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The Euro effect on trade: evidence in gravity equations using panel cointegration techniques*

Estrella Gómez and Cecilio Tamarit**

Abstract

In this paper we present new evidence on the effect of the Euro on trade. We use a data set containing all bilateral combinations in a panel of 26 countries covering the period 1967-2008. We estimate the equation using two sets of variables: a standard one and a second one built according to the criticisms stated by Baldwin and Taglioni (2006). We implement a new generation of tests that allow us to solve some of the problems derived from the non-stationary nature of the data usually present in the macroeconomic variables used in gravitational equations (GDP, trade). To this aim we use some panel tests that account for the presence of cross-section dependence as well as discontinuities in the non-stationary panel data series. We test for cointegration between the variables using panel cointegration tests, especially the ones proposed by Banerjee and Carrión-i-Silvestre (2004, 2010). We also efficiently estimate the long-run relationships using the CUP-BC and CUP-FM estimators proposed in Bai et al. (2009). The results obtained are in line with those of Bun and Klaassen (2007). We argue that the creation of the European Monetary Union is best interpreted as a culmination of a series of policy changes that have been increasing economic integration in Europe during over four decades.

Keywords: gravity models; trade; panel cointegration; common factors; structural breaks, cross-section dependence.

JEL classification: C12, C22, F15, F10.

Resumen

En este trabajo presentamos nueva evidencia del efecto del Euro sobre el comercio. Para ello utilizamos una base de datos que contiene todas las combinaciones bilaterales en un panel de 26 países para el periodo 1967-2008. Estimamos la ecuación de gravedad usando dos tipos de variables: la estándar en la literatura y la que recoge las críticas de Baldwin y Taglioni (2006), aplicando una nueva generación de contrastes que nos permiten resolver los principales problemas derivados de la naturaleza no estacionaria de las series. Con este propósito utilizamos algunos contrastes de panel que tienen en cuenta la presencia de dependencia cross-section así como de rupturas en las series. Para realizar el análisis de cointegración cabe destacar el uso del contraste de Banerjee y Carrion-i-Silvestre (2006, 2010). También estimamos de forma eficiente las relaciones de largo plazo mediante los estimadores CUPpc y CUPfm propuestos por Bai et al. (2009). Los resultados obtenidos están en línea con los de Bun y Klaassen (2007). Nuestro argumento es que la creación de la Unión Monteria Europea se debe interpretar como la culminación de un conjunto de cambios de política que han ido dando lugar a un proceso de integración económica en Europa durante las últimas cuatro décadas.

Palabras clave: modelos de gravedad; comercio; cointegración con paneles; factores comunes; cambios estructurales; dependencia.

Clasificación JEL: C12, C22, F15, F10.

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** E. Gómez: Universidad de Granada. C. Tamarit: Universidad de Valencia. Corresponding author: C. Tamarit, e-mail: cecilio.tamarit@uv.es.

1. Introduction

Gravity equations are a widely used tool in the empirical literature on trade determinants. They relate bilateral trade flows to country-specific characteristics of the exporters and importers such as economic size, and to bilateral characteristics such as trade frictions between the trading partners. An increasing number of papers have flourished trying to improve the robustness of the results either through new modelling techniques or using competing specifications. A part of these studies has been devoted to understanding the impact of trade frictions on international trade analyzing in a very detailed manner the impact of distance and geography, free trade agreements, WTO membership, and more recently, the effect of currency unions (CU) on trade .

The introduction of the euro has raised a new interest in measuring the impact of CU on trade flows. The argument that a CU promotes greater trade among members depends on the fulfilment of different potential benefits, namely, lower transaction costs, reduced exchange rate uncertainty, enhanced competition as well as more credibility and reputation of the macro policies. Many economists claim that the Euroarea was not an Optimum Currency Area (OCA) at the time of its creation, and therefore, that some degree of exchange rate flexibility could be a useful tool when facing prospective asymmetric shocks. However, 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 OCA, that means that even if the European Monetary Union (EMU) was not created as an OCA, it is moving 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 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 EMU.

In 1999 eleven countries of the EU adopted the euro as a common currency. Greece joined the EMU in 2001. Since then, also Slovenia, Cyprus, Malta, Slovakia and Estonia have joined the Euro 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 the 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) exchange rate stability. As pointed out by Faruquee (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 the trade effects through time, and finally, to measure the distribution of trade effects among member states.

Currently, 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. 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 Faruquee (2004), is at odds with the related literature that typically finds very little negative impact of exchange rate volatility on trade. Not surprisingly, the findings by Rose (2000) 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's (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 the 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 more relevant issue for EMU is about the possible change in the level of trade for member states over time, before and after the introduction of the single currency. In order to solve this second flaw, Glick and Rose (2002) and Frankel and Rose (2002) exploited the time series information using panel data. They obtained similar results[1] giving birth to a literature in search of "more reasonable" effects (Eicher and Henn, 2009)[2]. 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 of 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%). All in all, 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%[3], 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.

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 leading to

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[1] They found that adopting a monetary union doubles bilateral trade but, again EMU countries were not included.

[2] Quoting Eicher and Henn (2009): "Rose's estimate sparked a controversy out of which emerged an entire literature attempting to shrink the Rose effect".

[3] For a recent survey of the empirical literature, see Gómez and Milgram (2010).

a much reduced estimate of the Euro effect (3%) on bilateral trade compared to previous findings by other authors[4]. All in all, they employed methods that assume cross-section independence. The latter 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. the business cycles of other economies). Moreover, the dependency 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 unit root and cointegration tests have been proposed accounting for cross-sectional dependence in the form of common factors. See for example Breitung and Pesaran (2008) for an overview of the literature and Gengenbach et al (2009) for a comparison of panel unit root tests.

Therefore, in this paper we try to provide new evidence on the effect of the Euro using a data set that contains data on all bilateral combinations in a panel of 26 countries covering the period 1967-2008. We implement a new generation of tests that allow us to solve some of the problems derived from the non-stationary nature of the data usually present in macroeconomic variables used in gravitational equations (GDP, trade...). To this aim we use some panel tests that account for the presence of cross-section dependence as well as discontinuities in the non-stationary panel data series. More specifically, we implement the panel unit root and stationary tests proposed by Im, Pesaran and Shin (2003), Hadri (2000), Pesaran (2007) and Bai and Ng (2004) to test whether the variables entering the gravity model are non-stationary. We then test for cointegration between the variables using panel cointegration tests, with a special emphasis in the one proposed by Banerjee and Carrión-i-Silvestre (2010). Finally, the coefficients are efficiently estimated through the continuously updated estimator (CUP) of Bai et al. (2009). 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 is not accounted for .

The paper is organized as follows. Section 2 discusses the empirical literature on CU and trade, emphasizing the econometric approaches based on the gravity model. Section 3 presents a new econometric approach that overcomes some of the present problems in the current literature. Section 4 describes the data and discusses the empirical results. A final section concludes.

[4] Other papers that stress the importance of the non stationary nature of the series and that apply cointegration techniques are Faruquee (2004) and Fidrmuc (2009).

2. Methodological approaches to measuring the Euro effect on trade: a synthesis of previous studies and criticisms to the empirical application of the gravity equation

The gravity model has become very popular due to its success in explaining trade flows among countries. The gravity model, as Baldwin and Taglioni (2006) point out, has become a workhorse tool in many empirical fields related to trade. Tinbergen (1962) was the first to introduce the gravity equation for trade and although in the beginning the gravity model was criticized for its lack of theoretical underpinnings, now rests on a solid theoretical background. There are various theoretical foundations for gravity equations[5]. Anderson (1979) was the first to provide a microfoundation to the equation, whereas during the eighties, Bergstrand (1985) founded it in the traditional trade theory (the connection between factor endowment and bilateral trade). In the late 1970s and the 80s the emergence of the "New Trade Theory" also provided theoretical foundations to the gravity equation. Recent trade models that predict gravity equations in equilibrium include the Ricardian framework (Eaton and Kortum, 2002), the multilateral resistance framework (Anderson and van Wincoop, 2003), the heterogeneous firms models (Chaney, 2008) and the Heckscher-Ohlin framework (Deardorff, 1998). Therefore, as stated in Westerlund and Wilhelmsson (2006) 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 feed-back 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 due to the fact that heterogeneity among countries was not properly controlled for.

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[6]. The most popular approach for the estimation of the gravity model using panel data is to first make it linear by taking logarithms and then to estimate the resulting log-linear model by the fixed effects ordinary least squares (OLS). However, Santos-Silva and Tenreyro (2006) argue that the standard empirical methods used to estimate the gravity equation are inappropriate. According to them,

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[5] See, for instance, Feenstra, Markusen and Rose (2001)

[6] See Baldwin (1994), Egger (2002) or Nilson (2000). Moreover, as clearly explained by Westerlund and Wilhelmsson (2009), if we desire to measure the impact of a currency union on trade (which is the relevant case in this paper), 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.

two are the problems related to the practice of log-linearization. First, the empirical model in the presence of heteroskedasticity leads to both biased and inconsistent estimates. Second, log-linearization is incompatible with the existence of zeros in trade data. These problems related to the OLS estimation have been largely ignored by applied researchers as the methods more commonly used to solve them are not easy to implement.[7] More recently, Santos-Silva and Tenreyro (2006) propose an alternative estimation technique, the Poisson pseudo-maximum likelihood method that is robust to different patterns of heteroskedasticity and provides a natural way to deal with zeros in trade data. Westerlund and Wilhelmsson (2009) also study the effects of zero trade in the estimation of the gravity model. They propose a very similar alternative: estimating the model directly from its non-linear form using the fixed effects Poisson ML estimator with bootstrapped standard error.

