UK, USA, Japan and China) over the period 1990-2012 applying different cointegration methods;
- addressing the competitive devaluation issue.

3. OUTLINE

This thesis consists of the following three sections. Chapter 2 reviews the literature on different economic and, especially, econometric approaches to the estimation of international export elasticities. This chapter provides a discussion of some of the issues involved in the estimation of the price elasticities of the demand for exports according to the existing literature and its evolution over the years and some results. This is required in order to provide a general and detailed (although not exhaustive) outlook of the body of work.

Chapters 3 outlines the data used and the empirical setting of export elasticities and explains the econometric methodology applied to estimate the long-run export elasticity equation - i.e. the time series cointegration analysis using the vector error correction model (*VECM*).

In order to provide more robust estimations of long-run export elasticities, Chapter 4 refers to panel data techniques. In particular, it is based on the non-stationary panel pooled mean group (*PMG*) and the mean group (*MG*) models. The chapter reports the evidence found when allowing for inter-country and intra-country variations of short and long-run price-elasticity.

The discussion and the interpretation of the results as well as the conclusions of this study are presented in the final sections of the thesis.

Figure 3 briefly illustrates the research outline:

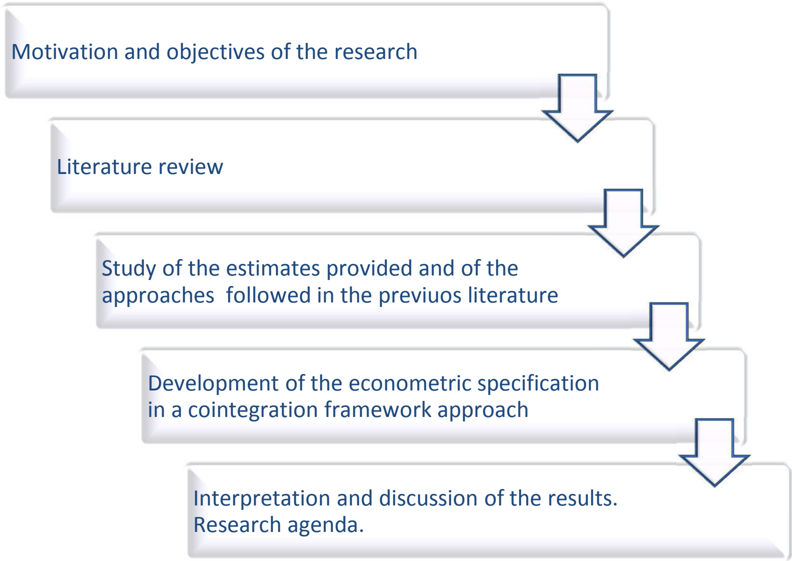

Figure 1.3. Research outline

CHAPTER 2

LITERATURE REVIEW

1. INTRODUCTION

The purpose of this section is twofold:

- to provide an overview of the previous theoretical and empirical literature within the international trade elasticities context;
- to act as a gateway to the methodology and the econometric specification applied in the present analysis.

Some preliminary remarks are necessary before starting.

First of all, in order to provide a fluent overview of the literature, it was decided to analyze the different studies proceeding by the main empirical and theoretical approaches followed and not by a chronological sequence: the existing literature, in fact, is very extensive - covering a period of over fifty years - and arranging the numerous researches by date would have made it very complicated to offer a general and overall outlook of the issue. In addition, taking into account the results of earlier empirical studies – equally important – an emphasis was reserved on the empirical contributions of the recent years.

Secondly, the econometric sophistication of time series goes hand in hand with the development of the international trade elasticities theories. For this reason, they are treated together.

Finally, this section is not to be considered exhaustive in including all the methodologies and cases studied up to now but rather it has to be read as a detailed summary which is intended to provide a background to the recent economic developments in times series econometrics and, in particular, in the estimation of international trade elasticities.

2. INTERNATIONAL TRADE ELASTICITIES: CONCEPT AND DEFINITION

Trade elasticities measure the responsiveness of demand or supply to changes in income, prices or other variables. The two main elasticities are the income elasticity and the price elasticity of demand. The income elasticity measures the percentage change in the quantity demanded resulting from a one-percent increase in income with E = elasticity, Q = quantity demanded, I = income and P = relative price (Escaith *et al.*, 2010):

$$E_I = \frac{\frac{\Delta Q}{Q}}{\frac{\Delta I}{I}} = \frac{I}{Q} * \frac{\Delta Q}{\Delta I}$$

The price elasticity measures the percentage change in the quantity demanded resulting from a one-percent increase in relative price:

$$E_P = \frac{\frac{\Delta Q}{Q}}{\frac{\Delta P}{P}} = \frac{P}{Q} * \frac{\Delta Q}{\Delta P}$$

2.1 The historical background

The estimation of trade elasticities has a very long history from both a theoretical and an empirical point of view. It is nonetheless firm that few papers cover all the issues raised in the econometric literature (Sawyer, Sprinkle, 1996).

2.1.1 The theoretical literature

The forerunner of the great amount of research concerning the estimation of trade elasticities is Orcutt (1950). Beginning with his paper, the large body of literature in this field has involved issues referring not only to how the elasticities are used and how they are determined but also to the development of the econometric specifications. These papers were first surveyed by Stern *et al.* (1976) and Goldstein and Khan (1985) and, since the 1970s, the literature has continuously evolved, entailing different issues related to trade elasticities.

The theoretical model underlying the estimation of trade elasticities is an imperfect substitutes model, that is, a model in which it is assumed that exports and imports are imperfect substitutes for domestically produced goods. Goldstein and Khan (1985) provide a detailed discussion of this model. In an imperfect substitutes model, the foreign demand for goods and services is determined by three main factors: foreign income, the prices of domestic goods and services, and the prices of goods and services that compete with domestic goods and services in the foreign market. Similarly, the domestic demand for foreign goods and services is determined by the country's income, the prices of foreign goods and services, and the prices of goods and services that compete with foreign goods and services in the domestic market.

The income elasticity of demand for imports measures to what extent changes in an importing country's income have an effect on changes in its imports. In the same way, the income elasticity of demand for exports measures to what extent changes in foreign countries' incomes affect the demand for exports.

Usually import and export elasticities with respect to income are positive, that is: an increase in a country's income leads it to buy more from foreign countries. An income elasticity of imports or exports that is equal to one implies that imports or exports increase at the same rate as income.

Divergences from this imply long-term imbalances in the global economy. Specifically, an income elasticity for imports greater than one implies that, at the margin, domestic consumers have a stronger preference for foreign goods than for domestic goods. This means that if prices do not adjust, imports increase more than proportionately to income growth. This case is particularly meaningful for countries that, on an international scale, experience a higher income growth rate (nota: emerging economies) since, compared to others, these countries will be encouraged to develop their demand of imports and this may possibly overweigh their exports: specifically, in many East Asian economies in which most of imports are used for re-exports, an increase in exports may entail, to some extent, a similar increase of imports. As a matter of fact, many of the imports into these countries are parts and components or capital goods that are used to assemble goods for re-export to the rest of the world. An exchange rate appreciation that reduces exports will also reduce the demand for imported goods that are used to produce exports (Thorbecke, 2010).

On the other hand, economic theory predicts that the volume of imported goods will decrease, while the volume of exported goods will increase, when the relative prices of a country's products decline, i. e. when its real exchange rate depreciates. The problem is: what happens to the value of exports and imports as a consequence of a country's real exchange rate depreciation? The answer depends upon the size of price elasticities of exports and imports.

2.1.2 The Imperfect Substitutes Model

Since the amount of export (and import) adjustments depends on the sensitivity to price and income variations, it is important to estimate the price and income elasticity of a country's exports. The theoretical basis of the empirical analysis is the Imperfect Substitutes Model. The basic assumption of the model is that neither exports nor imports are perfect substitutes for domestic products. Such a hypothesis is confirmed by empirical evidence. If domestic and foreign goods were perfect substitutes, a given country would be either an exporter or an importer. Since the world market is characterised by the presence of bilateral trade and the coexistence between imports and domestic production, the hypothesis of perfect substitution can be rejected (Algieri, 2004).

Moreover, a large body of empirical studies (Kravis and Lipsey (1978); Kravis and Lipsey) have shown that price differentials can be surprisingly large for the same product in different countries, as well as between the domestic and export prices of a given product in the same country. In other words, the *law of one price* (LOP) fails dramatically in practice, even for products that are usually traded in international markets. The LOP states that prices in different parts of the world for a given product should be the same when expressed in a common currency. All this said, the finite price elasticities of demand and supply that the imperfect substitutes models postulates can, therefore, be estimated for traded goods.

The imperfect substitutes model[2] (Goldstein and Khan 1985; Hooper and Marquez, 1995) of the home country's exports to, and imports from, the rest of the world (*) is formalized by a set of equations:

$M^d = \gamma\,(Y, P^M, P)$	$\gamma 1\,, \gamma 3 > 0, \gamma 2 < 0$	(1)
$X^d = \pi(Y^*\,e, P^X, P^*e)$	$\pi 1\,, \pi 3 > 0, \pi 2 < 0$	(2)
$M^s = \varphi[P^{M*}\,(1+S^*), P^*]$	$\varphi 1 > 0, \varphi 2 < 0$	(3)
$X^s = \xi[P^X\,(1+S), P]$	$\xi 1 > 0, \xi 2 < 0$	(4)
$P^M = P^{X*}\,(1+T)e$		(5)
$P^{M*} = P^X\,(1+T^*)/e$		(6)
$M^d = M^s e$		(7)
$X^d = X^s$		(8)

The eight equations identify the quantities of imports demanded by the home country (M^d), the quantity of exports demanded by the world from the home country (X^d), the quantity of imports supplied by the rest of the world to the home country (M^s), the quantity of the home country exports supplied to the rest of the world (X^s), the prices in domestic currency paid by the importers (P^M and P^{M*}) and the prices in domestic currency paid to the exporters (P^X and P^{X*}). The level of nominal

[2] This model is described is in Goldstein, and Khan (1985), Income and price effects in foreign trade.

income (Y, Y^*), the prices of domestic commodities produced within the regions (P, P^*), proportional tariffs (T, T^*), subsides to imports and exports (S, S^*) and the real exchange rates (e) are the explanatory variables.

Foreign demand for goods and services is determined by three main factors: foreign income, the prices of domestic goods and services, and the prices of goods and services that compete with domestic goods and services in the foreign market.

Similarly, the domestic demand for foreign goods and services is determined by the country's income, the prices of foreign goods and services, and the prices of goods and services that compete with foreign goods and services in the domestic market.

In the imperfect substitutes model the demand functions for exports and imports describe the quantity demanded as a function of the level of monetary income in the importing country, the imported product's own price, and the price of domestic substitutes. By considering a logarithmic utility function, the income (γ_1 and π_1) and price elasticity (γ_3 and π_3) of substitutes are assumed to be positive, while the price elasticity of the traded product is assumed to be negative (γ_2 and π_2).

Let us assume the demand function to be homogeneous of degree 0, equation 1 can be written in the following way:

$M^d = \gamma\ (Y/P\ ,\ PM/\ P) \qquad \gamma'_1>0,\ \gamma'_2<0$

where

Y/P = real income

and

$PM/\ P$ = real import price

Considering an n-country model, the symmetry between the demand function for imports and the demand function for exports vanishes. Imports compete, in fact, only with goods produced within the country. Exports compete both with goods produced in the imported country and with exported goods by third countries. The equation 2 is corrected with prices of competing goods.

$X^d\ /\ X^*$

$d = \pi\ (P^*\ X^d/P^*\ X^{*d})$

where X^{*d} is the demand for exports to the rest of the world from third countries.

The supply functions depend on the prices of exported and domestic goods and on subsidies. The price elasticities of exported and local commodities (φ_1 and ξ_1) are assumed to be positive, the price elasticities of substitutes (φ_2 and ξ_2) are supposed to be negative. The equilibrium conditions are represented by the last two equations. The implicit hypothesis is that prices move in order to equate demand and supply over time.

The imperfect substitutes model, by presenting both demand and supply side equations, allows to identify simultaneous relationships among quantities and prices. Orcutt (1950) and Goldstein and Kahn (1985), have highlighted this characteristic but, nonetheless, a multitude of time series works on export and import equations have considered the supply side only by assumption.

In the early 1990s, the standard methodology to estimate import (Eq. 1) and export demand (Eq. 2) was based on the assumption of an infinite supply-price elasticity for imports and exports (φ_1 in Eq. (3) and ξ_1 in Eq. (4)). Under this hypothesis, PM and PX were viewed as exogenous and thus estimated by single equations.[3]

Since the late 1990s, economic researchers have improved more and more their approach to the analysis and have applied cointegration analyzes or Fully-Modified-OLS methodologies to deal with simultaneity problems and to overcome endogeneity and serial correlation biases.

According to the literature examined in this study, therefore, the basic linear specifications for the export (2.1) demand function[4] can be expressed as follows:

$$Log\ X_t = a + b\ Log\ (PX/PXW)_t + c\ Log\ Y_t + \varepsilon_t \qquad (2.1)$$

where X_t = volume of exports, PX_t = export prices, PXW_t = world export price level, Y_t = world income and ε is an error term. The price elasticity is given by b[5].

A complete model will include other explanatory variables affecting demand besides income and prices. Houthakker and Magee (1969), for example, include control variables for domestic or

[3] If the supply elasticities were instead, less than infinite, the problem would be more difficult because one should either calculate the complete structural system of simultaneous equations or solve the reduced form for quantities and prices as functions of the exogenous variables in the system (Algieri, 2004).

[4] The import demand function is: Log Mt = a + b Log (PM/PD)t + c Log Yt+ εt where Mt = volume of imports, PMt = import prices, PDt = domestic price level, Yt = domestic income. The price elasticity is given by b.

[5] The income elasticity is given by *c*.

world GDP, to estimate the income elasticity of imports. Recently, Algieri (2011) included an unobserved component (UC) to model the export equation.

2.1.3 Relative prices and exchange rates: a complete pass-through

Thus far, enunciating the classical model provided by the literature, relative prices have always been mentioned as one of the most important variables of the export demand function. At this point, the question that could arise is how relative prices link to exchange rates and why recent models include real exchange rates rather than including (directly) relative prices of exported goods.

A last question to examine when talking about the export concerns, therefore, the relative prices. The main assumption made in this study (and according to the literature) is that there is a complete pass-through between relative prices and real exchange rates: that is, exchange rate fluctuations translate into proportional movements in the domestic price level and, therefore, pass-through is equal to one; this simplification offers the opportunity to gauge the price elasticity estimates using the real effective exchange rates without taking into account other factors that can determine divergent fluctuations and, most of all, without invalidating the estimates.

Exchange rate pass-through literature takes its roots from the aforementioned LOP and the Purchasing Power Parity (PPP) literature[6]. According to Anaya (2000), when using a 20-year time period, pass-through estimates for most countries are close to one supporting a long-term stable relationship. This kind of relation fits closely the purposes of the present study and, for this reason, the variable actually included in the model will be the real effective exchange rates.

2.1.4 The development of time series econometrics

Since the 1970s, the empirical literature evolved as a consequence of the rapid development of times series econometrics and of the need to consider the idea that trade flows do not respond instantly to changes in relative prices (and also in income and exchange rates). New theoretical and technical outcomes led to a vast number of papers beginning with Stern et al. (1976). Starting from the 1990s until today, the cointegration analysis and all the concepts related to it have become an important frame model.

As a result, the early specifications have experienced over fifty years of econometric sophistication, surveyed in Marquez (Marquez, 2002). In the last years, namely, Marquez (2002) or

[6] According to Anaya (2000), the Relative PPP (a weak version of the strong PPP) basically implies that the exchange rate and domestic and foreign price levels move proportionately to each other. Namely, strong version of PPP implies that $p_t = e_t p_t^*$ while the relative PPP implies pt = α et pt*where α is the real exchange rate or alternatively, is the home currency price level as a percentage of foreign.

Kwack *et al.* (2007) report some estimates for 8 Asian economies, including Hong Kong, the Philippines, whereas Cheung *et al*. (2009) estimate Chinese trade elasticities. The *new* models:

- include differences between short and long run elasticities;
- ponder the importance of heterogeneity between traded goods;
- study the stability of trade relationships;
- cope with endogeneity issues.

In particular, most of the researchers have tried to reduce endogeneity and this attempt is clear in all this recent (and vast) empirical literature. The effort consists mainly in introducing simultaneous equations and cointegration analysis. The main notion behind the cointegration analysis is that if a linear combination of a set of nonstationary variables (such as those in the import demand model) is stationary, those variables are said to be cointegrated. Indeed, recent developments in econometric literature have shown the non-stationarity of most macro data and this substantially invalidates the OLS, 2SLS and Instrumental Variable techniques results[7]. The Johansen (1988) Johansen – Juselius cointegration approach (1990) and the Engle – Granger (1987) two step approach have been used more and more to reveal the existence of long-run relationships and, in addition, produce empirical results that are not spurious (Marquez, 1990; Gagnon, 2003; Hooper, Johnson and Marquez, 1998; J.S. Mah, 2000).

2.1.5 The empirical model for trade elasticities

The empirical literature on trade elasticities goes back to at least Kreinin (1967) or Houthakker and Magee (1969) followed by Khan (1974), (1975), Goldstein and Khan (1976), (1978), Wilson and Takacs (1979), Warner and Kreinin (1983), Haynes and Stone (1983), Bahmani-Oskooee (1986), Marquez (1990) and Mah (1993). This plethora of studies of the past have all estimated trade elasticities using the OLS, 2SLS method or Instrumental Variables methods. (Bahmani-Oskooee, Niroomand, 1998).

It is indisputable that the empirical literature is vast and that most of the attention is addressed to the forecasting properties of the estimates. Most econometric estimations indicate that price elasticities fall in a range of 0 to –4.0, while income elasticities fall between 0.17 and 4.5[8]. Since the values of price elasticities vary considerably, the recent literature questions the effectiveness of real devaluation in affecting exports and imports. According to Rose (1990, 1991)

[7]When data are nonstationary, inferences based on the standard techniques are no longer valid because they suffer from the "spurious regression" problem, see Bahmani-Oskooee, Niroomand, (1998).

[8] Algieri B., (2004), *Price and Income Elasticities of Russian Exports*, The European Journal of Comparative Economics, Vol. 1, n. 2, 2004, pp. 175-193.

and Ostry and Rose (1992), a real depreciation does not impact significantly on the trade balance; Reinhart (1995), Senhadji and Montenegro (1998), Senhadji and Montenegro (1999) provide instead, strong support to the view that depreciations improve the trade balance. It seems that low econometric estimates of price elasticities are unreliable for the purpose of forecasting the effect of a depreciation, and there is a strong presumption that these elasticities lead to a considerable underestimation of its effectiveness (Algieri, 2004).

Modeling the time series behavior of imports and exports is a longstanding issue of economists as well as of econometricians; well along with their research, their main questions concern:

- the *type* of traded commodity, that is, if it is a homogeneous or a differentiated good;
- the *main purpose* to which the traded good is designed for, that is, if it is used as a factor of production or as a final product;
- the *institutional or legal structure* of the environment in which the trade takes place;
- the *aim* of the modeling analysis, or better, if the intention is to forecast or to test hypotheses;
- the *typology of data available*, that is, if data are annual, quarterly, monthly, etc.;
- the level of aggregation, that is, if the data are aggregated or disaggregated (and the entity of the disaggregation).

The appropriate model, indeed, relies on all the above mentioned factors.

3. ESTIMATING TIME SERIES: DIFFERENT APPROACHES AND EMPIRICAL CONTRIBUTIONS FOR TRADE ELASTICITIES.

The following sections survey some of the approaches used in the estimation of time series variables. The discussion of each topic will be illustrated by examples and empirical analyzes of selected references.

3.1 Distributed-Lag and Autoregressive Models approach

The estimation of price elasticities is a fundamental part of the econometric analysis of long-run relations. This category of analysis has been the focus of much theoretical and empirical research in economics. Where the variables in the long-run relation of interest are trend stationary, the general practice has been to de-trend the series and to model the de-trended series as stationary distributed lag or autoregressive distributed lag (ARDL) models.

In this section, after a brief illustration of the theoretical issues underlying the distributed-lag and the autoregressive models, the literature overview will examine how the study of trade elasticities of demand has been treated over the years through empirical contributions.

3.1.1 Modeling time lags and long-run relations

Time series data entail a variety of issues related to the fact that the regression model includes not only the current value but also the past values of the variables. The time passing between a cause and its effect is called a *lag*. The lag may be a specific time (e.g., three months, one year, etc.) but, in many cases, the effects of an economic cause are spread over many months, or even many years. In such cases, we have a *distributed lag*.

Lags occupy a central role in economics, principally when dealing with aggregate data.

When the model includes past values of the regressors (explanatory variables indicated by the *X*'s), these past values are called *lagged* values and, therefore, the regression analysis is called *distributed-lag model*.

Furthermore, the dynamic behaviour of an economy can reveal itself through a dependence of the current value of an economic variable on its own past values.

When the model includes past values of the dependent variable (*Y*) among the explanatory variables, it is called an *autoregressive model (AR)*.

We can present a general distributed-lag model with a finite lag of *k* time periods as:

$$Y_t = \alpha + \beta_0 X_t + \beta_1 X_{t-1} + \beta_2 X_{t-2} + \ldots + \beta_k X_{t-k} + u_t \qquad (2.2)$$

The coefficient β_0 is known as the short-run multiplier because it gives the change in the mean value of *Y* following the unit change in the *X* in the same time period. The coefficients β technically can be expressed as the partial derivatives of *Y* with respect to the *X*'s[9]:

$$\frac{\partial Y_t}{\partial X_{t-k}} = \beta_k$$

In this model, after *k* periods[10], we obtain the long-run (or distributed-lag) multiplier:

[9] β0 is the partial derivative of Y with respect to Xt, β1 with respect to Xt-1, β2 with respect to Xt-2, and so forth.

[10] In the model [2.2], if the explanatory (input) variable *X* undergoes a one-off unit change (impulse) in some period t, then the immediate impact on *Y* is given by β0; β1 is the impact on *Y* after one period, β2 is the impact after two

$$\sum_{i=0}^{k} \beta i = \beta_0 + \beta_1 + \beta_2 + \ldots + \beta_k = \beta \ .$$

The autoregressive model, instead, is expressed as:

$$Y_t = \alpha + \beta X_t + \gamma Y_{t-1} + u_t \qquad (2.3)$$

The autoregressive model actually describes the time path of the dependent variable in relation to its past values and for this reason it is properly known as a *dynamic model*.

We have to consider that the real world presents a mixture of short-run and long-run adjustments and that these adjustment times are necessary for the dependent variable (in our case, the export/import volumes) to respond to variations in the explanatory variables (relative prices and income). The fact that the dependence of a dependent variable *Y* on other variables (the *X*'s) is rarely instantaneous implies that any kind of analysis that involves time series data needs to consider such lapse of time (the so called *lag*).Very more often, indeed, the effect of a given cause is distributed over a certain period of years. Obviously, for this reason, a closer attention should be paid to the factors which account for distributed lag relationships in order to comprehend the economic theory underlying the nature of the lags and why they occur, first of all.

Generally, there are three main reasons[11] that are used to explain why lags occur:

1. *psychological reasons*, under which we include forces of habit and assumptions on the part of consumers that changes may be only temporary;
2. *technological reasons*, which include factors such as, in a general case, the lack of knowledge about possible substitutes[12];
3. *institutional reasons*, which include situations in which certain contractual items of expenditure or savings may need to be adjusted before shifts can be made in consumption patterns.

In addition, it is clear that the lapse in reactions can depend on the nature of the variation, that is, if the change is permanent or transitory.

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periods, and so on. The final impact on *Y* is βk and it takes place after *k* periods. Hence, it takes k periods for the full effects of the impulse to be realized. The sequence of coefficients (β0, β1, β2,…, βk) constitutes the impulse response function of the mapping from *X*t to *Y*t.

[11] Nerlove (1958).

[12] "Suppose the price of capital relative to labour declines, making substitution of capital for labour economically feasible. Of course, addition of capital takes time (the gestation period). Moreover, if the drop of price is expected to be temporary, firms may not rush to substitute capital for labor, especially if they expect that after the temporary drop the price of capital may increase beyond its previous level. Sometimes, imperfect knowledge also accounts for lags." (Gujarati, 1995).

3.1.2 The ARDL model

As above mentioned, another important case is when we find a dependence of the current value of an economic variable on its own past values. Precisely, models of how decision/policy makers' expectations are formed, and how they respond to changes in the economy, result in the value of *Y*t depending on lagged *Y*'s. So, an alternative way to capture the dynamic component of economic behaviour is to include lagged values of the dependent variable on the right-hand side of the regression together with the other exogenous variables.

In time-series econometric modeling, a dynamic regression will usually include both lagged dependent and independent variables as regressors:

$$Y_t = \alpha_0 + \alpha_1 Y_{t-1} + \ldots + \alpha_p Y_{t-p} + \beta_0 x_t + \beta_1 x_{t-1} + \ldots + \beta_k x_{t-k} + \varepsilon_t. \qquad (2.4)$$

The above model is called the *Autoregressive Distributed-Lag model*, known as ARDL (*p; k*). The values of *p* and *k* (i.e., how many lags of *Y* and *X* will be used) are chosen:

i. on the basis of the statistical significance of the lagged variables;

ii. so that the resulting model is well specified (e.g. it does not suffer from serial correlation).

3.1.3 Estimation issues

A distributed-lag model can be estimated by OLS but this approach leads to a certain number of problems. The first question is related to the length of the lag: even though some economical and/or theoretical considerations must be brought forward on the β's to avoid estimation problems, there is no way to know *a priori* what the maximum length of the lag is supposed to be.

Secondly, in time series data, the successive lags tend to be highly correlated (Gujarati, 1995) and this means that we are dealing with multicollinearity and, consequently, with imprecise estimation.

Finally, since results *are* sensitive to lag-lengths, the search for the optimum lag length can widen the dangerous doors of data mining. Data mining refers to all those activities that have in common, a search over different ways to process or package data econometrically with the purpose of making the model meet certain design criteria (Hoover and Perez, 2000). A ready-mix model does not exist so the general recommendation is to always follow (economic and econometric) theory as a guide for any model building.

3.1.4 Empirical contributions

Some studies[13] (Bahmani-Goswami, 2004) relied on bilateral trade data to provide strong evidence for the support of positive long-run relation between exchange rate and trade balance. In this sense, the main question is if currency devaluation can be used as a tool to correct trade balances.

When researches follow the traditional approach to estimate import demand and export demand elasticities using aggregate trade data, the problem is that a significant price elasticity with one trading partner could be more than compensated by an insignificant elasticity with another partner yielding an insignificant trade elasticity. This so-called "aggregation bias" problem requires estimating trade elasticities on a different basis. Indeed, a new body of the literature is emerging and it includes analyzes that estimate trade elasticities on bilateral basis.

Ketenci and Uz, (2011) use an ARDL approach to measure the impact of currency devaluation. The assessing of the impact is carried out using the real exchange rate. The model is applied between the EU and its eight major industrial trading partners and, furthermore, its six major trading regions[14]. The ARDL approach is used to determine whether the dependent and the independent variables are cointegrated. The ARDL approach involves two steps for estimating the long-run relationship:

i. the first step is to examine the existence of long-run relationship among all the variables in an equation;

ii. the second step (applied only if the first step showed a cointegration relationship) is to estimate the short-run and the long-run coefficients of the same equation.

Cointegration relations between the variables of bilateral import and export functions are found to be due not only to the strong relations between trade and real exchange rate but also to

[13] "First, some have followed standard textbook prescription by estimating the well-known Marshall-Lerner (ML) condition. The ML condition states that devaluation improves a country's trade balance if the sum of the price elasticities of that country's import and export demands is more than unity. Second, some have argued that estimating the ML condition is an indirect approach. Thus, they have adhered to a direct method of establishing a link between the trade balance and exchange rate (as well as other determinants) following a reduced-form modeling approach. In both the first and the second approaches, researchers have mostly used aggregate data and provided mixed results. Recently, a few studies have concentrated on employing disaggregated data. While the disaggregated approach on a bilateral basis is applicable to reduced-form trade-balance models, it cannot be applied to estimate the trade flow elasticities or the ML condition. This is due to the fact that import and export prices are not available on a bilateral basis to obtain bilateral trade volumes. The remedy here is to establish a direct link between a country's inpayments (value of exports) and outpayments (value of imports) and real exchange rate on a bilateral basis" (Bahmani-Goswami, 2004).

[14] Namely: Canada, China, Japan, Norway, Russia, Switzerland, Turkey and the US; the six major regions are: the NMCs, the CEECs, the EFTA, the NAFTA, the ASEAN and the DACs.

important relations between EU import and EU income and between EU export and partners' income. The authors also employ an Error Correction Model (ECM) to provide additional evidence of cointegration among the variables.

Since the data for import and export prices on a bilateral level are not available, the authors cannot estimate trade flow elasticities for determining the Marshall-Lerner (ML) condition in their model. Consequently, they establish a direct link between a country's value of import and export and real exchange rate on the bilateral basis. The direct method of determining whether currency depreciation is effective in increasing a country's inpayments from a trading partner consists in considering the export value (or inpayments) and determining how sensitive it is to a change in exchange rate; similarly, they consider import value (or outpayments) and try to determine its sensitivity directly to a change in exchange rate.

The authors employ the Akaike Information Criterion (AIC) in order to select the optimum lag length.

As for the diagnostic statistics, the authors' models pass all the following tests:

- the Lagrange Multiplier (LM) test to check for the serial correlation among the residuals;
- the Jarque-Bera statistic to test the normality of the residuals;
- White's test to check the heterosckedasticity of the residuals;
- Ramsey RESET test[15] to check the functional misspecification of each model.

The long-run elasticities of real exchange rate are used as a proxy of price elasticity for determining the ML condition but they result too low so that the long-run coefficient estimates do not provide any empirical evidence that the ML holds.

For the bilateral export demand equation, the expected positive sign of the long-run real exchange rate elasticities occurs only for three partners (on a total of 14) but it is not significant; in other three cases the real exchange rate is significant but has a negative (unexpected) sign. The negative sign means that there is an adverse effect of the currency depreciation on the bilateral export between the EU and the involved trade partners.

For the bilateral import demand equation, the expected negative sign of the long-run coefficients of the real exchange rate occurs in two cases and the coefficients are significant; this indicates that real depreciation of the euro decrease European imports from the regions involved while in other two cases the adverse effect is observed.

[15] Using the square of the fitted values and distributed as X^2 with 1 degree of freedom.

However, the real exchange rate elasticities are so low that it is impossible to make a conclusion about the effect of currency depreciation on bilateral trade and, according to these results, it actually can be concluded that EU's imports and exports are insensitive to exchange rate variations[16].

Yin and Hamori (2011) have recently analyzed China's import demand function employing ARDL and Dynamic Ordinary Least Squares (DOLS)[17] techniques in the estimation of the long-run coefficients of price and income elasticities. The aim of this study is to resolve the issue of trade balance from the perspective of China's policy making in the larger context of global imbalances problems.

The import demand model adopted derives from the imperfect substitution theory according to which the demand for real imports is a function of domestic income or GDP and relative price (import price index deflated by an index of domestic prices)[18].

The choice of the ARDL is motivated by the authors' consideration that, according to recent evidence, it possesses desirable small sample properties and can effectively correct for possible endogeneity of explanatory variables; they also include the estimates from DOLS, regarded as one of the most widely used estimators of cointegrating vectors in applied literature.

Before interpreting the estimated import demand equations, the existence of a long-run import demand relationship is analyzed. In order to do so, different bound tests are employed using the Johansen (1991, 1995) approach. Since there is strong evidence of the existence of a long-run relationship among the variables included in the long-run import demand model, they estimate the long-run cointegration relationship (long-run coefficients) for imports using the ARDL and DOLS single equation estimation methods.

The authors employ the Schwarz-Bayesian Information Criterion (SBIC) in order to select the optimum lag length.

The diagnostic tests carried out by the authors are the following:

- autocorrelation tests

[16] See also Bahmani-Goswami, 2004.

[17] See paragraph 3.3 of the current section for the DOLS empirical contributions.

[18] The traditional import demand model is expressed as: $ln\ M_t = \alpha_0 + \alpha_1 lnY_t + \alpha_2\ ln\ RP_t + \varepsilon_t$, where ln is the natural logarithmic form and ε_t is the error term. M_t denotes the volume of imports at time t, Y_t denotes real income at time t, and RP_t denotes the relative price (the import price index deflated by a GDP deflator) at time t. Generally, the hypothesized values of the coefficients of the explanatory variables are $\alpha_1 > 0$ and $\alpha_2 < 0$, which represent the income and price elasticities respectively of import demand.

- the Jarque-Bera statistic to test the normality of the residuals;
- heterosckedasticity test of the residuals.

The estimated residuals did not provide any significant evidence of serial correlation, non-normality or heterosckedasticity in the error term.

The DOLS estimates of the relative price coefficient are higher as compared to those from ARDL. Indeed, the estimated coefficients for income and relative price variables were found to be rather different when different estimation techniques were employed.

For what concerns the policy-making implications, they reported the following considerations: first of all, the estimated long-run elasticity is inelastic and approximately within the range of –0.5 to –1. Hence, it appears that China cannot depend on using its exchange rate policies to correct the balance of trade problem. The long-run price elasticity is statistically significant, suggesting that if the growth in inflation in China is related to the import price, then China's imports will increase (and the trade balance will get worse). Second, the growth in income has a significant and elastic impact on import demand in the long-run.

3.2 Phillips' Fully Modified Estimator Approach

Senhadji and Montenegro (1999) have conducted a cross-country analysis for a large number of developing and industrial countries of export demand equations to gauge export price and income elasticities. The technique implemented accounts the non-stationarity for the data and is derived from dynamic optimization. In this study, the export demand equation is expressed as follows:

$$log\,(x_t) = \gamma_0 + \gamma_1 log(x_{t-1}) + \gamma_2 log(p_t) + \gamma_3 log(gdpx_t^*) + e_t \qquad (2.5)$$

where x_t is real exports of the home country; p_t is the export price of the home country relative to the price of its competitors; and $gdpx_t^*$ is the activity variable defined as real GDP minus real exports of the home country's trading partners. Thus, this model is close to the standard export demand function except that the income variable is real GDP minus real exports of the trading partners, rather than the trading partners' GDP. The equation is estimated using both OLS and Phillips' Fully Modified estimator (FM) The fully modified OLS (FM-OLS) approach originally proposed by Phillips and Hansen (1990) which takes into account the non-stationarity in the data as well as potential endogeneity of the right-hand side variables and autocorrelation of the error term.

The FM estimation method is, indeed, an approach to regressions for time series taking advantage of data non-stationarity and cointegrating links between variables. Cointegrating links between variables lead to endogeneity of regressors: the FM estimator is designed to estimate cointegrating relationships by *modifying* OLS with corrections that take into account of endogeneity and serial correlation.

The long-run price and income elasticities are defined as the short-term price and income elasticities divided by one, minus the coefficient estimate of the lagged dependent.

The average long-run price and income elasticities are found to be approximately -1 and +1,5, respectively; the short-run price elasticities present an average of -0,21 and the average short-run income elasticity is 0,41. Thus, exports do react to both the trade partners' income and to relative prices. According to the results of the estimation, exports seem to be much more responsive to changes in relative prices in the long-run than in the short-run. In particular, among the 53 countries investigated, there are Italy, France, Japan, USA, UK and China (5 of the 6 countries analyzed in the present research) and the reported FM export price elasticities are:

Table 2.1 Export price elasticities using FM-OLS for Italy, France, USA, UK, Japan and China

Country	*OLS Price elasticity*	*FM Long-run Price elasticity*	*FM Short-run Price elasticity*
Italy	-0.07	-0.13	-0.05
France	-0.01	0.00	-0.02
USA	-0.19	-0.69	-0.03
UK	-0.16	-0.33	-0.12
Japan	-0.25	-1.33	-0.17
China	-0.78	-3.55	-0.63

Table 2.1 Export price elasticities using FM-OLS.
Source: Senhadji and Montenegro (1999).

