Long-run evidence on EMU trade using gravity equations

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

***Keywords*:** Gravity models; exports; euro; panel cointegration; structural breaks, cross-section dependence.

***JEL classification*:** C12, C22, F15, F10.

1. Introduction.

The effect that a currency union has on trade has been largely explored in the literature. Rose (2000) is one of the most cited articles in this field, and his prediction of a tripling of trade for a country when it joins a currency union has been revisited several times. Moreover, the creation of EMU has provided researchers a natural experiment to further investigate on this effect, thus renewing the debate and leading to improvements in both the specification and estimation of the gravity equation. Although initial estimates were found to be quite high, ranging from the approximate 2% in Glick and Rose (2002) to the 27% in Barr et al. (2003), more recent literature has considerably reduced this effect. Thus, as Berger and Nitsch (2008) claim, the creation of EMU is best interpreted as a continuation of a series of policy changes that have led over the last decades to greater economic integration among the EMU countries, the euro having just a residual effect. Other articles supporting this hypothesis are Bun and Klaasen (2007), Fidrmuc (2009), Gengenbach (2009) and Camarero et al. (2011). Finally, Hiller and Kruse (2010) provide an analysis of this integration process, revealing the most relevant dates in the integration process for each one of the EMU countries.

The solution proposed by Berger and Nitsch (2008) to properly analyze the long-run dimension of the process is to include a time trend in the specification. A further step is given by Bun and Klaasen (2007) with the introduction of country-pair specific time trends that capture the impact of all omitted trending variables with a coefficient that is allowed to vary for each pair of countries. Both articles show that the inclusion of a deterministic trend notably reduces, or even eliminates, the euro effect on trade; however, both ignore the potential existence of stochastic trends in the data. The euro is a long-run process; hence, we claim that long-run estimation methods are more appropriate to measure its effect. Therefore, the nonstationarity of variables or the existence of cointegration relationships among them should be controlled for to avoid biases and inconsistencies. For that reason, the use of cointegration techniques and the inclusion of time trends –both deterministic and stochastic- is a necessary step in the analysis of the euro effect.

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

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

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

The remainder of the article is organized as follows. In Section 2 we describe the data and the variables used in the analysis, as well as the methodology and tests employed. In Section 3 we present the results for the EMU as a whole. In Section 4, two analyses are accomplished; first, we estimate a gravity equation for each of the EMU-members vis-à-vis the other eleven partners and we study the euro effect country by country; second, the same strategy is replicated for the analysis of EMU members' exports to third countries. Finally, Section 5 concludes.

2. Data, methodology and empirical results.

2.1. Data and model

We include in our study all the countries that joined the EMU in 1999 plus Greece, which became a member in 2001. Belgium and Luxembourg are included as an only area, so the total number of individuals is 11[1]. The sample contains annual data and covers the period 1967-2008. Hence, we have a balanced panel with dimension N =110 (11x 10, all possible bilateral combinations of countries) and T = 42. The total number of observations is NT =4,620. In a second step, we study the ex-

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

ports of these 11 countries to 15 OECD countries that do not belong to the EMU[2] and China; so we have a panel with dimension N =176 (11x16) and T =42. Although the number of years available was higher, we have opted by restricting our sample to this period, in order to exclude the effects of the financial crisis that started in 2008. Following Baldwin and Taglioni (2006) critiques', the variables are introduced in nominal terms. Descriptive statistics are presented in Appendix A.

We use a specification of the gravity equation similar to the specification from the recent literature on the euro effect using nonstationary panels:

$$EXPORTS_{ijt} = \beta_1 GDP_{it} + \beta_2 GDP_{jt} + \delta_1 RTA_{ijt} + \delta_2 EURO_{ijt} + \eta_{ij} + \tau_{ij} \cdot t + \varepsilon_{ijt} \qquad (1)$$

