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ROSE EFFECT VERSUS BORDER EFFECT

by Gianluca Cafiso

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In 2008 all ECB publications feature a motif taken from the €10 banknote.

1 The views expressed are those of the author and do not necessarily reflect the views or policies of the European Central Bank. I thank Rodolfo Helg for the helpful comments and support, and the anonymous Referee for her remarks to a previous version of this paper.
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Abstract

This paper assesses the Euro's influence upon European trade by estimating two different indicators. The first is the so-called "Rose Effect", while the second is the "Border Effect". The former measures how much a country within a currency union trades more with its partners than with non-member countries, the latter measures the integration of a country with its trade partners. This study of the Euro's influence by means of the Border Effect is a novelty in the literature, it reveals that the Euro's influence upon trade is not so clear as papers focused only on the Rose Effect claim. This casts doubts about the consequences of the Euro introduction for the European Single Market. Both indicators are estimated by means of a gravity model for bilateral trade flows using a panel of manufacture exports among twenty-four OECD countries.

Keywords: Euro, European Integration, Trade, Rose Effect, Border Effect.

JEL Classification: F10, F14, F15.

Non-Technical Summary

One of the reason to introduce a common currency among the European Union countries was its expected positive effect upon intra EU trade. Then, as soon as an econometrically usable time period from the Euro's introduction was available, its influence upon trade has become object of intense empirical research. The methodology employed was firstly used by Andrew Rose (2000) for his study on the effects of currency unions upon trade. Even if there is no consensus on the magnitude of the Euro's effect, many have agreed that it is significantly positive for the countries which adopted it in 1999. This positive evidence is mostly founded on *Rose Effect* estimates which use aggregate and industry-level bilateral trade data. The explanation of a such positive effect recalls the economic rationale at the basis of the monetary integration process, which consists in a trade-costs reduction and higher competition within the European Single Market.

This paper shows that the Euro's influence upon trade is not so clear as empirical studies based exclusively on the Rose Effect claim. This is done through the comparison of the Euro's effect as detected by the Rose Effect and as detected by an alternative indicator. This alternative indicator is the *Border Effect*, which measures integration within a group of trade partners. At our knowledge, there is no other paper which checks the effect of currency unions on trade through the estimation of the Border Effect. Then, the estimation of the Border Effect is meant as an alternative way to test the effect of the Euro which serves as a check of the positive Rose Effect found in the literature and in our estimations too. The Rose Effect and the Border Effect should provide unanimous evidence about the Euro's influence upon trade if it is true that the positive effect of the Euro, as detected by the Rose Effect, is due to a trade-costs reduction. Indeed, the lower the trade costs are, the lower the Border Effect is, the higher the integration among the countries in the group is. However, the evolution of the Border Effect does not show higher integration in the time period considered and in this sense it rules out a trade-costs reduction as explanation of the larger trade observed after 1999.

The Border Effect is estimated through the same dataset used for the estimation of the Rose Effect to which we have added the National Trade observations necessary for its estimation. Indeed, the use of the same dataset for both indicators is crucial in our analysis in order to compare them in a consistent way. The sample includes trade flows among twenty-four OECD countries in the period 1993-2003. The econometrics used is gravity panel econometrics, the estimates of the Rose Effect and the Border

Effect are robust with respect to different techniques which we discuss in a specific section.

In the last section of this paper we provide a possible explanation of the conflict between the effect of the Euro found through the two different indicators. In a nutshell, we believe that the Euro has not caused a significant trade-costs reduction among the Euro Zone countries in accordance to what the Border Effect shows. The positive increase in exchanges caught by the Rose Effect is to be explained through something different than a trade-costs reduction as recent studies on firms' productivity and competitiveness claim.

I - Introduction

The analysis of the effects of Currency Unions upon trade is a recent strand of research which has grown much in the last years. The interest in this field of research has increased much because researchers were eager to predict how much the just-introduced Euro would have boosted trade among the European countries. Indeed, Politicians and economists affirmed that a single currency would have caused higher trade and/or integration among the European countries thanks to a significant reduction in trade costs (see Emerson et al. 1992). A reduction in trade costs caused by: i) the end of exchange rate volatility, ii) the facility to operate with only one currency iii) higher market transparency and market competition, iv) higher macroeconomic stability (European Commission 2003, Micco et al. 2003, and Frankel and Rose 2002). Then, there was a theoretical motivation to check for the trade effects of the Euro.

The first paper to engineer a direct estimation strategy for the effects of currency unions upon trade is Rose (2000). Andrew Rose estimates a cross-section gravity model on a sample of 186 developing and poor countries, he finds that trade flows among countries in a common currency union are on average three times higher than those among countries which are not into any. Even though the magnitude of the effect was too high to be believable for the European countries, Rose's technique has become the workhorse of this empirical literature. Such technique consists in the introduction of a dummy variable, which controls for membership in a currency union, in the widely-used gravity model for bilateral trade flows. For ease, many researchers call now "Rose" effect this way to assess the effect of currency unions upon trade.

The purpose of this paper is to check the common finding of a positive Rose Effect (alias, a positive effect of the Euro upon trade) through an alternative indicator named Border Effect in order to verify if the Euro has really eased trade exchanges by reducing trade costs. If it is so, the Border Effect evolution has to be consistent with the finding of a positive Rose Effect. The contribution of this work to the literature comes from this comparison which suggests that a positive Rose Effect, if it is not a spurious result, cannot be explained by a reduction in border-linked trade costs. We consider this an important finding since all the previous works tend to explain the Rose Effect via a trade costs reduction. This paper flows into that group of papers (i.e. Baldwin 2006, De Nardis, De Santis and Vicarelli 2008) which are against a naïve interpretation of the Rose Effect and that try to gain an insight on the Euro's influence through a more sophisticated analysis.

To fully understand the comparison abovementioned, it is necessary to make clear what the Border Effect is. It is an indicator which quantifies how much trade within a country is higher than that country's average trade with its representative trade partner. It is used as an indicator of integration among a group of trade partners and it has an interpretation in terms of trade costs (Engel and Rogers 1996, Nitsch 2000, Chen 2004). The link between the Border Effect as a measure of trade costs and the Border Effect as an indicator of integration is straightforward. Given the trade costs in which goods incur when they cross the border (Anderson and van Wincoop 2004, Emerson et al. 1992), nationals prefer the consumption of domestic goods which are relatively less expensive than imported goods. Then, if the Euro had a positive effect on internal European trade (by eliminating exchange costs and so reducing overall trade costs), this should have caused a decrease of the BE which reflects more integration among the Euro Zone countries.

It is to remember that the Euro was meant as the completion of the European Single Market Programme, but many border costs were eliminated before the exchange one. This basically means that convergence among the European countries is an historical process (Berger and Nitsch 2005). Hence, when we look at the trend of the Border Effect over the last decades, we expect that it is downwards sloping (and so is it). Consequently, to check whether or not the Euro has had a positive effect in terms of European Integration, we will look at the trend line of the Border Effect after 1999. If after 1999 it is more downwards sloping than before 1999, we will deduce a positive effect of the Euro in terms of European Integration. Furthermore, if the Euro has caused more integration after 1999, the trend line of the Border Effect of Denmark, Sweden and the United Kingdom should be somehow affected (the EU countries which did not join the Euro).

This paper is organized as follows. In section II we describe the data used and the transformations necessary to make them operational; section III contains static estimations of the Rose Effect; in section IV we discuss the estimation of the Border Effect; in section V we discuss some robustness checks which are not included in the paper; in section VI we draw the conclusions and provide an explanation of the clash between the Border and Rose Effect. Some robustness checks are in the two appendices, we decided to keep those in the paper for their interesting features.

Appendix I contains dynamic estimations of the Rose Effect, while appendix II contains a robustness check of the Border Effect.

