Is there a "euro effect" on trade? New evidence using gravity equations with panel cointegration techniques

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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 OECD countries covering the period 1967-2008. We estimate the equation using two sets of variables: first, one defined as it is standard in the gravity equation literature, and a second one built according to the criticisms stated by Baldwin and Taglioni (2006). Moreover, from a methodological point of view, we implement a new generation of tests that allow solving some of the problems derived from the non-stationary nature of the data (GDP, trade). To this aim we apply panel tests that account for the presence of cross-section dependence as well as discontinuities in the non-stationary panel data. We test for cointegration between the variables using panel cointegration tests, especially the ones proposed by Banerjee and Carrión-i-Silvestre (2010). We also efficiently estimate the long-run relationships using the continously updated estimators proposed in Bai et al. (2009). Our results challenge earlier estimates using standard panel data techniques and are in line with those of Bun and Klaassen (2007). We argue that, after controlling for cross-section dependence and deterministic trends and breaks in trade integration, the euro appears to generate far lower trade effects than predicted in previous studies.

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1. Introduction

The introduction of the euro has raised a new interest in measuring the impact of currency unions (CU) on trade flows. The very high estimates of trade induced by the creation of monetary unions found in the seminal papers by Rose (2000) and Frankel and Rose (2002) has led to the concept of "endogeneity" of Optimum Currency Areas (OCA) that means for the Euro-area that, even if the European Monetary Union (EMU) was not created as an OCA, it could be progressing in that direction (Frankel and Rose, 1998). Recent research surveyed by Rose and Stanley (2005) and Rose (2008) suggests that the introduction of the euro still has a sizable and statistically significant effect on trade among EMU members. Taking together all these estimates imply that 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 while Greece entered in 2001. Since then, also Slovenia, Cyprus, Malta, Slovakia and Estonia have joined the Euro-area while other members of the EU are "waiting and seeing", the so-called derogation countries. Moreover, the introduction of the euro was preceded by other stages of economic integration (Customs Union, European Monetary System and the Single Market), so 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 Faruqee (2004) the central questions at stake are the following: first, to ascertain the effects of EMU on the area's trade flows; second, to analyze the evolution of the trade effects through time, and finally, to measure the distribution of trade effects among member states.

In this paper we have tried to overcome some of the main flaws found in the standard empirical literature and recently outlined by Eicher and Henn (2011). First, 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 should include fixed country and time effects. The former account for trade policy measures including tariff and non-tariff barriers while the latter can capture business cycle effects. These are not random effects but instead deterministically associated with certain historical, political or geographical factors. In order to avoid the above-mentioned problems, in this paper we have accounted for Baldwin's critiques in the specification of the model as well as in the definition of the variables included in the estimation of the gravity model.

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

Third, Bun and Klaasen (2007) have stated that models measuring the effect of the euro on trade have omitted some variables, causing an upward bias in the trade benefits earlier esti-

mated. They find that the longer the data period considered, the higher the euro effect estimate. Thus this might be due to some misspecification of the time-series characteristics of the variables involved, namely the trends in trade flows over time. To correct for this bias they add a time-trend to their specification and allow it to have heterogeneous coefficients across country-pairs. Then they estimate long run relationships using first-generation panel cointegration techniques, that is, without considering dependence in the cross-section dimension.

Therefore, in this paper we try to provide new evidence on the effect of the euro using a data set that contains information on all bilateral combinations in a panel of 26 countries covering the period 1967-2008. We implement a new generation of tests that allows us to solve some of the problems derived from the non-stationary nature of the data used in gravitational equations (GDP, trade, etc). 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 series. More specifically, we implement the panel unit root and stationary tests proposed by 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, we apply the continuously updated estimator (CUP) of Bai et al. (2009) to efficiently estimate the regression coefficients. The results obtained are in line with Bun and Klaassen (2007) confirming a smaller euro effect than in other research papers, like for instance, Gil-Pareja et al (2008), where cross-section dependence and the non-stationary nature of the variables are not accounted for.

The 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 problems present in the current literature. Section 4 describes the data and discusses the empirical results. A final section concludes.

