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ESSAYS ON THE ESTIMATION OF THE EURO EFFECT ON TRADE

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To my parents

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Abstract

The European integration process has raised a wide debate about its cost and benefits since its very inception and the debate on its advantages and disadvantages has growing parallel to the Union. In this dissertation we aim to contribute to this literature in several manners. First, an empirical strategy to compare estimation methods is suggested and a thorough revision of the main factors affecting trade is performed, including the effect of the exchange rate (ER) level and volatility. Second, we improve the specification and estimation of the gravity equation, allowing for the presence of cross section dependencies, nonstationarities and structural breaks in the data as well as deterministic and stochastic trends. Finally, we investigate the impact of the euro both at the aggregate level and on each one of its members. We repeat this analysis for European Monetary Union (EMU) trade with third countries to explore the existence of potential diversion effects.

This thesis is structured into four chapters. In chapter 2, we focus on the study of the gravity equation, which is the main empirical tool employed in the literature to predict trade flows across countries. However, several problems related with its empirical application still remain unsolved. The unobserved heterogeneity, the presence of heteroskedasticity in trade data or the existence of zero flows, which make the estimation of the logarithm unfeasible, are some of them. This chapter provides a survey of the most recent literature concerning the specification and estimation methods for this equation. Using a dataset covering 80% of world trade, the most widely extended estimators are compared, showing that the Heckman sample selection model performs better overall for the specification of gravity equation selected. Furthermore, it is shown

that methods that do not properly treat the presence of zero flows on data exhibit noticeably worse performance than the rest in terms of efficiency. On the other hand, nonlinear estimators show more accurate results and are robust to the presence of heteroskedasticity.

In chapter 3 we address the impact that ER variables have on trade; an aspect frequently ignored in the estimation of the euro effect on trade. We estimate a gravity equation including the level and volatility of real exchange rate (RER) in order to capture the additional effect the euro could have had apart from the one coming from the elimination of the volatility. We find that the elimination of the volatility boosted export per se, especially before 1999. Then, the possibility to peg to the euro could boost trade with third countries and between those third countries. Our results show that the common currency has had a positive impact on intra-EMU exports. Though, it has reduced Eurozone's imports from third countries while it had not a significant impact on Eurozone's exports to other countries. Central and Eastern European (CEE) countries represent an exception since both their exports to EMU and imports from EMU have been boosted by euro. The analysis for individual EMU members reveals the existence of a good deal of variation in the effect of the euro across member countries. Concerning the impact of the euro over time, it significantly boosted intra-EMU trade starting in 1999, with this effect reaching its maximum in the 2003-2005 period.

In chapter 4, a further step is reached. We focus on the long-run estimation of the euro effect. We reduce the previous dataset to 26 OECD countries and we estimate the equation using two sets of variables: first, one defined as it is standard in the gravity equation literature, and a second one built according to the criticisms stated by Baldwin and Taglioni (2006) (BT henceforth). From a methodological point of view, we apply panel tests that account for the presence of cross-section dependence as well as discontinuities in the non-stationary panel data. We test for cointegration between the variables and efficiently estimate the long-run relationships using the new Continuously Updated Bias Corrected (CUP-BC) and Continuously Updated Fully Modified (CUP-FM) estimators proposed in Bai et al. (2009). These estimators assume that the cross sections in the model share common sources of non-stationary variation in the form of global stochastic trends. We argue that, after controlling for cross-section dependence, deterministic trends and breaks in trade integration, the euro generate lower trade effects than predicted in previous studies.

Finally, in chapter 5 we present evidence of the long-run effect of the euro focusing on trade of the twelve initial EMU countries from a double perspective. First, we pool all the bilateral combinations of trade flows among EMU countries in a panel cointegration gravity specification. Second, we estimate a gravity equation for each member vis-à-vis the other eleven partners. Whereas the joint gravity equation provides evidence on the aggregate effect of the euro on intra-European trade, by isolating the individual countries we assess which of the member countries have obtained a larger benefit from the euro. Moreover, this strategy permits to check the robustness of the aggregate results and to find possible asymmetries. Finally, we repeat both the aggregate and individual analysis for the bilateral trade of EMU members with third countries.

Resumen

El proceso de integración europeo ha constituido un objeto de estudio desde sus inicios, y el debate sobre sus ventajas e inconvenientes ha crecido de forma paralela a la Unión. En esta tesis contribuimos a la literatura anterior de varias maneras. En primer lugar, sugerimos una estrategia para comparar distintos métodos de estimación y llevamos a cabo una revisión exhaustiva de los principales factores que afectan al comercio, incluyendo entre ellos el efecto del tipo de cambio y la volatilidad. En segundo lugar, mejoramos la especificación y la estimación de la ecuación de gravedad, permitiendo la presencia de dependencia transversal y de no estacionariedad, así como de cambios estructurales en las series, e incluimos además tendencias deterministas y estocásticas. Finalmente, investigamos el impacto del euro tanto a nivel agregado como para cada uno de sus miembros, y repetimos este análisis para el comercio de la Unión Económica y Monetaria (UEM) con terceros países para descubrir la existencia de posibles efectos de desviación de comercio.

Esta tesis está estructurada en cuatro capítulos. El Capítulo 2 se centra en el estudio de la ecuación de gravedad, dado que es la herramienta empírica más utilizada en la literatura para predecir flujos de comercio entre países. Sin embargo, varios problemas relacionados con su aplicación permanecen aún sin resolver. La heterogeneidad no observada, la presencia de heteroscedasticidad en los datos o la existencia de ceros en la muestra –que imposibilita linealizar la ecuación usando logaritmos- son algunos de ellos. Por tanto, en este capítulo llevamos a cabo una revisión de la literatura más reciente relacionada con la especificación y los métodos de estimación para esta ecuación. Utilizando una muestra que cubre el 80% del comercio mundial, comparamos los estimadores más conocidos, concluyendo que el método de selección de Heckman es

el que presenta mejores resultados para la especificación seleccionada. Mostramos también que los métodos que no tratan correctamente la presencia de ceros en la muestra presentan peor comportamiento que el resto en términos de eficiencia. Por otra parte, los estimadores no lineales proporcionan resultados más precisos y son robustos a la presencia de heteroscedasticidad.

En el Capítulo 3 pasamos a analizar el impacto que tienen tanto el nivel del tipo de cambio como la volatilidad del mismo sobre el comercio; dos aspectos que generalmente no han sido tenidos en cuenta en la estimación del efecto del euro. Para ello, utilizamos una ecuación de gravedad incluyendo ambas variables de forma que el efecto del euro es aislado del impacto de la eliminación de la volatilidad, la desaparición de los costes de transacción y otros cambios asociados a la nueva moneda. Los resultados muestran que la reducción de la volatilidad del tipo de cambio ha impulsado por sí misma las exportaciones, especialmente antes de 1999. Por ello, la posibilidad de anclar otras monedas al euro podría incrementar tanto el comercio europeo con terceros países como el comercio entre esos terceros países. Los resultados muestran que la creación de una moneda común ha tenido un impacto positivo sobre las exportaciones de la UEM hacia el resto de países miembros. Por otra parte, aunque ha reducido las importaciones de la Eurozona provenientes de terceros países, no ha tenido un impacto significativo en las exportaciones hacia esos países. Los países de Europa Central y del Este constituyen una excepción dado que tanto sus importaciones como sus exportaciones se han visto incrementadas gracias al euro. El análisis individual para cada uno de los países revela la existencia de una notable variación del impacto del euro en los países miembros. En cuanto al efecto en el tiempo, el euro ha incrementado notablemente el comercio intraeuropeo desde 1999, alcanzando un máximo en el periodo de 2003-2005.

En el Capítulo 4 damos un paso más, centrándonos en la estimación a largo plazo del efecto del euro. Reducimos la muestra anterior a 26 países pertenecientes a la OCDE y estimamos la ecuación utilizando dos tipos de variables: en primer lugar, un grupo definidas como tradicionalmente se ha hecho en la literatura relacionada con la ecuación de gravedad; y en segundo lugar otro grupo definido teniendo en cuenta las críticas hechas en Baldwin y Taglioni (2006). Desde un punto de vista metodológico, aplicamos diversos tests para datos de panel robustos a la presencia de dependencia transversal, así como a la existencia de discontinuidades en datos no estacionarios. Comprobamos la existencia de cointegración entre las variables y estimamos de forma eficiente las relaciones de largo plazo utilizando los nuevos estimadores CUP-BC y CUP-FM propuestos por Bai et al. (2009). Estos estimadores asumen que las unidades en el panel comparten fuentes comunes de variación no estacionaria en forma de tendencias estocásticas globales. Afirmamos que, después de controlar la dependencia transversal y la existencia de tendencias deterministas y estocásticas, así como de cambios estructurales en la integración comercial, el euro muestra un efecto más moderado que en estudios anteriores.

Por último, en el Capítulo 5 examinamos el efecto a largo plazo del euro centrándonos en el comercio de los doce países originarios de la UEM desde una doble perspectiva. En primer lugar, analizamos todas las posibles combinaciones bilaterales de flujos de comercio entre los miembros de la UEM usando para ello una ecuación de gravedad estimada con técnicas de cointegración. En segundo lugar, realizamos el mismo análisis para el comercio de cada uno de los miembros de la UEM con el resto de miembros. El primer ejercicio proporciona evidencia del efecto agregado del euro en el comercio intraeuropeo; mientras que el segundo, al aislar a los países de forma individual, permite comprobar la robustez de los resultados y encontrar posibles asimetrías del efecto del euro sobre sus miembros. Finalmente, repetimos tanto el análisis agregado como el individual para el comercio de los países UEM con terceros países.

Chapter 1

General Introduction

Since the signing of the Treaty of Rome in 1956, the European Union (EU) has become one of the most integrated areas in the world context. Robert Schuman's words pronounced in 1950 "*Europe will not be made all at once, or according to a single plan. It will be built through concrete achievements which first create a de facto solidarity*" now becomes a reality. Throughout the process, several milestones can be highlighted. In 1968, a customs union was established, followed in 1979 by the European Monetary System. Both facts consolidated the process even despite the emerging Euroscepticism and led to the signing of the Single European Act in 1986. The Maastricht Treaty in 1992 confirmed the foundation of the EU and prepared the launch of the euro as a common currency in 1999. The process is still going on, and several waves of enlargement have taken place in the last few years. At the present time, 27 countries belong to the EU, of which 17 share the euro as common currency.

Nowadays, more than ten years after the advent of the euro, the debate has undergone a strong renewal due to the common currency crisis prevailing in Europe. This crisis has renewed all the issues raised 10 years ago by Eurosceptics about the difficulties to share a common currency without significant fiscal and political coordination. It is not trivial that joining the EMU involves in many cases the loss of national sovereignty. In fact, the scepticism is less concerned with commercial integration and more related to political issues. The relevance of the issue makes it a challenge that has generated a growing academic interest. The debate has been activated by the surprising results in Rose' (2000) article, which predicted a tripling of trade for those countries belonging to a Currency Union (CU). In parallel, the econometric approach to this issue has also changed over time as a result of a feedback process between theory and empirics. The development of modern econometric software has suggested the need for a revision of the previous results concerning the euro effect. To the extent that some European countries are still thinking about joining the EMU, it is important to have a robust evaluation of the benefits the euro had on trade and could still have. In light of the euro crisis, it could also be interpreted as a demonstration of the cost of non-participation.

In this dissertation, we aim to contribute to this debate providing evidence of the euro effect on trade and the appropriate methods to estimate it. To this end, the thesis is divided into four essays, each one with its own structure and framework, which are integrated into a coherent work with a logical progression from one to another. In chapter 2 a thorough review of different methods to estimate the gravity equation and relevant literature are presented. The performance of several linear and nonlinear estimators is compared using a three-dimensional dataset, analyzing their most relevant properties. Chapter 3 focuses on the effects of the level and volatility of ERs on trade, completely isolating the euro effect from other factors. In addition, the effect of EMU on Eurozone's trade with third countries is estimated. In chapter 4 we deal with some econometric problems that affect the long-run estimation of the equation; namely the nonstationarity of the variables, the presence of cross-correlation among the series and the existence of discontinuities in the time dimension. Finally, in chapter 5 we apply the previous methodology focusing exclusively on the twelve initial EMU countries from a double perspective. First, we pool all the bilateral combinations of export flows among EMU countries and next we estimate a gravity equation for each member vis-à-vis the other eleven partners. We repeat both the aggregate and individual analysis for the bilateral exports of EMU members to third countries.

We use a dataset that includes bilateral exports flows from 80 countries during the period 1967-2009, representing 80% of world trade. In chapter 4 we reduce this sample to 26 OECD countries and in chapter 5 we focus exclusively on 12 EMU countries. Our specification of the gravity equation is based on Anderson and van Wincoop's (2003) theoretical model (AvW 2003 henceforth). In chapter 2 this specification is completed with the inclusion of ER variables. Finally, in chapters 4 and 5 we also introduce country pair specific trends, structural breaks in the constant, the country pair specific trends and the cointegrating vector, and we allow for a common factor structure in the error term.

This thesis presents several novelties with respect to previous literature. First, we use an augmented gravity equation that explicitly takes into account the level and volatility of bilateral RER, the presence of regional trade agreements and the EMU. To the best of our knowledge, none of the previous studies have included such a complete specification in terms of ER variables and trade agreements, thus ignoring important aspects in the estimation of the euro effect. Additionally, most of the recent literature makes the choice to focus on a reduced sample of developed countries. By using a large dataset, we are able to study the impact of the euro on Eurozone's trade with third countries and between these third countries as well as the opportunity that the euro offers to other EU countries in terms of trade. Finally, this is the first time that panel cointegration techniques allowing for structural breaks and cross section dependence are applied to the estimation of the euro effect using the gravity equation.

The first logical step in the analysis is the selection of the appropriated tool to carry out the estimation. The gravity equation is widely recognized in trade literature to be one of the most successful empirical tools to predict trade flows. Hence, in the first chapter, we review the literature concerning this equation and perform a thorough revision of the main estimation problems and their corresponding solution in the short-run estimation. We summarize these problems into four groups. First, as claimed by AvW, the exclusion of the multilateral trade resistance (MTR) terms, which capture all the barriers that each country faces with its trading partners, leads to biased estimates. Second, the log-linearisation of the gravity equation changes the property of the error term, thus leading to inefficient estimation in the presence of heteroskedasticity, as noted by Santos-Silva and Tenreyro (2006). Although heteroskedasticity does not affect the parameter estimates -the coefficients should still be unbiased- it biases the variance of the estimated parameters and, consequently, the t-values cannot be trusted. Third, some characteristics in data such as regulation, political factors, technology, efficiency, etc. may differ from one country -or country pair- to another without being captured by the regressors. This unobserved heterogeneity leads to biased estimates if it is not properly controlled for. Finally, and since the logarithm of zero is unfeasible, the use of disaggregated datasets, in which over 50% of values are zero, creates a problem of selection bias and information loss. All these potential problems require a detailed inspection of the dataset before choosing the appropriate estimator. We review in this chapter the solutions that have previously been given to these questions by comparing them for the same dataset and we suggest an empirical strategy to be performed before selecting the appropriate estimator, including a battery of pre and post-estimation tests and several goodness-of-fit criteria. In addition, we provide evidence of the fact that methods that do not properly treat the presence of zero flows on data exhibit noticeably worse performance than others in terms of efficiency and that nonlinear estimators show more accuracy in the presence of heteroskedasticity. For the specific dataset employed, the Heckman sample selection model is the preferred estimation method. However, it should be noted that this result depends on the data characteristics and an appropriate analysis should be performed before estimating the equation.

Once the empirical tool is properly selected and the properties of the estimation methods are studied, we address in chapter 3 the impact that the ER level, ERV and euro have on Eurozone’s trade. Previous literature includes the works of Baldwin et al. (2005), Barr et al. (2003), Brouwer et al. (2008) or Gil-Pareja et al. (2008). Rose (2000), Rose and Engel (2002), Clark et al. (2004) and Tenreyro (2007) also take into account the ERV, though they do not explicitly focus on the euro effect. However, there are still several challenges when dealing with this issue. One of them is to separate the impact of the elimination of ERV on trade flows from other effects of the euro like the elimination of transaction costs and other permanent changes associated with the new currency. This implies computing the ERV for a sufficiently long period and large sample to capture differences in this variable among partners and time. Besides that, it is worth noticing that although ERV has received more attention, most of the related articles do not include the ER level in the specification. We show that these variables have a significant effect on trade and should be included in the estimation. Therefore, in this chapter we estimate a gravity equation including the level and volatility of the RER in order to capture the additional effect the euro could have had apart from the one coming from the elimination of the volatility. We find that the elimination of ERV boosted export per se, especially before 1999. Then, the possibility to peg to the euro could boost trade with third countries and between these third countries. Our results show that the common currency has had a positive impact on intra-EMU exports, though it has reduced Eurozone’s imports from third countries while it had not a significant impact on

Eurozone's exports to other countries. CEE countries represent an exception since both their exports to EMU and imports from EMU have been boosted by the euro. The analysis for individual EMU members reveals the existence of a good deal of variation in the euro effect across member countries since it has only boosted the exports of 4 of 12 EMU countries. Concerning its impact over time, the euro significantly boosted intra-EMU trade starting in 1999, with this effect reaching its maximum in the 2003-2005 period.

In chapter 4 a further step is reached. While initially the literature overlooked some crucial econometric issues regarding non-stationary series in panel estimation, more recent works have taken into account these aspects in the context of the gravity model. The novelty of this problem and the possibilities offered by the development of the econometric software raise a stimulating challenge in the econometric area that has promoted a renewal of the literature. Faruquee (2004), Bun and Klaasen (2007), Berger and Nitsch (2008), Fidrmuc (2009) or Gengenbach (2009) are some of the authors that have applied cointegration techniques to the estimation of the gravity equation and the euro effect. However, there are three additional aspects in this context that still remain unexplored. The first one relates to the fact that some widely used tests and estimators assume the absence of correlation across units in the panel. This assumption is not realistic, especially when the countries are neighbours or are involved in integration processes. A sizeable literature has been developing along these lines, but none of these works explicitly deals with the issue of cross-section dependence with the exception of Gengenbach (2009). The second aspect is noted by Bun and Klaasen (2007). They show that the longer the period considered, the higher the euro effect estimate. They attribute this fact to some misspecification of the time-series characteristics of the variables involved, namely the trends in trade flows over time. To correct for this bias they add a time-trend to their specification and allow it to have heterogeneous coefficients across country pairs. Finally, the third issue concerns the fact that, until the moment, the time series in the panel dataset have been assumed to be stable; something that, depending on the sample, may be quite unrealistic. Banerjee and Carrion-i-Silvestre (2010) point out that both the cointegrating vector and the deterministic components might change during the period analyzed, and if these structural breaks in the parameters of the model are not taken into account, inference concerning the presence of cointegration can be affected by misspecification errors.

Thus, in this chapter we contribute to the previous literature in several respects. First, we account for cross-section dependence among countries in the panel tests. We implement the panel unit root and stationary tests proposed by Pesaran (2004, 2007) and Bai and Ng (2004) and we test for cointegration between the variables using panel cointegration tests, with a special emphasis in the one proposed by Banerjee and Carrion-i-Silvestre (2010). Second, we allow for the existence of a break in the deterministic components (constant, trends and cointegrating vector) of the model as well as in the cointegration relationship, a major point to assess the effect of institutional changes in the relationship. Furthermore, since the trends included in the specification are country pair specific, the breaks in the trends are also allowed to have different coefficients for each country pair, therefore allowing for a higher degree of heterogeneity in the estimation. To the best of our knowledge, this is the first time that structural changes have been considered in the euro effect literature based on gravity equations. Finally, the estimation of the long-run relationship uses a methodology that not only efficiently estimates the coefficients but also is based on the common factors decomposition that assures a homogeneous econometric approach. We choose, for this purpose, the continuously updated estimators proposed by Bai et al. (2009); CUP-FM and CUP-BC. The presence of unit roots, dependencies and breaks is confirmed, thus revealing the need for using appropriate estimators. The results obtained confirm a smaller euro effect than other research papers, where cross-section dependence and the non-stationary nature of the variables are not accounted for.

Additionally, this essay also contributes to the existing literature addressing BT's critiques, known as 'gold, silver and bronze medal errors', that concern the specification of gravity models and the definition of the variables. To deal with the 'gold' error BT suggest the inclusion of time varying fixed effects in the specification. However, if doing so, we would not be able to explore cointegration between Gross Domestic Product (GDP) and exports, since the time varying fixed effects would absorb GDP. Instead of that, we include the above-mentioned country pair specific time trend, which captures all the unobserved heterogeneity through time, as well as country specific fixed effects. Moreover, the application of cointegration techniques implies the proper treatment of the time dimension, since it takes into account the long-run relationships among variables. We include exports as dependent variable and define all the variables in nominal terms to avoid 'silver' and 'bronze' errors and we compare the results using both sets of variables. As expected, we find that the estimation with incorrectly defined variables provides biased coefficients.

A natural extension of the previous essay is the study of the euro effect on each individual country applying cointegration techniques, since there is little evidence on the asymmetric effect of the euro on its members, as well as in trade with third countries. Faruquee (2004) provides a comparison of this effect by interacting country dummies with the EMU variable. Dwane et al. (2011) also perform this analysis, but focusing exclusively on Irish trade. In both cases the possibility of breaks is ignored and cross section dependencies are not modelled. On the other hand, the estimation of the euro effect on trade with third countries has received much less attention in the literature. Kelejian et al. (2011) include two dummy variables in the estimation to distinguish between imports and exports, finding positive results. Studies of Micco et al. (2003), Baldwin et al. (2005) and Gil-Pareja et al. (2008) also obtain results in this line.

Besides that, the test proposed by Banerjee and Carrion-i-Silvestre (2010) assumes the existence of one structural break common to all the countries included in the sample. Both the idea of having more information on each specific EMU country and the need for a more homogeneous sample to make the assumption in Banerjee and Carrion-i-Silvestre (2010) suitable lead us in chapter 5 to perform an empirical exercise using a reduced dataset that includes as exporters only EMU members. Having determined in chapter 4 the correct definition of variables, the correct specification and the appropriate estimation method, we perform in this chapter two analyses; one for the exports of each EMU country to the rest of members; the other for the exports of EMU members to 16 non EMU countries. We investigate the aggregate euro effect on internal and external European trade as well as the specific effect on each one of its members in a panel cointegration framework, allowing for structural breaks in the specification. Again, we employ Bai et al. (2009) CUP estimator, which is consistent in the presence of cross section dependencies, and we use a more homogeneous sample -more appropriate when the date of the break is unique. Our findings show that the euro has mainly affected intra-EMU trade, whereas the effect for third countries is not significant. Consistently with chapter 3, the inspection at the country level reveals that France, Italy and Belgium and Luxembourg are the countries that more benefited from the introduction of the euro when referring to intra-EMU trade. The effects for trade with third countries are in general more moderate; and with the exception of Greece, there is no evidence of diversion effects. Nevertheless, it is worth noticing that the rest of countries included in the estimation are OECD members and the commercial relationships between these countries and EMU members have been relatively stable since the OECD creation in 1960. Hence, it is not surprising that the introduction of the euro has not negatively affected these links. By contrast, the results in chapter 3 reveal that when the sample is extended to developing countries there is evidence of a reduction in EMU imports from third countries due to the introduction of the euro. Finally, the analysis of the structural breaks also sheds some light on the integration process. In the aggregate case the break is found in 1987 for intra-EMU case, a date that can be attributed to the effects of the Single European Act, which came into force in that year. The main commitment agreed on in this Treaty was the adoption of measures guided to the progressive establishment of a common market over a period that would conclude in 1992. For trade with third countries, the break takes place in 1989. We relate this date with the signing of the Plaza and Louvre Agreements, which were important milestones in the international economic context. Concerning the country specific results, different dates are found. For intra-EMU trade, the dates are close to the entering of each country in the EU, whereas for EMU trade with third countries the breaks are more related with the oil crisis in the 1973-1979 period.

Chapter 2

Comparing alternative methods to estimate gravity models of bilateral trade

Abstract. The gravity equation has been traditionally used to predict trade flows across countries. However, several problems related with its empirical application still remain unsolved. The unobserved heterogeneity, the presence of heteroskedasticity in trade data or the existence of zero flows, which make the estimation of the logarithm unfeasible, are some of them. This chapter provides a survey of the most recent literature concerning the specification and estimation methods of this equation. For a dataset covering 80% of world trade, the most widely extended estimators are compared, showing that the Heckman sample selection model performs better overall for the specification of gravity equation selected.

2.1. Introduction

The gravity model of trade, which was originally inspired by Newton's gravity equation, is based on the idea that trade volumes between two countries depend on their sizes in relation to the distance between them. In the last fifty years, this model has been widely used to predict trade flows.

The gravity equation appears to be highly effective at this point as proven at a very early date by the works of Linnemann (1966) and Leamer and Stern (1971). However, several controversies have arisen regarding the model. The theoretical framework was put into doubt and afterwards justified by Bergstrand (1989) for the factorial model, Deardorff (1998) for the Hecksher-Ohlin model, Anderson (1979) for goods differentiated according to their origin, and Helpman et al. (2008) in the context of firm heterogeneity. After some additional discussions concerning its specification in the nineties, the debate has now turned to the performance of different estimation techniques. New estimation problems concerning the validity of the log linearisation process of the gravity equation in the presence of heteroskedasticity and the loss of information due to the existence of zero trade flows have been recently explored.

Traditionally, the multiplicative gravity model has been linearised and estimated using Ordinary Least Squares (OLS) assuming that the variance of the error is constant across observations (homoskedasticity), or using panel techniques assuming that the error is constant across countries or country pairs. However, as pointed out by Santos-Silva and Tenreyro (2006), in the presence of heteroskedasticity, OLS estimation may not be consistent and nonlinear estimators should be used. Another challenge described in the literature concerns the zero values. Helpman et al. (2008) propose a theoretical foundation based on a model with heterogeneity of firms à la Melitz (2003) and an adapted Heckman procedure to predict trade taking into account these features. Recently, the works of Burger et al. (2009), Martin and Pham (2008), Martínez-Zarzoso (2011), Siliverstovs and Schumacher (2009) and Westerlund and Wilhelmsson (2009) have obtained divergent results when comparing alternative estimation methods.

This chapter reviews most estimation methods and problems and provides a survey of the literature related to this topic. The performance of several linear and nonlinear estimators is compared using a three-dimensional (*i, j, t*) dataset, analysing the most relevant properties of each one. To this end, a gravity equation based on AvW's theoretical model is used. Using this equation, the fit of different estimation procedures applied to a large dataset of bilateral exports for 80 countries (80% of world trade) over the 1980-2008 period is discussed. The fit of each method is compared through different measures, revealing the main advantages and disadvantages of each one. It is shown that methods that do not properly treat the presence of zero flows on data exhibit noticeably worse performance than others. On the other hand, nonlinear estimators show more accurate results and are robust to the presence of heteroskedasticity in data. Overall, the Heckman sample selection model is revealed to be the estimator with the most desirable properties, confirming the existence of sample selection bias and the need to take into account the first step (probability of exporting) to avoid the inconsistent estimation of gravity parameters.

The rest of the chapter is organised as follows. The next section briefly reviews the different theoretical foundations of the gravity equation to justify the election of the empirical specification of the gravity equation chosen. Section 2.3 compares in detail the different estimation methods available in the gravity literature. In Section 2.4, data are presented and the results of different estimations methods are discussed and compared according to different criteria. Conclusions are drawn in Section 2.5.

2.2. The gravity equation

The theoretical foundation of the gravity equation appeared seventeen years after its empirical specification. The first article providing a microfoundation of this equation was Anderson (1979) and was based on the Armington assumption of specialisation of each nation in the production of only one good. Bergstrand (1985) initially supported this hypothesis, completing the theoretical foundation with a more detailed explanation of the supply side of economies and the inclusion of prices in the equation.

A few years later, a new wave of developments came with what has been called “the new trade theory”. The main improvement was the replacement of the assumption of product differentiation by country of origin by the assumption of product differentiation among producing firms. In this line, Bergstrand (1990) provided a foundation based on Dixit and Stiglitz’s monopolistic competition assumption. In addition, he generalised the model by introducing prices and incorporating the Linder hypothesis. Helpman (1987) also derived a foundation relying on the assumption of increasing returns to scale where products were differentiated by firms, not only by country, and firms were monopolistically competitive. However, some years later Deardoff (1998) asserted that the gravity equation could be derived from standard trade theories, conciliating both the old and the new theories.

Later on, the “new new trade theory” insisted on the heterogeneity of firms regarding their exporting behaviour (Melitz 2003), thereby giving a theoretical foundation for the presence of zero trade flows in data. In this line, Helpman et al. (2008) generalised the empirical gravity equation by developing a two-stage estimation procedure that takes into account extensive and intensive margins of trade. They showed that the incorrect treatment of zero flows may lead to biased estimates and developed a complete framework to provide a rationale for the existence of these flows.

Regarding the specification, AvW propose an augmented version of the Anderson (1979) model based on the assumption of differentiation of goods according to place of origin. Their main contribution is the inclusion of multilateral resistance terms for the importer and the exporter that proxy for the existence of unobserved trade barriers. This model is interesting overall to the extent that the discussion of the multilateral resistance may matter for heteroskedasticity considerations. In this model, countries are representative agents that export and import goods. Goods are differentiated by place of origin and each country is specialised in the production of only one good. Preferences are identical, homothetic and approximated by a constant elasticity of substitution (CES) function.

The linear gravity equation estimated by AvW is as follows:

$$\begin{aligned} \ln X_{ij} = {} & k + \ln y_i + \ln y_j + (1 - \sigma)\rho \ln d_{ij} + (1 - \sigma)\ln b_{ij} + (1 - \sigma)\ln P_i \\ & + (1 - \sigma)\ln P_j + \varepsilon_{ij} \end{aligned} \tag{2.1}$$

where X_{ij} is the nominal value of exports from i to j; k is a positive constant, y_i and y_j are the nominal income of each country, generally proxied by its GDP, and d_{ij} is a measure of the bilateral distance between i and j, which are introduced to proxy for transport costs. b_{ij} is a dummy variable that takes value one if two countries share a border. Finally, the variables P_i and P_j are the multilateral resistance terms, defined as a function of each country’s full set of bilateral trade resistance terms. The variable of interest for AvW is b_{ij} since their objective is to estimate the trade effect of national borders. They apply their equation to regional data.

The multilateral price indices (P_i and P_j) are not observed and should be estimated. AvW use the observed variables in their model (distances, borders, and income shares) to obtain the multilateral trade resistance terms. Assuming symmetric trade costs, using 41 goods market-equilibrium conditions[1] and a trade cost function defined in terms of observables, they obtain the P_i and P_j terms. Although they argue that this method is more efficient than any other, it is highly data consuming and has not been frequently used by other authors.

An alternative solution is to include a remoteness variable to proxy for these multilateral trade resistance indexes:

$$Rm_i = \sum_j \frac{d_{ij}}{(y_j / y_{ROW})} \tag{2.2}$$

where the numerator would be the bilateral distance between two countries, and the denominator would be the share of each country's GDP in the rest of the world's GDP. Head and Mayer's (2000) remoteness variable describes the full range of potential suppliers to a given importer, taking into account their size, distance and relevant costs of crossing the border. Wei (1996), Wolf (1997), and Helliwell (1996) provide other examples of regressions including a remoteness variable. Alternatively, Feenstra (2002) proposes introducing importer and exporter fixed effects to account for the specific country multilateral resistance term. The coefficient of the dummies for the importer and the exporter should reflect the multilateral resistance for each country. Several

[1] Their sample contains the same 30 US states and 10 Canadian provinces that McCallum (1995) includes. There are 20 additional states, plus Columbia, which they aggregate into one. Hence, they finally have 41 equations.

studies using this approach are described in the Appendix (Table 2.A.1). Finally, Baier and Bergstrand (2009) suggest generating a linear approximation of the P_i and P_j terms by means of a first-order Taylor series expansion.

Concerning the proxy for supply and demand sizes, the common practice is to use importer's and exporter's GDP correspondingly. In some cases GDP per capita is also introduced as a proxy for capital-labour intensities.

Transaction costs are frequently proxied by geographical distance. However, it is commonly accepted that geographical distance may be a poor approximation[2]. Thus, this variable is often completed with other proxies for trade barriers specified as indicator variables. For instance, adjacency takes value one if trade partners share a common border, common language takes value one if both countries share a language, colonial links captures the effect of having had a common coloniser or having been colonised by another country in the past; religion takes value one when both countries have the same religion; access to water takes value one if a country has access to water, or Regional Trade Agreement (RTA) which assess the effect of RTAs on trade. All these factors affect international trade via transaction costs and complete the geographical distance variable in order to reflect the economic distance.

2.3. Summary of estimation methods

As mentioned above, interest in the last years has focused on estimation methods to accurately predict trade flows. In this section, a brief summary of some of the most important methods as well as a revision of related empirical literature (Table 2.1) are

[2] In addition, there is no single opinion about how distance should be measured. The most common measures are the great circle formula and the distance between the two principal cities. See Wei (1996), Wolf (1997), and Head and Mayer (2000) for further information.

presented.