The dependent variable is $EXPORTS_{ijt}$, defined as the log of the export flows from country i to country j in nominal terms. As Baldwin and Taglioni (2006) point out, the gravity equation is an expenditure function that explains uni-directional bilateral trade flows; hence we include exports instead of the average of exports and imports. GDP_{it} and GDP_{jt} are logarithms of the nominal GDPs - instead of real terms, according to Baldwin and Taglioni (2006)'s critiques- in the exporter and importer country respectively, obtained from the CHELEM – CEPII database. Additionally, two dummy variables have been built to include the effect of particular integration agreements on trade. Namely RTA_{ijt} which is 1 if both countries have a free trade agreement at time t and 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. When analyzing the euro effect with third countries, this variable takes value one when one of the countries involved in the trade flow uses the euro. Our purpose is to isolate the effects of EMU trying to control for other factors that may have an influence on exports but are not related to the monetary union. □*ij* is a comprehensive set of country-pair specific dummies that captures all those bilateral time-invariant unobserved characteristics. We do not include any term to capture the unobserved time effects since the estimators that we will use already include a common factor structure.

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

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

2.2. Panel unit root tests and cross-section dependence

Two important aspects should be taken into account prior to the estimation of the gravity equation. First, it is highly probable that the series are interrelated among them, since the countries in the sample are members of a highly integrated area; the European Union. For that reason we implement the Pesaran (2004) CD test. Under the null hypothesis of cross section independence the CD statistic converges to the standard normal distribution. This test is valid for N and T tending to ∞ in any order and that is particularly useful for panels with small T and large N. In addition, this test is also robust to possible structural breaks, which makes it especially suitable for our study. The results in the first column of Table 1 show that the null hypothesis of independence is strongly rejected both in the case of intra EMU exports and in the case of EMU exports to third countries; hence cross-section dependence should be considered when computing the panel data statistics.

The second important point is the presence of unit roots in the data, which if unaccounted for may lead to wrong conclusions and biased estimates. Thus we apply Pesaran CADF (2007) and Bai and Ng (2004) tests to control for both aspects simultenously: unit root and cross-section dependence. Pesaran suggests to augment the Im, Pesaran and Shin (2004) 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. The test is based on the mean of individual ADF t-statistics of each unit in the panel:

$$\Delta Y_{ij} = a_i + b_i Y_{i,t-1} + c_i \overline{Y}_{t-1} + d_i \Delta \overline{Y}_t + \varepsilon_{it} \qquad i = 1,\dots,N; \quad t = 1,\dots,T \tag{2}$$

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

$$\overline{CADF} = N^{-1}\sum_{i=1}^{N} CADF_i \tag{3}$$

where $CADF_i$ is the t-statistic of b_i in the previous regression. The second column of Table 1 summarizes the results of the Pesaran CADF test. The null hypothesis of non-stationarity is rejected in all cases.

Table 1. Pesaran's CD and CADF statistics

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

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

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

$$Y_{it} = D_{it} + F_t^{'}\pi_i + e_{it} \tag{4}$$

with $t = 1, \ldots, T$, $i = 1, \ldots, N$, 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 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 null 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. Table 2 shows the results of this test. The idiosyncratic component is found to be nonstationary for the GDP variables, though stationary for exports. The results of the factor component analysis point also in the same direction; the null hypothesis of independent stochastic trends cannot be rejected in any of the cases. Hence, we have enough evidence to conclude that the variables are non-stationary and that cross-section dependencies are present in our data.

Table 2- Panel Data Statistics based on Approximate Common Factor Models
Bai and Ng (2004) statistics

Panel A: Intra-EMU						
Bai and Ng (2006) statistics						
	Exports$_{ijt}$		***GDP$_{it}$***		***GDP$_{jt}$***	
	Test	p-value	Test	p-value	Test	p-value
Idiosyncratic ADF statistic	-0.438	0.33	4.856	0.99	4.856	0.99
	Test	$\hat{r}_1$	Test	$\hat{r}_1$	Test	$\hat{r}_1$
MQ test (parametric)	-40.016	5	-33.766	6	-33.766	6
MQ test (non-parametric)	-40.591	5	-35.338	6	-35.338	6
Panel B: Third countries						
Bai and Ng (2006) statistics						
	Exports$_{ijt}$		***GDP$_{it}$***		***GDP$_{jt}$***	
	Test	p-value	Test	p-value	Test	p-value
Idiosyncratic ADF statistic	-2.04	0.02	4.856	0.99	-2.259	0.01