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

The literature examining the impact of CU on trade is a burgeoning field of research. All in all, the diversity of existing estimates indicates the potential bias inherent in applied specifications. Although in the beginning the gravity model was criticized for its lack of theoretical underpinnings, now rests on a solid theoretical background[2]. Therefore, as stated in Westerlund and Wilhelmsson (2009) the focus of this line of research has shifted from its theoretical soundness towards the estimation techniques used.

The econometric approach has changed over time as a result of a feedback process between theory and empirics. In this abundant literature, the traditional approach has been to use *cross-section data*. However, it is generally accepted that the results obtained were suffering from a bias, as the heterogeneity among countries was not properly controlled for. Thus, Rose's (2000) initial estimates in a cross-sectional study suggested a tripling of trade. This result was quite striking, and as quoted by Faruqee (2004), is at odds with the related literature that typically finds very little negative impact of exchange rate volatility on trade. Not surprisingly, Rose's findings have received substantial revisions, and subsequent analysis

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

generally finds a smaller (albeit still sizable) effect of CU membership on trade. There are different reasons that make the implication of Rose (2000) work unclear. First, the sample countries were mostly smaller and poorer, not including the EMU ones. This has led to question whether the results apply to bigger countries such as 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 most relevant issue about EMU is the possible change in the level of trade for its member over time, before and after the introduction of the single currency. In order to solve this problem, a second string of literature started to use *panel data estimation techniques*, which permits more general types of heterogeneity[3]. However, Baldwin and Taglioni (2006) define what they call in this context "the gold medal error", also known as 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[4]. 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, 2006). Country dummies (for exporters and importers) only remove the average impact leaving the time dimension in the residuals, which leads to biased results. Therefore, time-invariant country dummies are not enough and a proper treatment of the time dimension is needed. Moreover, Baldwin and Taglioni (2006) also stress the importance of an omitted variable bias when the empirical specification does not account for unobserved determinants of bilateral trading relationships. This problem can be solved introducing bilateral heterogeneity.

In addition to the above-mentioned specification caveats, Baldwin and Taglioni (2006)

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[3] 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.

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

pointed out two additional minor problems, coined as "silver" and "bronze" medal errors. 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 deflator: all the prices in the gravity equation are measured in terms of a common numeraire, so there is no price illusion. However, many authors deflate trade flows and GDP using the US CPI (following Rose's example). As Baldwin et al. (2008) claim, fortunately, the bronze medal bias is eliminated by including time dummies which is the common practice.

Finally, concerning the estimation problems, Santos-Silva and Tenreyro (2006) argue that the standard empirical methods used to estimate the gravity equation (i.e. Ordinary Least Squares, OLS) are also inappropriate, even if these problems have been largely ignored by applied researchers, as the econometric methods commonly used to solve them were not easy to implement[5]. Glick and Rose (2002) and Frankel and Rose (2002) exploited the time series information using panel data. They obtained similar results[6] giving birth to a literature in search of "more reasonable" effects (Eicher and Henn, 2011). Micco et al. (2003) examined the dynamic impact of EMU on trade for 22 industrial countries using panel regressions

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[5] 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. 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.

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

based on a gravity model. Their findings suggest that EMU has fostered bilateral trade between 8% and 16% depending on the EMU membership of the countries and that the positive effect has been rising over time. Other studies, like Bun and Klaasen (2002) estimate a dynamic panel data model and distinguish between short (3.9%) and long-run effects (38%). 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%[7], which is smaller than previous evidence. However, these papers generally use smaller and shorter datasets than Rose's. When they focus on large panels, they find bigger estimates (over 100%). Therefore, the empirical literature is far from conclusive and we can infer that dataset dimensions, and, especially, econometric approaches, influence the results.

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

Bun and Klaasen (2007) constitutes a path-breaking study in this respect. They show that the residuals of the least squares dummy variables estimator (LSDV) exhibit trends over time.