Asian countries show significantly higher price elasticities than both industrial and developing countries. In addition, according to the authors' results, Asian countries benefit from higher income elasticities than the rest of the developing world, corroborating the general view that trade has been a powerful engine of growth for these economies.

3.3 DOLS approach

The Ordinary Least Squares (OLS) method provides estimates of the regression slopes that are consistent and converge at rate T where T is the sample size. When there is correlation between the regression error and the regressors, i.e. endogeneity, OLS estimates have an asymptotic bias which makes inference difficult. In order to overcome this bias, several methods have been proposed, and, besides the FM approach, one of these is the Dynamic OLS model (DOLS). The dynamic OLS (DOLS) approach proposed by Stock and Watson (1993) augments the original regression with lags of the first differences of the regressors. If the lag structure is chosen in a suitable way, the asymptotic bias is removed but the choice of lag remains an important practical issue and often researches do not have a practical guidance on how to choose them.

In the equation[19]:

$$Y_t = \alpha + \theta X_t + \varepsilon_t \qquad (2.6)$$

θ is the cointegration coefficient and it is the result of an OLS regression.

If X_t and Y_t are cointegrated, the OLS estimator in the regression of the cointegration coefficient in (2.6) will be inconsistent. Generally, the OLS estimator can lead to problems and to wrong results. For this reason, econometricians have developed alternative estimators able to measure the cointegration coefficient. One of these estimators is exactly the DOLS. The DOLS estimator is based on a modified version of equation (2.6) that includes past, present and future values of X_t:

$$Y_t = \beta_0 + \theta X_t + \sum_{j=-p}^{p} \delta j\, \Delta X_{t-j} + u_t \qquad (2.7)$$

Therefore in equation (2.7), the regressors are: X_t, ΔX_{t+p}, …ΔX_{t-p}. The DOLS estimator of θ is the OLS estimator of θ in equation (2.6).

If X_t and Y_t are cointegrated and the sample is numerous enough, then the DOLS estimator is efficient. Furthermore, since X_t and Y_t , being cointegrated, have a common stochastic trend, the DOLS estimator remains consistent even if X_t is endogenous.

The Dynamic OLS (DOLS) approach is used by Aziz and Li (2007) to estimate the export and import equations for China using quarterly data from 1995:Q1–2006:Q4. DOLS is chosen

[19] Stock and Watson, (2003).

because of its small sample property: indeed, Monte Carlo experiments show that with finite sample, DOLS performs well.

According to the authors, using aggregate data, export elasticity to foreign demand is +3.8, and to relative price is −1.6. These estimates are within the range of other studies (Goldstein and Khan, 1985) and satisfy the Marshall-Lerner condition. In the discussion of the results, the authors argue that great cautious is needed when using trade elasticities to estimate the response of the Chinese economy to price and demand shocks. Trade elasticities used in existing studies on such subjects vary widely and such variation reflects not only data and methodological issues involved in estimating elasticities for all countries, including developed countries, but also a continuous structural shift in how production is organized in China. China is shifting away from stereotypical processing trade that involves mostly assembling imported parts and components to domestically sourcing larger portions of the production chain (Aziz and Li, 2007). In conclusion, the fast changing structure of China's trade also raises questions about how much one can rely on these estimates, especially the interaction between exchange rate and trade composition changes. Any analysis that does not take into account these factors, continuing to be influenced by China's past trade structure could lead to erroneous outcomes.

The DOLS procedure has been also adopted by Caporale and Chui (1999) to estimate the long-run income and price elasticities of trade for 21 countries, using annual data over the period 1960-1992, in a cointegration framework. According to the authors, faster growing economies have high income elasticities of demand for their exports but lower import elasticities. For what concerns in particular Italy, Germany, France, Japan, USA and UK the export elasticities estimates are:

Table 2.2 Price and income export elasticities using DOLS for Italy, France, USA, UK, Japan and Germany.

Country	*DOLS Price elasticity*	*DOLS Income elasticity*
Italy	**-0.93**	2.21
France	**-0.08**	2.13
USA	**-0.63**	1.40
UK	**-0.19**	1.29
Japan	**-1.70**	2.91

Germany	**-0.11**	1.87

Table 2.2 Price and income export elasticities using DOLS.
Source: Caporale and Chui (1999).

The estimation for price elasticities in a few cases (among others, France and Germany) are not statistically significant.

3.4 Cointegration and Error Correction Model Approach

The empirical analyzes employed in Chapters 3 and 4 rely on the cointegration techniques and, therefore, a more particular and detailed attention will be addressed to this econometric topic.

The econometric modeling of time series data has seen remarkable growth in recent years. The advancements made in the analysis of times series models over the last three decades are partly due to the developments of theoretical models and partly due to the improvements in computational ability. In earlier years, the analysis of time series models was strictly limited mainly by the time available to execute repetitive calculations, but with the advances made in software development most of the models developed in the early 1970s, 1980s and 1990s have become standard in statistical software packages.

In particular, over the last twenty years, one of the time series modeling research directions has been the development of the theory of *cointegrated time series* modeling based primarily on the seminal paper of Engle and Granger (1987). The theory was further developed by many other authors such as: Johansen (1988, 1991), Stock and Watson (1991), Johansen and Juselius (1992) and Pesaran and Shin (1997).

Cointegrated modeling is applied when the series under investigation are nonstationary and to test this circumstance Dickey and Fuller (1979) developed the initial theory and methodology for the stationarity testing of a time series: since then the analysis of nonstationary time series data has generated considerable research interest. Actually, one of the most important assumptions of what can be called the "classical" econometric theory was that the observed data came from a stationary process, meaning a process whose means and variances are constant over time; thus, the time series y_t is stationary if, for all values and every time period, it is true that:

$E(y_t) = \mu$ (constant mean)

$var\ (y_t) = \sigma^2$ (constant variance)

$cov\ (y_t, y_{t+s}) = cov\ (y_t, y_{t-s}) = \gamma_s$ (covariance depends on *s* and not on *t*)

Actually, economies evolve, develop and change over time in both real and nominal terms and, therefore, forecasts are often badly wrong (Hendry and Juselius, 1999). At the end, the practical problem facing econometricians is to find any kind of relationship that survives long enough to be useful.

Since the late 1990s, it seems clear that stationarity assumptions must been treated with caution or even completely abandoned for most observable economic time series[20].

Summarizing, it can be said that:

i. when dealing with time series data, the assumption of stationarity for modeling and inference is crucial: indeed, when data means and variances are non-constant, observations come from different distributions over time, and this leads to difficult problems for empirical modeling;

ii. the effects of incorrectly assuming constant means and variances when that is false is potentially risky and can induce serious statistical mistakes;

iii. the sources of non-stationarity are many and can be can be due to evolution of the economy, legislative changes, technological change, and political disorders;

iv. empirical specifications can be *transformed* so that stationarity can become a valid assumption: that is, some forms of non-stationarity can be eliminated by transformations.

In conclusion, given that stationarity is an important issue when dealing with time series variables:, it must be clear that, if we are interested in estimating parameters or testing hypotheses in cases where the set of variables are not all stationary, standard OLS techniques are generally invalid and, thus, inappropriate.

Besides the stationarity issue, another question that arises when regressing a time series variable on another time series variable is that of the so-called spurious regression. This problem takes place when, due to the presence of a trend, it seems as if a relationship between two economic variables exists but this is not true: hence, the results (or their interpretation) are doubtful More precisely, *spurious regression occurs when a pair of independent series, but with strong temporal properties, is found apparently to be related according to standard inference in a Least Squares*

[20] "Intermittent episodes of forecast failure (a significant deterioration in forecast performance relative to the anticipated outcome) confirm that economic data are not stationary: even poor models of stationary data would forecast on average as accurately as they fitted, yet that manifestly does not occur empirically". Hendry and Juselius (1999).

regression[21]. Spurious regressions can be defined as *nonsense* correlations and may result when one nonstationary time series is regressed against one or more nonstationary time series. Data involving economic time series often tend to move in the same direction, reflecting an upward or a downward trend: the presence of this common trend is simply a spurious association between two time series and not a *true* association: indeed, despite the absence of any genuine long-run relationship between the underlying series, there is found a (spurious) relationship. In these cases, an extremely high R^2 value leads to erroneous considerations. In particular, let us have:

$$y_t = y_{t-1} + u_t \qquad u_t \sim iid(0,\sigma^2)$$

$$x_t = x_{t-1} + v_t \qquad v_t \sim iid(0,\sigma^2)$$

where u_t and v_t are serially and mutually uncorrelated.

$$y_t = \beta_0 + \beta_1 x_t + \varepsilon_t$$

since y_t and x_t are uncorrelated we should expect R^2 to tend to zero, but this is not the case: indeed, if two series are growing over time, they can be correlated even if the increments in each series are uncorrelated. These series have no relation to one another yet, when plotted, there seem to be a positive relationship between them.

As aforesaid, one of the aims of this section is to provide an overview of the studies that deal with cointegration techniques and to outline some important empirical contributions. In Chapter 3, export elasticities have been estimated applying a cointegration technique within an error correction framework[22].

Briefly, explaining the temporal behavior of a set of variables that, on the basis of economic theory share a relationship that holds in equilibrium, if we observe that there are deviations from this equilibrium for all periods for which we have observations, what we need is to examine are the properties of these deviations, i.e., of the disequilibrium errors. The errors can be small on average and remain so over time: that is, if the errors are viewed as random variables, they all have an expected value of zero and the same variance also. These characteristics are typical of a stationary time series. Nevertheless, this is clearly not the case of all macroeconomic variables. Consumers' expenditure and disposable income, for example, whether measured in real or nominal terms, are certainly not stationary series, but instead exhibit trends in their means over time. If these trends are

[21] Quotation from Phillips (1986).

[22] The explanation of the reasons that justify this choice and the detailed description of the model are provided in Chapter 3.

stochastic trends, such that the first difference of the series is trendless, the series in question are said to possess a single *unit root*. It may be the case that a series becomes stationary only after differencing more than once, in which case it has multiple unit roots. Economic theory and visual plots of the time series can provide prior information or hints as to whether a series is expected to have one (or more) unit root, to be certain of this circumstance is necessary to verify it empirically through appropriate tests[23].

When the variables are not only stochastically trended (and feasible to become stationary after differencing), but also have common stochastic trends, some suitable linear combination of these variables will be stationary even though the level of each series is not stationary: this is the case of cointegrated variables.

In order to define cointegrated variables, let us have:

$X_t \sim I(1)$

and $Y_t \sim I(1)$

but $Z_t = Y_t - \beta X_t \sim I(0)$,

then X_t and Y_t are cointegrated, i.e., there is a long-run relation between the variables and β is the cointegrating vector and expresses the equilibrium relationship between the series Y_t and X_t and u_t was the *departure* from the long-run equilibrium path.

One of the approaches used in literature to test how many cointegration relationships there are is the Johansen procedure. It comprises two tests: the "λ-max" test, for hypotheses on individual eigenvalues, and the "trace" test, for joint hypotheses. Further on in the present section and in Chapter 3, the Johansen procedure will be explained in detail.

If a single time series variable presents a single unit root then we need to take the first difference of the variable in order to obtain a stationary series. However, the problem is that we are not concerned in a single variable viewed separately from others but rather in the relationship *between* variables. For this reason, it is more useful to consider differencing within the context of a regression model. For example, let X be exports and I be income and consider the model

$$\Delta X_t = \beta \Delta I_t + u_t \qquad (2.8)$$

In (2.8) the change in exports from one period to the next is explained by the change in income in the same time window without reference to any equilibrium or long run relationship

[23] The test used in the present study is the ADF Unit Root test, *cfr*. Chapter 3.

between consumption exports and income that may exist: that is, the equation (2.8) does not have an equilibrium or long-run solution. If the assumption is that in equilibrium the variables become constant we want to impose

$$X_t = X_{t-1} = X_{t-2} = \ldots \text{ and } I_t = I_{t-1} = I_{t-2} = \ldots$$

on (2.8). Doing this, (neglecting the disturbance term) gives 0 = 0 and so does not provide an expression for *X* in terms of *I,* does not explain the impact of *I* on *X*. Such an expression would be, indeed, the equilibrium solution of (2.8).

If we suppose that the value of *X* in equilibrium is given by X^* and assume $X^* = f(I)$, we can try to find this long-run solution. It is necessary to introduce a variable that takes into account the level of *X* in period *t-1* relative to the equilibrium value of *X* for the same period $(f(I_{t-1}))$. This leads to the following equation:

$$\Delta C_t = \beta \Delta I_t + \theta\big(C_{t-1} - f(I_{t-1})\big) + u_t \qquad (2.9)$$

in which both short-run and long-run factors are allowed a role to play in determining how *X* is changed from its value in period *t-1*. The new variable is the period *t-1* discrepancy between actual *X* and equilibrium *X* and, given this, we expect θ to be a negative parameter because it makes the variables *return* to equilibrium (after going away from it).

If we assume that f is a linear function and write:

$X^*_t = f(I_t) = \lambda I_t$

we obtain

$\Delta X_t = \beta \Delta I_t + \theta (X_{t-1} - \lambda I_{t-1}) + u_t$

or

$\Delta X_t = \beta \Delta I_t + \theta X_{t-1} - \theta \lambda I_{t-1} + u_t$

suggesting that a regression of ΔX_t on ΔIt, X_{t-1} and I_{t-1} is required. This kind of econometric specification is called an *Error Correction Mechanism* (ECM).

Summarizing the main features of the ECM, it can be said that:

- the error correction term tells us the speed with which our model returns to equilibrium following an exogenous shock;

- it should be negatively signed, indicating a move back towards equilibrium, a positive sign indicates movement away from equilibrium;
- the coefficient should lie between 0 and 1, 0 suggesting no adjustment one time period later, 1 indicates full adjustment;
- the error correction term can be either the difference between the dependent and explanatory variable (lagged once) or the error term (lagged once), they are in effect the same thing.

In most of the more recent literature, time series econometrics in general and the estimation of trade elasticities in particular are modelled using cointegration techniques and error correction mechanisms.

3.4.1 Empirical contributions

The late 1990s was the cointegration analysis breakthrough: Bahmani-Oskooee and Niroomand (1998), Bahmani-Oskooee (1998), Bahmani-Oskooee and Brooks (1999), Marquez (1999) and many others employed a (then) so-called long-run method, i.e. a cointegration technique to estimate the long-run trade elasticities.

To establish whether there is a long-run equilibrium relation among the variables of import and export (standard) demand equations, Bahmani-Oskooee and Niroomand (1998) employ Johansen (1988) and Johansen and Juselius (1990) cointegration analysis based on the maximum-likelihood estimation procedure that provides the two tests statistics (λ-max and trace). According to their results, the M-L condition is satisfied for almost all of the 30 countries investigated, indicating that devaluations could improve the trade balance. Using annual data from 1960 to 1992, the long-run price elasticities are reported as follows:

Table 2.3 Export and import price elasticities using Johansen and Juselius (1990) cointegration technique for a sample of countries

Country	*Long-run Export price elasticities*	*Long-run Import price elasticities*
Italy	**-0.24**	-4.81
France	**-6.74**	-0.42
USA	**-1.60**	-0.34
UK	**-0.36**	-0.28

Japan	**-0.49**	-0.97
Germany	**-0.75**	-0.55

Table 2.3 Export and import price elasticities using Johansen and Juselius (1990) cointegration technique. Source: Bahmani-Oskooee and Niroomand (1998).

As it is easy to notice, some of these results are very high (e.g., Italy's import price elasticity and France's export elasticity) and surely need further investigation.

In the same year (1998), Bahmani-Oskooee investigated the trade elasticities in the Less Developed Countries[24] (LDCs) applying the Johansen and Juselius method and obtaining quite the same results: that is, that the M-L condition is met and that devaluations can actually improve the trade balance.

Marquez (1999) conducted a cointegration analysis for the estimation of long period import elasticities for Japan, Canada and the USA covering a secular period (1890-1992). The author claims that the estimates of trade elasticities and, in particular, of import price and income elasticities are very unstable and al the studies carried out in the previous four decades are far from being successful in such sense. According to the author, the dispersion of the estimates is large enough to question whether they are useful in studying international interdependencies. Reliance on century-long sample reveals, however, that, if the assumption of constancy of expenditure shares is avoided, the results can support the view that income and prices affect imports. Obviously, it can be argued that century-long fluctuations are not relevant because estimated elasticities are used just to translate predictions of prices and expenditures into predictions for imports (and, I would add, for exports) and this achieves more usefulness if the observations are more recent. On the other hand, though, "to predict" is not "to understand" and this can partially explain why there is a need for continual re-estimation.

Hooper and Marquez (2000) estimated and tested the stability of import and export elasticities relating the G7 countries to their respective income and prices. The period covered ranges from 1990 to 1996 and the quarterly data include goods and services. For what concerns the "G6" countries investigated in the present research the long-run (Table 2.4) and the short-run (Table 2.5) estimates are:

[24] The countries examined are namely: Greece, Korea, Pakistan, Philippine, Singapore and South Africa.

Table 2.4 Long-run Export and Import elasticities for Italy, France, USA, UK, Japan and Germany				
Country	*Long-run Export price elasticities*	*Long-run Export income elasticities*	*Long-run Import price elasticities*	*Long-run Import income elasticities*
Italy	**-0.9**	1.6	-0.4	1.4
France	**-0.2**	1.5	-0.4	1.6
USA	**-1.5**	0.8	-0.3	1.8
UK	**-1.6**	1.1	-0.6	2.2
Japan	**-1.0**	1.1	-0.3	0.9
Germany	**-0.3**	1.4	-0.06	1.5

Table 2.4 Long-run Export and Import elasticities. Source: Hooper P., Johnson K., Marquez J. (2000).

Table 2.5 Short-run Export and Import elasticities for Italy, France, USA, UK, Japan and Germany				
Country	*Short-run Export price elasticities*	*Short-run Export income elasticities*	*Short-run Import price elasticities*	*Short-run Import income elasticities*
Italy	**-0.3**	2.3	-0.0	1.0
France	**-0.1**	1.8	-0.1	1.7
USA	**-0.3**	1.8	-0.6	2.3
UK	**-0.2**	1.1	-0.0	1.0
Japan	**-0.5**	0.6	-0.1	1.0
Germany	**-0.1**	0.5	-0.2	1.0

Table 2.5 Short-run Export and Import elasticities. Source: Hooper, Johnson., Marquez. (2000).

Thus, as we can see, with exception of Italy, France and Germany (exceptions to the general case), price elasticities for exports and imports satisfy the M-L condition. Among the other main conclusions, according to the authors: trade elasticities are stable enough to help translate economic analyzes into policy recommendations and, for what concerns the USA, a real depreciation of the dollar would keep its external deficit from growing wider.

Algieri (2004) measured export demand elasticities for Russia using an Error Correction mechanism within a cointegration framework. The monthly data collected cover a period from 1993 to 2001. Even though this study does not involve the countries examined in the present research, it is interesting to highlight some of its results. In this case, indeed, Russia's long-run price and income elasticities have the expected sign and are highly significant. In particular, the long-run price elasticity is found to be -2.40 and, according to the author, this can be explained by the low proportion of high-technology goods exported by Russia. Furthermore, it emerges that there is a specialization in products that allow less differentiation and, thus, have to undergo to higher price competition from other countries.

Algieri (2010) conducts an analysis for five big Euro countries (France, Italy, Germany, the Netherlands and Spain) and for their three major competitors (Japan, UK and USA) using quarterly data from 1978 to 2009. The author modeled the export equations using an Unobserved Component (UC) model in order to capture underlying non-price competitiveness. Actually, traditional models specify two key determinants of exports: price competitiveness and foreign demand. Nonetheless, empirical evidence and studies suggest that these factors alone are not able to explain exhaustively export performances. The study, therefore, introduces non-price factors (the UC) of competitiveness. These non-price factors include, among others, aspects such as advanced technology, globalization of production and product quality. The methodology adopted allows to capture underlying changes in export performance giving non-price information and, doing so, it overcomes any misspecification.

4. Final Remarks and Summary Table

When dealing with trade elasticities, exchange rates and global imbalances, one of the main problems to face is definitely the definition of an explanatory framework that researches can agree on. This presumes that the first step is to broadly understand the dynamics that rotate around weak currencies and around the economic relations between the major exporting countries (Thorbecke, 210). The review presented in this chapter is selective rather than exhaustive, concentrating on what I regarded as the most important issues and studies related to the estimation of export price elasticities.

This section provides a summarized examination (Table 2.6) of the existing empirical literature on export equations. It is easy to notice that price elasticities vary across studies and according to the estimation technique adopted:

Table 2.6 Selected Long-run price elasticities: a comparison of studies and results

Author (year)	*Export Price Elasticity*	*Country*	*Estimation period*	*Level of aggregation*	**Model/Approach**
OECD (2010)	-0.60	USA		Exports of goods and services	Standard export equations
	-0.51	Euro Area	Quarterly data		
	-1.00				
Ca' Zorzi and Schnatz (2007)	0.61	Euro Area	1992:Q1 – 2006:Q1	Exports of goods and services	Standard export equations in ECM framework
Di Mauro and Maurin (2005)	0.58	Euro Area	1992-2003	Exports of goods and services	Standard export equations
	0.54	France			
	0.42	Germany			
	0.42	Italy			
OECD (2005)	-0.60	France	1982-2002 (quarterly data)	Exports of goods and services	Standard export equations in ECM framework
	-0.47	Germany			
	-0.60	Italy			
	-0.60	UK			
	-0.60	USA			
	-1.05	Japan			
	-1.5	China			
European Central Bank (2004)	-0.26	Euro Area	1991:Q1 – 2003:Q3	Extra-area exports of goods and services	Standard export equations in ECM framework
Banco de Espaňa (2003)	-0.41	France	1975:Q1 - 2001:Q1	Volume of manufacturing exports	Standard export equations in ECM framework
	-1.08	Germany			
	-0.42	Italy			
OECD (2000)	-0.81	France	1975 -1997 (semi-annual data)		Single equation approach in ECM framework (linear trend or no trends)
	-1.44	Germany			
	-0.98	Italy			
	-1.58	UK			
	-1.41	USA			
	-1.40	Japan			
	-0.68	China			
Senhadji and Montenegro (1999)	*OLS (short-run)*		1960-1993	Exports of goods and non-factor services	OLS and Fully Modified estimates
	-0.01	France			
	-0.07	Italy			
	-0.06	Spain			
	-0.16	UK			
	-0.19	USA			
	-0.25	Japan			
	-0.78	China			

	FM estimates				
	-0.02	France			
	-0.14	Italy			
	-0.18	Spain			
	-0.35	UK			
	-0.73	USA			
	-1.27	Japan			
	-3.13	China			
Caporale and Chui (1999)	-0.08 (DOLS) -0.04 (ARDL)	France	1960-1992 (annual)	Exports of goods and services	(1) DOLS, (2) ARDL.
	-0.11 (DOLS) -0.10 (ARDL)	Germany			
	-0.93 (DOLS) -0.47 (ARDL)	Italy			
	-1.93 (DOLS) -1.22 (ARDL)	Spain			
	-0.19 (DOLS) -0.29 (ARDL)	UK			
	-0.63 (DOLS) -1.36 (ARDL)	USA			
	-1.70 (DOLS) -0.19 (ARDL)	Japan			
Hooper et al. (1998)	-0.2	France	1970-1997 (for France, Germany, Italy); 1960s to 1994Q4 or 1997Q1 for the others	Exports of goods and services	Cointegration vectors and ECM
	-0.3	Germany			
	-0.9	Italy			
	-1.6	UK			
	-1.5	USA			
	-1.0	Japan			
Anderton (1991)	-0.32	Italy	1971:Q2 – 1998Q4	Manufacturing export volumes	Standard export demand
	-0.47	UK			
	-0.65	USA			
	-1.11	Japan			
	-0.27	Germany			

Table 2.6 Selected Long-run price elasticities: a comparison of studies and results.
Source: Author's elaboration on Algieri B. (2010).

Even though according to recent trade empirical[25] and economic geography studies trade price elasticities are supposed to be rather high, ranging from 3 to 11, price elasticity estimations at aggregate levels lead to lower values of around unity. Furthermore, as we have noticed in the literature review, these estimates vary through studies and techniques.

There can be possibly more than just one explanation to this discrepancy but the most likely could be misspecifications in the traditional trade equations as well as measurement errors in export and import prices.

According to the *Comparative analysis of export demand for manufactures in the Euro Area countrie*s conducted by the Banco de Españ̌a (2003), the differing long-term price elasticities seem

[25] See Erkel-Rousse and Mirza (2002) for further details.

to be explained, in general terms, by the productive specialisation of each country. In a context in which manufacturing industry specialisation continues to change dramatically and very unequally across countries, it is reasonable that the sensitivity of exports to price competitiveness should continue to be different in each country. However, it can also be expected that there will be a general trend of elasticities to decrease as countries make progress in improving their competitiveness in ways other than through leadership in costs and in selling prices. In particular, the response of a country's exports to changes in their determining variables largely explains the diversity of results perceived. Such differences will tend to diminish once the processes of commercial integration and productive structure development in the countries that started from a lesser developed level have come to an end.

In an overall perspective, low price elasticities estimates and the current debate on external trade balance adjustment are controversial.

Estimation and re-estimation of trade elasticities is, therefore, not only useful but necessary in order to overcome such drawbacks, to improve the structural modeling process and to keep up to date the trends of trade patterns and of industrial specialization. A continual re-estimation of export (and import) trade elasticities, in addition to a constant development of the empirical modeling techniques, will certainly lead to less puzzling outcomes and can be a solid support in reconsidering the evidence.

CHAPTER 3

ESTIMATING EXPORT ELASTICITIES USING A VECTOR ERROR CORRECTION MODEL

1. INTRODUCTION

Taking into account an imperfect substitute framework and according to the literature, in order to improve the trade balance it is necessary to reduce imports and/or increase exports (ad increase savings) at the same time; to do this, one of the feasible policies could be to work on real exchange rates; many questions can arise: is this kind of policy really effective? How effective is this policy? Is there a boundary value? What is the magnitude of variation supposed to be to have a significant impact on exports? Do exports, at an aggregate level, react differently depending on different factors such as the development status of the traders, the sector or the type of the exported good?

2. THE MARSHALL-LERNER CONDITION AND THE J-CURVE EFFECT

The Marshall-Lerner condition (ML condition, hereafter) is at the heart of the elasticities approach to the balance of payments. The condition seeks to answer the following question:

- (when) does a real devaluation (or a real depreciation)[26] of the currency improve the current account-balance of a country?

The ML condition states that a real devaluation or a real depreciation of the currency will improve the trade balance if the sum of the absolute values of elasticities of the demand for imports and the demand for exports with respect to the real exchange rate is greater than 1:

$$|EM| + |EX| > 1$$

where EM is the demand for imports elasticity and EX is the is the demand for export elasticity.

This condition rests on two fundamental assumptions: the first is that we start from a situation of balanced trade; the second is that the supply elasticities are infinite. This implies that, if

[26] We talk about real devaluation in fixed exchange rates and of real depreciation in floating exchange rates.

the initial situation is a trade deficit, then the ML condition is a necessary but not sufficient stability condition.

Even when the ML condition is met, and improvement ultimately occurs, it may be that at the beginning trade balance deteriorates before it subsequently improves. There is some support in theory for this pattern, known as the J-curve effect.

In theory, the impact of a real exchange rate depreciation on the trade balance is commonly believed to follow a J-curve. According to this view, a currency depreciation improves the trade balance in the long-run but worsens it in the short-run. The initial deterioration in the trade balance occurs because (*i*) currency depreciation increases import prices, while export prices are sticky in the sellers' currency and (*ii*) trade volume tends to respond slowly to a change in relative prices.

To better explain the J-curve phenomenon, it can be said that, at the moment of depreciation, there is a *price effect* due to higher prices of imported goods: since there can be some delays in transactions which have been ordered several months before, the value of imports increases in the short-run.

Later, when traders have had some time to change their strategy, they integrate their loss in competitiveness face to face to goods produced abroad. This produces a *quantity effect*: the volume of imports decreases while local production is probably increased to satisfy demand. In this way, adjustment of quantities traded are slower to adjust than are changes in relative prices. It is expected that the final effect in the long-run is a net improvement in the trade balance.

The phenomenon is named the J-curve effect because when a country's net trade balance is plotted on the vertical axis and time is plotted on the horizontal axis, the response of the trade balance to a devaluation or depreciation looks like the curve of the letter J.

There are numerous empirical studies exploring both whether currency depreciation leads in in the long run to trade balance improvement, and if so, whether a J-curve pattern occurs. These studies investigate different kind of economies such as developed countries, emerging East-European and Asian economies, as well as few developing African countries. Their findings are mixed and, as always, it is up to empirical evidence to support or reject the occurring of the J-curve effect (Pertrović and Gligorić, 2009).

According to the above mentioned theories, one of the policy options to improve the current account is depreciation, which involves the deliberate reduction in the value of a country's currency. This type of policy encourages consumers to alter the distribution of their spending: that is, it is based on an *expenditure switching* process. Expenditure switching, indeed, encourages consumers to switch *away from* imports to domestically produced products and this will lead to a fall in import demand. Contemporaneously, a fall in the exchange rate will, *ceteris paribus*, reduce export prices encouraging export demand.

It is easy to see that the main viewpoints of the debate on the results of an appropriate policy end up to two: those who believe in the positive effects of a currency depreciation on the trade balance and those who don't. If the exchange rate of an economy affects aggregate demand through its effect on export and import prices, policy makers may exploit this connection by deliberately altering exchange rates to influence the macro-economic environment.

Price elasticity estimates are therefore clearly fundamental not only for forecasting purposes and hence, for the implementation of a correct policy, but also for the *a posteriori* evaluation of its effectiveness.

3. EMPIRICAL BACKGROUND SETTINGS

3.1 Introduction

Without doubt, one of the most important issues in empirical research is to design a model which actually represents a certain economic phenomenon, aware of the fact that there is not just *one* right way to model an applied research.

For my purposes, it is necessary to underline that there are many problems that can possibly occur when modelling functions that involve time series economic variables and these issues certainly complicate and make more complex and challenging the process of model building.

One of the most common problems to face when dealing with times series data is the presence of a bilateral causal relationship between two or more variables. Certainly, for reasons that will be clarified in the following paragraphs, the *Vector Error Correction Models* (hereafter, VECM) are frequently applied in examining models that can suffer problems due to endogenous variables.

Indeed, to overcome such problems, in order to carry out the estimation of the long-run export elasticities for the seven countries under scrutiny, I apply a VECM analysis in which the interpretation of the estimates can be naturally classified into short-run and long-run effects.

Before explaining the reasons that are at the basis of this choice and before going into detail with the econometric specification, though, it will be definitely useful to illustrate the underlying empirical framework in which the VECM analysis is set. In order to do so, the next sections will briefly illustrate some of the most popular and used methodologies in the estimation of time series (and, more precisely, of trade elasticities) of the last years.

3.2 Simultaneous equation models

In the recent years, more and more econometricians and economists believe that the use of *vector autoregressive* (VAR) models for macro econometric analysis is one of the most valid alternatives to the common *simultaneous equation models*[27] (or SEMs) that were quite used up to twenty (and more) years ago. One of the reasons is that the simultaneous equations models often did not take into consideration the rich dynamic structure in time series data of quarterly or monthly frequency. Another reason concerns the aforementioned endogeneity issue in so far as the assumptions on the exogeneity of some variables are criticized because not fully supported by theoretical considerations. In contrast, in VAR models all observed variables are typically treated as *a priori* endogenous (Lütkepohl, 2005).

When talking about simultaneous equation models we assume that there is a two-way (or *simultaneous*) flow of influence among the economic variables: that is, one economic variable affects another economic variable and is, in turn, affected by it (Gujarati, 1995). Therefore, in these cases, the distinction between dependent and explanatory variables is not very useful: a more valuable distinction is that between endogenous and exogenous variables. The endogenous variables are those that are dependent one from another whereas the exogenous ones are those that are regarded as the real independent and non-stochastic variables.

This explains why, in this kind of model, the regressand in one equation may appear as a regressor in another equation of the system. Disregarding for endogeneity can lead to serious estimation problems such as biased and inconsistent estimators.

In the simultaneous equation models, there are as many equations as the number of endogenous variables and the parameter of each equation is estimated taking into account

[27] *Simultaneous equation models* (SEMs), also called *Structural equation models*, are multi-equation (multivariate) regression models. In these models, variables may influence one-another reciprocally, either directly or through other variables as intermediaries. These structural equations are meant to represent causal relationships among the variables in the model.

information resulting from the other equations present in the system. It is necessary to specify that the simultaneous equation models treat some variables as endogenous and others like exogenous *a priori* because this remark is a proper link to the VAR analysis.

3.3 VAR models

The Vector Autoregressive models, unlike the SEMs, do not distinguish *a priori* endogenous and exogenous variables because all variables are treated on an equal basis, i.e., all variables are (treated as) endogenous.

The term *autoregressive* is due to the fact that, among the regressors (right-hand side of the equation) there is the lagged value of the dependent variable while the term *vector* is obviously due to the fact that the model contains the vector of two or more variables.

VARs were introduced initially as a replacement of large scale macroeconometric models estimated usually by OLS and Instrumental Variable regressions, entailing of a huge set of equations estimated separately in which the parameters were then subtracted in other equations.

In a VAR[28], consisting of two variables X ,Y the path of Y is explained by the current and past realizations of X (and other variables additively) simultaneously the realizations of X rely on past and current realizations of Y.

The VARS are, therefore, suitably applied when we cannot reject the hypothesis that there exists a bilateral causality among the variables. From a more purely economic point of view, it can be said that these models are especially useful for describing the dynamic behaviour of economic (and financial) time series.

A general 2-variable VAR of order p, that is, a VAR(p) model with two variables (*Y* and *X*) presenting a bilateral causality, can be expressed as follows:

$$Y_t = \alpha + \sum_{j=1}^{p} \beta_j \, Y_{t\text{-}j} + \sum_{j=1}^{p} \gamma_j X_{t\text{-}j} + \varepsilon_{1t}$$

$$X_t = \alpha' + \sum_{j=1}^{p} \theta_j \, Y_{t\text{-}j} + \sum_{j=1}^{p} \lambda_j X_{t\text{-}j} + \varepsilon_{2t}$$

In this model, *p* stands for the number of lags and ε is the stochastic error term. If each equation contains the same number of lagged variables in the system, each equation can be estimated using OLS.

[28] VARs introduction in the literature was done by Sims (1980) in an influential paper '*Macroeconomics & Reality*'.

Some of the advantages of the VAR model are:

- it represents a simple method;
- there is no need to determine which variables are endogenous and which ones exogenous;
- OLS can be applied to each equation separately.

On the other hand, though, the simplicity of this method can represent its drawback. For example, if there are many lags in each equation, it may not be easy to interpret each coefficient.