	Test	$\hat{r}_1$	Test	$\hat{r}_1$	Test	$\hat{r}_1$
MQ test (parametric)	-38.804	6	-33.766	6	-25.995	6
MQ test (non-parametric)	-39.182	6	-35.338	6	-26.257	6

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

2.3. Evidence of structural breaks in the European Integration Process

The next step in our empirical strategy is to test whether GDP_{it}, GDP_{jt} and $EXPORTS_{ijt}$ are cointegrated using Banerjee and Carrión-i-Silvestre (2010) test[3]. They propose a panel test for the null hypothesis of no cointegration allowing for breaks both in the deterministic components and in the cointegrating vector that also accounts for the presence of cross-section dependence using factor models. It is worth noticing that inference concerning the presence of cointegration can be affected by misspecification if the existence of breaks is ignored. In Table 3 we present the results of the cointegration tests for the model with homogeneous structural breaks for the eight potential specifications discussed above. In the left-hand side, the results of the intra-EMU exports are shown, whereas the right hand side provides the results of EMU exports to third countries. Using the BIC_3 information criterion of Bai and Ng (2002)[4] we choose the specification 5 in both cases, which contains a constant and a trend and a structural break that affects them both simultaneously. In order to test for non-cointegration, we apply the statistics based on the accumulated idiosyncratic components, Z_j^*. We present the tests for all possible specifications; in all cases the null hypothesis of non-cointegration is rejected. The break is found to happen in 1987 -the year of the signature of the Single European Act (SEA) - for intra-EMU trade and in 1989 for EMU trade with third countries. Although the assumption of a common break for all country pairs might seem a little restrictive, however, the homogeneity of the sample -we include only EMU and OECD countries and China- is enough to find a reasonable break common to all country pairs.

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[3] See appendix B for further information about the test.

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

Table 3: Banerjee and Carrion (2010) BC cointegration tests

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

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

Finally, given that the existence of cointegration relationships is unambiguous, the next step is to estimate the long-run relationship in the form of a gravity equation. For this purpose, in the next section we will employ consistent techniques proposed by Bai et al (2009). We allow the coefficients of the trend as well as the coefficients of the structural breaks in the trend to be different for each pair of countries, thus introducing a higher degree of heterogeneity in the model[5].

3. Estimation of the gravity equation for the EMU

Traditional estimation methods as Ordinary Least Squares (OLS) or the Least Squares Dummy Variables (LSDV) approach present biases and inconsistencies in the presence of nonstationarities and cointegration relationships among the variables. The Fully Modified (FM) estimator of Phillips and Hansen (1990) and the Dynamic Ordinary Least Squares (DOLS) estimator proposed by Saikkonen (1991) and Stock and Watson (1993) are some of the alternatives employed in the literature. However, although both estimators consistently estimate the long-run parameters and correct for autocorrelation and endogeneity, they do not account for dependence. Since the Pesaran

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

CD test has revealed the existence of dependencies among the series, alternative estimators should be employed.

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

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

$$y_{it} = x_{it}^{'}\beta + e_{it} \tag{5}$$

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

$$x_{it} = x_{it-1} + \varepsilon_{it} \tag{6}$$

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

$$\hat{\beta}_{LS} = \left(\sum_{i=1}^{n} \sum_{t=1}^{T} x_{it} x_{it}^{'} \right)^{-1} \sum_{i=1}^{n} \sum_{t=1}^{T} x_{it} y_{it} \tag{7}$$