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

Therefore, they estimate the gravity equation allowing for country pair specific time trends to account for the observed trending behaviour in the residuals. Moreover, they analyze the non-stationary nature of the data as well as the cointegration relationships and obtain a much smaller estimate of the euro effect (3%) on bilateral trade[8]. 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. e. other economies business cycles'). Moreover, dependence is generated by construction as gravity models include bilateral trade flows together with aggregate national variables. Furthermore, the gravity model itself implies spatial dependence in the data due to the hypothesized effect of distance on trade. Several new panel unit root and cointegration tests have been proposed accounting for cross-sectional dependence in the form of common factors.[9]

More recent studies have insisted on the importance of accounting for the existence of trends in the data and its possible non-stationary nature. Historically, researchers have assumed stationary time series to estimate gravity models. However, if the variables are non-stationary, a different statistical setup needs to be used. As Faruqee (2004) claimed, estimating the impact of a monetary union on trade faces several econometric challenges. Recent literature shows that the results of the gravity models are sensitive to their proper specification (Egger and Pfaffermayr, 2003). However, properly specified models in panel data may have some caveats when data are non-stationary. If the non-stationary nature of the series is not considered, spurious regressions may appear. Although the spurious correlation problem is less impor-

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[8] Other papers that stress the importance of the non-stationary nature of the series and that apply cointegration techniques are Faruqee (2004) and Fidrmuc (2009).
[9] See for example Breitung and Pesaran (2008) for an overview of the literature and Gengenbach et al (2010) for a comparison of panel unit root tests.

tant in panels than in time series analysis, as the fixed effects estimator for non-stationary data is asymptotically normal (see Kao and Chiang, 2000), the results are biased. Correspondingly, panel cointegration techniques are used accounting for different possible estimation problems (endogeneity, cross-correlation or breaks). Therefore, a sound empirical strategy must proceed as follows: First, to determine the order of integration of the variables through panel unit root tests; second, to test for cointegration among the integrated variables using panel cointegration tests; finally, to use the panel cointegration estimators to provide reliable point estimates.

The contribution of our paper to the existing literature about the euro effect on trade is two-fold. First, unlike previous research, (excepting Eicher and Henn, 2011) we address Baldwin's critiques regarding the proper specification of gravity models and the definition of the variables, as we account for multilateral resistance, as well as country pair specific characteristics through unobserved bilateral heterogeneity. Second, we apply an econometric methodology comprising of a range of techniques to test and estimate efficiently in a non-stationary panel framework, solving endogeneity problems as well as possible biases posed by structural breaks and cross-section dependence.

3. Data, methodology and empirical results

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

framework allowing addressing most of the issues concerning both specification and estimation discussed in the previous section.

A first common problem in the context in panel non-stationary variables is that tests 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. A second generation of panel tests, in contrast, introduce different forms of dependence, solving the above- mentioned problem.

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

In the case of panel unit roots, several tests have been formulated based on factor models[10]. In particular, Bai and Ng (2004) account for the non-stationarity of the series coming either from the common factors, the idiosyncratic component or from both. Moreover, they consider the possible existence of multiple common factors as well as the existence of cointegration relationships among the series of the panel. Banerjee et al. (2004) stated that there is a tendency to over-reject the null of stationarity when cointegration is present. As the existence of cointegrating relations between trade series is a very plausible hypothesis in economic integrated areas, the proposal in Bai and Ng (2004) is the best approach in our case[11].

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

For the sake of comparison, we will also present the results obtained using Pesaran's (2007) approach. Similarly, we will also allow for dependence in the estimation of the cointegration relationships using the common factor approach of Bai and Ng (2004).

A second caveat appears when there are structural breaks in the time dimension of the panel. If there exist linear combinations of integrated variables that cancel out their common stochastic trends then, these series are said to be cointegrated. The economic translation is that these series share an equilibrium relationship. However, a commonly neglected phenomenon is that both, the cointegrating vector and the deterministic components might change during the period analyzed, and if we do not take account of these structural breaks in the parameters of the model, inference concerning the presence of cointegration can be affected by misspecification errors. Therefore, in this paper we propose the use of the tests developed in Banerjee and Carrion-i-Silvestre (2010). They generalize the approach in Pedroni (1999, 2004) to account for one structural break that may affect the long run relationship in a number of different ways (cointegrating vector and/or deterministic components). Moreover, they address the cross-section dependence issue by using the above-mentioned factor model approach due to Bai and Ng (2004) to generalize the degree of permissible cross-section dependence allowing for idiosyncratic responses to multiple common factors.

To sum up, we control for econometric issues usually neglected in earlier literature: first, we account for cross-section dependence among countries in the panel tests, both unit roots and cointegration. Second, we allow for the existence of a break in the deterministic components of the model as well as 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, the estimation of the long-run relationship uses a methodology that not

only efficiently estimates the coefficients but also is based on the common factors decomposition that assures a homogeneous econometric approach. We choose, for this purpose, the CUP Fully Modified (CUP-FM) and the CUP Bias Corrected (CUP-BC) estimators by Bai et al. (2009).