While the heterogeneity bias is controlled through the use of fixed-effects, a second kind of misspecification is related to dynamics. Both elements can be tackled using the system-GMM estimator proposed by Arellano and Bover (1995). The recent theoretical literature on international trade with heterogenous 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, there is now substantial evidence that the level of productivity of exporting firms is generally higher than that of non-exporting firms. The explanation lies in a self-selection mechanism, due to sunk costs associated with entry into foreign markets (Roberts and Tybout, 1997; Eichengreen and Irwin, 1997). Thus, 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.

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 Faruquee (2004) claimed, estimating the impact of a monetary union on trade faces several econometric challenges[8]. 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

[7] We can add a constant arbitrarily small to each observation on the dependent variable or just discard the zeros. The latter can create a sample selection problem as long as the zeros are not randomly distributed.

[8] As recently claimed by Fidrmuc (2009) standard gravity models include non-stationary variables characterized by cross-sectional correlations between country pairs, and as a result of that, standard panel unit root and cointegration tests are biased.

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 considering 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.

Moreover, following Baldwin and Taglioni (2008) we have taken into account the most common empirical errors in the estimation of the euro effect in previous literature when we have defined the variables of interest. These errors were coined as the "gold", "silver" and "bronze" medal errors by Baldwin (2006) and their effect was assessed later in Baldwin and Taglioni (2006). The gold medal is also called the "Anderson-van Wincoop (A-vW) misinterpretation" in the sense that A-vW developed a cross-section technique estimation to control for omitted variables with pair fixed effects. 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, Flam and Nordstrom, 2003, Gil-Pareja et al., 2008). 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.

Secondly, according to Baldwin and Taglioni (2008) the silver medal error arises when authors use the log of the sum instead of the sum of the logs in the bilateral trade term. The silver medal mistake will create no bias if bilateral 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, then 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 deflation: 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). As Baldwin and Taglioni (2008) claim, fortunately, the bronze medal bias is eliminated by including time dummies, which is the common practice.

The contribution of our paper to the existing literature about the Euro effect on trade is twofold. First, unlike previous research, (excepting Eicher and Henn, 2009) we address Baldwin's critiques regarding the proper specification of gravity models and the definition of the variables. 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 and cross-section dependence.

3. A new methodological proposal: using panel cointegration tests that allow for dependence and structural breaks

Although, in theory, the above-mentioned cointegration techniques provide a well-developed econometric framework to properly estimate gravity equations, some practical problems implicitly implied that most of the evidence obtained so far did not consider the non-stationary nature of the series. However, new developments in macro-econometrics allow estimating gravity models among countries using a variety of panel data tests based on the theory of cointegration.

First, as for the unit root tests, the pioneer research by Levin *et al.* (2002) and Im *et al.* (2003) proposed different versions of unit root tests in a panel setting, whereas Hadri (2000) built stationarity tests in panels. By adding this new dimension we increase the amount of information for each cross-section thus solving the problems related to the lack of power of univariate unit root tests when the root is close to one, especially in small samples[9] when the time dimension is restricted by the lack of availability of long and reliable time series data. Moreover, using shorter samples with rich information helps us to avoid a second serious problem arising from the fact that standard unit root tests are biased towards the non-rejection of the null hypothesis in the presence of structural breaks. Obviously, as we reduce the sample length, the probability of discontinuities in the series generated either by shocks or by institutional changes diminishes.

Although this first generation of tests is still being extensively used in the empirical literature, the first main drawback (common to all them) is that they assume the absence of correlation across the cross-sections of the panel. That is, the individual members of the panel (countries) are 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. O'Connell (1998) argued that the assumption of independence would bias the unit root tests in favour of variance stationarity. A second generation of panel tests, in contrast, allows for different forms of dependence (see Pesaran, 2003, 2004), solving the above- mentioned problem.

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[9] See Shiller and Perron (1985)

There are several alternative proposals formulated in the literature to overcome the cross-section dependency problem. First, Levin et al. (2002) suggest computing the test removing the cross-section mean. Although simple, this implies assuming, quite restrictively, that cross-section dependence is driven by one common factor with the same effect for all individuals in the panel data set. Second, Maddala and Wu (1999) propose obtaining the bootstrap distribution to accommodate general forms of cross-section dependence. Third, Breuer et al. (2002) also propose a panel unit root test that allows for contemporaneous correlation among the errors. Separate null and alternative hypotheses are tested for each panel member using the information captured through the variance-covariance matrix in a system estimated within a SUR framework and the critical values are obtained by bootstrap methods. Fourth, more recently, Pesaran (2007), Phillips and Sul (2003), Moon and Perron (2004) and Bai and Ng (2004) have suggested other proposals that are especially relevant when the dependence is pervasive, which is the most common case for integrated markets. They assume that the process is driven by a group of common factors, so that it is possible to distinguish between the idiosyncratic component and the common component. Although there are differences among the methods proposed, their driving idea is similar. Pesaran (2007), Phillips and Sul (2003), Moon and Perron (2004) focus on the extraction of the common factors that generate the cross correlations in the panel to assess the non-stationarity of the series, while in Bai and Ng (2004) the non-stationarity of the series can come either from the common factors, the idiosyncratic component or from both. Moreover, Pesaran (2007) and Phillips and Sul (2003) only consider the existence of one common factor, while in Moon and Perron (2004) and Bai and Ng (2006) there can be multiple common factors. Finally, Bai and Ng (2004) consider also the possibility 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. Moreover, Monte Carlo comparisons developed by Gengenbach et al. (2004) 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). Consequently, our choice to account for the existence of cross-section dependency is based on the Bai and Ng's (2004) approach.

A second caveat appears when there are discontinuities 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 affected by misspecification errors. Therefore, in this paper we propose the use of the tests developed in Banerjee and Carrion-i-Silvestre (2004, 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 the following econometric issues usually neglected in earlier literature: first, we account for cross-section dependence among countries in the panel tests. Second, we allow for the existence of a break in the cointegration relationship, a major point to assess the effect of institutional changes in the relationship. 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, we estimate efficiently the coefficients of the long-run relationships using the CUP FM estimators by Bai et al. (2009).

4. Data and empirical results

Our dataset contains annual data from 26 OECD countries and covers the period 1967-2008. The total number of country pairs or combinations from a sample of 26 countries is C(26, 2) = 325. Hence, we have a balanced panel with cross-section dimension N=325 and a time series span of T = 42, yielding a total number of observations NT = 13,650.

The dataset (see Appendix 1) includes the following variables, where upper case letters stand for nominal variables while lower case letters stand for variables in real terms. $TRADE_{ijt}$ is the log of the bilateral trade in goods between trading partners *i* and *j* at time t, defined as the *sum of the logs* of nominal imports and exports. Data for nominal imports and exports are obtained from the CHELEM – CEPII database, and are expressed in current dollars. $Trade_{ijt}$ stands for real bilateral trade, calculated as the sum of the logs of nominal bilateral exports and imports in US dollars deflated using the US CPI obtained from the IMF International Financial Statistics (IFS); gdp_{ijt} is the log of the product of bilateral real PPP-converted GDP in country *i* and *j* and GDP_{ij} is the log of the product of bilateral nominal GDP. Both are obtained from CHELEM-CEPII database. $GDPCAP_{ijt}$ (and $gdpcap_{ijt}$) measure the log of the product of countries' nominal (and real) GDP per capita, respectively. Population data used to construct $GDPCAP_{ij}$ are also obtained from CHELEM[10]. Additionally, two dummy variables have been built to include the effect of particular integration agreements on

[10] Additional information about the dataset can be obtained from Appendix 1.

trade. Namely FTA_{ijt} which is 1 if both countries have a free trade agreement at time t and finally, the key variable of interest $EURO_{ijt}$ which equals one 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 (see, in particular, Bun and Klaasen, 2007). The purpose is to isolate the effects of EMU on trade trying to control for other factors that may have an influence on trade flows but are not related to the monetary union. The gravity model predicts that bilateral trade flows should depend on factors such as economic size or "mass" (i. e. gravity variables related to economic size and population), distance, and other related considerations. Bearing this in mind the basic panel equation in the literature can be expressed as follows:

$$TRADE_{ijt} = \beta_1 GDP_{ijt} + \beta_2 GDPCAP_{ijt} + \delta_1 EMU_{ijt} + \delta_2 FTA_{ijt} + \eta_{ij} + \tau_{ij} \,.\, t + \lambda_t + \varepsilon_{ijt} \qquad (1)$$

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

The fixed effect (η_{ij}) is intended to capture all individual fixed factors, including unobservable characteristics, associated with a given country pair 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). Hence, the model does not include distance between countries as an explanatory variable and assume that country-pair specific fixed effects will account for the distance effect. Moreover, as we have previously stated, the econometric approach used in this paper accounts for spatial dependence properly.