3.4 VECM Analysis

Over the recent years, the cointegration and the error correction approaches have been studied intensely in the analyses of time series econometrics. In particular, the Vector Error Correction Model[29] (VECM) results very appealing for its distinctive and advantageous characteristics:

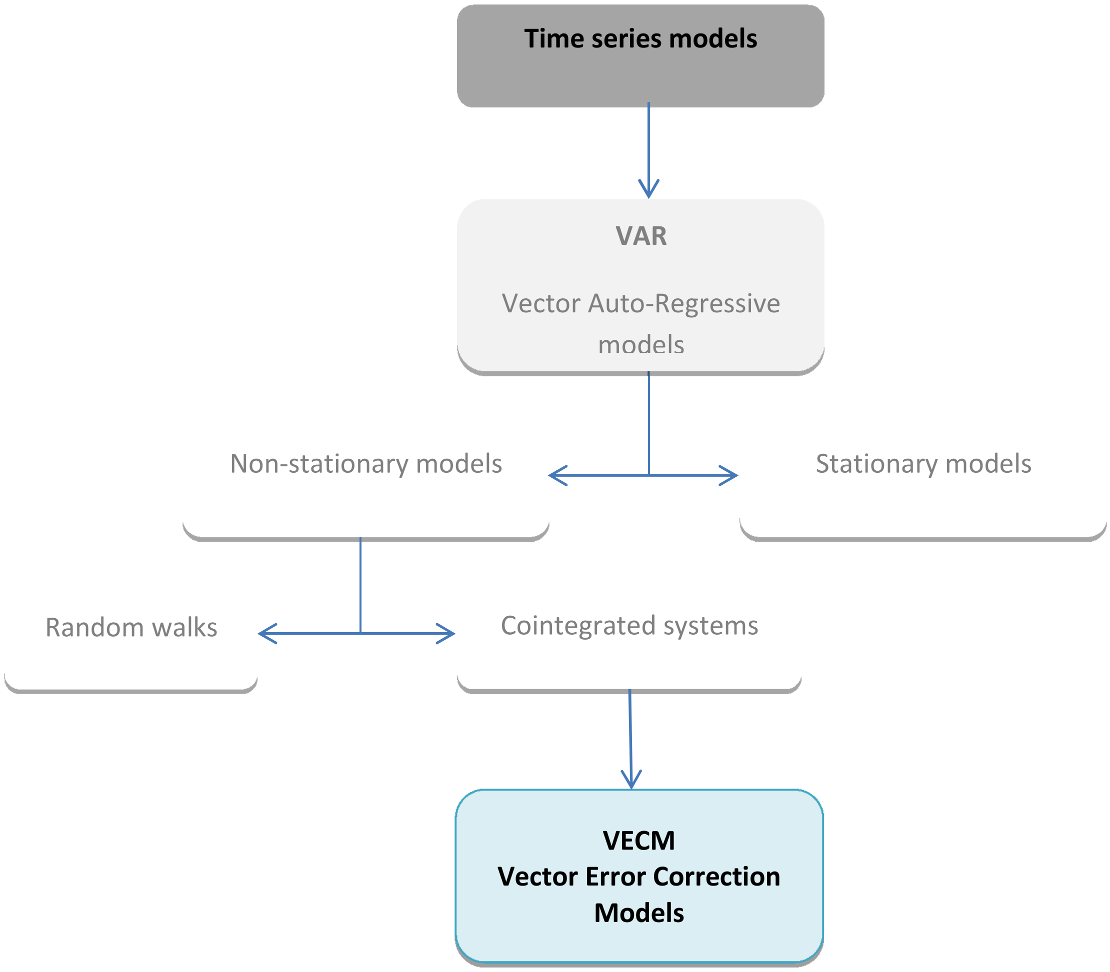
i. first of all, it allows the researcher to represent economic equilibrium relationships within a relatively rich time-series specification;
ii. secondly, it is structured in order to give the possibility to consider all the variables of the model endogenous;
iii. finally, it overcomes the old dichotomy between *(a)* structural models that faithfully represented macroeconomic theory but failed to fit the data, and *(b)* time-series models that were accurately tailored to the data but difficult if not impossible to interpret in economic terms[30].

Considering the VAR models as a starting point, it can be said that a VECM:

- is simply a VAR for variables that are stationary in their differences, i.e., I(1);
- a VECM is a VAR with an error correction term incorporated into the model;
- any VAR(p) model can be re-written as a VECM;
- a VECM is a restricted VAR model: the VECM specification restricts the long-run behaviour of the endogenous variables to converge to their long-run equilibrium relationships and allows the short-run dynamics.

[29] Also known as *Vector Error Correction Mechanism*.
[30] Cottrell A., Lucchetti, R:"J"., (2011), Gnu Regression, Econometrics and Time-series Library, see http://www.gnu.org/licenses/fdl.html, Department of Economics and Dipartimento di Economia, Università Politecnica delle Marche.

Figure 3.1 provides a schematic idea of where these models fit in a time series framework. The



Vector Error Correction Models (VECM) and are shown in the light blue square.

Figure 3.1. VECM scheme

Indeed, once the variables included in the VAR model are found to be cointegrated[31], we will use the VECM: it is important to remember that VECMs are used with non-stationary data and allow the short-term and long-term relationships to be modelled simultaneously *as long as the variables are cointegrated.*

If a non-stationary series Y_t must be differenced d times before it becomes stationary, then it is said to be integrated of order *d*: $Y_t \sim I(d)$. So if $Y_t \sim I(d)$ *then* $\Delta^d Y_t \sim I(0)$, an I(0) series is a stationary series whereas an I(1) series contains one unit root, e.g., $Y_t = Y_{t\text{-}1} + \varepsilon_t$.

31 For discussion on cointegration analysis see Chapter 2, paragraph 2.4.

When the concept of non-stationarity was first considered, a usual response was to independently take the first differences of a series of I(1) variables. The problem with this approach is that pure first difference models have no long-run solution: consider y_t and x_t both I(1). If, for example, we want to estimate the model

$$\Delta y_t = \beta \Delta x_t + \varepsilon_t \qquad (3.1)$$

this collapses to nothing in the long run. The definition of the long run that we use is where

$$y_t = y_{t-1} = y; \;\; x_t = x_{t-1} = x.$$

Hence, all the difference terms will be zero, i.e., $\Delta y_t = 0$; $\Delta x_t = 0$.

One way to avoid this problem is to use *both first difference and levels terms*, e.g.:

$$\Delta y_t = \beta_1 \Delta x_t + \beta_2 (y_{t-1} - \gamma x_{t-1}) + \varepsilon_t \qquad (3.2)$$

where $EC=(y_{t-1}-\gamma x_{t-1})$ is known as the *Error Correction (EC) term*: it is the error from a regression of y_t on x_t The error correction component simply says that Δy_t can be explained by the lagged value of the error correction term itself. Indeed, the EC term measures the speed at which prior deviations from equilibrium are corrected. It can also be thought of as an equilibrium error (or disequilibrium term) occurred in the previous period. If it is non-zero, the model is out of equilibrium and *vice versa*.

Lastly, given that y_t and x_t are cointegrated with cointegrating coefficient γ, then $(y_{t-1}-\gamma x_{t-1})$ will be I(0) even though the constituents are I(1). We can thus validly use OLS to estimate the model.

Any cointegrating relationship can be expressed as an equilibrium correction model. Basically, once showed that the variables under scrutiny are cointegrated, (i.e., there is a long-run equilibrium relationship between them) we can surely think that in the short-run there might be disequilibrium or, in other words, causality can be further sub-divided into long-run and short-run causality. This is the reason why the error term can be treated as the "equilibrium error" (Gujarati, 1995). This error term can be, hence, used to link the short-run behaviour of the variable to its long-run value. This procedure is known as the Error Correction Mechanism (Sargan, 1964) and was made popular by Engle and Granger (1987) in the late 1980s: the mechanism, practically, corrects for disequilibrium and offers the possibility to know the speed to reconcile equilibrium, i.e., it estimates the speed at which a dependent variable *Y* returns to equilibrium after a change in an independent variable *X*. Engle and Granger point out that a linear combination of two or more non-stationary series may be stationary. The stationary combination may be interpreted as the cointegration, or equilibrium relationship between the variables. This explains why the VECMs are very relevant in the modern econometrics: indeed, they represent the tie between times series analyses and economic theory, short-run and long-run.

In this framework, the long-run relations are now often separated from the short-run dynamics. The cointegration or long-run relations are often of particular interest because they can be associated with relations derived from economic theory. It is therefore useful to construct models which explicitly separate the long-run and short-run parts of a stochastic process (Lütkepohl, 2005). VECMs known also as *equilibrium correction models* offer an appropriate structure in this sense.

Error correction models can be used to estimate the following quantities of interest for all *X* variables:

- short term effects of *X* on *Y*
- long term effects of *X* on *Y* (long run multiplier)
- the speed at which *Y* returns to equilibrium after a deviation has occurred.

3.4.1. Advantages of using VECMs

The present analysis applies a cointegration technique based on VECMs to estimate the quarterly demand for exports in some economies present in the international trade scenario. The reasons of such choice relies on the several objective advantages already quoted[32] and on further systematic reasons highlighted as follows.

First of all, the VECM not only is a standard mechanism vastly used and established in the empirical literature but it also likely represents one of the main explanatory models for the theoretical counterpart applied in this study: that is, the imperfect substitutes model. The econometric specification, indeed, is based and is structured on the assumptions of the imperfect substitutes theory.

Secondly, the fact that in this model, relative prices and income enter endogenously into the demand system and that the cointegration rank is not assumed to be known *a priori* but subject to inference leads to significant improvements in the efficiency of the estimates and *suggests that the*

[32] *Cfr*. paragraph 2.3, Chapter 3.

VECM can be viewed as the "natural" econometric reference model for investigating dynamic demand systems[33].

Finally, in order to extend the present analysis in a forecasting perspective and investigate how the macroeconomic outcomes may differ over a certain period if a different policy had been pursued[34], the VECM approach represents the standard approach used in the empirical literature.

For all these reasons, the estimation of long-run and short-run export price elasticities based on VECMs seems to be perfectly suited.

4. A VECM ANALYSIS OF EXPORT ELASTICITIES FOR ITALY, GERMANY, FRANCE, USA, JAPAN, UK AND CHINA

4.1 Outline

The analysis covers a period of over twenty years, from 1990 to 2012, and is conducted for three countries of the Euro Area (EA), namely: Italy, Germany and France; for the three major competitors of the EA: UK, USA and Japan; lastly, for China, that represents the conversation piece of the current debate on devaluation, weak currencies and exchange rate misalignments. Actually, the countries can be classified within two wider groups or areas: the *G6*[35] (Italy, Germany, France, UK, USA and Japan) and China, as a representative of the BRICS[36].

The present section will provide the estimates of export price elasticities *separately* for each country under scrutiny. This decision relies on the fact that it seems necessary to clarify, specifically for each country, the development and the different steps of the econometric model. At the end of the discussion, summarizing tables will provide the results for each country.

In order to streamline the reading, the specific steps/phases of the modelling process will be illustrated in detail for the first country under study, that is, for Italy. A cross-reference to Appendix

[33] Quoted from: "A cointegrated VECM demand system for meat in Italy", L. Fanelli and M Mazzocchi, (2002), Applied Economics, 34, pgg 1593-1605.
[34] *Cfr*. paragraph 4, Chapter 1.
[35] Goldstein A.(2011), BRIC. Brasile , Russia, India, Cina alla guida dell'economia globale, Editore Il Mulino. According to Goldstein, G6 is more appropriate than G7 because Canada is "*sui generis*" for different reasons.
[36] The BRICS countries (Brazil, Russia, India, China and South Africa) are experiencing a very solid economic dynamism on the international context and especially for what concerns China, BRICS came to the spotlight because of the global imbalances issue and because, maybe for the first time in history, the question is not whether you are an industrialized or emerging country but if you are a deficit or a surplus country referring to the trade balance.

1 and to section 5.4 will be made for the corresponding steps/phases of the procedure and results of France, Germany, UK, USA, Japan and China.

When all countries' export price (and income) elasticities are estimated, I will provide an overall view of the short-run and long-run elasticities within summarizing tables, compare them with the findings of previous research (used as benchmark values) and discuss the results.

4.2 Data

The quantitative equations for the countries under investigation have been modeled using quarterly data ranging from January 1990 to January 2012 (1990:1-2012:1) collected from International Financial Statistics (IFS) and Datastream databases. The total number of observations is therefore, 89 for each series. All data have been indexed with base 2005=100. In order to achieve the estimates of export elasticity, the variables used in the econometric model are the natural logarithms of the original data. All the time series data are seasonally adjusted.

The econometric software package used throughout the VECM analysis is *Gretl*, version 1.9.5.

4.3 Variables

Having elicited the general outlines of the imperfect substitutes model (*cfr.* §2.1.2), I will now consider the empirical variables that have been used as the appropriate counterparts to the theoretical ones.

For what concerns the choice of the variables for the construction of the model, according to the literature[37], trade equations are specified with the value of exports as the dependent variable (Y).

The explanatory variables chosen are those used in the prevalent existing literature and comprise real exchange rates and the rest of the world's income. For what concerns the use of the real exchange rates, this choice is supported by the fact that, as commonly assumed in the empirical literature[38], there is the basic assumption that there exists a complete pass-through[39] between relative prices and exchange rates.

---

[37] Goldstein, M., Khan, M. S. (1985), Income and price effects in foreign trade, in: Jones R. W., Kenen P. B. Handbook of International Economics, Amsterdam, North Holland (1985), pp. 1042-1099.
[38] Goldstein, M., Khan, M. S. (1985), *op.cit.*.

Given all these issues, the chosen dependent (*Y*) variable is the value of country's exports and the explanatory (*X*) variables are the real exchange rates and the world income; a sketch of all the information on the data is presented in Table 3.1. The variables have been transformed in log form (*l*) in order to express the resultant coefficients as elasticities of the variable included in the model.

Table 3.1. List of variables and data information for Italy, France, Germany, UK, USA, Japan and China

	Country	Variable		Definition and Concept	Source	Unit
1	Italy	*IT_Rex*	*X*	Real effective exchange rate, Consumer Price Index based	Datastream	Index, 2005=100
		IT_Rex_ULC	*X*	Real effective exchange rate, Unit Labor Cost Index based	Datastream	Index, 2005=100
		IT_E	*Y*	Exports of goods & services (real, US$)	Datastream	Index, 2005=100
2	France	*FR_Rex*	*X*	Real effective exchange rate, Consumer Price Index based	Datastream	Index, 2005=100
		FR_Rex_ULC	*X*	Real effective exchange rate, Unit Labor Cost Index based	Datastream	Index, 2005=100
		FR_E	*Y*	Exports of goods & services (real, US$)	Datastream	Index, 2005=100
3	Germany	*GE_Rex*	*X*	Real effective exchange rate, Consumer Price Index based	Datastream	Index, 2005=100
		GE_Rex_ULC	*X*	Real effective exchange rate, Unit Labor Cost Index based	Datastream	Index, 2005=100
		GE_E	*Y*	Exports of goods & services (real, US$)	Datastream	Index, 2005=100
4	UK	*UK_Rex*	*X*	Real effective exchange rate, Consumer Price Index based	Datastream	Index, 2005=100
		UK_Rex_ULC	*X*	Real effective exchange rate, Unit Labor Cost Index based	Datastream	Index, 2005=100
		UK_E	*Y*	Exports of goods & services (real, US$)	Datastream	Index, 2005=100
5	USA	*US_Rex*	*X*	Real effective exchange rate, Consumer Price Index based	International Financial Statistics (IFS)	Index, 2005=100
		US_Rex_ULC	*X*	Real effective exchange rate, Unit Labor Cost Index based	International Financial Statistics (IFS)	Index, 2005=100
		US_E	*Y*	Exports of goods & services (real, US$)	Datastream	Index, 2005=100
6	Japan	*JP_Rex*	*X*	Real effective exchange rate, Consumer Price Index based	Datastream	Index, 2005=100

[39] For further details, see Chapter 2, section 2.1.5.

		JP_Rex_ULC	*X*	Real effective exchange rate, Unit Labor Cost Index based	Datastream	Index, 2005=100
		JP_E	*Y*	Exports of goods & services (real, US$)	Datastream	Index, 2005=100
7	China	*CH_Rex*	*X*	Real effective exchange rate, Consumer Price Index based	Datastream	Index, 2005=100
		CH_E	*Y*	Exports of goods & services (real, US$)	Datastream	Index, 2005=100
8	World	*y*	*X*	Income (real GDP, US$)	Datastream	Index, 2005=100

Table 3.1 List of variables and data information for Italy, France, Germany, UK, USA, Japan and China

5. ESTIMATES

5.1 Introduction

In the same way as most of the main European countries, since 1998, Italy is experiencing serious macroeconomic imbalances that need to be addressed. In particular, macroeconomic developments in the area of export performance deserve attention as, Italy has been losing external competitiveness in the last decade, due to both cost and non-costs factors, and has been hit hard by the financial crisis. Given the high level of public debt, improving the growth potential and competitiveness should be key priorities. Mainly due to exposure to competition by emerging countries, Italy's share in world export markets in sectors in which it specialises declined quite considerably in the 2000s[40.] A probably unfavourable product specialisation and geographical destination of exports also explain decreasing competitiveness. With export products that are rather similar to those of emerging economies, Italy has been exposed more than other Euro Area countries to increasing global competition. Italy's exports are also disadvantaged by their still relatively low penetration into fast-growing emerging markets, especially in Eastern Asia. The relatively small size of the Italian firms also probably plays a key role in hampering the reorientation of exports towards new and distant markets.

[40] European Commission (2012).

Figure 3.1 Italian exports, 1998:Q1-2012:Q3

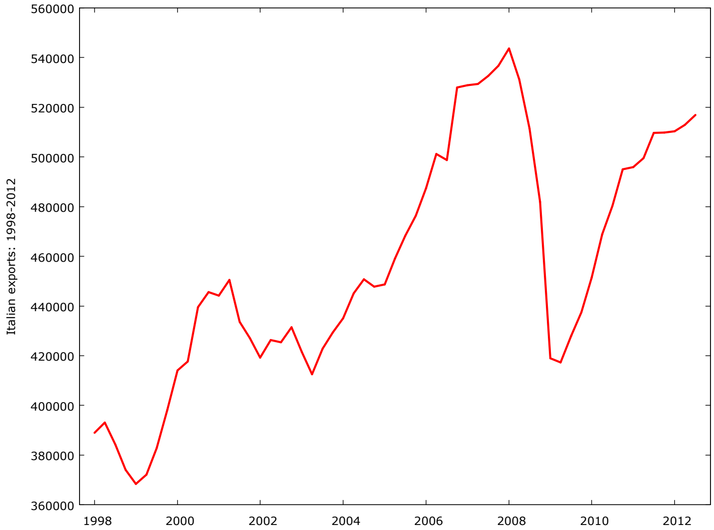


Figure 3.1 Italian exports, 1998:1-2012:3.
Source: Author's own elaboration via Datastream database.

5.2 Export elasticity estimates for Italy

Once collected the data, the first critical step of the analysis is to visually inspect it: indeed, the variables' dynamics can be *informally* investigated through the visual plots and through the autocorrelation function (ACF) correlogram.

The information provided at this first informal level can give a first suggestion of what kind of issue must be addressed: in particular, it can give an indication whether the time series is stationary or not. For example, figure 3.4 shows the ACF correlogram of the Italian exports (expressed in logarithm) l_IT_E up to 30 lags[41]. What we can see is that the ACF correlogram starts at a very high value (0,9662 at lag 1) and becomes smaller very gradually. Even at lag 13, the autocorrelation coefficient measures still about 0,5. This kind of pattern is generally an indication that the time series is non-stationary (Gujarati, 1995). Autocorrelation would be zero at any lag greater than zero if the stochastic process is purely random. The confidence interval ($\pm$ 1,96) is delimitated by the solid blue lines. The same consideration suits the ACF correlograms of Germany, France, USA, Japan, UK and China (see Appendix 1.A).

[41] In practice, the maximum length of lags to be use is up to one-third of the sample size.

Figure 3.4 ACF of exports for Italy

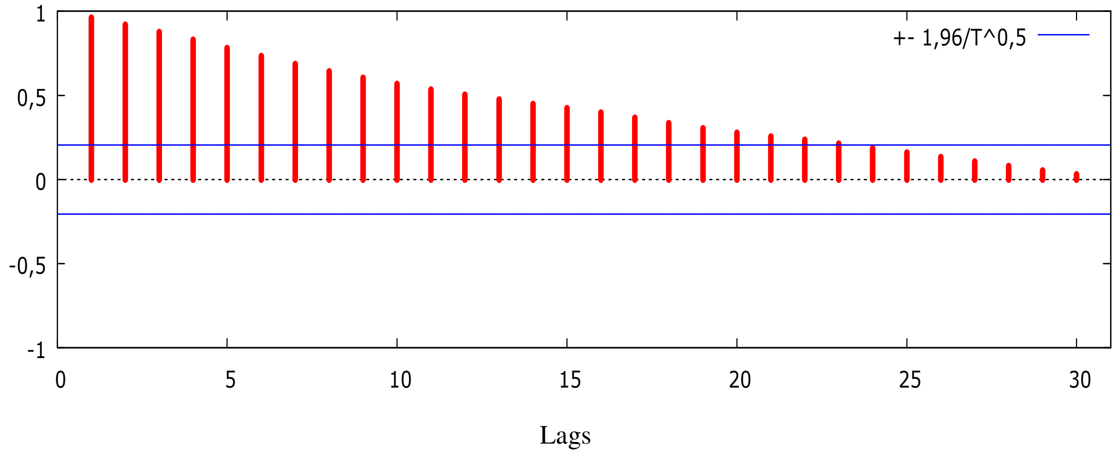


Figure 3.4. ACF correlogram of l_IT_E, 1990:1-2012:1. Source: Author's elaboration.

In the same way, a visual plot of the data (here expressed in logs) is usually one of the first steps in the analysis of any time series (figure 3.5 and 3.6):

Figure 3.5

Figure 3.5 CPI-based and ULC-based Exchange rate, exports and world income plots for Italy

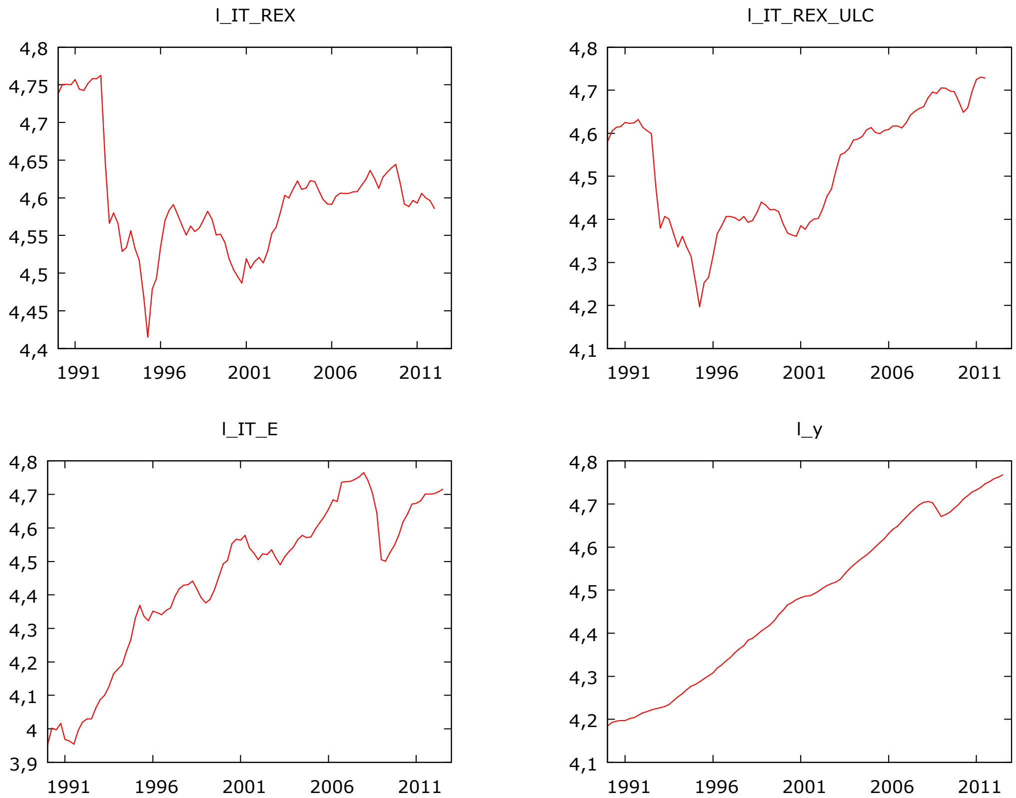


Figure 3.6 CPI-based and ULC-based exchange rate, exports and world income jointed plot for Italy

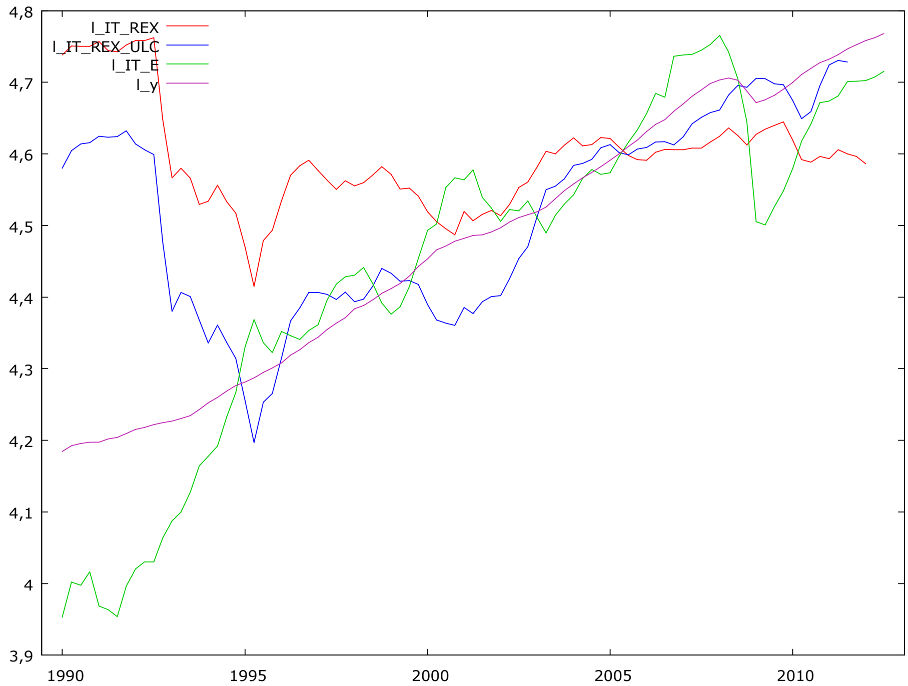


Figure 3.6

Summary statistics are described as follows:

Table 3.3. Descriptive statistics of variables for Italy

Variable	Mean	Median	Minimum	Maximum	Std. Dev.	Var. coeff.	Asymmetry	Curtosis
l_IT_E	4,44171	4,50553	3,9527	4,76552	0,23398	0,05268	-0,6981	-0,569
l_IT_REX	4,59489	4,59178	4,4151	4,76226	0,07352	0,016	0,69256	0,51748
l_IT_REX_ULC	4,51389	4,55472	4,19705	4,73057	0,13832	0,03064	-0,1873	-1,2213

Table 3.3. Descriptive statistics, Sample period: 1990:1 - 2012:3. Source: Author's own elaboration on Datastream and IFS Databases.

Since at a first visual and informal level[42] each series seems to be non-stationary, I proceed to a formal level of investigation and implement a unit root test for each variable.

5.2.1 Unit Root Test of stationarity

To test whether the series are non-stationary I implement the Augmented Dickey-Fuller (ADF) test (1981). Under the null hypothesis (H_0) there is the presence of a unit root. The statistic-test is the tau-test (τ) for which the critical values are those tabulated by MacKinnon through Monte

[42] See Appendix 1. for the visual plots of the time series variables.

Carlo simulations[43]. *Gretl* software package prints out the p-value based on MacKinnon's approximation to the distribution of the τ - statistics: the p-value is, therefore, the value on which I figure the decision of rejecting or accepting the null hypothesis of the presence of a unit root.

The ADF tests in which lags of Δ*y*, are added to avoid the problem of serial correlation of the residuals[44]: indeed, the τ -statistics, computed as ordinary t-statistics, remain *asymptotically valid* in the presence of serial correlation when this is done.

Running the ADF test in levels, Italy's exports, the exchange rate based on the Consumer Price Index and the exchange rate based on the Unit Labor Cost Index series are shown to be all *I(1),* i.e., integrated of order 1. The tests were executed for three different cases: (i) τ_c, with a constant, (ii) τ_{ct} with a constant and trend and (iii) τ_{nc} without a constant. The results are sensitive to the case applied because the critical values change in the three cases.

Afterwards, the test was repeated using the first differences of each series: the series, differenced only once are shown to be stationary (the null hypothesis of a unit root can be rejected) so it can be affirmed that the original time series (the one in levels) are integrated of order 1, *I(1).* The ADF test implemented on original time series (*l_IT_E, l_IT_REX* and *l_IT_REX_ULC*) for Italy in levels and in first differences are summarized in Table 3.4[45]:

Table 3.4. ADF Unit root tests: comparative settings for Italy

Variable	Variant	τ –statistic ADF level	p-value*	τ –statistic ADF first difference	p-value*
l_IT_E	Constant, no trend τ_c	-1,44672	0,5608	-6,24789	0,00000
	Constant and trend τ_{ct}	-2,0832	0,5547	-6,25815	0,00000
	No constant τ_{nc}	1,57885	0,9725	-5,96656	0,00000
l_IT_REX	Constant, no trend τ_c	-2,69871	0,07419	-4,15522	0,00078
	Constant and trend τ_{ct}	-3,39395	0,05212	-3,62381	0,02779
	No constant τ_{nc}	-0,11459	0,6443	-4,2796	0,00000

[43] MacKinnon J. G. (2010).

[44] "The 'augmented' Dickey-Fuller, or ADF, tests, in which lags of Δy, are added […] so as to whiten the residuals". MacKinnon, *op cit.*

[45] The results of the ADF tests for the other countries under investigation are given in Appendix 2.A.

l_IT_REX_ULC	Constant, no trend τ_c	-0,83832	0,8077	-3,81311	0,00279
	Constant and trend τ_{ct}	-3,08087	0,1108	-3,6079	0,02909
	No constant τ_{nc}	0,32101	0,7783	-3,58096	0,00034

Table 3.4. ADF Unit root tests: comparative settings. MacKinnon (1996) critical values for the null hypothesis H_0 = presence of a unit root. Asymptotic p-values.*

As we can see, the variables in levels are non-stationary so the null hypothesis (H_0 = there is a unit root) cannot be rejected for each series and for each variant of the test (τ_c, τ_{ct}, τ_{nc}). In order to achieve the stationarity of the series, I use the first differences of this variable and then test them for the presence of a unit root. The asymptotic p-values of the tests implemented in the three variants show that we can now reject the null hypothesis of a presence of a unit root. Once non-stationarity is established, a long-run equilibrium relationship can be reasonably expected. For this reason, the next step is to gauge the cointegration relationship. The estimation of the cointegration relationship requires the execution of a test to determine the number of cointegrating vectors present in the system, in other words, the *cointegrating rank* of the system. To determine the number of cointegrated vectors the Johansen approach is followed.

5.2.2 Johansen Cointegration Tests

This is an important phase since it provides the required information to subsequently implement the VECM analysis with, at least, the appropriate rank. The two Johansen tests for cointegration are used to establish the rank are the "λ-max" test, for hypotheses on individual eigenvalues, and the "trace" test, for joint hypotheses.

Two test statistics are, thus, used to test the number of cointegrating vectors, based on the characteristic roots. For both the null (H_0) is: *at most* r cointegrating vectors.

The trace statistics:

$$\lambda_{trace}(r) = -T\sum_{i=r+1}^{k} \ln(1-\hat{\lambda}_i)$$

where the alternative (H_1) is: at most k cointegrating vectors.

It looks at the trace of A(1) = the sum of eigenvalues. If there is no cointegration, then all 1-$\hat{\lambda}_i$ are zero and trace of A(1) = 0. The test is ran in sequence: start from the null of *at most* 0 cointegrating vectors up to *at most* k cointegrating vectors against the alternative.

Lambda-max statistics:

$$\lambda_{max}(r, r+1) = -T\ln(1-\hat{\lambda}_{r+1})$$

where the alternative (H_1) is: at most *r+1* cointegrating vectors. It tests rank *r+1* by testing if $\hat{\lambda}_{r+1}$ is zero.

Italy's model presents three variables so the tests can be specified as follows:

Rank	*Trace test*		*λ-max test*	
	H_0	H_1	H_0	H_1
0	c = 0	c = 3	c = 0	c = 1
1	c = 1	c = 3	c = 1	c = 2
2	c = 2	c = 3	c = 2	c = 3

where c = cointegrating vectors.

Neither of these test statistics follows a chi-square distribution in general; asymptotic critical values are those tabulated with Doornik's (Doornik, 1998) gamma approximation distribution.

When r = 0 there are no cointegrating vectors. If there are k variables in the system of equations, there can be a maximum of k-1 cointegrating vectors.

The results for the model that includes l_IT_REX[46] are:

Rank	Eigenvalue	Trace test	p-value	Lmax test	p-value
0	0,21118	30,560	[0,0405]	20,164	[0,0672]
1	0,11414	10,397	[0,2559]	10,302	[0,1963]
2	0,0011112	0,094509	[0,7585]	0,094509	[0,7585]

Table 3.5: Johansen Tests l_IT_REX

The values are examined one row at a time starting from the first row and, as it can be seen, the trace test has a p-value lower than the 5% so we can reject the null of r = 0 (even though the *λ*-

[46] See Appendix 2.B and 2.C for the Johansen cointegration tests for Germany, France, USA, Japan UK and China.

max is not that straightforward[47]): this means we can stop and consider r = 1. Indeed, looking at the p-value related to the null: r =1, this cannot be rejected. The same can be said for the second model that comprises l_IT_REX_ULC:

Rank	Eigenvalue	Trace test	p-value	Lmax test	p-value
0	0,21251	29,826	[0,0497]	19,829	[0,0751]
1	0,10448	9,9974	[0,2861]	9,1588	[0,2795]
2	0,010052	0,83853	[0,3598]	0,83853	[0,3598]

Table 3.6: Johansen Tests for l_IT_REX_ULC

5.2.3 Estimates

Once ascertained that the variables are integrated of order I (1), the VECM analysis can be executed. For each country there are two different models respectively for the two types of real exchange rates: one based on the Consumer Price Index (CPI), model A, and the other on the Unit Labor Cost (ULC), model B.

A) l_IT_E = f (l_IT_REX, l_y)

Table 3.7. Export elasticity estimates using VECM for Italy. CPI-based real exchange rate.

Country	*Long-run Price elasticity*	*Short-run Price elasticity*	*Long-run Income elasticity*	*Short-run Income elasticity*	*ECM Speed of Adj.*
Italy	**-0,72**	**-0,05**	1,01	0,70	**-0,25**
s.e.	*-0,141100*	*0,123866*	*-0,053981*	*0,815929*	*0,056667*

Table 3.7:VECM system, 4 lags. Obs.: 1990:1-2012:1 (T = 85); Cointegration rank =1; Exchange rates on Consumer Price Index bases. Source: Own estimations on Datastream and IFS databases. Notes: Aggregation level: value of goods and services; Index 2005=100.

These results[48] indicate that the long-run export price and income elasticities estimates are, respectively: -0,72 and +1,01.

The short-run export price and income elasticities estimates are, respectively: -0,05 and +0,7.

[47] If results of the two test statistics are not consistent the suggestion is to use the trace statistics: the trace test, indeed, is likely to pick up the correct value of r and have good power. For further details, see K. Juselius lecture notes, http://www.econ.ku.dk/okokj.

[48] See Appendix 3 for detailed estimation outputs related to all the countries under investigation.

The Error Correction term coefficient, which is expected to be negative and to lie between 0 and 1, is - 0,25; it is statistically significant and exhibits the expected negative sign. As aforesaid, it indicates the speed at which the variables return to equilibrium after departing for the equilibrium path (after a shock, for example).

The Durbin-Watson test is: 1,96 while the Adjusted R2 is 0,45.