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

$$e_{it} = \lambda_i^{'} F_t + u_{it} \tag{8}$$

where F_{it} is an $r \times 1$ vector of latent common factors, λ_i is an $r \times 1$ vector of factor loadings and u_{it} is the idiosyncratic error. If both F_t and u_{it} are stationary, then e_{it} is also stationary. In this case, a consistent estimator of the regression coefficients can still be obtained even when the cross-section dependence is ignored. Though, it is crucial to note that when F_t is $I(1)$, if $F_t = F_{t-1} + \eta_t$ then e_{it} is $I(1)$ and the pooled OLS in (7) is not consistent. This is why Bai et al. (2009) develop the case of non-stationary common factors, aiming at achieving consistent estimators. Let the true model in vector form be

$$y_i = x_i \beta^0 + F^0 \lambda_i^0 + u_i \tag{9}$$

where

$$y_i = \begin{bmatrix} y_{i1} \\ y_{i2} \\ \cdots \\ y_{iT} \end{bmatrix}, \; x_i = \begin{bmatrix} x_{i1}^{'} \\ x_{i2}^{'} \\ \cdots \\ x_{iT}^{'} \end{bmatrix}, \qquad F = \begin{bmatrix} F_1^{'} \\ F_2^{'} \\ \cdots \\ F_T^{'} \end{bmatrix}, \quad u_i = \begin{bmatrix} u_{i1} \\ u_{i2} \\ \cdots \\ u_{iT} \end{bmatrix} \tag{10}$$

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 $\widetilde{\beta}_{LSFM}$ and the bias corrected version that is identical. The estimators are consistent and the limiting distributions are normal. However, in the majority of the cases, the factors F_t are unobserved and the LSFM estimator is infeasible. In this case F_t should be estimated along with β by minimizing the objective function

$$S_{nT}(\beta, F, \Lambda) = \sum_{i=1}^{n} (y - x_i \beta - F\lambda_i)^{'} (y - x_i \beta - F\lambda_i) \tag{11}$$

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

$$\hat{\beta}\left(\sum_{i=1}^{n} x_i^{'} M_F x_i\right)^{-1} \sum_{i=1}^{n} x_i^{'} M_F x_i \tag{12}$$

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

$$S_{nT}(\beta, F) = \sum_{i=1}^{n} (y - x_i\beta)^{'} M_F (y - x_i\beta) \tag{13}$$

the continuously-updated estimator (CUP) for (β, F) would be

$$(\hat{\beta}_{Cup}, \hat{F}_{Cup}) = \arg\min_{\beta, F} S_{nT}(\beta, F) \tag{14}$$

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

$$\hat{\beta} = \left(\sum_{i=1}^{n} x_i^{'} M_F x_i\right)^{-1} \sum_{i=1}^{n} x_i^{'} M_F x_i \tag{15}$$

$$\hat{F} V_{nT} = \left[\frac{1}{nT^2} \sum_{i=1}^{n} (y_i - x_i\hat{\beta})(y_i - x_i\hat{\beta})^{'}\right] \hat{F} \tag{16}$$

where $M_{\hat{F}} = I_T - T^{-2}\hat{F}\hat{F}^{'}$ since $\hat{F}^{'}\hat{F}/T^2 = I_r$, and V_{nT} is a diagonal matrix consisting of the r largest eigenvalues of the matrix inside the brackets, arranged in decreasing order. The estimator is obtained solving for $\hat{\beta}$ and $\hat{F}$ using (15) and (16) and it is consistent for β, although it still has a bias derived from having to estimate F_t. The authors correct this bias using two fully-modified estimators. The first one directly corrects the bias of $\hat{\beta}_{Cup}$ and is denoted $\hat{\beta}_{CupBC}$. The second one makes the correction in each iteration and is denoted $\hat{\beta}_{CupFM}$.

We present in next section the results of the CUP estimation using the methodology of Bai et al. (2009), as well as the Bai FM results for the sake of comparison. We have selected the specification according to the results of Banerjee and Carrión-i-Silvestre (2010) tests. In order to account for the existence of incidental trends (intercept and/or trend), Bai et al. (2009) define accordingly the projection matrix *M* considered above for demeaned and/or detrended variables. We concentrate the deterministic components by filtering the five variables in the equation before estimating the long-run parameters. Among the deterministic components we include the constant, the country pair specific trends, the common break in the constant and the common break in the country pair specific trends. The number of common factors for the estimation is selected according to Principal Components Factor Analysis (PCA henceforth).