2.3.1. Linear methods

Since the logarithm of zero is not defined, truncation and censoring methods have been proposed in the literature to treat the problem of the existence of zero flows in data. However, these procedures reduce efficiency due to the loss of information and may lead to biased estimates due to the omission of data. Furthermore, as Westerlund and Wilhelmsson (2009) point out, the elimination of trade flows when zeros are not randomly distributed leads to sample selection bias.

In addition, a panel framework permits recognising how the relevant variables evolve through time and identifying the specific time or country effects. Over the last years, researchers such as Egger (2000), Rose and van Wincoop (2001), Mátyás (1998), Egger and Pfaffermayr (2003, 2004), Glick and Rose (2002), Brun et al. (2002), and Melitz (2007) have turned towards panel data[3]. Two main techniques are employed to fit data depending on the a priori assumptions. The fixed effects estimator assumes the existence of an unobserved heterogeneous component constant over time that affects each individual (pair of countries) of the panel in a different way. By contrast, the random effects model imposes no correlation between the individual effects and the regressors, implicitly assuming that the unobserved heterogeneous component is strictly exogenous. Under the null hypothesis of zero correlation, the random effects model is more efficient. However, if the null is rejected, only the fixed effects model provides

[3] See Appendix A for further information.

consistent estimators[4].

2.3.2. Nonlinear methods

As Santos-Silva and Tenreyro (2006) points out, the log-linearisation of the gravity equation changes the property of the error term, thus leading to inefficient estimations in the presence of heteroskedasticity. If the data are homoskedastic, the variance and the expected value of the error term are constant but if they are not -as usually happens with trade data-, the expected value of the error term is a function of the regressors. The conditional distribution of the dependent variable is then altered and OLS estimation is inconsistent. Heteroskedasticity does not affect the parameter estimates; the coefficients should still be unbiased, but it biases the variance of the estimated parameters and, consequently, the t-values cannot be trusted. Hence, the recent literature concerning estimation techniques have opted to use nonlinear methods as well as two parts models for estimating the gravity equation.

Among nonlinear estimation methods, the most frequently used are Nonlinear Least Squares (NLS), Feasible Generalised Least Squares (FGLS), Heckman sample selection model and Gamma and Poisson Pseudo Maximum Likelihood (GPML and PPML). Santos-Silva and Tenreyro (2006) claim that NLS is inefficient since it gives more weight to observations with larger variance and is not robust to heteroskedasticity. Martínez-Zarzoso (2011) propose Feasible Generalised Least Squares (FGLS) as the most appropriate model if the exact form of heteroskedasticity in data is ignored since it weighs the observations according to the square root of their variances and is robust to any form of heteroskedasticity. Manning and Mullahy (2001) propose Gamma Pseudo

[4] The Hausman test provides a method for testing the adequacy of the random effect model. If the null is rejected, the random effects model is not consistent. However, it is important to note that this result does not imply that the fixed effect model is adequate.

Maximum Likelihood (GPML). In this case the conditional variance of the dependent variable is assumed to be proportional to its conditional mean. This estimator therefore assigns less weight to observations with a larger conditional mean. Martínez-Zarzoso et al. (2007) computes the performance of this estimator, finding it to be adequate in the presence of heteroskedasticity, although it shows less accuracy when zero trade values are present. Finally, Poisson Pseudo Maximum Likelihood (PPML) is similar to GPML, but assigns the same weight to all observations. Santos-Silva and Tenreyro (2006) point out that this is the most natural procedure without any further information on the pattern of heteroskedasticity.

In addition, two-step estimation methods have also been proposed to estimate the gravity equation. This is the case of Heckman sample selection model. In the first step, a Probit equation is estimated to define whether two countries trade or not and in a second step, the expected values of the trade flows, conditional on that country trading, are estimated using OLS. In order to identify the parameters on both equations, a selection variable is required. This exclusion variable should affect only the decision process; hence, it should be correlated with a country's propensity to export but not with its current levels of exports. Some examples in the literature are the common language and common religion variable (Helpman et al. 2008), governance indicators of regulatory quality (Shepotylo 2009), or the historical frequency of positive trade between two countries (Bouet et al. 2008). Alternatively, Linders and de Groot (2006) or Haq et al. (2010) include the same variables in both equations, imposing the normality of the error in both equations as an identification condition, which implies a zero covariance between them. The advantage of a sample selection model comes from the fact that the decision on whether to trade or not and the decision on how much to trade are not modelled as completely independent. The model allows for some positive correlation between both error terms to better reflect the real decision process. For further information on this topic see Egger et al. (2011).

Helpman et al. (2008) extends Heckman's estimation method to also take into account the bias associated with the heterogeneity of firms. The authors develop a complete theoretical framework from which they obtain an empirical specification of the gravity equation. Their model accounts for firm heterogeneity, trade asymmetries and fixed trade costs, suggesting that the decision to export (extensive margin) and the volume of exports (intensive margin) are not independent variables. The model allows both positive and zero trade flows between countries to be predicted and it also allows exports to vary according to the destination country. Helpman et al. (2008) describe a varying distribution of firms where each firm is bounded by a marginal exporter who breaks even by exporting to another country. The underlying idea is that if at least one firm in the country is productive enough to export, country-level exports in that case will be positive. Hence, zero exports are originated by countries where firms are not productive enough to export profitably. In this manner, information that would normally require firm-level data is extracted from country-level data.

They argue that controlling for both the extensive margin and the sample selection would completely eliminate the bias in the estimation. The results confirm their theoretical predictions, showing that the omission of a measure of firms' heterogeneity leads to substantial biases in the estimation. They prove the robustness of their results using religion instead of common language as exclusion variable. Most articles employing the Helpman et al. (2008) methodology apply it to a cross-section dataset. Application of the methodology in a panel framework still requires further research and goes beyond the scope of this chapter.

Every method has advantages and disadvantages and it cannot be asserted that any one of them absolutely outperforms the others. For that reason, it has become a frequent practice in the literature to include several estimation methods for the same database. In the next section, an empirical exercise comparing these methods is presented.

Table 2.1. Summary of estimation methods

Estimation method	**Advantages**	**Disadvantages**	**References**
Truncated OLS	- Simple	- Loss of information (zero flows) - Biased coefficients	Linders and de Groot (2006); Westerlund and Wilhelmsson (2009); Martin and Pham (2008)
OLS ($1+T_{ij}$)	- Simple - It deals with the zero trade flows problem	- Biased coefficients	Linneman (1966), Bergeijk and Oldersma (1990); Wang and Winters (1991); Baldwin and DiNino (2006)
Tobit (censored regression)	- Simple - It deals with the zero trade flows problem	- Same set of variables to determine the probability of censoring and the value of the dependent variable - Lack of theoretical foundation	Soloaga and Winters (2001); Anderson and Marcouiller (2002); Baldwin and DiNino (2006); Schiavo (2007); Martin and Pham (2008)
Panel fixed effects	- Simple - It controls for unobserved heterogeneity	- Loss of information (constant terms dropped) - Elimination of zero flows - Sample selection bias	Mátyás (1998); Egger (2000); Glick and Rose (2002); Egger and Pfaffermayr (2003); Micco et al. (2003); Andrews et al. (2006); Henderson and Millimet (2008)
Heckman two-step	- Different set of variables to determine the probability of censoring and the value of the dependent variable - It provides a rationale for zero trade flows	- Difficult to find an identification restriction - Exclusion variables are required	Bikker and de Vos (1992); Linders and de Groot (2006); Martin and Pham (2008)

Table 2.1. Summary of estimation methods

Estimation method	**Advantages**	**Disadvantages**	**References**
PPML (Poisson Pseudo Maximum Likelihood)	- It deals with the zero trade flows problem - Robust to heteroskedasticity - All observations weighted equally - The mean is always positive	- It may present limited-dependent variable bias when a significant part of the observations are censored	Westerlund and Wilhelmsson (2009); Siliverstovs and Schumacher (2009); Liu (2009); Shepherd and Wilson (2009); Martínez-Zarzoso (2011); Santos-Silva and Tenreyro (2006); An and Puttitanun (2009)
NLS (Nonlinear Least Squares)	- It deals with the zero trade flows problem	- More weight to observations with a larger variance (inefficiency). - Not robust to heteroskedasticity - Sample selection bias	Santos-Silva and Tenreyro (2006)
FGLS (Feasible Generalised Least Squares)	- It deals with the zero trade flows problem - Robust to heteroskedasticity	- The variance covariance matrix should be estimated first	Martínez-Zarzoso et al. (2007)
GPML (Gamma Pseudo Maximum Likelihood)	- It deals with the zero trade flows problem - Robust to heteroskedasticity	- Less weight to observations with a large conditional mean (less prone to measurement errors)	Martínez-Zarzoso (2011)
Helpman, Melitz and Rubinstein (2008)	- It provides a rationale for zero trade flows - Unbiased estimates	- Difficult to estimate - Additional data is required (exclusion variables)	Helpman et al. (2008); Santos-Silva and Tenreyro (2008)

2.4. Comparing estimation methods for a baseline gravity equation

The new workhorse in the estimation of the gravity equation is still unclear. Econometric estimation presents some challenges that remain unsolved as of yet. First, the exclusion of the multilateral trade resistance terms leads to biased estimates due to the omission of variables. AvW claimed that this misspecification invalidates the estimation. Second, taking logarithms and estimating by OLS in the presence of heteroskedasticity leads to inconsistent estimates as noted by Santos-Silva and Tenreyro (2006). Third, there are some aspects that may differ from one country to another but are not reflected by the regressors (i.e. regulation, political factors, technology, e-business, port efficiency, etc.). This unobserved heterogeneity should be controlled for to obtain unbiased estimates. Finally, if two countries do not trade in a given year the value of their trade would be represented by a zero in the dataset. Since the logarithm of zero is unfeasible, some information would be lost. This problem is becoming more important due to the use of disaggregated data, in which over 50% of values is zero.

2.4.1. Data and model

The sample covers bilateral exports of 80 countries over the 1980-2008 period. All the countries of the EU15, the CEE new European members, and 6 Middle East and North African (MENA) countries (Morocco, Tunisia, Egypt, Turkey, Israel and Algeria) as well as most OECD countries are included[5]. The total number of observations should be 176,960 but is reduced to 157,080 due to missing data. Data were collected from several sources, including the CHELEM-International Trade database, the CEPII database and

[5] Table 2.B.2 in Appendix B lists the countries included.

the World Bank[6].

For the sake of comparison, a gravity equation based on AvW's theoretical model will be used:

$$\ln X_{ijt} = \beta_1 \ln y_{it} + \beta_2 \ln y_{jt} + \beta_3 contig_{ij} + \beta_4 smctry_{ij} + \beta_5 \ln d_{ij} + \eta_{ij} + \eta_{it} + \eta_{jt} + \varepsilon_{ijt} \tag{2.3}$$

The dependent variable is the logarithm of the volume of exports in current dollars from country *j* to *i*, obtained from the CHELEM-CEPII database. $\ln y_{it}$ and $\ln y_{it}$ are the logarithms of nominal GDP in each country whose effect on trade is expected to be positive. $contig_{ij}$ (Contiguity), *comla* (Common language) and *smctry* (Same country) are dummy variables that take value one when two countries share a border, a language, or were the same country in the past, correspondingly. In all cases, the coefficient is expected to be positive. d_{ij} is a variable representing the geodesic distance between *i* and *j* and is obtained from the CEPII database. According to Egger and Pfaffermayer (2003), country pair specific fixed effects, η_{ij}, as well as time varying fixed effects for the importer and the exporter, η_{it}, η_{jt}, are included in the estimation in order to capture any importer or exporter time varying characteristics. These terms correct biases that arises from the fact that we are not estimating a cross-section but a panel (see BT). Due to the inclusion of these dummies, GDP terms are dropped from the estimation. However, as first noticed by Neyman and Scott (1948), the estimation of a Tobit and Probit models with fixed effects is inconsistent due to the incidental parameter problem. Hence, fixed effects are not included in these two models.

[6] The CHELEM database is previously refined using a 7-step procedure. Bilateral trade data is harmonised using reports on each of the countries involved in the transaction.

2.4.2 Results

Before estimating equation (2.3), some specification tests were conducted. First, the Likelihood Ratio (LR) and the Lagrange Multiplier (LM) tests on time and individual effects were performed. In both cases, the null hypothesis of no fixed effects is rejected. The standard F-test for the joint significance of individual and time dummies confirms this result, so it can be concluded that unobserved heterogeneity is present and OLS estimation yields biased and inconsistent estimates. A simple analysis of the residuals and the fitted values confirms the presence of heteroskedasticity in the regression (see Figure 2.2 in Appendix C). Hence, estimation with a nonlinear method is required.

Table 2.2 reports the estimation outcomes resulting from the different techniques employed. The dependent variable is the logarithm of exports in all cases except for Poisson regression, in which this variable is introduced in levels.

Overall, the estimation techniques seem to affect the magnitude but not the sign of the parameters for most gravity variables. As expected, both the exporter and importer GDP increases exports regardless of the estimation method used, while the distance reduces exports. Other gravity variables are also highly significant, and proximity (either in history or in space) tends to increase exports. Belonging to a RTA also increases trade, although it shows a moderate effect. The main differences among estimators are revealed in the magnitude of coefficients. Whereas the Heckman and panel methods show results that are more in line with the related literature, the incorrect treatment of zeros is observed to lead to an overestimation of coefficients in the Tobit and OLS estimation. These differences suggest the existence of a substantial bias in the estimation of the Tobit and OLS methods. On the other hand, PPML shows the lowest coefficients; a result that is in line with Santos-Silva and Tenreyro (2006) and

Siliverstovs and Schumacher (2009). The goodness of fit measures also reveal the existence of significant differences among the methods compared.

Columns 2 and 3 show the results for OLS adding a constant and Tobit estimates correspondingly. In both cases, the zero flows in the dependent variable are assumed to take a value of one, which is not theoretically consistent. In fact, the visual inspection of the kernel estimates reveals that Tobit coefficients are strongly biased, whereas OLS estimators present more variance than the others.

Other alternatives in the literature that do not artificially modify the dependent variable simply propose discarding the zero flows from the estimation. These are the cases reported in the first, sixth and seventh columns. The first column shows the results for the truncated OLS estimation. Most variables have the expected sign, and are highly statistically significant, though the effect of a RTA on trade is predicted to be negative, contrary to expectations. Furthermore, as mentioned before, the OLS estimation is inconsistent due to the presence of unobserved heterogeneity. Column six shows the results for the panel estimation assuming fixed effects and column seven allows the heterogeneous component to be distributed randomly. The coefficients are also significant and show the expected sign.

The last column shows the results for the PPML estimation. In this case, the dependent variable is introduced in levels instead of logarithms. Although the sign and significance are quite similar to the other estimators, PPML notably reduces the magnitude of the coefficients as well as the standard errors. Santos-Silva and Tenreyro (2006) claim that this is the preferred estimation method in the presence of heteroskedasticity.

However, none of the above methods explains the presence of zero flows. Indeed, these observations are simply dropped or censored at one. Since these procedures may lead to sample selection bias when the zeros in the sample are not random, one of the alternative solutions proposed in the literature is to use a Heckman sample selection model. While other methods treat zero flows as inexistent, Heckman considers them to be unobserved. The outcomes from the first step (Probit equation) are reported in column 4. Following Helpman et al. (2008), common language is used as an excluded variable since this variable is expected to affect the probability of exporting, but not the size of exports. Column 5 reports the results for the second step. The inverse Mills ratio is highly significant, thus confirming the existence of a sample selection bias.

Several goodness-of-fit criteria have been used in order to compare estimation methods. First, the predicted over the real value of exports in a specific year (2008) is plotted for different techniques and the dispersions of the results (Figures 2.3 to 2.9 in Appendix C) are compared. Second, the graphs of the univariate kernel density estimation are examined to gain a more accurate idea of the bias and the variance of the distribution of the predicted values in each case (Figure 2.1). Finally, Table 2.3 shows the results of three goodness-of-fit functions: the bias, the mean squared error (MSE) and the absolute error loss.

Table 2.2. Results for alternative estimation methods

	Truncated OLS	**OLS (1+X)**	**Tobit**	**Probit**	**Heckman**	**Panel fixed**	**Panel random**	**PPML**
Dependent variable	$\ln X_{ijt}$	$\ln(1+X_{ijt})$	$\ln(1+X_{ijt})$	$Pr(X_{ijt}>0)$	$\ln X_{ijt}$	$\ln X_{ijt}$	$\ln X_{ijt}$	X_{ijt}
$\ln y_i$			1.431***	0.0907***				
			(0.024)	(0.035)				
$\ln y_j$			1.513***	0.104***				
			(0.023)	(0.0342)				
$contig_{ij}$	0.129***	-0.482***	0.0462	-0.327			0.225***	0.413***
	(0.030)	(0.082)	(0.402)	(0.289)			(0.068)	(3.53e-10)
$comla_{ij}$	0.929***	2.221***	2.355***	1.606***			1.071***	0.244***
	(0.018)	(0.049)	(0.245)	(0.175)			(0.052)	(3.40e-10)
$smctry_{ij}$	0.626***	0.609***	0.609	-0.869**			0.712***	0.007***
	(0.048)	(0.147)	(0.599)	(0.375)			(0.094)	(6.30e-10)
$\ln d_{ij}$	-1.318***	-1.943***	-1.866***	-0.873***			-1.330***	-0.644***
	(0.008)	(0.024)	(0.074)	(0.063)			(0.021)	(1.59e-10)
RTAboth	-0.0625***	-0.779***	0.0436	0.757***	0.336***	0.337***	0.292***	0.441***
	(0.017)	(0.046)	(0.070)	(0.130)	(0.0147)	(0.0382)	(0.014)	(3.74e-10)
Inverse Mills Ratio					0.617***			
					(0.0908)			
Constant	14.64	-11.50***	11.50***	2.777***	5.314	16.76***	14.22***	14.91***
	()	(2.362)	(0.676)	(0.649)	(6.147)	(2.478)	(0.787)	(1.46e-07)

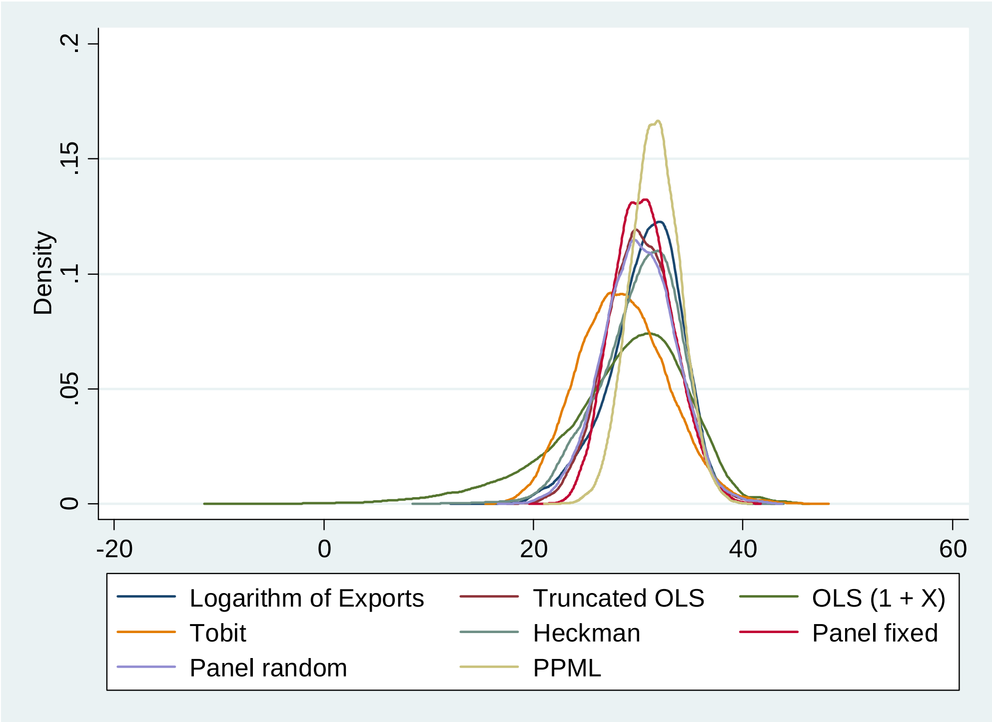
*Notes:**** denotes significant at 1%. Robust standard errors in parenthesis . All specifications except Tobit and Probit include importer and exporter time varying effects.

The main advantage to the last function, as suggested by Martínez-Zarzoso (2011), is that over- and under-estimations are not cancelled out. It is defined as follows:

$$L(X_{ijt}, \hat{X}_{ijt}) = \left| X_{ijt} \quad \hat{X}_{ijt} \right| \tag{2.4}$$

Figure 2.1 plots the kernel density estimates of the distributions of the predicted values from each method, as well as the observed data. The logarithm of exports is normally distributed and slightly right skewed. A one-to-one comparison of the methods reveals that almost all the estimators are slightly left skewed and present a bias with different magnitudes. The distribution of fixed PPML notably differs from all others in kurtosis (it shows a positive and high kurtosis and hence a smaller variance), whereas the rest tend to be platykurtic (higher variance). However, it exhibits a stronger bias. Hence, although it shows a smaller variance, the prediction is very poor for low trade values, which are overestimated. The plot of individual graphs for a cross-section (Figure 2.3 to 2.9 in Appendix C) and the different measures of goodness of fit in Table 2.3 confirm this result. On the other hand, Tobit and OLS adding a constant show a very high variance, which is related to the fact that both methods treat the zeros in the sample in an incorrect manner, thus forcing the observations to have no theoretical justification. Overall, the distribution of Heckman, truncated OLS and panel random effects seem to be closest to the real distribution.

Figure 2.1. Kernel densities of different estimators



Concerning the other goodness-of-fit criteria employed, the outcomes in Table 2.3 confirm the abovementioned results. Heckman is the preferred estimation method regarding the MSE and absolute error loss criteria, followed by Pooled OLS and panel random effects; whereas Tobit, OLS with a modified dependent variable and panel fixed effects estimation obtain the worst results.

Table 2.3. Goodness of fit

	Bias	MSE	Error loss
Truncated	**7.95E-11**	2.415	1.111
OLS (1+X)	-1.069	9.955	2.200
Tobit	-1.667	8.104	2.303
Heckman	8.86E-11	**0.950**	**0.623**
Panel fe	-4.61E-11	13.315	2.915
Panel re	-0.079	2.476	1.139
PPML	1.221	5.403	1.553

Notes: Bold values indicate the preferred estimation method in each case.

2.5. Concluding remarks

The gravity model is considered one of the most successful empirical frameworks in international economics. It has become a successful tool for the evaluation of trade policies or the calculation of trade potential associated with regional integration. However, a more detailed analysis of the theoretical underpinnings, the use of larger datasets and improvements in statistical and econometric software have highlighted new problems in estimating the gravity equation.

This chapter has provided an in-depth review of recent developments in the literature on estimation methods for the gravity equation, finding that there are at least two problems related to the log linearisation of the gravity equation that require further research as there is no consensus about the optimal method to solve them. First, the exclusion of the multilateral trade resistance terms defined by AvW, as well as the unobserved heterogeneity present in trade data leads to biased estimates due to misspecification. One usual procedure to solve this problem is to log linearise the model and to estimate it by OLS with fixed effects. However, the heteroskedasticity intrinsic to the log-linear formulation of the gravity model can result in biased and inefficient estimates when applying OLS. Second, the logarithm of zero is unfeasible. As a result, the presence of zero trade flows in data means that these observations must either be dropped or replaced by an arbitrary positive value, leading to sample selection bias and loss of information. This problem is becoming increasingly important due to the use of disaggregated datasets in which over 50% of values are zero.

An empirical exercise to compare several techniques with a dataset covering 80% of world trade has been conducted. The equation is based on the AvW specification of the gravity equation, allowing for different assumptions about the unobserved heterogeneity component. After applying several criteria to test goodness of fit, it is argued that ad hoc methods are not appropriate for estimating the gravity equation since they provide biased and inefficient estimates. On the other hand, although the use of PPML has been proposed by several authors in the literature, it does not behave so well for an aggregate dataset in the presence of unobserved heterogeneity. This chapter suggests that the Heckman sample selection model is the preferred estimation method within nonlinear techniques when data are heteroskedasticity and contain a significant proportion of zero observations.

Appendix A

Table 2.A.1. Articles using fixed effects or random effects in the estimation of the gravity equation

Article	Effects included	Data	Dependent variable
Mátyás (1998)	- Importer, exporter and time effects	11 countries; 1982-1994	Exports
Rose and van Wincoop (2001)	- Importer, exporter and time effects	200 countries; data at five-year intervals between 1970 and 1995	Bilateral trade
Glick and Rose (2002)	- Country-pair fixed effects - Symmetric country pair effects.	217 countries; 1948-1997	Real bilateral trade
Baltagi et al. (2003)	- Importer, exporter and time effects - Country-pair fixed effects - Importer-time effects - Exporter-time effects	EU15, USA and Japan with their 57 most important trading partners; 1986–1997	Real bilateral exports
Micco et al. (2003)	- Time effects - Country-pair fixed effects	22 developed countries; 1992 - 2002	Bilateral trade
De Benedictis and Vicarelli (2005)	- Country-pair fixed effects - Dynamic effects (Arellano and Bond estimator)	Each of former 11 Eurozone countries to 32 importer countries; 1991-2000	Exports
Cheng and Wall (2005)	- Country-pair fixed effects - Time effects	29 countries; 1982, 1987, 1992, and 1997	Real exports
Fratianni and Hoon-Oh (2007)	- Country-pair and time fixed effects - Random effects	143 countries; 1980-2003	Real bilateral imports
Ruiz and Vilarrubia (2007)	- Importer, exporter and time effects - Exporter-period and importer-period dummies (annual, triennial and quinquennial)	205 countries; 1948-2005	Bilateral trade
Cafiso (2008)	- Country-pair and time fixed effects	24 OECD countries (sectors 15-37, ISIC Rev. 3); 1993-2003	Exports
Fidrmuc (2008)	- Country-pair and time effects	19 OECD countries; 1980-2002	Bilateral trade flows

Table 2.A.1. Articles using fixed effects or random effects in the estimation of the gravity equation

Article	**Effects included**	**Data**	**Dependent variable**
Henderson and Millimet (2008)	- Importer, exporter and time effects - Country-pair fixed effects	US data. 25 two-digit SIC industries;1993 and 1997	Nominal value of exports
Hoon-Oh and Selmier II (2008)	- Country-pair fixed effects - Random effects	859 pairs; 1980–2001	Imports
Kavallari et al. (2008)	- Random effects	German imports of olive oil from 14 exporting countries; 1995-2006	Imports
Bussière and Schnatz (2009)	- Country-pair fixed effects	61 countries; 1980-2003	Bilateral trade
Yu (2010)	- Fixed effects	157 countries; 1962–1998	Exports

Table 2.B.1. **Articles related to the problem of zero-flows and heteroskedasticity**

Article	**Data**	**Estimation methods**	**Dependent variable**	**Simulation studies**
Santos-Silva and Tenreyro (2006)	136 countries; 1990	PPML, NLS, GPML, OLS, ET-tobit, OLS ($y > 0.5$) OLS (y+1)	Trade	- PPML, NLS, GPML OLS; OLS(y + 1); truncated OLS ET-tobit. - Four different patterns of heteroskedasticity
Martínez-Zarzoso (2011)	3 datasets: 1) 180 countries; 1980-2000 2) 47 countries; 1980-1999 3) 65 countries; data for every 5 years over 1980-1999	FGLS, GPML, Poisson, Heckman	Exports	- OLS, NLS, GPML, PPML and FGLS
Helpman et al. (2008)	158 countries; 1970-1997	HMR, NLS, semi-parametric, non-parametric	Exports	No
Martin and Pham (2008)	136 countries; 1990	Truncated OLS, ET-Tobit, PPML, Heckman ML, Heckman 2SLS	Bilateral trade	- Truncated OLS, OLS (y+1), truncated NLS, censored NLS, GPML, PPML, truncated PPML, ET- Tobit, Poisson-Tobit, Heckman
Santos-Silva and Tenreyro (2008)	158 countries; 1986	HMR, NLS, semi-parametric, non-parametric, GPML	Exports	No
Burger et al. (2009)	138 countries; 1996-2000	OLS, PPML, ZIPPML, BPPML	Exports	No

Table 2.B.1. **Articles related to the problem of zero-flows and heteroskedasticity**

Article	**Data**	**Estimation methods**	**Dependent variable**	**Simulation studies**
Siliverstovs and Schumacher (2009)	22 OECD countries; 1988-1990. Disaggregated data: 25 three-digit ISIC Rev.2 industries	OLS, PPML	Trade	No
Westerlund and Wilhelmsson (2009)	EU and other developed countries; 1992-2002	OLS, fixed effect PPML	Nominal imports	- OLS, truncated OLS, OLS (y+1), PPML - Two patterns of heteroskedasticity
Yu (2010)	157 countries 1962–1998	OLS, fixed effects, IV, PPML	Exports	No

Appendix B

Table 2.B.2. List of countries included in the sample

Albania	Gabon	Paraguay
Algeria	Germany	Peru
Argentina	Greece	Philippines
Australia	Hong Kong	Poland
Austria	Hungary	Portugal
Bangladesh	Iceland	Romania
Belarus	India	Russian Federation
Belgium and Luxembourg	Indonesia	Saudi Arabia
Bolivia	Ireland	Singapore
Bosnia and Herzegovina	Israel	Slovakia
Brazil	Italy	Slovenia
Brunei Darussalam	Japan	South Korea
Bulgaria	Kazakhstan	Spain
Cameroon	Kenya	Sri Lanka
Canada	Kyrgyzstan	Sweden
Chile	Latvia	Switzerland
China	Libyan Arab Jamahiriya	Taiwan
Colombia	Lithuania	Thailand
Côte d'Ivoire	Macedonia	Tunisia
Croatia	Malaysia	Turkey
Czech Republic	Mexico	Ukraine
Denmark	Morocco	United Kingdom
Ecuador	Netherlands	United States
Egypt	New Zealand	Uruguay
Estonia	Nigeria	Venezuela
Finland	Norway	Vietnam
France	Pakistan	

Appendix C

Cross-validation for the different estimation methods in year 2008

Figure 2.2. Heteroskedasticity in data. Distribution of errors

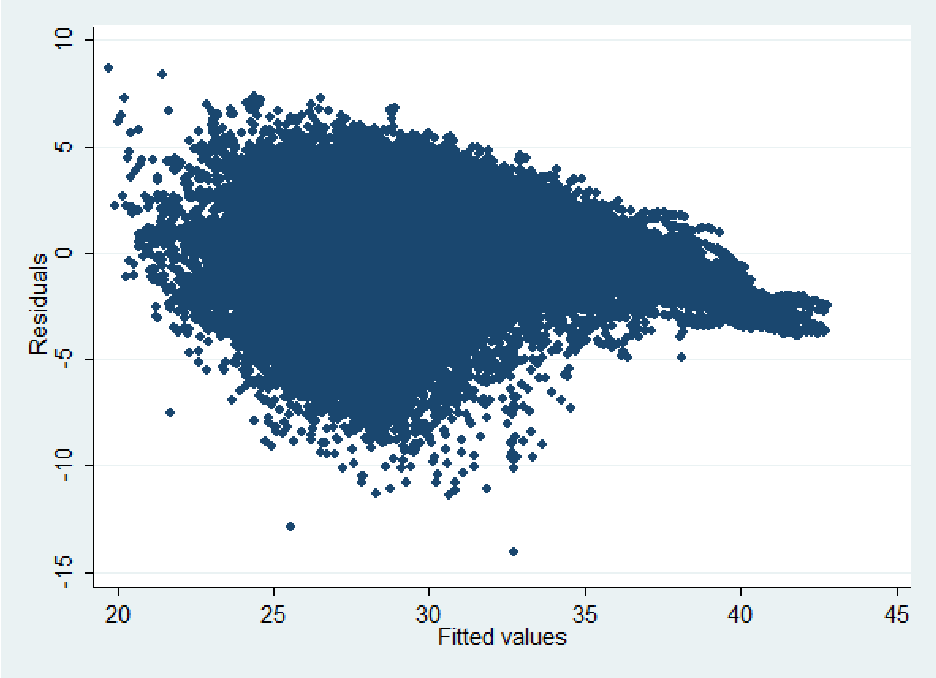


Figure 2.3. Truncated OLS

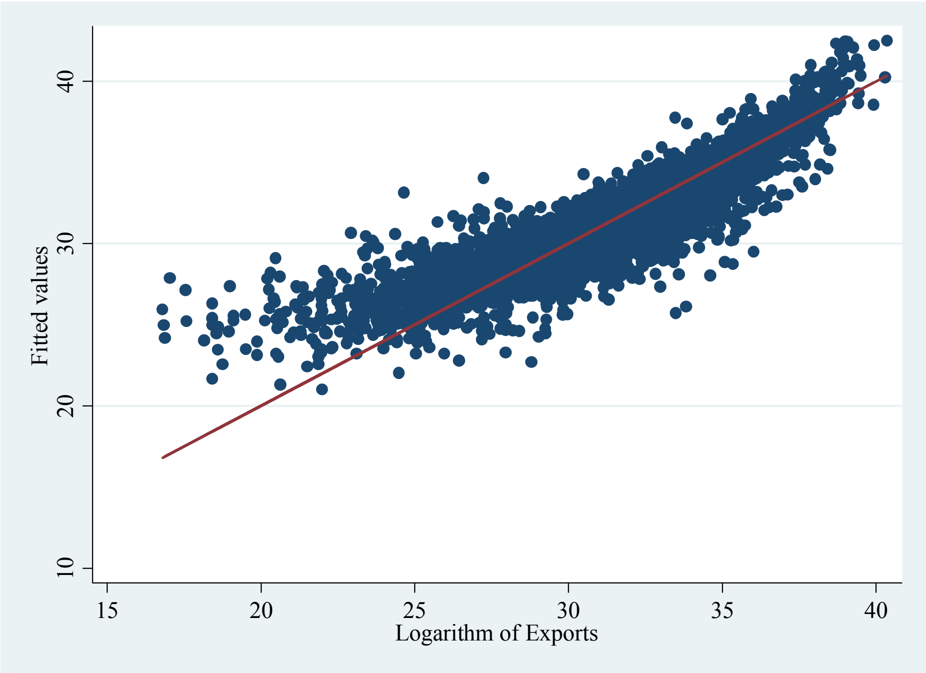


Figure 2.4. OLS (1+X)