The following graph plots the residuals of the system jointly:

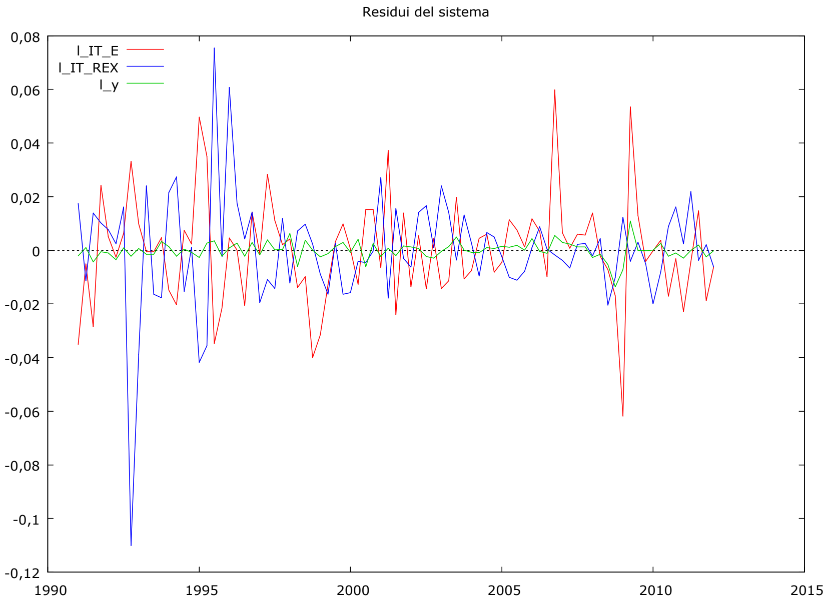


Figure 3.7. Source: Author's own elaboration.

The statistic used to test the presence of serial correlation is the Ljung-Box Q-statistic. In general the test on serial correlation using the Q-statistic has the null hypothesis of "no serial correlations" (up to the lags used for the test, which here are 4):

Serial correlation diagnostic test:

Equation 1:
Ljung-Box Q' = 0,407565 with p-value = 0,982

Equation 2:
Ljung-Box Q' = 4,27906 con p-value = 0,37

Equation 3:
Ljung-Box Q' = 0,10923 con p-value = 0,999

Hence, the test indicates that there is no serial correlation since you cannot reject the null.

Italy presents a long-run price elasticity of -0,72 (Table 3.7), which confirms the general findings. It is interesting to notice that the long-run export price elasticity estimates provided in the present research for Italy, using two different techniques[49] but the same sample period and the same econometric specification, are almost exactly the same (-0,71)[50]. The main difference concerns the dependent variable: in the first case, I use the value of exports as the dependent variable which includes both price and quantity variations; in the second case, exports are expressed in volumes and, hence, entail only quantity variations.

B) l_IT_E = f (l_IT_REX_ULC, l_y)

Table 3.8. Export elasticity estimates using VECM for Italy. ULC-based real exchange rate.

Country	*Long-run Price elasticity*	*Short-run Price elasticity*	*Long-run Income elasticity*	*Short-run Income elasticity*	*ECM Speed of Adj.*
Italy	**-0,48**	**-0,12**	**1,32**	**0,60**	**-0,28**
s.e.	*-0,083110*	*0,111288*	*-0,065614*	*0,810375*	*0,065019*

Table 3.8:VECM system, 4 lags. Obs.: 1990:1-2012:1(T = 85); Cointegration rank =1; Exchange rates on Unit Labor Cost index bases. Source: Own estimations on Datastream and IFS databases. Notes: Aggregation level: value of goods and services; Index 2005=100.

These results show that the long-run export price and income elasticities estimates are, respectively: -0,48 and + 1,32.

The short-run export price and income elasticities estimates are, respectively: -0,12 and +0,6.

The Error Correction term coefficient is -0,28; it is statistically significant and exhibits the expected negative sign.

The Durbin-Watson test is: 1,95 while the Adjusted R2 is 0,46.

The evidence is that the short-run price elasticities are noticeably smaller than the long-run price elasticities in both the cases considered: indeed, in general, the long-run elasticities are roughly twice as high as the short-run elasticities. In the first model, though, the difference between long-run and short-run estimates is much greater.

[49] Distributed-Lag model (*cfr*. Appendix 4) and VECM (cfr.§5, Chapter 3).
[50] *Cfr*. Chapter 3, section 4.2.

5.3. Export elasticity estimates for Germany, France, USA, Japan, UK and China.

In this section I present the estimation results for all the seven countries under scrutiny: having estimated the countries' export price[51] elasticities separately, I provide an overall view of the short-run and long-run elasticities within two summarizing tables. In the following sections, I compare and discuss the results both with the findings of previous studies and with my findings. Reiterating the VECM equations[52] for Germany, France, USA, Japan, UK and China, I obtain the following results:

Table 3.9. Export elasticity estimates using VECM for Italy, Germany, France, USA, Japan, UK and China. CPI-based real exchange rate.

Country	*Long-run Price elasticity*	*Short-run Price elasticity*	*Long-run Income elasticity*	*Short-run Income elasticity*	*ECM Speed of Adj.*
Italy	**-0,72**	**-0,05**	**1,01**	**3,86**	**-0,25**
s.e.	*0,141100*	*0,123866*	*0,053981*	*0,815929*	*0,056667*
Germany	**-0,58**	**-0,25**	**2,56**	**4.32**	**-0,11**
s.e.	*0,26930*	*0,135190*	*0,15096*	*1,10649*	*0,093337*
France	**-1,41**	**-0,27**	**0,10**	**3,63**	**-0,12**
s.e.	*0,407000*	*0,153284*	*0,100790*	*0,681037*	*0,027807*
USA	**-1,21**	**-0,03**	**1,38**	**2,63**	**-0,23**
s.e.	*0,190390*	*0,140229*	*0,068277*	*0,980787*	*0,034103*
Japan	**-0,55**	**-0,23**	**1,34**	**1,66**	**-0,28**
s.e.	*0,123420*	*0,086874*	*0,082945*	*1,319900*	*0,076372*
UK	**-0,84**	**-0,07**	**1,60**	**0,02**	***0,03***
s.e.	*0,233410*	*0,140872*	*0,118090*	*1,392200*	*0,044925*
China	**-1,95**	**-0,27**	**5,58**	**1,87**	**-0,27**
s.e.	*0,307520*	*0,254649*	*0,140780*	*0,229376*	*0,114734*

Table 3.9: VECM system, 4 lags. Obs.: 1990:1-2012:1(T= 85); Cointegration rank =1; Exchange rates on Consumer Price Index bases. Source: Own estimations on Datastream and IFS databases. Notes: Aggregation level: value of goods and services; Index 2005=100.

[51] The income elasticities have been also reported for a broad comprehension of the whole issue.

[52] *l_Country_E = f (l_Country_REX, l_y)* and *l_Country_E = f (l_Country_REX_ULC, l_y).*

Table 3.10. Export elasticity estimates using VECM for Italy, Germany, France, USA, Japan, UK and China. ULC-based real exchange rate.

Country	*Long-run Price elasticity*	*Short-run Price elasticity*	*Long-run Income elasticity*	*Short-run Income elasticity*	*ECM Speed of Adj.*
Italy	**-0,48**	**-0,12**	**1,32**	**0,60**	**-0,28**
s.e.	*0,083110*	*0,111288*	*0,065614*	*0,810375*	*0,065019*
Germany	**-0,21**	**-0,17**	**2,06**	**3,06**	**-0,19**
s.e.	*0,106450*	*0,116483*	*0,074202*	*0,828905*	*0,063510*
Francia	**-1,20**	**-0,32**	**1,06**	**3,21**	**-0,23**
s.e.	*0,152520*	*0,114040*	*0,071855*	*0,602382*	*0,042725*
USA	**-0,57**	**0,08**	**1,26**	**2,17**	**-0,34**
s.e.	*0,108390*	*0,119547*	*1,259200*	*1,009050*	*0,048355*
Japan	**-0,37**	**-0,06**	**1,39**	**-0,32**	**-0,34**
s.e.	*0,105960*	*0,084542*	*0,099315*	*1,897500*	*0,087257*
UK	**-0,26**	**-0,09**	**1,64**	**0,73**	**-0,08**
s.e.	*0,099019*	*0,139602*	*0,084836*	*1,605630*	*0,094353*
China	-	-	-	-	-
s.e.	-	-	-	-	-

Table 3.10: VECM system, 4 lags. Obs.: 1990:1-2012:1(T= 85); Cointegration rank =1; Exchange rates on Unit Labor Cost index bases. Source: Own estimations on Datastream and IFS databases. Notes: Aggregation level: value of goods and services; Index 2005=100.

6. DISCUSSIONS OF THE RESULTS

The VECM results reveal both correspondences (with the estimates of the existing literature, the price elasticities ranging from -0.21 in the case of Germany to -1.95 in the case of China) and discrepancies with the current debate on weak currencies and with the concrete situation of some of the countries examined like Germany:

Figure 3.8. German exports to Asia, EU and USA.

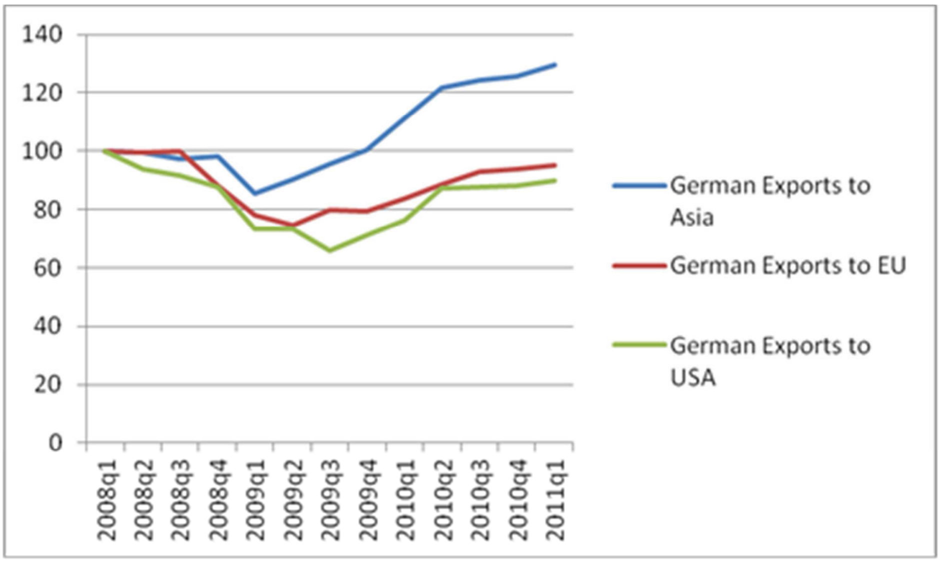


Figure 3.8 Germany: exports to Asia, EU and USA. Index, 2008Q1 = 100
Note: Growth refers to that of real GDP. Exports are of goods and for 2011Q1 refer to January and February 2011.
Source: Deutsche Bundesbank; OECD, National Accounts database, author's elaboration.

In general, it is interesting to notice, indeed, that the results report long-run export price elasticity estimates lower than unity in most cases (Table 3.9). In particular, Germany (-0,11), UK (-0,84) and Japan (-0,55), three of the major exporting countries, present very low values that do not justify policies applied on exchange rates to promote growth through trade balances' surplus and, furthermore, diverge from the line of reasoning of the currency war issues in which they are involved. For what concerns the estimates reported for USA (-1,21) and China (-1,95), they are certainly more plausible and consistent with some results of previous studies but still far from the current debate.

These considerations assume even more importance when looking at the estimates provided using the exchange rate on Unit Labor Cost basis (Table 3.10): all countries, indeed, present very low long-run price elasticity estimates that unlikely can meet any currency depreciation policy effectiveness.

The aggregation level can possibly be a cause of these low estimates and probably higher disaggregation would lead to more reasonable estimates that could explain how effective devaluation policies are.

The short-run elasticity estimates confirm the general theory as they are smaller than the long-run estimates and this is mainly due to the adaptation period[53].

One of the reasons why the VECM was adopted was because it comprised the possibility to identify the speed of adjustment to the already mentioned "*equilibrium path*". The last columns of Tables 3.9 and 3.10 reveal *how much* of the error has been corrected and, at the same time, *how long* will it take to reach the equilibrium value. For example, Italy's error correction term, -0,25 (Table 3.9), indicates that 25% of the error has been corrected and that it will take other four periods to reach the equilibrium value: since quarterly data has been used, this means that the adjustment will be pursued within one year[54]. The highest speed of adjustment is registered by Japan (-0,28) while the country that will need more time to reach equilibrium is France (-0,12). All values are significant and of the expected sign with exception of UK (+0,03).

Finally, with the exception of France (Table 3.9), the long-run income elasticity estimates are, within limits, higher than the long-run price elasticity ones meaning that export demand is, according to these outcomes, more income elastic than price elastic. A high income elasticity of export demand means that an increase in world income will increase export demand of a country substantially: c*eteris paribus*, this will improve the balance of trade.

6.1 Evidence for the Marshall-Lerner condition

International trade estimates are an important contribution for any analysis of the aggregate effects of changes in income and relative prices but, obviously, they do not guarantee that a particular result will in fact occur in response to the above mentioned changes. It is also clear that the knowledge of elasticity magnitudes is important to deal with the unavoidable (and, to a certain degree, predictable) changes they produce on a country's trade balance and level of income and employment.

[53] *Cfr*. Chapter 2.
[54] 1/0,25 = 4. 4 *(3 months) = 1 year.

Trade elasticities could be (and actually are) used by policy-makers to estimate the exchange rate variation that would be required to eliminate or reduce trade balance deficits and, in the same way, to take decisions on currency depreciation or appreciation[55].

This analysis presented in this chapter focuses on the long-run price elasticities, using a cointegration VECM to estimate export functions. One of the important results to be achieved was to uncover evidence for the ML condition, that is to verify if a country (e.g. China), weakening its currency, could accomplish a competitive depreciation, actually improving its trading position. Reminding that the condition states that the elasticities (in absolute values) of exports and imports must sum to greater than one for a depreciation to be effective, it is necessary to test the condition empirically. When price elasticities are obtained for both exports and imports along with the corresponding standard errors the following test can be executed:

$$t = \frac{|Ex| + |Em| - 1}{\sqrt{\hat{\sigma}_x{}^2 + \hat{\sigma}_m{}^2}}$$

where *Ex* is the is the export elasticity and *Em* is the imports elasticity, and $\hat{\sigma}$ is the corresponding standard error.

Since the present study estimates a single elasticity, it does not fulfil the criteria for the formal ML condition test illustrated above[56]. Nevertheless, it is still possible to test the ML condition through this type of specification[57]: indeed, the condition is met if the single elasticity exceeds one. For example, Felipe *et al.* (2010) use only exports to claim that China's exports hold the ML condition. Therefore, using the single export elasticity estimates provided by the VECM, it is possible to identify the countries that can surely rely on weak currencies to improve their trade balance in terms of exports. Considering the VECM estimates using exchange rates on Consumer Price Index bases:

- the estimates of China (-1.95), France (-1.41) and USA (-1.21) undoubtedly meet the Marshall-Lerner condition so it certain that the exchange rate has a significant impact on these three countries; indeed, even in the lack of a formal test, the condition holds being the single elasticity (in absolute value) already major than one;

[55] De Vanssay X. (2003).
[56] The ML condition will be formally tested in Chapter 4 using panel data estimates.
[57] Bahmani *et al.* (2013).

- UK (-0.84) and Italy (-0.72) give more ambiguous results: in this case, the import price elasticities are required in order to state that that the condition is met;
- Germany (-0.58) and Japan (-0.55) give low results (relatively to the boundary value of one) but, even though it seems improbable that as for UK and Italy, these results are ambiguous.

It is interesting to notice that China is the country that reports the highest price elasticity value, confirming the widespread fear that the country is operating a “competitive devaluation” in order to enhance its exports.

When considering the VECM estimates using exchange rates on Unit Labor Cost index bases, the scenario changes substantially: indeed, the only country that can claim to meet the Marshall-Lerner condition is France (-1.20) while the remaining six countries give very lower results: in particular, Japan (-0.37) and UK (-0.26) are the countries that, most likely, cannot claim to benefit from a devaluation or depreciation. In support to the part of the literature that argues that undervaluation facilitates growth among developing countries and stresses the role of relative prices as instruments of industrial policy in the process of economic convergence we find that exports are price elastic in China, France and USA: hence, competitive undervaluation may trigger growth (Rodrik, 2008). Nevertheless, considering the other four countries of the sample, it can be said that, effects of exchange rate policies on exports seem to be limited. In order to obtain a sustainable and stabilized export growth, trade policies, which are based on diversification of exported products and production of technology-intensive goods, have to be developed.

6.2 Final remarks

In the present debate on Asian currencies (appreciation against depreciation policies), it can be said that the overall effects of devaluation or depreciation are mixed. In my study, three countries representing the almost 43% of the sample can actually claim that the Marshall-Lerner condition is met. This result is consistent with that of a recent survey (Bahmani *et al.*, 2013) on the empirical tests of the Marshall-Lerner condition in literature: according to the survey, the Marshall-Lerner condition is met in about 62% of the cases even though, the sum of the absolute values of the point estimates is significantly greater than one in the 30% of the total.

The controversial debate on global trade imbalances (especially between the EA and China and between the USA and China), anyway, underlines the role of exchange rates and of exchange rate misalignments that are perceived as the origin of a series of economic disorders both on domestic and on global basis. These problems are complex and involve a variety of issues such as economic stability and competitiveness.

Generally, the debate focuses on the valuation of the Chinese currency, Renminbi (hereafter, RMB) and on how China, artificially[58] depressing its currency's value and promoting policies that tend to depreciate the RMB or to keep it weak, increases its surplus and generates global imbalances without improving the effective non-price competitiveness factors. It is straightforward that China (and other rapidly developing countries) has enormously increased its share in trade[59] in the recent years even though industrialized countries have better performances and are more competitive[60] and the common opinion is that this positive trend is due to exchange rate policies.

In any direction the question always turns, though, to the role of China's real exchange rate, how this can explain its economic performance on the international markets and to the possibility that a RMB appreciation could reduce global imbalances. This is surely the point of view of some

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[58] In the sense that China is deliberately manipulating its exchange rate to obtain a competitive advantage. If it is true that every country's aim has always been its development and the achievement of always more profitable economic processes, it is also established that, at the moment, every country wants to grow as fast it can to try to overcome the financial crisis, and one of the ways to achieve this goal could be, a real depreciation, entailing a decrease in the price of labor with respect to other countries.

When trade becomes unbalanced, deficit countries need to raise interest rates to reduce demand for imports and exports, as well as reducing wages to increase competitiveness. Actually, there is no tangible self-regulating system that can lead to quick fixes or that can restore global growth and reduced wages are largely a response to higher unemployment.

Both American and European trade partners are particularly concerned with the self-protecting policies carried out by Asian emerging countries in order to overcome the financial crisis. This kind of behavior is seen as a potentially damaging dynamic that can lead to a global currency war. Indeed, the results from several specifications indicate that a real exchange rate appreciation will surely increase the value of the country's imports but the value of exports can either increase or decrease. A real depreciation is expected to stimulate growth by the expansion of exports and the contraction of imports. More recent literature (Gupta, Mishra, Sahay, 2007) focuses on the negative effects: a sudden stop or reversal of capital inflows during a crisis can slow down growth and the slowdown may be worse if the currency crisis is accompanied by a banking crisis or by competitive devaluation in other countries.

In general, the global rebalancing entails bilateral adjustments that proceed by steps and it is difficult to outline the whole process unmistakably. Internal and external rebalancing are, actually, the two sides of a same problem. The economies face the problem, on foreign and/or domestic basis, according to what they consider a priority. First of all, some economists and policy-makers delimitate the issue identifying two central economies, United States and China, while others think that all Asian economies are involved.

The USA urge an increase in Asian consumption spending without considering the need for more saving in the U.S.; on the other hand, Asian countries think that more saving is needed in the U.S. without considering that this implies an increase in spending in other countries to support global demand. Some consider exchange rate adjustments fundamental, while others do not consider them important (Eichengreen, Rua, 2010).

Finally, mentioning China, for example, its real exchange rate against the dollar has improved and, at the same time, wage growth and inflation have proceeded far faster in China than in America. China's real appreciation against other emerging Asian markets (against which it competes for export), though, has been far less, as those countries have also seen substantial inflation in prices and wages.

It is evident that there is a compound puzzlement; this confusion obviously leads to multiple (and, potentially, contradictory) explanations that need to be unraveled.

[59] These results, however, are very difficult to interpret because more than half of Chinese exports are classified as "transformation trade", that is, instrumental goods that are imported and processed by China to then be re-exported. Goldstein A. (2011).

[60] Goldstein A. (2011), *op. cit.*

deficit countries like USA. To further complicate the problem, China's undervaluing currency policies have potential implications not only for the industrialized economies but even for the so-called emerging ones and for the developing ones: indeed, some are afraid that a weak Chinese currency (that potentially brings to an increase in Chinese exports) will have a negative outcome as for their industrialization process; others think that the power of China's growth is the driving force of those economies and, therefore, any development that disadvantages China will penalize them too.

In a general perspective, the main and perhaps obvious conclusion beyond question is that even though the estimation of trade elasticities is far from being a new field of study, there is a need for continual estimation of trade elasticities and, due to the importance of the issue, it must be treated with great caution. Indeed, in spite of the large body of literature and of the development of the econometric specifications, there are still areas where the state of knowledge is rather inadequate. More specifically, for what concerns the responsiveness of exports to changes in the exchange rates, there is the possibility to contribute with further studies because still little is well-established.

CHAPTER 4

ESTIMATING EXPORT PRICE ELASTICITY USING NON-STATIONARY PANEL MODELS. EVIDENCE FROM A SAMPLE OF COUNTRIES

1. INTRODUCTION

Price elasticity estimation is one of the most important, controversial and intriguing topics in International trade.

This chapter describes the application of non-stationary panel techniques for time series to evaluate export price elasticities of a selected sample of countries. I first illustrate the motivation and the objectives of such analysis. Secondly, I outline the empirical background of the technique and present an empirical model; at last, I discuss the results.

The empirical implementation of a model on a cross-country time series sample poses two main challenges. First, although the model defines a long-run relationship among exports and its fundamentals, the equilibrium may be achieved *gradually* in the long run[61]. Hence, in the empirical analysis, the process of a short-run adjustment must complement the long run equilibrium model. This has been already accomplished (*cfr*. Chapter 3) applying a VECM to estimate the long-run export price elasticities as well as the short-run values and the speed of adjustment to the equilibrium. However, a constraint of this approach is that it does not allow a comparison between different countries. It is reasonable to argue that countries differ regarding, for example, market imperfections (e.g. labor or product market rigidities), monetary arrangements or different access to the international goods and capital markets. Thus, it is important to take into account the very likely possibility of heterogeneity across countries. For these reasons, I have implemented some panel time series methodologies which can provide more extensive avenues to approach the estimation of long-run export elasticities.

As illustrated (*cfr*. Chapter 2), the empirical literature on trade elasticities using aggregate data is vast and has gone through sophistication in estimation over the years, ranging from OLS and Distributed lag models to the more recent cointegration techniques. In particular, the cointegration literature has focused on the estimation of long-run relationships between *I(1)* variables using both

[61] The export elasticity may not always be in equilibrium at every point in time due to imperfections, rigidities or regulations.

time series and panel models. Within such framework, non-stationary panel econometrics represents a relatively new field of research, placed on the frontier of applied econometrics: in fact, the theory starts in early 1990s but most research goes back just to the past decade. These models have received, indeed, a lot of attention only in recent years. Scholars hope to combine the best of two worlds: the method of dealing with non-stationary data from the time series analysis and the increased data and power of cross-section analysis (Podestà, 2002; Hsiao, 2007; Baltagi, 2008; Bonham, 2013). The applications of time series methods applied to panels regard especially panel unit root tests, panel cointegration tests and the estimation of long-run average relations. Examples from real exchange rate literature include Frankel and Rose (1996), Jorion and Sweeney (1996), MacDonald (1996), Wu (1996), O'Connell (1998), Pedroni (1999), Maddala and Wu (2000), Groen and Lombardelli (2004), Smith et al. (2004), Pesaran *et al.* (1999, 2000, 2007), Binder and Offermanns (2007), Caporale et al. (2009), Thorbecke (2010, 2012, 2013).

Nevertheless, few researchers (Kubota, 2009; Jovanovic, 2012) use these methods to assess international trade elasticities so there is not much accessible literature yet (Eberhardt, 2011). The novelty of this analysis, therefore, is application of non-stationary panel time series techniques using aggregate trade data for a selected sample of countries.

2. EMPIRICAL SETTING AND DATA

In this study, the imperfect substitutes model proposed by Goldstein and Khan (1985) is followed. The major assumption of this model is that neither imports nor exports are perfect substitutes for domestic goods. Exports are imperfect substitutes in world markets for goods and services produced by others, or for third countries' exports. The conventional demand theory says that, the consumer is postulated to maximize utility subject to a budget constraint. In the vein of much research on this subject, (*cfr*. Chapter 2) we proceed using aggregate data. Therefore $X_{i,t}$ refers to total exports of country i at time t, while the relative-prices variable is gauged by the real exchange rate ($REX_{i,t}$) and the size of foreign demand by the world income (Y_t^w). In this respect, export demand function is specified as a function of the real exchange rate and the rest-of-world real incomes:

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

where $X_{i,t}$ is the exports of goods and services of country I at time t, $REX_{i,t}$ is the real exchange rate of country i at time t and Y_t^w is the world income at time t.

The Real Exchange Rate is given as:

$$REX_{i,t} = \frac{CPI_{it}}{CPI_{RoW,t}} * E_{it}$$

where CPI_{it} is the Consumer Price Index of domestic goods and services in country i at time t and $CPI_{RoW,t}$ is Consumer Price Index in the Rest of the World at time t. The nominal exchange rate is the domestic currency price of one unit of foreign currency. The *REX* is constructed so that increase stands for real appreciation.[62]

Given the log-linear form of the equation (1), β_1 is the real exchange rate elasticity of export demand and β_2 is the real foreign income elasticity. Based on the theory, it is expected that β_1 has a negative sign, implying an increase in demand with the depreciation of countries' currency and β_2 has a positive sign, indicating that exports rise as world income increases.

The model estimations are based on quarterly data between the years 1990:Q1 and 2012:Q1, collected from Datastream databases, the reference year being 2005=100. The sample of countries comprises Italy, Germany and France, UK, USA and Japan. In other words, we refer to the group of G7, but Canada.[63] In addition, we also consider China that represents the conversation piece of the current debate on devaluation, weak currencies and exchange rate misalignments because of its growing role as production hub and its consequent increased role as trade player. All the time series data are in real terms and are seasonally adjusted. Descriptive statistics and visual plots are illustrated in Chapter 3 and in Appendix 1.

3. ECONOMETRIC METHODOLOGY

Equation (1) is estimated using panel data techniques. The benefits from using panel data estimation are various. In panel data estimation, variations over both the cross-section and time series dimensions are considered jointly. This brings the advantage of using all the information available which are not detectable in pure cross-sections or in pure time series data, blending the inter-unit differences with the intra-unit dynamics. Other advantages are:

[62] The REX is based on the Consumer Price Index, and increase stands out for real appreciation (i.e. loss of competitiveness). It is constructed as a weighted average of the real exchange rates against selected countries. For further details, Durand *et al*. (1998).

[63] The so called *G6* countries (Italy, Germany, France, UK, USA and Japan) and China, as a representative of the BRICS. According to Goldstein (2011), G6 in some analyses, is more appropriate than G7 because Canada is "*sui generis*" for different reasons.

- more accurate inference of model parameters. Panel data usually contain more degrees of freedom and more sample variability than cross-sectional data which may be viewed as a panel with T =1, or time series data which is a panel with N = 1, hence improving the efficiency of econometric estimates;
- simplifying computation and statistical inference. Panel data involve at least two dimensions: a cross-sectional dimension and a time series dimension. Under normal circumstances one would expect that the computation of panel inference would be more complicated than cross-sectional or time series data. However, in certain cases, the availability of panel data actually simplifies both computation and inference (Hsiao 2007).

In addition to this, panel data estimation provides improved coefficient estimates by increasing the power of the tests (Maddala and Wu, 2000).

From an empirical perspective, the analysis is carried out by performing the panel unit root test proposed by Levin, Lin and Chu (2002) and the panel cointegration test developed by Westerlund (2007). After performing these two tests, we proceed by using the Mean Group (MG) estimator developed by Pesaran, Smith and Im (1996) and the Pooled Mean Group (PMG) method proposed by Pesaran et al. (1999).

3.1 Stationarity and Cointegration Tests

Before testing for the existence of a cointegrating relationships between the variables, it is required to examine the stochastic properties of each series. Since it is generally accepted that the commonly used unit root tests like the Dickey-Fuller (DF), augmented Dickey-Fuller (ADF) tests lack power in distinguishing the unit root null from stationary alternatives, and that using panel data unit root tests is one way of increasing the power of unit root tests based on a single time series, we use the homogeneous Levin, Lin and Chu (2002) panel unit root test.

This test assumes that each individual unit in the panel shares the same AR(1) coefficient, but allows for individual effects, time effects and possibly a time trend. Lags of the dependent variable are introduced to allow for errors' serial correlation. The test may be viewed as a pooled Dickey-Fuller test, or an ADF test when lags are included, with the null hypothesis of nonstationarity (*I(1)* behavior). The *t*-statistic converges to the standard normal distribution: therefore, the standard normal critical values are used in testing the hypothesis.

Table 4.1 shows the LLC unit root test for exports and exchange rate. Under the null hypothesis the series are non-stationary (H_0: series is *I(1)*):

Table 4.1. Levin Lin Chu test for all variables of the models in the panel

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

In the case of exports, the estimated coefficient of the one-year lagged variable is -0,0307 and the LLC test allows to accept the hypothesis of non-stationarity with a high level of statistical significance (the p-value is about 0.82). The same applies for the exchange rate, whose estimated coefficient of the one-year lagged variable is -0,0691, with a p-value around 0.19 (Table 4.1).

After non-stationarity has been ascertained, the next step is to estimate the cointegration relationship. This is done by implementing the test proposed by Westerlund (2007). I use the pooled test (Pt) which pools information over all the cross-sectional units. Considering the following model:

$$\Delta Y_{i,t} = \alpha_i + \beta_{i1} * \Delta Y_{i,t-1} + \beta_{i2} * \Delta Y_{i,t-1} + \ldots + \beta_{ip} * \Delta Y_{i,t-p} + \gamma_{i0} * \Delta X_{i,t} + \gamma_{i1} * \Delta X_{i,t-1} + \ldots + \gamma_{ip} * \Delta X_{i,t-p}$$

$$+ \beta_i (Y_{i,t-1} - \gamma_i * X_{i,t-1}) + u_{i,t} \qquad (2)$$

β_i provides an estimate of the speed of error-correction towards the long-run equilibrium $Y_{i,t} = -(\gamma_i/\beta_i)*X_{i,t}$ for that series. Pt test statistics pool information over all the cross-sectional units to test H_0: $\beta_i=0$ for all I versus H_1: $\beta_i<0$ for all i. Rejection of H_0 should, therefore, be taken as rejection of cointegration for the panel as a whole. The underlying idea is to test for the absence of cointegration by determining whether the individual panel members are error correcting. The test is very flexible and allows for an almost completely heterogeneous specification of both the long-run and short-run parts of the error correction model, where the latter can be determined from data. The series are also allowed to be of different length. The test has limiting normal distribution and is consistent. The results for all countries are reported in Table 4.2. The evidence shows that the H_0 of no cointegration is rejected and, therefore, a cointegrating relationship between exports and its fundamentals in the panel data exist.

Table 4.2 Westerlund ECM panel cointegration test

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

4. Panel estimation of the long-run relationship

Having found that the cointegrating relationship exists, we go on to estimate the export demand function using non-stationary panel methods. The estimation of the long-run export elasticity is made performing the Pooled Mean Group estimator (PMG) by Pesaran Shin and Smith (1999) and the Mean Group estimator (MG) by Pesaran, Smith and Im (1996). Both are non-stationary time series techniques for heterogeneous panels.

There are many alternative methods for multi-country estimation, which allow for different degrees of parameter heterogeneity across countries. At one extreme, the fully heterogeneous-coefficient model imposes no cross-country parameter restrictions. This specification can be estimated on a country-by-country basis, provided the time-series dimension of data is sufficiently large. As cross-country dimension is large, the mean of long and short-run coefficients across countries can be estimated consistently by the unweighted average of the individual country coefficients. This is the MG method introduced by Pesaran, Smith, and Im (1996). At the other extreme, the fully homogeneous coefficient model requires that all slope and intercept coefficients be equal across countries. This is the simple “pooled” estimator.

In ‘between two extremes’, there are a variety of estimators. The PMG method developed by Pesaran, Shin, and Smith (1999), restricts the long-run coefficients to be the same across countries but allows the short-run coefficients and the speed of adjustment to be country-specific. The PMG also generates consistent estimates of the mean of short-run coefficients across countries by taking the unweighted average of the individual country coefficients (provided that the cross-sectional dimension is large). In *I(1)* panels this estimator "allows for mix of cointegration and non cointegration" (Eberhardt, 2011).

The MG fits parameters as averages of the N individual group regressions and assumes homogeneity across countries for the long-run coefficients. This method has been employed by considering the following equation:

$$\Delta\log X_{i,t} = \delta_i + \beta'_1 \Delta\log REX_{i,t} + \beta_2 \Delta\log Y^w_{i,t} + \mathrm{v}_{i,t} \qquad (3)$$

(i=1,..,7; t= 1,..,89)

which is derived from eq. (1). In order to achieve the stationarity of the series, variables in equation (3) are in first differences (Δ), as they are non-stationary in level (cfr. §3.1).[64] By taking first differences we also control for unobserved fixed effects.

Deriving the empirical specification from equation (1) as well, the PMG model is expressed as:

$$\Delta\log X_{i,t} = \delta_i + \beta_{i1}\,\Delta\log REX_{i,t} + \lambda_i\,(\theta\, REX_{i,t-1}, - X_{i,t-1}) + \beta_{i2}\,\Delta\log Y^w_{i,t} + \mathrm{v}_{i,t} \qquad (4)$$

with $\mathrm{v}_{it} \sim \mathrm{iidN}(0, \sigma_i^2)$

In equation (4), the coefficients β_i are short-run parameters which, like σ_i^2, differ across countries. The error-correction term λ_i also differs across i, while the long-run parameter θ is constant across the groups. This estimator is quite appealing when studying small sets of arguably 'similar' countries rather than large diverse macro panels (Eberhardt, 2011).[65]

In choosing among these estimators there is a general trade-off[66] between consistency and efficiency. For my purposes, applying both MG and PMG offers the best available compromise in the search for consistency and efficiency. Indeed, the PMG is particularly useful when the long-run is given by country-independent equilibrium conditions, whereas the short-run adjustment depends on country characteristics such as financial development and relative price flexibility. In other

[64] The MG offers the opportunity to get only one short run and long run elasticities simply by averaging the estimations of each individual country. This is an advantage to use panel data instead of time series.
[65] The main requirements for the validity of both these methodologies are such that: *(i)* there exists a long-run relationship among the variables of interest and, *(ii)* the dynamic specification of the model be augmented such that the regressors are exogenous and the resulting residual is not serially correlated.
[66] "The comparison of the asymptotic properties of PMGE and MGE can be put in the general trade-off between consistency and efficiency. If the long-run coefficients are equal across countries, then the PMGE will be consistent and efficient while the MGE will only be consistent. If the long-run coefficients are not equal across countries, then the PMG estimates will be inconsistent while the MGE will be still a consistent estimate of the mean of long-run coefficients across countries. The long-run homogeneity restrictions can be tested by Hausman or likelihood ratio tests to compare the PMGE and MGE of the long run coefficients.". (Kubota, 2009).

words, the PMG predicts not only a common long-run equilibrium relationship but also short-run dynamics of each single country.