3.1. Data

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 NT =13,650.

Following the discussion in section 2, one of the contributions of the paper is to perform the analysis and the estimation of the gravity equation for the euro effect using two sets of variables. In the first one, that we call "Baldwin-variables" and use upper-case letters, the series have been computed as suggested by Baldwin and Taglioni (2006). The second set of variables, defined as it is commonly done in mainstream gravity literature, is called "standard-variables" and we use lower-case letters to represent them.

Therefore, the dataset includes the following variables where moreover, upper-case letters also 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_{ij} is the log of the product of bilateral nominal GDP in country i and j and gdp_{ijt} is the log of the product of bilateral real PPP-converted 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. Additionally, two dummy variables have been built to include the effect of particular integration agreements on trade. Namely RTA_{ijt} which is 1 if both countries have a free trade agreement at time t is constructed using World Trade Organization (WTO) data, and finally the key variable of interest $EURO_{ijt}$ which equals 1 if both trading partners belong to the euro area in year t and zero otherwise. To the extent that these agreements are made or dissolved during the sample period, this variable is distinct from the time-invariant country-pair fixed effect.

The formal model that we estimate comes from the gravity equation, and in particular, we follow the traditional specification from the recent literature on the euro effect using non-stationary panels (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:

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

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[12].

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)[13]

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, whereas an example of the second could be due to country-specific variables such as institutional characteristics, factor endowments, and cultural aspects that may also change over time and that can be captured by

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[12] Later in the analysis, we will include additional deterministic trends in equation (1), that correspond to structural breaks in either the constant, the trend or both.

[13] 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.

country specific time trends[14]. 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). 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.

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

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

3.2.1. Testing the null hypothesis of cross-section independence

In this subsection we test the null hypothesis of non-correlation against the alternative hy-

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[14] 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.

pothesis 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 Cross-section Dependence (CD) statistic of Pesaran (2004) converges to the standard normal distribution. The results in Table 1 show that 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 and CADF statistics

Variable	CD dependence test	CADF panel unit root test
gdp_{ijt}	37.011^{***}	-2.223
GDP_{ijt}	49.095^{***}	-2.058
$gdpcap_{ijt}$	40.382^{***}	-2.099
$GDPCAP_{ijt}$	57.275^{***}	-2.043
$Trade_{ijt}$	30.515^{***}	-2.101
$TRADE_{ijt}$	32.515^{***}	-2.346

*** denotes rejection at 1% level.

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

Once we have found the presence of dependence in the variables, we study the order of integration of the variables. Several procedures to test for unit roots in panels are already available in the literature, from the early works of Levin et al (2002). However, these first generation tests were based on the unrealistic assumption of cross-section independence[15]. Therefore, we follow Pesaran (2007) and Bai and Ng (2004) and specify the unit root tests allowing for cross-sectional dependence as driven by a common factor model, so that it is possible to distinguish between the idiosyncratic component and the common component. 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[16] while the other alternative can consider several ones. We implement both tests in this section. The results obtained from the Pesaran Cross-Sectionally Augmented ADF (CADF) test are reported in Table 1 concluding in favour of non-stationarity, with a critical value of -2.50 at a 5% confidence level and statistics that vary from -2.043 to -2.346.

In addition to the previous evidence, we also apply the test based on the approximate common factor models of Bai and Ng (2004). This is a suitable approach when cross-correlation is pervasive, as in this case. Furthermore, this approach controls for cross-section dependence given by cross-cointegration relationships, potentially possible among our group of

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

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

countries and variables — see Banerjee et al. (2004). The Bai and Ng (2004) approach decomposes the $Y_{i,t}$, time series as follows:

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

with t = 1, . . . , T , i = 1, . . . , 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 use a non-parametric and a parametric correction respectively to account for additional serial correlation. Both statistics have a non-standard limiting distribution. They test the hypothesis of $k_1 = m$ against the alternative $k_1 < m$ for m starting from $\hat{k}$. The procedure ends if at any step $k_1 = m$ cannot be rejected. The results from the application of the Bai and Ng (2004) statistics are summarized in Table 2. 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 2.
Panel Data Statistics based on Approximate Common Factor Models

Panel A: Variables defined following standard literature
Bai and Ng (2006) statistics

	$trade_{ijt}$		gdp_{ijt}		$gdpcap_{ijt}$	
	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

	$TRADE_{ijt}$		GDP_{ijt}		$GDPCAP_{ijt}$	
	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.