The time effects (λ_t and $\tau_{ij} \,.\, t$) are intended to capture both common and individual time developments with respect to bilateral trade across all trading partners in the panel. An example of the first could be the special case of a linear time trend in trade that captures the increasing global integration process for all country-pairs[11], whereas an example of the second could be due to country-specific variables such as institutional characteristics, factor

[11] 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.

endowments, and cultural aspects that may also change over time and that can be captured by country specific time trends. Therefore, the approach that we follow to account for trend effects is very flexible, as in Bun and Klaasen (2007) and considers both, the time dimension and the heterogeneous behavior (coefficients) across country-pairs.

The set of coefficients δ_1 and δ_2 represents the effect of EMU and any free trade agreements on trade between member states relative to their country peers (including extra-area trade), after controlling for the effects of economic size and population, time-invariant (fixed) individual country-pair effects (i. e., distance, common language, etc.), common time effects (i. e., transport costs reductions, world trade liberalization..) and time dependent country-pair effects (productivity, capital intensity, etc.). Therefore, the parameter of interest is δ_1 and the difference in trade 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 the EMU trade effects at the area-wide level as well as cross-country differences.

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

In this paper we present an alternative testing procedure to deal with the problem of cross-section dependence. We first suggest to 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 statistics that account for such dependence when required.

In this paper, we test whether the time series are cross-section dependent and then apply panel data methodology accounting for cross-section dependence.

4.1.1. Testing the null hypothesis of cross-section independence

In this subsection we test the null hypothesis of non correlation against the alternative hypothesis of correlation using the approach suggested by Pesaran (2004). He designs a test statistic based on the average of pair-wise Pearson's correlation coefficients $\hat{\rho}_j$, j = 1, 2,…..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 1 show that the Pesaran's CD statistic strongly rejects the null hypothesis of independence, 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 1. Pesaran's CD statistic

Variable	**Test**
gdp_{ijt}	37.011***
GDP_{ijt}	49.095***
$gdpcap_{ijt}$	40.382***
$GDPCAP_{ijt}$	57.275***
$Trade_{ijt}$	30.515***
$TRADE_{ijt}$	32.515***

*** denotes rejection at 1% level.

4.1.2. Panel data unit root and stationarity tests with cross-section dependence

We should start the analysis studying 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 and Lin (1992, 1993) finally published as Levin, Lin and Chu (2002), to the Im, Pesaran and Shin-IPS (2003) tests. In this section we have applied in addition to the IPS panel unit root test, the LM tests for the null of stationarity proposed by Hadri (2000) with heterogeneous and serially correlated errors due to its better power. These last tests can be considered the panel version of the KPSS tests applied in a univariate context. The two statistics are called $\eta\mu$ for the null of stationarity around a constant and $\eta\tau$ when the null is stationarity around a deterministic trend.

As we discussed in a previous section, these first generation tests were based under the unrealistic assumption of cross-section independence. Some authors have proposed different alternatives in order to relax this hypothesis. They have tried to deal with this problem by considering correlations across units as nuisance parameters, and therefore removing them from the model. O'Connell (1998) and Levin *et al.* (2002) suggest subtracting the cross-section mean from the data. This is the approach we use, although its main drawback is that it assumes that the effect of cross-section dependence is the same for all individuals. We apply the tests to the logarithm of both real and nominal bilateral GDP and GDP per capita as well as to the two definitions of the dependent variable: $TRADE_{ij}$ and $Trade_{ij}$. After a visual inspection of the series we have decided to include both, a trend and a constant, in the specification of the test. The number of lags is chosen using the Akaike and Schwarz information criteria.

The results of the tests applied to the variables involved are presented in Table 2. As we expected, the results unambiguously show that all the series are non-stationary. First, our results for the IPS statistics take values from -0.119 to 0.899. The critical value in this case at 5% conficence level is -2.31 (see Im et al. 2003, table 2), so the IPS test fails to reject the null hypothesis of a unit root for all the series. Additionally, when we implement Hadri's test, the null hypothesis of stationarity can be easily rejected in the two cases (with and without time trend) at 1%, so that all the panel variables can be considered non-stationary. This reinforces our initial intuition of the existence of a unit root in the data.

Table 2. Panel unit root tests

Variable	IPS	Pesaran CADF	Hadri (trend)	Hadri (no trend)
gdp_{ijt}	-0.119	-2.223	122.784***	207.235***
GDP_{ijt}	0.899	-2.058	103.014***	148.488***
$gdpcap_{ijt}$	0.3613	-2.099	123.368***	200.025***
$GDPCAP_{ijt}$	0.826	-2.043	105.993***	151.527***
$Trade_{ijt}$	-1.358	-2.101	65.665***	142.536***
$TRADE_{ijt}$	-1.358	-2.346	65.665***	142.536***

*** denotes rejection at 1% level. All variables are in logarithms. One lag is selected for real GDP and real GDP per capita and two lags for the rest of variables using AIC and BIC. For Hadri tests, Newey-West bandwidth selection using Bartlett kernel is used to determine maximum lags, which is one in all variables. Cross-sectional means are removed.

Next we follow Pesaran (2007) and Bai and Ng (2004) and specify the cross-sectional dependencies as driven by a common factor model, so that it is possible to distinguish between the idiosyncratic component and the common component. Although there are differences among the methods proposed, their driving idea is similar. While Pesaran (2007) focuses on the extraction of the common factors that generate the cross correlations in the panel to assess the non-stationarity of the series, in Bai and Ng (2004) the non-stationarity of the series can come either from the common factors, the idiosyncratic component or from both. Moreover, Pesaran (2007) only considers the existence of one common factor while the other alternative can consider several ones. The main advantage of this method is its simplicity to compute while its drawback is that the behavior of the idioysncratic component is to some extent neglected being assumed its stationarity[12]. We implement both tests in this section. The results obtained from the Pesaran CADF test are reported in Table 2 and confirm previous findings, with a critical value of -2.50 at a 5% confidence level and statistics that varies from -2.043 to -2.223.

[12] See Appendix 2.

In addition to the previous evidence, a complementary test is the one based on the approximate common factor models of Bai and Ng (2004). This is a suitable approach when cross-correlation is pervasive, as the analysis with Pesaran's (2004) dependence test has revealed. Furthermore, this approach controls for cross-section dependence given by cross-cointegration relationships, where the time series in the panel might be cross-cointegrated — see Banerjee et al. (2004). The Bai and Ng (2004) approach decomposes the $Y_{i,t}$, time series as follows:

$$Y_{i,t} = D_{i,t} + F_t' \pi_i + e_{i,t},$$

$t = 1, \ldots, T$, $i = 1, \ldots, N$, where $D_{i,t}$ 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 $e_{i,t}$ 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 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 uses 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 3. Panel A of the table corresponds to the variables defined as it is standard in the gravity equations literature. In panel B, in contrast, the variables have been defined following Baldwin's critiques.

Table 3. Panel Data Statistics based on Approximate Common Factor Models Bai and Ng (2004) statistics

Panel A: Variables defined following standard literature **Bai and Ng (2006) statistics**						
	TRADE		***GDP***		***GDPCAP***	
	Test	p-value	Test	p-value	Test	p-value
Idiosyncratic ADF statistic	-0.8773	0.190	-11.673	0.000	-3.275	0.0005
	Test	$\hat{r}_1$	Test	$\hat{r}_1$	Test	$\hat{r}_1$
MQ test (parametric)	-3.733	1	-34.672	6	-35.646	6
MQ test (non-parametric)	-2.373	1	-34.933	6	-36.717	6

Panel B: Variables defined following Baldwin's critique **Bai and Ng (2006) statistics**						
	TRADE		***GDP***		***GDPCAP***	
	Test	p-value	Test	p-value	Test	p-value
Idiosyncratic ADF statistic	-2.625	0.004	-5.277	0.000	-3.113	0.000
	Test	$\hat{r}_1$	Test	$\hat{r}_1$	Test	$\hat{r}_1$
MQ test (parametric)	-36.737	4	-25.495	6	-26.369	6
MQ test (non-parametric)	-37.165	4	-23.346	6	-25.607	6

Concerning the idiosyncratic component, the results of the panel ADF unit root tests point to the rejection of the unit root hypothesis, with the only exception of the variable trade in the standard definition (panel A). In the case of the factor component, all the GDP variables have a total of six factors, whereas the trade variables have just one, in panel A, and four in panel B. It should be noted that our identification of the number of factors when the variables are defined following the standard tradition, are similar to those found by Gengenbach (2009). 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 2 can the null hypothesis of independent stochastic trends be rejected.