Since we cannot accept one of the two methods *a priori*, which of them is more appropriate will be decided on statistical grounds. The following sub-sections illustrate and discuss the results of the estimation of the export elasticities using the MG and the PMG methods. We will start by presenting the PMG estimations as the model can be considered as the MG with a constraint.

4.1 Pooled Mean Group (PMG) Estimation

As already said, the PMG restricts the long-run coefficients to be the same across countries, but allows the short-run coefficients (including the speed of adjustment) to be country-specific. The estimation is provided for the sample of seven countries over the period 1990-2012. Table 4.3 presents the results for the common long-run equilibrium elasticity and the individual short-run dynamics:

Table 4.3 Estimation of the export function of seven countries. Results from Pooled Mean Group Estimator (1990:Q1-2012:Q1)

	Coef.	Std. Err.	Z	P>|z|	[95% Conf. Interval]	
LR						
log(REX)	-0,8906	0,1350	-6,6	0	-1,1551	-0,6260
log(Y^w)	1,0813	0,0646	16,74	0	0,9547	1,2079
Italy - SR						
ec	-0,1297	0,0335	-3,87	0	-0,1954	-0,0641
Δlog(REX)	-0,3261	0,0878	-3,71	0	-0,4982	-0,1539
Δlog(Y^w)	3,9644	0,3951	10,03	0	3,1900	4,7388
intercept	0,4606	0,1533	3,01	0,003	0,1602	0,7609
Japan - SR						
ec	-0,1516	0,0344	-4,4	0	-0,2191	-0,0841
Δlog(REX)	*0,0482*	0,0732	0,66	0,511	-0,0953	0,1916
Δlog(Y^w)	7,1225	0,7090	10,05	0	5,7327	8,5122
intercept	0,5184	0,1372	3,78	0	0,2494	0,7874
France -SR						
ec	-0,0648	0,0175	-3,71	0	-0,0990	-0,0305
Δlog(REX)	-0,3225	0,1258	-2,56	0,01	-0,5690	-0,0759
Δlog(Y^w)	3,0207	0,3279	9,21	0	2,3780	3,6634
intercept	0,2251	0,0760	2,96	0,003	0,0763	0,3740
UK - SR						
ec	-0,0365	0,0146	-2,5	0,012	-0,0652	-0,0079
Δlog(REX)	-0,2337	0,0988	-2,36	0,018	-0,4274	-0,0400

$\Delta\log(Y^w)$	4,0029	0,6411	6,24	0	2,7464	5,2595
intercept	0,1130	0,0525	2,15	0,031	0,0101	0,2158
China - SR						
ec	-0,0345	0,0161	-2,14	0,032	-0,0661	-0,0029
$\Delta\log(REX)$	*0,0371*	0,0608	0,61	0,542	-0,0820	0,1561
$\Delta\log(Y^w)$	2,9605	0,5728	5,17	0	1,8378	4,0832
intercept	0,1176	0,0605	1,94	0,052	-0,0010	0,2363
Germany - SR						
ec	-0,0280	0,0153	-1,83	0,067	-0,0579	0,0019
$\Delta\log(REX)$	*-0,1888*	0,1795	-1,05	0,293	-0,5406	0,1630
$\Delta\log(Y^w)$	3,2094	0,5918	5,42	0	2,0495	4,3692
intercept	0,0935	0,0579	1,62	0,106	-0,0200	0,2069
USA - SR						
ec	-0,0469	0,0131	-3,59	0	-0,0725	-0,0213
$\Delta\log(REX)$	-0,2282	0,0750	-3,04	0,002	-0,3753	-0,0811
$\Delta\log(Y^w)$	2,5566	0,3928	6,51	0	1,7867	3,3266
intercept	0,1672	0,0516	3,24	0,001	0,0660	0,2684

*Significance level: '***' = 0.001; '**' = 0.01; '*' = 0.05; '.' = 0.1; ' ' = 1.*
Obs = 616; Numbero of Groups = 7; Obs per Group = 88
Log Likelihood = 1512.67
Source: elaborations on Datastream data.

The long-run homogenous price elasticity is negative and statistically significant: the estimated value is -0.89. This result suggests that, in the long-run, the exports are price inelastic for all the countries of the panel. It also shows that the panel is slightly foreign-income elastic in the long-run (being the elasticity around unity) and highly elastic in the short-run. Income elasticity ranges from 2.5 in the case of USA to 7.12 for the Japan. With regards to the speed of adjustment term, we can notice that it presents very low estimations, ranging from -0.03 (China and Germany) to -0.15 (Japan).

A by-product of PMG estimations regards the Marshall-Lerner (ML) condition for the entire sample of countries. This condition sets the boundary value beyond which currency depreciation policies can be considered effective around unity (>1 in absolute value when considering only export elasticities). The test has formally been executed using the estimated results reported in table 4.3 for what concerns the exports and those in the Appendix 5 for imports. When price elasticities are obtained for both exports and imports along with the corresponding standard errors the following test can be executed:

Table 4.4 Marshall-Lerner condition test

PMG export			PMG import		
LR		*Se*	**LR**		*Se*
Log (*REX*)	0,890565	(0,134974)	Log (*REX*)	0,261178	(0,135425)
Critical value: 1.96					
t = 0.52					

We can say that the sum of the absolute values of the import and export elasticities are significantly greater than one if the corresponding t-ratio is greater than the critical value of 1.96: since t= 0.52, the M-L condition is not met. This is consistent with Bahmani et al. (2013),[67]

4.2 Mean Group (MG) Estimation

Table 4.5 reports the results obtained when using the MG estimator. The countries have a statistically significant negative coefficient for the real exchange rate (*REX*) which varies between 0.52 (Japan) and 2.04 (France) in the long-run and between 0.19 (USA) and 0.33 (Italy) in the short-run. UK and China are the only exceptions in the long-run estimates, as they present statistically non-significant coefficients, while Japan, China and Germany present statistically non-significant coefficients in the short-run.

The aggregate export demand, consistent with the range of results of the literature (*cfr*. Chapter 2), is found to be real exchange rate inelastic both in the long-run and short-run with the only two exceptions of France (long-run exchange rate elasticity = |2.04|) and USA (long-run exchange rate elasticity = |1.77|). It is foreign income (Y^w) elastic both in the long-run and in the short-run. The results of price-elasticity found in table 4.5 indicate that the Marshall-Lerner condition holds for France and USA. Phrased differently, in case of France and USA, the condition of Marshall-Lerner is robust to the method used for estimating the export function.

Interesting insights also come from the dynamics towards the long-run equilibrium. The error correction speed of adjustment term is found to be high in the case of Japan and Germany (-0.23 and -0.33 respectively) meaning that they reach their long-run equilibrium faster with respect to the other countries of the panel. At the opposite side, the speed of adjustment is very low (-0.05) for USA. All this implies that USA will take 60 time units (quarterly) to get the equilibrium elasticity, whereas Germany and Japan will be in the long-run equilibrium in less than a year (9

[67] "The Marshall-Lerner condition does not hold in a large fraction of the cases in which it is claimed to do so. This has strong implications for future analyses of trade and exchange-rate policy".

months for Germany).[68] Lastly, it is worth noticing that Italy presents exactly the same long-run price elasticity ($|0.72|$) found using the time series VECM model (*cfr*. Chapter 3, Table 3.9).

Table 4.5 Estimation of the export function of seven countries.
Results from Mean Group Estimator (1990:Q1-2012:Q1)

	Coef.	Std. Err.	Z	P>|z|	[95% Conf. Interval]	
Italy – LR						
log(REX)	-0,7249	0,2217	-3,27	0,001	-1,1594	-0,2905
log(Y^w)	0,9768	0,0947	10,32	0	0,7913	1,1624
Italy – SR						
Ec	-0,1218	0,0344	-3,54	0	-0,1893	-0,0544
Δlog(REX)	-0,3283	0,0899	-3,65	0	-0,5045	-0,1520
Δlog(Y^w)	4,0579	0,4101	9,89	0	3,2541	4,8617
Intercept	0,3950	0,1802	2,19	0,028	0,0417	0,7482
Japan – LR						
log(REX)	-0,5254	0,1469	-3,58	0	-0,8133	-0,2375
log(Y^w)	1,3637	0,0975	13,98	0	1,1726	1,5549
Japan – SR						
Ec	-0,2331	0,0501	-4,66	0	-0,3313	-0,1350
Δlog(REX)	*0,0619*	0,0743	0,83	0,405	-0,0837	0,2075
Δlog(Y^w)	6,9404	0,7197	9,64	0	5,5299	8,3510
Intercept	*0,1251*	0,2245	0,56	0,577	-0,3149	0,5652
France –LR						
log(REX)	-2,0405	0,5828	-3,5	0	-3,1828	-0,8982
log(Y^w)	1,0052	0,1682	5,98	0	0,6754	1,3349
France -SR						
ec	-0,0764	0,0248	-3,08	0,002	-0,1251	-0,0277
Δlog(REX)	-0,2626	0,1334	-1,97	0,049	-0,5241	-0,0012
Δlog(Y^w)	3,0248	0,3332	9,08	0	2,3716	3,6779
intercept	0,6982	0,2514	2,78	0,005	0,2055	1,1910
UK - LR						
log(REX)	*-0,1159*	0,3412	-0,34	0,734	-0,7846	0,5529
log(Y^w)	1,4688	0,1706	8,61	0	1,1345	1,8031
UK - SR						
ec	-0,0990	0,0472	-2,1	0,036	-0,1915	-0,0065
Δlog(REX)	-0,2270	0,1020	-2,23	0,026	-0,4270	-0,0271
Δlog(Y^w)	3,9665	0,6597	6,01	0	2,6735	5,2594

[68] 1/0,05 = 20. 20 *(3 months) = 60 months.

intercept	*-0,1837*	0,2268	-0,81	0,418	-0,6283	0,2609
China - LR						
log(REX)	*-0,2207*	0,3009	-0,73	0,463	-0,8104	0,3690
log(Y^w)	1,5546	0,1527	10,18	0	1,2554	1,8538
China - SR						
ec	-0,1175	0,0455	-2,58	0,01	-0,2067	-0,0284
Δlog(REX)	*0,0430*	0,0623	0,69	0,49	-0,0791	0,1650
Δlog(Y^w)	3,1020	0,6107	5,08	0	1,9050	4,2989
intercept	*-0,1951*	0,1897	-1,03	0,304	-0,5669	0,1768
Germany - LR						
log(REX)	-0,6702	0,1759	-3,81	0	-1,0150	-0,3254
log(Y^w)	2,0309	0,0534	38,03	0	1,9263	2,1356
Germany - SR						
ec	-0,3287	0,0677	-4,86	0	-0,4613	-0,1961
Δlog(REX)	*0,1100*	0,1775	0,62	0,536	-0,2380	0,4579
Δlog(Y^w)	3,0716	0,5455	5,63	0	2,0023	4,1408
intercept	*-0,5654*	0,3704	-1,53	0,127	-1,2914	0,1605
USA - LR						
log(REX)	-1,7666	1,1816	-1,5	0,135	-4,0825	0,5494
log(Y^w)	1,3541	0,2893	4,68	0	0,7870	1,9212
USA - SR						
ec	-0,0502	0,0305	-1,65	0,1	-0,1100	0,0096
Δlog(REX)	-0,1921	0,0810	-2,37	0,018	-0,3508	-0,0333
Δlog(Y^w)	2,6022	0,4052	6,42	0	1,8081	3,3964
intercept	0,3195	0,1563	2,04	0,041	0,0132	0,6258

*Significance level: '***' = 0.001; '**' = 0.01; '*' = 0.05; '.' = 0.1; ' ' = 1.*

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

Source: elaborations on Datastream data.

A final discussion to be made regards the short-run elasticity. It has already been argued that both PMG and MG estimators allow short-run dynamics to differ across countries. By comparing the results, it emerges that the export price elasticity estimates are very similar in the short-run: in absolute values, they range from 0.23 (USA) to 0.33 (Italy). In this sense, the used-method does not impact on the results when estimating the short-run elasticity. On the contrary, the long-run equilibrium elasticity is different across countries in MG and they also differ from the long-run elasticity estimated with the PMG for the entire pool of countries. Based on the results of broadly similar short-run elasticity and different long-run elasticity, it is easy to expect that the speed of adjustment is sensitive to the method.

Table 4.6 EC estimates - Speed of Adjustment

	MG	PMG
Italy	-0,12	-0,13
Japan	-0,23	-0,15
France	-0,08	-0,06
UK	-0,10	-0,04
China	-0,12	-0,03
Germany	-0,33	-0,03
USA	-0,05	-0,05

These results probably can be used to argue that for some countries (i.e. USA, Italy and France) the constraint of a common long-run equilibrium elasticity is more reasonable than for others (China, UK, Japan and, in particular, Germany) because their pooled and average speed of adjustment are similar. Finally, Germany shows the largest discrepancy between the MG and the PMG estimates: when using MG, Germany reaches its long-run equilibrium elasticity in 9 months; when parameters are pooled and group-specific dynamics are allowed, Germany would reach the long-run value – common for all the countries of the panel – in more than 8 years.

With regards to the ML condition, according to the long-run equilibrium estimates displayed in table 4.5, it is possible to argue that this condition is met when considering France and USA, but no conclusion may be drawn for the other five countries. Therefore, it is hard to offer a generalization of the results for what concerns the relationship between exchange rate devaluation and trade balance. Indeed, the results of this study could suggest that there is no simple, consistent relationship between trends in the trade balance and trends in real exchange rates and so estimates need to be used with great caution.[69]

A further step of the study concerns the evidence about the best performing method, provided the sample of countries under scrutiny. At this end, we compare the MG and PMG by applying the likelihood ratio (LR) test. The LR test is commonly used to evaluate the difference between *nested* models once these are estimated. One model is considered nested in another if the first model can be generated by imposing restrictions on the parameters of the second. In this case, the restricted model is the PMG, while the unrestricted one is the MG. Under the H_0 hypothesis

[69] Additionally, if we start from a situation of imbalanced trade (deficit or surplus in the trade balances) and if the aim is to reduce or eliminate the deficit through variations in real exchange rates, the Marshall-Lerner condition is not sufficient as one of the assumptions of this condition is that we start from a situation of balanced trade. This analysis examines surplus but also deficit countries and this makes the interpretation of the results more complicated.

there is the common long-run equilibrium, while the alternative hypothesis is that the long-run elasticities differ from one country to another (as assumed by the MG estimator). According to the results of the LR test displayed in table 4.7, we reject the null hypothesis and, in choosing between the two models, the evidence suggests that MG fits better the data than PMG. According to this, it is argued that the countries are not constrained to have a common long-run elasticity.

Table 4.7 Likelihood-ratio test

H_1: no constraints (MG)	H_0= common long-run (PMG)
LR chi2(12) = 44.0	Prob > chi2 = 0.000

4.3 Testing for structural breaks

The test for cointegration with regime shifts proposed by Gregory and Hansen (1996) is applied in order to examine the possibility of regime shifts in our models. The GH test considers cases where the intercept and/or slope coefficients have a single break of unknown timing. When considering the long-run relationship between exports and exchange rates, the GH procedure allows to identify possible breaks: when this occurs, it tests the null hypothesis of absence of change in the long-run relationship. Under the alternative, instead, there is a shift towards a new long-run equilibrium (Gregory and Hansen, 1996). We account for structural change using three different kinds of breaks.[70] The first presents a level shift in the long-run relationship, which can be modeled as a change in the intercept, while the slopes are constant. This implies that the equilibrium equation has shifted in a parallel fashion. This model is known as the *Level shift model* and the relative results are reported in Table 4.8[71]. In the second case, called *Level shift with trend model*, a time trend is introduced into the level shift model (Table 4.9). Another possible structural change allows the slope vector to shift as well. This permits the equilibrium relation to rotate and also allows a shift. This is the third test applied, known as the Regime shift model (Table 4.10). Gregory and Hansen (1996) proposed an extension of the ADF, Zt and Za tests for cointegration. The structural changes are reported for each test (Tables 4.8, 4.9 and 4.10). These test statistics detect the stability

[70] The G-H test (Gregory and Hansen, 1996), defines a dummy variable to model structural change:

$\varphi_{t\tau} = \begin{cases} 0 & \text{if } t \leq [n\tau] \\ 1 & \text{if } t > [n\tau] \end{cases}$ where the unknown parameter $\tau \in (0,1)$ denotes the timing of the change point and [ ] denotes integer part.

[71] The statistical software package used in this section is Stata 10.

of cointegration over time in the presence of structural changes in the form of shifts, level shifts with trend and regime shifts as we defined above.

Table 4.8 Gregory-Hansen Test for Cointegration with Regime
Shifts Model: Change in Level Number of obs = 89
Lags= 2 chosen by Akaike criterion Maximum Lags = 5

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

Table 4.9 Gregory-Hansen Test for Cointegration with Regime Shifts Model: Change in Regime Number of obs = 89
Lags= 2 chosen by Akaike criterion Maximum Lags = 5

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

Table 4.10 Gregory-Hansen Test for Cointegration with Regime Shifts Model: Change in level and trend Number of obs = 89
Lags= 2 chosen by Akaike criterion Maximum Lags = 5

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

According to the tests, with the exception of USA, Japan and Germany, the other four countries present structural changes (in level and in trend) in 2008, likely triggered by the financial crisis whereas UK presents a change in level even in 2007, the year of the onset of the crisis. USA presents structural changes (in level and in trend) beginning in the last quarter of 2001, probably due to the World Trade Center terrorist attack and the *dot.com* crisis. The accession of China to the

World Trade Organization in 2001 and that of the Chinese Taipei in 2002 surely are the reasons of the structural changes in level and in regime showed from the third quarter of 2001 to the first quarter of 2003.

Nevertheless, the results show that the long-run elasticity does not change before and after the structural breaks: the calculated statistics are in all cases lower than the asymptotic critical values at 1%, 5% and 10%.

5. FINAL REMARKS

In this chapter, the export demand elasticities of real exchange rate and foreign income are estimated by using non-stationary time series data. One of the objectives of this study is the investigation and application of some panel data methods.

In empirical analysis, LLC panel unit root test and Westerlund panel cointegration test are performed. By finding evidence in favor of the cointegration relationship between variables, two models, MG and PMG, are estimated for total exports by using the data of the G6 countries and China.

One interesting result obtained from the empirical analysis is the estimated error correction term: it makes us lean towards choosing the MG as the most appropriate method to estimate long-run export elasticity. This choice is also supported by the LR post-estimation test, ran to compare MG with PMG.

The conclusions that can be drawn from the empirical results of the MG model are twofold. First of all, I find that aggregate export demand is foreign income elastic both in the long-run and in the short-run. This can be interpreted as growth in trade partner countries may affect a country's export positively and significantly. This result is consistent with the expectation and the evidence provided by others.

Secondly, the analysis shows that aggregate export demand is real exchange rate inelastic both in the long-run and short-run for five of the seven countries included in the study. The two exceptions are France, with a long-run exchange rate elasticity = $|2.04|$ and USA, with a long-run exchange rate elasticity = $|1.77|$.

This gives support to the hypothesis that the exchange rate policies may not be successful in promoting export growth**.** The results obtained from the PMG model confirm that export demand is foreign income elastic (both in the short-run and in the long-run) while it appears to be price inelastic in the long-run, reporting an estimate less than one, although rather high ($|0.89|$).

Low price elasticities may not be surprising, especially when dealing with aggregate data and when referring to developing countries[72]. Nevertheless, these results are puzzling in the light of the debate on currency devaluation, that seems to imply that exports are highly price elastic. Indeed, according to the results, the effects of exchange rate policies on exports seems to be fairly limited, hence, in order to obtain a sustainable and stabilized export growth, trade policies, which are based on diversification of exported products and production of technology-intensive goods, have to be developed rather than currency policies.

[72] "…export growth may be more dependent to factors like foreign demand, production capacity, productivity, diversification of exported goods and production of technology-intensive goods rather than price changes." (Coşar, 2002).

FORTHCOMING RESEARCH

There are many areas for further research and, mainly, the effort will be to deepen the study of the determinants of trade price elasticities in order to develop and implement an econometric model free from specification errors and able to capture a range of (underlying) variables. Given that the impact of macroeconomic policies based on time series techniques is traditionally analyzed at the level of the overall economy or for highly aggregated sectors, another study could entail the investigation of what happens when there is a very high good/service disaggregation level: that is, if disaggregation changes the overall results. In particular, to deepen the understanding of the role of exchange rate policies and currency manipulation in China's trade imbalance, future work should investigate whether further disaggregation (examining commodity types more finely or conducting sectoral level analyses) can yield greater insights into Chinese trade behavior.

Finally, it would be interesting to study the trade dynamics of the so-called globalization. The exchange rate is the key relative price in international finance; the rapid pace of globalization in goods and asset markets has only enhanced the importance of this variable. Globalization and global supply chains have certainly changed the way trade responds to relative price changes. In particular, higher imported content in exports is likely to lower the sensitivity of trade to changes in the exchange rate: examining whether the re-export issue implicates significant findings. In the case of the Chinese trade flows, for instance, there is some reason to believe that the conventional elasticities approach is insufficient.

CONCLUSIONS

The question of misaligned currencies in real terms is important in both academic and policy debate because it may reflect distortions in relative prices attributed to devaluation policies. Along this line of research, the study of export price-elasticities becomes relevant as it sheds some lights on their determining factors so that policymakers could attempt to implement the required adjustments.

This work contributes to long-dated debate among economists concerning (i) the impact that weak currencies may have on economic growth by promoting exports, and (ii) the role of real exchange rate which is to be meant as the relative price that would drive the international adjustment of countries.

Despite the plethora of studies conducted in this field of research, there is an evident discrepancy in the estimated values of price-elasticities. The literature on trade elasticity developed in the last fifty years have provided, indeed, a high heterogeneity in results which makes difficult any interpretation. The sample of studies quoted in the present research confirms that the high variability in the estimates change for numerous reasons. The factors that yield different elasticity estimates are basically related to: differences in sample periods, differences in models/approaches (OLS, ARDL, DOLS, ECM, MG, PMG, etc.) and to differences in levels of aggregation.

From an empirical perspective, the techniques used in the past before the introduction of the cointegration approach have often left behind a number of issues such as the response lags (Stern et al., 1976). In the less recent theoretical studies, indeed, it is assumed that prices (and quantities) adjust instantaneously to some given exogenous change; realistically, however, it will take time for adjustment to take place. This means that the related policies implemented on the basis of these predictions (size and time patterns) were, at worst, erroneous. The introduction of explanatory lagged variables takes into account these issues but implies other questions such as multicollinearity. A big contribution in overtaking some of these issues is given, as aforesaid, by the cointegration (time series and panel) techniques: the introduction of concepts and tools associated with cointegrated data has profoundly altered the econometrics (Hendry and Juselius, 2000). Nevertheless, application of cointegration analysis requires careful thought about model specification and interpretation to be sure to avoid forecast failure.

From a policy-making perspective, the high variability of the estimates makes it very difficult to appraise the actual effects of changes of exchange rates (and income) and, it can be thought that, in spite of its self-evident importance, the estimated elasticities are used just to translate (Marquez, 1999) predictions of prices and incomes into predictions for exports (and imports). It is also true that a large dispersion can undermine the usefulness of these estimates in the analyses of international interdependencies, exchange rate misalignments, global imbalances and, in general, in the measurement of policy effectiveness. This is one of the main reasons why policy-making is now, more than ever, a challenging task especially for the economies that have been hit hard by the financial crisis and that are endeavouring a gradual recovery and explains why this field of study is still unsaturated and why there is still the need to identify one or more factors of variability and/or instability: any result in this sense contributes to fill a gap in the empirical literature. A reliable estimate of the level of exchange rate misalignment and of trade (exports, specifically) elasticities gauges the severity of the problem and contributes to formulating the appropriate policy response whereas an imprecise estimate makes it difficult to comprehend the extent and the importance of the problem and to articulate a suitable policy.

In this study the main goal was to complement and improve upon the existing literature on export elasticities. To this end, I have reviewed the literature and estimated short and long-run export elasticities, focusing on the role of exchange rate variations. A traditional time series VECM and a non-stationary panel for time series data accounting for conditional long-run homogeneity in dynamic panel (PMG) and for complete heterogeneity (MG) have been applied and the results have been discussed and compared.

The objective of Chapter 2 was twofold. It acts as a gateway to the methodology and to the econometric specification applied in the present analysis and provides an overview of the previous empirical and theoretical literature within the international trade elasticities context. Due to the great number of studies related to this issue, the different contributions have been analysed considering the main empirical and theoretical approaches with an emphasis on trade elasticities theories and with an overview of selected empirical contributions of the recent years. This section has to be read as a detailed summary that provides a background to the recent economic developments in times series econometrics and, in particular, to the estimation of international trade elasticities. However, it is not exhaustive.

The results of export elasticities estimates from VECM methodology are reported in Chapter 3. My first goal was to complement the existing literature by estimating the time series for a sample of countries (Italy, Germany, France, UK, USA, Japan and China) covering the period from 1990 to 2012. I find that there is no strong evidence of the Marshall-Lerner condition except for China, France and USA and, with respect to this, that the benefit on exports of currency devaluation remains an open question for future analyses of exchange rate policies.

This part of the work also explains the main advantages of using the above mentioned methodology and provides a detailed explanation of the development of the VECM technique in the empirical literature. Nevertheless, in order to benefit of the advantages of using panel time series data and to combine the best of two worlds (the method of dealing with non-stationary data from the time series and the increased data and power of cross-section analysis), I proceed by estimating the long-run elasticities using panel techniques, blending inter-unit differences and intra-unit dynamics.

Chapter 4, indeed, describes the application of a non-stationary panel technique for time series. Although the model assumes a long-run relationship among exports and its fundamentals, the export elasticity may not always be in equilibrium at every point in time: in fact, the equilibrium may be achieved gradually in the long-run. Hence, in the empirical analysis, the process of a short-run adjustment complements the long-run equilibrium model. The two models proposed are the PMG and the MG: PMG allows the short-run coefficients to differ but constrains the long run coefficients to be the same for all groups. MG assumes complete heterogeneity, that is, it imposes no constraints on any of the parameters and allows to estimate a separate equation for each group. The estimated elasticities in both models show that exports are price inelastic for all countries except for France and USA. In fact, although conventional wisdom holds that a trade surplus could be achieved by weakening the associated currency, reviewing the standard economic analyses and considering the results of the present research, probably the exchange rate effect is more complicated: in this study, referring to the long-run price elasticity estimates using VECM, 70% of results are lower than unity in absolute values; this percentage increases when using the MG method (71.43%). The PMG long-run price elasticity estimate for all the countries of the panel is also lower than unity in absolute values (-0.89). Statistically significant estimates of price elasticities lower than 1 lead to economic observations that apparently do not meet currency depreciation and global imbalances issues as they do not meet the ML condition. Additionally, one interesting thing to notice is that the long-run export price elasticity estimates provided for Italy, using the three different techniques in same sample period, are (almost) exactly the same (about -0.7). Lastly, according to the results, the ML condition is not met even when formally tested: this has, obviously, strong implications for future analyses of trade and exchange rate policy. While it is very common to think that “a competitive real exchange rate is at the heart of the authorities’ development strategy" (Eichengreen, 2008), according to the evidence provided by the study, it should be noted that this sort of exchange rate policy will unlikely produce the desired effects as the exports seem to be price inelastic.

In conclusion, the evidence of this study shows that countries that target real exchange rates "competitiveness" may not actually achieve their goal. This study also confirms that export price elasticities in the literature are, to a certain extent, puzzling due to their high variability across sample periods and econometric methodologies. While the last chapter of this dissertation employs the most advanced empirical setting based on non-stationary panel data, the evidence provided is in the vein and spirit of the past related empirical evidence, as export price-elasticises are found to be generally less than unity. In some ways, the use of a method placed on the frontier of applied econometrics does yet not help to understand why world's leading exporters focus on competitive devaluations as reported by America's Treasury in the semi-annual report to Congress of October 2013: "*Within the euro area, countries with large and persistent surpluses need to take action to boost domestic demand growth and shrink their surpluses. Germany has maintained a large current account surplus throughout the euro area financial crisis, and in 2012, Germany's nominal current account surplus was larger than that of China. Germany's anemic pace of domestic demand growth and dependence on exports have hampered rebalancing at a time when many other euro-area countries have been under severe pressure to curb demand and compress imports in order to promote adjustment.[…].Treasury will continue to monitor closely exchange rate developments, with particular attention to the need for greater RMB appreciation*".

REFERENCES

Algieri B. (2004), Price and Income Elasticities of Russian Exports, The European Journal of Comparative Economics, Vol. 1, n. 2, 2004, pp. 175-193.

Algieri B. (2010), Modelling export equations using an unobserved component model: the case of the Euro Area and its competitors", Empirical Economics, Vol. doi:10.1007/s00181-0, pp. 1- 45.

Anaya J.A. (2000), Exchange Rate Pass-through and Partial Dollarization: Is there a Link? Center For Research On Economic Development And Policy Reform, Stanford University, Working Paper, No. 81.

Anderton, R. (1991), Underlying trends in manufacturing export shares for six major industrialized countries, Bull Econ Res 43(2), pp.169–178.

Aquino A. (1990), Squilibri e riaggiustamenti nel sistema dei pagamenti internazionali: elementi di analisi teorica ed esperienze dei paesi industriali, Giappichelli Editore.

Aziz J., Li X. (2007), China's Changing Trade Elasticities, IMF Working Papers, WP/07/266.

Bahmani-Oskooee M., Niroomand F., (1998), Long run elasticities and the Marshall-Lerner condition revisited, Economic letters 61, pp. 101-109.

Bahmani M., Goswami G. (2004), Exchange rate sensitivity of Japan's bilateral trade flows, Japan and the World Economy, 16, pp. 1-15.

Bahmani et al. (2013) "Empirical tests of the Marshall-Lerner condition: a literature review", Journal of Economic Studies, Vol. 40 Iss: 3, Pages 411 – 443.

Baltagi B.H, Kao C. (2000), Nonstationary panels, cointegration in panels and dynamic panels: A survey, Center for Policy Research, Working Paper No. 16.

Baltagi, B. H. (2008). Econometric Analysis of Panel Data (4th ed.). New York: John Wiley and Sons.

Banco de Espaňa (2003), Comparative analysis of export demand for manufactures in the Euro Area countries, Econ Bull July, pp. 67–75.

Bayoumi, T. et al. (2011), Euro area export performance and competitiveness. International Monetary Fund.

Beffy P. O., Ollivaud P., Richardson P., Sédillot F. (2006), New OECD methods for supply-side and medium term assessments: a capital services approach, OECD Economics Department Working Paper, No. 482.

Binder M. and Offermanns C.J. (2007), International investment positions and exchange rate dynamics: a dynamic panel analysis, Discussion Paper, Series 1: Economic Studies, No 23/2007.

Bonham, C. S. et al. (2013), Estimating Demand Elasticities in Non-Stationary Panels: The Case of Hawaii Tourism. Working Paper-University of Hawaii Economic Research Organization, University of Hawaii at Manoa.

Broda C., Weinstein D. E. (2006), Globalization and the gains from variety, The Quarterly Journal of Economics.

Bussièr et al. (2013), Estimating Trade Elasticities: Demand Composition and the Trade Collapse of 2008-09, American Economic Journal: Macroeconomics 2013, 5(3): 118-151.

Ca' Zorzi M., Schnatz B. (2007), Explaining and forecasting Euro Area exports which competitiveness indicator performs best? ECB Working Paper, 833.

Caporale G.M., Chui M.K.F. (1999) Estimating income and price elasticities of trade in a cointegration framework, Rev. Int. Econ 7, pp. 192–202.

Caporale et al. (2009), International Financial Integration and Real Exchange Rate Long-run Dynamics in Emerging Countries: Some Panel Evidence, Brunel university West London, Working Paper n° 09-16

Coşar, E. E. (2002). Price and income elasticities of Turkish export demand: a panel data application. Central Bank Review, 2, 19-53.

Costantini, Destefanis (2009), Cointegration analysis for cross-sectionally dependent panels: The case of regional production functions, Economic Modelling, Volume 26, Issue 2, March 2009, Pages 320‑327.

Cottrell A., Lucchetti, R:"J"., (2011), Gnu Regression, Econometrics and Time-series Library, see http://www.gnu.org/licenses/fdl.html Department of Economics and Dipartimento di Economia, Università Politecnica delle Marche.

De Grauwe, P. (2003): The economics of monetary union. 5th ed., Oxford University Press.

Deutsche Bundesbank; OECD, National Accounts Database, www.oecd.org.

De Vanssay X. (2003), The Marshall-Lerner Condition, An Encyclopedia of Macroeconomics, Edward Elgar Publishing, pp. 461-464.

Di Mauro F., Maurin L. (2005) Euro Area export performance and price competitiveness in: Competitiveness and the export performance of the Euro Area. ECB Occasional Paper Series, 30, pp 12–24.

Durand M., Madaschi C., Terribile F. (1998), Trends in OECD Countries' international competitiveness: the influence of emerging market economies, OECD Economics Department Working Paper, No. 195.

Eberhardt M. (2011). "Panel time-series modeling: New tools for analyzing xt data," United Kingdom Stata Users' Group Meetings 2011 22, Stata Users Group.

Eichengreen B., Rua G. (2010), Exchanges rates and global rebalancing, ADBI Working Paper Series, No. 278.

Erkel-Rousse H., Mirza D., (2002), Import price elasticities: reconsidering the evidence, Canadian Journal of Economics, Vol. 35, No. 2.

Escaith H., Lindenberg N., Miroudot S. (2010), International supply chains and trade elasticity in times of global crisis, World Trade Organization Economic Research and statistic divisions, Staff working paper.

European Commission (2004), Quarterly report of the Euro Area, Vol. 3, no. 1. European Commission, Directorate General for Economic and Financial Affairs.

European Central Bank, ECB (2004), Understanding the impact of the external dimension on the euro area: trade, capital flows and other international macroeconomic linkages, no. 12. ECB, Frankfurt am Main.

European Central Bank, ECB (2005), Competitiveness and the export performance of the Euro Area, Occasional paper Series, No. 30, ECB Frankfurt am Main.

European Commission (2012), In-depth review for Italy, Commission staff working document, SWD, n° 156.

Fanelli L., M Mazzocchi M. (2002), A cointegrated VECM demand system for meat in Italy, Applied Economics, Vol. 34, Issue 13, pp. 1593-1605.

Feenstra R. C. (1994), New product varieties and the measurement of International prices, The American Economic Review, Vol. 84 (1), pp. 157-177.

Felipe J. *et al.* (2010), Why has China succeeded? And why it will continue to do so, Asian Development Bank, Mimeograph. Working paper, Levy Economics Institute of Bard College.

Ferreira, J.E. (2006) : Effects of fundamentals on the exchange rate: A panel analysis for a sample of industrialised and emerging economies, Department of Economics Discussion Paper, University of Kent, No. 06,03

Frankel J. A., Rose A. K. (1996), Currency Crashes in Emerging Markets: Empirical Indicators, NBER Working Paper No. 5437.

Goldstein, M., Khan, M. S. (1976), Large versus small price changes and the demand for imports, IMF Staff papers, 23, pp. 200-225.

Goldstein, M., Khan, M. S. (1985), Income and price effects in foreign trade, in: Jones R. W., Kenen P. B. Handbook of International Economics, Amsterdam, North Holland (1985), pp. 1042-1099.

Gregory, Hansen (1996), Residual-based tests for cointegration in models with regime shifts, Journal of Econometrics 70, Pages 99- 126.

Groen, Lombardelli, (2004), "Real exchange rates and the relative prices of non-traded and traded goods: an empirical analysis," Bank of England working papers 223, Bank of England.

Gujarati D. N. (1995), Basic Econometrics, third edition, McGraw- Hill, Inc.

Gupta P., Mishra D., Sahay R. (2007), Behavior of output during currency crises, Journal of International Economics 72, pp. 428–450.