II - Data Description

The data used are mainly extracted from OECD Statistics, the time period is 1993-2003 for the following reasons. Firstly, we want to avoid a possible structural break in 1993 when the recording system of trade flows changed after 1992 due to the removal of EU internal customs. Secondly, no major integration action has been enforced from 1993 until the introduction of the Euro in 1999; this should allow us to estimate more accurately the Rose Effect after 1999.[1] Our dataset comprises the following variables and control dummies:

- Manufacture export between 24 OECD countries (sectors 15-37, ISIC Rev. 3). The countries are: Austria, Australia, Belgium-Luxembourg[2], Canada, Switzerland, Germany, Denmark, Spain, Finland, France, Greece, Iceland, Ireland, Italy, Japan, South-Korea, Mexico, the Netherlands, Norway, New Zealand, Portugal, Sweden, the United Kingdom, the United States. Export figures are extracted from OECD-Stan Bilateral Trade dataset in nominal US dollars.[3]
- Nominal GDPs in national currency, from OECD Economic Outlook 2005.
- Real Effective Exchange Rates from OECD Economic Outlook 2005 to account for the competitiveness of each country with respect to the others.
- Distance between countries, calculated through the Great Circle Formula, is extracted from the CEPII-Distance dataset.
- Control Variables for: 1) *Contiguity,* a dummy which controls for geographical contiguity between two trade partners; 2) *Language,* a dummy which controls for common spoken languages.

To make the data operational, we converted nominal exports and GDPs in real terms. This conversion is not as straightforward as it could seem. Indeed, it has to be decided if to estimate the model from the demand side (exports) or from the supply side (imports). Since we want to check if the Euro's introduction has favored

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[1] We acknowledge that more recent observations would have been useful. Unfortunately, at the date this version was completed, the OECD data used were available only up to 2003.

[2] Data for Belgium and Luxembourg are recorded together for the so-called Belgium-Luxembourg Economic Union (BLEU).

[3] The group of *EZ countries* includes: Austria, Belgium-Luxembourg, Germany, Spain, Finland, France, Greece, Ireland, Italy, Netherlands, Portugal. The group of *EU-non-EZ countries* includes: Denmark, Sweden and the United Kingdom. The group of *non-EU countries* includes: Australia, Canada, Switzerland, Iceland, Japan, South-Korea, Mexico, Norway, New Zealand, the United States.

European firms' export into countries which adopted the Euro, we decided to carry out our analysis focusing on the supply-side. Coherently, we divided export figures by the country-specific Producer Price Index.[4] GDPs are divided by country-specific GDPs deflators.

The end of the exchange rate volatility among the Euro Zone (EZ) countries has modified the EZ exporters' competitiveness in other EZ markets with respect to non-EZ exporters. For this reason, we included variables which account for the evolution of the competitive position of each exporter with respect to the others. In addiction, trade flows from country *i* to country *j* are likely to be affected by changes in the importer's competitive position as well, then we include a competitiveness index for the importer too.[5] The rationale is basically the same which supports the introduction of the Multilateral Resistance Terms, it is not the direct value of country *j*'s currency with respect to country *i*'s which matters, but country *j*'s relative exchange cost with respect to all its trade partners. As competitive indicators we use the Real Effective Exchange Rates.

III - Rose Effect Estimations

In this paper we only discuss estimations of the Rose Effect (RE) obtained using the full sample of twenty-four OECD countries. The motivation to consider bilateral trade not only within the EU is twofold. First, we want to check not only how the Euro has affected trade among the EU countries, but also how it has affected external trade towards the other OECD countries. Secondly, estimations on the OECD sample provide clearer evidence of the Euro's effect. This is so because the OECD sample is much larger and it includes trade flows across different Regional Trade Agreements (RTAs) and countries in different regional economic areas.[6]

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[4] We chose the Producer Price Index (PPI) and not the Consumer Price Index (CPI) or the GDP deflator because we consider only manufacture exports. Both the CPI and GDP deflator are computed on a larger basket of goods than the manufacture one. In Glick and Rose (2002) and Micco et al. (2003) trade figures are divided by a common price index (the US PPI); we deem this choice to be more appropriate if the analysis is focused on the demand-side since all the values are converted in the reference numeraire for the international consumer. On the contrary, we use country-specific indexes to maintain the time comparability for each exporter.

[5] To wit, kept constant the competitiveness of the exporter, if the importer's competitiveness changes (for instance its national currency depreciates or its price level increases) this is likely to affect its purchasing power into international markets by causing lower imports. At the same time, if the importer's currency value does not change with respect to the exporter's one, but appreciates with respect to a third country (another of its trade partners) this could cause lower exports from *i* to *j* , since *j* could substitute the now-relative-more-expensive imports from *i* with now-relatively-cheaper imports from the third country.

[6] The OECD24 countries which fall in a RTA are: EU Countries, Iceland and Norway in the EEA (European Economic Area); USA, Canada and Mexico in the NAFTA (North American Free Trade Area); Switzerland, Iceland and Norway in the EFTA (European Free Trade Area); Australia and New

In Figure 1 we plot the yearly Overall Total Export among the OECD24 countries, while Total Export by Country is in Figure 2 and Figure 3. Total Export by Country is computed as:

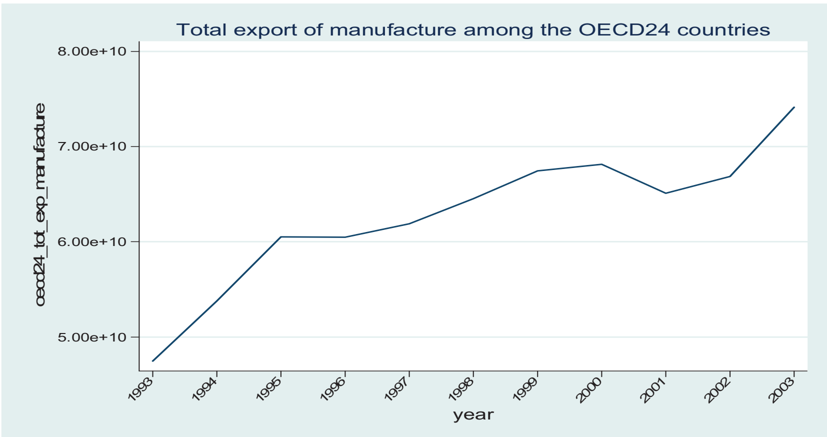
$$\exp_y^i = \sum_{j=1}^{23} \exp_y^{ij}$$

where $\exp_y^{ij}$ is export from i to j in the y^{th} year, while Overall Total Export is:

$$\exp_y^{OECD24} = \sum_{i=1}^{24} \exp_y^i \ .$$

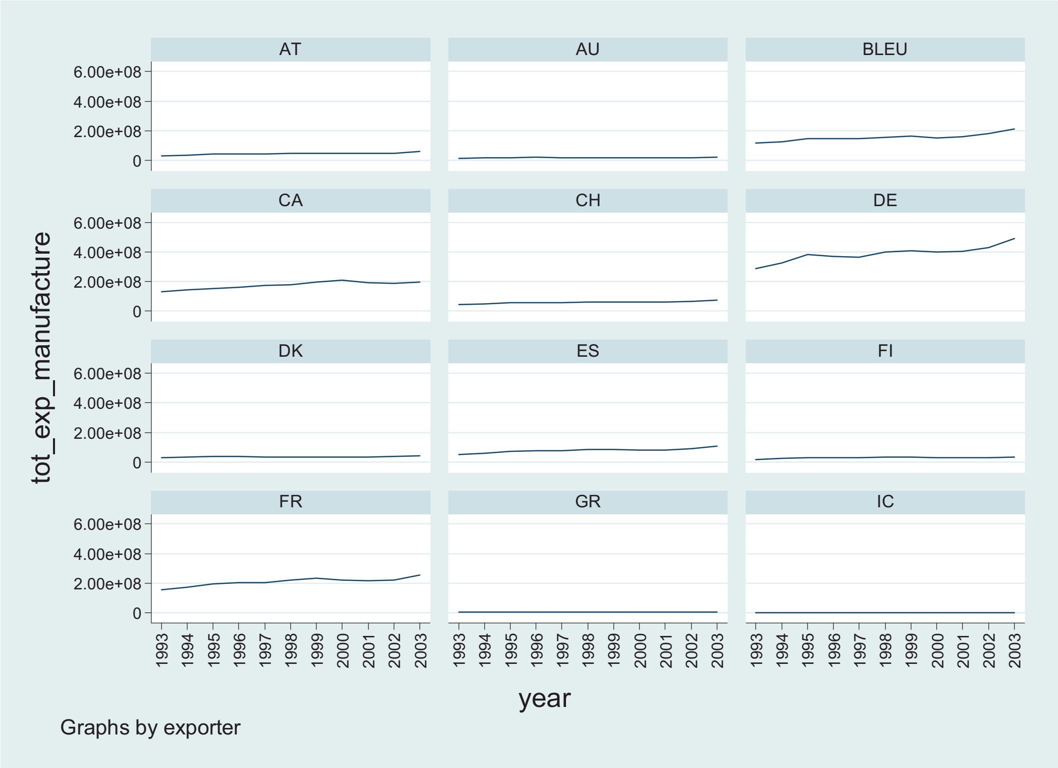
Figure 1 shows that even though there has been a slowdown in 1995 and a temporary decrease after 1999, the amount of export among the OECD24 countries has increased during the period considered. The graphs of Total Export by Country show that the USA is the major exporter of manufacture goods among the OECD24 countries; Germany, Japan, France and Italy follow.

Figure 1 - Total export among the OECD24 countries.



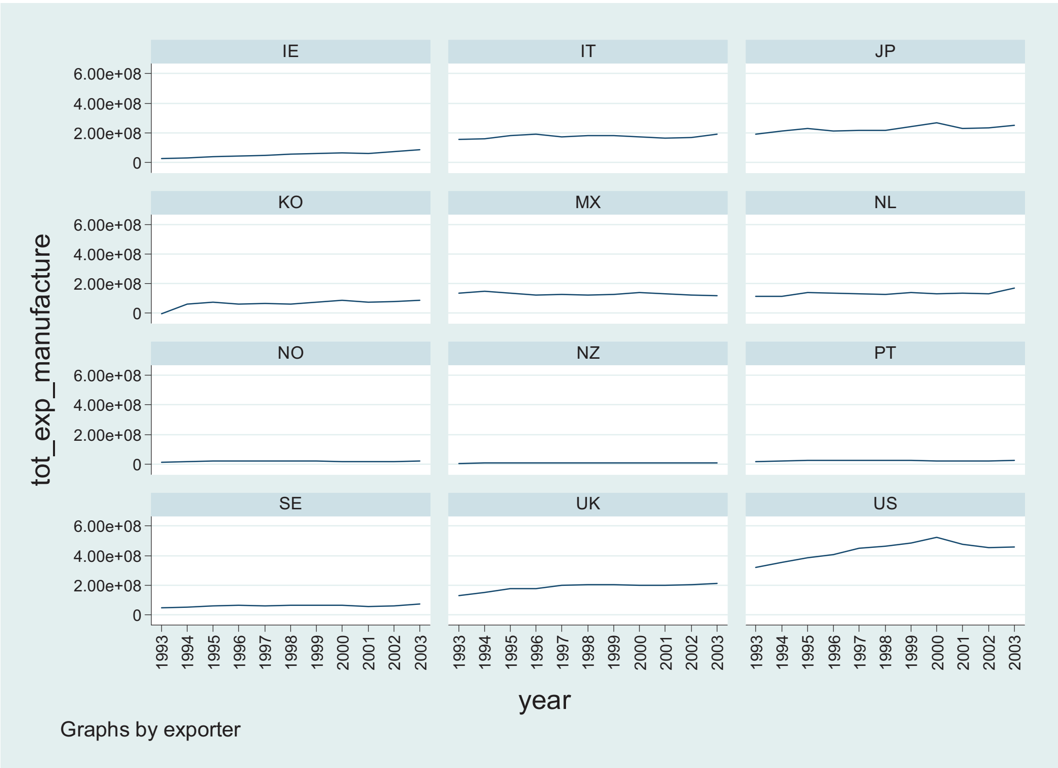
Zealand in the CER (Closer Economic Relationship); Mexico and South Korea in the PTN (Protocol to Trade Negotiations) and GSTP (General System of Trade Preferences). The use of pair-specific dummies automatically accounts for any RTA between the two countries in the pair when the relation holds for the entire time range. On the contrary, controls for RTAs have to be included in the regression if there is a regime switching during the time range or if country-specific dummies are used. Even though some RTAs are time-varying within the 1999-2003 time range, the variation is minimum because many of those came into force in 1994.

Figure 2 - Total Export by exporter.



Exporter: Austria (AT), Australia (AU), Belgium-Luxembourg (BLEU), Canada (CA), Switzerland (CH), Germany (DE), Denmark (DK), Spain (ES), Finland (FI), France (FR), Greece (GR), Iceland (IC).

Figure 3 - Total Export by exporter.



Exporter: Ireland (IE), Italy (IT), Japan (JP), South Korea (KO), Mexico (MX), the Netherlands (NL), Norway (NO), New Zealand (NZ), Portugal (PT), Sweden (SE), the United Kingdom (UK), the United States (US).

We consider both a Direct and an Indirect effect of the Euro. The former (Rose Effect) is caught by a dummy called $EZ2_t$ which is equal to one if both countries have the Euro as their national currency, the latter (Trade Diversion) is caught by the $EZ1_t$ dummy which is equal to 1 when just one country in the pair uses the Euro. The second dummy serves to check if the Euro has caused Trade Diversion. Indeed, a regional agreement (in this case a currency regional agreement) can cause a diversion in trade from countries out the agreement towards countries into (see Markusen et al. 1995 for a theoretical discussion, or European Commission 1997 for a more empirical one). Throughout this paper we refer to the Rose Effect (RE) as that one caught by the $EZ2_t$ dummy, while to Trade Diversion (TD) as the effect caught by the $EZ1_t$ dummy. The estimated equation is:

$$ex_{ij,t} = \alpha_0 + \alpha_{ij} + \alpha_t + \beta_1 y_{i,t} + \beta_2 y_{j,t} + \beta_3 comp_{i,t} + \beta_4 comp_{j,t} + \beta_5 EZ2_{ij,t} + \beta_6 EZ1_{ij,t} + \varepsilon_{ij,t}, \quad (1)$$

where the dependent variable is export from country *i* to country *j* at time t, while the regressors are respectively: $y_{i,t}$, the exporter's GDP at time *t*; $y_{j,t}$, the partner's GDP at time *t*; $comp_{i,t}$, the exporter's competitiveness indicator at time *t*; $comp_{jt}$, the partner's competitiveness indicator at time *t*; $EZ2_{ij,t}$, the Euro Effect control at time *t*; $EZ1_{ij,t}$, the Trade Diversion control at time *t*; α_{ij} , the pair-specific dummies which account for the Multilateral Resistance Terms; α_t and α_0 , the time dummy and the constant term. Since we have included pair–specific dummies to account for the Multilateral Resistance Terms (see Anderson and van Wincoop 2003) and a year dummy as well, the model is a two-way fixed effects model (Baltagi 2005). The estimation output is in Table 1.[7]

The GDP's marginal effect is lower than the competitiveness indicator's and the competitiveness indicator is significant both for the exporter and the partner, all the covariates are correctly signed. The Rose Effect (RE) is about 10.2% ($[e^{0.098}-1]*100$ = 10.2%) if estimated alone, and about 17.7% ($[e^{0.163}-1]*100$ = 17.7%) if estimated simultaneously with the Trade Diversion (TD); the TD is about 10.9% ($[e^{0.104}-1]*100$ = 10.9%). Both the RE and the TD are statistically significant. Hence, we may deduce a 17.7% increase of trade among the Euro Zone (EZ) countries during 1999-2003 with respect to 1993-1998, and a 10.9% increase of trade between the EZ countries and the rest of the OECD24 countries during 1999-2003 with respect to 1993-1998. We point

[7] It is to notice that in eq. (1) there are no time-invariant covariates. This is so because their estimation is not feasible given the perfect multicollinearity with the group dummies or their elimination if within-group transformations are used.

out that these values of the RE are realistic and literature-consistent (see Flam and Nostrom 2003, or Micco et al. 2003).