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

4.2. Panel cointegration

The econometric methodology we use to analyze long-run relationships among the variables of our panel is based on cointegration techniques. Two approaches can be adopted to estimate the parameters in the panel. In the homogeneous case, we restrict the β parameters to be the same for all the countries in the panel, that is, $\beta_{11} = \beta_{12} = ... = \beta_{1N}, \beta_{21} = \beta_{22} = ...\beta_{2N}$, etc.

In the heterogeneous panel case, this restriction is lifted and the slope coefficients may differ between countries. This possibility makes the use of the heterogenous panel methodology especially interesting in this case, because we expect to find diversity of results.

We have applied tests for cointegration both in the homogeneous and heterogeneous case. More specifically, the panel tests that have been implemented in this paper are: first, the DF and ADF-type tests proposed by Kao (1999) for the null hypothesis of no cointegration in homogeneous and heterogeneous panels; second, the panel cointegration test proposed by McCoskey and Kao (2001) for the null of cointegration in heterogeneous panels, based on Harris and Inder (1994) LM test developed for time series[13]. The application of the LM test makes it necessary to use an efficient estimation technique of cointegrated variables. Kao and Chiang (2000) recommend the fully modified (FM) estimator of Phillips and Hansen (1990) and the dynamic ordinary least squares (DOLS) estimator as proposed by Saikkonen (1991) and Stock and Watson (1993). The DOLS estimator is specially suited for this case: the relation linking trade and GDP and GDPpc should allow for the presence of adjustment costs, since neither exports (imports) react immediately to changes in foreign demand because of the presence of investment plans, capacity constraints, and therefore, they have to be accounted for by the inclusion of lagged variables.

The results of the panel cointegration tests based on the DOLS residuals are presented in Table 4, for the homogeneous case and Table 6, for the heterogeneous one. We consider two alternative models (model 1 and model 2), the first one specified without dummies whereas the second includes two dummies that correspond to those pairs of countries in a monetary union (EMU_{ijt}) and to those in a trade agreement (FTA_{ijt}). Moreover, in both cases, we construct the variables for $TRADE_{ij}$, (GDP_i x GDP_j) and ($GDPCAP_i$ x $GDPCAP_j$) following different approaches. We call (a) to the variables that correct for the critiques of Baldwin (2008) and (b) stands for the variables that are constructed as commonly accepted in the gravity equations literature.

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[13] The econometric procedures necessary to calculate the tests and estimate the coefficients have been kindly provided by S. McCoskey and C. Kao. In addition, we have used the program NPT 1.1 (see Chiang and Kao, 2002). All computations have been made in GAUSS 9.

Concerning the homogeneous case, we apply the panel tests proposed by Kao (1999) and that are different versions of the times series Dickey Fuller tests for non-cointegration. The analysis of this section starts with the tests proposed by Kao (1999), based on the OLS residuals and assuming as the null hypothesis the absence of cointegration. The DF_ρ^* and the DF_t^* statistics are not dependent on the nuisance parameters, and are computed under the assumption of endogeneity of the regressors. Alternatively, he defines a bias-corrected serial correlation coefficient estimate and, consequently, the bias-corrected test statistics and calls them DF_ρ and DF_t. Finally, he also proposes an ADF-type test for the null of no cointegration. The results of applying these tests are presented in Table 4. From the information in Table 4, we can reject the null of no cointegration in all cases independently of the variables definition and the inclusion or absence of dummies.

Table 4. Homogeneous Panel Cointegration Tests Kao (1999) DF and ADF Tests. DOLS estimation

MODEL: TRADE = a + b(GDPi x GDPj) + c(GDPper $capita_i$ x GDPper $capita_j$).
Model 1a: No dummies. Baldwin version.
Model 1b: No dummies. Standard version.
Model 2a: with dummies. Baldwin version. No trend.
Model 2b: with dummies. Standard version. No trend

Test	Model 1a Baldwin	Model 1b Standard	Model 2a Baldwin	Model 2b Standard
DF_ρ	-40.47^{***}	-48.34^{***}	-32.12^{***}	-40.09^{***}
$DF_{t\rho}$	-22.87^{***}	-22.19^{***}	-16.52^{***}	-21.77^{***}
$DF_{\rho*}$	-90.81^{***}	91.08^{***}	-77.16^{***}	-90.57^{***}
$DF_{t\rho*}$	-22.06^{***}	-21.48^{***}	-17.90^{***}	-21.24^{***}
ADF	-18.00^{***}	-18.13^{***}	-17.03^{***}	-17.96^{***}

Note: the three asterisks denote rejection of the null hypothesis of noncointegration at 1. The tests statistics are distributed as N(0;1).

The results for the heterogeneous panel are presented in Table 6. In this case we apply again the ADF test for the null of non-cointegration as well as McCoskey and Kao (2001) LM test for the null of cointegration. Also in this case, we easily reject non-cointegration with the ADF test and cannot reject the existence of a long-run relationship using the LM test. The Monte Carlo experiments reported by McCoskey and Kao (2001) point at the larger power of the LM test if compared with other residual based tests for cointegration in heterogeneous panels under the null of no cointegration.

Concerning the parameters estimates, we show in Table 5 the results of models 1 and 2 using four different estimation techniques together with the t-values in parentheses: OLS, bias-corrected OLS, fully modified (or FM) and dynamic OLS (DOLS). According to McCoskey and Kao (2001) the most appropriate estimation methods are the fully modified and the dynamic OLS, as they correct for endogeneity and autocorrelation. Only a few conclusions can be drawn from this simple estimation, where the presence of a trend has not been considered. First, for the majority of the models and estimation techniques, the coefficient of GDP (in its different definitions) was significant and its value around 0.7. Second, the results for the different definitions of GDPCAP are less homogeneous both in sign and in magnitude. Finally, concerning the parameter of the dummy variable in model 2, its value is around 0.2 (with the exception of the FM estimation in model 2b), relatively large. However, we must take into account that at this stage the model does not contain a trend.

Table 5. Homogeneous panel OLS, OLS bias corrected, FM and DOLS

Coefficient estimates (model 1)

Variables	Model 1a OLS	Model 1a Bias-adj.	Model 1a FM	Model 1a DOLS	Model 1b OLS	Model 1b Bias-adj.	Model 1b FM	Model 1b DOLS
GDP	**0.158**	**0.252**	**0.688**	**0.740**	**0.262**	**0.356**	**0.693**	**0.743**
	(3.24)	(1.94)	(5.15)	(4.99)	(4.38)	(2.31)	(4.39)	(4.25)
GDPpc	**1.035**	**0.892**	0.172	**0.754**	**0.937**	**0.792**	0.165	**0.745**
	(16.36)	(5.48)	(1.03)	(4.07)	(12.10)	(4.12)	(0.84)	(3.41)

Note: t-values in parentheses. Significant coefficients in bold.

Coefficient estimates (model 2)

Variables	Model 2a OLS	Model 2a Bias-adj.	Model 2a FM	Model 2a DOLS	Model 2b OLS	Model 2b Bias-adj.	Model 2b FM	Model 2b DOLS
GDP	**0.517**	**0.752**	**0.835**	**0.740**	0.087	0.162	**0.708**	**0.745**
	(7.77)	(4.70)	(5.10)	(4.07)	(1.35)	(1.04)	(4.44)	(4.22)
GDPpc	**-0.185**	**-0.450**	0.014	-0.223	**1.123**	**0.991**	0.105	**0.707**
	(-2.55)	(-2.63)	(0.08)	(-1.15)	(13.65)	(5.11)	(0.84)	(3.21)
EMU	**0.221**	**0.261**	-0.006	**0.357**	-0.035	0.02	**0.962**	**0.284**
	(5.58)	(3.59)	(-0.08)	(4.34)	(-0.91)	(0.29)	(13.18)	(3.51)
RTA	**0.186**	**0.206**	**0.539**	**0.234**	**0.201**	**0.236**	**0.892**	**0.198**
	(6.92)	(5.58)	(14.23)	(5.57)	(7.71)	(6.59)	(24.28)	(4.86)

Note: t-values in parentheses. Significant coefficients in bold.

The main caveat of the tests applied up to now is that they do not consider the presence of cross-section dependence among the members of the panel. Trying to solve this problem, new statistics have been proposed in the literature such as those described in the introduction of the section. Moreover, the existence of structural breaks in the cointegrating relationships biases the results in panel settings, as it has been described in Banerjee and

Carrion-i-Silvestre (2004, 2006, 2010). They propose an extension of the Gregory and Hansen (1996) approach. In addition, they use the common factors to account for dependence.

Table 6. Heterogeneous panel LM and ADF cointegration tests results

Panel tests	Model 1a Baldwin	Model 1b Standard
LM	**-13.999*****	**-13.701*****
ADF	**-12.232*****	**-11.856*****

Notes: (a) The tests and the models have been estimated using COINT 2.0 in GAUSS 3.24 using the procedures provided by S. McCoskey and C. Kao. (b) The critical values at 1% (***), 5% (**), and 10% (*) for the LM tests are the following: with one regressor, 0.549, 0.3202 and 0.233; 0.372, 0.21, and 0.167 with two regressors; 0.275, 0.159 and 0.120 with three (Harris and Inder, 1994). The critical value for the panel LM test is 1.64.