Hamilton J. D. (1994), Time Series Analysis, Princeton University Press.

Hamori S., Yin F. (2011), Estimating the import demand function in the autoregressive distributed lag framework: the case of China, Economics bulletin, vol. 31, issue 2.

Harberger, A. C. (1953), A structural approach to the problem of import demand, American Economic Review, 43, pp. 148-159.

Harberger, A. C. (1957), Some evidence on the international price mechanism, Journal of Political economy, 65, pp. 506-521.

Helliwell J. F., Padmore T. (1985), Empirical studies of macroeconomic interdependence, Handbook of International Economics, Vol. 2, pp. 1108-1151.

Hendry D. F., Juselius K. (1999), Explaining Cointegration Analysis: Part I, Nuffield College, Oxford, European University Institute, Florence.

Hendry D. F., Juselius K. (2000), Explaining Cointegration Analysis: Part II Discussion Papers Department of Economics, University of Copenhagen.

Hickman, B. G., Lau L. J. (1973), Elasticities of substitution and export demands in a World Trade model, European Economic Review, 4, pp. 347-380.

Holtemöller, Oliver; Zeddies, Götz (2010), Has the Euro increased international price elasticities? Working Paper, IWH - Diskussions papiere, No. 2010,18.

Hooper P., Johnson K., Marquez J. (1998), Trade elasticities for G-7 countries. International Finance discussion papers, no. 609.

Hooper P., Johnson K., Marquez J. (2000), Trade elasticities for the G-7 Countries, Princeton Studies in International economics, no. 87.

Hoover K. D., Perez S. J. (2000), Three attitudes towards data mining, Journal of Economic Methodology 7:2, 195–210 2000.

Hsiao (2007), Panel data analysis—advantages and challenges, Sociedad de Estadística e Investigación Operativa.

Imbs, J., & Mejean, I. (2010), Trade elasticities: a final report for the European Commission (No. 432). Directorate General Economic and Monetary Affairs (DG ECFIN), European Commission.

Jorion, P., Sweeney, R., (1996). Mean reversion in real exchange rates: evidence and implications for forecasting. Journal of International Money and Finance. 15 (4), 535–550.

Joseph A. M.(2004), Qualitative Research Design: An Interactive Approach.

Ketenci N. and Uz I., (2011), Bilateral and regional trade elasticities of the EU, Empirical Economics, 2011, Vol. 40, issue 3, pp. 839-854.

Kravis, I., Lipsey, R. (1978), Price behaviour in the light of balance of payments theories. Journal of International Economics, 8, 193-246.

Kravis, I., Lipsey, R. (1983), Toward an explanation of national price levels, International Finance Section, Dept. of Economics, Princeton University (Princeton, N.J.), Vol. 52.

Kreinin M. E. (1967), Price elasticities in international trade, Review of Economics and Statistics, 49, pp. 510-516.

Kreinin M. E. (1977), The effect of exchange rate changes on the prices and volume of foreign trade, IMF Staff Papers, 24, pp. 207-329.

Krueger A. O. (1984), Trade policies in developing countries, Handbook of International Economics, Vol. 1, pp. 519-569.

Kubota, M. (2009), Real exchange rate misalignments: theoretical modelling and empirical evidence. Discussion Papers in Economics. York: University of York.

Levich R. M. (1985), Empirical studies of exchange rates: price behavior, rate determination and market efficiency, Handbook of International Economics, Vol. 2, pp. 980-1036.

Lütkepohl H. (2005), Structural Vector Autoregressive Analysis for Cointegrated Variables EUI Working Paper ECO No. 2005/2.

MacKinnon J. G. (2010), Critical Values for Cointegration Tests, Queen's Economics Department Working Paper No. 1227, Queen's University.

Maddala G. S., Wu S. (2000), A comparative study of unit root tests with panel data and a new simple test, Oxford Bulletin of Economics and Statistics, Special Issue.

Marquez J., McNeilly C (1988), Income and price elasticities for exports of developing countries, Review of Economics and Statistics, Vol. 7, pp.306-314.

Marquez J. (1999), Long-period trade elasticities for Canada, Japan, and the United States, Review of International Economics, 7 (1), pp. 102-116.

Narayan, P. K. (2008), The purchasing power parity revisited: new evidence for 16 OECD countries from panel unit root tests with structural breaks. Journal of International Financial Markets, Institutions and Money, 18(2), 137-146.

Nerlove M. (1958), Distributed Lags and Demand Analysis for Agricultural and Other Commodities, Agricultural Marketing Service, United States Department of Agriculture.

OECD (2000), Modelling manufacturing export volumes equations: a system estimation approach, by Murata K, Turner D, Rae D, Le Fouler L. OECD Economics Department working papers, no. 235. OECD Publishing.

OECD (2005), The new OECD international trade model. Economics Department working papers, no. 440. OECD Publishing, by Pain N, Mourougane A, Sédillot F, Le Fouler.

OECD (2010), The OECD'S new global model. Economics Department working papers, no. 768. OECD Publishing, by Hervé K, Pain N, Richardson P, Sédillot F, Beffy P.

Orcutt G. H. (1950), Measurement of price elasticities in International trade, The Review of Economics and Statistics, Vol. XXXII, No. 2, pp. 117-132.

Pedroni P. (1999), Critical values for cointegration tests in heterogeneous panels with multiple regressors, Oxford Bulletin of Economics and Statistics, Special Issue.

Pertrović P., Gligorić M., (2009), Exchange rate and trade balance: J-curve effect, Panoeconomicus, 2010, 1, pp.23-41.

Pesaran, M. Hashem, Ron P. Smith, and Kyung So Im. (1996), "Dynamic Linear Models for Heterogenous Panels." in Laszlo Mátyás, and Patrick Sevestre. (eds.), Econometrics of Panel Data. Boston; Dordrecht and London: Kluwer Academic.

Pesaran, M. Hashem. (1997), "The Role of Economic Theory in Modeling the Long Run." Economic Journal, 107:178-191.

Pesaran, M. Hashem. (2007), "A Simple Panel Unit Root Test in the Presence of Cross-Section Dependence." Journal of Applied Econometrics, 22(2): 265-312.

Pesaran, M. Hashem, Yongcheol Shin, and Ron P. Smith. (1999), "Pooled Mean Group Estimation of Dynamic Heterogeneous Panel." Journal of the American Statistical Association, 94: 621-634.

Pesaran, M. Hashem, Yongcheol Shin, and Richard J. Smith. (2000). "Bounds Testing Approaches to the Analysis of Level Relationships." University of Cambridge, DAE Working Paper 9907.

Pesaran H. and Shin Y. (1997), An Autoregressive Distributed Lag Modelling Approach to Cointegration Analysis, M. Trinity College, Cambridge, England.

Phillips P.C.B. (1986), Understanding spurious regressions in econometrics, Journal of Econometrics, 33, pp. 331-340.

Phillips P.C.B., Hansen B.E. (1990), Statistical inference in instrumental variables regression with I(1) processes, Review of Economic Studies 57, pp. 99-125.

Podestà F. (2002), "Recent developments in quantitative comparative methodology: The case of pooled time series cross-section analysis." DSS Papers Soc., 3-02.

Reinhart C. (1995), Devaluation, relative prices and international trade, Staff Papers, Vol. 42, pp.290-312.

Rodrik, Dani. 2008. “The Real Exchange Rate and Economic Growth: Theory and Evidence.” Brookings Papers on Economic Activity, Fall Conference Draft.

Sawyer, W. C. and Sprinkle R. L. (1996), The demand for imports and exports in the U.S.: A survey, Journal of Economics and Finance, Vol. 20, No. 1, pp.147-178.

Senhadji A. S., Montenegro C. E. (1999), Time series Analysis of Export Demand Equations: a Cross-Country Analysis, IMF Staff Papers Vol. 46, No. 3., International Monetary Fund.

Shigeyuki H. and Yoichi M. (2009), Empirical Analysis of Export Demand Behavior of LDCs: Panel Cointegration Approach, Kobe University MPRA Paper No. 17316.

Sims C. A. (1980), Macroeconomics and reality, Econometrica, vol. 48, n° 1 pp. 1-48.

Stern R. M., et al. (1976), Price elasticities in International Trade, Trade Policy Research Centre, London.

Stock, J. H., Watson M.W. (2003), Introduction to econometrics, 1st edition, published by Pearson.

Thorbecke W. (2010), How elastic is Asian Demand for consumption goods?, ADB Institute Working Paper Series, No. 260.

Turner D., Richardson P., Rauffet S. (1996), Modelling the supply side of seven major OECD economies, OECD Economics Department Working Paper, No. 167.

U.S. Department of the Treasury Office of International Affairs (2013), Report to Congress on International Economic and Exchange Rate Policies.

Wooldridge, J. M. (2002a), Econometric Analysis of Cross Section and Panel Data, Cambridge, MA: MIT Press.

Wooldridge, J. M. (2002b), Introductory Econometrics, A Modern Approach, Mason, OH: South-Western, second ed.

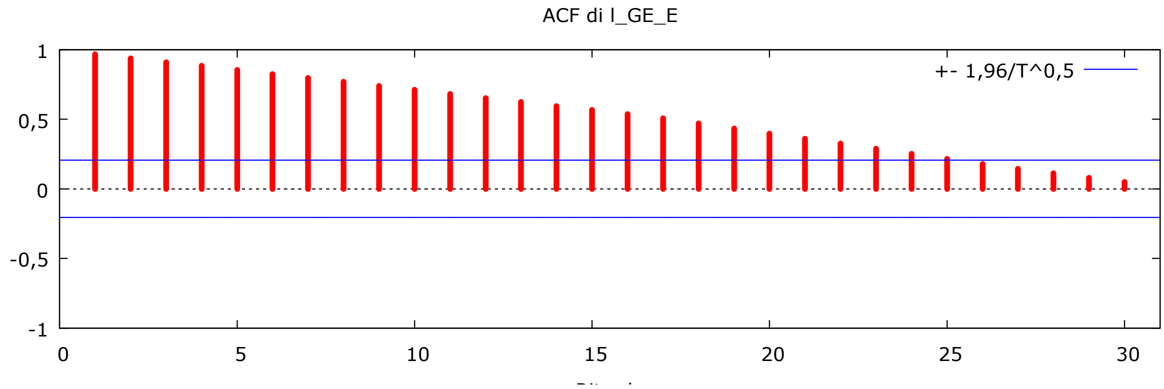
Wu, Y. (1996) Are real exchange rates nonstationary? Evidence form a panel-data test, Journal of Money, Credit and Banking, 28, 54-63.

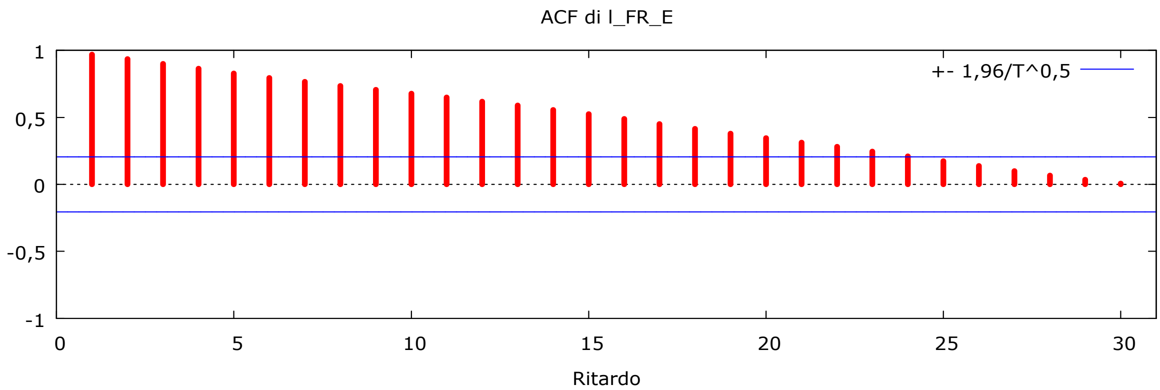
Yalta, A. T., Yalta A. Y. (2007), 'GRETL 1.6.0 and its numerical accuracy', Journal of Applied, Econometrics 22: 849–854.

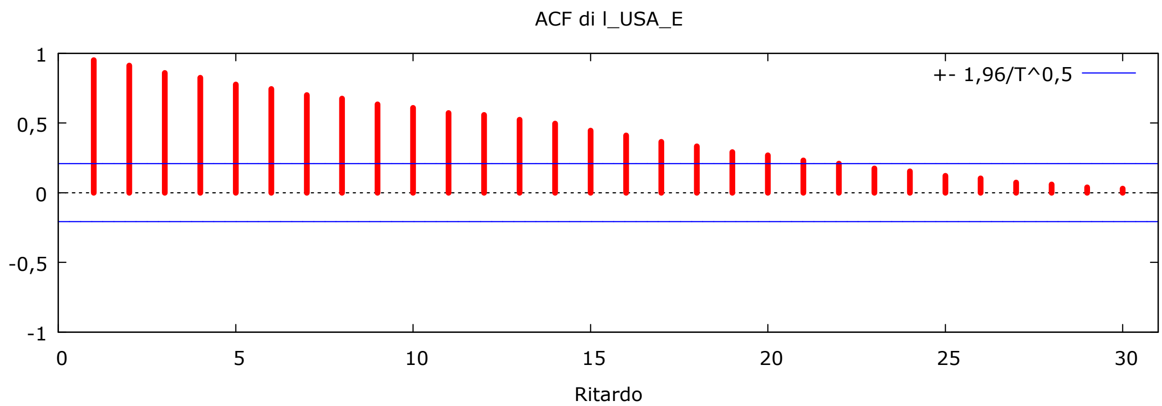
APPENDIX

APPENDIX 1

1.A) Visual inspection of the data for Germany, France, USA, Japan, UK and China:







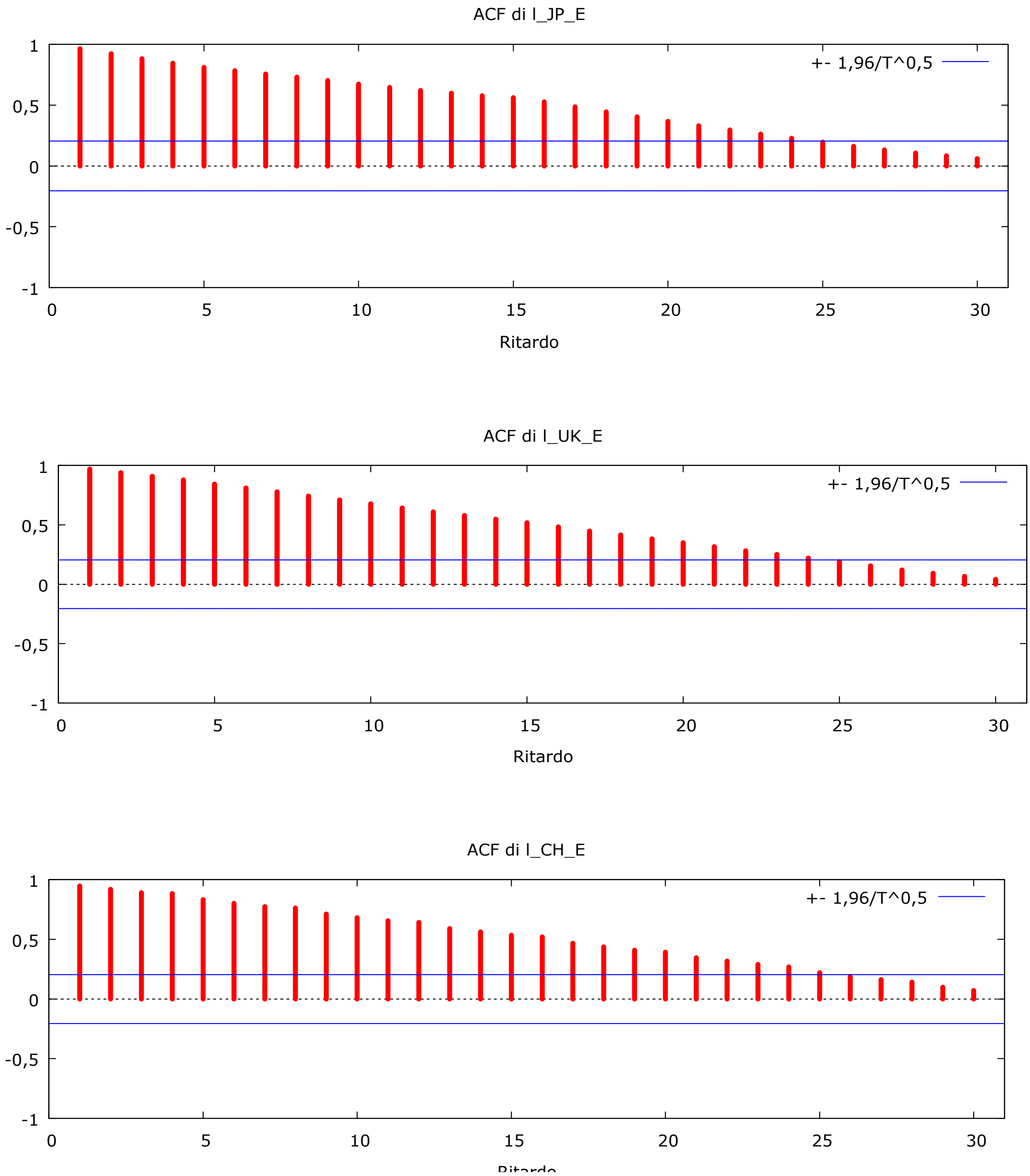
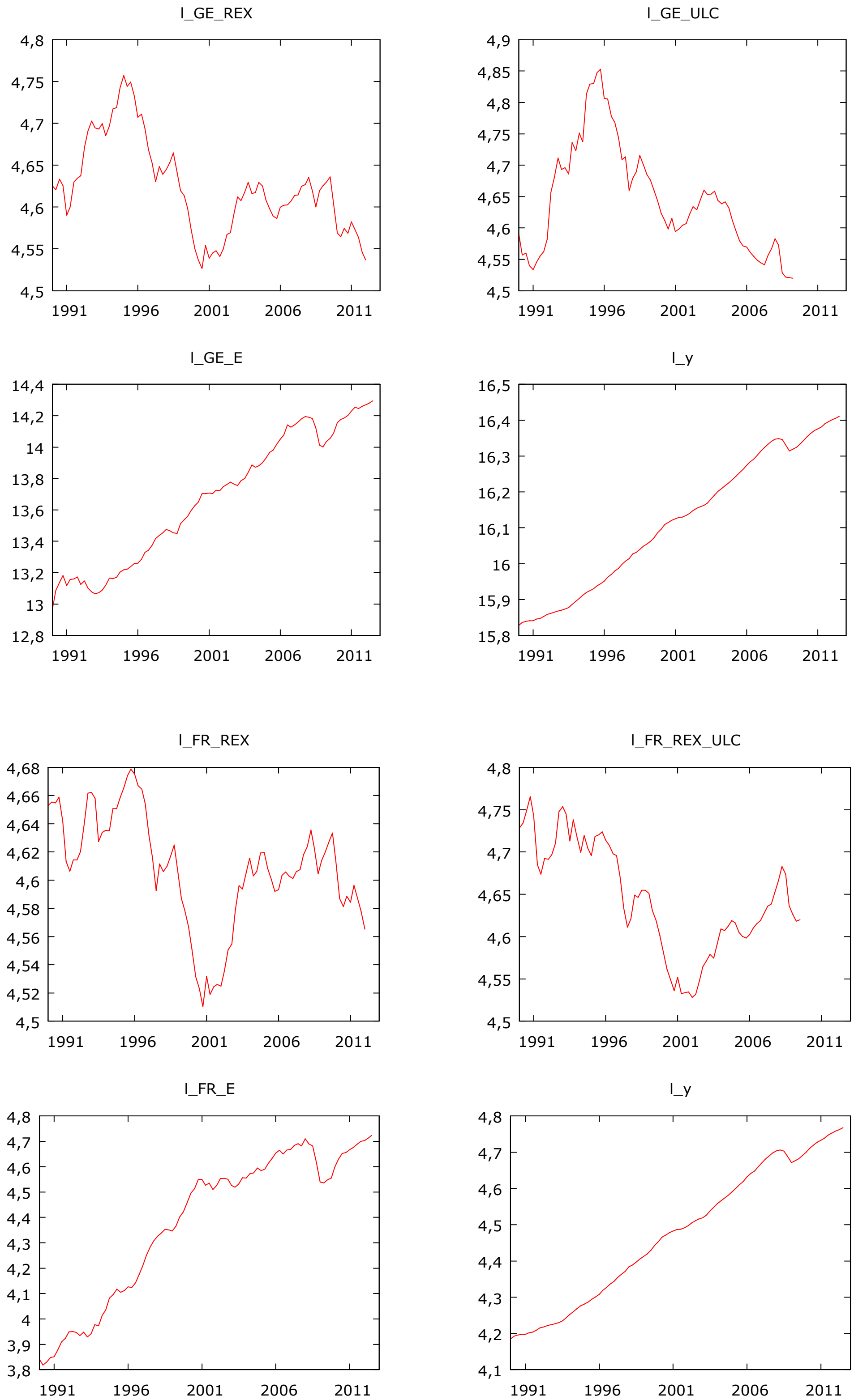
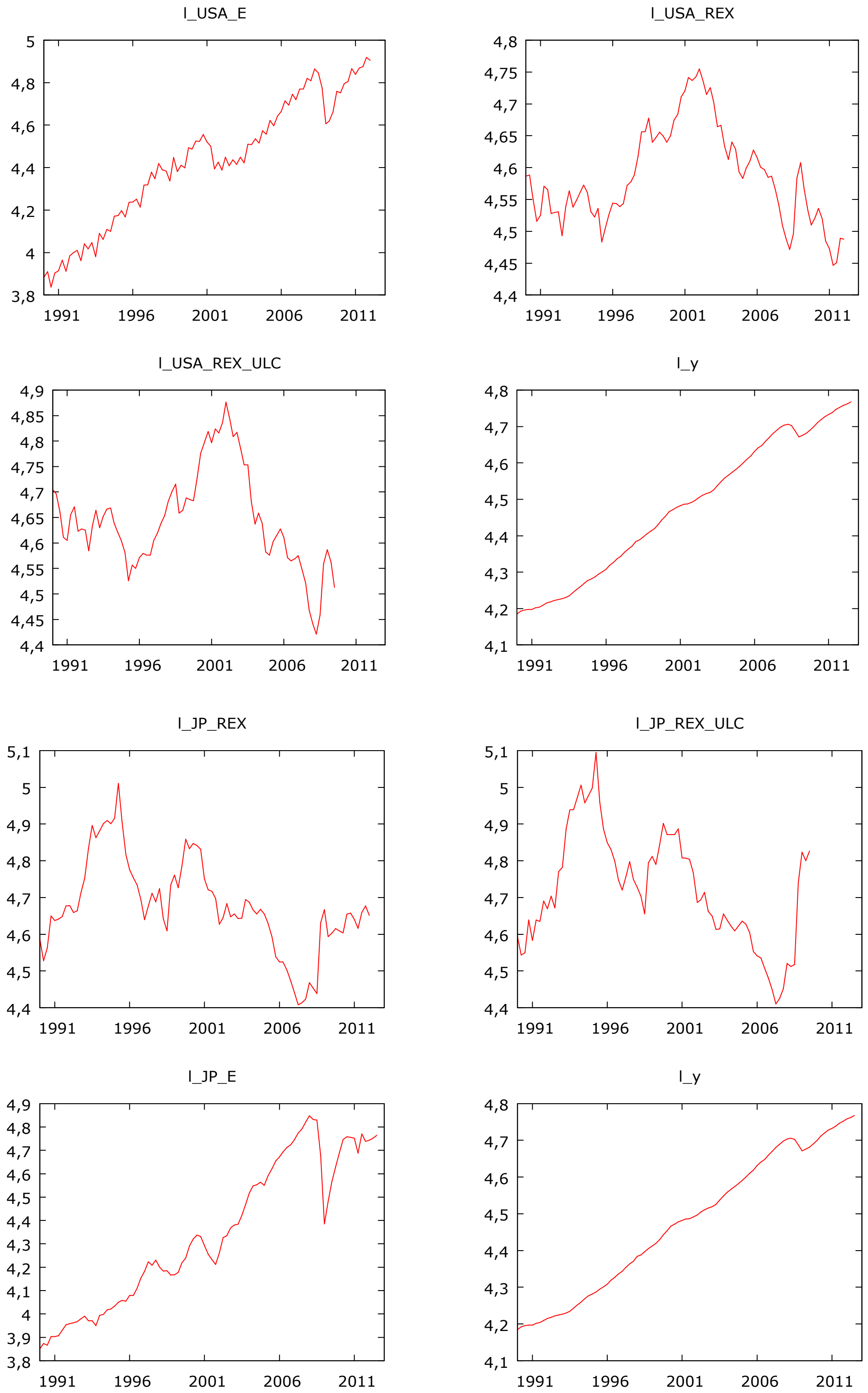


Figure 1.A: ACF for exports. Source: Author's elaboration on Datastream and IFS Databases.

1.B) Plots of world income and of CPI-based and ULC-based exchange rates and exports for Germany, France, USA, Japan, UK and China:




l_USA_E
5 4,8 4,6 4,4 4,2 4 3,8
1991 1996 2001 2006 2011
l_USA_REX
4,8 4,75 4,7 4,65 4,6 4,55 4,5 4,45 4,4
1991 1996 2001 2006 2011
l_USA_REX_ULC
4,9 4,85 4,8 4,75 4,7 4,65 4,6 4,55 4,5 4,45 4,4
1991 1996 2001 2006 2011
l_y
4,8 4,7 4,6 4,5 4,4 4,3 4,2 4,1
1991 1996 2001 2006 2011
l_JP_REX
5,1 5 4,9 4,8 4,7 4,6 4,5 4,4
1991 1996 2001 2006 2011
l_JP_REX_ULC
5,1 5 4,9 4,8 4,7 4,6 4,5 4,4
1991 1996 2001 2006 2011
l_JP_E
4,9 4,8 4,7 4,6 4,5 4,4 4,3 4,2 4,1 4 3,9 3,8
1991 1996 2001 2006 2011
l_y
4,8 4,7 4,6 4,5 4,4 4,3 4,2 4,1
1991 1996 2001 2006 2011

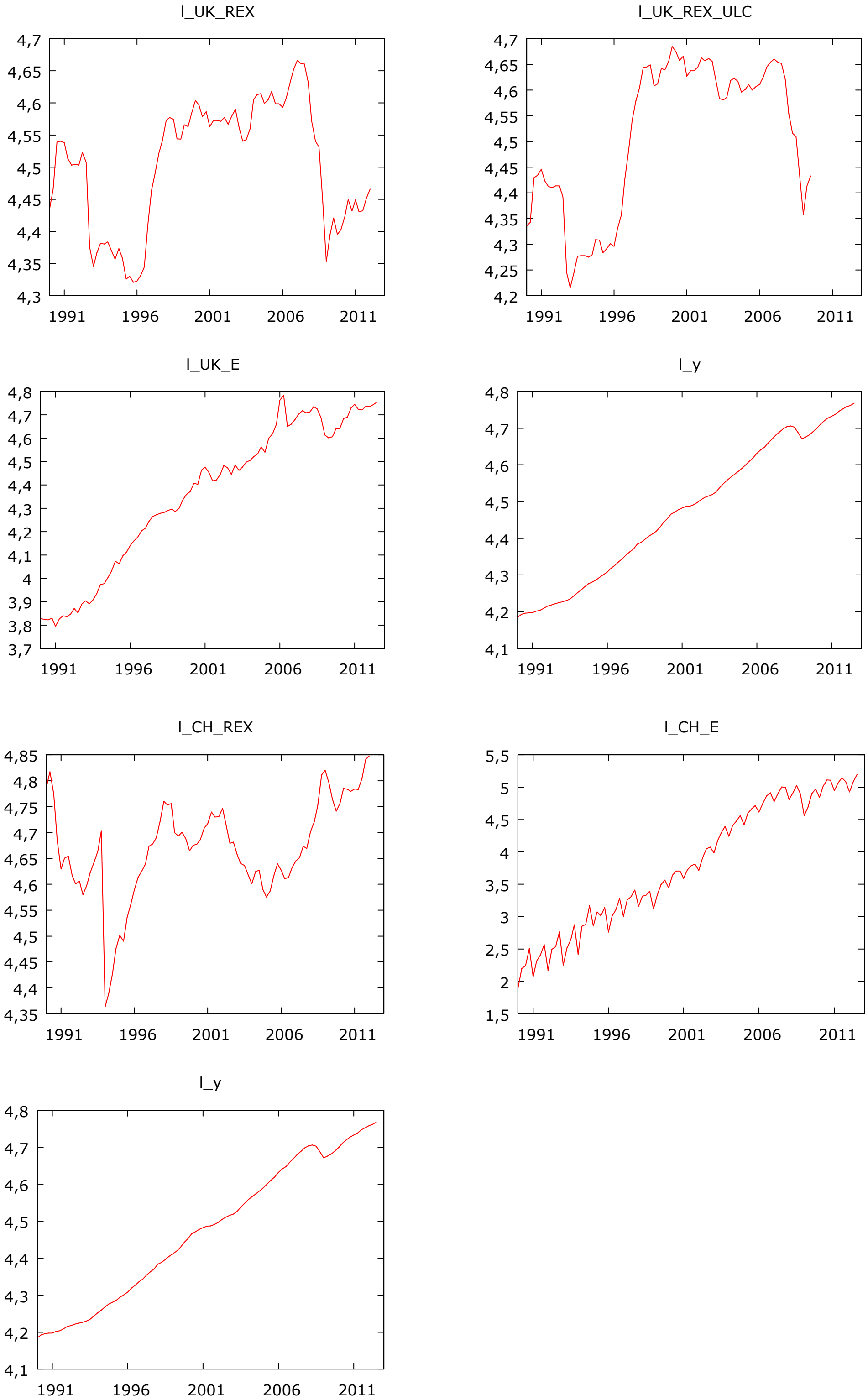
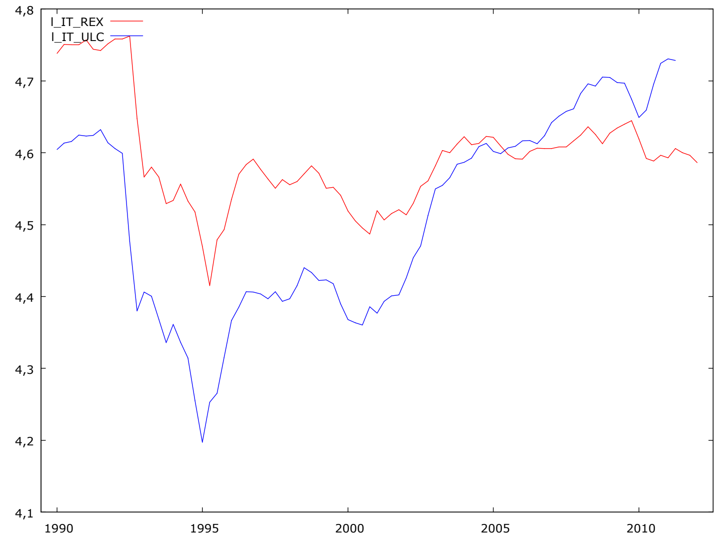


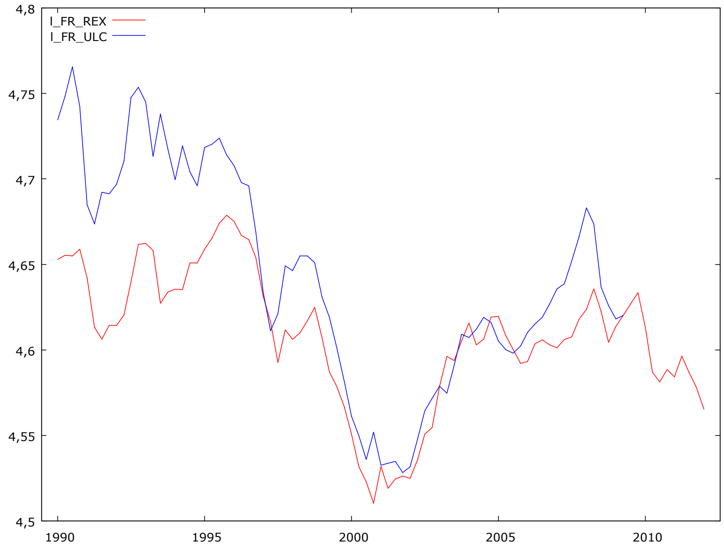
Figure 1.B: Graphs of the time series variables.
Source: Author's elaboration on Datastream and IMF Databases.

1.C) Plots of the real effective exchange rates (CPI and ULC based) for Italy,Germany, France, USA, Japan, UK and China:

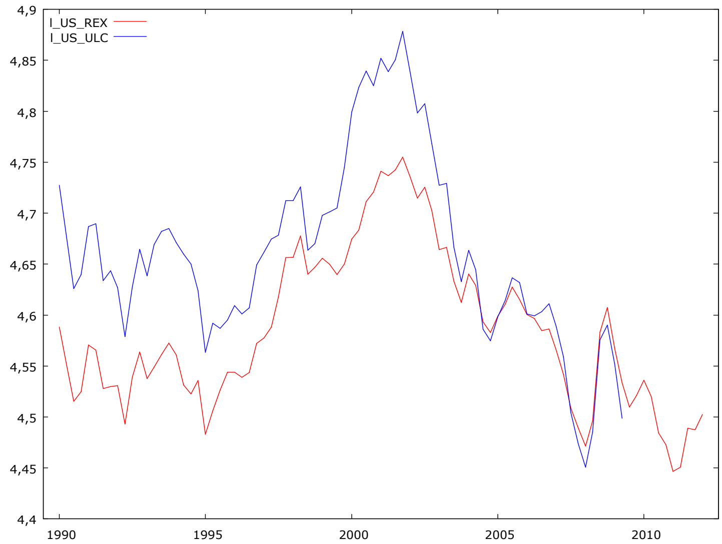


Italy

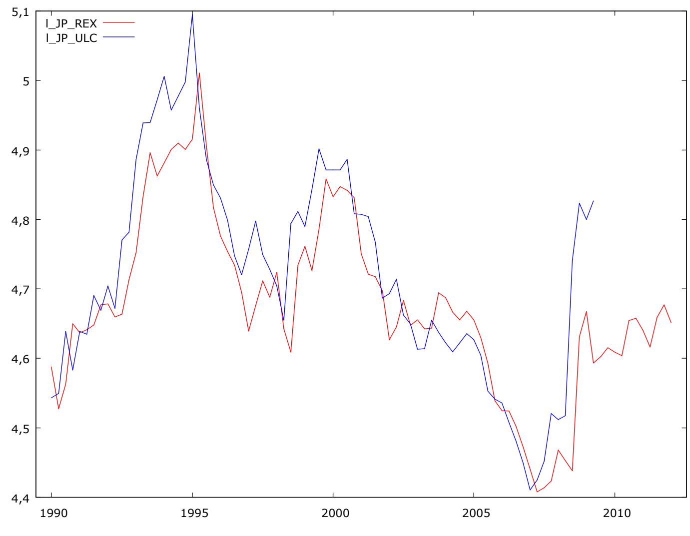
Germany



France

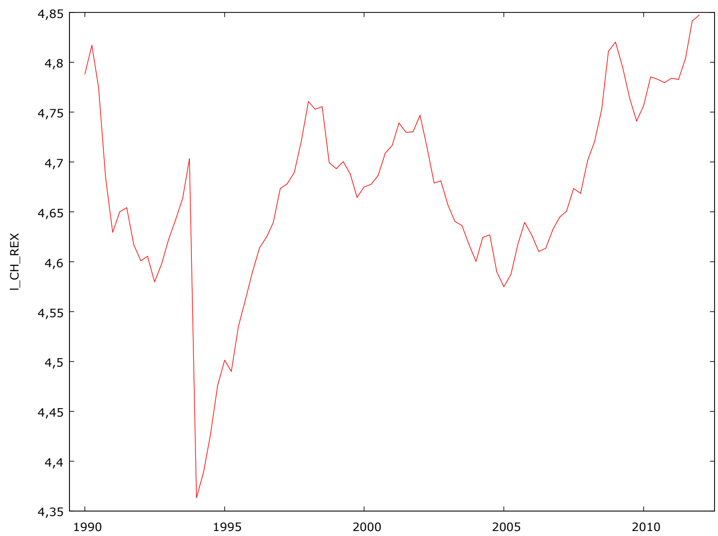


USA



Japan

UK



China

Figure 1.C: Plot of the variables and Real Effective Exchange Rate based on CPI and on ULC.
Source: Author's elaboration.