Table 1 – Estimation of the RE and TD.

	(1) WGE	s.e.	(2) WGE	s.e.
LOG_GDP_EXP	0.548	(0.079)**	0.541	(0.078)**
LOG_GDP_PAR	0.259	(0.051)**	0.253	(0.051)**
LOG_REER_EXP	-0.640	(0.101)**	-0.608	(0.101)**
LOG_REER_PAR	0.728	(0.072)**	0.756	(0.072)**
EZ2_t	0.098	(0.011)**	0.163	(0.015)**
EZ1_t			0.104	(0.015)**
Observations	6069		6069	
Groups	552		552	
Ov. R-squared	0.492		0.492	

- Countries included: OECD24. - Pair and Year dummies included
- Robust standard errors in parentheses - * significant at 5%; ** significant at 1%

In the previous estimate we have set 1999 as the start year for the Euro's effect, an alternative strategy is to let the data say when it started. This is done through a yearly $EZ2_t$ dummy (EZ2_199, EZ_1994, …, EZ_2003). Estimates of the yearly RE and TD are reported in Table 2 and displayed in Figure 4. The RE arises in 1999 and strengthens year by year as one would expect. Hence, the Euro's Effect has a credible timing which encourages to consider this estimation as a clear proof of a positive effect of the Euro. As regards the TD, it is significant since 2000 and its pattern is believable and consistent with the evolution of export plotted in Figure 1.

On the whole, the evidence in this section is consistent with the findings in the literature, namely on the basis of the Rose Effect the Euro has been having a significant and positive effect upon European trade since its introduction in 1999. In the next section we check this finding through the analysis of the Border Effect evolution.

Figure 4 - Yearly RE.

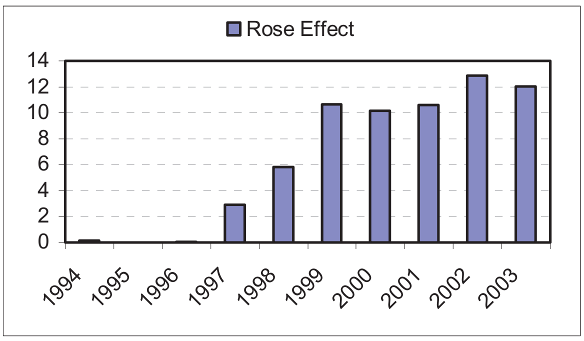


Table 2 – Yearly RE and TD.

	(1) WGE	s.e.	(2) WGE	s.e.
LOG_GDP_EXP	0.543	(0.079)**	0.537	(0.079)**
LOG_GDP_PAR	0.254	(0.052)**	0.248	(0.052)**
LOG_REER_EXP	-0.635	(0.101)**	-0.603	(0.101)**
LOG_REER_PAR	0.733	(0.073)**	0.765	(0.072)**
EZ2_1994	0.001	(0.033)	0.010	(0.043)
EZ2_1995	-0.005	(0.031)	-0.007	(0.043)
EZ2_1996	0.000	(0.030)	-0.014	(0.042)
EZ2_1997	0.028	(0.030)	0.041	(0.041)
EZ2_1998	0.058	(0.030)	0.068	(0.044)
EZ2_1999	0.106	(0.029)**	0.136	(0.039)**
EZ2_2000	0.101	(0.032)**	0.168	(0.042)**
EZ2_2001	0.105	(0.031)**	0.200	(0.042)**
EZ2_2002	0.128	(0.035)**	0.224	(0.045)**
EZ2_2003	0.120	(0.035)**	0.216	(0.046)**
EZ1_1994			0.014	(0.043)
EZ1_1995			-0.001	(0.045)
EZ1_1996			-0.024	(0.043)
EZ1_1997			0.017	(0.042)
EZ1_1998			0.015	(0.045)
EZ1_1999			0.045	(0.040)
EZ1_2000			0.103	(0.042)*
EZ1_2001			0.149	(0.043)**
EZ1_2002			0.151	(0.044)**
EZ1_2003			0.153	(0.045)**
Observations	6069		6069	
Number of pair	552		552	
Ov. R-squared	0.493		0.492	

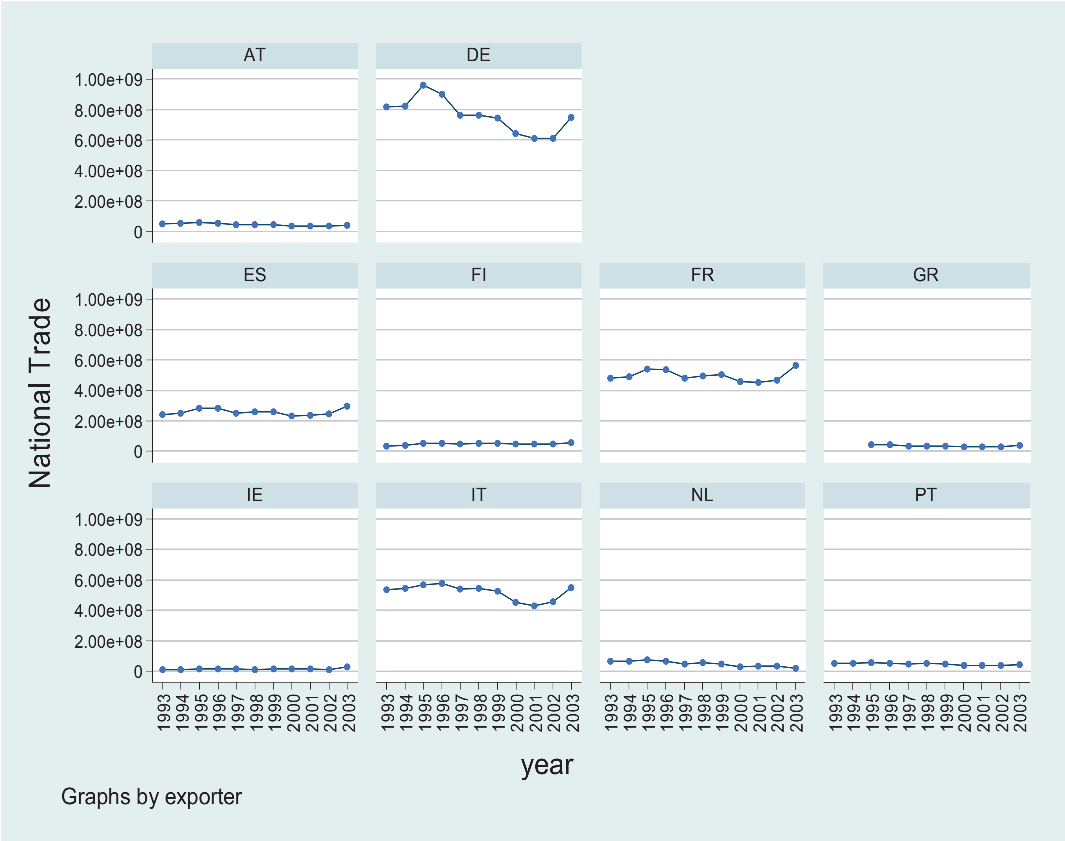
- Countries included: OECD24 - Pair and Year dummies Included.
- Robust standard errors in parentheses - * significant at 5%; ** significant at 1%

IV - Border Effect Estimations

In our analysis it is fundamental to use the same dataset to estimate both the Rose Effect (RE) and the Border Effect (BE), otherwise the coherence between the Euro's effect detected through these two indicators could be diverted somehow. Nonetheless, we need National Trade (NT) observations to estimate the BE. Country *i* ' s NT is equal to its total production of a certain good less its total export of that good. It is therefore the amount which is traded-consumed within the national borders of the country where its production has taken place; this definition of NT is due to Wei (1996). Then, the dataset we use to estimate the BE includes NT observations, this is the only difference with respect to the one used for the RE. In Figure 5 we display the amount of NT for each EZ country.