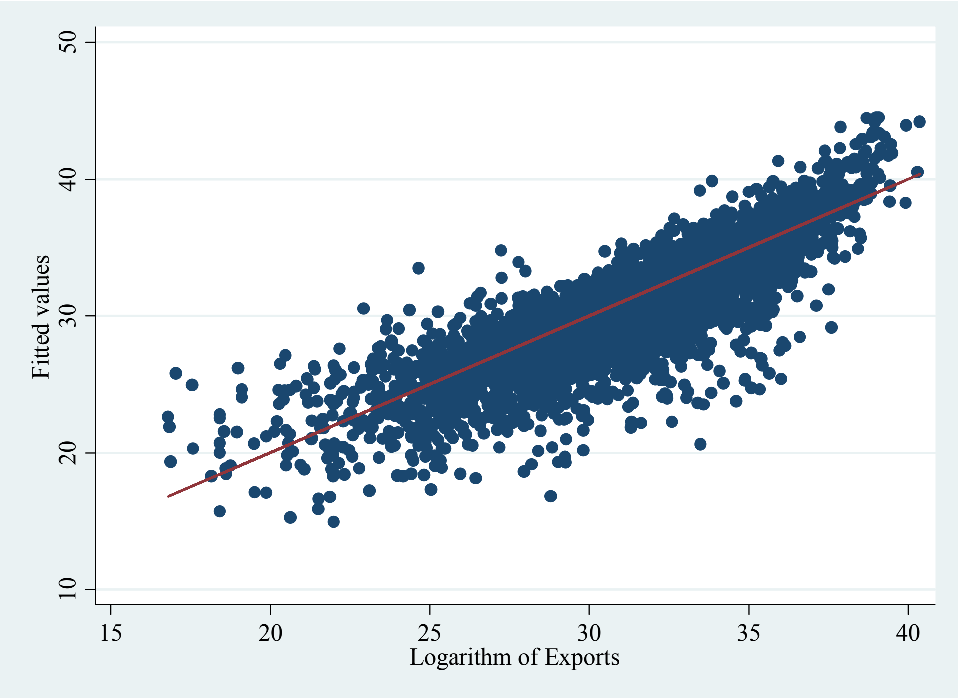


Figure 2.5. Tobit

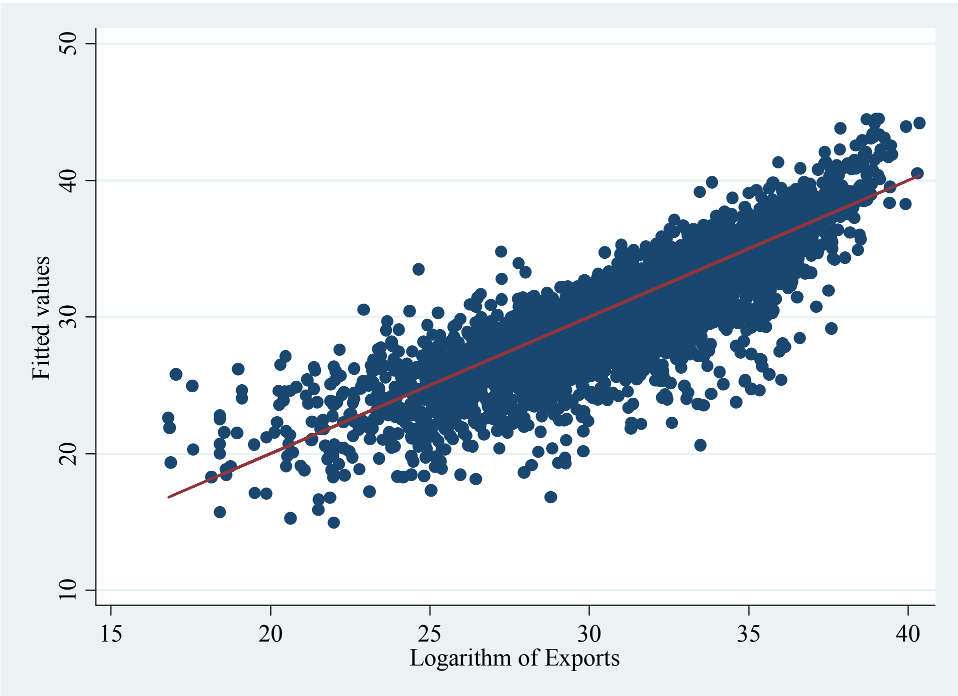


Figure 2.6. Heckman model

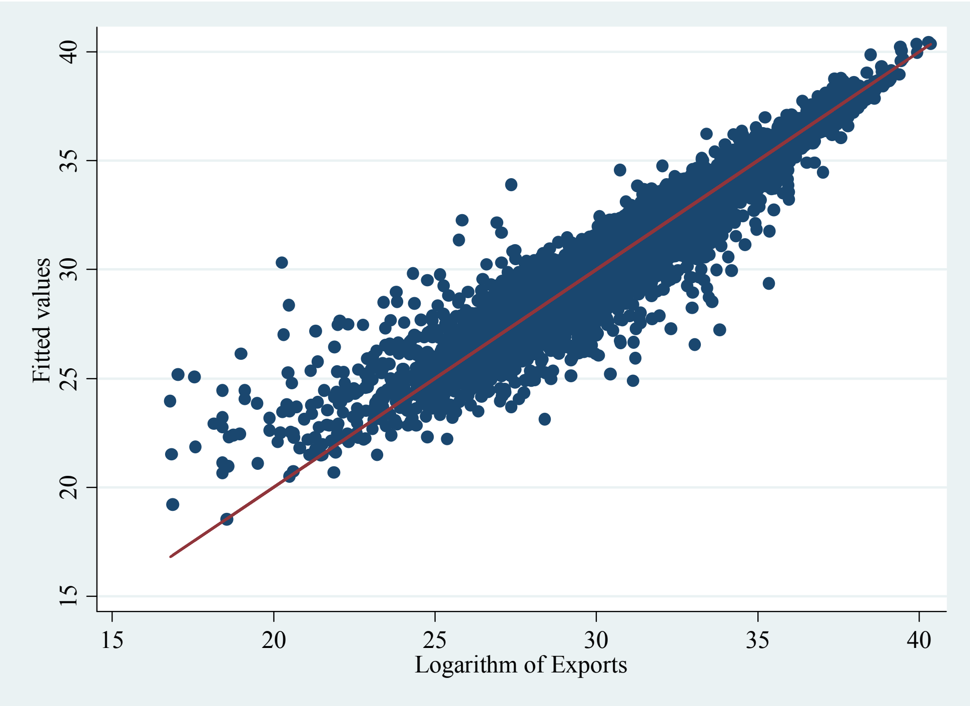


Figure 2.7. Panel fixed effects

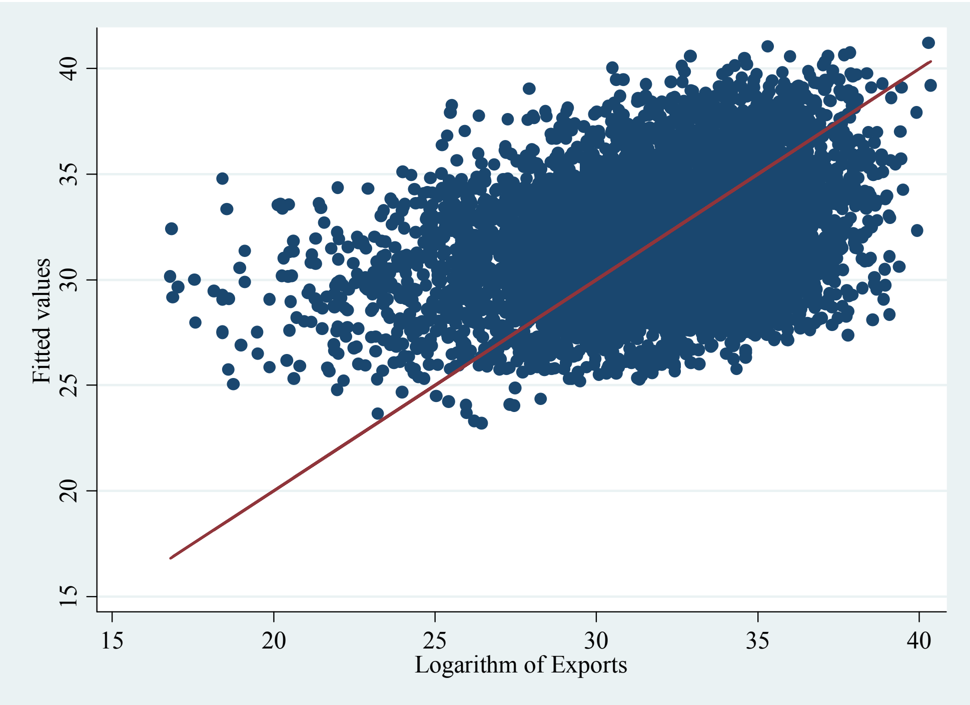


Figure 2.8. Panel random effects

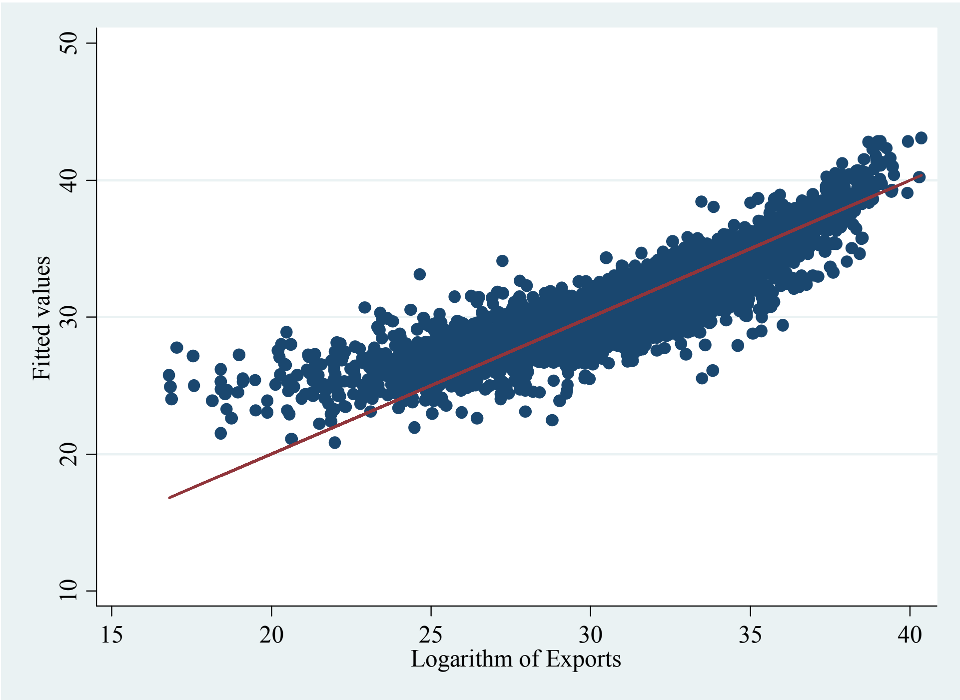
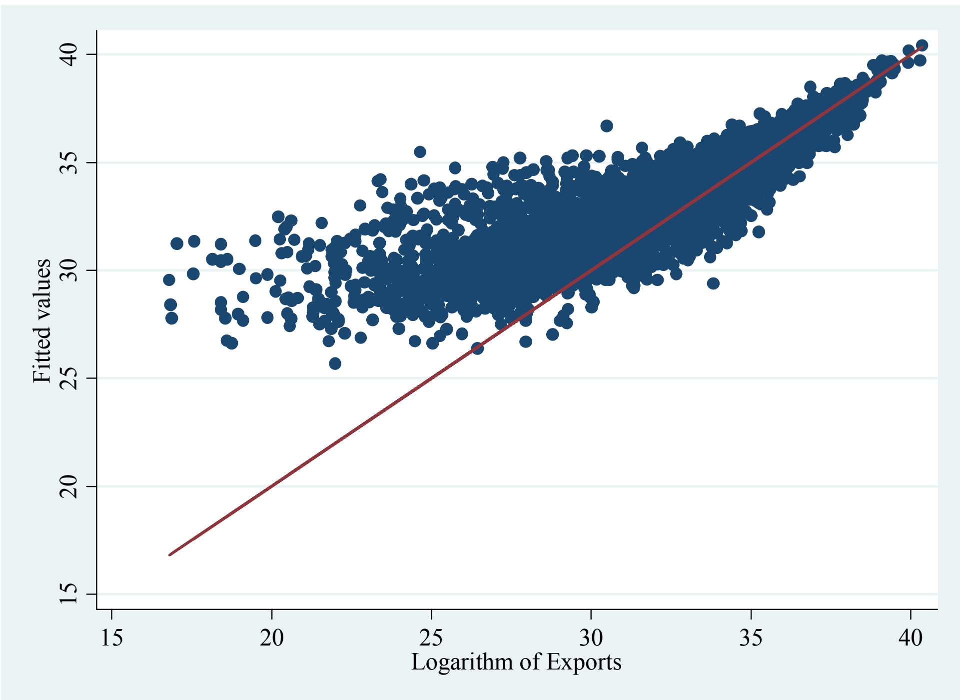


Figure 2 9. Poisson Pseudo Maximum Likelihood



Chapter 3
Ten years after: Did euro boosted trade?

Abstract. In this chapter we investigate the impact of the euro adoption on trade. To this end, we estimate a gravity equation including the RER level and volatility to capture the additional effect the euro could have had apart from the one coming from the elimination of ERV. We use a large sample of 80 countries during the period 1967-2009, covering 80% of world trade. Concerning the volatility issue, we find that the elimination of volatility boosted export per se, especially before 1999. Then, the possibility to peg to the euro could boost the trade of third countries and between these third countries. The common currency has had a positive impact on EMU exports to other EMU countries, though it has reduced Eurozone's imports from third countries while it had not a significant impact on Eurozone's exports to other countries. CEE countries represent an exception since both their exports to EMU and imports from EMU have been boosted by euro. The analysis for individual EMU members reveals the existence of a good deal of variation in the effect of the euro across member countries since the euro has only boosted the exports of 4 of 12 EMU countries. Concerning the impact of euro over time, the euro significantly boosted intra-EMU trade starting in 1999, with the effect reaching its maximum in the 2003-2005 period.

3.1. Introduction

More than ten years after the advent of the euro, the common currency has thrown in 2010 its worse crisis. The costs of loosing independence in monetary policy in case that members face asymmetric shocks has appeared clearly during the crisis faced by some countries like Greece. The crisis has raised all the issues raised 10 years ago by Eurosceptics about the difficulties to share a common money without significant fiscal and political coordination. To convince Eurosceptics of the benefits of the common currency, the defenders of a unique European currency bet on large positive effect of the euro on trade and investment. Due to the policy relevance of the issue, in particular for the European countries that are still thinking about joining the EMU, it is important to have a robust evaluation of the benefits the euro had on trade and could still have since the debate is in the air. Our objective is to assess the effect of EMU on trade among the members of the Eurozone and between the Eurozone's members and other countries. To this end, we use the longer post euro period available at this date (our period ranges from 1967 to 2009) and try to control for all the possible effects the introduction of the euro had.

The main motives to expect large trade effect when adopting the euro were based on the beliefs that elimination of transaction costs and elimination of ERV should promote trade. Transaction costs could vary from 13.1-19€ billion according to the Emerson report (Emerson, 1992) and could represent a 0.3-0.4% of GDP of exporters. Then, the impact was expected to be rather large. The expectations concerning the gains to obtain from the elimination of ERV were less clear-cut. It might exist financial instruments to hedge against ER risks. Though, these instruments are costly. In sum, the ex-ante effects of the elimination of ERV and transaction costs were difficult to evaluate. Other potential gains were underlined but even more difficult to quantify. For instance, the euro may increase the degree of transparency of policies and make transactions among members similar to national ones. To change the money is a decision perceived as irrevocable, what in turn reduces the uncertainty and may increase all transactions including trade. Few years after the introduction of the euro, Baldwin et al. (2008) and Baldwin et al. (2005) have added another optimistic contribution to the debate: the effect of the elimination of ERV when this volatility is small could have lead to a biggest increase in trade than a similar reduction when volatility is high. That is, they suggested a nonlinear relation between ERV and trade. They argue that for lowest levels of the volatility, the elimination of this risk will lower export costs below a threshold that would allow more small firms to export. Since small firms represent a most important share of European firms, this would have a positive and non-proportional effect on trade[7]. Gil-Pareja et al. (2008) used this argument to explain why they found that monetary agreements among OECD countries have boosted trade, even when the elimination of volatility is controlled for.

From the entry of the euro, the euro first appreciated comparing to the dollar and then, went depreciating until 2009. Now it fluctuates around its original level. The evolution of the real competitiveness followed a similar evolution (European Commission, 2010). Obviously, the overall competitiveness of the EMU affects the euroland trade and can bring changes in the price elasticities of import and export and substitution. All in all, the world demand and supply of euro drive its level. But, the weight of the euro in the Central Bank assets of foreign countries and in debt has increased. This is a way to diversify the variability of their ER and the value of their assets. This strategy in turn, affects the level of the euro

[7] A basic result of this model is that a reduction in ERV raises both the sales per exporting firm (intensive margin) and the number of exporting firms (extensive margin), because the minimum size-class of firms that export falls as volatility decreases. Berthou and Fontagné (2008) and Esteve-Pérez et al. (2011) offer an empirical verification of this proposal using respectively data for French and Spanish firms.

and, afterwards, the level of trade of the Eurozone. Then, we argue that the ER level should be taken into account to quantify in a more realistic way the impact of the euro on trade.

We use an augmented gravity equation that explicitly takes into account the level and volatility of the bilateral ER, the presence of RTAs and the EMU. By this way, we are able to separate the specific trade effect of the euro from the trade impact of trade and monetary agreements. We discuss the fit of different estimation procedures in terms of error losses applied to a large dataset of bilateral exports for 80 countries (93% of world trade) over the period 1967-2009.

To anticipate our most important findings, our study confirm that the common currency has had a positive impact on EMU exports, additionally to the fact that the elimination of ERV boosted export per se. This result reinforces the conclusions obtained by De Nardis (2004), which used a shorter period. The analysis for individual EMU members reveals the existence of a good deal of variation in the effect of the euro across member countries. Besides that, we provide evidence that the EMU has contributed to the expansion of some CEE countries exports. Finally, our study shows that the estimation technique leads to similar results for a basic model of trade flows. Additionally, it does not have a crucial impact on the conclusions concerning the effect of exchange-rate regimes – defined by the RER level and volatility – on exports.

The rest of the chapter is organized as follows. In the next section we review the findings and the non solved challenges of the empirical literature that study the impact of euro trade effects. Section 3.3 presents the empirical model and the data. In section 3.4 the estimation method is detailed. We comment the results in sections 3.5 and 3.6. Finally, some conclusions are provided in Section 3.7.

3. 2. Literature Review

Due to the success of the gravity equation to accurately reproducing real trade flows, this empirical model has been widely used by empirical researchers to study the sensibility of trade flows to a wide range of variables. Due to the relevance of the issue, the effect of monetary arrangement and of the common currency has also been studied in this framework. Table 3.A.1 in Appendix A provides a review of most recent literature on this topic. Contemporaneous of the beginning of the euro, the pioneer and famous article of Rose (2000) concluded that currency unions could triple trade among members. It became a challenge to confirm or detract these extraordinarily optimistic results and numerous studies have focused on the question of the euro impact on trade after them. Concretely, the study of Rose (2000) was based on a cross-section study involving heterogeneous countries and currency unions from 1970 with very different contexts than the Eurozone; then economists thought that Rose's results were overestimating the effects of a currency union. But the idea that the euro could have a significant impact on trade gained weight in the debate and several studies after that have founded impacts quite smaller than Rose did, but still economically large.

Since the euro was adopted, and data for trade has become available, numerous studies have re-examined the question of the ex-post effects of the creation of the EMU on trade. There are several challenges when dealing with this issue. One of them is to separate the impact of the elimination of ERV on trade flows from other euro effects as the elimination of transaction costs and other permanent changes associated with the new currency. This implies computing the ERV for all country pairs and a sufficiently long period and large sample to capture differences in this variable among partners and time. Examples of works analyzing the impact of EMU on trade and taking into account ERV are numerous and varied in their estimation methods and datasets; Table 3.A.1 in appendix summarizes most of them. An important specificity of the EMU is that all members are members of the Common market, and then EU effects should be distinguished from EMU effect. Unfortunately, many studies do not take into account the additional effect of other RTA[8]. Another branch of gravity literature controls for the effect of RTAs and take into account ERV but do not explicitly focus on the euro effect. Some examples of this kind of studies are Rose (2000), Rose and Engel (2002), Clark et al. (2004) and Tenreyro (2007). These authors measure the impact of ERV as the reduction in trade flows provoked by the increase in volatility by one standard deviation around its mean. The results range from 4-6% in Tenreyro (2007) to 13% in Rose (2000). Concerning the RTA coefficient, the results obtained differ slightly from one another. Tenreyro (2007) finds a negative influence of RTA on trade of 45%, while the rest of authors obtain positive and significant results that varies from 32% to 145%.

Recent empirical research includes ERV as well as the euro impact and the effect of RTAs on trade. Gil-Pareja et al. (2008) include a sample of 25 OECD countries for the period 1950-2004 to study the impact of monetary agreements on trade. They conclude that these agreements have a similar effect on bilateral trade among member countries as RTA once volatility is controlled for. De Nardis (2004) suggest that political and economic context of the euroland economies that pre-existed or accompanied the formation of EMU may have given rise to an independent increase in the share of intra-area trade that would bias estimates if trends are not taken into account. He explicitly addresses the issue of the persistence of trade using a dynamic panel framework. He concludes that euro had a short-run effect of 9-10 per cent on intra-EMU trade and a 19% in the long run. They qualify this impact of small. This is a very interesting contribution but their estimations are based on a

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[8] Baldwin et al. (2005), Barr et al (2003), Brouwer et al. (2008) Dell'Ariccia (1999) and De Grauwe and Skuderlny (2000) all study the EMU effects controlling for ERV.

very short post euro period (1980-2000) that does not allow them to fully address this question[9]. Other examples are Baak (2004), Maliszewska (2004), and Bussière et al. (2008).

Micco et al. (2003) offer a very complete study to quantify the early effects of the euro with data for the 1982-2002 period. They take into account the level of the real exchange rate (RER) and their sample includes 22 industrialized countries. Their results suggest that the euro had a noticeable impact on trade (between 4 and 16%), even at this early stage. Furthermore, EMU countries seem to have increased their trade with non-EMU countries. Though, they do not take into account ERV, and then it is not possible to know if this variation is only due to the elimination of volatility or to other effects. Flam and Nostrom (2006) also takes into account the RER level but do not control for ERV or RTA. They find rather large effects of euro on trade –17% for the 99-2001 period and 28% for 2002-2006 –. Note that the timing they found is different from the ones of Micco et al. (2003) and Belke and Spies (2008) whose results suggest that the euro boosted trade more at the beginning than at the end of the period they consider. Thus, the periods used by the latter are shorter than the one of the former.

Baldwin et al. (2008) argue that using time-invariant dummies for countries will leave a time-varying component in the errors that may bias the studies of this kind. This would explain why authors find a larger euro effect when they use longer datasets. Obviously the longer the dataset is the worse job a time-invariant dummy does in capturing the time-varying policy changes. To solve this problem, they interact EU dummies with time-varying indicators of the integration of the EU and the EMU with some indicators of the financial and monetary integration to take into account the progressive achievements of these agreements. They still find a positive and highly significant, but small – about 2%– EMU

[9] In this line, Bun and Klaassen (2007) and Berger and Nitsch (2008) find that the time trend reduces or makes the euro effect disappear. Though, they do not take into account ERV.

effect on trade while they find EMU to be trade diverted.

Results in almost all cases show a positive EMU effect on trade, though notably smaller than that predicted by Rose (2000), ranging from 2.6% for the most pessimistic to 112% for the most optimistic. Most of them conclude that EMU has had a positive impact on trade flows with non-EMU countries and that there are still potential trade increases associated with EMU enlargement and within EMU members. Furthermore, there is asserting consensus to consider that joining a monetary arrangement has an additional effect apart from the mere reduction of ERV.

Now, data are available for a sufficiently long post-euro period to have a more precise ex-post evaluation of the euro effect. The methodological debate about the estimations of the gravity model has also evolved rapidly in recent years and this re-examination of the euro trade effect should take these considerations into account. Additionally, most of the recent literature makes the choice to focus on a reduced sample of developed countries which probably is more accurate when one focus on the effect of the common currency on the euroland trade. Though, considering a larger sample is more accurate to study the question of the impact of the euro on its trade with non EMU countries and the opportunity that euro offers to other EU countries in terms of trade. For that reason it is important to use an estimation technique that allows dealing with the heterogeneity of countries. Finally, most of the articles above mentioned do not include the ER level or the ERV in the specification. We claim that these variables have a significant effect on trade and should be included. Until the moment, any of the articles includes at the same time the level and volatility of the ER, as well as dummies for EMU and other RTAs. As far as we concern, this is the first work studying jointly these effects with alternative estimation methods.

3.3. Methodology: the gravity model

3.3.1. Baseline empirical model

We consider the augmented version of the Anderson (1979) model proposed by AvW. This model is overall interesting to the extent that the discussion of the multilateral resistance may matter for heteroskedasticity considerations. For instance, as GDP increases, remote countries with higher trade costs will tend to diversify their production, becoming less open to trade. However, if they are located near to other countries their specialization is likely to be higher, and trade flows in that case will become more frequent. This divergence in trade patterns can thus lead to a variance that is a function of one of the regressors (level of income). Although the specification proposed by AvW has become very standard in the gravity literature, its estimation is not straightforward since it includes two multilateral resistance terms[10] -one for the importer and one for the exporter-, which are not observed. In this sense, Feenstra (2002) propose to include importer and exporter fixed effects to account for the specific country multilateral resistance term. The coefficient of the dummies for the importer and the exporter should reflect the multilateral resistance of each country. Besides that, Egger and Pfaffermayr (2003) propose the inclusion of country pair fixed effects in order to capture all those bilateral characteristics that are specific to each pair of countries.

A second aspect that requires attention is the introduction of a distance variable to proxy for transaction costs. It is commonly accepted that geographical distance may be a poor approximation of all the economic barriers for international trade. In equation (3.1),

[10] These 'multilateral resistance' variables are denoted by P_i and P_j and capture the fact that bilateral trade flows do not only depend on bilateral trade barriers but also on trade barriers across all trading partners.

transaction costs (t_{ij}) are proxied by bilateral distance[11], d_{ij}. However, this variable should be reinforced in order to control for other factors that may affect trade:

- Contiguity. This variable takes value 1 if trade partners share a common border. Its effect on trade is expected to be positive.

- Common language: sharing a language should make all transaction easier and costless.

- Same country: this variable reflects the fact that one country has been divided. It is especially important in our dataset, since several countries of the former URSS are included.

Concerning the proxy for supply and demand sizes, the common practice is to use GDP for the importer and for the exporter. In some cases GDP per capita is also introduced as a proxy for capital-labour intensities (not only factor endowments of a country).

We end up with the following baseline model:

$$\begin{aligned} \ln X_{ijt} &= \beta_1 \ln y_{it} + \beta_2 \ln y_{jt} + \beta_3 contig_{ij} + \beta_4 smctry_{ij} + \beta_5 \ln d_{ij} + \\ &+ \eta_{ij} + \eta_i + \eta_j + \lambda_t + \varepsilon_{ijt} \end{aligned} \tag{3.1}$$

The dependent variable is the logarithm of the volume of exports in constant dollars from country *j* to *i*. $\ln y_{it}$ and $\ln y_{it}$ are the logarithms of real PPP-converted GDPs in each country; their effect on trade is expected to be positive. $contig_{ij}$ (Contiguity), *comla* (Common language) and *smctry* (Same country) are dummy variables that take value 1 when two countries share a border, a language, or were the same country in the past, respectively. In all cases, the coefficients are expected to be positive. d_{ij} is a variable representing the

[11] There is not a unique opinion about how distance should be measured. The most common measures are the great circle formula and the distance between the two principal cities. See Wei (1996), Wolf (1997), and Head and Mayer (2000) for further information.

geodesic distance between *i* and *j* and is obtained from CEPII database. It is expected to have a negative influence on trade. η_i and η_j are country specific fixed effects, η_{ij} are the country pair specific effects and λ_t denotes the time effects. Data are collected from several sources, including CHELEM-International Trade database for the export values and GDP, CEPII's database for gravity variables, World Bank data and IFM Statistical Yearbook for prices indexes.

Our sample includes 80 countries. It includes all the countries of the EU15 and the CEE new European members, 8 Middle East and North African (MENA) countries (Morocco, Libya, Tunisia, Egypt, Turkey, Saudi Arabia, Israel and Algeria), all the OECD and 18 Asian countries[12]. The period considered ranges from 1967 to 2009. Hence, the total possible number of observations is 271,760. The available number of observations is reduced to 209,448 due to the presence of zero flows.

3.3.2. Specification with exchange rate variables

Traditionally the ER regime has not received enough attention in the gravity literature. Both, the level and volatility of ER are variables affecting international trade via export price; however, articles using cross-section data have not focused on these variables, since they were unable to capture variations in the ER level. Thus, panel data is the appropriate framework to evaluate the effect of ER level on exports. If the Marshall-Lerner condition is fulfilled, which is generally the case when considering long-run elasticities, a real appreciation has a negative impact on exports through a decrease in competitiveness (demand effect) or a comparative increase of profitability of traded good sector against non-traded goods (supply effect). Even when market structures are taken into account (for

[12] Table 2.B.1 in chapter 2 lists the countries included.

instance when they give rise to pricing to market strategies), an appreciation in the RER leads to a worsening of the competitive position of the economy, and consequently to a rise in imports and a fall in exports. This fact is now well documented and it is robust to the use of alternative measurement strategies even if aggregate demand and supply elasticities also depend on the structure of specialization in each country. Thus, it is surprising that so much empirical models do not take into account the RER level. Another way to take into account the fluctuations of RERs is to use data in current prices in a common currency at current ERs; though, as Flam (2009) points out, the inconvenient of this method is that we cannot separate the effect of GDP from the effect of RER.

The impact of ERV on trade is more controversial, both in theory and empirical analysis. In theory, an increase in volatility could either increase or decrease trade, depending on the risk aversion of firms or on the shape of the production functions. Looking at empirical analysis suggests that the measured effects of ERV on trade can be either very low and little significant or significantly negative though minor in magnitude. McKenzie (1999) points out that the elasticity of trade flows to ERV can be positive or negative, and the results depend on the precise measure of volatility, the estimation technique and the sectors and countries concerned. The impact of ERV actually differs according to the countries under study: Sauer and Bohara (2001) show a negative impact of this variable on African and Latin American exports and a non-significant impact on Asian exports and on developed countries exports; Frankel et al. (1995) evidence a significant negative impact on trade flows across Asian countries on a cross-section basis; Rose (2000) finds it to be a significant and systematic impediment to trade for an extensive sample of countries and Gil-Pareja et al. (2008) find a statistically significant negative effect on trade. Tenreyro (2007) develops an instrumental-variable (IV) version of the PPML estimator to deal with the endogeneity and the measurement error of ER variability estimator. Results indicate that

NER variability has not a significant impact on trade flows. Mukherjee and Pozo (2011) analyze the real ERV effect on trade using semiparametric regression methods. They find that large ERV depresses trade, but the impact of uncertainty on trade volume fades as volatility grows.