In the appendix we summarize the procedure applied in Banerjee and Carrion-i-Silvestre (2004, 2006, 2010) to test for non-cointegration in models that allow for up to six alternative specifications for breaks affecting both the deterministic elements and the long-run relationships.

In Table 7 we present the results of the tests for non-cointegration Z_j^* for the model with homogeneous structural breaks. The models correspond to all the potential specifications. Using the BIC information criterion, we choose model 3 in the case of the Baldwin variables and model 1 for the standard specification. Model 3 contains a constant and a trend and the structural break affects them both simultaneously, whereas model 1 includes a constant, no trend and the break occurs in the constant term. With the two databases and using again the BIC information criterion, we found six factors in the panel. In order to test for non-cointegration in the two panels, we apply the statistics based on the accumulated idiosyncratic components, Z_j^*. We present the tests for all possible model specifications. With all of them the null hypothesis of non-cointegration can be rejected. Concerning the time of the break, for the variables constructed following Baldwin's critiques, we find the break in 1987, whereas for the standard variables the break is found in 1998.

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).

Table 7. Banerjee and Carrion (2010) BC cointegration tests

Model	Baldwin model			Standard model		
	Z_j^*	r	r_1	Z_j^*	r	r_1
1	-15.50	6	1	-19.67	6	1
2	-13.55	6	1	-15.12	6	1
3	-15.32	6	1	-12.11	6	1
4	-17.76	6	1	-21.03	6	1
5	-23.00	6	1	-19.07	6	1
6	-17.91	6	1	-11.93	6	1

4.3. 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.

A first look at the parameters estimates is given in the second part of table 5, that extends the specification tested for and estimated following McCoskey and Kao (2001) to include the two dummies, RTA and EMU. We provide again the two alternative sets of variables, under the heading “Model 2”. In the first specification, using the Baldwin’s critiques variables, we find that GDP per capita is non-significant in neither the FM nor the DOLS estimation. As for the dummies, the EMU appears to be relevant only in the DOLS case, with a parameter value of 0.35. The existence of regional trade agreements is, in contrast, significant in both cases. Using the FM estimators the magnitude is 0.54, twice as much as the one obtained with DOLS (0.23). Concerning the GDP, the value of the parameter coincides with the one found in model 1 (0.74), the FM estimation being slightly larger (0.835).

The second specification uses the standard definitions of the variables in the literature. Looking at the last two columns (that correspond again to the FM and DOLS estimations) all the parameters are significant, except for GDP per capita in the FM case. The GDP coefficients are in line with those obtained in model 1, all around 0.7. However, the most striking result concerns the magnitude of the dummies, very large (0.9) when using the FM technique. A more plausible result is obtained with the dynamic OLS method: the EMU coefficient value is 0.28, whereas the RTA dummy has a parameter of 0.198.

The FM and DOLS estimators, although consistently estimate the long-run parameters and correct for autocorrelation and endogeneity, do not account for dependence. Moreover, 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 cointegrated 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. Although this procedure yields a consistent estimator of β, the estimator is asymptotically biased. To account for this bias, the authors construct two estimators that deal with endogeneity and serial correlation and re-center the limiting distribution around zero. The first one, CupBC, estimates the asymptotic bias directly. The second, denoted CupFM, modifies the data so that the limiting distribution does not depend on nuisance parameters. Both are "continuously-updated" (Cup) procedures and require iteration till convergence. The estimators are $\sqrt{nT}$ consistent and enable the use of standard tests for inference. Finally, the approach is robust to mixed I(1)/I(0) factors as well as mixed I(1)/I(0) regressors.

Bai et al. (2009) consider the following model:

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

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

$$x_{it} = x_{it-1} + \varepsilon_{it}$$

x_{it} is a set of k non-stationary regressors, β is a $k \times 1$ 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 pooled least squares estimator of β is as follows:

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

Although his estimator is, in general, T consistent, there is an asymptotic bias due to the long-run correlation between e_{it} and ε_{it}. This 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}:

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

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 both 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. Bai and Kao (2006) considered a two-step fully-modified estimator (2sFM). First, they use the pooled OLS to obtain a consistent estimate of β. The residuals are then used to construct a fully-modified (FM) estimator as in Phillips and Hansen (1990): the nuisance parameters induced by cross-section correlation are dealt with just like serial correlation by suitable estimation of the long-run covariance matrices. Thus, the 2sFM treats the I(0) common shocks as part of the error processes.

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 is not consistent. This is why Bai et al. (2009) develop the case of non-stationary common factors, aiming at achieving consistent estimators.

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. In this case, the LSFM estimator is infeasible. Thus, F should be estimated along with β by minimizing the objective function:

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

subject to the constraint $T^{-2}F'F = I_r$ and $\Lambda'\Lambda$ is positive definite.

Although F is not observed when estimating β, and similarly, β is not observed when estimating F, the unobserved quantities can be replaced by initial estimates and iterate until convergence. Once the covariate matrix is concentrated out, it is possible to define:

$$S_{nT}(\beta, F) = \frac{1}{nT^2} \sum_{i=1}^{n} (y_i - x_i\beta) M_F (y_i - x_i\beta)$$

where M is the projection matrix.

The continuously-updated estimator (Cup) for (β, F) is defined as $(\hat{\beta}_{Cup}, \hat{F}_{Cup}) = \underset{\beta, F}{\arg\min}\, S_{nT}(\beta, F)$. This is the solution to the following two nonlinear equations

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

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

where M_F is the projections matrix and V_{nT} is a diagonal matrix consisting of the r largest eigenvalues of the matrix inside the brackets, arranged in decreasing order. The estimator is obtained by iteratively solving for $\hat{\beta}$ and $\hat{F}$ using the previous equations and it is nonlinear although linear least squares estimation is involved in each iteration. The estimator $\hat{\beta}_{Cup}$ is consistent for β, although it still has a bias derived from having to estimate F. 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 8 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 Carrión-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. 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. Among those deterministic components we have also included the common structural break.

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 8, where we have also included the LSDV estimation results and the Bai FM estimator 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 CupFM and the CupBC. These results are presented in the last two columns of the table. Although the Least Squares estimator LSDV 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 8, where we present the results obtained when the variables have been constructed accounting for the critiques by Baldwin and Taglioni (2006). The model has been estimated with six common factors, as was derived from the Banerjee and Carrion-i-Silvestre (2010) analysis. The dummy variable that should capture the EMU effect is non-significant independently of the estimation technique we apply. Thus, we show the results for the model with just RTA as a dummy variable. Also in this case, its significance can be questioned, especially for the CupBC estimator (last column of the table). Using this estimator, the two GDP measures have positive parameters smaller than one. RTA, the regional trade agreements dummy, is negative and non-significant. The reason that may explain this outcome is that we found that the common structural break occurs around 1987, the date when the Single Market was born. Thus, the majority of the bilateral effects represented by this dummy could have been already captured by the structural break. In addition, the results are quite similar no matter the estimator chosen, with the exception of LSDV. We should note that this estimator is shifted away from zero due to an asymptotic bias induced by the cross-section dependence. The three estimated coefficients obtained using LSDV are much larger than with the other estimators due to the above mentioned upward bias. This bias may lead to the conclusion that the effects of the integration agreements are much larger and positive. The Bai FM estimator, in contrast, corrects for the presence of dependence and assumes stationary common factors. However, Bai et al. (2009) strongly recommend the use of the CupFM and CupBC when there is dependence and the common factors are non-stationary.

Table 8. Cup estimation of the long-run parameters 1967-2008

Variables	LSDV	Bai FM	CupFM	CupBC
Baldwin variables definition				
GDP_{ijt}	**1.77**	**0.78**	**0.79**	**0.81**
	(77.94)	(13.09)	(10.75)	(11.15)
$GDPCAP_{ijt}$	**0.46**	-0.02	**0.30**	**0.27**
	(12.59)	(-0.32)	(3.99)	(3.72)
RTA	**1.34**	**-0.39**	**-0.04**	-0.008
	(18.31)	(-12.14)	(-6.68)	(-1.43)
EMU	---	---	---	---
Standard variables definition				
gdp_{ijt}	**0.80**	**0.53**	**-1.01**	**-0.95**
	(83.08)	(14.33)	(-10.18)	(2.84)
$gdpcap_{ijt}$	**1.40**	**1.08**	**2.81**	**2.84**
	(53.66)	(18.68)	(26.94)	(27.43)
RTA	**-0.13**	**-0.15**	**0.03**	**0.02**
	(-4.16)	(-19.09)	(5.62)	(4.35)
EMU	**-0.43**	**0.26**	**-0.09**	**-0.07**
	(-8.07)	(12.78)	(-5.88)	(-4.80)

The lower files of Table 8 contain the results obtained when we use the variables defined as they commonly are in the empirical literature. We again transform them to account for the deterministic components and the structural break, that this time is found in 1998, in the eve of the creation of the EMU. In this case, the parameters obtained differ both in size and sign from those predicted by the literature. In particular, the GDP variables are larger and, in the case of the Cup estimators, negative for bilateral GDP. The LSDV and the Bai FM estimators provide correctly signed parameters, but are again relatively large. The dummies are also incorrectly signed in the majority of the cases and also large. The reason behind these striking results may have two origins: first, the Baldwin critiques, already mentioned above in the paper, and the inclusion of a common structural change, that may capture at least partially the effects of regional trade agreements and monetary integration.