4. Results

4.1 Intra EMU trade

In a first step, equation (1) is estimated including exports flows among the 11 EMU countries included in the sample. Table 4 shows the results of the estimation using CUP-FM and CUP-BC estimators. Bai FM is also included for the sake of comparison. As expected, the exporter and importer GDPs have a positive influence on exports in all cases. The importer GDP has higher effect than the exporter's, indicating that demand has greater influence on exports than supply. The RTA coefficient is positive and statistically significant in all cases. It is worth noticing that this variable is already capturing the effect of joining to the European Free Trade Agreement (EFTA) or the European Union (EU). The EMU coefficient is also positive and highly significant, and its magnitude is 0.13 and 0.15 using both CUP estimators; this implies that the adoption of the euro has increased exports between EMU members by about 13% and 16% respectively[6]. Table C.1 in Appendix summarizes previous estimates of euro effect in the literature. As remarked in the table, our results reduce the initial optimistic coefficients and are in line with more recent literature.

Table 4
Estimation of the long-run parameters 1967-2008 for intra EMU trade

[6] To interpret dummy coefficients as a percentage change we need to apply a simple transformation to the coefficient obtained, 100*[EXP(α)-1].

Variables	Bai FM	CupBC	CupFM
GDP_{it}	**0.35**	**0.50**	**0.50**
	(13.43)	(19.07)	(18.56)
GDP_{jt}	**0.93**	**1.17**	**1.17**
	(35.16)	(44.45)	(44.45)
RTA	**0.19**	**0.12**	**0.12**
	(13.88)	(8.78)	(9.06)
EMU	**0.19**	**0.13**	**0.15**
	(15.15)	(11.12)	(12.25)

Notes: Bold letters indicate significance at a 5% level. The specification 5 is estimated with 2 common factors according to Principal Components Analysis. The t-statistic is reported in parenthesis. The common structural break takes place in 1987. EMU takes value one when both countries belong to EMU. The bandwidth parameter is 0.25 according to Silverman's rule of thumb. Results with a different number of factors and/or bandwidth are available under request

Next, we proceed to the analysis of each country separately. To assess which members have obtained larger benefits from joining the euro and to find possible asymmetries, we have constructed 11 additional sub-panels in which the exporter is each one of the EMU countries and the exporters are the other 10 remaining members. Hence, we have 11 sub-panels with dimension T =42 and N =10. The empirical strategy followed for each one of these sub-panels is analogous to the one previously employed. In a first step, we have checked the existence of dependencies among the series, as well as the nonstationarity of the variables[7]. Since we have found evidence of both facts, in a second step we have tested the existence of cointegration relationships among the variables using the Banerjee and Carrión-i-Silvestre (2010) test. The results are again positive and the specification 5 is selected among all possible specifications according to the BIC_3 criterion of Bai and Ng (2002) in all cases with the exception of Ireland, which also has a break in the cointegrating vector (specification 8). Table 5 shows the coefficient of our variable of interest, as well as the date of the break for each country. According to the CUP estimator, the euro effect is found to be negative and significant in the cases of Finland, Ireland and Greece, non-significant for the Netherlands; and positive and significant for the rest of the countries. A tentative explanation for the negative sign

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

could be the fact that these countries got used to depreciate their currency to foster exports before 1999, while after the introduction of the euro they could not use this strategy anymore. More specifically, Finland faced a commercial crisis after the demolition of the URSS, his main commercial partner, which implied consecutive devaluations after 1990. In the case of Greece, the "hard drachma" policy[8] adopted in 1995 implied a notable appreciation of the drachma during the period 1995-1997. Later on, when Greece joined the ERM in 1998 the currency experienced a devaluation of 12.3%.