When we consider only two countries, country *i* 's BE indicates how much country *i*'s NT is higher than its export to country *j*. If we consider more than two countries, country *i*'s BE indicates how much its NT is higher than its average export towards the other countries; namely, towards its *representative* trade partner.[8] In the following subsections we first explain what characterizes the estimation of the Border Effect, in subsection IV.b we discuss our estimations.

Figure 5 – National Trade by EZ country.



IV.a - Econometrics of the Border Effect

The estimation of the Border Effect has some specific characteristics, here we explain two of them. The BE is estimated through a gravity equation in which dummies catch the National Trade (NT) observations; the NT observations are outliers with respect to the series of bilateral trade flows.[9] If one estimates country *i* 's BE only with respect to country *i* ' s main trade partner (country *j*), she gets the lowest possible BE value for country *i*. While, if she estimates country *i* ' s BE with respect to its two main trade partners (country *j* and country *z*), she gets a higher BE value. Indeed, the less important country *i* 's trade partners are included (country *z* less important that *j*), the higher the BE value is. This happens because the ratio of NT

[8] Since we consider bilateral flows measured in monetary values, the BE indicates how much the value of NT is higher than the value of export towards the representative partner.

[9] Whenever we include in the gravity equation dummies which account for block effects (such as RTAs dummy or Contiguity dummy) we set them equal to zero for the NT observations. This is common practice in gravity estimations (see Nitsch 2000 and Helliwell 1997).

over export towards the representative partner increases. Hence, any EU country's BE with respect to the group of its EU partners is lower than its BE with respect to a larger sample of trade partners (the EU sample is nested in the larger) because the EU as a whole is the main trade partner of any European country (see Eurostat 2003). For this reason, when we want to estimate the BE, we have to include only the countries in which we are interested; we therefore included only the EU countries in the BE estimations.

Another important characteristic of the BE estimation is that we cannot use pair-specific dummies to proxy the Multilateral Resistance Terms. Indeed, if we do so, the NT observations will be caught by two different dummies and the BE cannot be estimated. Hence, after many attempts, we eventually concluded that for the objective of this paper (and probably whenever researchers are interested in the evolution of the BE) pair-specific dummies are not suitable to model the Multilateral Resistance Terms.[10]

IV.b - Estimations of the Border Effect

We now discuss the estimation of the gravity equation which provides us with an estimate of the Border Effect (BE). In this gravity equation the Multilateral Resistance Terms are modelled through Country-Specific Time-Invariant dummies. To infer the effect of the Euro through the evolution of the BE, we study the BE on a sample of data which includes only the European countries for the reason abovementioned. Then, the average BE ex-ante and ex-post 1999, both for the Euro Zone (EZ) countries and for Denmark, Sweden and the UK (EU non EZ countries), is estimated through the following equation:

$$\begin{aligned} ex_{ij,t} = {} & \alpha_0 + \alpha_i + \alpha_j + \alpha_t + \beta_1 y_{i,t} + \beta_2 y_{j,t} + \beta_3 comp_{i,t} + \beta_4 comp_{j,t} + \beta_5 dist_{ij} + \beta_6 contig_{ij} \\ & + \beta_7 lang_{ij} + \beta_8 \text{NT\_EZ\_ante99} + \beta_9 \text{NT\_EZ\_post99} + \beta_{10} \text{NT\_EUnoEZ\_ante99} + \beta_{11} \text{NT\_EUnoEZ\_post99} + \varepsilon_{ij,t} \end{aligned} \quad (2)$$

The dependent variable $ex_{ij,t}$ is export from country i to country j at time t. The covariates are respectively: $y_{i,t}$, the exporter's GDP at time t; $y_{j,t}$, the partner's GDP at time t; $comp_{i,t}$, exporter's competitiveness indicator at time t; $comp_{j,t}$, partner's competitiveness indicator at time t; $dist_{ij}$, distance between country i and j; $contig_{ij}$,

[10] For the same reason the estimation of the BE in a dynamic panel framework seems not to be econometrically feasible. In fact, to perform dynamic panel regressions it is necessary to define the group indicator on which to calculate the first difference of the equation (GMM procedures) or to include group dummies (Least Squares Dummy Variables Corrected). But, as explained, we cannot estimate the BE whenever a pair-dummy is included in the regression or the pair is taken as group indicator.

Contiguity control; $lang_{ij}$, Common Language control; NT_{ij}_..., National Trade dummies.[11]

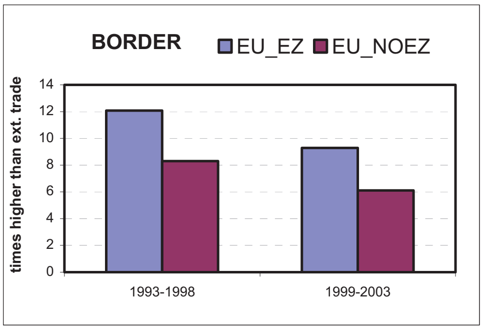
The regression output is in Table 3, while the histogram of the estimated BEs is in Figure 6. The BE reduction between 1993-1998 and 1999-2003 is about 23.6% ($[e^{2.22} - e^{2.49}]/e^{2.49}$) for the EZ countries and 26.6% for the EU-non-EZ countries ($[e^{1.80} - e^{2.11}]/e^{2.11}$). Hence, the percentage decrease is larger for the EU-non-EZ countries, while we expected a stronger BE decrease for the EZ countries in case of a positive and relevant effect of the Euro. This first evidence against a significant Euro's effect is strengthened by the results of the Wald test of linear restrictions. The linear restriction "EZ countries' BE equal before and after 1999" is not rejected (p-value 0.104), while the hypothesis "EU-non-EZ countries' BE equal before and after 1999" is rejected (p-value 0.002); both the hypothesis "EZ countries' BE equal to EU-non-EZ countries' before 1999" and "EZ countries' BE equal to EU-non-EZ countries' after 1999" are rejected (p-value 0.001 and 0.004).

Table 3 - BE by group and time interval.

	LSDV	s.e.
LOG_GDP_EXP	0.914	(0.173)**
LOG_GDP_PAR	0.661	(0.158)**
LOG_REER_EXP	-0.864	(0.229)**
LOG_REER_PAR	-0.009	(0.210)
LOG_DISTANCE	-0.776	(0.030)**
CONTIGUITY	0.367	(0.036)**
LANGUAGE	0.329	(0.061)**
B_eu_ez_post99	2.229	(0.144)**
b_eu_ez_ante99	2.491	(0.116)**
b_eu_noez_post99	1.809	(0.092)**
b_eu_noez_ante99	2.116	(0.085)**
Observations	2143	
Deg.ofFree.	2095	
R-squared	0.95	

- Countries included: EU14
- Exporter and Partner plus Year dummies included
- Robust standard errors in parentheses
- * Significant at 5%, ** significant at 1%

Figure 6 - BE by group and time interval.



To have a clearer view of the BE evolution, we need to estimate the yearly BE. We rule out the 1993 NT observations for all the countries to avoid perfect multicollinearity with the year dummies, and we estimate the yearly BE on a sample which includes bilateral trade only among the EZ countries. As abovementioned, to

[11] It is to notice that to estimate eq.(2), measures of inter and intra-national distance are necessary. The BE estimates are known to be dependent upon the specific measure of distance used; for a detailed discussion see Nitsch (2000), Head and Mayer (2002). We took measures of distance from the CEPII Distances dataset (www.cepii.fr). We checked the sensitiveness of our results to different measures, as expected the choice affects a lot the magnitude, but not the evolution. Roughly speaking, the change from a measure to another shifts upwards or downwards a plot which in all the cases looks alike.

assign a positive effect to the Euro in terms of trade integration, we look at the trend of the BE. If the Euro had a positive effect, we expect to observe a BE evolution as that one drawn in Figure 7. The estimate of the yearly BE is displayed in Table 4 and plotted in Figure 8.