Taking equation (3.1) as starting point, we estimate three additional specifications (equations 3.2, 3.3 and 3.4) to control for the effects of ER, ERV and trade and currency agreements.

In the second specification we measure the sensitiveness of exports to exchange-rate regimes introducing the RER level and volatility:

$$\begin{aligned} \ln X_{ijt} &= \beta_1 \ln y_{it} + \beta_2 \ln y_{jt} + \beta_3 contig_{ij} + \beta_4 smctry_{ij} + \beta_5 \ln d_{ij} + \beta_6 \ln rer_{ijt} \\ &+ \beta_7 vol_{ijt} + \eta_{ij} + \eta_i + \eta_j + \lambda_t + \varepsilon_{ijt} \end{aligned} \tag{3.2}$$

where:

rer_{ijt} is the real exchange rate, computed using CPI and defined as the relative price of *j* to *i* (an increase therefore signals a real depreciation of the currency of *i* compared to *j*).

vol_{ijt} is a measure of volatility, defined as the standard deviation of the rate of change of the volatility of the monthly RERs for a given year *t*, computed as:

$$vol_{ijt} = \sqrt{\mathrm{var}(|\ln rer_{ij\tau} - \ln rer_{ij\tau-1}|)_{\{\tau=1\rightarrow 12\}}} \tag{3.3}$$

where $rer_{ij\tau}$ is the monthly real exchange rate of *j* to *i* for month τ in year t.

3.3.3. Specifications with RTA and EMU variables

In this section we control for the effects of trade agreements[13] and EMU on bilateral trade. We include four dummies indicating if one or both trade partners have a trade or a monetary agreement. Then, we capture how the common currency and RTA affect exports to countries belonging to the agreement and exports to third countries:

$$\begin{aligned}\ln X_{ijt} &= \beta_1 \ln y_{it} + \beta_2 \ln y_{jt} + \beta_3 contig_{ij} + \beta_4 smctry_{ij} + \beta_5 \ln d_{ij} + \beta_6 \ln rer_{ijt} \\ &+ \beta_7 vol_{ijt} + \beta_8 RTAone_{ijt} + \beta_9 RTAboth_{ijt} + \beta_{10} EMUone_{ijt} + \beta_{11} EMUboth_{ijt} \\ &+ \eta_{ij} + \eta_i + \eta_j + \lambda_t + \varepsilon_{ijt}\end{aligned} \quad (3.4)$$

where:

RTAone and *RTAboth* take value one when one or both countries have a regional trade agreement, and zero otherwise. We intend to capture possible creation or diversion effects; a positive sign for *RTAboth* would imply that belonging to a RTA has a creation effect while a negative sign for *RTAone* could indicate a diversion effect for exports or imports.

EMUone and *EMUboth* take value one when one or both countries respectively belong to EMU, and zero otherwise. *EMUboth* allows assessing the effect of EMU on exports inside the Eurozone. A positive sign would indicate a positive effect of the common currency on EMU exports to the Eurozone, apart from the effect of the non-tariff regime among these members and once the effect of the elimination of ERV is controlled for. A positive effect for *EMUone* would indicate that the euro has favoured exports and imports between the Eurozone and third countries.

[13] Some articles related are Frankel et al. (1995), Sapir (2001), Soloaga and Winters (2001), Greenaway and Milner (2002), Martínez-Zarzoso and Nowak-Lehman (2003), Fratianni and Oh (2007) or Oh and Selmier II (2008).

Finally, in a fourth specification, we disentangle the effect of EMU on the exports and imports of Eurozone members. To this end, we substitute *EMUone* by two dummies (*EMUimp* and *EMUexp*) to distinguish among the cases in which only the exporter or the importer belongs to the EMU.

$$\begin{aligned}\ln X_{ijt} &= \beta_1 \ln y_{it} + \beta_2 \ln y_{jt} + \beta_3 contig_{ij} + \beta_4 smctry_{ij} + \beta_5 \ln d_{ij} + \beta_6 \ln rer_{ijt} \\ &+ \beta_7 vol_{ijt} + \beta_8 RTAone_{ijt} + \beta_9 RTAboth_{ijt} + \beta_{10} EMUimp_{ijt} + \beta_{11} EMU\exp_{ijt} \\ &+ \beta_{12} EMUboth_{ijt} + \eta_{ij} + \eta_i + \eta_j + \lambda_t + \varepsilon_{ijt}\end{aligned} \tag{3.5}$$

where *EMUimp* (*EMUexp*) takes value 1 if the importer (exporter) involved in the trade flow belongs to the EMU and zero otherwise. A negative sign of the variable *EMUimp* would imply a diversion effect of EMU; EMU countries would be substituting their imports from the rest of the world by imports of EMU countries. A negative sign for *EMUexp* would indicate a geographic reallocation of exports of the members of the Eurozone in detrimental of third countries.

3.4. Estimation methods

The new workhorse in the estimation of the gravity equation is still unclear. Among the econometric challenges that remain unsolved until the moment, at least four could be remarked. First, the exclusion of the multilateral trade resistance terms leads to omitted variable bias. Second, the presence of heteroskedasticity in trade data leads to inconsistent estimates when estimating by OLS (Santos-Silva and Tenreyro 2006). Third, some aspects affecting trade are not reflected by the regressors. For instance, regulation, port efficiency, e-business, political factors, technology, etc. may differ from one country to another. This unobserved heterogeneity is not easily quantifiable, but should be controlled for to obtain unbiased estimates. Finally, the existence of zeros in the dataset provokes a loss of information since the logarithm of zero is unfeasible, and thus leads to biased estimates. This problem is becoming more important due to the use of disaggregated data, in which over a 50% of values is zero[14].

Every method presents important advantages and disadvantages and it cannot be asserted that any of them outperforms absolutely the others. For that reason, it is frequent in the literature to include several estimation methods using the same database. In our dataset the percentage of zeros represents only a 10% of the total, so we will focus on the first three problems. We will include fixed effects to avoid the problem of omitted variable bias; we will use panel estimation methods to solve the issue of unobserved heterogeneity; and finally, we will compare these results with a PPML estimator, which deals with the problem of heteroskedasticity in trade data. Results are reported in Appendix (Table 3.B.1).

Previously, we have conducted some specification tests. If unobserved heterogeneity is present, OLS estimation yields biased and inconsistent estimates. Hence, we have tested the existence of fixed effects using Likelihood Ratio (LR) and Lagrange Multiplier (LM) tests on time and individual effects. In both cases, we reject the null hypothesis of no fixed effects. In addition, the standard F-test for the joint significance of individual and time dummies confirms our results. We conclude that OLS results are biased and inconsistent, and should not be used as estimation method in this case.

In order to choose between fixed and random effect models, we have performed a Hausman test. Under the null hypothesis, the random effect model is assumed to be consistent and efficient. In all cases we reject the null and conclude that the random effect

[14] Recently, the problem of the zero flows has been revisited. The literature distinguishes several methods of dealing with that problem. Truncation (elimination) or censoring methods have been widely used. However, these methods have not a strong theoretical support and do not guarantee consistent estimates. Alternative solutions are Tobit estimation, Poisson Pseudo Maximum Likelihood estimation, Nonlinear Least Squares (NLS), Feasible General Least Squares (FGLS) and Helpman et al. (2008) procedure.

model is not appropriate; consequently we use the within fixed effects estimator. We include vectors of fixed effects for exporter, importer and time and we apply the within transformation to each pair of countries. Hence, we are controlling for country, country pair and time fixed effects. We have implemented the White's general test in OLS regressions and the modified Wald statistic for groupwise heteroskedasticity in the fixed effect models to test for the presence of heteroskedasticity. In all cases the null hypothesis of homoskedasticity is rejected, thus we compute robust standard errors.

To compare among the three estimation methods, we compute a loss function as in Martínez-Zarzoso (2011), based on the comparison of the absolute error loss:

$$L(X_{ijt}, \hat{X}_{ijt}) = \left| X_{ijt} - \hat{X}_{ijt} \right| \qquad (3.6)$$

The main advantage of this method is that over and underestimations are not cancelled out. We also apply the bias and mean squared error (MSE) criteria to compare different models; results can be found in the appendix (Table 3.B.2). Comparing the different results for all specifications leads us to conclude that differences among pooled OLS and panel fixed effects are not very large, whereas PPML estimates show larger losses.

3.5. Impact of EMU on exports

Turning to our results, we first describe briefly the results of the baseline model (equation 3.1), then analyze the effect of the RER level and volatility on exports (equation 3.2). Next we discuss the results of two additional specifications (equations 3.4 and 3.5) that takes into account the effect of RTA and EMU on exports. We report these results in Table 3.1. Finally, Table 3.2 presents the results of several robustness checks.

3.5.1 Baseline model

The results for the baseline model are in line with related literature. As expected, both the exporter and importer GDP increase exports regardless of the estimation method used. The estimated coefficients for GDP are close to one, which is the expected order of magnitude, and the distance coefficient is near to minus one. The use of country pair fixed effects does not allow identifying the effects of other gravity variables included in the model. However, pooled OLS results in Table 3.B.1 show that these coefficients are also highly significant, and proximity (either in history or in space) tends to increase exports.

3.5.2 Effect of Exchange rate

The analysis of the impact of RER level and volatility on exports (equation 3.2) show that the RER has the expected positive sign –depreciation leads to an increase in bilateral exports– and a moderate effect. In the most complete version of the model (equation 3.5), as well as in all the other specifications ran as robustness checks, RER displays a similar positive sign. In the same line, other estimation methods of equation (3.5) lead to a positive coefficient of similar magnitudes, though notably reduced when using PPML (see Table 3.B.1 in Appendix B).

The second column in Table 3.1 shows that ERV has an important detrimental effect on exports, which is significant at the 1% level. This result implies that if ERV were to rise by one standard deviation, trade would fall by about 5.79%[15]. This result is in line with those of De Nardis and Vicarelli (2003), which find a reduction of 4.04%, Gil-Pareja et al. (2008),

[15] This interpretation was first proposed by Rose (2000). It consists of reducing volatility by an amount equal to its standard deviation. In our case, the standard deviation of vol_{ij} is 0.115 and the estimate of its parameter –0.49. Hence, the, the increase in bilateral trade following a reduction of volatility from one standard deviation to zero would be given by [(exp-0.49x (VOL-0.115)/e-0.49xVOL)-1]x100, where VOL is the sample mean of the volatility. The result obtained is 5.79%.

Table 3.1. From a baseline model to an empirical model to assess EMU effect

	(1)	(2)	(3)	(4)
Dependent variable	$\ln X_{ijt}$	$\ln X_{ijt}$	$\ln X_{ijt}$	$\ln X_{ijt}$
$\ln y_i$	1.401***	1.308***	1.276***	1.282***
	(0.0488)	(0.057)	(0.058)	(0.059)
$\ln y_i$	1.277***	1.413***	1.376***	1.370***
	(0.0478)	(0.060)	(0.060)	(0.060)
$\ln rer_{ijt}$		0.051***	0.051***	0.052***
		(0.014)	(0.013)	(0.013)
vol_{ijt}		-0.492***	-0.447***	-0.447***
		(0.064)	(0.063)	(0.063)
RTAone			0.214***	0.214***
			(0.035)	(0.035)
RTAboth			0.395***	0.393***
			(0.034)	(0.035)
EMUone			-0.058*	
			(0.030)	
EMUimp				-0.086**
				(0.042)
EMUexp				-0.03
				(0.032)
EMUboth			0.165***	0.166***
			(0.045)	(0.045)
Constant	-0.179	-0.468	0.231	0.232
	(0.755)	(0.921)	(0.937)	(0.937)
R-squared	0.792	0.806	0.808	0.808

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. Robust standard errors in parenthesis. *EMUimp* takes value one when the importer belongs to the EMU and *EMUexp* when the exporter belongs to the EMU. All the regressions include country pair fixed effects, importer and exporter specific fixed effects and time effects. Control variables (Contiguity, Common Language, Same Country, Distance) were included but are dropped when using within transformation.

(1.5%), Rose (2000) (13%), or De Grauwe and Skuderlny (2000) (16,9%). Unlike Clark et al. (2004) and Gil-Pareja et al. (2008), we find that the negative relationship between volatility and trade is robust when introducing country year fixed effects. Gains from anchoring to one money are assumed to be larger when the elasticity of trade to ERV is higher. Our results confirm that there is a potential for an increase in international trade by reducing ERV. This could be an argument for some developing countries to anchor their currency on the yen, dollar or euro. To the extent that this sensibility calculated for the world average may accurately represents the sensibility of EMU exports to ERV, this allows us to calculate the effect that the elimination of the volatility of the old currencies had on EMU exports. Indeed, the volatility among the partners of the Eurozone has been of 3.12% before the adoption of the euro. Then, the elimination of the volatility would have led to an increase of 1.33% of EMU exports. This is an important impact but rather far from the optimistic Rose's proposal.

3.5.3 RTA and EMU effects

Turning to our most important variable, results in Table 3.1 confirm that EMU has a positive and significant impact on exports once we control for the presence of RTA agreements, RER level and ERV in addition to other traditional gravity variables. The panel fixed effects estimation concludes that EMU members export 17%[16] more than other countries do[17]. Additional estimations ran as robustness checks (see next section) put the effect of EMU in a rise of trade between 13% and 21% and the rise of trade due to RTA between 47% and 54%. Unlike Clark et al. (2004) and Gil-Pareja (2008), we find that the positive impact of RTA is robust when introducing country year fixed effects and the variable does not loose significance. We explain that difference by the size of our sample; it contains more country pairs with RTA that does not belong to the euroland which allows us to a better distinction among the effects of the different kind of agreements.

It is striking that we find evidence of a positive impact of the euro on trade once the elimination of volatility is controlled for. This can improbably be explained only by the elimination of transaction costs. Then, this is an additional proof of the non linear

[16]To interpret dummy coefficients as a percentage change we apply the following transformation to the coefficient obtained: 100*(EXP(α)-1].

[17] When we use the alternative estimation methods, results are more optimistic concerning RTA (around 49%) and more pessimistic concerning EMU (15%). Thus, we are able to conclude that there is a positive impact of EMU on trade once the elimination of the volatility is controlled for and the presence of RTA is taken into account.

relationship between ERV and trade. Baldwin et al. (2005 and 2008) offers an attractive explanation for the fact that the elimination of hedging costs associated with ERV before the introduction of the euro translates in an extensive increase of trade. Unfortunately, the full verification of this fact would imply to rely on a measure of the intensive and extensive margin of trade, and this issue is beyond the goal of this study.

Trade agreements may also have a diversion effect on trade with third countries by reducing imports since it artificially reduces the price of imports coming from members. Though, the impact of exports is less clear. We find that belonging to a RTA also has a positive and significant impact on exports to non-members countries. Turning to the effect of EMU on trade with non-members, we find evidence of diversion effect; EMU seems to have reduced imports from third countries. On the contrary, the common currency has not had a significant effect on EMU exports.

3.6. Additional results

We have re-estimated the third specification in Table 3.1 introducing some changes to check the robustness of the results obtained and to bring new elements to the discussion; results are displayed in Table 3.2.

3.6.1. Volatility impact on trade. Before versus after the euro

The introduction of the euro not only limits ERV when one of the partners uses the euro but also affects volatility among third countries. Then, it could be the case that the introduction of the euro changes the overall sensibility of the traders to ERV. To test this hypothesis, we have introduced a new set of dummies. VOL67-98 reports the value of volatility for the period 1967-1998 and zero otherwise; and VOL99-09 reports the value of volatility for the period 1999-2009. Results reported in column 2 show that the effect of volatility was significantly detrimental to trade before the introduction of the euro but does not have a significant impact after the introduction of the common currency. This is an important result since it points out important collateral effect of the new currency.

3.6.2. Volatility impact per period. Euroland versus non euroland

Of course, the coordination among EU members is not a story that started in 1999. The common policy concerning politics, social, trade and numerous norms of convergence had previously reinforced all the links among members and made transactions among their members less risky. Then, volatility could affect in a different way future members of the EMU than other countries. The ERV can also be disentangled into the ERV of future EMU members and non-EMU countries. In column 3 we have alternatively split each of the volatility variables into three variables. We interact *VOL67-98* and *VOL99-09* with three different dummies depending if one, both or none of the countries belong to EMU. The volatility is detrimental in all cases, though non-significant for exports to non EMU countries in the post-euro period.

3.6.3. The joint effect of RTA and EMU

Most of EMU countries were already members of the same RTA for a long period. Then, the dummy EMU could be capturing not only the financial and monetary integration of the euroland but some progressive deepening in the integration of the common market. In sum, the EMU coefficient may be overestimated since RTA captures the average effect of very different trade agreements while the EU represents the most integrated region. To be sure this is not the case; in column 4 we replace the variables RTA and EMU by two dummies: RTAxEMU that takes the value one when both trade partners joined the EMU for all the period and RTAxnoEMU when trade partners are members of a RTA (the EU or another agreement) but did not join the EMU. The results confirm that exports of EMU countries involved in a RTA are larger than exports of countries involved in another type of RTA. Other RTA also have a positive effect for exports of their members, similar to the one obtained in previous estimations.

Table 3.2. Robustness checks

	Vol per period	**Vol per period and EMU**	**RTAxEMU**
Dependent variable	$\ln X_{ijt}$	$\ln X_{ijt}$	$\ln X_{ijt}$
$\ln rer_{ijt}$	0.052***	0.052***	0.052***
	(0.013)	(0.013)	(0.013)
vol_{ijt}			-0.458***
			(0.063)
Vol 67-98, No EMU		-1.549***	
		(0.171)	
Vol 67-98, EMUone		-0.929***	
		(0.205)	
Vol 67-98, EMUboth		-2.681***	
		(0.956)	
Vol 99-09, No EMU		0.161	
		(0.541)	
Vol 99-09, EMUone		-0.045	
		(0.068)	
Vol 67-98	-1.401***		
	(0.141)		
Vol 99-09	-0.023		
	(0.058)		
RTAone	0.206***	0.203***	0.187***
	(0.035)	(0.035)	(0.037)
RTAboth	0.388***	0.389***	
	(0.034)	(0.034)	
EMUone	-0.078*	-0.052	-0.107***
	(0.042)	(0.043)	(0.040)
EMUimp	-0.021	0.004	-0.035
	(0.033)	(0.034)	(0.032)
EMUexp	0.198***	0.180***	-0.035
	(0.046)	(0.046)	(0.032)
RTAxEMU			0.122**
			(0.054)
RTAxnoEMU			0.316***
			(0.037)
Constant	0.154	0.135	0.328
	(0.935)	(0.935)	(0.924)
R-squared	0.808	0.824	0.823

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. Robust standard errors in parenthesis . Control variables (Exporter and Importer GDP, Contiguity, Common Language, Same Country, Distance) are included but not reported for the sake of clarity. All specifications include country pair, exporter, importer and time effects.

3.6.4. Euro impact on Eurozone trade with other EU members

In Table 3.3 we analyze the effects of the euro on EMU trade with other European countries not belonging to the EMU. The first row reports the euro effect on trade with Denmark, Sweden and UK. Though positive, the effect is non-significant for exports and imports from these countries. In the second row we provide evidence of the euro effect when trading with CEE countries[18]. Our result are in line with Cieslik et al. (2009), revealing a positive and statistically significant coefficient of this variable, which indicates that it has strongly contributed to the expansion of some CEE countries exports.

Table 3.3. The impact of EMU on other EU countries

	Exports	**Imports**
Dependent variable	$\ln X_{ijt}$	$\ln X_{ijt}$
EMU - EU15	0.02	0.032
	(0.067)	(0.058)
EMU -EU25	0.308^{***}	0.311^{***}
	(0.057)	(0.072)

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. Robust standard errors in parenthesis . Control variables (Exporter and Importer GDP, Contiguity, Common Language, Same Country, Distance) are included but not reported for the sake of clarity. All specifications include country pair, exporter, importer and time effects.

3.6.5. Euro impact on each EMU country

The analysis for individual EMU members reveals the existence of a good deal of variation in the effect of the euro across member countries. In related literature, this impact is found to be particularly high for Spain; see, for example, Gil-Pareja et al. (2005), Baldwin and Di Nino (2006) and Baldwin et al. (2008) or Micco et al. (2003).

Table 3.4 shows in the first column the impact of EMU on each country when trading

[18] Bulgaria, Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Romania, Slovakia, Slovenia.

with other EMU countries. The second (third) column shows this impact over exports to (imports from) third countries. Our results show a positive and large coefficient for Spain when trading with other EMU countries, followed by Portugal, Italy, Belgium and Luxembourg. This effect is found to be negative and significant for Ireland. Turning to trade with third countries, it is shown that Germany and Ireland have reduced both their exports to and their imports from third countries following the introduction of the euro, whereas for Spain the effect has been positive in both cases.

Table 3.4. EMU effect by country

	EMUboth	**EMUexp**	**EMUimp**
Austria	0.123	0.024	-0.189
	(0.081)	(0.086)	(0.116)
BL	0.322***	0.003	0.04
	(0.071)	(0.064)	(0.115)
Finland	-0.022	0.074	-0.409**
	(0.068)	(0.071)	(0.163)
France	0.098	-0.232***	-0.137
	(0.074)	(0.069)	(0.087)
Germany	0.032	-0.309***	-0.309***
	(0.067)	(0.071)	(0.098)
Greece	-0.036	0.079	0.039
	(0.099)	(0.109)	(0.134)
Ireland	-0.358***	-0.512***	-0.419***
	(0.073)	(0.096)	(0.110)
Italy	0.275***	-0.023	-0.048
	(0.071)	(0.064)	(0.085)
Netherlands	0.045	-0.046	-0.02
	(0.065)	(0.064)	(0.088)
Portugal	0.332***	0.215**	0.189
	(0.112)	(0.109)	(0.133)
Spain	0.660***	0.130*	0.271***
	(0.087)	(0.072)	(0.091)

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. Robust standard errors in parenthesis . Control variables (Exporter and Importer GDP, Contiguity, Common Language, Same Country, Distance) are included but not reported for the sake of clarity. BL stands for Belgium and Luxembourg. All specifications include country pair, exporter, importer and time effects.

Hence, it is shown that there are important differences across countries regarding the impact of EMU on trade, and the exercise performed in this section allows us to better understand the aggregate coefficients displayed in Table 3.1. However, the fact that only four over twelve euroland members have benefited from an 'extra' gain in terms of trade after controlling for the gains obtained from the elimination of volatility may be explained by conditions shared by all four or by the fact that the effect year by year could differ from one country to another.

3.6.6. Euro impact over time

To study the euro effect on trade more precisely, we estimate the single currency's year-by-year impact. Table 3.5 presents the results on the evolution of EMU effects over time. It can be appreciated that the euro significantly boosted intra-EMU trade starting in 1999, with the effect reaching its maximum in the 2003-2005 period. Micco et al. (2003) perform a similar exercise and their results point in the same direction. However, their sample only includes the 1992-2002 period.

Table 3.5. EMU impact over time

	EMUboth
Dependent variable	$\ln X_{ijt}$
1999	0.303***
	(0.046)
2000	0.192***
	(0.047)
2001	0.126***
	(0.048)
2002	0.174***
	(0.049)
2003	0.262***
	(0.049)
2004	0.235***
	(0.049)
2005	0.205***
	(0.050)
2006	0.131***
	(0.051)
2007	0.108**
	(0.054)
2008	0.034
	(0.054)
2009	0.155***
	(0.055)

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. Robust standard errors in parenthesis. Control variables (Exporter and Importer GDP, Contiguity, Common Language, Same Country, Distance) are included but not reported for the sake of clarity. All specifications include country pair, exporter, importer and time effects.

3.7. Conclusions

We have estimated a gravity equation including the RER level and volatility, the existence of trade agreements and the introduction of the euro for a large sample of 80 countries and a long period 1967-2009. As far as we know, this is the more complete gravity equation and larger period used to this end in the literature. We use panel and Poisson estimators with country, country pair and year fixed effects to estimate our empirical model. We find that panel techniques with country and year fixed effects minimizes the error loss compared to alternative estimations procedures. Compared to other studies, we use a more heterogeneous sample and longer period which allow us to confirm some previous results and add new ones. We confirm that the common currency has had a positive impact on EMU exports to other EMU countries, though it has reduced imports from third countries to the Eurozone. In addition, we find that the elimination of the ERV boosted export per se. In particular our study highlights the detrimental effect of ERV on exports. This is not a specificity of the Eurozone since our estimations are based on a large sample and a long period. Then, the possibility to peg to the euro could boost the trade of third countries and between these third countries. On the other hand, EMU countries have clearly loose the possibility to adjust with their ER. Thus one should be cautious when comparing the benefits of the EMU with the gains the countries could obtained from depreciation, this is an important element to take into account.

Besides that, our sensitivity analysis on the volatility impact on trade shows that it was significantly detrimental to trade before the introduction of the euro but it does not have a significant impact after the introduction of the common currency. In addition, we show that exports of EMU countries involved in a RTA are larger than exports of countries involved in another type of RTA.

Finally, the analysis of the euro effect for individual EMU members reveals the existence of a good deal of variation in the effect of the euro across member countries, showing the higher coefficients for Spain, Italy, Portugal, Belgium and Luxembourg. This effect is found to be negative and significant for Ireland. Moreover, we provide evidence that EMU has contributed to the expansion of some CEE countries exports.

This work can be extended in several directions. First, further research is needed to explain the differences in the euro effect across EMU members. Comparative analysis of the impact of ER on the extensive and intensive margin of trade for several countries could shed more light on the European process, allowing us to have a better understanding of its benefits and disadvantages. Secondly, it would be interesting to analyze the properties of the time series of the panel. The presence of nonstationarities and structural breaks in the data is rather probable given the long-run period under analysis and the use of more adequate estimators could improve the conclusions obtained.

Appendix A

Table 3.A.1. EMU effect estimation in the literature

Article	Data	Estimation method	RTAs/EMU dummies	ER and ERV
Baldwin et al. (2005)	1991-2002; 15 EU countries, Australia, Canada, Norway, Japan and USA	- Panel fixed effects	- No RTA dummies - EMU dummies	- No ER level - ERV: Annual variance of the weekly nominal ER return
Barr et al (2003)	1978- 2002; 17 countries	- Panel fixed effects - IV estimation	- No RTA dummies - EU and EMU dummies	- No ER level - ERV: standard deviation of monthly logarithm changes in bilateral ER for the year prior to observation date
Belke and Spies (2008)	1991-2004; OECD members and Romania, Bulgaria, Estonia, Latvia, Lithuania and Slovenia	- Pooled OLS - Panel fixed effects - Fixed Effects Vector Decomposition - Hausman-Taylor	- No RTA dummies - EMU, EU and EFTA dummies	- RER level - No ERV
Brouwer et al. (2008)	1990-2004; 29 countries	Panel fixed effect	- No RTA dummies - EMU and EU dummies	- Nominal ER (NER) level - ERV: standard deviation of the monthly percentage changes in the RER within a year

Table 3.A.1. EMU effect estimation in the literature

Article	Data	Estimation method	RTAs/EMU dummies	ER and ERV
Bun and Klaasen (2007)	1967-2002; 19 countries	- Dynamic OLS	- RTA dummies - Euro dummy	- No ER level - No ERV
Bussière et al. (2008)	1980-2003; 61 countries	- OLS - Panel fixed and random effects - Dynamic OLS	- RTA dummies - EU dummy	- RER level - No ERV
Cieslik et al. (2009)	1993-2007; OECD and CEE countries	- Panel fixed and random effects - Hausman-Taylor	- No RTA dummies - EMU and EU dummies	- No ER level - ERV: standard deviation of the first difference of natural logarithm of the bilateral RER
De Grauwe and Skuderlny (2000)	1962-1995; 14 EU members (Greece excluded)	- Panel fixed effects	- No RTA dummies - EMU dummy	- RER level - ERV: variance of the monthly nominal ER return between the currencies of countries *i* and *j* in year *t*
De Nardis and Vicarelli (2003)	1980-2000; 11 EU exporter countries; 30 importer countries	- Dynamic OLS	- RTA dummies - Euro dummy	- No ER level - ERV: standard deviation of the first difference of the natural logarithm of the bilateral NER
De Nardis et al. (2008)	1988-2004; 23 developed countries	- Panel OLS - System GMM	- No RTA dummies - EMU and EU dummies	- No ER level - ERV: standard deviation of the first difference of the monthly natural logarithm of the bilateral NER
Flam and Nostrom (2006)	1999-2005; 20 OECD countries	- Panel fixed effects	- No RTA dummies - Euro dummies	- RER level and trade-weighted RER level - No ERV (dropped because an insignificant effect was found)

Table 3.A.1. EMU effect estimation in the literature

Article	Data	Estimation method	RTAs/EMU dummies	ER and ERV
Gil-Pareja et al. (2008)	1950-2004; 24 OECD countries	- Panel fixed effects	- RTA dummies - EMU dummies	- No ER level - ERV: variance of the first difference on the monthly natural logarithm of the bilateral NER
Maliszewska (2004)	1992-2002; 22 developed countries	- Panel fixed effects	- RTAs dummies - EMU and EU dummies	- No ER level - No ERV
Micco et al. (2003)	1992-2002; 22 developed countries	- Panel fixed effects	- RTA dummies - EMU dummies	- NER level - No ERV
Rose (2000)	1970, 1975, 1980, 1985, 1990; 186 countries	- OLS	- CU dummy	- ERV: standard deviation of the first difference of the monthly natural logarithm of the bilateral NER
Spies and Marques (2009)	1991-2003; 204 countries	- Panel fixed effects - Fixed Effects Vector Decomposition	- RTA dummies - EU dummies	- RER level - No ERV

Appendix B

Table 3.B.1. Alternative estimations methods

	Pooled OLS	**Panel fixed effects**	**Poisson fixed effects**
Dependent variable	$\ln X_{ijt}$	$\ln X_{ijt}$	X_{ijt}
$\ln y_i$	1.208***	1.282***	0.981***
	(0.027)	(0.059)	(6.60e-10)
$\ln y_i$	1.175***	1.370***	0.867***
	(0.027)	(0.060)	(6.94e-10)
$contig_{ijt}$	0.050*		
	(0.030)		
$comla_{ij}$	0.634***		
	(0.049)		
$smctry_{ij}$	-1.329***		
	(0.007)		
$\ln d_{ij}$	0.050*		
	(0.030)		
$\ln rer_{ijt}$	0.038***	0.052***	0.0255***
	(0.010)	(0.013)	(3.65e-10)
vol_{ijt}	-0.362***	-0.447***	-0.573***
	(0.076)	(0.063)	(3.33e-09)
RTAone	0.410***	0.214***	-0.134***
	(0.018)	(0.035)	(4.02e-10)
RTAboth	0.126***	0.393***	0.403***
	(0.016)	(0.035)	(3.71e-10)
EMUone	0.380***	-0.086**	0.127***
	(0.022)	(0.042)	(5.09e-10)
EMUimp	0.447***	-0.03	0.167***
	(0.018)	(0.032)	(4.99e-10)
EMUexp	-0.042	0.166***	0.146***
	(0.033)	(0.045)	(4.76e-10)
Constant	10.340***	0.232	
	(0.336)	(0.937)	
R-squared	0.842	0.808	

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. Robust standard errors in parenthesis. EMUimp takes value one when the importer belongs to the EMU and EMUexp when the exporter belongs to the EMU. All the regressions include country pair, importer, exporter and time effects.

Table 3.B.2. Goodness of fit. Equation (3.5)

	Bias	**MSE**	**Absolute error loss**
Pooled OLS	**1.28E-09**	**2.366**	**1.095**
Panel fixed effects	3.54E-09	5.170	1.725
Poisson fixed effects	-0.672	7.996	2.802

Notes: Bold values indicate the preferred estimation method in each case.

Appendix C

3.C.1. Panel techniques

In contrast to cross-section regressions, a panel framework allows to recognize how the relevant variables evolve through time and to identify the specific time or country effects (institutional, economical, cultural time-invariant or population-invariant factors). During the last years, researchers have turned towards panel data (cross-section gravity models for several consecutive years). Egger (2000); Rose and van Wincoop (2001); Mátyás (1998); Egger and Pfaffermayr (2003, 2004); Glick and Rose (2002); Brun et al. (2002) and Melitz (2007) constitute some examples.

Fixed effect models assume that the unobserved heterogeneous component in the regression is constant over time and include dummy variables for importer and exporter in the regression to control for this heterogeneity. According to Baltagi et al. (2003) fixed effects can be classified into two groups; main and interaction effects. The first term makes reference to the usual fixed exporter, importer and time effects (η_i, η_j, η_t), whereas the second refers to controls for country pair fixed effects (η_{ij}) and exporter and importer specific time-varying effects (η_{it}, η_{jt}).

Nevertheless, the introduction of country or country pair dummies implies high computational costs. In addition, any explanatory variable constant across time in each country (or pair of countries) will be perfectly collinear with the fixed effects, and dropped from the model. Hence, the impact of some interesting variables such as land area, common language or common borders cannot be measured.

Some authors have opted to assume that the unobserved component of the regression is distributed randomly. Random effect models impose zero correlation between the individual effects and the regressors, implicitly assuming that the unobserved heterogeneous component is strictly exogenous. Under the null hypothesis of zero correlation, random effect model is more efficient. However, if the null is rejected, only fixed effect model provides consistent estimators.