We can compare our results with previous findings in this literature. In particular, Gengenbach (2009) presents a summary of the main results that he obtains using the Cup estimator and Pesaran's (2006) CCEP estimator and compares them with those found by Bun and Klaassen (2002, 2007) LSDV, DOLS and ADL estimates. He also considers not only the trended but also the non-trended versions of the specified relationships. We should note that all the variables have the standard definitions used in the gravity equation literature, so that a direct comparison may only be adequate in the case of the second specification that we consider. In addition, none of the studies allow for structural breaks in the relationships. Bearing all this in mind, the first common element in his results is the presence of inverted signs in many of the long-run parameters estimates, although this outcome is less frequent in the case of the dummies. Second, some of the GDP parameters are larger than one (notably in the LSDV estimator and in the Cup estimator with no trends). Therefore, our results are in line with previous evidence and can be summarized as follows.

First, there exists a long-run relationship linking trade and the gravity equation variables but the system exhibits cross-section dependence and non-stationary common factors, that 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 CupFM and CupBC estimators proposed by Bai et al. (2009).

Concerning the results, the variables that have been constructed following Baldwin's critiques provide estimations of the long-run parameters compatible with the theory, that is, correctly signed. Moreover, the inclusion of a good specification of the deterministic elements of the model, such as intercept, trends and structural breaks, seems to be enough to capture a process of economic integration that has been gradual in general with some significant milestones, such as the Single Market or the creation of the Euro area. This explains the little relevance of the dummy variables in the long-run estimated relationship. 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 is simply a continuation of a long-run trend, probably linked to the broader set of EU's economic integration policies and that euro has only a residual effect.

5. Summary and concluding remarks

The last decade has witnessed an upsurge of interest in the estimation of EMU effects on trade. In a seminal paper Rose (2000) provided provocative estimates of the trade effects of CU, suggesting a tripling of trade, which seems to be unrealistic. Therefore, the subsequent literature focused on refining previous analyses either through different econometric methodologies or reducing the size of the panels. In this paper we have tried to overcome some of the main flaws found in the standard empirical literature.

First, Baldwin's (2006) 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 (global trend and general equilibrium considerations) and unobserved bilateral heterogeneity (country pair specific characteristics). In the same vein, Egger (2000) suggests that the proper econometric specification of the gravity model in most applications would be one of fixed country and time effects. The former can be trade policy measures including tariff and non-tariff barriers while the latter are business cycle effects. These are not random but deterministically associated with certain historical, political, geographical or other factors. In order to avoid former problems, in this paper we have accounted for Baldwin's critiques in the specification of the model as well as the constructions of the variables included for 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 (trade, GDP) as well as the cross-sectional correlation between the elements (countries) of the panel normally neglected in the empirical works. 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 through 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 apart from Gengenbach (2009).

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

In this paper we try to fill the above mentioned gaps. Using a relatively novel and complete data set that includes 26 OECD countries from 1967 to 2008, we estimate gravity equations 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 in the time dimension. Third, the appropriate Cup-BC and Cup-FM estimators are used to estimate the long-run relationships.

To the best of our knowledge, there has never been an attempt to jointly incorporate Baldwin's critiques, the hypothesis of cross-sections dependence and structural breaks in the time domain within the estimation of a gravity equation on possibly non-stationary series. This approach allow 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 continuation, or culmination, of a series of policy changes that have led over the last four decades to greater economic integration among the countries that now constitute the EMU. 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 fades out in line with most recent empirical literature.

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Appendix 1: List of countries included in the sample and data sources

The countries included in the study are Australia, Austria, Belgium and Luxembourg, Canada, Chile, Denmark, Finland, France, Germany, Greece, Iceland, Ireland, Italy, Japan, South Korea, Mexico, 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. Hence, we have a balanced panel with dimension N=325 (all possible bilateral combinations of countries) and T= 42. The total number of observations is 13,650. Data for nominal imports and exports are obtained from the CHELEM – CEPII database, and are expressed in current dollars. We deflate both flows using US CPI, obtained from the IMF International Financial Statistics (IFS). Real bilateral trade is calculated as the sum of the logarithms of nominal bilateral exports and imports in US dollars.

GDP_{ij} is the product of bilateral real PPP-converted GDP in country *i* and *j* and gdp_{ij} is the nominal product of bilateral nominal GDP. Both are obtained from CHELEM-CEPII database. Population data used to construct $GDPCAP_{ij}$ and $gdpcap_{ij}$ are also obtained from CHELEM. The FTA_{ij} dummy is constructed using World Trade Organization (WTO) data.

Variable	**Description**	**Source**
$TRADE_{ijt}$	Sum of the logs of nominal exports and imports.	Author's calculation from CHELEM databse
$Trade_{ijt}$	Sum of the logs of real exports and imports, deflated using US CPI	Author's calculation from CHELEM database and IMF International Finance Statistics
GDP_{ijt}	Log of the product of bilateral nominal GDP.	CHELEM database
gdp_{ijt}	Log of the product real PPP-converted GDP	CHELEM database
$GDPCAP_{ijt}$	Log of the product of bilateral nominal GDP divided by total population.	CHELEM database
gdp_{ijt}	Log of the product real PPP-converted GDP divided by total population.	CHELEM database
FTA_{ij}	Dummy variable that takes value 1 if both countries belong to a FTA	WTO
EMU_{ijt}	Dummy variable that takes value 1 if both countries belong to EMU	Author's calculation

Appendix 2: Panel unit root and stationarity tests

a) Im, Pesaran and Shin-IPS test (2003)

Early ADF-type panel unit root tests assumed that all members of the panels have the same autoregressive parameter. This is the case of the Levin, Lin and Chu-LLC (2002), Harris and Tzavalis (1999) or Breitung (2000) tests. Hence, given the model:

$$y_{ijt} = \alpha_i + \rho_i y_{it-1} + \varepsilon_{it}; i = 1,...N; t = 1,...T$$

where $\varepsilon_i \sim iid(0,\sigma^2)$, this type of tests propose the null $H_0 : \rho_i = 0$ for all *i* versus the homogeneous alternative $H_A : \rho_i = \rho < 0$. Whereas the null may have sense in some cases, the alternative imposes a very restrictive assumption and it is very unlikely to hold when using country level data. IPS take a step forward, allowing each panel or group of panels to have a different coefficient. They propose a heterogeneous alternative, given by:

$$H_1 : \rho_i < 0; i = 1,...N; \rho_i = 0, i = N_1 + 1,..., N$$

which implies that a fraction of panels are stationary. Thus, instead of pooling the data, separate unit root tests are used for the *N* cross-section units. In order to assure consistency of the test, $N_1 / N \rightarrow \kappa > 0$ as $N \rightarrow \infty$ should be assumed.

Hence, IPS is a group mean test based on individual ADF tests. In a first step, individual ADF regressions are calculated with an error term that is allowed to be serially correlated and heteroskedastic, though cross-sectionally independent:

$$y_{ij} = \alpha_i + \rho_i y_{i,t-1} + \sum_{k=1}^{K_i} \gamma k_i \Delta y_{i,t-k} + \varepsilon_{it}$$

To test for the null hypothesis of a unit root, the IPS t-bar statistic is computed as the average of the individual ADF statistic:

$$\bar{t} = N^{-1} \sum_{i=1}^{N} t_{\rho_i}$$

where t_{ρ_i} is the individual t-statistic of testing $H_0 : \rho_i = 0$ in the previous ADF regression. Assuming a fixed *N*, it can be demonstrated that, as $T \rightarrow \infty$

$$t_{\rho_i} \rightarrow \frac{\int_0^1 W_{iZ}\, dW_{iZ}}{\left[\int_0^1 W_{iZ}^2\right]^{1/2}} = t_{iT}$$

Where W_{iZ} is the projection residual of a standard Brownian motion on a continuous function. Next, IPS assumes that t_{iT} is *iid* and has finite mean and variance. Then, by the central limit theorem

$$\frac{\sqrt{N}\left(N^{-1}\sum_{i=1}^{N} t_{iT} - E\left[t_{iT} \mid \rho_i = 1\right]\right)}{\sqrt{Var\left[t_{iT} \mid \rho_i = 1\right]}} \Rightarrow N(0,1)$$

as $N \rightarrow \infty$. Hence

$$t_{IPS} = \frac{\sqrt{N}(\bar{t}) - E\left[t_{iT} \mid \rho_i = 1\right]}{\sqrt{Var\left[t_{iT} \mid \rho_i = 1\right]}} \Rightarrow N(0,1)$$

as $T \rightarrow \infty$ followed by $N \rightarrow \infty$ sequentially. IPS tabulate in their paper the values $E\left[t_{iT} \mid \rho_i = 1\right]$ and $Var\left[t_{iT} \mid \rho_i = 1\right]$, which are obtained by stochastic simulation using 50,000 replications.