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

3.3. Panel cointegration

As in the case of the unit root tests, the main caveat of the first generation panel cointegration tests is that they do not consider the presence of cross-section dependence among the members of the panel. However, as the majority of the empirical evidence on the euro effect that uses cointegration uses this approach, we have also applied these methods for the sake of comparison and present some of the estimation results in Table 5[17].

Trying to solve the problem of cross-section dependence, new statistics have been also designed to test for cointegration, using factor models in a fashion similar to the one proposed by Bai and Ng (2004) for unit root testing. Moreover, as the existence of structural breaks in the cointegrating relationships biases the results in panel settings - see Banerjee and Carrion-i-Silvestre (2010) - they propose an extension of the Gregory and Hansen (1996) approach using common factors to account for dependence.

Banerjee and Carrion-i-Silvestre (2010) propose 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):

$$y_{i,t} = D_{i,t} + x'_{i,t}\delta_{i,t} + u_{i,t} \tag{2}$$

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

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[17] In particular, we have applied the panel cointegration tests proposed by Kao (1999) and McCoskey and Kao (1998). The complete results are available from the authors upon request.

$$D_{i,t} = \mu_i + \beta_i t + \sum_{j=1}^{m_i} \theta_{i,j} DU_{i,j,t} + \sum_{j=1}^{m_i} \gamma_{i,j} DT_{i,j,t}, \tag{3}$$

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).

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

Model 1. No linear trend - $\beta_i = \gamma_{i,j} = 0 \ \forall \ i, j$ in (3) - and constant cointegrating vector. In this

case, the model only considers the presence of multiple level shifts.

Model 2. Stable trend - $\beta_i \neq 0 \ \forall i$ and $\gamma_{i,j} = 0 \ \forall \ i, j$ in (3) - and constant cointegrating vector.

Also considers only multiple level shifts.

Model 3. Changes in level and trend - $\beta_i \neq \gamma_{i,j} \neq 0 \ \forall \ i, j$ in (3) - and constant cointegrating

vector. The model considers multiple level and trend shifts.

Model 4. No linear trend - $\beta_i = \gamma_{i,j} = 0 \ \forall \ i, j$ in (3) - but the presence of multiple structural

breaks affects both the level and the cointegrating vector of the model.

Model 5. Stable trend $\beta_i \neq 0 \ \forall i$ and $\gamma_{i,j} = 0 \ \forall \ i, j$ in (3) - with the presence of multiple struc-

tural 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 im-

posed on the parameters of (3).

The common factors are estimated following the method proposed by Bai and Ng (2004).

They first compute the first difference of the model; then, they take the orthogonal projec-

tions 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 idiosyn-

cratic 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 us-

ing the ADF statistic.

The null hypothesis of a unit root can be tested using the pseudo *t*-ratio $t_{\tilde{e}_i}^j(\lambda_i)$, $j = c, \tau, \gamma$.

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 by τ and, finally, γ

refers to the models with a time trend with changing trend (Models 3 and 6).

When common (homogeneous) structural breaks are imposed to all the units of the panel

(although with different magnitudes), we can compute the statistic for the break dates, where

the break dates are the same for each unit, using the idiosyncratic disturbance terms[18].

In Table 3 we present the results of the tests for non-cointegration Z_j^* for the model with

homogeneous structural breaks for the six potential specifications discussed above. Using

the BIC information criterion, we choose model 3 in the case of the Baldwin variables and

model 1 for the standard variables definition. 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 alternative-variable

definitions and using again the BIC information criterion, we find six factors in the panel. In

order to test for non-cointegration in the two cases, 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.

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[18] As described in equations (2) and (3), a heterogeneous version of the test is also possible, authough the homogeneous case is the more adequate for the particular case of the gravity model and the estimation of the parameters in the long-run relationship.

Table 3.

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

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).