3.C.2. Poisson Pseudo Maximum Likelihood (PPML)

The log-linearization of the gravity equation changes the property of the error term and thus conduces to inefficient estimations in the presence of heteroskedasticity. If the data are homoskedastic, the variance and the expected value of the error term are constant, but if not (as usual in trade data) the expected value of the error term is a function of the regressors. The conditional distribution of the dependent variable is then altered and OLS estimation is inefficient. Heteroskedasticity does not affect the parameter estimates; the coefficients should still be unbiased, but it biases the variance of the estimated parameters and, consequently, the t-values cannot be trusted. The source of heteroskedasticity is not unique; the variance of the error term may vary with the regressors, with the dependent variable or with some other variable that has been omitted. For instance, Santos Silva and Tenreyro (2006) argue that the variance of the error term is correlated with the countries' GDP and with the measure of distance. Nonlinear estimation methods have been recently suggested in the literature to tackle this issue. Among these techniques, the most frequently used are Nonlinear Least Squares, Feasible Generalized Least Squares, Heckman sample selection model and Gamma and Poisson Pseudo Maximum Likelihood.

Chapter 4

Is there a ‘euro effect’ on trade? New evidence using panel cointegration techniques

Abstract. In this chapter we present new evidence on the effect of the euro on trade. We use a dataset containing all bilateral combinations in a panel of 26 OECD countries covering the period 1967-2008. We estimate the equation using two sets of variables: first, one defined as it is standard in the gravity equation literature, and a second one built according to the criticisms stated by BT. From a methodological point of view, we apply panel tests that account for the presence of cross-section dependence as well as discontinuities in the non-stationary panel data. We test for cointegration between the variables using the test proposed by Banerjee and Carrion-i-Silvestre (2010). We also efficiently estimate the long-run relationships using the CUP-BC and CUP-FM estimators proposed in Bai et al. (2009). Our results challenge earlier estimates using standard panel data techniques and are in line with those of Bun and Klaassen (2007). We argue that, after controlling for cross-section dependence and deterministic trends and breaks in trade integration, the euro generates lower trade effects than predicted in previous studies.

4. 1. Introduction

The introduction of the euro has raised a new interest in measuring the impact of currency unions (CU) on trade flows. The very high estimates of trade induced by the creation of monetary unions found in the seminal papers by Rose (2000) and Frankel and Rose (2002) has led to the concept of 'endogeneity' of Optimum Currency Areas (OCA) that means for the euro area that, even if the EMU was not created as an OCA, it could be progressing in that direction (Frankel and Rose, 1998). Recent research surveyed by Rose and Stanley (2005) and Rose (2008) suggests that the introduction of the euro still has a sizable and statistically significant effect on trade among EMU members. Taking together all these estimates imply that the EMU has increased trade by about 8%-23% percent in its first years of existence. This issue can be very relevant for prospective new members of the EMU.

In 1999 eleven countries of the EU adopted the euro as a common currency while Greece entered in 2001. Since then, also Slovenia, Cyprus, Malta, Slovakia and Estonia have joined the euro area while other members of the EU are 'waiting and seeing', the so-called derogation countries. Moreover, the introduction of the euro was preceded by other stages of economic integration (Customs Union, European Monetary System and the Single Market), so EMU effect has to be analyzed as an on-going process with a time dimension. It might be interesting to investigate whether there is an additional benefit of a common currency over (relative) ER stability. As pointed out by Faruqee (2004) the central questions at stake are the following: first, to ascertain the effects of EMU on the area's trade flows; second, to analyze the evolution of trade effects over time, and finally, to measure the distribution of trade effects among member states.

In this chapter we have tried to overcome some of the main flaws found in the standard empirical literature and recently outlined by Eicher and Henn (2011). First, BT's critiques regarding the proper specification of gravity models in large panels to prevent omitted variable bias point out the need to simultaneously account for multilateral resistance and unobserved bilateral heterogeneity. We have accounted for BT's critiques in the specification of the model as well as in the definition of the variables included in the estimation of the gravity model.

Second, more recently, Fidrmuc (2009) and Bun and Klaasen (2007) have outlined the importance of considering the possible non-stationary nature of the variables included in the gravity equation, as well as the cross-sectional correlation between the elements (countries) of the panel, both aspects normally neglected in the empirical applications. While initially the literature overlooked some crucial econometric issues regarding non-stationary series in panel estimation, more recent works have taken into account these aspects using non-stationary panel data techniques. A sizeable literature has been developing along these lines, but none of these works explicitly deals with the issue of cross-section dependence with the exception of Gengenbach (2009).

Third, Bun and Klaasen (2007) have stated that models measuring the effect of the euro on trade have omitted some variables, causing an upward bias in the trade benefits earlier estimated. They find that the longer the data period considered, the higher the euro effect estimate. Thus this might be due to some misspecification of the time-series characteristics of the variables involved, namely the trends in trade flows over time. To correct for this bias they add a time-trend to their specification and allow it to have heterogeneous coefficients across country pairs. Then they estimate long-run relationships using first-generation panel cointegration techniques, that is, without considering dependence in the cross-section dimension.

Therefore, in this chapter we try to provide new evidence on the effect of the euro using a dataset that contains information on all bilateral combinations in a panel of 26 countries covering the period 1967-2008. We implement a new generation of tests that allows us to solve some of the problems derived from the non-stationary nature of the data used in gravitational equations. More specifically, we implement the panel unit root and stationary tests proposed by Pesaran (2004, 2007) and Bai and Ng (2004) to test for the presence of cross-section dependence as well as discontinuities in the non-stationary series. We then test for cointegration between the variables using panel cointegration tests, with a special emphasis in the one proposed by Banerjee and Carrion-i-Silvestre (2010). Finally, we apply the Bai et al. (2009) CUP estimator to efficiently estimate the regression coefficients. The results obtained are in line with Bun and Klaassen (2007) confirming a smaller euro effect than in other research papers, like for instance, Gil-Pareja et al. (2008), where cross-section dependence and the non-stationary nature of the variables are not accounted for.

The chapter is organized as follows. Section 4.2 discusses the empirical literature on CU and trade, emphasizing the econometric approaches based on the gravity model. Section 4.3 presents a new econometric approach that overcomes some of the problems present in the current literature. Section 4.4 describes the data and discusses the empirical results. A final section concludes.

4.2. Previous studies and criticisms to the empirical application of the gravity equation for measuring the euro effect on trade

The literature examining the impact of CU on trade is a burgeoning field of research. All in all, the diversity of existing estimates indicates the potential bias inherent in applied specifications. Although in the beginning the gravity model was criticized for its lack of theoretical underpinnings, now rests on a solid theoretical background[19]. Therefore, as stated in Westerlund and Wilhelmsson (2009) the focus of this line of research has shifted from its theoretical soundness towards the estimation techniques used.

The econometric approach has changed over time as a result of a feedback process between theory and empirics. In this abundant literature, the traditional approach has been to use cross-section data. However, it is generally accepted that the results obtained were suffering from a bias, as the heterogeneity among countries was not properly controlled for. Thus, Rose's (2000) initial estimates in a cross-sectional study suggested a tripling of trade. This result was quite striking, and as quoted by Faruqee (2004), is at odds with the related literature that typically finds very little negative impact of ERV on trade. Not surprisingly, Rose's findings have received substantial revisions, and subsequent analysis generally finds a smaller (albeit still sizable) effect of CU membership on trade. There are different reasons that make the implication of Rose (2000) work unclear. First, the sample countries were mostly smaller and poorer, not including the EMU ones. This has led to question whether the results apply to bigger countries such as EMU members. Second, the cross-sectional analysis included in Rose (2000) provides a comparative benchmark across members of a monetary union against third countries but the most relevant issue about EMU is the possible change in the level of trade for its member over time, before and after the introduction of the single currency. In order to solve this problem, a second string of literature started to use panel data estimation techniques, which permits more general types of heterogeneity[20]. However, BT define what they call in this context 'the gold medal error', also known as the 'Anderson-van

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[19] See, for instance, Feenstra et al. (2001).

[20] Moreover, as clearly explained by Westerlund and Wilhelmsson (2009), if we desire to measure the impact of a currency union on exports (which is the relevant case in this chapter), while simultaneously controlling for country pair propensity to trade, it is easier under a panel data framework by means of a country pair fixed effect term. For a single cross-section, these controls can only depend on observed country pair attributes such as common language, and estimates can thus be biased if there is additionally an unobserved component to the country pair propensity to trade.

Wincoop misinterpretation' in the sense that AvW developed a cross-section estimation technique to control for omitted variables with pair fixed effects[21]. However, this technique has been generalized to the panel data framework by many authors without considering the time dimension (see, for example, Glick and Rose, 2002 or Flam and Nordstrom, 2006). Country dummies (for exporters and importers) only remove the average impact leaving the time dimension in the residuals, which leads to biased results. Therefore, time-invariant country dummies are not enough and a proper treatment of the time dimension is needed. Moreover, BT also stress the importance of an omitted variable bias when the empirical specification does not account for unobserved determinants of bilateral trading relationships. They suggest the inclusion of time varying fixed effects in the specification. However, if doing so, we would not be able to explore cointegration between GDP and exports, since the time varying fixed effects would absorb GDP. Instead, and following Bun and Klaasen (2007), we include in our specification a country pair specific time trend which captures all the unobserved heterogeneity through time, as well as country specific fixed effects. Furthermore, the application of cointegration techniques implies the proper treatment of the time dimension, since it takes into account the long-run relationships among variables.

Besides the above-mentioned specification caveats, BT pointed out two additional minor problems, coined as 'silver' and 'bronze' medal errors. The silver medal error concerns the definition of the dependent variable. As BT point out, the gravity equation is an expenditure function that explains the value of spending by one nation on the goods produced by another nation; it explains uni-directional bilateral trade. Most gravity models, however, work with the average of the two way exports and frequently the averaging procedure is wrong. The problem arises when authors use the logarithm of the sum instead of the sum of the logarithms in the bilateral trade term. The silver medal mistake will create no bias if bilateral

[21] See Anderson and van Wincoop (2003).

trade is balanced. However, if nations in a currency union tend to have larger than usual bilateral imbalances, as it has been the case in the Eurozone, the silver medal misspecification leads to an upward bias as the log of the sum (wrong procedure) overestimates the sum of the log (correct procedure). Finally, the bronze medal mistake concerns the price deflator: all the prices in the gravity equation are measured in terms of a common numeraire, so there is no price illusion. However, many authors deflate trade flows and GDP using the US CPI (following Rose's example). In this article we include exports as dependent variable and we define all the variables in nominal terms to avoid silver and bronze errors.

Concerning the estimation problems, Santos-Silva and Tenreyro (2006) argue that the standard empirical methods used to estimate the gravity equation (i.e. OLS) are inappropriate. Glick and Rose (2002) and Frankel and Rose (2002) exploited the time series information using panel data, giving birth to a literature in search of 'more reasonable' effects (Eicher and Henn, 2011). Micco et al. (2003) examined the dynamic impact of EMU on trade for 22 industrial countries using panel regressions based on a gravity model. Their findings suggest that EMU has fostered bilateral trade between 8% and 16% depending on the EMU membership of the countries and that the positive effect has been rising over time. Other studies, like Bun and Klaasen (2002) estimate a dynamic panel data model and distinguish between short (3.9%) and long-run effects (38%). Rose and Stanley (2005) perform a meta analysis of the results of 34 studies, and find a combined estimate of the trade effect between 30% and 90%, which is smaller than previous evidence. However, these papers generally use smaller and shorter datasets than Rose's. When they focus on large panels, they find bigger estimates (over 100%). Therefore, the empirical literature is far from conclusive and we can infer that dataset dimensions, and, especially, econometric approaches, influence the results.

While the heterogeneity bias is controlled through the use of fixed-effects, a second kind of misspecification is related to dynamics. The recent theoretical literature on international trade with heterogeneous firms (Bernard et al., 2003; Melitz, 2003; Helpman et al., 2004) has been largely based on evidence that, in a sector, the behaviour of firms can be highly heterogeneous, both concerning their productivity and their involvement in international transactions. In particular, the existence of sunk costs borne by exporters to set up distribution and service networks in the partner country may generate inertia in bilateral trade flows, especially among EMU countries, where there is also accumulation of invisible assets such as political, cultural and geographical factors characterizing the area and influencing the commercial transactions taking place within it.

Bun and Klaasen (2007) constitutes a path-breaking study in this respect. They show that the residuals of the Least Squares Dummy Variables estimator (LSDV) exhibit trends over time. Therefore, they estimate the gravity equation allowing for country pair specific time trends to account for the observed trending behaviour in the residuals. Moreover, they analyze the non-stationary nature of the data as well as the cointegration relationships and obtain a much smaller estimate of the euro effect (3%) on bilateral trade[22]. However, they employ methods that assume cross-section independence, and this is an assumption unlikely to hold in bilateral trade data. As recently stated by Fidrmuc (2009), cross-correlation is likely to be present in gravity models because foreign trade is strongly influenced by the global economic shocks (i.e. other economies business cycles). Moreover, dependence is generated by construction as gravity models include bilateral trade flows together with aggregate national variables. Furthermore, the gravity model itself implies spatial dependence in the data due to the hypothesized effect of distance on trade. Several new panel

[22] Faruqee (2004) and Fidrmuc (2009) are other papers that stress the importance of the non-stationary nature of the series and that apply cointegration techniques.

unit root and cointegration tests have been proposed accounting for cross-sectional dependence in the form of common factors[23].

More recent studies have insisted on the importance of accounting for the existence of trends in the data and its possible non-stationary nature. Historically, researchers have assumed stationary time series to estimate gravity models. However, if the variables are non-stationary, a different statistical setup needs to be used. As Faruqee (2004) claimed, estimating the impact of a monetary union on trade faces several econometric challenges. Recent literature shows that the results of the gravity models are sensitive to their proper specification (Egger and Pfaffermayr, 2003). However, properly specified models in panel data may have some caveats when data are non-stationary. If the non-stationary nature of the series is not considered, spurious regressions may appear. Although the spurious correlation problem is less important in panels than in time series analysis, as the fixed effects estimator for non-stationary data is asymptotically normal (see Kao and Chiang, 2000), the results are biased. Correspondingly, panel cointegration techniques are used accounting for different possible estimation problems (endogeneity, cross-correlation or breaks). Therefore, a sound empirical strategy must proceed as follows: first, to determine the order of integration of the variables through panel unit root tests; second, to test for cointegration among the integrated variables using panel cointegration tests; finally, to use the panel cointegration estimators to provide reliable point estimates.

The contribution of this chapter to the existing literature about the euro effect on trade is twofold. First, unlike previous research, (excepting Eicher and Henn, 2011) we address BT's critiques regarding the proper specification of gravity models and the definition of the variables, as we account for multilateral resistance, as well as unobserved bilateral

[23] See for example Breitung and Pesaran (2008) for an overview of the literature and Gengenbach et al. (2010) for a comparison of panel unit root tests.

heterogeneity. Second, we apply an econometric methodology comprising of a range of techniques to test and estimate efficiently in a non-stationary panel framework, solving endogeneity problems as well as possible biases posed by structural breaks, country pair specific trends and cross-section dependence.

4.3. Data, methodology and empirical results

Bun and Klaasen (2007) showed the importance of a correct specification of the gravity model including not only deterministic trend components but also stochastic trends derived from the non-stationary nature of the macro-variables involved. However, some practical problems implied that most of the evidence obtained so far did not considered nonstationarity. New developments in macroeconometrics have been recently extended to the panel framework allowing addressing most of the issues concerning both specification and estimation discussed in the previous section.

A first common problem in the context of panel non-stationary variables is that some widely used tests assume the absence of correlation across the cross-sections of the panel. That is, the individual members of the panel (countries) are considered independent. This assumption is not realistic and, therefore, cannot be maintained in the majority of the cases, especially when the countries are neighbours or are involved in integration processes. More recently, a second generation of panel tests, in contrast, introduce different forms of dependence, solving the above-mentioned problem.

Although there are several alternative proposals formulated in the literature to overcome the cross-section dependence problem, when dependence is pervasive –as in economic integrated areas- the best alternative is the use of factor models. This consists of assuming that the process is driven by a group of common factors, so it is possible to distinguish between the idiosyncratic component and the common component.

In the case of panel unit roots, several tests have been formulated based on factor models[24]. In particular, Bai and Ng (2004) account for the non-stationarity of the series coming either from the common factors, the idiosyncratic component or from both. Moreover, they consider the possible existence of multiple common factors as well as the existence of cointegration relationships among the series of the panel. Banerjee et al. (2004) stated that there is a tendency to over-reject the null of stationarity when cointegration is present. As the existence of cointegrating relations between trade series is a very plausible hypothesis in economic integrated areas, the proposal in Bai and Ng (2004) is the best approach in our case[25]. For the sake of comparison, we will also present the results obtained using Pesaran's (2007) approach. Similarly, we will also allow for dependence in the estimation of the cointegration relationships using the common factor approach of Bai and Ng (2004).

A second caveat appears when there are structural breaks in the time dimension of the panel. If there exist linear combinations of integrated variables that cancel out their common stochastic trends then, these series are said to be cointegrated. The economic translation is that these series share an equilibrium relationship. However, a commonly neglected phenomenon is that both the cointegrating vector and the deterministic components might change during the period analyzed, and if we do not take account of these structural breaks in the parameters of the model, inference concerning the presence of cointegration can be

[24] Namely, Pesaran (2007), Phillips and Sul (2003), Moon and Perron (2004) and Bai and Ng (2004).
[25] Moreover, using Monte Carlo methods, Gengenbach et al. (2010) and Jang and Shin (2005) show that, for all the specifications considered in their simulation experiments, the test in Bai and Ng (2006) has more power than those by Moon and Perron (2004) and Pesaran (2007), and better empirical size than that of Phillips and Sul (2003).

affected by misspecification errors. Therefore, in this chapter we propose the use of the tests developed in Banerjee and Carrion-i-Silvestre (2010). They generalize the approach in Pedroni (1999, 2004) to account for one structural break that may affect the long-run relationship in a number of different ways (cointegrating vector and/or deterministic components). Moreover, they address the cross-section dependence issue by using the above-mentioned factor model approach due to Bai and Ng (2004) to generalize the degree of permissible cross-section dependence allowing for idiosyncratic responses to multiple common factors.

To sum up, we control for econometric issues usually neglected in earlier literature: first, we account for cross-section dependence among countries in the panel tests, both unit roots and cointegration. Second, we allow for the existence of a break in the deterministic components (constant, trend and cointegrating vector) of the model as well as in the cointegration relationship, a major point to assess the effect of institutional changes in the relationship. Furthermore, since the trend included in the specification is country pair specific, the break in the trend is also allowed to have different coefficients for each country pair, therefore allowing for a higher degree of heterogeneity in the estimation. To the best of our knowledge, this is the first time that structural changes have been considered in the euro effect literature based on gravity equations. Finally, the estimation of the long-run relationship uses a methodology that not only efficiently estimates the coefficients but also is based on the common factors decomposition that assures a homogeneous econometric approach. We choose, for this purpose, the Bai et al. (2009) CUP-FM and CUP-BC estimators.

4.3.1. Data

The countries included in the study are Australia, Austria, Belgium and Luxembourg (as an unique area), Canada, Chile, Denmark, Finland, France, Germany, Greece, Iceland, Ireland, Italy, Japan, South Korea, Mexico, the Netherlands, New Zealand, Norway, Poland, Portugal, Spain, Sweden, Switzerland, United Kingdom and United States.

The dataset contains annual data from these 26 OECD countries and covers the period 1967-2008. Although the number of years available was higher, we have opted by restrict our sample to this period to exclude the effects of the financial crisis that started in 2008. Hence, we have a balanced panel with dimension N=650 (all possible bilateral combinations of countries) and T= 42. The total number of observations is NT =27,300.

Following the discussion in section 4.2, one of the contributions of this chapter is to perform the analysis and the estimation of the gravity equation for the euro effect using two sets of variables. In the first one, that we call "Baldwin-variables" and use upper-case letters, the series have been computed as suggested by BT. The second set of variables, defined as it is commonly done in mainstream gravity literature, is called "standard-variables" and we use lower-case letters to represent them.

Concerning the dependent variable, and following BT's critiques, we include exports as dependent variable instead of the average of exports and imports, as it is frequently made in the literature. As BT points out, the gravity equation is an expenditure function that explains uni-directional bilateral trade flows. De Benedictis and Taglioni (2011) also reinforce this point, arguing that the choice of the dependent variable should be driven by theoretical considerations, which privilege the use of uni-directional trade data[26]. In addition, mistakes derived from a wrong averaging are avoided in this fashion. Hence, $EXPORTS_{ijt}$ is the log of the export flows from country i to country j in nominal terms[27]-according to BT's critiques-, obtained from the CHELEM – CEPII database and expressed in current dollars. $exports_{ijt}$ stands for the logarithm of real exports in US dollars, deflated using the US CPI obtained from the IMF International Financial Statistics (IFS); GDP_{it} and GDP_{jt} are the logs of the nominal GDPs in the exporter and importer country respectively and gdp_{it} and gdp_{jt} are the logs of the real PPP-converted GDPs in each country. Both are obtained from CHELEM-CEPII database. Additionally, two dummy variables have been built to include the effect of particular integration agreements on trade. Namely RTA_{ijt} which is 1 if both countries have a regional trade agreement at time t is constructed using World Trade Organization (WTO) data, and finally the key variable of interest $EURO_{ijt}$ which equals 1 if both trading partners belong to the euro area in year t and zero otherwise. To the extent that these agreements are made or dissolved during the sample period, this variable is distinct from the time-invariant country pair fixed effect.

The formal model that we estimate comes from the gravity equation, and in particular, we follow the traditional specification from the recent literature on the euro effect using non-stationary panels. The purpose is to isolate the effects of EMU on exports trying to control for other factors that may have an influence on exports flows but are not related to the monetary union. The gravity model predicts that bilateral exports should depend on factors such as economic size or 'mass', distance, and other related considerations. Bearing this in mind the basic panel equation in the literature can be expressed as follows:

[26] See De Benedictis and Taglioni (2011), p. 71.
[27] Since we include OECD countries, the total number of zero observations represents only the 0.2% of total flows (64 observations). We have replaced these zero flows by 0.01.

$$EXPORTS_{ijt} = \beta_1 GDP_{it} + \beta_2 GDP_{jt} + \delta_1 EMU_{ijt} + \delta_2 RTA_{ijt} + \eta_{ij} + \tau_{ij} \,.\, t + \varepsilon_{ijt} \quad (4.1)$$

where η_{ij} is a country specific fixed effect, $\tau_{ij} \,.\, t$ is a country pair specific time trend and ε_{ijt} is the error term[28].

The fixed effect, η_{ij}, is intended to capture all individual fixed factors, including unobservable characteristics associated with a given pair of countries that have affected bilateral trade flows historically. These time invariant factors include geographical distance, area, common language, common border, etc. The advantage of fixed effects estimation over directly including these specific measures is controlling for omitted variables bias as a whole at the expense of isolating the individual contribution of each of the variables considered (Micco et al. 2003)[29].

The country pair specific time trend, $\tau_{ij} \cdot t$, is intended to capture all country pair specific omitted trending variables, for instance, institutional characteristics, factor endowments, and cultural aspects that may change over time.[30] Therefore, the approach that we follow to account for trend effects is very flexible and considers both, the time dimension and the heterogeneous behaviour (coefficients) across country pairs. Potential bias due to the existence of common time effects is also controlled through the use of common factors; hence, time effects are not included in the specification.

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[28] Later in the analysis, we will include additional deterministic trends in equation (4.1), which correspond to structural breaks in the constant, the trend or both.

[29] Hence, the model does not include distance between countries as an explanatory variable since the country pair specific fixed effects will account for the distance effect. Moreover, as we have previously stated, the econometric approach used in this chapter accounts for spatial dependence properly.

[30] Country-pair specific variables, such as transport costs or tariff, can vary over time due to technical progress in transport and telecommunications or to the trade liberalization process, generating trends in trade that must be accounted for.

The set of coefficients δ_1 and δ_2 represents the effect of any RTAs and EMU on member states' exports to their country peers (including extra-area trade). Therefore, the parameter of interest is δ_2 and the difference in exports before and after the introduction of the euro is used to identify this coefficient.

The next subsections are devoted to the presentation of the empirical results, comprising panel estimates of EMU trade effects at the area-wide level as well as cross-country differences.

4.3.2. Panel unit root, stationarity tests and cross-section dependence

We use a testing procedure that deals with the problem of cross-section dependence. First, we compute the test statistic by Pesaran (2004) to assess whether the time series in the panel are cross-section independent. Then, we proceed in a second stage to compute unit root statistics that account for such dependence when required.

Pesaran (2004) designs a test statistic based on the average of pair-wise Pearson's correlation coefficients $\hat{\rho}_j$, $j = 1, 2, \ldots n$, $n = N(N - 1)/2$, of the residuals obtained from an autoregressive (AR) model. We estimate an autoregressive model to isolate cross-section dependence from the autocorrelation that might be driving the individual time series. Under the null hypothesis of cross-section independence the CD statistic of Pesaran (2004) converges to the standard normal distribution. The results in Table 4.1 show that Pesaran's CD statistic strongly rejects the null hypothesis of independence both in real and nominal exports; so that cross-section dependence has to be considered when computing the panel data statistics if misleading conclusions are to be avoided. Note that, according to Pesaran (2004) the CD test is valid for N and T tending to ∞ in any order and that it is particularly useful for panels with small T and large N. Moreover, this test is also robust to possible structural breaks, which makes it especially suitable for our study.

Table 4.1. Pesaran's CD and CADF statistics

Variable	**CD dependence test**	**CADF panel unit root test**
gdp_{it}	-0.239	0.392
GDP_{it}	-0.146	-1.235
gdp_{jt}	-0.239	0.392
GDP_{jt}	-0.146	-1.235
$Exports_{ijt}$	115.911***	2.331
$EXPORTS_{ijt}$	105.136***	-0.964

Notes: *** denotes rejection at 1% level. All variables are in logarithms. Upper case letters stands for nominal (Baldwin) variables and lower case for real (standard) variables. One lag is selected for GDP; two lags for standard exports and one for Baldwin exports. Trend and constant are included in all cases.

Once we have found evidence of dependence, we study the order of integration of the variables. Several procedures to test for unit roots in panels are already available in the literature, from the early works of Levin et al. (2002). However, these first generation tests were based on the unrealistic assumption of cross-section independence[31]. Therefore, we follow Pesaran (2007) and Bai and Ng (2004) and specify the unit root tests allowing for cross-sectional dependence as driven by a common factor model, so that it is possible to distinguish between the idiosyncratic component and the common component. Pesaran (2007) [32] suggests augmenting the Im et al. (2003) test with the cross-sectional averages of lagged levels and their first differences of the individual series to proxy the common factors between the cross-sectional units. The test is based on the mean of individual Augmented Dickey-Fuller (ADF) t-statistics of each unit in the panel:

[31] Empirical evidence using Levin et al. (2002) Im et al. (2003) and Hadri (2000) tests following the suggestions of O'Connell (1998) and Levin et al. (2002) to correct for the independence bias are available from the authors upon request.

[32] The main advantage of this method is its simplicity to compute while its drawback is that the behaviour of the idiosyncratic component is to some extent neglected being assumed its stationarity.

$$\Delta Y_{ij} = a_i + b_i Y_{i,t-1} + c_i \overline{Y}_{t-1} + d_i \Delta \overline{Y}_t + \varepsilon_{it}; i = 1,...,N; t = 1,...T \tag{4.2}$$

where $\overline{Y}_{t-1} = N^{-1}\sum_{t=1}^{N} Y_{it}$ and $\Delta\overline{Y}_t = N^{-1}\sum_{t=1}^{N}\Delta Y_{it} = \overline{Y}_t - \overline{Y}_{t-1}$ and $\varepsilon_{it} \sim iid(0,\sigma^2)$. The null hypothesis assumes that all series are non-stationary, whereas the alternative considers that some (but not all) of them are stationary. The average of the *N* individual *CADF* t-statistic is employed to test the null

$$\overline{CADF} = N^{-1}\sum_{i=1}^{N} CADF_i \tag{4.3}$$

where $CADF_i$ is the t-statistic of b_i in the previous regression. The results obtained from the Pesaran CADF test are reported in Table 4.1 concluding in favour of non-stationarity, with a critical value of -2.50 at a 5% confidence level.

The second test, proposed by Bai and Ng (2004), is a suitable approach when cross-correlation is pervasive, as it is the case. Furthermore, this method controls for cross-section dependence given by cross-cointegration relationships, potentially possible among our group of countries and variables — see Banerjee et al. (2004). The Bai and Ng (2004) approach decomposes the Y_{it}, as follows:

$$Y_{it} = D_{it} + F_t^{'} \pi_i + e_{it} \tag{4.4}$$

with $t = 1, . . . , T$, $i = 1, . . . , N$, where D_{it} denotes the deterministic part of the model — either a constant or a linear time trend — F_t is a (r x1)-vector that accounts for the common factors that are present in the panel, and e_{it} is the idiosyncratic disturbance term, which is assumed to be cross-section independent. Unobserved common factors and idiosyncratic disturbance terms are estimated using principal components on the first difference model.

For the estimated idiosyncratic component, they propose an ADF test for individual unit roots and a Fisher-type test for the pooled unit root hypothesis ($P_{\hat{e}}$), which has a standard normal distribution. The estimation of the number of common factors is obtained using the panel Bayesian information criterion (BIC) information criterion as suggested by Bai and Ng (2002), with a maximum of six common factors. Bai and Ng (2004) propose several tests to select the number of independent stochastic trends, k_1 in the estimated common factors, $\hat{F}_t$. If a single common factor is estimated, they recommend an ADF test whereas if several common factors are obtained, they propose an iterative procedure to select k_1: two modified Q statistics (MQ_c and MQ_f) that use a non-parametric and a parametric correction respectively to account for additional serial correlation. Both statistics have a non-standard limiting distribution. They test the hypothesis of $k_1 = m$ against the alternative $k_1 < m$ for m starting from $\hat{k}$. The procedure ends if at any step $k_1 = m$ cannot be rejected. The results from the application of the Bai and Ng (2004) statistics are summarized in Table 4.2. Panel A of the table corresponds to the variables defined as it is standard in the gravity equations literature. In panel B, by contrast, the variables have been defined following BT's critiques.

Concerning the idiosyncratic component, the results of the panel ADF unit root tests clearly point to the rejection of the unit root hypothesis; however, the results of the unit root analysis of the factor component for all the variables analyzed point to nonstationarity. In none of the cases presented in Table 4.2 can the null hypothesis of independent stochastic trends be rejected.

Thus, the main conclusion is that the variables are non-stationary. Moreover, its source is not variable-specific, but associated to the common factors.

Table 4.2. Panel Data Statistics based on Approximate Common Factor Models.

Bai and Ng (2004) statistics

Panel A: Variables defined following standard literature						
	$exports_{ijt}$		gdp_{it}		gdp_{jt}	
	Test		Test		Test	
Idiosyncratic ADF statistic	-3.542***		-3.505***		-3.505***	
	Test	$\hat{r}_1$	Test	$\hat{r}_1$	Test	$\hat{r}_1$
MQ test (parametric)	-2.716	1	-36.314	6	-36.314	6
MQ test (non-parametric)	-4.155	1	-36.240	6	-36.240	6
Panel B: Variables defined following BT's critiques						
	$EXPORTS_{ijt}$		GDP_{it}		GDP_{jt}	
	Test		Test		Test	
Idiosyncratic ADF statistic	-3.387***		-1.849***		-1.849***	
	Test	$\hat{r}_1$	Test	$\hat{r}_1$	Test	$\hat{r}_1$
MQ test (parametric)	-34.968	4	-21.987	6	-21.987	6
MQ test (non-parametric)	-32.057	4	-23.343	6	-23.343	6

Notes: *** denotes rejection at 1% level. The tests on the factors are asymptotically independent of the tests on the idiosyncratic errors. MQ_c and MQ_f use a non-parametric and a parametric correction respectively to account for additional serial correlation. Both statistics have a non-standard limiting distribution.