1.D) Descriptive statistics of the variables for Italy, Germany, France, USA, Japan, UK and China:

Variable	Mean	Median	Minimum	Maximum	Std. Dev.	Var. coeff.	Asymmetry	Curtosis
l_IT_E	4,44171	4,50553	3,9527	4,76552	0,23398	0,05268	-0,6981	-0,569
l_IT_REX	4,59489	4,59178	4,4151	4,76226	0,07352	0,016	0,69256	0,51748
l_IT_REX_ULC	4,51389	4,55472	4,19705	4,73057	0,13832	0,03064	-0,1873	-1,2213
l_GE_REX	4,62409	4,61947	4,52634	4,75729	0,05522	0,01194	0,48539	-0,3082
l_GE_ULC	4,64505	4,63618	4,51961	4,85281	0,08648	0,01862	0,63491	-0,3521
l_GE_E	13,6691	13,7079	12,966	14,2954	0,41004	0,03	-0,0564	-1,4158
l_FR_REX	4,60781	4,60996	4,5102	4,6787	0,0399647	0,00867326	-0,536431	-0,0906746
l_FR_ULC	4,64535	4,63667	4,52829	4,76559	0,0644898	0,0138826	-0,0466113	-0,990702
l_FR_E	4,37743	4,52569	3,81819	4,72364	0,289851	0,0662149	-0,597264	-1,1416
l_USA_E	4,42494	4,43628	3,83698	4,91811	0,297136	0,0671502	-0,22772	-0,957968
l_USA_REX	4,58586	4,57261	4,44657	4,75483	0,0755895	0,0164832	0,442474	-0,57782
l_USA_ULC	4,64524	4,63453	4,42118	4,87602	0,0970655	0,0208957	0,317851	-0,0705412
l_JP_REX	4,67713	4,65937	4,40757	5,01057	0,12631	0,0270058	0,127295	0,0203702
l_JP_ULC	4,72141	4,71402	4,41037	5,09498	0,153618	0,0325364	0,0978264	-0,613977
l_JP_E	4,34199	4,29294	3,85003	4,84774	0,308248	0,0709924	0,122019	-1,36302
l_UK_REX	4,50463	4,53914	4,32082	4,66645	0,0977882	0,0217084	-0,380447	-1,10351
l_UK_ULC	4,50898	4,58395	4,21509	4,68491	0,145935	0,0323654	-0,515539	-1,2759
l_UK_E	4,35984	4,44467	3,7943	4,78379	0,313297	0,0718598	-0,42604	-1,14532
l_CH_REX	4,66973	4,6733	4,36335	4,84749	0,0945882	0,0202556	-0,704007	1,04158
l_CH_E	3,78938	3,71494	1,88834	5,1992	0,971298	0,256321	-0,135171	-1,30559
l_y	4,47521	4,48611	4,18454	4,76789	0,18733	0,0418595	-0,0475016	-1,37176

Table 1.D: Descriptive statistics; Sample period: 1990:1 - 2012:3. Source: Author's elaboration on Datastream and IFS Databases.

APPENDIX 2

2.A) ADF Unit Root Tests of Stationarity for Germany, France, USA, Japan, UK and China:

Variable	Variant	τ –statistic ADF level	p-value*	τ –statistic ADF first difference	p-value*
l_GE_X	Constant, no trend (τc)	-0,43514	0,9009	-4,98326	0,000353
	Constant and trend (τct)	-2,34585	0,4084	-4,9514	0,000000
	No constant (τnc)	3,24399	0,9998	-0,51066	0,0001
l_GE_REX	Constant, no trend (τc)	-1,87141	0,3462	-4,03643	0,001232
	Constant and trend (τct)	-2,75233	0,2154	-4,06476	0,007022
	No constant (τnc)	-0,565855	0,4724	-4,03214	0,000000
l_GE_REX_ULC	Constant, no trend (τc)	-1,85793	0,3527	-4,49938	0,0001
	Constant and trend (τct)	-2,01076	0,5949	-4,57413	0,00109
	No constant (τnc)	-0,371096	0,5511	-4,5205	0,000000
l_FR_E	Constant, no trend (τc)	-1,973	0,2991	-5,06985	0,000000
	Constant and trend (τct)	-1,09011	0,9293	-5,39984	0,000000
	No constant (τnc)	2,77432	0,9988	-1,68596	0,086950
l_FR_REX	Constant, no trend (τc)	-1,99762	0,2881	-6,74981	0,000000
	Constant and trend (τct)	-2,16242	0,5101	-6,70888	0,000000
	No constant (τnc)	-0,63333	0,4433	-6,74735	0,000000
l_FR_REX_ULC	Constant, no trend (τc)	-1,85078	0,3561	-6,41244	0,000000
	Constant and trend (τct)	-1,88406	0,6628	-6,41638	0,000000
	No constant (τnc)	-0,61705	0,4504	-6,41262	0,000000
l_USA_E	Constant, no trend (τc)	-0,03903	0,9539	-3,76467	0,003310
	Constant and trend (τct)	-3,12653	0,1	-3,75923	0,018650
	No constant (τnc)	2,5127	0,9974	-2,26578	0,022650
l_ USA _REX	Constant, no trend (τc)	-1,05267	0,7364	-2,43965	*0,130800*
	Constant and trend (τct)	-0,87599	0,9571	-7,63994	0,000000
	No constant (τnc)	-0,34589	0,5608	-2,43891	0,014260

l_ USA _REX_ULC	Constant, no trend (τc)	-1,70342	0,4296	-6,40555	0,000000
	Constant and trend (τct)	-1,70458	0,7496	-6,39371	0,000000
	No constant (τnc)	-0,52161	0,491	-6,41927	0,000000
l_JP_E	Constant, no trend (τc)	-0,6961	0,846	-3,96515	0,001611
	Constant and trend (τct)	-2,11656	0,536	-3,95401	0,010130
	No constant (τnc)	2,43926	0,9968	-7,04455	0,000000
l_JP_REX	Constant, no trend (τc)	-2,06844	0,2577	-3,51542	0,007630
	Constant and trend (τct)	-3,93359	0,01082	-3,43752	0,046490
	No constant (τnc)	-0,04026	0,6694	-3,54234	0,000391
l_ JP_REX_ULC	Constant, no trend (τc)	-1,14231	0,7012	-2,79499	0,058950
	Constant and trend (τct)	-3,18786	0,0868	-3,64337	0,026280
	No constant (τnc)	0,43145	0,8069	-2,393	0,016170
l_UK_E	Constant, no trend (τc)	-2,00246	0,286	-4,0191	0,001315
	Constant and trend (τct)	-1,40523	0,8598	-4,45287	0,001746
	No constant (τnc)	1,90304	0,9868	-1,6311	*0,097180*
l_UK_REX	Constant, no trend (τc)	-2,0531	0,2642	-4,25818	0,000520
	Constant and trend (τct)	-2,19001	0,4946	-4,21154	0,004235
	No constant (τnc)	-0,21805	0,608	-4,28032	0,000000
l_ UK_REX_ULC	Constant, no trend (τc)	-1,68906	0,4369	-5,79992	0,000000
	Constant and trend (τct)	-1,40487	0,8599	-5,86825	0,000000
	No constant (τnc)	0,170702	0,7358	-5,83235	0,000000
l_CH_E	Constant, no trend (τc)	-1,01863	0,7489	-3,19051	0,020570
	Constant and trend (τct)	-1,33535	0,8789	-3,60704	0,029160
	No constant (τnc)	1,29709	0,9513	-2,70457	0,006640
l_CH_REX	Constant, no trend (τc)	-2,05186	0,2646	-9,05044	0,000000
	Constant and trend (τct)	-2,95042	0,1523	-9,19407	0,000000
	No constant (τnc)	0,101098	0,7122	-9,10316	0,000000

l_y	Constant, no trend (τc)	-0,18459	0,9381	-4,08991	0,001640
	Constant and trend (τct)	-3,03175	0,1234	-4,0673	0,009920
	No constant (τnc)	3,27755	0,9998	-2,28942	0,022080

Table 2.A: ADF Unit root tests: comparative settings. MacKinnon (1996) critical values for the null hypothesis H_0 = presence of a unit root. Asymptotic p-values.*

2.B) Johansen Cointegration Tests using CPI-base exchange rates for Italy, Germany, France, USA, Japan, UK and China:

Country	Rank	Eigenvalue	Trace test [p-value]	Lmax test [p-value]
Italy	0	0,21118	30,560 [0,0405]	20,164 [0,0672]
	1	0,11414	10,397 [0,2559]	10,302 [0,1963]
	2	0,0011112	0,094509[0,7585]	0,094509 [0,7585]
Germany	0	0,15163	27,272 [0,0192]	20,554 [0,0164]
	1	0,052168	6,7175 [0,3555]	6,6973[0,2837]
	2	0,00016158	0,02020[0,9304]	0,02020[0,9243]
France	0	0,21128	26,576 [0,1153]	20,175 [0,0670]
	1	0,055034	6,4012 [0,6527]	4,8115 [0,7644]
	2	0,018528	1,5897 [0,2074]	1,5897 [0,2074]
USA	0	0,40456	59,212 [0,0000]	44,068 [0,0000]
	1	0,15169	15,144 [0,0550]	13,983 [0,0535]
	2	0,013559	1,1604 [0,2814]	1,1604 [0,2814]
Japan	0	0,18276	23,949 [0,2092]	17,155 [0,1706]
	1	0,075777	6,7939 [0,6074]	6,6981 [0,5335]
	2	0,0011260	0,095760 [0,7570]	0,095760 [0,7570]
UK	0	0,15558	24,484 [0,1865]	14,374 [0,3487]
	1	0,11128	10,110 [0,2773]	10,028 [0,2143]
	2	0,00096505	0,082069 [0,7745]	0,082069 [0,7745]
China	0	0,16735	20,782 [0,3820]	15,567 [0,2616]
	1	0,038535	5,2149 [0,7848]	3,3403 [0,9118]
	2	0,021813	1,8746 [0,1709]	1,8746 [0,1709]

Table 2.B: Johansen Cointegration Tests for model with exchange rates based on the Consumer Price Index

2.C) Johansen Cointegration Tests using ULC-base exchange rates for Italy, Germany, France, USA, Japan, UK and China:

Country	Rank	Eigenvalue	Trace test [p-value]	Lmax test [p-value]
Italy	0	0,21251	29,826 [0,0497]	19,829 [0,0751]
	1	0,10448	9,9974 [0,2861]	9,1588[0,2795]
	2	0,010052	0,83853 [0,3598]	0,83853[0,3598]
Germany	0	0,14262	27,700 [0,0167]	17,541[0,0526]
	1	0,047153	10,159 [0,1124]	5,5063 [0,4171]
	2	0,039988	4,6523 [0,0352]	4,6523 [0,0368]
France	0	0,34297	44,522 [0,0004]	31,502 0,0007]
	1	0,12225	13,020 [0,1142]	9,7799 [0,2315]
	2	0,042284	3,2403 [0,0718]	3,2403 [0,0718]
USA	0	0,43533	54,402 [0,0000]	42,864 [0,0000]
	1	0,12784	11,538 [0,1828]	10,259 [0,1991]
	2	0,016919	1,2798 [0,2579]	1,2798 [0,2579]
Japan	0	0,19847	24,401 [0,1899]	16,593 [0,1996]
	1	0,082638	7,8086 [0,4933]	6,4690 [0,5613]
	2	0,017703	1,3396 [0,2471]	1,3396 [0,2471]
UK	0	0,21313	26,628 [0,1138]	17,977 [0,1343]
	1	0,10687	8,6512 [0,4057]	8,4765 [0,3400]
	2	0,0023261	0,17466 [0,6760]	0,17466 [0,6760]

Table 2.C: Johansen Cointegration Tests for model with exchange rates based on the Unit Labor Cost index.

APPENDIX 3

3.A) Export price and income elasticities estimates using VECM. Summarizing tables:

Country	*Long-run Price elasticity*	*Short-run Price elasticity*	*Long-run Income elasticity*	*Short-run Income elasticity*	*ECM Speed of Adj.*
Italy	**-0,72**	**-0,05**	**1,01**	**3,86**	**-0,25**
s.e.	*-0,141100*	*0,123866*	*-0,053981*	*0,815929*	*0,056667*
Germany	**-0,58**	**-0,25**	**2,56**	**4.32**	**-0,11**
s.e.	*0,26930*	*0,135190*	*0,15096*	*1,10649*	*0,093337*
France	**-1,41**	**-0,27**	**0,10**	**3,63**	**-0,12**
s.e.	*0,407000*	*0,153284*	*0,100790*	*0,681037*	*0,027807*
USA	**-1,21**	**-0,03**	**1,38**	**2,63**	**-0,23**
s.e.	*0,190390*	*0,140229*	*0,068277*	*0,980787*	*0,034103*
Japan	**-0,55**	**-0,23**	**1,34**	**1,66**	**-0,28**
s.e.	*0,123420*	*0,086874*	*0,082945*	*1,319900*	*0,076372*
UK	**-0,84**	**-0,07**	**1,60**	**0,02**	***0,03***
s.e.	*0,233410*	*0,140872*	*0,118090*	*1,392200*	*0,044925*
China	**-1,95**	**-0,27**	**5,58**	**1,87**	**-0,27**
s.e.	*0,307520*	*0,254649*	*0,140780*	*0,229376*	*0,114734*

Table 3.A: VECM system, 4 lags. Obs.: 1990:1-2012:1(T= 85); Cointegration rank =1; Exchange rates on Consumer Price Index bases. Source: Own estimations on Datastream and IFS databases. Notes: Aggregation level: value of goods and services; Index 2005=100.

Country	*Long-run Price elasticity*	*Short-run Price elasticity*	*Long-run Income elasticity*	*Short-run Income elasticity*	*ECM Speed of Adj.*
Italy	**-0,48**	**-0,12**	**1,32**	**0,60**	**-0,28**
s.e.	*-0,083110*	*0,111288*	*-0,065614*	*0,810375*	*0,065019*
Germany	**-0,21**	**-0,17**	**2,06**	**3,06**	**-0,19**
s.e.	*0,106450*	*0,116483*	*0,074202*	*0,828905*	*0,063510*
Francia	**-1,20**	**-0,32**	**1,06**	**3,21**	**-0,23**
s.e.	*0,152520*	*0,114040*	*0,071855*	*0,602382*	*0,042725*
USA	**-0,57**	**0,08**	**1,26**	**2,17**	**-0,34**
s.e.	*0,108390*	*0,119547*	*1,259200*	*1,009050*	*0,048355*
Japan	**-0,37**	**-0,06**	**1,39**	**-0,32**	**-0,34**
s.e.	*0,105960*	*0,084542*	*0,099315*	*1,897500*	*0,087257*
UK	**-0,26**	**-0,09**	**1,64**	**0,73**	**-0,08**
s.e.	*0,099019*	*0,139602*	*0,084836*	*1,605630*	*0,094353*
China	-	-	-	-	-
s.e.	-	-	-	-	-

Table 3.A.(1): VECM system, 4 lags. Obs.: 1990:1-2012:1(T= 85); Cointegration rank =1; Exchange rates on Unit Labor Cost index bases. Source: Own estimations on Datastream and IFS databases. Notes: Aggregation level: value of goods and services; Index 2005=100.

3.B) VECM Estimates for Germany:

Model A with l_GE_REX:

VECM system, 4 lags
1991:1-2011:4 (T = 84)
Cointegration rank = 1

l_GE_X	1,0000
	(0,00000)
l_GE_REX	-0,57633
	(0,26930)
l_y	+2,5645
	(0,15096)

These results indicate that the long-run export price and income elasticities estimates are, respectively: -0,58 and +2,56.

The short-run export price and income elasticities estimates are, respectively: -0,25 and +4.32.

The Error Correction term coefficient is -0,11, it is statistically significant and it exhibits the expected negative sign. As aforesaid, it indicates the speed at which the variables return to equilibrium after departing for the equilibrium path (after a shock, for example). Probably, the positive sign indicates that the variable did not depart from equilibrium but rather has still not reached it.

The Durbin-Watson test is: 1,87 while the Adjusted R^2 is 0,42.

The following graph plots the residuals of the system:

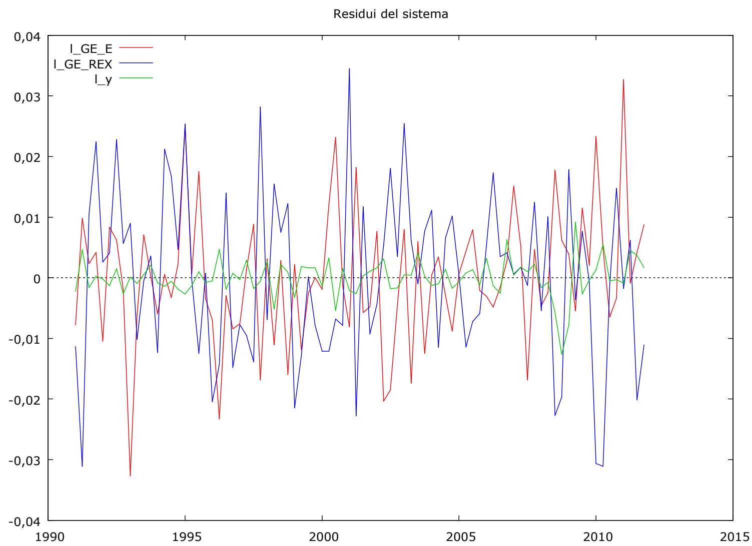


Figure 3.B: Germany, residuals. Source: Author's elaboration.

Model B with l_GE_REX_ULC:

VECM system, 4 lags
1991:1-2009:3 (T = 75)
Cointegration rank = 1

l_GE_X	1,0000
	(0,00000)
l_GE_REX_ULC	-0,21059
	(0,10645)
l_y	+2,0562
	(0,074202)

These results indicate that the long-run export price and income elasticities estimates are, respectively: -0,21 and + 2,06.

The short-run export price and income elasticities estimates are, respectively: -0,17 and +3,06.

The Error Correction term coefficient is -0,19, it is statistically significant and it exhibits the expected negative sign.

The Durbin-Watson test is: 1,98 while the Adjusted R^2 is: 0,47.

Serial correlation test:

Equation 1:

Ljung-Box Q' = 1,67946 with p-value = 0,794

Equation 2:

Ljung-Box Q' = 1,24542 with p-value = 0,871

Equation 3:

Ljung-Box Q' = 0,610665 with p-value = 0,962

The null hypothesis of no serial correlation cannot be rejected (critical value for alpha= 0,05: 0,71).

3.C) VECM Estimates for France:

Model A with l_FR_REX:

VECM system, 4 lags
1991:1-2012:1 (T = 85)
Cointegration rank = 1

l_FR_E	1,0000
	(0,00000)
l_FR_REX	- 1,4069
	(0,40700)
l_y	+0,99785
	(0,10079)

These results indicate that the long-run export price and income elasticities estimates are, respectively: -1, 41 and +1,0.

The short-run export price and income elasticities estimates are, respectively: -0,27 and +3,63.

The Error Correction term coefficient is - 0,12, which is statistically significant and exhibits the expected negative sign.

The Durbin-Watson test is: 2,0 while the Adjusted R^2 is 0,44.

Serial correlation test:

Equation 1:

Ljung-Box Q' = 1,50554 with p-value = 0,826

Equation 2:

Ljung-Box Q' = 0,0826793 with p-value = 0,999

Equation 3:

Ljung-Box Q' = 0,623104 with p-value = 0,96

The null hypothesis of no serial correlation cannot be rejected (critical value for alpha= 0,05: 0,71).

The following graph plots the residuals of the system:

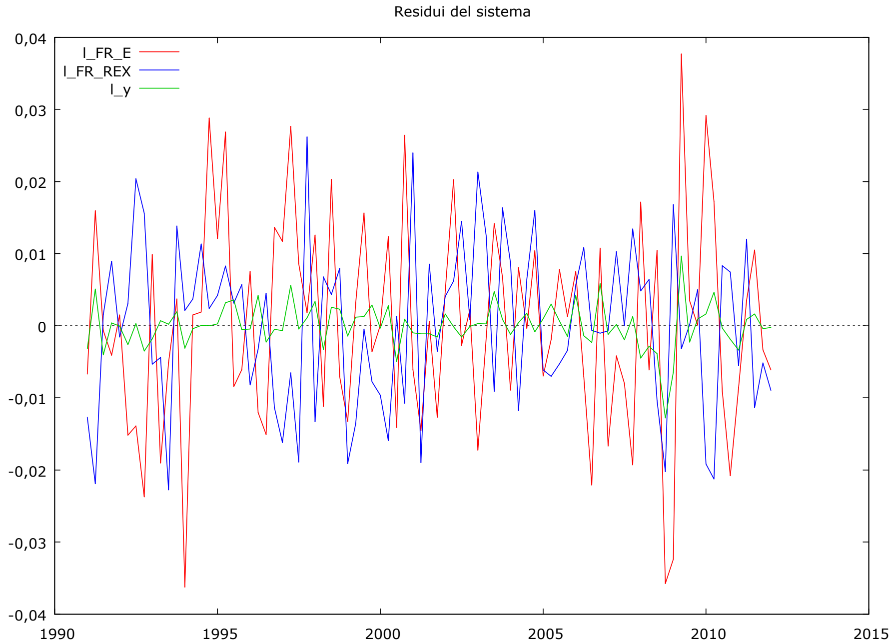


Figure 3.C. France, residuals. Source: Author's elaboration.

Model B with l_FR_REX_ULC:

VECM system, 4 lags
1991:1-2009:3 (T = 75)
Cointegration rank = 1

l_FR_E	1,0000
	(0,00000)
l_FR_REX_ULC	-1,2041
	(0,15252)
l_y	+1,0556
	(0,071855)

These results indicate that the long-run export price and income elasticities estimates are, respectively: -1,20 and +1,06 .

The short-run export price and income elasticities estimates are, respectively: -0,32 and +3,21.

The Error Correction term coefficient is -0,23; it is statistically significant and exhibits the expected negative sign.

The Durbin-Watson test is: 2,1 while the Adjusted R^2 is 0,56.

The following graph plots the residuals of the system for each variable:

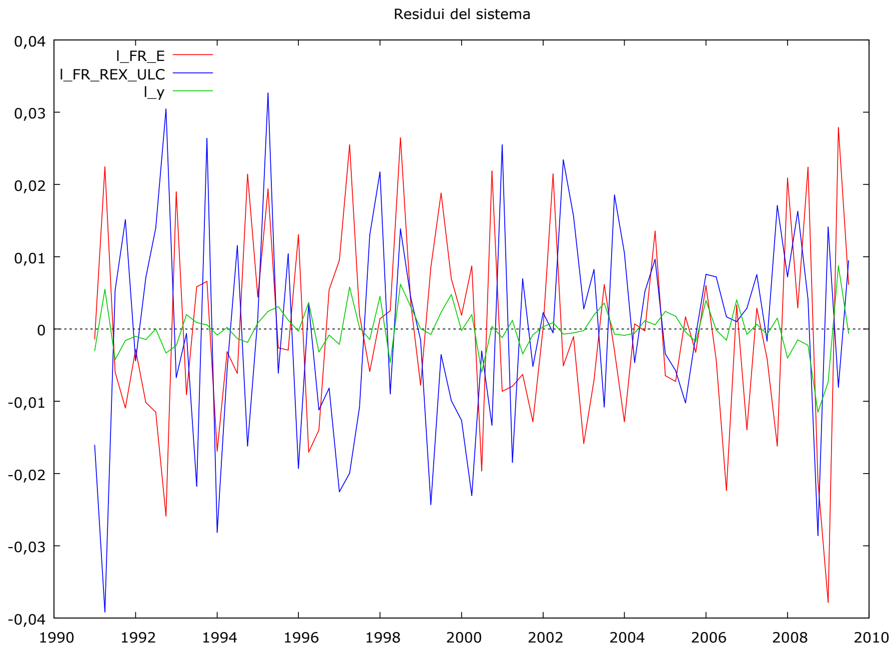


Figure 3.C(1). France, residuals. Source: Author's elaboration.

3.D) VECM Estimates for USA:

Model A with l_USA_REX:

VECM system, 4 lags
1991:1-2012:1 (T = 85)
Cointegration rank = 1

l_USA_E	1,0000
	(0,00000)
l_USA_REX	-1,2057
	(0,19039)
l_y	+1,3804
	(0,068277)

These results indicate that the long-run export price and income elasticities estimates are, respectively: -1,21 and +1,38.

The short-run export price and income elasticities estimates are, respectively: +0,03 and +2,63.

The Error Correction term coefficient is - 0,23; it is statistically significant and exhibits the expected negative sign.

The Durbin-Watson test is: 1,68 while the Adjusted R^2 is 0,73.

The following graph plots the residuals of the system for each variable:

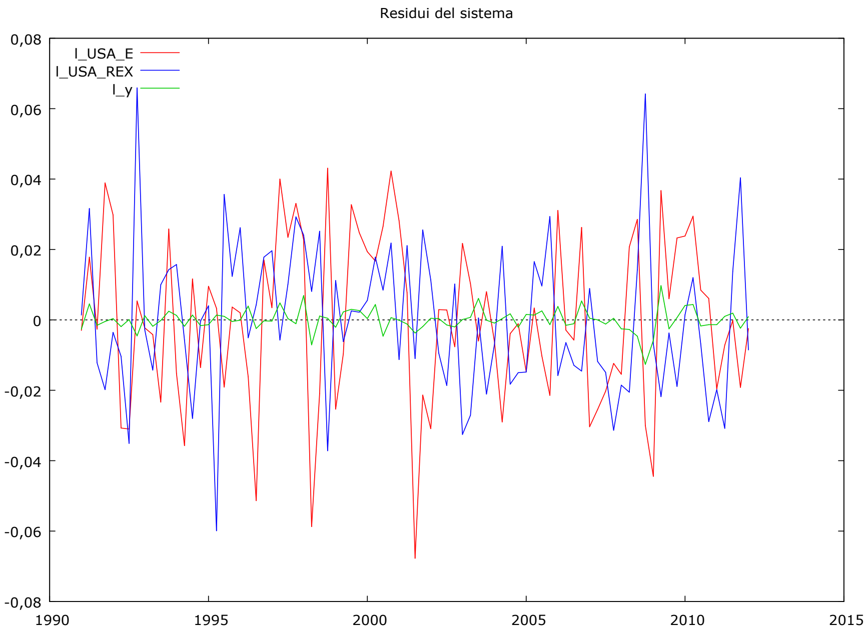


Figure 3.D. USA, residuals. Source: Author's elaboration.

Serial correlation test:

Equation 1:
Ljung-Box Q' = 4,7949 with p-value = 0,309

Equation 2:
Ljung-Box Q' = 2,13921 with p-value = 0,71

Equation 3:
Ljung-Box Q' = 0,825071 with p-value = 0,935

The null hypothesis of no serial correlation of the Q-statistic test cannot be rejected.

Model B with l_USA_REX_ULC:

VECM system, 4 lags
1991:1-2009:3 (T = 75)

Cointegration rank = 1

l_USA_E	1,0000
	(0,00000)
l_USA_REX_ULC	-0,56960
	(0,10839)
l_y	+1,2592
	(0,061851)

These results indicate that the long-run export price and income elasticities estimates are, respectively: -0,57 and +1,26.

The short-run export price and income elasticities estimates are, respectively: +0,08 and +2,17. The short-run price elasticity presents, as we can see, a positive unexpected sign.

The Error Correction term coefficient is -0,34; it is statistically significant and exhibits the expected negative sign.

The Durbin-Watson test is: 1,68 while the Adjusted R^2 is 0,73.

The following graph plots the residuals of the system for each variable:

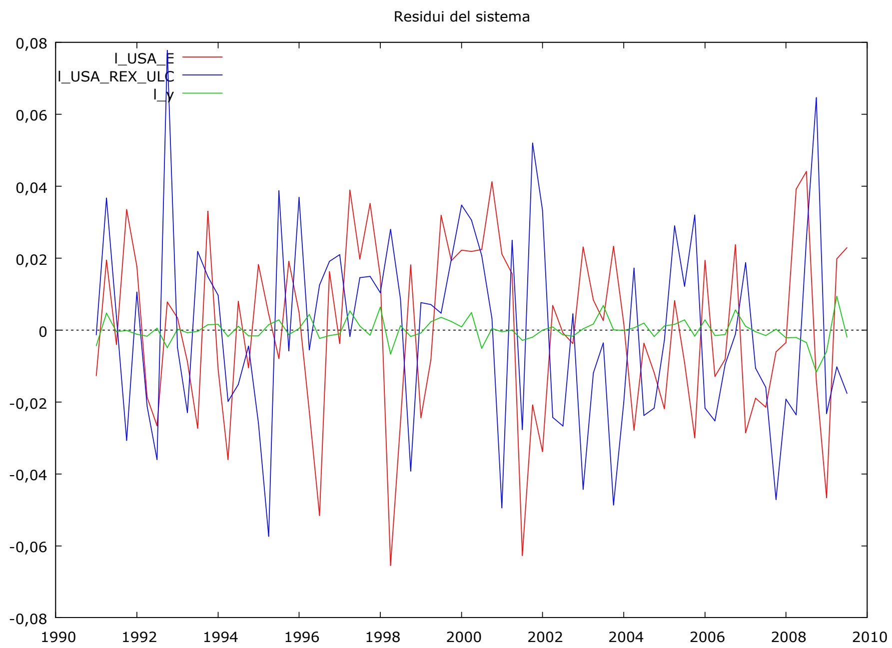


Figure 3.D (1). USA, residuals. Source: Author's elaboration.

3.E) VECM Estimates for Japan:

Model A with l_JP_REX:

VECM system, 4 lags
1991:1-2012:1 (T = 85)
Cointegration rank = 1

l_JP_E	1,0000
	(0,00000)
l_JP_REX	-0,54663
	(0,12342)
l_y	+1,3425
	(0,082945)

These results show that the long-run export price and income elasticities estimates are, respectively: -0,55 and +1,34.

The short-run export price and income elasticities estimates are, respectively: +0,23 and +1,66.

The Error Correction term coefficient is -0,28; it is statistically significant and exhibits the expected negative sign.

The Durbin-Watson test is: 2,02 while the Adjusted R^2 is 0,57.

Serial correlation test:

Equation 1:
Ljung-Box Q' = 0,144175 with p-value = 0,998

Equation 2:
Ljung-Box Q' = 0,201967 with p-value = 0,995

Equation 3:
Ljung-Box Q' = 0,262408 with p-value = 0,992

The null hypothesis of no serial correlation cannot be rejected (critical value for alpha= 0,05: 0,71).

The following graph plots the residuals of the system for each variable:

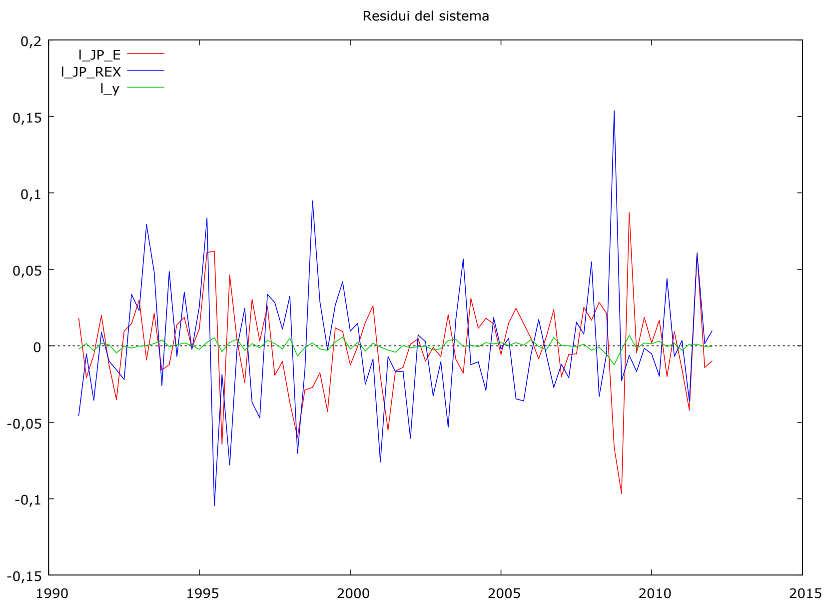


Figure 3.E. Japan, residuals. Source: Author's elaboration.

Model B with l_JP_REX_ULC:

VECM system, 4 lags
1991:1-2009:3 (T = 75)
Cointegration rank = 1

l_JP_E	1,0000
	(0,00000)
l_JP_REX_ULC	-0,37445
	(0,10596)
l_y	+1,3936
	(0,099315)

These results show that the long-run export price and income elasticities estimates are, respectively: -0,37 and +1,39.

The short-run export price and income elasticities estimates are, respectively: -0,06 and +0,32.

The Error Correction term coefficient is -0,34; it is statistically significant and exhibits the expected negative sign.

The Durbin-Watson test is: 1,95 while the Adjusted R^2 is 0,51.

The following graph plots the residuals of the system for each variable:

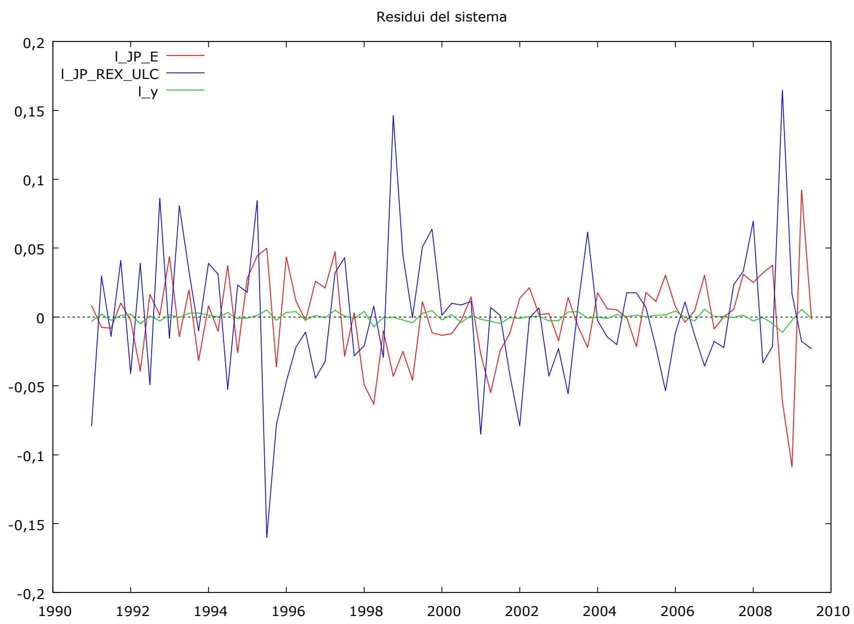


Figure 3.E (1). Japan, residuals. Source: Author's elaboration.