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

Table 5
Country comparison of the EMU effect. Intra EMU trade

Country	Bai FM	CupFM	CupBC
Austria	**0.28**	**0.20**	**0.20**
1994	(10.78)	(7.93)	(7.90)
Belgium	**0.35**	**0.25**	**0.36**
1993	(15.67)	(12.56)	(18.02)
Finland	**-0.32**	**-0.30**	**-0.44**
1987	(-11.15)	(-11.03)	(-15.79)
France	**0.14**	**0.31**	**0.24**
1974	(8.69)	(16.23)	(13.99)
Germany	**0.21**	**0.07**	**0.18**
1974	(13.92)	(6.42)	(15.98)
Greece	-0.06	**-0.07**	**-0.09**
1980	(-1.54)	(-2.01)	(-2.57)
Ireland	**-0.25**	**0.14**	**-0.06**
1973	(-8.66)	(5.63)	(-2.16)
Italy	**0.19**	**0.26**	**0.21**

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

1985	(11.26)	(14.80)	(12.15)
Netherlands	**0.11**	**-0.13**	-0.02
1975	(6.87)	(-7.76)	(-1.20)
Portugal	**0.43**	**0.15**	**0.34**
1984	(11.66)	(4.19)	(9.91)
Spain	**0.35**	**0.13**	**0.18**
1989	(9.08)	(3.63)	(5.01)

Notes: The specification 5 is estimated with 1 or 2 common factors according to PC factor analysis. The t-statistic is reported in parenthesis. The year of the break is indicated below the name of each country. Bold letters indicate significance at a 5% or 10% level. Bandwidth parameter is 0.25 according to Silverman's rule of thumb. Results with a different number of factors and/or bandwidth are available under request.

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

4.2. EMU trade with third countries

The third objective in our analysis is to assess the euro effect on trade with non-EMU countries. We have included the same EMU exporters, but now we focus on their exports to 16 non-member countries. Now EMU_{ijt} takes value one when one of the countries (not the two) involved is an EMU-member.

For the estimation of the aggregated effect we have a panel with N = 176 individuals and T =42 years. In addition, for the estimation of the effect for each EMU country we have constructed 11 additional sub-panels with dimension N = 16 and T =42. We have performed the same empirical strategy to check the existence of nonstationarity and cointegration, obtaining again positive evidence[9]. Table 6 shows the results of the analysis for the aggregated database. As before, there are no substantial differences between BaiFM and CUP estimators. Both importer and exporter GDP show a positive and significant coefficient, as expected, and again the importer GDP has higher importance. The structural break for the aggregated dataset is found to happen in 1989, very close to the signature of the Single European Act. In this case the countries included are less related to EMU process; hence this break-date may be related to the Plaza (1985) and Louvre (1987) agreements, which were signed with the objective to stabilize the international currency markets. Belonging to a RTA has an unambiguous positive effect on exports. Although EMU shows now a positive but not significant coefficient at the aggregate level, a closer inspection of the EMU effect on third countries reveals in Table 7 that at this effect is generally significant in each individual case.

Table 6
Estimation of the long-run parameters 1967-2008. EMU with third countries

Variables	Bai FM	CupFM	CupBC
Dependent variable: $EXPORTS_{ijt}$			
$\boldsymbol{GDP_{it}}$	**0.53**	**0.53**	**0.53**
	(9.11)	(9.09)	(9.04)
$\boldsymbol{GDP_{jt}}$	**0.70**	**0.71**	**0.71**
	(11.36)	(11.36)	(11.42)
RTA	**0.29**	**0.30**	**0.29**
	(4.52)	(4.62)	(4.53)
EMU	0.04	0.04	0.04
	(0.84)	(0.87)	(0.77)

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

Table 7 shows the coefficient of EMU variable and the date of the break for each sub-panel. Break dates are found to happen in dates very close to the oil shocks in 1973 and 1979 or the Plaza and Louvre agreements, which makes more sense since these facts are more prone to affect international trade (not only EMU countries). In two cases – Finland and Italy - the break is found

[9] Results available upon request.

already in the ninetees.