4.3.3. Panel cointegration

As in the case of the unit root tests, the main caveat of the first generation panel cointegration tests is that they do not consider the presence of cross-section dependence among the members of the panel[33]. Trying to solve this problem, new statistics have been also designed to test for cointegration using factor models in a fashion similar to the one proposed by Bai and Ng (2004) for unit root testing. Moreover, as the existence of structural breaks in the cointegrating relationships biases the results in panel settings - see Banerjee

[33] We have also applied the panel cointegration tests proposed by Kao (1999) and McCoskey and Kao (1998) for the sake of comparison. The complete results are available from the authors upon request.

and Carrion-i-Silvestre (2010) - they propose an extension of the Gregory and Hansen (1996) approach using common factors to account for dependence. Banerjee and Carrion-i-Silvestre (2010) propose a panel test for the null hypothesis of no cointegration allowing for breaks both in the deterministic components and in the cointegrating vector and also accounts for the presence of cross-section dependence using factor models. They define a (*m* x 1) vector of non-stationary stochastic process, $Y_{i,t} = (y_{i,t}, x_{i,t}^{'})$ whose elements are individually I(1) with the following Data Generating Process:

$$y_{it} = D_{it} + x_{it}^{'}\delta_{it} + u_{it} \tag{4.5}$$

The general functional form for the deterministic term $D_{i,t}$ is given by:

$$D_{it} = \mu_i + \beta_i t + \sum_{j=1}^{m_i}\theta_{ij} DU_{ijt} + \sum_{j=1}^{m_i}\gamma_{ij} DT_{ijt} \tag{4.6}$$

where DU_{ijt} =1 and DT_{ijt} = (t - T_{it}^{b}) for $t > T_{it}^{b}$ and 0 otherwise, $T_{it}^{b} = \lambda_{ij}^{b} T$ denotes the timing of the *j*-th break, $j = 1,\ldots, m_i$, for the i-th unit, $I = 1,\ldots, N$, $\lambda_{ij}^{b} T \in \Lambda$, being Λ a closed subset of (0,1). The cointegrating vector is a function of time so that

$$\delta_{it} = \begin{cases} \delta_{i1}\, T_{i0}^{c} < t \le T_{i1}^{c} \\ \delta_{i2}\, T_{i1}^{c} < t \le T_{i2}^{c} \\ \cdots\cdots \\ \delta_{ij}\, T_{ij-1}^{c} < t \le T_{ij}^{c} \\ \cdots\cdots \\ \delta_{in_i+1}\, T_{in_i}^{c} < t \le T_{in_i+1}^{c} \end{cases} \tag{4.7}$$

with $T_{i0}^{C} = 0$ and $T_{in_i+1}^{C} = T$, where $T_{ij}^{C} = \lambda_{ij}^{C} T$ denoting the *j*-th time of the break, $j = 1,...,n_i$, for the *i*-th unit, $i = 1,...,N$, for the i – th unit, $i = 1,\ldots,N$, $\lambda_{ij}^{C} \in \Lambda$.

Banerjee and Carrion-i-Silvestre (2010) propose eight different model specifications:

Model 1. Constant term, no linear trend - $\boldsymbol{\theta_{ij} = \beta_i = \gamma_{ij} = 0}$ $\forall ij$ in (4.6) – and constant cointegrating vector.

Model 2. Stable trend - $\boldsymbol{\theta_{ij} = 0; \beta_i \neq 0}$ and $\boldsymbol{\gamma_{ij} = 0}$ $\forall ij$ in (4.6) – and constant cointegrating vector.

Model 3. Constant term with shifts; stable trend - $\boldsymbol{\theta_{ij} \neq 0; \beta_i \neq 0; \gamma_{ij} = 0}$ $\forall ij$ (4.6) – and constant cointegrating vector. The model considers multiple level shifts.

Model 4. Constant term, trend and changes in trend, - $\boldsymbol{\theta_{ij} = 0; \beta_i \neq \gamma_{ij} \neq 0}$ $\forall ij$ in (4.6) – and constant cointegrating vector. The model considers multiple trend shifts.

Model 5. Changes in constant and trend - $\boldsymbol{\theta_{ij} \neq 0; \beta_i \neq 0}$ and $\boldsymbol{\gamma_{ij} \neq 0}$ $\forall ij$ in (4.6) – and constant cointegrating vector. The model considers multiple trend and level shifts.

Model 6. No trend, constant term with shifts - $\boldsymbol{\theta_{ij} \neq 0; \beta_i = 0}$ $\forall ij$ in (4.6) – and changes in the cointegrating vector.

Model 7. Constant term, trend; changes in the level - $\boldsymbol{\theta_{ij} \neq 0; \beta_i \neq 0}$ $\forall ij$ in (4.6) – and changes in the cointegrating vector.

Model 8. Constant term, trend; changes in the level and the trend - $\boldsymbol{\theta_{ij} \neq 0}$; $\boldsymbol{\beta_i \neq 0}$ and $\boldsymbol{\gamma_{ij} \neq 0}$ $\forall ij$ in (4.6) – and changes in the cointegrating vector

The common factors are estimated following the method proposed by Bai and Ng (2004). They first compute the first difference of the model; then, they take the orthogonal projections and estimate the common factors and the factor loadings using principal components.

In any of these specifications, Banerjee and Carrion-i-Silvestre (2010) recover the idiosyncratic disturbance terms ($\tilde{e}_{i,t}$) through cumulation of the estimated residuals and propose testing for the null of no cointegration against the alternative of cointegration with break using the ADF statistic.

The null hypothesis of a unit root can be tested using the pseudo *t*-ratio $t^{j}_{\tilde{e}_i}(\lambda_i)$, $j = c, \tau, \gamma$. The models that do not include a time trend (Models 1 and 6) are denoted by c. Those that include a linear time trend with stable trend (Models 2, 3 and 7) are denoted by τ and, finally, γ refers to the models with a time trend with changing trend (Models 4, 5 and 8).

When common (homogeneous) structural breaks are imposed to all the units of the panel (although with different magnitudes), we can compute the statistic for the break dates, where the break dates are the same for each unit, using the idiosyncratic disturbance terms[34].

[34] As described in equations (4.5) and (4.6), a heterogeneous version of the test is also possible, although the homogeneous case is the more adequate for the particular case of the gravity model and the estimation of the parameters in the long-run relationship.

Table 4.3. Banerjee and Carrion (2010) BC cointegration tests

	Standard model			Baldwin model		
Model		r	r_1	Z_j^*	r	r_1
1	-12.78***	6	6	-5.66***	6	6
2	-7.43***	6	6	-5.59***	6	6
3	-15.18***	6	6	-7.72***	6	6
4	-8.08***	6	6	-6.19***	6	6
5	**-16.12*****	**6**	**6**	**-15.88*****	**6**	**6**
6	-18.60***	6	6	-10.02***	6	6
7	-22.74	6	6	-16.09	6	6
8	-18.21	6	6	-15.97	6	6

Notes: *** denotes rejection at 1% level. We use BIC criterion to select the correct specification. Specification 5 is selected in both cases; it includes a constant, a trend and a break in both components. The break takes place in 1990 for the standard model and in 1989 for the Baldwin model. The null of no cointegration is rejected in all cases. r_1 is the number of independent stochastic trends underlying the r common factors; r is the total number of factors allowed in the specification.

In Table 4.3 we present the results of the tests for non-cointegration Z_j^* for the model with homogeneous structural breaks and the eight potential specifications discussed above. Using the BIC_3 information criterion[35], proposed by Bai and Ng (2002) we choose specification 5 both in standard and Baldwin model, which contains a constant and a trend and the structural break affects them both simultaneously. In the two cases, we apply the statistics based on the accumulated idiosyncratic components, Z_j^* to test for non-cointegration. We present the tests for all possible model specifications. With all of them the null hypothesis of non-cointegration is rejected. Concerning the time of the break, for the variables constructed following BT's critiques we find the break in 1989, whereas for the standard variables the break is found in 1990.

[35] This criterion is more appropriated than BIC since it takes into account the panel nature of the problem by including the *N* dimension in the calculation of the function. See Bai and Ng (2002) for further information.

The next step of the analysis is to estimate the long-run relationship in the form of a gravity equation. For this purpose, we will use efficient techniques proposed by Bai et al. (2009).

4.3.4. Estimation of the gravity equation

Once the different tests applied have provided us with evidence of cointegration, either considering a stable relationship or instabilities, we should obtain the long-run estimates using consistent techniques. Kao and Chiang (2000) recommended the fully modified (FM) estimator of Phillips and Hansen (1990) and the dynamic ordinary least squares (DOLS) estimator as proposed by Saikkonnen (1991) and Stock and Watson (1993). However, although both estimators consistently estimate the long-run parameters and correct for autocorrelation and endogeneity, any of the two account for dependence[36]. This fact is very relevant in this study as we found in the PANIC analysis due to Bai and Ng (2004) that the common factors were non-stationary. Bai et al. (2009) consider the problem of estimating the cointegrating vector in a panel data model with non-stationary common factors. The presence of common sources of non-stationarity leads naturally to the concept of cointegration. In addition, by putting a factor structure one can deal with other sources of correlation and with large panels, as it is our case.

Bai et al. (2009) treat the common $I(1)$ variables as parameters. These are estimated jointly with the common slope coefficients β using an iterated procedure. The estimators are $\sqrt{nT}$ consistent and enable the use of standard tests for inference. The approach is robust to mixed $I(1)/I(0)$ factors as well as mixed $I(1)/I(0)$ regressors.

[36] We have estimated the cointegration vectors by fully modified (FM) and dynamic ordinary least squares (DOLS). We have omitted most of these results from the text, although they are available upon request.

They consider the following model:

$$y_{it} = x_{it}^{'}\beta + e_{it} \tag{4.8}$$

where for $i = 1,\dots,n$, $t = 1,\dots,T$, y_{it} is a scalar,

$$x_{it} = x_{it\ 1} + \varepsilon_{it} \tag{4.9}$$

x_{it} is a set of k non-stationary regressors, β is a k x 1 vector of the common slope parameters, and e_{it} is the regression error. They assume that e_{it} is stationary and *iid* across i.

The cross-section pooled least squares estimator of β would be:

$$\hat{\beta}_{LS} = \left(\sum_{i=1}^{n} \sum_{t=1}^{T} x_{it} x_{it}^{'} \right)^{-1} \sum_{i=1}^{n} \sum_{t=1}^{T} x_{it} y_{it} \tag{4.10}$$

Although this estimator is, in general, T-consistent, there is an asymptotic bias due to the long-run correlation between e_{it} and ε_{it}. The asymptotic bias can be estimated and a panel fully-modified estimator can be developed as in Phillips and Hansen (1990) to achieve $\sqrt{nT}$ consistency and asymptotic normality. In addition, they model cross-section dependence by imposing a factor structure on e_{it}:

$$e_{it} = \lambda_i^{'} F_t + u_{it} \tag{4.11}$$

where F_{it} is an $r \times 1$ vector of latent common factors, λ_i is an $r \times 1$ vector of factor loadings and u_{it} is the idiosyncratic error. If both F_t and u_{it} are stationary, then e_{it} is also stationary. In this case, a consistent estimator of the regression coefficients can still be obtained even when the cross-section dependence is ignored. Though, it is crucial to note that when F_t is $I(1)$, if

$F_t = F_{t-1} + \eta_t$, then e_{it} is I(1) and the pooled OLS in (4.10) is not consistent. This is why Bai et al. (2009) develop the case of non-stationary common factors, aiming at achieving consistent estimators. Let the true model in vector form be

$$y_i = x_i \beta^0 + F^0 \lambda_i^0 + u_i \tag{4.12}$$

where

$$y_i = \begin{bmatrix} y_{i1} \\ y_{i2} \\ \cdots \\ y_{iT} \end{bmatrix}, \quad x_i = \begin{bmatrix} x_{i1}^{'} \\ x_{i2}^{'} \\ \cdots \\ x_{iT}^{'} \end{bmatrix}, \quad F = \begin{bmatrix} F_1^{'} \\ F_2^{'} \\ \cdots \\ F_T^{'} \end{bmatrix}, \quad u_i = \begin{bmatrix} u_{i1} \\ u_{i2} \\ \cdots \\ u_{iT} \end{bmatrix} \tag{4.13}$$

When the common factor F_t is observed, they propose what can be considered the panel version of the Phillips and Hansen (1990) statistic, a linear estimator that they call $\tilde{\beta}_{LSFM}$ and the bias corrected version that is identical. The estimators are consistent and the limiting distributions are normal. However, in the majority of the cases, the factors F_t are unobserved and the LSFM estimator is infeasible. In this case F_t should be estimated along with β by minimizing the objective function:

$$S_{nT}(\beta, F, \Lambda) = \sum_{i=1}^{n} (y - x_i \beta - F\lambda_i)^{'} (y - x_i \beta - F\lambda_i) \tag{4.14}$$

subject to the constraint $T^{-2} F^{'} F = I_r$ and $\Lambda^{'}\Lambda$ is positive definite, where $\Lambda = (\lambda_1, ..., \lambda_n)^{'}$ is an *n* x *r* matrix. After concentrating out λ, the least squares estimator for β for a given F is then

$$\hat{\beta} = \left(\sum_{i=1}^{n} x_i^{'} M_F x_i \right)^{-1} \sum_{i=1}^{n} x_i^{'} M_F x_i \qquad (4.15)$$

Although F is not observed when estimating β, and β is not observed when estimating F, unobserved quantities can be replaced by initial estimates and iterate until convergence. Defining

$$S_{nT}(\beta, F) = \sum_{i=1}^{n} (y - x_i\beta)^{'} M_F (y - x_i\beta) \qquad (4.16)$$

the CUP estimator for (β, F) would be

$$(\hat{\beta}_{Cup}, \hat{F}_{Cup}) = \arg\min_{\beta, F} S_{nT}(\beta, F) \qquad (4.17)$$

More precisely, $(\hat{\beta}_{Cup}, \hat{F}_{Cup})$ is the solution to the following two nonlinear equations

$$\hat{\beta} = \left(\sum_{i=1}^{n} x_i^{'} M_{\hat{F}} x_i \right)^{-1} \sum_{i=1}^{n} x_i^{'} M_{\hat{F}} x_i \qquad (4.18)$$

$$\hat{F} V_{nT} = \left[\frac{1}{nT^2} \sum_{i=1}^{n} (y_i - x_i\hat{\beta})(y_i - x_i\hat{\beta})^{'} \right] \hat{F} \qquad (4.19)$$

where $M_{\hat{F}} = I_T - T^{-2}\hat{F}\hat{F}^{'}$ since $\hat{F}^{'}\hat{F}/T^2 = I_r$, and V_{nT} is a diagonal matrix consisting of the r largest eigenvalues of the matrix inside the parenthesis , arranged in decreasing order. The estimator is obtained solving for $\hat{\beta}$ and $\hat{F}$ using (4.18) and (4.19) and it is consistent for β, although it still has a bias derived from having to estimate F_t. The authors correct this bias using two fully-modified estimators. The first one directly corrects the bias of $\hat{\beta}_{Cup}$ and is denoted $\hat{\beta}_{CupBC}$. The second one makes the correction in each iteration and is denoted $\hat{\beta}_{CupFM}$.

We present in Table 4.4 the results of the CUP estimation using the methodology of Bai et al. (2009). We have based our estimation on the results previously obtained using the Banerjee and Carrion-i-Silvestre (2010) tests concerning not only the cointegration tests, but also the deterministic specification of the chosen model. Bai et al. (2009) consider extensions of their estimators when the assumptions about the deterministic components are relaxed.

Table 4.4. CUP estimation of the long-run parameters 1967-2008

	LSDV	**Bai FM**	**CUP-FM**	**CUP-BC**
Variables defined following standard literature				
	$exports_{ijt}$	$exports_{ijt}$	$exports_{ijt}$	$exports_{ijt}$
gdp_{it}	1.18***	0.86***	0.85***	0.85***
	(72.57)	(30.90)	(30.37)	(30.34)
gdp_{jt}	1.10***	1.00***	1.01***	1.01***
	(67.51)	(27.50)	(27.62)	(27.59)
RTA	0.81***	0.34***	0.24***	0.23***
	(16.29)	(9.41)	(6.69)	(6.52)
EMU	-0.09	-0.23***	-0.27***	-0.27***
	(-0.78)	(-3.52)	(-4.13)	(-4.10)
Variables defined following BT's critiques				
	$EXPORTS_{ijt}$	$EXPORTS_{ijt}$	$EXPORTS_{ijt}$	$EXPORTS_{ijt}$
GDP_{it}	1.17***	0.67***	0.64***	0.64***
	(64.00)	(27.14)	(25.54)	(25.37)
GDP_{jt}	1.08***	0.79***	0.78***	0.78***
	(59.66)	(27.18)	(26.34)	(26.29)
RTA	0.79***	0.33***	0.22***	0.22***
	(13.41)	(7.55)	(5.22)	(3.36)
EMU	0.56***	0.26***	0.17**	0.16**
	(4.23)	(3.36)	(2.23)	(2.07)

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. t-statistic in parenthesis . The specification 5 is estimated with 2 common factors according to PCA. The common structural break takes place in 1989 for the Baldwin model and 1990 for the standard. The bandwidth parameter is 0.20 for Baldwin variables and 0.18 for standard variables according to Silverman's rule of thumb.

In order to account for the existence of incidental trends (intercept and/or trend), they define accordingly the projection matrix *M* considered above for demeaned and/or detrended variables. We concentrate the deterministic components before we estimate the long-run parameters. As we have mentioned above, among the deterministic components we include the constant, the country pair specific trends, the common break in the constant and the common break in the country pair specific trends[37]. The number of common factors for the estimation is selected according to Principal Components Factor Analysis (PCA).Therefore, once we have performed this transformation we are able to apply the Bai et al. (2009) estimators to the two definitions of the variables. The results are shown in Table 4.4, where we have also included the LSDV estimation results and the Bai and Ng (2006) two-step fully-modified estimator (Bai FM henceforth) for the sake of comparison. However, it should be noted that the only estimators that are consistent when the common factors are non-stationary are the CUP-FM and the CUP-BC. These results are presented in the last two columns of the table. Although the LSDV estimator is the most commonly applied in the gravity equation literature, the parameters obtained are biased when the common factors are non-stationary. The size of this bias is shown in Bai et al. (2009) and this may explain earlier results in the applied literature.

Let us first analyze the upper part of Table 4.4, where we present the results obtained when we use the variables defined as they commonly are in the empirical literature. We transform them to account for the deterministic components and the structural break found in 1990, at the eve of the creation of the EMU. In this case, the GDP variables are positive and significant in all cases, though larger than expected, which reveals the existence of a bias due to the incorrect definition of variables. The RTA dummy is positive and significant;

[37] Note that this implies that in the model specification of the gravity equation in expression (4.1) above, we have filtered all the variables of the deterministic components.

however, the EMU dummy is incorrectly signed in all cases. The reason behind this striking result should be attributed to the BT's critiques, already mentioned above in this chapter.

The lower files of Table 4.4 contain the results obtained when the variables are correctly defined, constructed according to BT's critiques. In this case, the EMU dummy is correctly signed and significant. The CUP-BC and CUP-FM estimators provide lower results than LSDV and Bai FM, which confirm our theoretical predictions of the need for accounting dependence and nonstationarities. We should note that the LSDV estimator is shifted away from zero due to an asymptotic bias induced by the cross-section dependence. The RTA coefficient is again positive and significant and its effect is also notably reduced when using the proper estimators.

Concerning the GDP variables, the values obtained are around 0.65 and 0.8 for the exporter and importer respectively. The importer GDP shows a higher coefficient than the exporter GDP, which reflects the fact that demand forces have greater influence on trade than supply forces. Again, the two significant estimated coefficients obtained using LSDV are larger than those obtained with the other estimators due to the above-mentioned bias. The Bai FM estimator, by contrast, corrects for the presence of dependence and assumes stationary common factors. However, Bai et al. (2009) strongly recommend the use of the CUP-FM and CUP-BC when there is dependence and the common factors are non-stationary. The common structural break occurs in this case in 1989, a year which is very close to the date of the signing of the Single European Act (1987).

Therefore the main empirical findings can be summarized as follows: first, there exists a long-run relationship linking trade and the gravity equation variables in a system that exhibits cross-section dependence and non-stationary common factors, which cancel-out in cointegration. Second, there are some significant instabilities that can be identified using panel cointegration tests that also account for the common factors. Third, the existence of dependence and non-stationary common factors makes it necessary to use consistent estimators, notably the CUP-FM and CUP-BC estimators proposed by Bai et al. (2009). The best results are obtained using the variables constructed à la Baldwin. All in all, the unrealistically high effects of the euro on trade found in previous empirical literature are notably reduced when the trend of the integration process is accounted for. Our results are in line with the most recent literature started with Bun and Klaasen (2007), Fidmurc (2009), Gengenbach (2009) and Berger and Nitsch (2008). They show that the increase in trade within the euro area should be viewed as a continuation of a long-run trend, probably linked to the broader set of EU's economic integration policies and institutional changes.

4. 4. Summary and concluding remarks

In this chapter we have tried to fill the gaps present in the previous literature on euro effects on trade. Using a dataset that includes 26 OECD countries from 1967 to 2008, we estimate a gravity equation through a cointegration approach fully allowing for cross-section dependence. The analysis consists of three steps. First, unit root tests for cross-sectionally dependent panels are applied. Second, the existence of a cointegration relationship among the variables of a proper specification of the gravity equation is tested. In this exercise we account both for dependence in the cross-section dimension and discontinuities in the deterministic and the cointegrating vector. Third, the appropriate CUP-BC and CUP-FM estimators are used to estimate the long-run relationships.

To the best of our knowledge, this is the first attempt to jointly incorporate in the estimation of a gravity equation for the assessment of the euro effect the following aspects:

first, we include BT's critiques in terms of model specification and variables' construction and we include country pair specific trends; second, we account for the existence of cross-section dependence as well as structural breaks in the time domain; and third, we consider the non-stationary nature of the series involved in the analysis. This approach allows us to put the adoption of the euro by EMU members in historical perspective. We argue that the creation of the EMU is best interpreted as a progression of policy changes that have contributed to greater economic integration among EMU countries over the last decades with some significant milestones, such as the Single European Act or the creation of the euro area. We find strong evidence of a gradual increase in trade intensity between European countries as well as pervasive cross-section dependence. Once we control for both, dependence and this (breaking) trend in trade integration, the effect of the formation of the EMU is notably reduced in line with most recent empirical literature.

Appendix A

The Bai FM, CUP-BC and CUP-BC estimators are constructed by making corrections for endogeneity and serial correlation to the OLS estimator in (4.10). These corrections involve kernel density estimation of the long-run covariance matrix. However, when the length of time series is short the estimate of the long-run covariance matrix may be sensitive to the length of the bandwidth. Thus, in this section we check the robustness of our results to alternative choices of the bandwidth parameter. We compare our previous results with those obtained with the number selected by Bai et al. (2009) in their Monte Carlo simulations.

Table 4.A.1. Robustness checks with alternative bandwidth

	LSDV	**Bai FM**	**CUP-FM**	**CUP-BC**
Variables defined following standard literature				
Dependent variable	$exports_{ijt}$	$exports_{ijt}$	$exports_{ijt}$	$exports_{ijt}$
gdp_{it}	1.18***	0.78***	0.79***	0.79***
	(72.57)	(24.25)	(24.74)	(25.00)
gdp_{jt}	1.10***	1.12***	1.12***	1.12***
	(67.51)	(26.72)	(26.88)	(27.36)
RTA	0.81***	0.03	0.11***	0.10***
	(16.29)	(0.75)	(2.70)	(2.58)
EMU	-0.09	-0.77***	-0.42***	-0.52***
	(-0.78)	(-10.01)	(-5.63)	(-6.94)
Variables defined following BT's critiques				
Dependent variable	$EXPORTS_{ijt}$	$EXPORTS_{ijt}$	$EXPORTS_{ijt}$	$EXPORTS_{ijt}$
GDP_{it}	1.17***	0.46***	0.52***	0.51***
	(64.00)	(16.33)	(19.20)	(18.92)
GDP_{jt}	1.08***	0.71***	0.82***	0.83***
	(59.66)	(21.65)	(25.64)	(26.05)
RTA	0.79***	-0.01	0.08**	0.05
	(13.41)	(-0.13)	(1.83)	(1.14)
EMU	0.56***	0.51***	0.35***	0.46***
	(4.23)	(5.82)	(4.22)	(5.61)

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. t-statistic in parenthesis . The specification 5 is estimated with 2 common factors according to PCA. The common structural break takes place in 1989 for the Baldwin model and 1990 for the standard. The bandwidth parameter is 5 following Bai et al. (2009).

In Table 4.4 we have shown the results setting the bandwidth at 0.20 for the Baldwin variables and at 0.18 for the standard variables, calculated according to Silverman's rule of thumb. In their article, Bai et al. (2009) perform the Monte Carlo simulations using a bandwidth of 5. Hence, we present in Table 4.A.1. the results when the selected bandwidth parameter is 5. It can be appreciated that there are not substantial differences in the magnitude, significant or sign of the coefficients.

Chapter 5

EMU and intra-European trade. Long-run evidence using gravity equations

Abstract. In this chapter we present evidence of the long-run effect of the euro on exports for the twelve initial EMU countries for the period 1967-2008 from a double perspective. First, we pool all the bilateral combinations of export flows among the EMU countries in a panel cointegration gravity specification. Second, we estimate a gravity equation for each EMU member vis-à-vis the other eleven partners. Whereas the joint gravity equation provides evidence on the aggregate effect of the euro on intra-European exports, by isolating the individual countries we assess which of them have obtained a larger benefit from the euro. Moreover, this strategy permits to check the robustness of the aggregate results and to find possible asymmetries. Finally, we repeat both the aggregate and individual analysis for the bilateral exports of EMU members to third countries. From an econometric point of view, we apply panel cointegration techniques based on factor models that account for cross-dependence and structural breaks.

5. 1. Introduction

The effect that a currency union has on trade has been largely explored in the literature. Rose (2000) is one of the most cited articles in this field, and his prediction of a tripling of trade for a country when it joins a currency union has been revisited several times. Moreover, the creation of the EMU has provided researchers a natural experiment to further investigate on this effect, thus renewing the debate and leading to improvements in both the specification and estimation of the gravity equation. Although initial estimates were found to be quite high, - 26% in Micco et al. (2003) or 27% in Barr et al. (2003) - more recent literature has considerably reduced this effect. In this line, Berger and Nitsch (2008) claim that the creation of the EMU should be interpreted as a continuation of a series of policy changes that have led over the last decades to greater economic integration among EMU countries. Other articles supporting this hypothesis are Bun and Klaasen (2007), Fidrmuc (2009), Gengenbach (2009). Hiller and Kruse (2010) provide an analysis of this integration process, revealing the most relevant dates in the integration process for each one of the EMU countries.

Traditionally, the model used to estimate the euro effect on trade has been the gravity equation with a euro dummy that takes value one if the countries involved in the trade flow belong to the euro area. However, the long-run nature of the European integration process requires a proper specification and estimation of this equation to avoid biases due to the omission of variables. For that reason, Berger and Nitsch (2008) propose to include a time trend in the specification. A further step is given by Bun and Klaasen (2007) with the introduction of country pair specific time trends that capture the impact of all omitted trending variables with a coefficient that is allowed to vary for each pair of countries. Both articles show that the inclusion of a deterministic trend notably reduces, or even eliminates, the euro effect on trade; however, both ignore the potential existence of stochastic trends in the data. In this chapter we argue that the use of cointegration techniques and the inclusion of time trends –both deterministic and stochastic– is a necessary step in the analysis of the euro effect. Given that its establishment is a long-run process, the nonstationarity of variables or the existence of cointegration relationships among them should be controlled to avoid biases and inconsistencies.

There is still another important caveat in the literature. Frequently the cointegrating relationship is assumed to be stable. Nevertheless, failure to account for the existence of changes in the cointegration relationship and/or the deterministic components affects inference on cointegration analysis, thus leading to wrong conclusions. The standard tests may not reject the null hypothesis of no cointegration when it is false, thus reducing the power of the test. As far as we know, Camarero et al. (2011) and Mancini-Griffoli and Pauwels (2006) are the only articles allowing for the possibility of structural breaks in the data when estimating the gravity equation using cointegration techniques. Camarero et al. (2011) find the break date in 1989. In the case of Mancini-Griffoli and Pauwels (2006) the break date is found in the first quarter of 1999 and three alternative specifications of the gravity equation are estimated using DOLS and an Error Correction Model (ECM). However, these estimators do not correct for cross-section dependence. Since the Pesaran CD statistic reveals the existence of these dependencies, we claim that robust estimators should be employed. We use Banerjee and Carrión-i-Silvestre (2010) cointegration test to properly specify the equation and the break is found to happen in 1987.

Finally, there is little evidence on the asymmetric effect of the euro on its members and in trade with third countries. Faruquee (2004) provides a comparison of this effect on euro-area members by interacting country dummies with the EMU variable. His results show that the Netherlands and Spain are the countries that have obtained the greatest benefits from joining the EMU, while Ireland, Finland and Portugal are the countries with the lowest benefits. Dwane et al. (2011) also perform an analysis of this effect, but they focus on Irish trade. In both cases the possibility of breaks is ignored and cross section dependencies are not modeled. The estimation of the euro effect on trade with third countries has received much less attention in the literature. Kelejian et al. (2011) give evidence of this effect including two dummy variables in the estimation to distinguish between imports and exports, finding positive results. Studies of Micco et al. (2003), Baldwin et al. (2005) and Gil-Pareja et al. (2008) also obtain results in this line. In this chapter, we investigate the aggregate euro effect on internal and external European trade as well as the specific effect on each one of its members in a panel cointegration framework, allowing for structural breaks in the specification. We employ Bai et al. (2009) CUP estimator, which is consistent in the presence of cross section dependencies, and we use a more homogeneous sample –more appropriate when the date of the break is unique. We repeat this analysis for trade of EMU members with third countries. To the best of our knowledge, estimators robust to cross section dependencies and structural breaks have never been applied before to the estimation of the euro effect.

Summing up, the contribution of this chapter to the existent literature is twofold. From an econometric point of view, we improve the specification and estimation of the gravity equation, allowing for the presence of cross section dependencies, nonstationarities and structural breaks in data as well as deterministic and stochastic trends. From an analytical point of view, we investigate the impact of the euro both at the aggregate level and on each one of its members. In addition, we repeat the analysis for EMU exports to third countries to explore the existence of potential diversion effects.

The remainder of the chapter is organized as follows. In section 5.2 we describe the data and the variables used in the analysis, as well as the methodology and tests employed. In section 5.3 we present the results for the EMU as a whole. In section 5.4, two analyses are accomplished; first, we estimate a gravity equation for each EMU member vis-à-vis the other eleven partners and we study the euro effect country by country; second, the same strategy is replicated for the analysis of EMU members' exports to third countries. Finally, section 5.5 concludes.

5.2. Data, methodology and empirical results

5.2.1. Data and model

We include in our study all the countries that joined the EMU in 1999 plus Greece, which became a member in 2001. Belgium and Luxembourg are included as a unique area, so the total number of individuals is 11[38]. The sample contains annual data and covers the period 1967-2008. Hence, we have a balanced panel with dimension N =110 (11x 10, all possible bilateral combinations of countries) and T = 42. The total number of observations is NT =4,620. In a second step, we study the exports of these 11 countries to 15 OECD countries that do not belong to the EMU[39] plus China; so we have a panel with dimension N =176 (11x16) and T =42. Although the number of years available was higher, we have opted by restricting our sample to this period to exclude the effects of the financial crisis that started in 2008. Following BT's critiques, the variables are introduced in nominal terms. Descriptive statistics are presented in Appendix A.

[38] Austria, Belgium and Luxembourg, Finland, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal, Spain.

[39] Australia, Canada, Chile, Denmark, Iceland, Japan, South Korea, Mexico, New Zealand, Norway, Poland, Sweden, Switzerland, United Kingdom and United States.

We use the specification of the gravity equation defined in chapter 4:

$$EXPORTS_{ijt} = \beta_1 GDP_{it} + \beta_2 GDP_{jt} + \delta_1 RTA_{ijt} + \delta_2 EURO_{ijt} + \eta_{ij} + \tau_{ij} \cdot t + \varepsilon_{ijt} \qquad (5.1)$$

The dependent variable is $EXPORTS_{ijt}$, defined as the logarithm of the export flows from country i to country j in nominal terms. GDP_{it} and GDP_{jt} are logarithms of the nominal GDPs – instead of real terms, according to BT's critiques– in the exporter and importer country respectively, obtained from the CHELEM – CEPII database. RTA_{ijt} is a dummy variable that takes value 1 if both countries have a regional trade agreement at time t constructed using World Trade Organization (WTO) data, and $EURO_{ijt,}$ is also a dummy that takes value 1 if both trading partners belong to the euro area in year t and zero otherwise. When analyzing the euro effect with third countries, this variable takes value one when one of the countries involved in the trade flow uses the euro. Our purpose is to isolate the effects of EMU trying to control for other factors that may have an influence on exports but are not related to the monetary union. Egger and Pfaffermayr (2003) show that the gravity model is very sensitive to the proper specification. They propose the inclusion of country pair fixed effects to capture all those bilateral characteristics that are specific to each pair of countries. Following them, we include η_{ij}, a comprehensive set of country pair specific dummies that captures all those bilateral time-invariant unobserved characteristics. We do not include any term to capture the unobserved time effects since the estimators that we will use already include a common factor structure.

Finally, as we did in chapter 4, and following Bun and Klaasen (2003, 2007) we include the term $\tau_{ij} . t$; a time trend with a coefficient that is allowed to vary for each pair of countries in the sample to capture the impact of all country pair specific omitted trending variables.