Serial correlation test:

Equation 1:

Ljung-Box Q' = 1,12126 with p-value = 0,891

Equation 2:

Ljung-Box Q' = 4,54334 with p-value = 0,337

Equation 3:

Ljung-Box Q' = 0,910473 with p-value = 0,923

The Q-statistic[73] has the null hypothesis of "no serial correlations" (up to the lags used for the test, which here are 4). Hence, each p-value indicates that there is no serial correlation since you cannot reject the null.

3.F) VECM Estimates for UK:

Model A with l_UK_REX:

VECM system, 4 lags

[73] The critical value is 0,710723 with alpha = 0,05.

1991:1-2012:1 (T = 85)
Cointegration rank = 1

l_UK_E	1,0000
	(0,00000)
l_UK_REX	-0,83580
	(0,23341)
l_y	+1,5977
	(0,11809)

These results show that the long-run export price and income elasticities estimates are, respectively: -0,84 and +1,60.

The short-run export price and income elasticities estimates are, respectively: +0,07 and +0,02.

The Error Correction term coefficient is + 0,03, but it is not statistically significant and does not exhibit the expected negative sign. The Durbin-Watson test is: 1,99.

The following graph plots the residuals of the system for each variable:

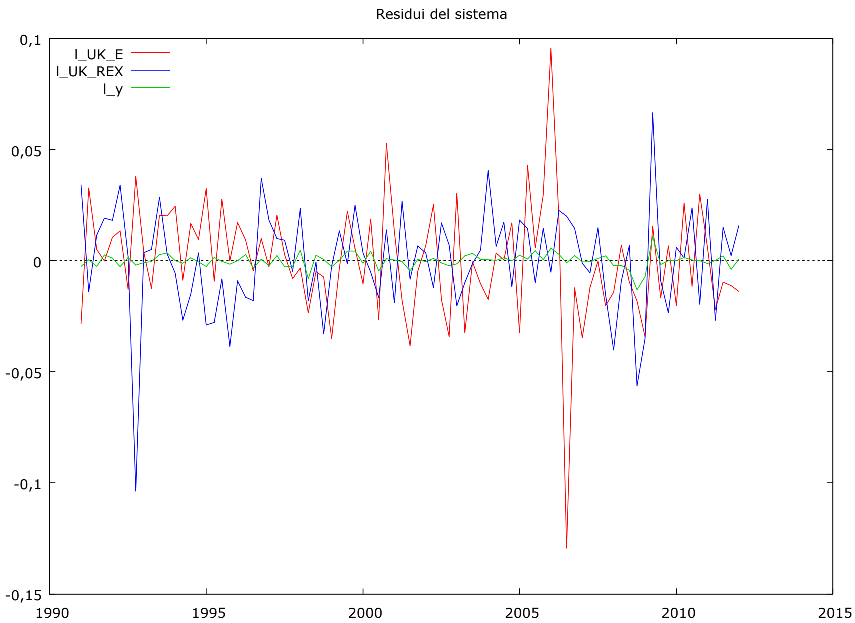


Figure 3.F. UK, residuals. Source: Author's elaboration.

Serial correlation test:

Equation 1:
Ljung-Box Q' = 0,17032 with p-value = 0,997

Equation 2:
Ljung-Box Q' = 0,269179 with p-value = 0,992

Equation 3:
Ljung-Box Q' = 0,802158 with p-value = 0,938

The null hypothesis of no serial correlation test cannot be rejected.

Model B with l_UK_REX_ULC:

VECM system, 4 lags
1991:1-2009:3 (T = 75)
Cointegration rank = 1

l_UK_E	1,0000
	(0,00000)
l_UK_REX_ULC	-0,26215
	(0,099019)
l_y	+1,6438
	(0,084836)

These results show that the long-run export price and income elasticities estimates are, respectively: -0,26 and +1,64.

The short-run export price and income elasticities estimates are, respectively: -0,09 and +0,73.

The Error Correction term coefficient is -0,07, it is statistically significant and it exhibits the expected negative sign. The Durbin-Watson test is 1,99.

The following graphs plot the residuals of the system for each variable:

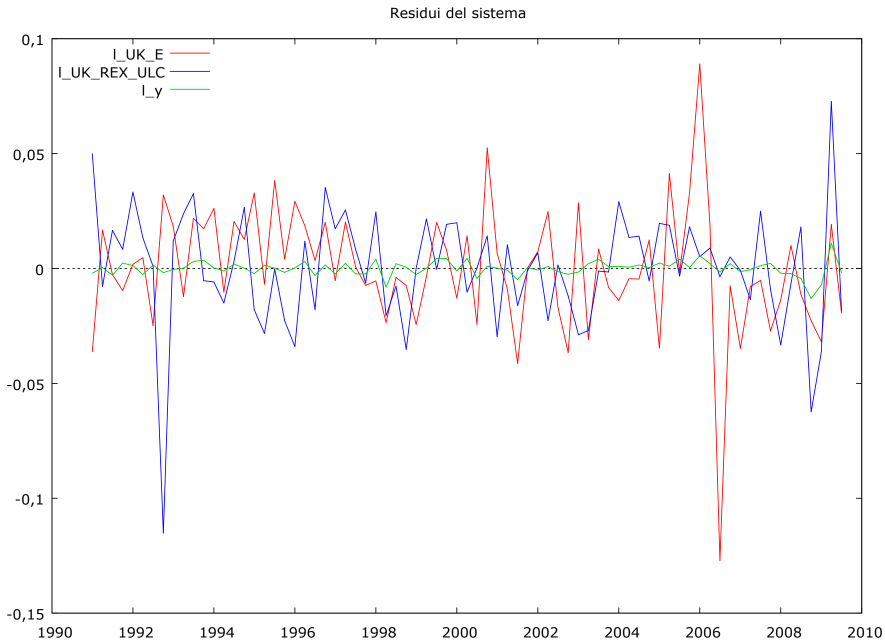


Figure 3.F(1). UK, residuals. Source: Author's elaboration.

Serial correlation test:

Equation 1:
Ljung-Box Q' = 0,143582 with p-value = 0,998

Equation 2:
Ljung-Box Q' = 0,97994 with p-value = 0,913

Equation 3:
Ljung-Box Q' = 0,86938 with p-value = 0,929

According to the Q-statistic serial correlation test, the null hypothesis of no serial correlation can be accepted.

3.G) VECM Estimates for China:

Model with l_CH_REX:

VECM system, 4 lags
1991:1-2012:1 (T = 85)
Cointegration rank = 1

l_CH_E	1,0000 (0,00000)

l_CH_REX -1,9483
(0,30752)
l_y +5,5794
(0,14078)

These results show that the long-run export price and income elasticities estimates are, respectively: -1,95 and +5,58.

The short-run export price and income elasticities estimates are, respectively: +0,27 and +1,87.

The Error Correction term coefficient is -0,27; it is statistically significant and exhibits the expected negative sign. The Durbin-Watson test is: 1,03 and the R2 is 0,82.

Serial correlation test:

Equation 1:
Ljung-Box Q' = 20,8791 with p-value = 0,000335
Equation 2:
Ljung-Box Q' = 0,940484 with p-value = 0,919
Equation 3:
Ljung-Box Q' = 0,4779 with p-value = 0,976.
The following graph plots the residuals of the for each variable:

Figure 3.G. China, residuals. Source: Author's elaboration.

APPENDIX 4

4. A) Preliminary results

Initially, to roughly test the sensitivity of the variables to variations, I analysed the responses of exports to their main determinants, estimating the export demand price elasticity for Italy using IMF quarterly data for the sample period 1990-2010 and applying static and dynamic models.

I used a log-linear demand function that expresses the elasticity of the dependent variables with respect to the independent variable. The variables involved were: Italian export volumes (*LXVOL*, dependent variable), real effective exchange rate (*LREX*), and foreign income (*Y*).

Implementing an OLS regression (*cfr.* § 4 B), I obtained the price (-0.76) and income (1.77) elasticities[74]. Further on, to take into account the adjustment time necessary for the dependent variable (export volumes) to respond to variations in the explanatory variables (that is, relative prices expressed as exchange rates and income), I moved on from a static model to a simple dynamic model introducing time lags in the variables. Precisely, I implemented two different distributed-lag (D-L models) models including variables with different time lags.

The long-run price and income elasticities are defined as the short-term price and income elasticities divided by one, minus the coefficient estimate of the lagged dependent variable[75].

In the first model (D-L model 1), besides the explanatory variables considered in the OLS regression, I introduced a third variable that is the lagged (-1) dependent variable and I obtained the following results:

- the long-run export price elasticity is equal to:

  0.24/ (1 – 0.66) = - **0.71**

- the long-run export income elasticity is equal to:

  0.59/ (1 – 0.66) = + 1.74

When I reiterated the regression (D-L model 2) including other lagged variables and corrections for seasonality I obtained the following results:

- the long-run export price elasticity is equal to:

  0.39 /(1 – 0.76) = **- 1.62**

[74] The results of the preliminary study are tabulated in Appendix 4.

[75] Hamilton (1994).

- the long-run export income elasticity was equal to:
  0.65/ (1 – 0.76) = + 2.71

Obviously, these results (summarized in Table 3.2) were very approximate and raw but still, they represented a solid starting point for my purposes. Additionally, this preliminary analysis gave me the possibility to compare the results provided by an approach ordinarily used in the past estimation (and before econometric sophistication and development) of trade elasticities with the more developed ones used in the present study and reported in the following sections of this chapter[76].

Methodology	Export price elasticity	
OLS	-0.76	
Distributed Lag model (1)	SR -0.25	LR- 0.71
Distributed Lag model (2)	SR -0.39	LR - 1.62

Table 3.2. Export price elasticities; Italy,1990:1-2012:1. Source: Author's own elaborations.

4. B) OLS regression:

Dependent Variable: LXVOL
Method: Least Squares
Date: 10/11/11 Time: 12:05
Sample: 1990Q1 2010Q2
Included observations: 82

	Coefficient	Std. Error	t-Statistic	Prob.
LREX	-0.759651	0.033227	-22.86257	0.0000
LY	1.767712	0.033900	52.14518	0.0000

R-squared	0.958656	Mean dependent var	4.465548
Adjusted R-squared	0.958140	S.D. dependent var	0.212015
S.E. of regression	0.043378	Akaike info criterion	-3.413642
Sum squared resid	0.150532	Schwarz criterion	-3.354941
Log likelihood	141.9593	Hannan-Quinn criter.	-3.390074
Durbin-Watson stat	1.425083		

[76] *Cfr*. Section 4.4, Table 3.9.

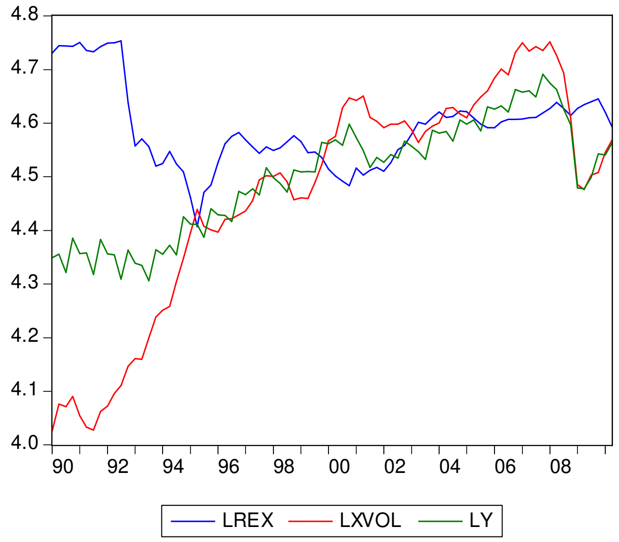


Figure (a) Italian exports, 1990:1-2010:2.
Source: Author's own elaboration on IMF database.

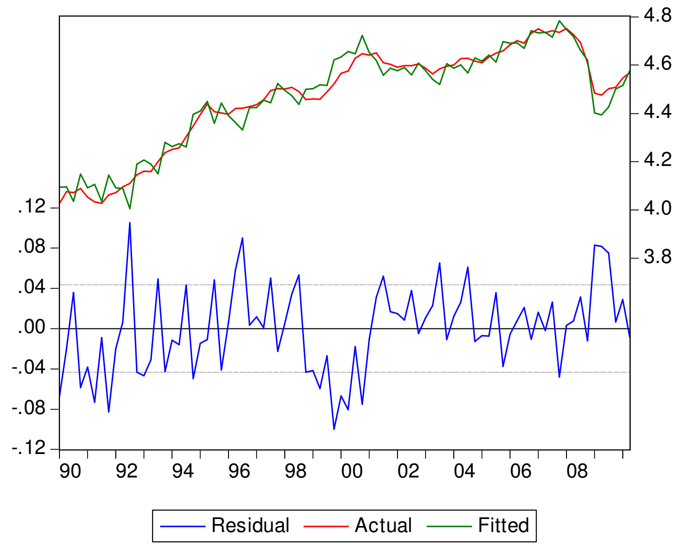


Figure (b) Italian exports, 1990:1-2010:2.
Source: Author's own elaboration on IMF database.

4. C) D-Lag regression (I):

Dependent Variable: LXVOL
Method: Least Squares
Date: 10/11/11 Time: 12:18
Sample (adjusted): 1990Q2 2010Q2

Included observations: 81 after adjustments

	Coefficient	Std. Error	t-Statistic	Prob.
LREX	-0.247260	0.036296	-6.812389	0.0000
LY	0.587837	0.077059	7.628388	0.0000
LXVOL(-1)	0.663006	0.042674	15.53658	0.0000

R-squared	0.989676	Meandependentvar	4.470971
Adjusted R-squared	0.989412	S.D. dependentvar	0.207534
S.E. of regression	0.021355	Akaike info criterion	-4.818705
Sum squaredresid	0.035572	Schwarzcriterion	-4.730021
Log likelihood	198.1575	Hannan-Quinn criter.	-4.783124
Durbin-Watson stat	1.480203		

4. D) D-Lag regression (II):

Dependent Variable: LXVOL
Method: Least Squares
Date: 10/11/11 Time: 12:37
Sample (adjusted): 1990Q4 2010Q2
Included observations: 79 after adjustments

	Coefficient	Std. Error	t-Statistic	Prob.
LREX	-0.393816	0.091207	-4.317833	0.0000
LY	0.645816	0.079587	8.114544	0.0000
LXVOL(-1)	0.762457	0.062030	12.29177	0.0000
LY(-3)	-0.234194	0.072735	-3.219836	0.0019
LREX(-1)	0.220542	0.100133	2.202485	0.0308
@SEAS(3)	0.029475	0.005615	5.249206	0.0000

R-squared	0.992193	Meandependentvar	4.481038
Adjusted R-squared	0.991658	S.D. dependentvar	0.200046
S.E. of regression	0.018271	Akaike info criterion	-5.094059
Sum squaredresid	0.024370	Schwarzcriterion	-4.914101
Log likelihood	207.2153	Hannan-Quinn criter.	-5.021962
Durbin-Watson stat	1.517391		

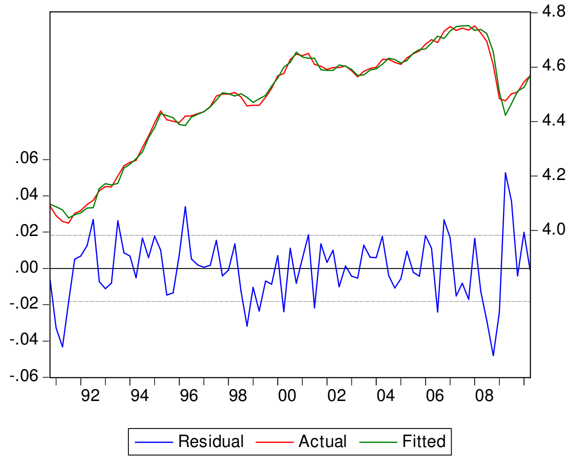


Figure (c) Italian exports, 1990:4-2010:2.
Source: Author's own elaboration on IMF database.

APPENDIX 5

5. A) Descriptive statistics for Italy, France, Germany, UK, Japan, USA and China

Table 5 a. Descriptive statistics, 1990:Q1 - 2012:Q1

Country	Variable	Mean	Median	Minimum	Maximum	Var. coeff.	Asymmetry	Curtosis	Std. Dev.
Italy	import__ita	83,187	88,902	36,076	129,286	30,141	0,362	-0,060	-1,292
	gdp_ita	92,60	96,80	0,000	104,2	0,123	0,132	-5,090	36,070
France	import_fra	82,167	88,362	47,749	117,953	23,516	0,286	-0,112	-1,451
	gdp_fra	92,00	94,20	0,763	106,0	0,102	0,111	-0,162	-1,480
Germany	import_ger	85,329	83,507	46,792	137,839	28,218	0,331	0,271	-1,203
	gdp_ger	95,70	98,40	0,000	111,4	0,132	0,138	-4,406	30,215
UK	import_uk	77,274	81,022	39,989	117,281	24,067	0,311	-0,164	-1,420
	gdp_uk	86,80	88,40	0,629	107,5	0,146	0,168	-0,245	-1,347
Japan	import_jap	88,830	88,730	59,805	111,692	14,042	0,158	-0,191	-0,978
	gdp_japan	82,00	94,70	0,000	105,1	0,351	0,428	-1,887	1,672
USA	import_usa	75,420	80,586	32,260	109,932	26,170	0,347	-0,255	-1,355
	gdp_usa	87,20	89,20	0,622	108,1	0,147	0,169	-0,267	-1,352
China	import_chi	64,915	44,890	9,199	152,597	47,899	0,738	0,446	-1,352
	gdp_china	83,60	67,90	0,241	194,7	0,491	0,587	0,750	-0,623

Table 5 b. Descriptive statistics of imports and gdp variables for Italy, France, Germany, UK, Japan, USA and China. Sample period 1990:Q1 - 2012:Q1.

Table 5a shows the main descriptive statistics of variables included in import functions. The variables are expressed in index numbers based 2005 = 100.

All the variables are non-stationary as shown by plots in figure 5a and in figure 5b and from Levin Lin Chu test in table 5b (we have the same findings for exchange rate – see chapter 3). In addition, the Westerlund test for cointegration between imports and exchange rate is executed in table 5c and cointegration is not-rejected. Table 5g shows the estimates of import functions for Italy, Japan, France, UK, China, Germany and USA over 1990-2012. The expected sign for both the covariates is positive. We find low values of price elasticity, but they are not significative for four countries (Japan, China, Germany and USA). Moreover, when income elasticity is positive, it is also significative (for six countries except China). Long-run elasticity is positive and significative (0.2612). We use this value to test the Marshall-Lerner condition.

5. B) Levin Lin Chu test for Imports, exchange rate and GDP

Table 5 b. Levin Lin Chu test for Imports, exchange rate and GDP

Levin-Lin-Chu test for imports		
Pooled ADF test (1 lag)	N,T = (7,89)	Obs = 609
	coefficient	-0,0425
	p-value	**0,8127**
Levin-Lin-Chu test for exchange rate		
Pooled ADF test (1 lag)	N,T = (7,89)	Obs = 609
	coefficient	-0,0691
	p-value	**0,1857**
Levin-Lin-Chu test for GDP		
Pooled ADF test (1 lag)	N,T = (7,88)	Obs = 602
	coefficient	0,0127
	p-value	**1,0000**

5. C) Westerlund test for imports and exchange rate.

Table 5 c. Westerlund ECM panel cointegration test for imports.

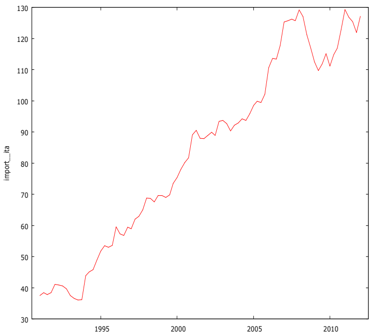
Results for H_0 = NO COINTEGRATION
With 7 series and 1 covariate

Test for cointegration between import and exchange rate - lags(1)

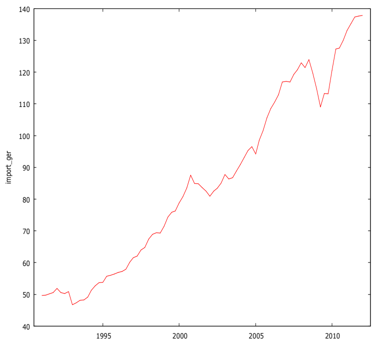
Statistic value	Z-value	p-value
-4,120	-2,372	0,009

5. D) Plots of imports for Italy, France, Germany, UK, Japan, USA and China

Italy



Germany



France

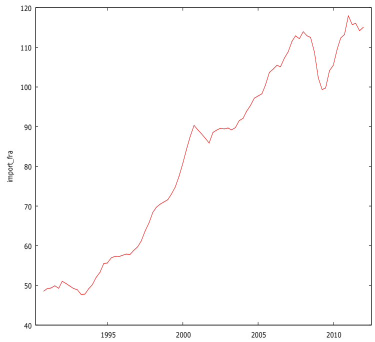
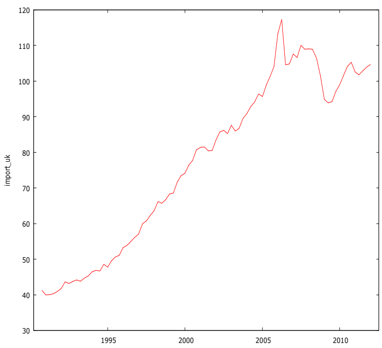
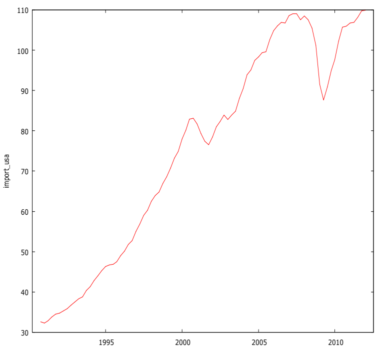


Fig.5.a.

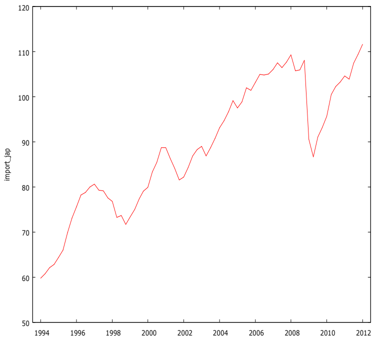
UK



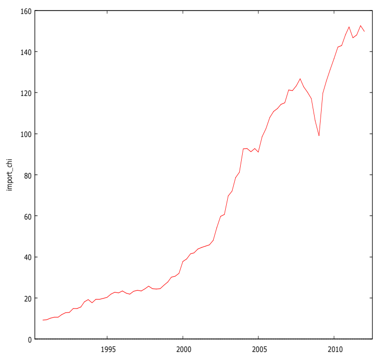
USA



Japan



China



Imports for Italy, Germany, France, UK, USA, Japan and China; 1990:Q1 - 2012:Q1. Source: Elaboration on Datastream databases. Reference year 2005=100.

5. E) Plots of GDP for Italy, France, Germany, UK, Japan, USA and China

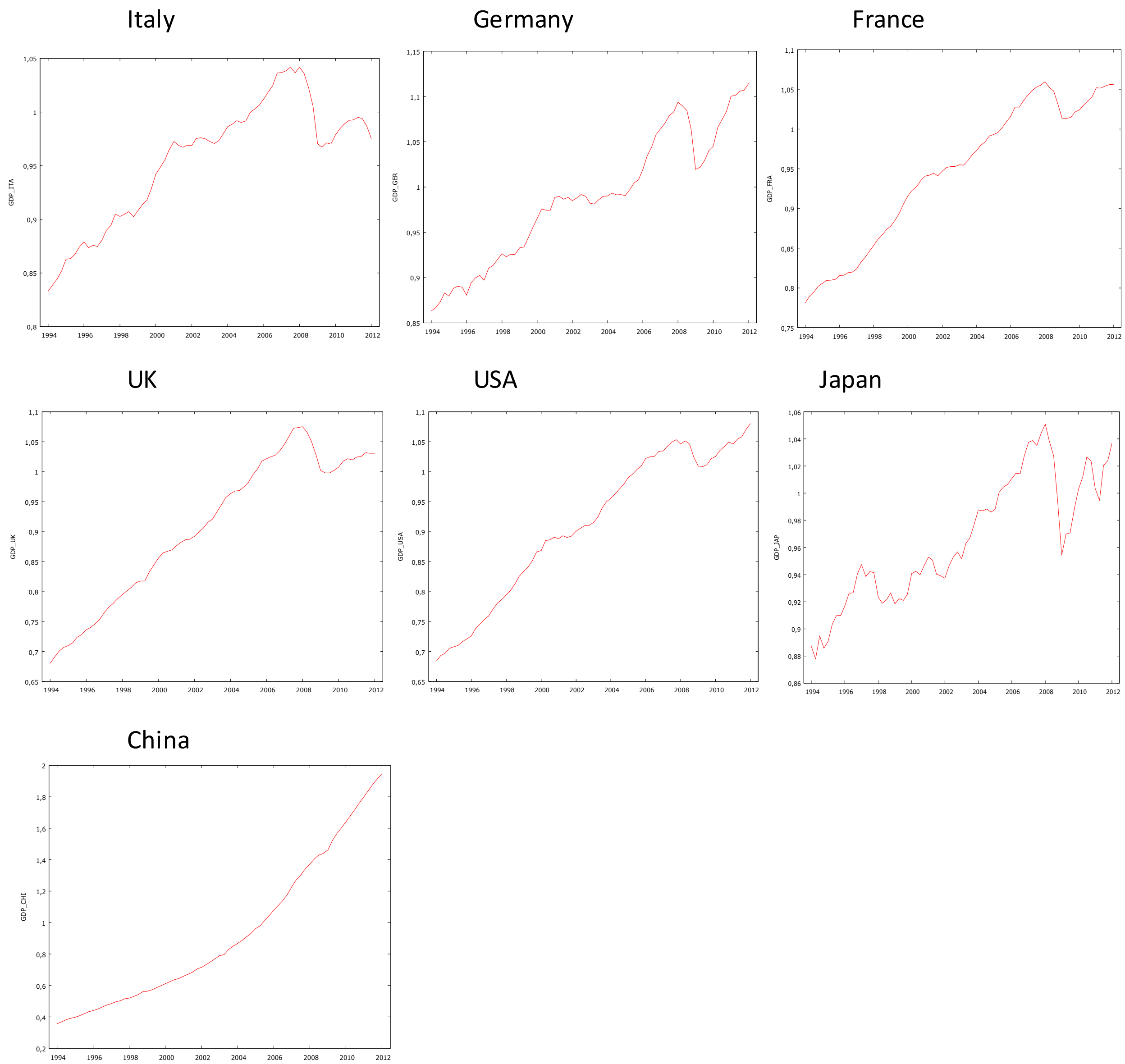


Fig.5.b. GDP for Italy, Germany, France, UK, USA, Japan and China; 1990:Q1 - 2012:Q1. Source: Elaboration on Datastream databases. Reference year 2005=100.

5. G) PMG estimation of import elasticities for Italy, France, Germany, UK, Japan, USA and China, 1990-2012:

Table 5 g. Estimation for import - Pooled Mean Group Estimator

	Coef.	Std. Err.	z	P>|z|	[95% Conf. Interval]	
LR						
log(exchange_rate)	**0,2612**	0,1354	1,93	0,054	-0,0043	0,5266
Italy - SR						
ec	-0,0185	0,0099	-1,87	0,061	-0,0379	0,0009
Δlog(exchange_rate)	0,3694	0,1736	2,13	0,033	0,0290	0,7097
Δlog(gdp)	0,1598	0,0547	2,92	0,003	0,0526	0,2669
intercept	-0,8561	0,3168	-2,7	0,007	-1,4771	-0,2352
Japan - SR						
ec	-0,2708	0,0431	-6,28	0	-0,3553	-0,1862
Δlog(exchange_rate)	*-0,0169*	0,0611	-0,28	0,783	-0,1367	0,1030
Δlog(gdp)	1,0474	0,1553	6,75	0	0,7431	1,3517
intercept	-12,8305	1,9398	-6,61	0	-16,6323	-9,0286
France -SR						
ec	*-0,0077*	0,0065	-1,18	0,237	-0,0204	0,0051
Δlog(exchange_rate)	-0,3905	0,1621	-2,41	0,016	-0,7083	-0,0727
Δlog(gdp)	0,0427	0,0181	2,36	0,018	0,0072	0,0782
intercept	-0,2210	0,1078	-2,05	0,04	-0,4323	-0,0097
UK - SR						
ec	-0,0172	0,0080	-2,16	0,03	-0,0328	-0,0016
Δlog(exchange_rate)	0,1867	0,0896	2,08	0,037	0,0112	0,3622
Δlog(gdp)	0,0441	0,0152	2,9	0,004	0,0143	0,0740
intercept	-0,1890	0,0888	-2,13	0,033	-0,3631	-0,0150
China - SR						
ec	*-0,0113*	0,0310	-0,36	0,716	-0,07212	0,049552
Δlog(exchange_rate)	*0,1109*	0,1278	0,87	0,386	-0,13963	0,361332
Δlog(gdp)	*-0,0012*	0,0454	-0,03	0,978	-0,09015	0,087689
intercept	*0,0721*	0,2897	0,25	0,803	-0,4957	0,63998
Germany - SR						
ec	-0,1305	0,0387	-3,37	0,001	-0,20632	-0,05467

Δlog(exchange_rate)	*-0,1285*	0,1624	-0,79	0,429	-0,44678	0,189871
Δlog(gdp)	0,5525	0,1606	3,44	0,001	0,237747	0,86719
intercept	-3,0478	0,8905	-3,42	0,001	-4,79324	-1,30245
USA - SR						
ec	-0,1144	0,0512	-2,24	0,025	-0,21466	-0,01414
Δlog(exchange_rate)	*0,0130*	0,1003	0,13	0,897	-0,18365	0,209562
Δlog(gdp)	0,2269	0,1182	1,92	0,055	-0,00483	0,458634
intercept	-1,7748	0,9665	-1,84	0,066	-3,66923	0,119538

Obs = 572; Numbero of Groups = 7; Obs per Group: max = 71; min = 84
Log Likelihood = 1272.91
Source: elaboration on Datastream databases

Verbale della riunione del: 10-10-2013

VERBALE DEL COLLEGIO DEI DOCENTI
SCUOLA DI DOTTORATO IN
SCIENZE ECONOMICHE E AZIENDALI

Il Collegio dei Docenti si riunisce, nella sede di Arcavacata - Aula Seminari - Cubo 0/C, alle ore 09:00 di giovedì 10/10/2013, per discutere in merito ai seguenti punti all'ordine del giorno:

1. Ammissione all'esame finale dei dottorandi del XXVI Ciclo del Dottorato in Scienze Economiche e Aziendali;
2. Ammissione al III° anno di corso dei dottorandi del XXVII Ciclo del Dottorato in Scienze Economiche e Aziendali;
3. Ammissione al II° anno di corso dei dottorandi del XXVIII Ciclo del Dottorato in Scienze Economiche e Aziendali;
4. Approvazione verbale del Collegio dei Docenti del 31-07-2013;
5. Varie ed eventuali.

Il Collegio, nella seduta odierna, è così composto:

N°	Membri del Collegio		P	AG	AI
1	AGOSTINO	Mariarosaria	X		
2	AIELLO	Francesco	X		
3	ALGIERI	Bernardina		X	
4	AQUINO	Antonio			X
5	BRUNI	Sergio			X
6	CARDAMONE	Paola	X		
7	CARIOLA	Alfio	X		
8	CARNEVALE	Concetta			X
9	COSTABILE	Massimo	X		
10	COSTANZO	G. Damiana			X
11	CRISTIANO	Elena			X
12	DE GIOVANNI	DOMENICO	X		
13	DE PAOLA	Maria	X		
14	D'ORIO	Giovanni		X	
15	DRAGO	Danilo			X
16	FABBRINI	Giuseppe	X		
17	FERRARI	Sonia		X	
18	FERRARO	Olga		X	

Il Segretario
(Dott. Giovanni Dodero)

Il Presidente
(Prof.ssa Patrizia Ordine)

Verbale della riunione del: **10-10-2013**

19	INFANTE	Davide	X		
20	LANZA	Andrea	X		
21	LA ROCCA	Maurizio	X		
22	LECCADITO	Arturo			X
23	LOMBARDO	Rosetta	X		
24	LUBERTO	Gaetano			X
25	MANNARINO	Lidia			X
26	MAROZZI	Marco	X		
27	MASSABO'	Ivar	X		
28	MAZZOTTA	Romilda	X		
29	MAZZUCA	Maria			X
30	MICELI	Gaetano	X		
31	MONTEFORTE	Daniele			X
32	NISTICO'	Rosanna	X		
33	ORDINE	Patrizia	X		
34	PASTORE	Patrizia			X
35	PILUSO	Fabio	X		
36	PIRRA	Marco			
37	PUNTILLO	Pina	X		
38	PUPO	Valeria		X	
39	RAIMONDO	Maria A.	X		
40	RICCIARDI	Antonio			X
41	RICOTTA	Fernanda	X		
42	RIJA	Maurizio		X	
43	ROSE	Giuseppe	X		
44	RUBINO	Franco E.			X
45	RUSSO	Emilio		X	
46	SCOPPA	Vincenzo	X		
47	SICOLI	Graziella	X		
48	SILIPO	Damiano B.			X
49	SILVESTRI	Antonella	X		
50	SMIRNOVA	Janna	X		
51	SUCCURRO	Marianna	X		
52	TRIVIERI	Francesco		X	
53	VELTRI	Stefania	X		

Il Segretario
(Dott. Giovanni Dodero)

Il Presidente
(Prof.ssa Patrizia Ordine)

Verbale della riunione del: **10-10-2013**

Presiede la seduta il Direttore del Collegio **Prof.ssa Patrizia Ordine**. Assume le funzioni di Segretario verbalizzante il **Dott. Giovanni Dodero**.

1. AMMISSIONE ALL'ESAME FINALE DEI DOTTORANDI DEL XXVI CICLO DEL DOTTORATO IN SCIENZE ECONOMICHE E AZIENDALI

Il Presidente invita i dottorandi Nadia Cosentino, Francesca Gioia, Fabiola Montalto e Alessia Via a partecipare alla riunione e ad esporre al Collegio dei Docenti il loro lavoro di Tesi, chiedendo agli stessi di focalizzare la loro esposizione sui principali risultati ottenuti e/o sugli aspetti innovativi emersi nel suddetto lavoro.

Segue la discussione delle tesi che avviene a porte aperte in ordine alfabetico.

La discussione delle Tesi termina alle ore 11:00 ed il Collegio si riunisce a porte chiuse per esprimere un giudizio sui lavori presentati, iniziando dalla dottoranda Alessia Via.

OMISSIS

Candidata: Dott.ssa Alessia Via

Il Presidente comunica ai Membri del Collegio le valutazioni espresse sul lavoro di Tesi della dottoranda Alessia Via dal Supervisore e dai Membri della Commissione di Valutazione Finale, come di seguito riportato:

Valutazione Finale della Tesi di Dottorato

Supervisore: Prof. Antonio Aquino
Il Prof. Antonio Aquino esprime una valutazione finale positiva sul lavoro di Tesi.

Commissione:
La Prof.ssa Mariarosaria Agostino esprime una valutazione finale positiva sul lavoro di Tesi.
Il Prof. Davide Infante, nonostante alcune riserve espresse su alcuni punti del lavoro di Tesi, ritiene che la candidata possa essere ammessa all'esame finale.
La Prof.ssa Rosanna Nisticò esprime una valutazione finale positiva sul lavoro di Tesi.

Dopo ampio ed approfondito dibattito, il Collegio, unanime, ammette la candidata Dott.ssa Alessia Via all'esame finale.

OMISSIS

Non essendovi null'altro da deliberare, il Presidente scioglie la seduta alle ore 14:45.

Il Segretario
(Dott. Giovanni Dodero)
Giovanni Dodero

Il Presidente
(Prof.ssa Patrizia Ordine)