As a drawback of this test, T is implicitly assumed to be equal for all cross-section units, and therefore, it can only be applied to balanced panels.

b) Hadri (2000)

Most of the procedures to test the order of integration of the variables take as null hypothesis the existence of a unit root. By contrast, Hadri (2000) test's is based on the homogeneous null of stationarity (or trend-stationary) versus the heterogeneous alternative that at least one of the panels contains a unit root. His proposal is a panel version of the univariate stationarity KPSS test (Kwiatkowski et al. 1992). The test is based on two alternative regressions:

$$y_{it} = \alpha_i + \sum_{t=1}^{T} \upsilon_{it} + \varepsilon_{it}$$

and

$$y_{it} = \alpha_i + \gamma_i t + \sum_{t=1}^{T} \upsilon_{it} + \varepsilon_{it}$$

where $\sum_{t=1}^{T}\upsilon_{it}$ is a random walk defined through $\upsilon_{it} = \upsilon_{i,t-1} + r_{it}$, and ε_{it}, r_{it} are white noise processes.

The statistic is a residual based Lagrange Multiplier test, with the following expression:

$$LM = \frac{N^{-1}\sum_{i=1}^{N} T^{-2}\sum_{t=1}^{T}\hat{S}_{it}^{2}}{\hat{\sigma}_{\varepsilon}^{2}}$$

where $\hat{S}_{it}$ is the partial sum process of the residuals,

$$\hat{S}_{it}^{2} = \sum_{j=1}^{t}\hat{\varepsilon}_{ij}$$

and $\hat{\sigma}_{\varepsilon}$ is a consistent estimate of the long-run variance of ε_{it}:

$$\hat{\sigma}_{\varepsilon} = \lim_{T\to\infty} T^{-1}E(S_{iT}^{2}), i = 1,\ldots,N$$

LM statistic is distributed as standard Normal under the null hypothesis. This test requires the panels to be strongly balanced. The disturbances are allowed to be homoskedastic across the panel or heteroskedastic across units.

Since there has to be strong evidence against stationarity to conclude in favor of the non-stationarity of the panel; it may be interesting to jointly apply both kinds of tests to obtain more robust conclusions about the panel properties.

c) Pesaran CADF (2007)

Both, IPS (2003) and Hadri (2000) tests belong to the first generation of panel unit root tests, which assume that the error term is cross-sectionally independent. However, several factors as unobserved or omitted common factors or residual interdependence may induce to cross-section dependence. The application of these tests to series characterized by cross-sectional dependencies leads to size distortions and low power. Regarding this, Pesaran suggests to augment IPS test with the cross-sectional averages of lagged levels and their first differences of the individual series (CADF statistics) to proxy the common factors between the cross-sectional units. Analogously to IPS test, it is based on the mean of individual ADF t-statistics of each unit in the panel:

$$\Delta y_{ij} = a_i + b_i y_{i,t-1} + c_i \bar{y}_{t-1} + d_i \Delta \bar{y}_t + \varepsilon_{it}; i = 1,..., N; t = 1,...T$$

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

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

where $CADF_i$ is the t-statistic of b_i in the previous regression.

To avoid size distortions, Pesaran suggest the use of a truncated version of the CADF statistics which has finite first and second order moments. Hence, provided that $\Pr[-P_1 < CADF_i < P_2]$ is sufficiently large, values of $CADF_i$ smaller than $-P_1$ or larger than P_2 are replaced by the respective bounds, where P_1 and P_2 are positive constants. The exact critical values of the t-bar statistic, as well as values for P_1 and P_2 are given by Pesaran (2007).

Appendix 3: Panel cointegration tests

a) Homogeneous panels: Kao (1999) tests with the null of non-cointegration

The various tests summarized in this section are residual-based tests under the null hypothesis of non-cointegration and using OLS estimators. The tests are based on regressing a non-stationary variable on a vector of non-stationary variables and may suffer the spurious regression problem. However, after appropriate normalizations, converge in distribution to random variables with normal distributions.

Kao (1999) proposes two sets of specifications for the DF test statistics. The first set depends on consistent estimation of the long-run parameters, while the second one does not. Under the null hypothesis of no cointegration, the residual series e_{it} should be non-stationary. The model has varying intercepts across the cross-sections (the fixed effects specification) and common slopes across i.

The DF test can be calculated from the estimated residuals as:

$$\hat{e}_{it} = \rho \hat{e}_{it-1} + v_{it}$$

The null hypothesis of non-stationarity can be written as $H_0 \; : \; \rho = 1$.

Kao constructs new statistics whose limiting distributions, $N(0,1)$, are not dependent on the nuisance parameters, that are called DF_{ρ}^{*} and DF_{t}^{*} (where it is assumed that both regressors and errors are endogenous). Alternatively, he defines a bias-corrected serial correlation coefficient estimate and, consequently, the bias-corrected test statistics and calls them DF_{ρ} and DF_{t}. In this case, the assumption is the strong exogeneity regressors and the errors.

Finally, Kao (1999) also proposes an ADF type of regression and an associated ADF statistic.

b) Pedroni (1998) tests for non-cointegration in heterogeneous panels with multiple regressors

The tests proposed in Pedroni (1998) allow for heterogeneity among individual members of the panel, including heterogeneity in both the long-run cointegrating vectors and in the dynamics. Consequently, Pedroni (1998) allows for varying intercepts and varying slopes.

In these tests, the null hypothesis is that for each member of the panel the variables involved are not cointegrated and the alternative that for each member of the panel there exists a single cointegrating vector. Moreover, this vector need not be the same in all cases. This fact makes the tests especially interesting, since very frequently the cointegrating vectors are not strictly homogeneous.

Pedroni (1998) proposes seven tests. Of these tests, four are based on pooling along the *within-dimension*, and three are based on pooling along the *between-dimension.* Thus, the former statistics pool the autoregressive coefficients across different members for the unit root tests on the estimated residuals, while the latter are based on estimators that simply average the individually estimated coefficients for each member i. The distinction is reflected in the autoregressive coefficient, ρ_i, of the estimated residuals under the alternative of cointegration: in the within-dimension statistics, the tests presume a common value for ρ, whereas in the between-dimension statistics, they don't. Thus, the between-dimension introduces an additional source of heterogeneity across the individual members of the panel.

Pedroni (1998) refers to the *within-dimension* based statistics as *panel cointegration statistics*, whereas the *between-dimension* based statistics are called *group mean panel cointegration statistics*. In both cases presents the panel version of the Phillips and Perron ρ and t-statistics, as well as a ADF-type test. The seventh test is a non-parametric variance ratio test only present in the panel cointegration statistics.

The statistics are constructed using the residuals of the cointegrating regression in combination with various nuisance parameter estimators. All the statistics are normally distributed, so that the critical value to consider is -1.95 with the exception of the panel variance tests, which is positive and the reference value is 1.95.

c) Panel cointegration tests: testing the null of cointegration

McCoskey and Kao (2001) propose a residual-based panel test of the null hypothesis of cointegration. This test is an extension of the Langrange Multiplier (LM) test and the Locally Best invariant (LBI) test for a MA unit root in the time series literature. A similar approach has been proposed for the time series case by Harris and Inder (1994). Under the null, the asymptotics are those of the estimation of a cointegrated relationship, instead of the asymptotics of the spurious regression.

For this test it is necessary to use an efficient estimation technique of cointegrated variables. Specifically, the authors recommend to use the fully modified (FM) estimator of Phillips and Hansen (1990) and the dynamic least squares (DOLS) estimator. They also show that the estimators are asymptotically normally distributed with zero means.

The model, that allows for varying slopes and intercepts:

$$y_{it} = \alpha_{it} + x_{it}^{'}\beta_i + e_{it}, i = 1,...,N, t = 1,...,T,$$

$$x_{it} = x_{it-1} + \varepsilon_{it}$$

$$\begin{aligned} e_{it} &= \gamma_{it} + u_{it}, \\ \gamma_{it} &= \gamma_{it-1} + \theta u_{it} \end{aligned}$$

The null hypothesis of cointegration is equivalent to $\theta = 0$.

d) Banerjee and Carrion-i-Silvestre (2004, 2010) panel cointegration tests with breaks and dependence

Banerjee and Carrion-i-Silvestre (2010) propose panel tests for the null hypothesis of no cointegration allowing for breaks both in the deterministic components and in the cointegrating vector. In addition, they tackle cross-section dependence using factor models.