The BE keeps decreasing all along the time period considered. We expected this as outcome of the deepening integration among the European countries. However, the evolution of the BE year after year does not support the hypothesis of a positive effect of the Euro. Indeed, if we plot the trend of the BE in two different periods (1993-98 and 1999-03), the trend line is less downward sloped in the second period (check the slope coefficients in Figure 8). This means that border-linked trade costs have decreased more quickly before the Euro introduction than after.

Table 4 - Yearly BE, EZ countries.

	LSDV	s.e.
LOG_GDP_EXP	1.015	(0.222)**
LOG_GDP_PAR	0.661	(0.152)**
LOG_REER_EXP	-1.417	(0.484)**
LOG_REER_PAR	0.935	(0.479)
LOG_DISTANCE	-0.704	(0.043)**
CONTIGUITY	0.320	(0.041)**
LANGUAGE	0.181	(0.072)*
border1994	2.631	(0.241)**
border1995	2.728	(0.270)**
border1996	2.673	(0.279)**
border1997	2.553	(0.286)**
border1998	2.446	(0.281)**
border1999	2.408	(0.286)**
border2000	2.336	(0.309)**
border2001	2.350	(0.303)**
border2002	2.289	(0.308)**
border2003	2.273	(0.340)**
Observations	1309	
Deg.ofFree.	1255	
R-squared	0.95	

- Countries included: EZ11. - Robust st.d errors in par.ses.
-* significant at 5%; ** significant at 1%.
- Exp.r and Par.r time-invariant plus Year dummies included.

Figure 7 - Supposed BE dynamics.

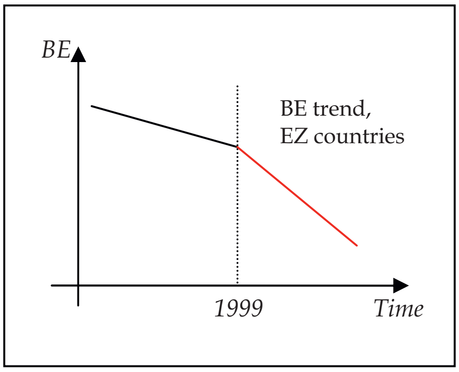
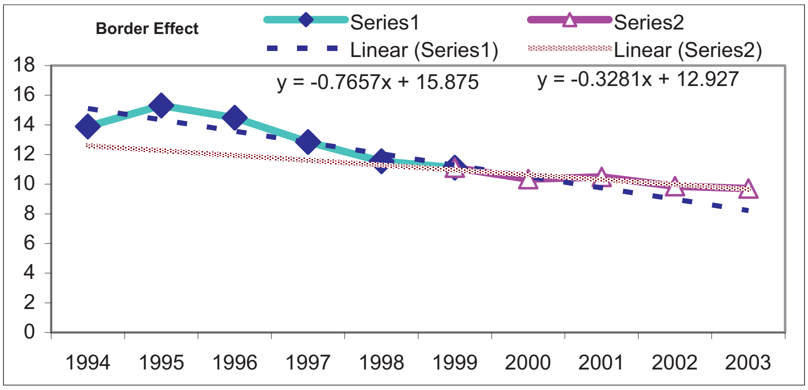


Figure 8 - Yearly BE, EZ countries.



It is worthwhile to compare the yearly BE of the EZ countries with the yearly BE of the European Union countries which do not use the Euro (EU-non-EZ countries). As previously explained, the EU-non-EZ countries could have been suffering from Trade Diversion since the Euro's introduction. If this was the case, the BE of the EU-non-EZ countries should have increased. To test this hypothesis we estimated the yearly BE for the two different groups of countries; the estimation output is in Table 5 and plotted in Figure 9.

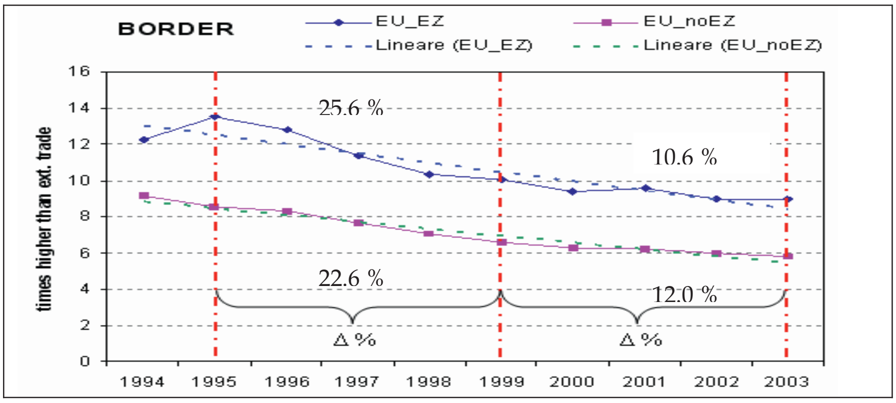
Denmark, Sweden and UK's integration in the European market seems not to have been affected by the Euro's introduction, it keeps on deepening all along the period considered. We calculated the BE change after and before 1999 for the two groups of countries, the EZ countries' BE has decreased by 25.6% in the period 1995-99 and by 10.6% in the period 1999-2003, while EU-non-EZ countries' BE decreased by 22.6% in 1995-99 and by 12.0% in 1999-03. Both groups of countries seem not to have been affected by the Euro's introduction in 1999, and their BE has decreased more in the pre-Euro period than after its introduction (see Figure 9).

Table 5 - Yearly BE, EU14 countries.

	LSDV	s.e.		LSDV	s.e.
LOG_GDP_EXP	0.929	(0.175)**	LOG_REER_EXP	-0.874	(0.230)**
LOG_GDP_PAR	0.677	(0.160)**	LOG_REER_PAR	-0.018	(0.211)
LOG_DISTANCE	-0.772	(0.030)**	LANGUAGE	0.329	(0.061)**
CONTIGUITY	0.370	(0.036)**			
b_eu_ez94	2.506	(0.220)**	b_eu_noez94	2.213	(0.126)**
b_eu_ez95	2.604	(0.246)**	b_eu_noez95	2.141	(0.142)**
b_eu_ez96	2.548	(0.260)**	b_eu_noez96	2.113	(0.143)**
b_eu_ez97	2.426	(0.271)**	b_eu_noez97	2.030	(0.138)**
b_eu_ez98	2.337	(0.267)**	b_eu_noez98	1.949	(0.125)**
b_eu_ez99	2.307	(0.271)**	b_eu_noez99	1.884	(0.139)**
b_eu_ez00	2.240	(0.296)**	b_eu_noez00	1.834	(0.155)**
b_eu_ez01	2.259	(0.288)**	b_eu_noez01	1.825	(0.160)**
b_eu_ez02	2.192	(0.299)**	b_eu_noez02	1.785	(0.159)**
b_eu_ez03	2.194	(0.326)**	b_eu_noez03	1.756	(0.169)**
Observations	2131				
Deg.ofFree.	2067				
R-squared	0.95				

- Countries included: EU14
- Exporter and Partner time-invariant dummies plus Year dummies included
- Robust standard errors in parentheses
- * significant at 5%; ** significant at 1%

Figure 9 – Yearly BE for the EZ and EU-noEZ group.



The estimations discussed in this section show a BE unaffected by the Euro. Given the pictured link between the RE and the BE, we did not expect this finding. Indeed, a positive RE should involve a decreasing BE if the RE is caused by a trade-costs reduction. Then, we can deduce that either the Euro did not cause a significant reduction of trade costs or that something else, for which we could not control in our estimations, offset the trade costs reduction. We discuss this in the conclusions where we provide an explanation of the clash between the Rose and the Border effect.