Table 7
Country comparison of EMU effect with third countries

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

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

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

5. Summary and concluding remarks

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

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

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

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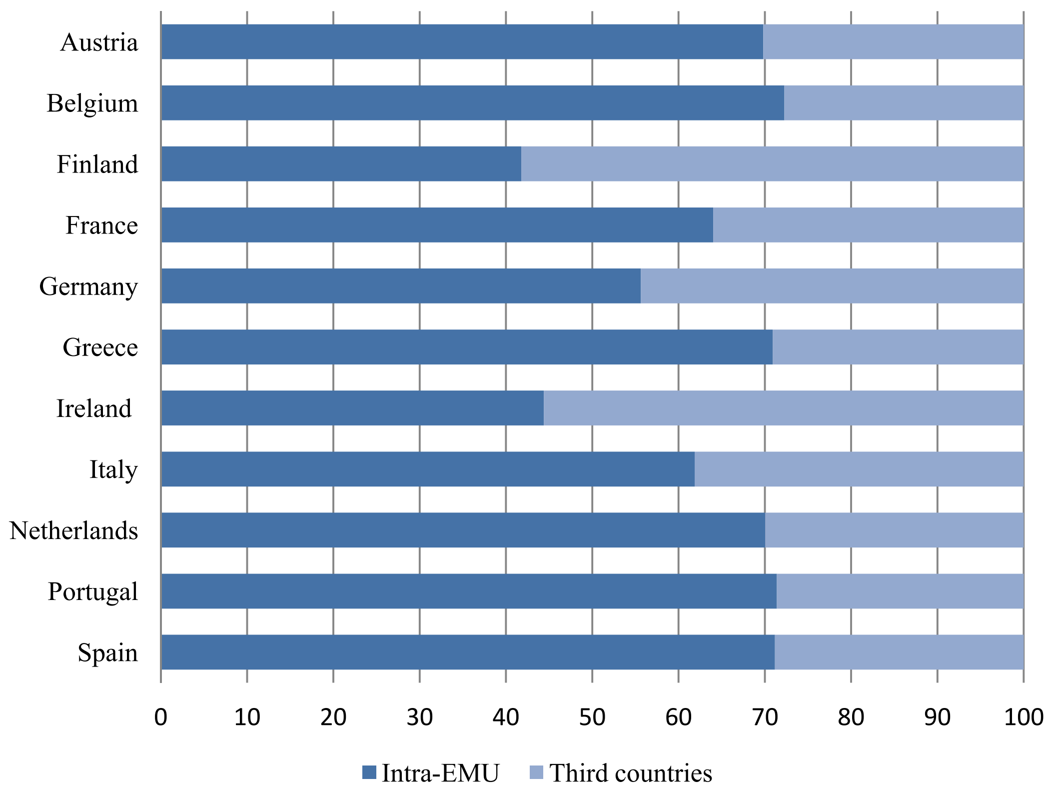
APPENDIX A. Descriptive statistics

Table A1. Descriptive statistics

Intra EMU					
Panel A: Basic statistics					
Variable	**Mean**	**Std. Dev.**	**Min**	**Max**	**Observations**
Log of Exports	34.24	2.31	26.39	39.38	4620
Log GDP of exporters	11.91	1.52	8.09	15.11	4620
Log GDP of importers	11.91	1.52	8.09	15.11	4620
RTA	0.59	0.49	0	1	4620
EMU	0.23	0.42	0	1	4620
Third countries					
Panel A: basic statistics					
Variable	**Mean**	**Std. Dev.**	**Min**	**Max**	**Observations**
Log of Exports	32.91	2.87	0	39.13	7392
Log GDP of exporters	11.91	1.52	8.09	15.11	7392
Log GDP of importers	11.96	1.87	6.03	16.47	7392
RTA	0.15	0.36	0	1	7392
EMU	0.23	0.42	0	1	7392

Fig. A1 Exports by destination



Note: Three main partners of each country are sorted in order of importance. Source: own elaboration based on CHELEM database.