5.2.2. Panel unit root tests and cross-section dependence

As we noted in chapter 4, there are two important aspects should be taken into account prior to the estimation of the gravity equation. First, it is highly probable that the series are interrelated among them, since the countries in the sample are members of a highly integrated area; the EU. For that reason we implement the Pesaran (2004) CD test. Under the null hypothesis of cross section independence the CD statistic converges to the standard normal distribution. This test is valid for N and T tending to ∞ in any order and that is particularly useful for panels with small T and large N. In addition, this test is also robust to possible structural breaks, which makes it especially suitable for our study. The results in the first column of Table 5.1 show that the null hypothesis of independence is strongly rejected both in the case of intra-EMU exports and in the case of EMU exports to third countries; hence cross-section dependence should be considered when computing the panel data statistics.

The second important point is the presence of unit roots in the data, which if unaccounted for may lead to wrong conclusions and biased estimates. We apply Pesaran CADF (2007) and Bai and Ng (2004) tests to control for this aspect[40]. The second column of Table 5.1 summarizes the results of the Pesaran CADF test. The null hypothesis of cross-section independence is rejected in all cases.

[40] See chapter 4 for further information on Pesaran (2004) CD, Pesaran CADF (2007) and Bai and Ng (2004) tests.

Table 5.1. Pesaran's CD and CADF statistics

Intra-EMU	**CD dependence test**	**CADF panel unit root test**
GDP_{it}	-0.01	-2.361
GDP_{jt}	-0.01	-2.361
$Exports_{ijt}$	36.76^{***}	-2.273
Third countries	**CD dependence test**	**CADF panel unit root test**
GDP_{it}	-0.01	-2.361
GDP_{jt}	-0.33	-2.334
$Exports_{ijt}$	26.82^{***}	-2.312

Notes: *** denotes rejection at 1% level. All variables are in logarithms. One lag is selected according to AIC and BIC criteria. Trend and constant are included in all cases.

The second test, proposed by Bai and Ng (2004), is a suitable approach when cross-correlation is pervasive, as it is the case. Furthermore, this method controls for cross-section dependence given by cross-cointegration relationships, potentially possible among our group of countries and variables — see Banerjee et al. (2004). Table 5.2 shows the results of this test. The idiosyncratic component is found to be non-stationary for the GDP variables, though stationary for exports. The results of the factor component analysis point also in the same direction; the null hypothesis of independent stochastic trends cannot be rejected in any of the cases. Hence, we have enough evidence to conclude that the variables are non-stationary and that cross-section dependencies are present in our data.

Table 5.2. Panel Data Statistics based on Approximate Common Factor Models.

Bai and Ng (2004) statistics

Intra-EMU						
	Exports$_{ijt}$		***GDP$_{it}$***		***GDP$_{jt}$***	
	Test	*p-value*	*Test*	*p-value*	*Test*	*p-value*
Idiosyncratic ADF statistic	-0.438	0.33	4.856	0.99	4.856	0.99
	Test	$\hat{r}_1$	*Test*	$\hat{r}_1$	*Test*	$\hat{r}_1$
MQ test (parametric)	-40.016	5	-33.766	6	-33.766	6
MQ test (non-parametric)	-40.591	5	-35.338	6	-35.338	6
Third countries						
	Exports$_{ijt}$		***GDP$_{it}$***		***GDP$_{jt}$***	
	Test	*p-value*	*Test*	*p-value*	*Test*	*p-value*
Idiosyncratic ADF statistic	-2.04	0.02	4.856	0.99	-2.259	0.01
	Test	$\hat{r}_1$	*Test*	$\hat{r}_1$	*Test*	$\hat{r}_1$
MQ test (parametric)	-38.804	6	-33.766	6	-25.995	6
MQ test (non-parametric)	-39.182	6	-35.338	6	-26.257	6

Notes: $\hat{r}_1$ is the number of independent stochastic trends underlying the r common factors. The tests on the factors are asymptotically independent of the tests on the idiosyncratic errors.

5.2.3. Evidence of structural breaks in the EMU process

The next step in our empirical strategy is to test whether GDP_{it}, GDP_{jt} and $EXPORTS_{ijt}$ are cointegrated using Banerjee and Carrion-i-Silvestre (2010) test[41]. They propose a panel test for the null hypothesis of no cointegration allowing for breaks both in the deterministic components and in the cointegrating vector that also accounts for the presence of cross-section dependence using factor models. It is worth noticing that inference concerning the presence of cointegration can be affected by misspecification if the existence of breaks is ignored. In Table 5.3 we present the results of the tests for non-cointegration for the model with homogeneous structural breaks. In the left-hand side, the results of the intra-EMU exports are shown, whereas the right hand side provides the results for EMU exports to

[41] See chapter 4 for further information on this test.

third countries. Using the BIC_3 information criterion of Bai and Ng (2002) we choose the specification 5 in both cases, which contains a constant, a trend and a structural break that affects them both simultaneously. In order to test for non-cointegration, we apply the statistics based on the accumulated idiosyncratic components, Z_j^*.

Table 5.3. Banerjee and Carrion (2010) BC cointegration tests

	Intra-EMU			Third countries		
Model	Z_j^*	**r**	**r_1**	Z_j^*	**r**	**r_1**
1	-5.52	6	6	-1.43	6	6
2	0.31	6	6	-2.34	6	6
3	-6.45	6	6	-2.89	6	6
4	-0.68	6	6	-3.36	6	6
5	**-2.85**	**6**	**6**	**-7.62**	**6**	**6**
6	-6.31	6	6	-3.30	6	6
7	-4.34	6	6	-9.44	6	6
8	-4.20	6	6	-8.24	6	6

Notes: Bold values indicate the preferred specification. Model is selected according to BIC_3. The model includes a constant, a trend and a break in both components in 1987 for intra-EMU trade and 1989 for trade with third countries. The null of no cointegration is rejected in all cases. r_1 is the number of independent stochastic trends underlying the r common factors; r is the total number of factors allowed in the specification.

We present the tests for all possible specifications; in all cases the null hypothesis of non-cointegration is rejected. The break is found to happen in 1987 -the year of the signing of the Single European Act (SEA) - for intra-EMU trade and in 1989 for EMU trade with third countries. Although the assumption of a common break for all country pairs might seem a little restrictive, however, the homogeneity of the sample -we include only EMU and OECD countries and China- is enough to find a reasonable break common to all country pairs.

Finally, given that the existence of cointegration relationships is unambiguous, the next step is to estimate the long-run relationship in the form of a gravity equation. For this purpose, in the next section we will employ consistent techniques proposed by Bai et al. (2009). We allow the coefficients of the trend as well as the coefficients of the structural breaks in the trend to be different for each pair of countries, thus introducing a higher degree of heterogeneity in the model[42].

5.3. Estimation of the gravity equation for the EMU

Traditional estimation methods as OLS or LSDV present biases and inconsistencies in the presence of nonstationarities and cointegration relationships among the variables. As noted in chapter 4, the FM estimator of Phillips and Hansen (1990) and the DOLS estimator proposed by Saikkonen (1991) and Stock and Watson (1993) are some of the alternatives employed in the literature. Since any of these estimators account for dependence and the Pesaran CD has revealed the existence of dependencies among the series, we present in this section the results of the CUP estimation using the methodology of Bai et al. (2009), as well as the Bai FM results for the sake of comparison. We have selected the specification according to the results of Banerjee and Carrion-i-Silvestre (2010) tests. In order to account for the existence of incidental trends (intercept and/or trend), Bai et al. (2009) define accordingly the projection matrix M considered above for demeaned and/or detrended variables. We concentrate the deterministic components by filtering the five variables in the equation before estimating the long-run parameters. Among the deterministic components we include the constant, the country pair specific trends, the common break in the constant and the common break in the country pair specific trends. The number of common factors for the estimation is selected according to PCA henceforth.

[42] See Bun and Klaasen (2003, 2007) for further information.

5.4. Results

5.4.1 Intra-EMU trade

In a first step, equation (5.1) is estimated including exports flows among the 11 EMU countries included in the sample. Table 5.4 shows the results of the estimation using CUP-FM and CUP-BC estimators. Bai FM is also included for the sake of comparison. As expected, the exporter and importer GDPs have a positive influence on exports in all cases. The importer GDP has higher effect than the exporter, indicating that demand has greater influence on exports than supply.

Table 5.4. Estimation of the long-run parameters for intra-EMU trade

Variables	**Bai FM**	**CUP-BC**	**CUP-FM**
Dependent variable: $EXPORTS_{ijt}$			
GDP_{it}	0.35***	0.50***	0.50***
	(13.43)	(19.07)	(18.56)
GDP_{jt}	0.93***	1.17***	1.17***
	(35.16)	(44.45)	(44.45)
RTA	0.19***	0.12***	0.12***
	(13.88)	(8.78)	(9.06)
EMU	0.19***	0.13***	0.15***
	(15.15)	(11.12)	(12.25)

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. t-statistic in parenthesis . The specification 5 is estimated with 2 common factors according to PCA. The common structural break takes place in 1987. EMU takes value one when both countries belong to EMU. The bandwidth parameter is 0.25 according to Silverman's rule of thumb. Results with a different number of factors and bandwidth are available under request.

The RTA coefficient is positive and statistically significant in all cases. It is worth noticing that this variable is already capturing the effect of joining the European Free Trade Agreement (EFTA) or the EU. The EMU coefficient is also positive and highly significant, and its magnitude is around 0.15 using CUP estimators; this implies that the adoption of the euro has increased exports between EMU members by 16%[43]. Table 5.B.1 in Appendix B summarizes previous estimates of euro effect in the literature. As it can be appreciated, our results reduce the initial optimistic coefficients and are in line with more recent literature.

Next, we proceed to the analysis of each country separately. To assess which members have obtained larger benefits from joining the euro and to find possible asymmetries, we have constructed 11 additional sub-panels in which the exporter is each one of the EMU countries and the exporters are the 10 remaining members. Hence, we have 11 sub-panels with dimension T =42 and N =10. The empirical strategy followed for each one of these sub-panels is analogous to the one previously employed. In a first step, we have checked the existence of dependencies among the series, as well as the nonstationarity of the variables[44]. Since we have found evidence of both facts, in a second step we have tested the existence of cointegration relationships among the variables using the Banerjee and Carrion-i-Silvestre (2010) test. The results are again positive and the specification 5 is selected among all possible specifications according to the BIC_3 criterion in all cases but Ireland, which also has a break in the cointegrating vector (specification 8). Table 5.5 shows the coefficient of our variable of interest, as well as the date of the break for each country. According to the CUP estimator, the euro effect is found to be negative and significant in the cases of Finland, Ireland and Greece, non-significant for the Netherlands; and positive and significant for the rest of the countries. These results are consistent with those obtained in chapter 3. A tentative explanation for the negative sign could be the fact that these countries got used to depreciate their currency to foster exports before 1999, while after the introduction of the euro they could not use this strategy anymore. More

[43] As noted in chapter 3, to interpret dummy coefficients as a percentage change it is necessary to apply a simple transformation to the coefficient obtained, 100*(EXP(α)-1].

[44] For the sake of brevity, PANIC, CD and Pesaran CADF results are reported only for aggregate sample. Individual results are available upon request.

specifically, Finland faced a commercial crisis after the demolition of the URSS, his main commercial partner, which implied consecutive devaluations after 1990. In the case of Greece, the "hard drachma" policy[45] adopted in 1995 implied a notable appreciation of the drachma during the period 1995-1997. Later on, when Greece joined the ERM in 1998 the currency experienced a devaluation of 12.3%.

Table 5.5. Country comparison of the EMU effect. Intra-EMU trade

Country	**Bai FM**	**CUP BC**	**CUP FM**
	Dependent variable: $EXPORTS_{ijt}$		
Austria	0.28***	0.20***	0.20***
1994	(10.78)	(7.90)	(7.93)
BL	0.35***	0.25***	0.36***
1993	(15.67)	(12.56)	(18.02)
Finland	-0.32***	-0.30***	-0.44***
1987	(-11.15)	(-11.03)	(-15.79)
France	0.14***	0.31***	0.24***
1974	(8.69)	(16.23)	(13.99)
Germany	0.21***	0.07***	0.18***
1974	(13.92)	(6.42)	(15.98)
Greece	-0.06*	-0.07**	-0.09***
1980	(-1.54)	(-2.01)	(-2.57)
Ireland	-0.25***	0.14***	-0.06**
1973	(-8.66)	(5.63)	(-2.16)
Italy	0.19***	0.26***	0.21***
1985	(11.26)	(14.80)	(12.15)
Netherlands	0.11***	-0.13***	-0.02
1975	(6.87)	(-7.76)	(-1.20)
Portugal	0.43***	0.15***	0.34***
1984	(11.66)	(4.19)	(9.91)
Spain	0.35***	0.13***	0.18***
1989	(9.08)	(3.63)	(5.01)

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. t-statistic in parenthesis. BL stands for Belgium and Luxembourg. The specification 5 is estimated with 1 or 2 common factors according to PCA. The year of the break is indicated below the name of each country. Bandwidth parameter is 0.25 according to Silverman's rule of thumb. Results with a different number of factors and bandwidth are available under request.

[45] See Hochreiter and Tavlas (2004) for further information.

On the other hand, all the EMU founding members but the Netherlands have obtained high benefits from its membership, ranging from the 19% in Germany to the 28% in Belgium and Luxembourg. Though a priori a higher coefficient may be expected for Germany, it should be noted that two of their main commercial partners do not belong to the EMU[46]. Hence, it is not surprising that the euro effects have been more moderate in this case. As before, it should be noted that the inclusion of the correct specification (constant, trend and structural breaks), as well as the RTA dummy, capture most of the euro effect, though reducing the high coefficients previously found in the literature.

A closer inspection of the dates of the break provides additional evidence of the integration process in each country. In many cases, the date of the break is very close to the year of EU membership. This is the case of Austria, which became a member in 1995; Greece, in 1981; Ireland in 1973; Portugal in 1986 and Spain in 1986. Belgium-Luxembourg seems to be more affected by the signature of the Maastricht Treaty in 1992, and Finland and Italy present dates more related with the Single European Act (1987). France, Germany and the Netherlands, in contrast, show a break at the very beginning of the period, which makes sense since all of them are founding members. For Germany the date may also be attributed to the Ostpolitik, which implied the normalization of relations between the Federal Republic of Germany and Eastern Europe. For France, the date coincides with the year in which this country abandoned the fixed parity for a free floating.

5.4.2. EMU trade with third countries

The third objective in this chapter is to analyse the euro effect on trade with non-EMU countries. We have included the same EMU exporters, but now we focus on their exports to 16 countries that do not belong to the EMU. Now EMU_{ijt} takes value one when one of

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[46] See Table 5.A.3 in Appendix.

the countries (not the two) involved is an EMU member.

For the estimation of the aggregate effect we have a panel with N = 176 individuals and T =42 years. In addition, for the estimation of the effect for each EMU country we have constructed 11 additional sub-panels with dimension N = 16 and T =42. We have performed the same empirical strategy to check the existence of nonstationarity and cointegration, obtaining again positive evidence[47]. Table 5.6 shows the results of the analysis for the aggregate database. As before, there are no substantial differences between Bai FM and CUP estimators. Both importer and exporter GDP show a positive and significant coefficient, as expected, and again the importer GDP has higher effect. The structural break for the aggregate dataset is found to happen in 1989, very close to the signature of the Single European Act. In this case, the countries included are less related to EMU process; hence this break-date may be related to the Plaza (1985) and Louvre (1987) agreements, which were signed with the objective to stabilize the international currency markets. Belonging to a RTA has an unambiguous positive effect on exports. Although EMU shows now a positive but non-significant coefficient at the aggregate level, a closer inspection of EMU effect on third countries reveals inTable 5.7 that this effect is generally significant in each individual case.

Table 5.7 shows the coefficient of EMU variable and the date of the break for each sub-panel. In this case break dates are found to happen in dates very close to the oil shocks in 1973 and 1979 or the Plaza and Louvre agreements, which makes more sense since these facts are more prone to affect international trade and not only EMU countries. In two cases – Finland and Italy – the break is found already in the nineties.

[47] Results are available upon request.

Table 5.6. Estimation of the long-run parameters. EMU with third countries

Variables	Bai FM	CUP BC	CUP FM
Dependent variable: $EXPORTS_{ijt}$			
GDP_{it}	0.53***	0.53***	0.53***
	(9.11)	(9.04)	(9.09)
GDP_{jt}	0.70***	0.71***	0.71***
	(11.36)	(11. 42)	(11.36)
RTA	0.29***	0.29***	0.30***
	(4.52)	(4.53)	(4.62)
EMU	0.04	0.04	0.04
	(0.84)	(0.77)	(0.87)

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. t-statistic in parenthesis . The specification 5 is estimated with 1 common factor according to PCA. The common structural break takes place in 1989. Bandwidth parameter is 0.25 according to Silverman's rule of thumb. Results with a different number of factors and bandwidth are available under request.

Although no evidence of trade diversion is shown, it is worth noticing that the rest of countries included in the estimation are OECD members and the commercial relationships between these countries and EMU members have been relatively stable since the OECD creation in 1960. Hence, it is not surprising that the introduction of the euro has not affected negatively these links. In fact, the results in chapter 3 revealed that when the sample is extended to developing countries, evidence of a reduction in EMU imports from third countries due to the introduction of the euro appears. The individual inspection of the coefficients reveals a negative effect only in the case of Greece. For Italy, the coefficient is now not significant, in contrast with the large and significant coefficient previously obtained, revealing the EMU-oriented export behaviour of this country. Portugal, although obtaining a positive coefficient in both cases, also exhibits this commercial pattern, with the euro fostering its exports to third countries more than its intra-EMU exports. The opposite case is represented by Austria, Finland, Ireland and the Netherlands, which seem to have obtained higher benefits from the euro on their external exports than on their intra-EMU commercial relationships. Belgium and Luxembourg, France, Germany and Spain show a similar effect in both cases, the euro having an equilibrated effect on their internal and external exports. All in all, Austria, Belgium and Luxembourg and the Netherlands are the countries that have more benefited of the euro in their trade with third countries, a fact that may be explained by the traditionally export openness of these countries.

Table 5.7. Country comparison of EMU effect with third countries

Variables	Bai FM	CUP BC	CUP FM
Dependent variable: $EXPORTS_{ijt}$			
Austria	0.33***	0.29***	0.37***
1980	(10.99)	(9.68)	(12.25)
BL	0.30***	0.30***	0.31***
1984	(9.19)	(9.21)	(9.32)
Finland	0.13***	0.13**	0.14***
1993	(2.16)	(2.23)	(2.34)
France	0.17***	0.17***	0.18***
1985	(5.11)	(5.17)	(5.47)
Germany	0.13***	0.12***	0.14***
1978	(5.82)	(5.21)	(6.07)
Greece	-0.19**	-0.22***	-0.29***
1984	(-2.14)	(-2.47)	(-3.32)
Ireland	0.23***	0.23***	0.26***
1981	(3.99)	(3.82)	(4.37)
Italy	-0.02	-0.01	-0.00
1994	(-0.46)	(-0.21)	(-0.01)
Netherlands	0.09***	0.27***	0.30***
1985	(2.86)	(8.78)	(9.81)
Portugal	0.08**	0.08**	0.08**
1984	(1.67)	(1.79)	(1.73)
Spain	0.13***	0.16***	0.18***
1974	(3.84)	(4.85)	(5.25)

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. t-statistic in parenthesis . BL stands for Belgium and Luxembourg. The specification 5 is estimated with 1, 2 or 3 common factors according to PCA. The year of the break is indicated below the name of each country. Bandwidth parameter is 0.25 according to Silverman's rule of thumb. Results with a different number of factors and bandwidth are available under request.

5.5. Summary and concluding remarks

In this chapter we contribute to the existent literature concerning the euro effect with the application of an estimation method that is consistent in the presence of nonstationarities and dependencies in the data. We use two different datasets; the first one includes exports flows among 11 EMU countries from 1967 to 2008 and the second includes exports from 11 EMU countries to 15 OECD non EMU countries and China during the same period. We estimate a gravity equation through a cointegration approach fully allowing for cross-section dependence. The analysis consists of three steps. First, unit root tests for cross-sectionally dependent panels are applied. Second, the existence of a cointegration relationship among the variables of a proper specification of the gravity equation is tested. In this exercise we account both for dependence in the cross-section dimension and discontinuities in the time dimension. Third, consistent estimation methods (CUP-BC and CUP-FM), that model the dependencies in the data using common factors, are used to estimate the long-run relationships.

Our specification allows for cross-sections dependencies and structural breaks in the time domain as well as nonstationarities in the variables. We find strong evidence of a gradual increase in trade intensity between European countries as well as pervasive cross section dependence. Once we control for both, dependence and this (breaking) trend in trade integration, the effect of the formation of the EMU is reduced in line with most recent empirical literature. Concerning intra-EMU exports, Belgium and Luxembourg, France and Italy are the countries more benefited from the introduction of the euro. The effects for exports to third countries are in general more moderate; and, with the exception of Greece, there is no evidence of diversion effects.

The introduction of a structural break in the specification is an important main contribution of this thesis. In the aggregate case the break is found in 1987 for intra-EMU trade and in 1989 for EMU trade with third countries. We attribute the cause of the intra-EMU break date to the effects of the Single European Act, which came into effect in that year. The main commitment agreed in this Treaty was the adoption of measures aimed to the progressive establishment of a common market over a period that would conclude on 1992. Hence, it is not surprising that it implied a significant change in the trading relationships of EMU countries. For trade with third countries, we relate the break date with the signature of the Plaza and Louvre Agreements, which were important milestones in the international economic context. Concerning the country-specific results, different break-dates are found. For intra-EMU trade, the dates are close to their EU membership, whereas for EMU trade with third countries the breaks are more related with the oil crisis in the 1973-1979 period.

Appendix A

Table 5.A.1. Descriptive statistics

Intra-EMU

Variable	Mean	Std. Dev.	Min	Max	Observations
Log of Exports	34.24	2.31	26.39	39.38	4620
Log GDP of exporters	11.91	1.52	8.09	15.11	4620
Log GDP of importers	11.91	1.52	8.09	15.11	4620
RTA	0.59	0.49	0	1	4620
EMU	0.23	0.42	0	1	4620

Third countries

Variable	Mean	Std. Dev.	Min	Max	Observations
Log of Exports	32.91	2.87	0	39.13	7392
Log GDP of exporters	11.91	1.52	8.09	15.11	7392
Log GDP of importers	11.96	1.87	6.03	16.47	7392
RTA	0.15	0.36	0	1	7392
EMU	0.23	0.42	0	1	7392

Note: Export and GDP variables are expressed in billion US$.

Table 5.A.2. Exports by destination, 1967-2008 (billion US$)

	Intra-EMU		**Third countries**	
	Value	%	Value	%
Austria	93.64	69.81%	40.49	30.19%
Belgium and Luxembourg	327.76	72.27%	125.81	27.73%
Finland	34.14	41.76 %	47.60	58.24 %
France	401.27	64.01 %	225.60	35.99 %
Germany	705.58	55.64 %	562.52	44.36 %
Greece	14.24	70.93 %	5.83	29.07 %
Ireland	59.95	44.39 %	75.09	55.61 %
Italy	303.60	61.87 %	187.13	38.13 %
Netherlands	317.20	70.03 %	135.72	29.97 %
Portugal	38.42	71.39 %	15.40	28.61 %
Spain	160.76	71.18 %	65.10	28.82 %
Total	2,456.56	62.30%	1,486.29	37.70%

Notes: Proportion of intra-EMU and external trade for each EMU country. Source: own elaboration according to CHELEM database.

Table 5.A.3. Main commercial partners of EMU members

	Intra-EMU	**Third countries**
Austria	Germany	Switzerland
	Italy	USA
	France	United Kingdom
BL	Germany	United Kingdom
	France	USA
	Netherlands	Switzerland
Finland	Germany	Sweden
	France	United Kingdom
	Netherlands	USA
France	Germany	United Kingdom
	Italy	USA
	BL	Switzerland
Germany	France	USA
	Italy	United Kingdom
	Netherlands	Switzerland
Greece	Germany	United Kingdom
	Italy	USA
	France	Sweden
Ireland	Germany	United Kingdom
	BL	USA
	France	Japan
Italy	Germany	USA
	France	United Kingdom
	Spain	Switzerland
Netherlands	Germany	United Kingdom
	BL	USA
	France	Sweden
Portugal	Spain	United Kingdom
	Germany	USA
	France	Sweden
Spain	France	United Kingdom
	Germany	USA
	Italy	Switzerland

Notes: The tree main partners of each country are sorted in order of importance. BL stands for Belgium and Luxembourg. Source: own elaboration based on CHELEM database.

Appendix B

Table 5.B.1. Euro effect in previous literature

Article	Period	Countries	Estimation method	EMU coefficient
Barr et al. (2003)	1978-2002	17 EU countries	IV	0.21***
Baldwin et al. (2005)	1985-2002	20 OECD countries	OLS	0.57***
Berger and Nitsch (2008)	1950-2003	18 EU countries	Fixed effects	0.053
Brouwer et al. (2008)	1990-2004	25 EU countries + 4 OECD	Fixed effects	0.144***
Bun and Klaasen (2007)	1967-2002	15 EU countries + 5 OECD	Fixed effects	0.032***
De Nardis et al. (2008)	1988-2004	13 EU + 10 OECD	GMM	0.04*
Eicher and Hehn (2011)	1950-2000	177 countries	Fixed effects	0.339***
Faruqee (2004)	1992-2002	22 industrial countries	DOLS	0.073**
Flam and Nordstrom (2006)	1995-1998 2002-2005	20 OECD countries	Fixed effects	0.165*** 0.232***
Gengenbach (2009)	1967-2002	15 EU countries + 5 OECD	CUP	0.075***
Gil-Pareja et al. (2008)	1950-2004	25 OECD countries	Fixed effects	0.326***
Gomes et al. (2006)	1980-2005	22 industrialized countries	Fixed effects	0.210***
Kelejian et al. (2011)	1991-2006	15 EU countries + 4 OECD	2SLS	0.041*
Micco et al. (2003)	1992-2002	15 EU countries	Fixed effects	0.084***
Serlenga and Shin (2007)	1960-2001	15 EU countries	Fixed effects	0.22*

Chapter 6

Conclusions

The European Union is the outcome of an integration process which started in 1950 with six original states: France, Germany, Italy, the Netherlands, Belgium and Luxembourg. From the beginning, this process has led economists to pay more attention to the development of theoretical considerations and empirical approaches to better understand the role of regional integration and its effects on international trade. Recent advances in econometric techniques have notably contributed to this understanding and have sparked renewed interest in the appropriateness of the estimation methods employed to this end.

In this dissertation, we aim to contribute to this literature in several manners. First of all, in chapter 2 a thorough review of the gravity equation and relevant literature is presented. The performance of several linear and nonlinear estimators is compared using a three-dimensional dataset, analyzing the most relevant properties of each one and revealing their main advantages and disadvantages. Chapter 3 has focused on the effects of ER level and volatility on trade, completely isolating the euro effect from other possible factors affecting trade (RTA, ERV). In addition, the EMU effect on third countries is estimated. In chapter 4 a further step is reached, and we deal with some econometric problems that affect the long run estimation of the equation; namely the nonstationarity of the variables, the presence of cross-correlation among the series and the existence of discontinuities in the time dimension. Finally, in chapter 5 we apply the previous methodology focusing exclusively on the twelve initial EMU countries from a double perspective. First, we pool all the bilateral combinations of export flows among the EMU countries and next we estimate a gravity equation for each EMU member vis-à-vis the other eleven partners. This strategy permits to check the robustness of the aggregate results and to find possible asymmetries. We repeat both the aggregate and individual analysis for the bilateral exports of EMU members to third countries.

Hence, the main contributions may be summarized as follows: first, an empirical strategy to compare estimation methods is suggested and a thorough revision of the main factors affecting trade is performed. The effect of ER level and volatility on trade is analyzed. Second, we improve the specification and estimation of the gravity equation, allowing for the presence of cross section dependencies, nonstationarities and structural breaks in the data as well as deterministic and stochastic trends. Finally, we investigate the impact of the euro both at the aggregate level and on each one of its members and we repeat this analysis for EMU trade with third countries to explore the existence of potential diversion effects. The novelty of this thesis lies in several aspects: to the best of our knowledge, none of the previous studies have included such a complete specification in terms of ER variables and trade agreements, thus ignoring important aspects in the estimation of the euro effect. Furthermore, the use of a large dataset allow us to study the impact of the euro on Eurozone’s trade with third countries and between these third countries as well as the opportunity that the euro offers to other EU countries in terms of trade, an aspect that remain unexplored until the moment. In addition, this is the first time that panel cointegration techniques allowing for structural breaks and cross section dependence are applied to the estimation of the euro effect using the gravity equation.

The results obtained can be summarized as follows. First, concerning the estimation, it is shown that methods that do not properly treat the presence of zero flows on data exhibit noticeably worse performance than others in terms of efficiency. On the other hand, nonlinear estimators show more accurate results and are robust to the presence of heteroskedasticity. It is worth noticing that, although the use of PPML has been proposed by several authors in the literature, it does not behave so well for an aggregate dataset in the presence of unobserved heterogeneity. The new estimators proposed by Bai et al. (2009) -CUP-BC and the CUP-FM- show a good performance and take into account the existence of cross-section dependence and non-stationary common factors. Regarding the specification, our results show the importance of including a country pair specific time trend to capture all country pair omitted trending variables. Besides that, failure to account for the existence of changes in the cointegration relationship and/or the deterministic components affects inference on cointegration analysis, thus leading to wrong conclusions.

With respect to the euro effect on trade, our results show that the common currency has had a positive impact on EMU exports. There is strong evidence of a gradual increase in trade intensity between European countries as well as pervasive cross section dependence. Once that the presence of dependence and (breaking) trends in trade integration are controlled, the effect of the formation of the EMU is reduced in line with most recent empirical literature and the euro effect is predicted to be small. The analysis of the euro effect for individual EMU members reveals the existence of a good deal of variation in the effect of the euro across member countries. Concerning intra-EMU trade, Belgium and Luxembourg, France and Italy are the countries that more benefited from the introduction of the euro. The effects for trade with third countries are in general more moderate. When analyzing EMU trade with the initial 80 countries in the sample, some evidence of diversion effects is shown. However, when focusing on OECD countries, this evidence disappears in all cases but Greece. This is a logical result, given that commercial relationships between EMU and OECD countries are stronger than EMU links with the rest of countries. The analysis of the structural breaks also sheds some light on the integration process. In the intra-EMU case the break is found in 1987, the year in which the Single

European Act came into force, implying the adoption of measures guided to the progressive establishment of a common market over a period that would conclude in 1992. For trade with third countries, the break takes place in 1989, a date more related with the signing of the Plaza and Louvre Agreements, which were important milestones in the international economic context.

Finally, and concerning the volatility issue, our results show a detrimental effect of the ERV on trade and therefore suggest that there is a potential for an increase in international trade by reducing this volatility. Then, the possibility to peg to the euro could boost trade from third countries and among them. However, since EMU countries have clearly lost the possibility to adjust with their ER, one should be cautious when comparing the benefits of the EMU with the gains the countries could obtained from depreciation.

Several lines of research remain open after this investigation. The estimation of long-run relationships among economic variables using panel data techniques is a novel field of research, and many improvements are still required. Hence, a first possible extension would be to study the performance of Pesaran (2006) estimator. As it has been mentioned, Bai et al. (2009) estimator assumes that the cross sections in the model share common sources of non-stationary variation in the form of global stochastic trends. Alternatively, Pesaran (2006) proposes a number of estimators, referred to as Common Correlated Effects (CCE) estimators, in which the unobserved factors and the individual-specific errors are allowed to follow arbitrary stationary processes. The basic idea is to filter the individual-specific regressors by means of cross-section averages such that asymptotically the differential effects of unobserved common factors are eliminated. The main advantage of this procedure with respect to CUP estimators is that it can be computed by least squares applied to auxiliary regressions where the observed regressors are augmented with cross-sectional averages of the dependent variable and the individual-specific regressors. An interesting exercise would be to compare the performance of both CCE and CUP estimators allowing for structural breaks and country pair specific trends in the specification.

In addition, the current economic crisis provides a natural experiment to evaluate the strengths and weaknesses of the EMU and to shed some lights on the benefits of this integration process. Due to the policy relevance of the issue, in particular for the European countries that are still thinking about joining the EMU, it is important to have a robust evaluation of the benefits the euro had on trade and could still have since the debate is in the air.

Finally, this work has focused on the euro effect on trade in goods at the aggregate level. An extension that would help to complete the comprehension of integration processes would be to replicate the analysis for services and FDI flows. Moreover, the use of sector and firm level data would help to disentangle additional effects that remain unnoticed at the aggregate level.

Capítulo 6

Conclusiones

La Unión Europea es el resultado de un proceso de integración que se inició en 1950 con seis países: Francia, Alemania, Italia, Holanda, Bélgica y Luxemburgo. Desde sus inicios, este proceso ha llevado a los economistas a prestar una mayor atención al desarrollo de consideraciones teóricas y aproximaciones empíricas que permitan comprender de forma más profunda el papel de la integración regional y sus efectos en el comercio internacional. Los recientes avances en técnicas econométricas han contribuido notablemente a esta compresión y han despertado un interés renovado en la adecuación de los métodos de estimación empleados para este fin.