Let $Y_{i,t} = (y_{i,t}, x_{i,t}^{'})$ be a $(m \times 1)$-vector of non-stationary stochastic process whose elements are individually I(1) with the following Data Generating Process (DGP) in structural form:

$$\begin{aligned} y_{i,t} &= D_{i,t} + x_{i,t}^{'}\delta_{i,t} + u_{i,t} \\ u_{i,t} &= F_t^{'}\pi_i + e_{i,t} \\ (I - L)F_t &= C(L)w_t \\ (I - \rho_i L)e_{i,t} &= H_i(L)\varepsilon_{i,t} \\ x_{i,t} &= x_{i,t-1} + G_t^{'}\varsigma_i + \Xi(L)v_{i,t} \\ G_t &= \Gamma(L)\varpi_t, \end{aligned} \tag{1}$$

$i = 1,...,N,\ \ t = 1,...,T_{\cdot j \cdot}$, where $C(L) = \sum_{j=0}^{\infty} C_j L^j$. The general functional form for the deterministic term $D_{i,t}$ is given by:

$$D_{i,t} = \mu_i + \beta_i t + \sum_{j=1}^{m_i} \theta_{i,j} DU_{i,j,t} + \sum_{j=1}^{m_i} \gamma_{i,j} DT_{i,j,t}, \tag{2}$$

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

$$\delta_{i,t} = \begin{cases} \delta_{i,1} T^c_{i,0} < t \le T^c_{i,1} \\ \delta_{i,2} T^c_{i,1} < t \le T^c_{i,2} \\ \cdots\cdots \\ \delta_{i,j} T^c_{i,j-1} < t \le T^c_{i,j} \\ \cdots\cdots \\ \delta_{i,n_i+1} T^c_{i,n_i} < t \le T^c_{i,n_i+1} \end{cases} \tag{3}$$

with $T^C_{i,0} = 0$ and $T^C_{i,n_i+1} = T$, where $T^C_{i,j} = \lambda^C_{i,j} T$ denoting the *j*-th time of the break, $j = 1,\ldots,n_i$, for the *i*-th unit, $i = 1,\ldots,N$, for the $i-$th unit, $i = 1,\ldots,N$, $\lambda^C_{i,j} \in \Lambda$.

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

Model 1. No linear trend **-** $\boldsymbol{\beta_i = \gamma_{i,j} = 0}$ $\forall i,j$ in (2) – and constant cointegrating vector - $\delta_{i,j} = \delta_i$ $\forall j$ in (3). In this case, the model only considers the presence of multiple level shifts.

Model 2. Stable trend - $\boldsymbol{\beta_i \neq 0}$ $\forall i$ and $\boldsymbol{\gamma_{i,j} = 0}$ $\forall i,j$ in (2) - and constant cointegrating vector - $\delta_{i,j} = \delta_i$ $\forall j$ in (3). In this case, the model only considers the presence of multiple level shifts.

Model 3. Changes in level and trend - $\boldsymbol{\beta_i \neq \gamma_{i,j} \neq 0}$ $\forall i,j$ (2) – and constant cointegrating vector - $\delta_{i,j} = \delta_i$ $\forall j$ in (3). In this case, the model only considers the presence of multiple level and trend shifts.

Model 4. No linear trend **-** $\boldsymbol{\beta_i = \gamma_{i,j} = 0}$ $\forall i,j$ in (2) – but the presence of multiple structural breaks affects both the level and the cointegrating vector of the model.

Model 5. Stable trend $\boldsymbol{\beta_i \neq 0}$ $\forall i$ and $\boldsymbol{\gamma_{i,j} = 0}$ $\forall i,j$ in (2) – with the presence of multiple structural breaks, that affect both the level and the cointegrating vector of the model.

Model 6. Changes in the level, trend and in the cointegrating vector. No constraints are imposed on the parameters of (2) and (3).

Banerjee and Carrion (2010) assume strictly exogenous stochastic regressors[14]. The common factors are estimated following the method proposed by Bai and Ng (2004). They

[14] In addition, Banerjee and Carrion (2006) suggest using the $DOLS$ estimation method proposed by Stock and Watson (1993) to account for endogeneity. The lags can be chosen using an information criterion.

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.

They estimate the common factors as in Bai and Ng (2004). They compute the first differences:

$$\Delta y_{i,t} = \Delta D_{i,t} + \Delta x'_{i,t}\delta_{i,t} + \Delta F'_t \pi_i + \Delta e_{i,t}$$

and take the orthogonal projections:

$$\begin{aligned} M_i \Delta y_i &= M_i \Delta F \pi_i + M_i \Delta e_i \\ &= f\pi_i + z_i, \end{aligned}$$

with $M_i = I - \Delta x_i^d (\Delta x_i^{d'} \Delta x_i^d)^{-1} \Delta x_i^{d'}$ being the idempotent matrix, and $f = M_i \Delta F$ and $z_i = M_i \Delta e_i$. They estimate the common factors and the factor loadings using principal components. The estimated principal component of $f = (f_2, f_3, \ldots, f_T)$, denoted as $\tilde{f}$ is $\sqrt{(T-1)}$ times the r eigenvectors corresponding to the first r largest eigenvalues of the matrix $y^* y^*$? , where $y_i^* = M_i \Delta y_i$. Then, the estimated residuals are $\tilde{z}_{i,t} = y_{i,t}^* - \tilde{f}_t \tilde{\pi}_i$. It is possible to recover the idiosyncratic disturbance terms through cumulation, so that $\tilde{e}_{i,t} = \sum_{j=2}^{t} \tilde{z}_{i,j}$ and test the unit root hypothesis $(\alpha_{i,0} = 0)$ using the ADF regression equation

$$\Delta \hat{e}_{i,t}(\lambda_i) = \alpha_{i,0}\hat{e}_{i,t-1}(\lambda_i) + \sum_{j=1}^{k} \alpha_{i,j} \Delta \hat{e}_{i,t-j}(\lambda_i) + \varepsilon_{i,t}.$$

The null hypothesis of a unit root can be tested using the pseudo *t*-ratio $t_{\tilde{e}_i}^j(\lambda_i)$, $j = c, \tau, \gamma$ for testing $\alpha_{i,0} = 0$ above. The models that do not include a time trend (Models 1 and 4) are denoted by *c*. Those that include a linear time trend with stable trend (Models 2 and 5) are denoted τ and, finally, γ refers to the models with a time trend with changing trend (Models 3 and 6).

When the number of factors *r=1*, we can use an ADF-type equation to analyze the order of integration of the common factors F_t , whereas when *r>1* we should use one of the two statistics proposed by Bai and Ng (2004) to fix the number of common stochastic trends. λ_i is the break fraction parameter, in most cases unknown.

The individual statistics for the idiosyncratic disturbance terms can be pooled to define panel data cointegration tests. There are several possible tests, depending on whether the break point is known or unknown and the degree of heterogeneity desired.

When the breaks are known, the panel data cointegration test is based on the average of the individual cointegration statistics:

$$Z_j(\lambda) = \sum_{i=1}^{N} t_{\hat{e}_i}^{j}(\lambda_i), \qquad j = c, \tau, \gamma,$$

where $\lambda = (\lambda_1^{'}, \lambda_2^{'}, \ldots, \lambda_N^{'})'$ for $j = c, \tau$ and $\lambda = (\lambda_1, \lambda_2, \ldots, \lambda_N)'$ for $j = \gamma$. The limiting distribution is described in Banerjee and Carrion-i-Silvestre (2010).

When the breaks are unknown and are allowed to be individual specific (heterogeneous), the breaks can be estimated following the procedure in Bai and Carrion-i-Silvestre (2009). It consists of minimizing the sum of square residuals over all possible break dates in the model written in first differences. Using the estimated breaks, the factors can be obtained as described above and, then, the standardized statistic can be constructed.

When common (homogeneous) structural breaks are imposed to all the units of the panel (although with different magnitudes), we can compute the $Z_j(\lambda)$, $j = c, \tau$ statistic for the break dates, where the break dates are the same for each unit, using the idiosyncratic disturbance terms. Thus, the null hypothesis of non-cointegration for the idiosyncratic terms is:

$$Z_j^* = \inf_{\lambda \in \Lambda^m} \left(\frac{N^{-1/2} Z_j(\lambda) - \Theta_j^{\tilde{e}} \sqrt{N}}{\sqrt{\Psi_j^{\tilde{e}}}} \right), j = c, \tau$$

where $\hat{\lambda} = \arg\min_{\lambda \in \Lambda^m} \left[\left(N^{-1/2} Z_j(\lambda) - \Theta_j^{\tilde{e}} \sqrt{N} \right) \left(\Psi_j^{\tilde{e}} \right)^{-1/2} \right]$ The limiting distribution of $Z_j^*(\lambda)$, $j = c, \tau$,

$$Z_j(\lambda) \Rightarrow N(0,1), \qquad j = c, \tau$$

Ivie

Guardia Civil, 22 - Esc. 2, 1º
46020 Valencia - Spain
Phone: +34 963 190 050
Fax: +34 963 190 055

Website: http://www.ivie.es
E-mail: publicaciones@ivie.es