3.4. Estimation of the gravity equation

Once the different tests applied have provided us with evidence of cointegration, either considering a stable relationship or instabilities, we should obtain the long-run estimates using consistent techniques. Kao and Chiang (2000) recommended the fully modified (FM) estimator of Phillips and Hansen (1990) and the dynamic ordinary least squares (DOLS) estimator as proposed by Saikkonnen (1991) and Stock and Watson (1993). The DOLS estimator is specially suited for the present case because the relationship linking trade, GDP and GDPpc should allow for the presence of adjustment costs, since neither export nor imports react immediately to changes in foreign demand due to the presence of investment plans, capacity constraints. Therefore, in order to account for this the inclusion of lagged variables is highly recommended.

However, although the FM and DOLS estimators consistently estimate the long-run parameters and correct for autocorrelation and endogeneity, do not account for

dependence[19]. This fact is very relevant in this study as we found in the Panel Analysis of Nonstationarity in Idiosyncratic and Common components (PANIC) 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, CUP-BC, estimates the asymptotic bias directly. The second, denoted CUP-FM, 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 $i = 1,...,n,\ t = 1,...,T,\ y_{it}$ is a scalar,

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

x_{it} is a set of k non-stationary regressors, β is a $k\ x\ 1$ vector of the common slope parameters,

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[19] We have also estimated the cointegration vectors by least squares dummy variables (LSDV), fully modified (FM) and dynamic ordinary least squares (DOLS). We have omitted most of these results from the text, although they are available upon request. Some of the results for the case of the standard-variables model are presented in Table 5 for the sake of comparison.

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 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 Ng (2006) considered a two-step fully-modified estimator (2sFM).

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 Least

Square Fully Modified (LSFM) estimator is infeasible. Thus, F_t should be estimated along

with β by minimizing the objective function, the unobserved quantities can be replaced by

initial estimates and iterate until convergence through the CUP estimator for (β, F), defined

as $(\hat{\beta}_{Cup}, \hat{F}_{Cup}) = \arg\min_{\beta, F} S_{nT}(\beta, F)$. The estimator $\hat{\beta}_{Cup}$ is consistent for β, although it still has

a bias derived from having to estimate F_t. The authors correct this bias using two fully-

modified estimators. The first one directly corrects the bias of $\hat{\beta}_{Cup}$ and is denoted $\hat{\beta}_{CupBC}$.

The second one makes the correction in each iteration and is denoted $\hat{\beta}_{CupFM}$.

We present in Table 4 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 accord-

ingly 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

breaks[20].

Therefore, once we have performed this transformation we are able to apply the Bai et al.

(2009) estimators to the two definitions of the variables. The results are shown in Table 4,

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[20] Note that this implies that in the model specification of the gravity equation in expression (1) above, we have filtered the three variables (trade, GDP and GDP per capita) of the deterministic components and the structural breaks, with the exception of the dummies RTA and EMU.

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 CUP-FM and the CUP-BC. These results are presented in the last two columns of the table. Although the LSDV estimator is the most commonly applied in the gravity equation literature, the parameters obtained are biased when the common factors are non-stationary. The size of this bias is shown in Bai et al. (2009) and this may explain earlier results in the applied literature.

Table 4.

CUP estimation of the long-run parameters 1967-2008

Variables	LSDV	Bai FM	CUP-FM	CUP-BC
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)
Baldwin variables definition				
GDP_{ijt}	**1.91**	**1.54**	**1.47**	**1.47**
	(64.39)	(89.90)	(81.81)	(86.89)
$GDPCAP_{ijt}$	**1.04**	**0.69**	**0.82**	**0.81**
	(19.72)	(23.23)	(26.44)	(27.61)
RTA	0.01	**0.19**	**0.16**	**0.16**
	(0.09)	(4.82)	(4.82)	(4.74)
EMU	0.10	-0.10	-0.07	-0.06
	(0.56)	(-1.03)	(-0.74)	(-0.69)

Note: Bold letters indicate significance at a 5% level.

Let us first analyze the upper part of Table 4, where we present the results obtained when we use the variables defined as they commonly are in the empirical literature. We transform them to account for the deterministic components and the structural break found in 1998, at 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.