Appendix B: Banerjee and Carrion-i-Silvestre (2010) test

Banerjee and Carrion-i-Silvestre (2010) propose a panel test for the null hypothesis of no cointegration allowing for breaks both in the deterministic components and in the cointegrating vector and also accounts for the presence of cross-section dependence using factor models. They define a *(m x 1)* vector of non-stationary stochastic process, $Y_{it} = (y_{it}, x_{it}^{'})$ whose elements are individually $I(1)$ with the following Data Generating Process:

$$y_{it} = D_{it} + x_{it}^{'}\delta_{it} + u_{it} \tag{A.1}$$

The general functional form for the deterministic term D_{it} is given by:

$$D_{it} = \mu_i + \beta_i t + \sum_{j=1}^{m_i}\theta_{ij} DU_{ijt} + \sum_{j=1}^{m_i}\gamma_{ij} DT_{ijt} \tag{A.2}$$

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

$$\delta_{it} = \begin{cases} \delta_{i1} T_{i0}^{c} < t \le T_{i1}^{c} \\ \delta_{i2} T_{i1}^{c} < t \le T_{i2}^{c} \\ \cdots \\ \delta_{ij} T_{ij-1}^{c} < t \le T_{ij}^{c} \\ \cdots \\ \delta_{in_i+1} T_{in_i}^{c} < t \le T_{in_i+1}^{c} \end{cases} \tag{A.3}$$

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

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

Model 1. Constant term, no linear trend - $\boldsymbol{\theta_{ij} = \beta_i = \gamma_{i,j} = 0}$ $\forall i, j$ in (A.2) – and constant cointegrating vector.

Model 2. Stable trend - $\boldsymbol{\Theta_{ij} = 0; \beta_i \neq 0}$ $\forall i$ and $\boldsymbol{\gamma_{i,j} = 0}$ $\forall i, j$ in (A.2) – and constant cointegrating vector.

Model 3. Constant term with shifts; stable trend - $\boldsymbol{\Theta_{ij} \neq 0; \beta_i \neq 0; \gamma_{i,j} = 0}$ $\forall i, j$ (A.2) – and constant cointegrating vector. The model considers multiple level shifts.

Model 4. Constant term, trend and changes in trend, - $\boldsymbol{\Theta_{ij} = 0; \beta_i \neq \gamma_{i,j} \neq 0}$ $\forall i, j$ in (A.2) – and constant cointegrating vector. The model considers multiple trend shifts.

Model 5. Changes in constant and trend - $\boldsymbol{\Theta_{ij} \neq 0; \beta_i \neq 0}$ and $\boldsymbol{\gamma_{i,j} \neq 0}$ $\forall i, j$ in (A.2) – and constant cointegrating vector. The model considers multiple trend and level shifts.

Model 6. No trend, constant term with shifts - $\boldsymbol{\Theta_{ij} \neq 0; \beta_i = 0}$ $\forall i, j$ in (A.2) – and changes in the cointegrating vector.

Model 7. Constant term, trend; changes in the level - $\boldsymbol{\Theta_{ij} \neq 0; \beta_i \neq 0}$ $\forall i, j$ in (A.2) – and changes in the cointegrating vector.

Model 8. Constant term, trend; changes in the level and the trend - $\boldsymbol{\Theta_{ij} \neq 0; \beta_i \neq 0}$ and $\boldsymbol{\gamma_{i,j} \neq 0}$ $\forall i, j$ in (A.2) – and changes in the cointegrating vector

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

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

The null hypothesis of a unit root can be tested using the pseudo *t*-ratio $t^{j}_{\widetilde{e}_i}(\lambda_i)$, $j = c, \tau, \gamma$. The models that do not include a time trend (Models 1 and 6) are denoted by *c*. Those that include a linear time trend with stable trend (Models 2, 3 and 7) are denoted by τ and, finally, γ refers to the models with a time trend with changing trend (Models 4, 5 and 8). When common (homogeneous) structural breaks are imposed to all the units of the panel (although with different magnitudes), we can compute the statistic for the break dates, where the break dates are the same for each unit, using the idiosyncratic disturbance terms.

Appendix C

Table C.1. Euro effect in previous literature

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

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