En esta tesis contribuimos a la literatura existente de varias maneras. En primer lugar, el capítulo 2 presenta una revisión exhaustiva de la ecuación de gravedad y de la literatura relacionada. El comportamiento de varios estimadores lineales y no lineales es comparado usando una base de datos de tres dimensiones, analizando las propiedades más relevantes de cada uno y poniendo de manifiesto sus principales ventajas e inconvenientes. El capítulo 3 se centra en los efectos del tipo de cambio y la volatilidad, aislando el efecto del euro de otros factores que pueden afectar al comercio. Asimismo, se analiza el efecto que la UEM ha tenido sobre terceros países. En el Capítulo 4 damos un paso más, afrontando algunos de los problemas econométricos que afectan a la estimación de largo plazo de la ecuación, como son la no estacionariedad de las variables, la presencia de dependencia transversal en los datos y la existencia de rupturas en la dimensión temporal. Finalmente, en el Capítulo 5 aplicamos la metodología anterior centrándonos exclusivamente en los doce países iniciales de la UEM desde una doble perspectiva. En primer lugar, analizamos todas las combinaciones bilaterales posibles de flujos de comercio entre los miembros de la UEM usando para ello una ecuación de gravedad estimada con técnicas de cointegración. Posteriormente realizamos el mismo análisis para el comercio de cada uno de los países de la UEM con el resto de miembros. El primer ejercicio proporciona evidencia del efecto agregado del euro en el comercio intraeuropeo; mientras que el segundo, al aislar a cada país de forma individual, permite comprobar la robustez de los resultados y encontrar posibles asimetrías en el efecto del euro sobre sus miembros. Finalmente, replicamos el análisis tanto agregado como individual para el comercio de los países de la UEM con terceros países.

Por tanto, las principales contribuciones de la tesis pueden resumirse en las siguientes: en primer lugar, proponemos una estrategia empírica para la comparación de métodos de estimación y llevamos a cabo una revisión exhaustiva de los principales factores que afectan al comercio, incluyendo el tipo de cambio y la volatilidad entre ellos. En segundo lugar, mejoramos la especificación y estimación de la ecuación de gravedad, permitiendo la presencia de dependencia transversal, no estacionariedad y cambios estructurales en los datos, así como de tendencias deterministas y estocásticas. Por último, investigamos el impacto del euro tanto a nivel agregado como para cada uno de sus miembros de forma individual y repetimos este análisis para el comercio de la UEM con terceros países para descubrir la existencia de posibles efectos de desviación de comercio. La novedad de esta tesis reside en varios aspectos: hasta el momento, ninguno de los estudios anteriores había incluido una especificación tan completa en términos de variables de tipo de cambio y acuerdos comerciales, por lo que se ignoraban aspectos importantes en la estimación del euro. Por otra parte, el uso de una base de datos amplia nos permite analizar el impacto del euro en el comercio de la Eurozona con terceros países y el comercio de esos terceros países entre sí, así como la oportunidad que el euro ofrece a otros países de la Unión Europea en términos de comercio. Además, esta es la primera vez que se aplican técnicas de cointegración en datos panel permitiendo la presencia de cambios estructurales y dependencia transversal a la estimación del euro usando la ecuación de gravedad.

Los resultados obtenidos conciernen principalmente a tres aspectos. En primer lugar, con respecto a la estimación, se demuestra que los métodos que no tratan correctamente la presencia de ceros en la muestra presentan peor comportamiento que el resto en términos de eficiencia. Por otra parte, los estimadores no lineales arrojan resultados más precisos y son robustos a la presencia de heteroscedasticidad. Es importante remarcar que, pese a que el uso del estimador PPML ha sido propuesto por varios autores en la literatura, su comportamiento no es óptimo para datos agregados en presencia de heterogeneidad no observada. Los nuevos estimadores propuestos por Bai et al. (2009) -CUP-BC y CUP-FM- son eficientes y tienen en cuenta la existencia de dependencia transversal y la posibilidad de factores comunes no estacionarios. Con respecto a la especificación, nuestros resultados ponen de manifiesto la importancia de incluir una tendencia temporal cuyo coeficiente no esté restringido, sino que pueda variar para cada par de países, de forma que se recojan todas las posibles variables omitidas de comportamiento tendencial que son específicas en cada caso. Aparte de esto, el hecho de no tener en cuenta la existencia de cambios en la relación de cointegración y/o los componentes deterministas afecta a la inferencia en el análisis, conduciendo a conclusiones erróneas.

En segundo lugar, con respecto al efecto del euro sobre el comercio, los resultados muestran que la creación de una moneda común ha tenido un impacto positivo sobre las exportaciones de la UEM. Hay evidencia de un incremento gradual en la intensidad del comercio entre los países europeos y, una vez que se controla la presencia de dependencia y tendencias en la integración, el efecto de la formación de la UEM sobre el comercio disminuye, en línea con la literatura relacionada, y se predice un impacto del euro reducido. El análisis del efecto del euro para cada uno de los miembros de la UEM revela la existencia de una notable variación sobre los países miembros. En relación al comercio intraeuropeo, Francia, Italia, Bélgica y Luxemburgo son los países que más se han beneficiado de su introducción. Los efectos para el comercio con terceros países son en general más moderados. Al analizar el comercio de la UEM con los 80 países inicialmente incluidos en la muestra se pone de manifiesto la existencia de efectos de desviación de comercio. Sin embargo, al centrar el análisis en el comercio con los países pertenecientes a la OCDE, estos efectos desaparecen en todos los casos salvo el de Grecia. Este es un resultado lógico, puesto que los lazos comerciales entre la UEM y los países de la OCDE son más estrechos que los lazos de la UEM con el resto de países. El análisis de los cambios estructurales también facilita la comprensión del proceso de integración. El caso del comercio intraeuropeo el cambio aparece en 1987, fecha en la que el Acta Única Europea entra en vigor, implicando la adopción de medidas orientadas al progresivo establecimiento de un mercado común a lo largo de un periodo que concluiría en 1992. Para el comercio con terceros países, el cambio aparece en 1989, una fecha más próxima a los Acuerdos del Plaza y el Louvre, que fueron hitos importantes en el contexto económico internacional.

Por último, en relación al tema de la volatilidad del tipo de cambio, nuestros resultados muestran un efecto perjudicial de la misma sobre el comercio y por tanto sugieren que aún hay potencial para un aumento del comercio internacional a través de la reducción de dicha volatilidad. En consecuencia, una posible paridad de otras monedas con el euro podría impulsar tanto el comercio con terceros países como entre ellos. Sin embargo, dado que los países de la UEM han renunciado a la posibilidad de ajustar su situación mediante el tipo de cambio, conviene debe ser prudente a la hora de comparar los beneficios de la UEM con las ganancias que los países podrían obtener de una depreciación.

Al finalizar esta tesis se abren varias líneas de investigación. Por un lado, la estimación de relaciones de largo plazo entre variables económicas con datos de panel es un campo de investigación novedoso, que permite aún muchas mejoras. Una primera extensión sería el estudio del comportamiento del estimador propuesto por Pesaran (2006). Como ha sido mencionado, el estimador de Bai et al. (2009) asume que las unidades del panel comparten fuentes comunes de variación no estacionaria en forma de tendencias estocásticas globales. Alternativamente, Pesaran (2006) propone una serie de estimadores, conocidos como estimadores de efectos comunes correlacionados (CCE), en los que se permite a los factores no observados y los a errores individuales seguir un proceso estacionario arbitrario. La idea consiste básicamente en filtrar los regresores que son específicos del individuo mediante el uso de las medias de corte transversal de forma que asintóticamente se eliminan los efectos de los factores comunes no observados. La principal ventaja de este procedimiento con respecto a los estimadores CUP es que puede ser calculado aplicando mínimos cuadrados a regresiones auxiliares en las que los regresores observados son aumentados con las medias transversales de la variable dependiente y de los regresores específicos de cada individuo. La comparación del comportamiento de los estimadores CUP y CCE permitiendo la presencia de cambios estructurales y tendencias específicas para cada par de países en la estimación constituiría un ejercicio empírico de gran atractivo.

Asimismo, la crisis económica actual sin duda proporciona un experimento natural para evaluar las fortalezas y debilidades de la UEM, y permite arrojar luz sobre los beneficios de este proceso de integración. Debido a la relevancia política de este tema, sobre todo para los países europeos que aún están pensando en unirse a la UEM, es importante tener una evaluación sólida de los beneficios que el euro ha tenido y puede aún tener.

Por último, este trabajo se ha centrado en el efecto del euro sobre bienes a nivel agregado. Una extensión que contribuiría a complementar la comprensión de los procesos de integración sería replicar el análisis para el comercio de servicios o los flujos de inversión directa extranjera. Asimismo, el uso de datos nivel de sector y empresa permitiría desentrañar efectos adicionales que a nivel agregado no son percibidos.

Bibliography

An, G. and Puttitanun, T. (2009). ‘Revisiting McCallum's Border Puzzle’, *Economic Development Quarterly,* 23, 167-170.

Anderson, J.E. and van Wincoop, E. (2003). ‘Gravity with gravitas: a solution to the border puzzle’, *American Economic Review*, 93, 170-192.

Anderson, J.E. (1979). ‘A theoretical foundation for the gravity equation’, *American Economic Review,* 69, 106-116.

Anderson, J.E. and Marcouiller, D. (2002). ‘Insecurity and the pattern of trade: An empirical investigation’, *Review of Economic and Statistics*, 84, 342-352.

Andrews, M., Schank, T. and Upward, R. (2006). ‘Practical fixed effects estimation methods for the three-way error components model’, *Stata Journal*, 6, 461-481.

Baak, S. (2004). ‘Exchange Rate Volatility and Trade among the Asia Pacific Countries’, Econometric Society Far Eastern Meetings 724.

Bai, J., Kao, Ch., and Ng, S. (2009). ‘Panel cointegration with global stochastic trends’, *Journal of Econometrics*, 149, 82-99.

Bai, J. and Ng, S. (2002). ‘Determining the number of factors in approximate factor models’, *Econometrica*, 70, 191-221.

Bai, J. and Ng, S. (2004). ‘A PANIC attack on unit roots and cointegration’, *Econometrica,* 72, 1127-1177.

Bai, J. and Ng, S. (2006). ‘Evaluating latent and observed factors in macroeconomics and finance’, *Journal of Econometrics,* 131, 507-537.

Baier, S. and Bergstrand, J. (2009). ‘Bonus vetus OLS: A simple approach for addressing the border-puzzle and other gravity issues’, *Journal of International Economics*, 77, 77-85.

Baldwin, R. and Di Nino, V. (2006). ‘Euros and zeros: The common currency effect on trade in new goods’, Discussion Paper No. 5973, CEPR.

Baldwin, R., Skuderlny F. and Taglioni, D. (2005). ‘Trade Effects of the euro: Evidence from Sectoral Data’, Working Paper No. 446, ECB.

Baldwin, R. and Taglioni, D. (2006). ‘Gravity for dummies and dummies for gravity equations’, Working Paper No. 12516, NBER.

Baldwin, R., Di Nino, V., Fontagné, L., De Santis, R. and Taglioni, D. (2008). ‘Study on the impact of the euro on trade and Foreign Direct Investment’, Economic Papers No. 321, European Economy.

Baltagi, B.H., Egger, P. and Pfaffermayr, M. (2003). 'A generalized design for bilateral trade flow models', *Economic Letters*, 80, 391-397.

Banerjee, A. and Carrion-i-Silvestre, J.L. (2010). 'Cointegration in panel data with breaks and cross-section dependence', mimeo, previously published as Working Paper Series No. 591, ECB.

Banerjee, A., Marcellino, M. and Osbat, C. (2004). 'Some cautions on the use of panel methods for integrated series of macro-economic data', *Econometrics Journal*, 7, 322-340.

Barr, D., Breedon, F. and Miles, D. (2003). 'Life on the Outside: Economic Conditions and Prospects Outside Euroland', *Economic Policy*, 37, 573-613.

Belke, A. and Spies, J. (2008) 'Enlarging the EMU to the east: what effects on trade?' *Empirica*, 35, 369-389.

Bergeijk, P. and Oldersma, H. (1990). 'Detente, market-oriented reform and German unification: Potential consequences for the World Trade System', *Kyklos,* 43, 599-609.

Berger, H. and Nitsch, V. (2008). 'Zooming out: The trade effect of the Euro in historical perspective', *Journal of International Money and Finance*, 27, 1244-1260.

Bergstrand, J.H. (1985). 'The gravity equation in international trade: some microeconomic foundations and empirical evidence', *Review of Economics and Statistics*, 67, 474-481.

Bergstrand, J.H. (1989). 'The generalized gravity equation, monopolistic competition, and the factor-proportions theory in international trade', *Review of Economics and Statistics*, 71, 143-153.

Bergstrand, J.H. (1990). 'The Heckscher-Ohlin-Samuelson model, the Linder hypothesis and the determinants of bilateral intra-industry trade', *The Economic Journal*, 100, 1216-1229.

Bernard A., Eaton J., Jensen B., and Kortum, S. (2003). 'Plants and productivity in international trade', *American Economic Review*, 93, 1268-1290.

Berthou, A. and Fontagné, L. (2008). 'The euro and the intensive and extensive margins of trade: Evidence from French firm level data', Working Paper No 2008-06, CEPII.

Bikker, A. and de Vos, A. (1992). 'An international trade flow model with zero observations: an extension of the Tobit model', *Brussels Economic Review*, 135, 379-404.

Bouet, A., Mishra, S. and Roy, D. (2008). 'Does Africa trade less than it should, and if so, why?: The role of market access and domestic factors', Discussion Paper No 770, International Food Policy Research.

Breitung, J. and Pesaran, M.H., (2008). 'Unit roots and cointegration in panels', in *The Econometrics of Panel data: Fundamentals and Recent Developments in Theory and Practice,* Matyas, L. and Sevestre, P. (ed), Springer.

Brouwer, J., Paap, R. and Viaene, J.M. (2008). 'The trade and FDI effects of EMU enlargement', *Journal of International Money and Finance*, 27, 188-208.

Brun, J.F., Carrère, C., Guillaumont, P. and de Melo, J. (2002). 'Has distance died? Evidence from a panel gravity model', *World Bank Economic Review*, 19, 99-120.

Bun, M.J. and Klaassen, F.J. (2002). 'Has the euro increased trade?' Discussion Paper No. 02-108/2, Tinbergen Institute.

Bun, M.J. and Klaassen, F.J. (2007). 'The euro Effect on trade is not as large as commonly thought', *Oxford Bulletin of Economics and Statistics*, 69, 473-496.

Burger, M.J., van Oort, F.G. and Linders, G.M. (2009). 'On the specification of the gravity model of trade: zeros, excess zeros and zero-inflated estimation', *Spatial Economic Analysis*, 4, 167-190.

Bussière, M., Fidrmuc, J. and Schnatz, B. (2008). 'EU Enlargement and Trade Integration: Lessons from a Gravity Model', *Review of Development Economics*, 12, 562-576.

Bussière, M. and Schnatz, B. (2009). 'Evaluating China's integration in world trade with a gravity model based benchmark', *Open Economies Review*, 20, 85-111.

Cafiso, G. (2008). 'The euro's influence upon trade: Rose effect versus Border effect', Working Paper No. 941, ECB.

Camarero, M., Gómez, E., Tamarit, C. 2011. 'Is there a "euro effect" on trade? New evidence using gravity equations with panel cointegration techniques', mimeo.

Cieslik, A., Michałek, J.J. and Mycielski, J. (2009). 'Trade effects of the euro adoption in Central and Eastern Europe' European Trade Study Group Annual Conference, Rome, 12-14 September 2009.

Clark, P., Tamirisa, N. and Wei, S.J., (2004). 'A New Look at Exchange Rate Volatility and Trade Flows', Occasional Paper No. 235, IMF.

Cheng, I.H. and Wall, H.J. (2005). 'Controlling for heterogeneity in gravity models of trade and integration', *Federal Reserve Bank of St. Louis Review*, 87, 49-63.

Deardorff, A. (1998). 'Determinants of bilateral trade: Does gravity work in a neoclassical world?', in *The Regionalization of the World Economy,* Frankel J.A. (ed), University of Chicago Press.

De Benedictis, L. and Taglioni, D. (2011). 'The Gravity Model in International Trade', in *The trade impact of European Union Preferential policies: An analysis through gravity models*, De Benedictis, L. and Salvatici, L. (ed), Springer.

De Benedictis, L. and Vicarelli, C. (2005). 'Trade Potentials in Gravity Panel Data Models', *Top Economic Analysis Policy,* 5, 1386-1417.

De Grauwe, P. and Skuderlny, F. (2000). 'Inflation and Productivity Differentials in EMU', Discussion Papers No. 0015, Center for Economic Studies, Katholieke Universiteit Leuven.

De Nardis, S. (2004). 'Currency unions and trade: The special case of EMU', *Review of World Economics*, 140, 625-649.

De Nardis, S. and Vicarelli, C. (2003). 'The Impact of euro on trade: The (early) effect is not so large', Working Paper No. 31, Institute for Studies and Economic Analyses.

De Nardis, S., De Santis, R. and Vicarelli, C. (2008). The euro's Effects on trade in a dynamic setting, *European Journal of Comparative Economics*, 5, 73-85.

Dell'Ariccia, G. (1999). 'Exchange Rate Fluctuations and Trade Flows: Evidence from the European Union', *IMF Staff Papers*, 46, 315-334.

Dwane, C., Lane, P. and McIndoe, T. (2011). 'Currency Unions and Irish external trade', *Applied Economics*, 42, 2393-2397.

Egger, P. (2000). 'A Note on the proper econometric specification of the gravity equation', *Economic Letters*, 66, 25-31.

Egger, P. and Pfaffermayr, M. (2003). 'The proper panel econometric specification of the gravity equation: A three-way model with bilateral interaction effects', *Empirical Economics*, 28, 571-580.

Egger, P. and Pfaffermayr, M. (2004). 'Distance, trade and FDI: A SUR Hausman-Taylor approach', *Journal of Applied Econometrics*, Vol. 19, 227-246.

Egger, P., Larch, M., Staub, K. and Winkelmann, R. (2011). 'The Trade Effects of Endogenous Preferential Trade Agreements', *American Economic Journal: Economic Policy*, 3, 113-143.

Eicher, T.S. and Henn, C. (2011). 'One money, one market. A revised benchmark', *Review of International Economics*, 19, 419-435.

Emerson, M., Gros, D., Italianer, A., Pisany-Ferry, J., and Reichenbach, H. (1992). 'One Market, One Money: An Evaluation of the Potential Benefits and Costs of Forming an Economic and Monetary Union, Oxford University Press.

Esteve-Pérez, S., Gil-Pareja, S., Llorca-Vivero, R. and Martínez-Serrano, J.A. (2011) 'The impact of the euro on firm export behaviour: does firm size matter?', *Oxford Economic Papers*, 63, 355-374.

European Comission (2010). 'Surveillance of intra-euro-area competitiveness and imbalances', European Economy 1/2010. ISSN 0379-0991.

Faruqee, H. (2004). 'Measuring the trade effects of EMU', Working Paper No. 04/154, IMF.

Feenstra, R.C. (2002). 'Border effects and the gravity equation: consistent methods for estimation', *Scottish Journal of Political Economy*, 49, 491-506.

Feenstra, R., Markusen, J. and Rose, A.K. (2001). 'Using the gravity equation to differentiate among alternatives theories of trade', *Canadian Journal of Economics*, 34, 430-447.

Fidrmuc, J. (2009). 'Gravity models in integrated panels', *Empirical Economics*, 37, 435-446.

Flam, H. and Nordstrom, H. (2006). 'Euro effects on Intensive and extensive margins of trade', Working Paper No. 1881, CESifo.

Flam, H. (2009). 'The Impact of the Euro on Trade and Investment: A Survey of Theoretical and Empirical Evidence', SIEPS Report 2009:8.

Fratianni, M. and Hoon-Oh, C.H. (2007). 'On the relationship between RTA expansion and openness', Working Paper No. 13, Kelley School of Business.

Frankel, J.A., Stein, E. and Wei, S.J. (1995). 'Trading blocs and the Americas: The natural, the unnatural, and the super-natural', *Journal of Development Economics*, 47, 61-95.

Frankel, J.A. and Rose, A.K. (1998). 'The endogeneity of the Optimum Currency Area criteria', *Economic Journal,* 108, 1009-1025.

Frankel, J.A. and Rose, A.K. (2002). 'An estimate of the effect of common currencies on trade and income', *Quarterly Journal of Economics*, 117, 437-466.

Fratianni, M., and Oh, C. H. (2007). 'On the Relationship Between RTA Expansion and Openness', Working Papers No. 13, Kelley School of Business.

Gengenbach, C. (2009). 'A panel cointegration study of the Euro effect on trade', mimeo.

Gil-Pareja, S., Llorca-Vivero, R., Martínez-Serrano, J. and Oliver-Alonso, J. (2005). 'The Border Effect in Spain', *The World Economy*, 28, 1617-1631.

Gil-Pareja, S., Llorca-Vivero, R. and Martínez-Serrano, J. (2008). 'Trade effects of monetary agreements: evidence for OECD countries', *European Economic Review* 52, 733-755.

Glick, R. and Rose, A.K. (2002). 'Does a currency union affect trade? The time-series evidence', *European Economic Review*, 46, 1125-1151.

Gomes, T., Graham, Ch., Helliwell, J., Kano, T., Murray, J. and Schembri, L. (2006). 'The euro and trade: Is there a positive effect?', mimeo.

Gregory, A.W. and Hansen, B.E. (1996). 'Residual-based tests for cointegration in models with regime shifts', *Journal of Econometrics*, 70, 99-126.

Greenaway, D., and Milner, C. (2002). 'Regionalism and gravity', *Scottish Journal of Political Economy*, 49, 574-585.

Hadri, K. (2000). ‘Testing for stationarity in heterogeneous panel data’, *Econometrics Journal*, 3, 148-161.

Haq, Z., Meilke, K. and Cranfield, J. (2010). ‘Does the gravity model suffer from selection bias?’, Working Paper No. 90884, Canadian Agricultural Trade Policy Research Network.

Head, K. and Mayer, T. (2000). ‘Non-Europe: The magnitude and causes of market fragmentation in the EU’, *Review of World Economies*, 127, 284-314.

Heckman, J.J. (1979). ‘Sample selection bias as a specification error’, *Econometrica*, 47, 153-161.

Helliwell, J.F. (1996). ‘Do national borders matter for Quebec’s trade?’, *Canadian Journal of Economics*, 29, 507-522.

Helpman, E. (1987). ‘Imperfect competition and international trade: evidence from fourteen industrial countries’, *Journal of Japanese and International Economies*, 1, 62-81.

Helpman, E., Melitz, M. and Rubinstein, Y. (2008). ‘Estimating trade flows: Trading partners and trading volumes’, *Quarterly Journal of Economics*, 123, 441-487.

Helpman E., Melitz M., Yeaple, S. (2004). ‘Export versus FDI with heterogeneous firms’, *American Economic Review*, 94, 300-316.

Henderson, D.J. and Millimet, D.L. (2008). ‘Is gravity linear?’, *Journal of Applied Econometrics*, 23, 137-172.

Hiller, S and Kruse, R. (2010). ‘Milestones of European Integration: Which matters most for Export Openness?’ Research Paper No. 27, School of Economics and Management, University of Aarhus.

Hochreiter, E. and Tavlas, G. (2004). ‘On the Road Again: An Essay on the Optimal Path to EMU for the New Member States’, *Journal of Economic Policy Modeling*, 26, 793-816.

Hoon-Oh, CH. and Selmier II, W.T. (2008). ‘Expanding international trade beyond the RTA border: The case of ASEAN's economic diplomacy’, *Economic Letters*, 100, 385-387.

Hooper, P., Johnson, K., and Marquez, J. (1998). ‘Trade elasticities for G-7 countries’, International Finance Discussion Paper No. 609, Board of Governors of the Federal Reserve System.

Im, K., Pesaran M.H. and Shin, Y. (2003). ‘Testing for unit roots in heterogeneous panels’, *Journal of Econometrics,* 115, 53-74.

Jang, M.J. and Shin, D.W. (2005). 'Comparison of panel unit root tests under cross-sectional dependence', *Economics Letters,* 89, 12-17.

Kao, C. (1999). 'Spurious regression and residual-based tests for cointegration in panel data', *Journal of Econometrics*, 90, 1-44.

Kao, C. and Chiang, M.H. (2000). 'On the estimation and inference of a cointegrated regression in panel data', *Advances in Econometrics*, 15, 179-222.

Kavallari, A., Maas, S. and Schmitz, P.M. (2008). 'Explaining German imports of olive oil: evidence from a gravity model', European Association of Agricultural Economists International Congress, August 26-29, 2008, Ghent, Belgium.

Kelejian, H., Tavlas, G.S. and Petroulas, P. (2011). 'In the neighborhood: The trade effects of the euro in a spatial framework', *Regional Science and Urban Economics,* 42, 314-322.

Leamer, E. and Stern, R. (1971). 'Quantitative international economics', *Journal of International Economics* 1, 359-361.

Levin, A., Lin, C. and Chu, C. (2002). 'Unit root tests in panel data: Asymptotic and finite-sample properties', *Journal of Econometrics*, 108, 1-24.

Linnemann, H. (1966). 'An Econometric study of international trade flows', Dissertation, Netherlands School of Economics.

Linders, G.M. and de Groot, H. (2006). 'Estimation of the gravity equation in the presence of zero flows', Discussion Paper No. 06-072/3, Tinbergen Institute.

Liu, X. (2009). 'GATT/WTO promotes trade strongly: Sample selection and model specification', *Review of International Economics*, 17, 428-446.

McCallum, J. (1995) 'National Borders Matter: Canada-U.S. Regional Trade Patterns', *American Economic Review*, 85, 615-23.

Maliszewska, M.A. (2004). 'New Member States Trading Potential Following EMU Accession: A Gravity Approach'. Studies and Analysis No. 286, CASE – Center for Social and Economic Research.

Mancini-Griffoli, T. and Pauwels, L. (2006). 'Is There a euro effect on trade? An application of end-of-sample structural break tests for panel data'. Working Paper No. 04, The Graduate Institute of International Studies.

Manning, W.G. and Mulahy, J. (2001). 'Estimating log models: to transform or not to transform?', *Journal of Health Economics*, 20, 461-494.

Martin, W. and Pham, C.S. (2008). 'Estimating the gravity equation when zero trade flows are frequent', Working Paper No. 9453, MPRA.

Martínez-Zarzoso, I., and Nowak-Lehman, F. (2003). 'Augmented Gravity Model: An Empirical Application to Mercosur-European Union Trade Flows', *Journal of Applied Economics*, 6, 291-316.

Martínez-Zarzoso, I. (2011). 'The Log of Gravity Revisited', *Applied Economics*, 45, 311-327.

Mátyás, L. (1998). 'The gravity model: Some econometric considerations', *The World Economy*, 21, 397-401.

McCoskey, S. and Kao, C. (1998). 'A residual-based test of the null of cointegration in panel data', *Econometric Reviews*, 17, 57-84.

McKenzie, M. (1999). 'The Impact of Exchange Rate Volatility on International Trade Flows', *Journal of Economic Surveys*, 13, 71-106.

Melitz, M. (2003). 'The impact of trade on intra-industry reallocations and aggregate industry productivity', *Econometrica*, 71, 1695-1725.

Melitz, M. (2007). 'North, South and distance in the gravity equation', *European Economic Review*, 51, 971-991.

Micco, A., Stein, E. and Ordoñez, G. (2003). 'The currency union effect on trade: early evidence from EMU', *Economic Policy*, 18, 315-356.

Moon, H.R. and Perron, B. (2004). 'Testing for a unit root in panels with dynamic factors', *Journal of Econometrics*, 122, 81-126.

Mukherjee, D. and Pozo, S. (2011). 'Exchange-rate volatility and trade: a semiparametric approach', *Applied Economics*, 43, 1617-1627.

Oh, C. H., and Selmier II, W. T. (2008). 'Expanding international trade beyond the RTA border: The case of ASEAN's economic diplomacy', *Economics Letters*, 100, 385-387.

O'Connell, P. (1998). 'The overvaluation of purchasing power parity', *Journal of International Economics*, 44, 1-19.

Pedroni, P. (1999). 'Critical values for cointegration tests in heterogeneous panels with multiple regressors', *Oxford Bulletin of Economics and Statistics*, 61, 653-678.

Pedroni, P. (2004). 'Panel cointegration; Asymptotic and finite sample properties of pooled time series tests with an application to the Purchasing Power Parity hypothesis', *Econometric Theory*, 20, 597-625.

Pesaran, M.H. (2004). 'General diagnostic tests for cross section dependence in panels', Working Papers in Economics No. 435, University of Cambridge.

Pesaran, M.H. (2006). 'Estimation and inference in Large Heterogeneous Panels with a Multifactor Error Structure', *Econometrica*, 74, 967-1012.

Pesaran, M.H. (2007). 'A simple panel unit root test in the presence of cross section dependence', *Journal of Applied Econometrics*, 22, 265-312.

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

Phillips, P.C. and Sul, D. (2003). 'Dynamic panel estimation and homogeneity testing under cross-section dependence', *Econometrics Journal*, 6, 217-260.

Rose, A.K. (2000). 'One money, one market: the effect of common currencies on trade', *Economic Policy*, 15, 7-46.

Rose, A.K. (2008). 'Is EMU becoming an Optimum Currency Area? The evidence on trade and business cycle synchronization', mimeo.

Rose, A. and Engel, C. (2002). 'Currency Unions and International Integration', *Journal of Money, Credit and Banking*, 34, 1067-1089.

Rose, A.K. and Stanley, T.D. (2005). 'A Meta-Analysis of the Effect of Common Currencies on International Trade', *Journal of Economic Surveys*, 19, 347-365.

Rose A.K. and van Wincoop, E. (2001). 'National money as a barrier to international trade: The real case for currency union', *American Economic Review*, 91, 386-390.

Ruiz, J. and Vilarrubia, J.M. (2007). 'The wise use of dummies in gravity models: export potentials in the Euromed region', Working Paper No. 0720, Bank of Spain.

Saikkonnen, P. (1991). 'Asymptotically efficient estimation of cointegration regressions', *Econometric Theory*, 7, 1-21.

Santos-Silva J.M. and Tenreyro, S. (2006). 'The Log of Gravity', *Review of Economics and Statistics*, 88, 641-658.

Santos-Silva, J.M. and Tenreyro, S. (2008). 'Trading partners and trading volume: Implementing the Helpman-Melitz-Rubinstein model empirically', Discussion Paper No. 0935, CEPR.

Sapir, A. (2001). 'Domino effects in Western European Regional Trade', *European Journal of Political Economy*, 17, 377-388.

Sauer, C., and Bohara, A. K. (2001). 'Exchange rate volatility and export: regional differences between developing and industrialized countries', *Review of International Economics*, 9, 133-152.

Schiavo, S. (2007). 'Common currencies and FDI flows', *Oxford Economic Papers*, 59, 536–560.

Shepherd, B. and Wilson, J.S. (2009). 'Trade facilitation in ASEAN member countries: Measuring progress and assessing priorities', *Journal of Asian Economics*, 20, 367-383.

Siliverstovs, B. and Schumacher, D. (2009). 'Estimating gravity equations: to log or not to log?', *Empirical Economics*, 36, 645-669.

Shepotylo, O. (2009). 'Gravity with zeros: estimating trade potential of CIS countries', Discussion Paper No. 16, Kyiv School of Economics.

Serlenga, L. and Shin, Y. (2007). 'Gravity models of intra-EU trade: application of the CCEP-HT estimation in heterogeneous panels with unobserved common time-specific factors', *Journal of Applied Economics*, 22, 361-381.

Soloaga, I. and Winters, A. (2001). 'Regionalism in the nineties: what effect on trade?', *North American Journal of Economics and Finance*, 12, 1-29.

Spies, J. and Marques, H. (2009). 'Trade effects of the Europe agreements: A theory-based gravity approach', *Journal of International Trade and Economic Development*, 18, 11-35.

Stock, J. and Watson, M. (1993). 'A simple estimator of cointegrating vectors in higher order integrated systems', *Econometrica*, 61, 783-820.

Tenreyro, S. (2007). 'On the Trade Impact of Nominal Exchange Rate Volatility', *Journal of Development Economics*, 82, 485-508.

Wang, Z.K. and Winters, L.A. (1991). 'The trading potential of Eastern Europe', Discussion Paper No. 610, CEPR.

Wei, S.J. (1996). 'Intra-national versus international trade: How stubborn are nations in global integration?', Working Paper No. 5531, NBER.

Westerlund, J. and Wilhelmsson, F. (2009). 'Estimating the gravity model without gravity using panel data', *Applied Economics*, 43, 641-649.

Wolf, H.C. (1997). 'Patterns of intra- and inter-state trade', Working Paper No. 5939, NBER.

Yu, M. (2010). 'Trade, democracy, and the gravity equation', *Journal of Development Economics*, 91, 289-300.