In Table 5 we compare our results with previous findings in this literature. First, we include summary-results from Gil-Pareja et al. (2008). We have chosen this paper as a benchmark of mainstream classical panel gravity equation on the euro effect on trade. Second, follow Gengenbach (2009) who presents a summary of the main results that he obtains using the CUP estimator and Pesaran's (2006) Common Correlated Effects Pooled estimator (CCEP) and compares them with those found by Bun and Klaassen (2002, 2007) LSDV and DOLS . We concentrate on the trended versions from the different papers that use panel cointegration techniques. These papers are not directly comparable, as they use different databases. However, all of them have in common that the definition of the variables is the standard one in the literature. Bearing this caveat 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). Third, concerning the traditional panel paper of Gil-Pareja et al. (2008), they obtain a quite large effect for EMU and, more importantly, this effect is bigger than the one found for previous economic integration stages. Recent papers using cointegration techniques show a much smaller EMU effect. Moreover, once the trend of the integration process is taken into account, the final effect of EMU is non-existent or just a minor one. Therefore, our results

are in line with this last group of empirical papers. Additionally, in an attempt to refine the previous output, we repeat the exercise with the variables constructed "à la Baldwin".

Table 5.
Standard-variable results and comparison with previous results in the literature

Variables	Gil-Pareja et al. (2008)	Bun and Klaasen (2007)		Faru-qee (2004)	This article		Gengenbach (2009)		This article	
	Classic Panel tech-niques	Panel cointegration techniques								
		Cross-section independence					Cross-section dependence			
		LSDV	DOLS	DOLS	LSDV	DOLS	CUP	CCE	CUP_{BC}	CUP_{FM}
GDP_{ijt}	1.20	0.70	0.94	1.74	**0.80**	**0.74**	0.36	0.33	**-0.95**	**-1.01**
$GDPCAP_{ijt}$	1.15	-0.23	-0.49	-0.73	**1.40**	**0.71**	0.18	0.091	**2.84**	**2.81**
RTA	0.41	0.16	0.05	0.012	**-0.13**	**0.28**	0.016	0.011	**0.02**	**0.03**
EMU	0.53	0.032	0.034	0.073	**-0.43**	**0.198**	0.075	0.006	**-0.07**	**-0.09**

Note: Bold letters indicate significance at a 5% level.

The lower files of Table 4 contain the results obtained with the variables constructed according to Baldwin and Taglioni's critiques (2006). The model has been estimated with six common factors, as was derived from the Banerjee and Carrion-i-Silvestre (2010) analysis. We should first mention that the estimates obtained are very similar; no matter the estimator chosen the EMU dummy is non-significant. In contrast, the RTA one is significant with the only exception of LSDV.

Concerning the GDP variables, the values obtained are around 1.5 and 0.8, respectively. The only discrepancy is found using the LSDV estimator. We should note that this estimator is shifted away from zero due to an asymptotic bias induced by the cross-section dependence. The two significant estimated coefficients obtained using LSDV are much larger than with the other estimators due to the above-mentioned upward bias. 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 CUP-FM and CUP-BC when there is dependence and the common factors are non-stationary. When using CUP-BC estimator

(last column of the table) the two GDP measures have positive parameters of 1.47 and 0.81, whereas RTA, the regional trade agreements dummy parameter is 0.16 and highly significant. The reason that may explain this relatively small value is that we found that the common structural break in both the trend and the intercept occurs in 1987, the date when the Single Act was approved. Thus, the majority of the bilateral effects represented by this dummy could have been already captured by the structural breaks and the remaining deterministic components already included in the regression. We convey with Gengenbach (2009) about the importance of a proper specification of the deterministic components in the gravity equation.

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

4. Summary and concluding remarks

In this paper we try to fill the gaps present in the previous literature on euro effects on trade. Using a 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, this is the first attempt to jointly incorporate Baldwin's critiques (in terms of model specification and variables' construction), the hypothesis of cross-sections dependence and structural breaks in the time domain within the estimation of a gravity equation on non-stationary series. This approach allows us to put the adoption of the euro by EMU members in historical perspective. We argue that the creation of the EMU is best interpreted as a 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.

Concerning the results, the variables that have been constructed following Baldwin's critiques (both in terms of multilateral resistance and unobserved bilateral heterogeneity) 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, as well as a dummy variable on trade agreements, 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 non-significance of the EMU dummy in the long-run estimated relationship.

Acknowledgements

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