V - Robustness Checks

The results discussed in this paper have gone through many robustness checks. These are not included in the text, but we shortly mention them in this section. The checks consisted in using different specifications of the gravity equation, samples of countries and data to check the robustness of the Rose and the Border Effect (BE) estimates. On the whole, the checks proved our findings robust. The magnitude of the estimates change by applying a specification instead of another, but the indicators' evolution remains alike. Consequently, our conclusions are unaffected.[12]

In the first group of checks we estimated the Rose Effect (RE) in a sample which includes only the European countries before the May 2004 enlargement. These countries have been for a long time subject to a similar institutional treatment and the bulk of them joined the Euro in 1999. Then, we thought it convenient to check our results for this more-selected group in order to rule out any suspect of bias caused by the inclusion of other countries. Secondly, we checked both the RE and the BE estimates by using different specifications of the Multilateral Resistance Terms (see Cheng and Wall 2005). With respect to the estimations discussed in the text, we used country-specific time-invariant dummies to estimate the RE and country-specific

[12] All the robustness checks are available upon request.

time-varying dummies (as suggested by Baldwin 2006) to estimate the BE. We thought this to be an important check. Indeed, there is not consensus about the right specification of the gravity equation and on how to account for factors that cannot be caught by any specific control.

Thirdly, we used a large sectoral dataset to check both the RE and the BE sector by sector. This was done to get sure that the results discussed in the paper are not dependent upon aggregation and not limited to aggregate analyses. For the majority of the sectors considered, the sectoral RE and BE have a similar pattern to the RE and BE estimated using aggregate data. Hence, sectoral and aggregate estimations are unanimous about the pattern of the RE and BE in the period considered. Forth, we used different measures of distance to find out how much the BE estimates are sensitive to any specific one. As mentioned in footnote 11, the change from one measure to another shifts upwards or downwards a plot which in all the cases looks alike.

We have dedicated the two appendices of this paper to the discussion of some other robustness checks which we deemed of greater interest. Appendix I contains dynamic estimations of the RE, while in Appendix II we discuss an intuitive check of the yearly BE estimate in Table 4. We refer the reader to those appendices for a detailed discussion.

VI - Conclusions

Our main finding is that, on the basis of the Rose Effect, the Euro has influenced European trade after 1999, while there is no effect of the Euro according to the evolution of the Border Effect. The most-quoted theoretical explanation of the Rose Effect is founded on the removal of border-linked costs (this is one of the main argument of the Optimum Currency Area literature, see Baldwin and Wyplosz 2004), but we deem the Border Effect to be the true indicator of border-linked costs. Then, given the evolution of the estimated Border Effect, we are prone to believe that there was no acceleration in the reduction of border-linked costs (with respect to the historical tendency) after the Euro's introduction. Consequently, we can either believe that the Rose Effect does not catch the effect of the Euro (it is a spurious outcome) or that the effect of the Euro caught by the Rose Effect is due to something else than a reduction in border-linked trade costs.

Let us begin considering the first option. Berger and Nitsch (2005) and Mongelli et al. (2005) show that economic integration was rising steeply just before Euro's introduction. Their thesis is that pro-trade adjustments to pre-Euro integration take time, then it could be that the lagged effects of Single Market measures show up in the post-1999 data and get blurred with the trade effects of the Euro. This would not be a problem for Rose Effect estimations if only all EU members introduced Single Market measures at the same time, but EU members differ widely on their pace of implementing EU directives.[13] Even though we have used a dynamic specification of the gravity equation and controlled for common shocks -through the year dummies- and country-specific shocks, one can not be completely sure that this catches everything but the effect of the Euro. Indeed, it may be quite difficult to effectively disentangle the effects of European integration from the Euro's because European integration is a work in progress as the Border Effect evolution shows (as well as Berger and Nitsch 2005, and Mongelli et al. 2005). Then, there is a chance that the Euro's effect detected is nothing more than the delayed and differential effect of pro-trade directives as the BE evolution suggests.

On the contrary, if we believe that the Rose Effect correctly catches the increase in exchanges due to the Euro's introduction, the question to answer is what else than a trade costs reduction could have caused such increase. Answers founded on output variations or changes in the cross-countries competitive position are not believable since we controlled for those. In our estimations, we found that trade between the EZ countries and other nations rose with the Euro's introduction, but not as much as trade among the EZ countries (see Table 2 and 3). This result is intriguing and tends to reject the cost-reduction explanation of the Rose Effect because in that case there should have been trade diversion. The Euro would have been akin to a discriminatory liberalisation and this should have reduced the exports of non-Euro nations to the Euro Zone as Baldwin (2006) suggests. According to Baldwin (2006), an explanation of the positive Rose Effect not founded on costs reduction and not conflicting with trade creation can be found in Baldwin and Taglioni (2004). Their basic intuition is simple. Most European firms were not engaged in trade before the Euro's introduction, they sold only in their local markets due to a variety of reasons, one of which is aversion to exchange rate uncertainty. Such uncertainty is easily faced by large companies but to small and medium firms it is as a real barrier. Since monetary unions eliminate this uncertainty, the number of EZ firms engaged in export to other EZ markets increased after the Euro's introduction. Consequently, we

[13] The Internal Market Scoreboard gives an account of the difference in implementing EU directives among the EU countries. More information and the Scoreboard itself can be found at "http://ec.europa.eu/internal_market/score/index_en.htm"

observe a trade-flows increase after 1999. Furthermore, more transparency and cost/price comparability as well as higher ease to access other national markets after the Euro's introduction could have worked in the same manner as the removal of exchange rate uncertainty just discussed.

On the whole, we are prone to believe that the Rose Effect is not a spurious outcome because it emerges in many papers which use different econometric techniques, datasets and time intervals, and because it is always clearly detected in 1999. It is unlikely that all of us made the same mistake and that it emerges always around 1999 by chance. Consequently, we deem more likely the second case just discussed. Namely, the Rose Effect catches an increase in exchanges linked to the Euro's introduction, but that increase can not be explained through an acceleration in the reduction of border-linked trade costs among the countries which adopted the Euro.

Appendix I - Dynamic Estimations of the Rose Effects.

Trade flows are observed to be persistent over time. There are different explanations for this, one is *habit formation* in consumption and another is based on internationalization and sunk costs linked to export in a foreign market (Evans 2003, Eichengreen and Irwin 1997). Then, we could need to include lagged values of the dependent variable to write the model in a correct econometric way (Bun and Klassen 2002). In this section we discuss dynamic estimations of the RE obtained through two different estimators, these serve as robustness checks of the static ones in section III. The first estimator used is a General Method of Moments (GMM) estimator, the second is the Least Squares Dummy Variables Corrected estimator; we start with the GMM estimator.[14]

In the group of GMM estimators available, we chose the Arellano-Bond estimator because of the low coefficient of the lagged dependent variable.[15] The model which we estimate is:

$$ex_{ij,t} = \alpha_0 + \alpha_{ij} + \alpha_t + \beta_0 ex_{ij,t-1} + \beta_1 y_{i,t} + \beta_2 y_{j,t} + \beta_3 comp_{i,t} + \beta_4 comp_{j,t} + \varepsilon_{ij,t} \quad . \qquad (3)$$

In any Arellano-Bond estimation there is need to define the Instruments Matrix. This matrix includes the lagged levels of the dependent variable by default, the researcher can also include as additional instruments the regressors in the equation to estimate. However, these can be included only if they are strictly exogenous or predetermined (a regressor x is *Predetermined* when $E[x_{i,t}v_{i,s}]=0$ for $t < s$, t and s = $1 \ldots T$; *Strictly Exogenous* when $E[x_{i,t}v_{i,s}]=0$ for $\forall\, t$ and $\forall\, s$). The definition of each regressor as strictly exogenous, predetermined or excluded instrument should be based on the Difference Sargan Test (DST) and the economic nature of the variable itself.[16] Unfortunately, the DST is not always reliable (rejects too often) when the data are affected by heteroskedasticity as argued by Arellano and Bond (1991).[17] For this

[14] Given the longitudinal and time dimension of our panel, GMM estimators are probably the most suitable, see Judson and Owen (1999) for a discussion about the choice of the dynamic estimator.

[15] An alternative is the Blundel-Bond estimator more suited in case of a high coefficient of the lagged Dependent Variable (see Bond 2002).

[16] The DST checks the validity of the additional over-identifying restrictions due to shifting from a model with less to a model with more restrictions (namely, more instruments included).

[17] According to the DST, we should have included in the Instruments Matrix either the Competitiveness indicators or the GDPs variables as predetermined, or both as strictly exogenous. Only when we impose a very small number of restrictions (around 60) the Sargan (1958) test (and not the Difference Sargan Test) does not reject the hypothesis of valid instruments (see Bond 2002 for an example of instruments selection by means of the DST). Though, the estimated coefficients are not gravity-consistent if those settings are used.

reason, we did not follow the conclusions of the DST, instead we defined the regressors according to a reasonable economic rationale. Consequently we thought to have two possible settings: a) past values of the competitiveness indicators and of the GDPs do not influence current bilateral exports, in this case both variables are predetermined and their past values can be included in the Instruments Matrix; b) only past values of the competitiveness indicators do not influence current bilateral exports, while GDP is endogenous and it has to be left out.

We run estimations for both settings, in either the Arellano-Bond test for second order autocorrelation rejects the null, so we did not doubt the consistency of the Arellano-Bond estimator. We performed both ONE-Step and TWO-Step GMM estimations (see Arellano 2003, appendix A.7). We point out that the TWO-Step standard errors are computed in accordance to the Windmeijer finite-sample correction (otherwise they are seriously biased, see Windmeijer 2005).[18]

The estimate of the yearly RE is in Table 6. We choose regression 2 as our preferred for the following reasons. First, the estimation of regression 1 and 2 is more efficient than that of regression 3 and 4 because the GMM estimator increases its efficiency when more restrictions are imposed (in regression 1 and 2 the Instruments-Matrix includes two more instruments). Second, since the TWO-Step robust estimation is more efficient than the ONE-Step robust, we prefer regression 2 to regression 1. We acknowledge that defining the GDP as predetermined is questionable, however the results are not driven by this choice. Indeed, the coefficients are slightly different across estimates.

On the basis of the estimation output, the RE arises in 1998 and not in 1999 as it was in the static regressions in section III (compare Table 6 with Table 2). From a cross-regressions comparison, it turns out that the RE is present and strongly significant in 1999, it weakens in 2000 but it strengthens again in 2001.

[18] The Two-Step procedure allows computing the optimal Weights-Matrix of the GMM estimator (see Hansen 1982) on the basis of the consistent estimate run at the first step. Nonetheless, this estimate at the second step is not efficient in case of heteroskedasticity. Then, it is common practice to apply the Windmeijer (2005) correction.

Table 6 - Yearly RE, Dynamic Panel Estimation, Arellano-Bond estimator.

	RCPs and GDPs predetermined		RCPs predetermined, GDPs out	
	ONE-Step (1)	TWO-Step (2)	ONE-Step (3)	TWO-Step (4)
1st lag LOG_EXPij	0.391	0.382	0.316	0.321
	(0.039)**	(0.041)**	(0.050)**	(0.052)**
LOG_GDP_EXP	0.553	0.496	0.762	0.676
	(0.097)**	(0.095)**	(0.165)**	(0.174)**
LOG_GDP_PAR	0.109	0.106	-0.012	0.091
	(0.078)	(0.081)	(0.114)	(0.130)
LOG_REER_EXP	-0.528	-0.445	-0.677	-0.540
	(0.131)**	(0.127)**	(0.169)**	(0.183)**
LOG_REER_PAR	0.724	0.741	0.823	0.716
	(0.103)**	(0.113)**	(0.120)**	(0.154)**
EZ_1995	0.000	0.016	-0.005	0.004
	(0.020)	(0.022)	(0.021)	(0.019)
EZ_1996	-0.012	0.004	-0.014	-0.007
	(0.019)	(0.020)	(0.021)	(0.020)
EZ_1997	0.014	0.027	0.014	0.018
	(0.020)	(0.020)	(0.022)	(0.020)
EZ_1998	0.048	0.064	0.044	0.044
	(0.023)*	(0.023)**	(0.027)	(0.024)
EZ_1999	0.065	0.073	0.068	0.064
	(0.021)**	(0.023)**	(0.024)**	(0.024)**
EZ_2000	0.047	0.054	0.055	0.051
	(0.025)	(0.026)*	(0.028)	(0.028)
EZ_2001	0.072	0.083	0.075	0.067
	(0.025)**	(0.026)**	(0.028)**	(0.028)*
EZ_2002	0.084	0.095	0.089	0.066
	(0.029)**	(0.028)**	(0.033)**	(0.031)*
EZ_2003	0.070	0.097	0.074	0.070
	(0.027)*	(0.029)**	(0.032)*	(0.031)*
Observations	4965	4965	4965	4965
N° of pair	552	552	552	552
Number of instruments	273	273	165	165

- Countries included: OECD24
- Pair and Year dummies included
- Robust standard errors in parentheses
- *significant at 5%; ** significant at 1%

The second estimator we discuss in this section belongs to the group of Bias-Corrected Estimators. These correct the Bias of the Fixed-Effects panel estimator. We implement the Estimator defined by Bruno (2005) built on the basis of Kiviet's corrections (Kiviet 1995) of the Least Squares Dummy Variables estimator (LSDV Corrected, LSDVC). The estimate of the yearly RE is in Table 7.

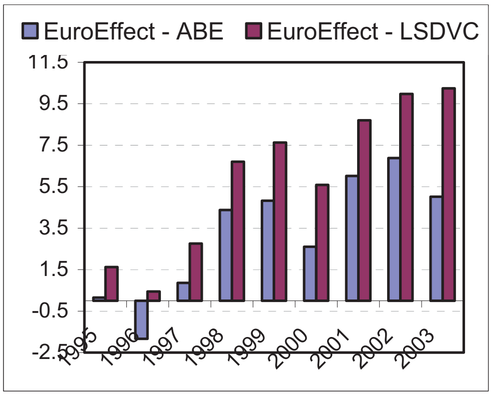
On the basis of the bootstrapped standard errors, it turns out to be not statistically significant. However, by plotting the yearly RE obtained through the Arellano-Bond estimator with that obtained through the LSDVC estimator in Figure 10, it appears that they have a similar evolution, although the level of the RE yielded by the LSDVC estimator is persistently lower.

Table 7 – Yearly RE, Dynamic Panel Estimation, LSDVC estimator.

	LSDVC	Boot.-s.e.
1st lag LOG_EXPij	0.543	(0.012)**
LOG_GDP_EXP	0.532	(0.031)**
LOG_GDP_PAR	0.116	(0.042)**
LOG_REER_EXP	-0.446	(0.054)**
LOG_REER_PAR	0.647	(0.068)**
EZ_1995	0.001	(0.031)
EZ_1996	-0.018	(0.037)
EZ_1997	0.008	(0.034)
EZ_1998	0.042	(0.033)
EZ_1999	0.047	(0.035)
EZ_2000	0.025	(0.029)
EZ_2001	0.058	(0.037)
EZ_2002	0.066	(0.033)*
EZ_2003	0.048	(0.036)
Observations	5517	
Number of pair	552	

- Dataset: OECD24
- Pair and Year dummies included
- Bootstrap Standard errors in parentheses, obtained through 50 repetitions
- * significant at 5%; ** significant at 1%

Figure 10 – Yearly RE, Arellano-Bond and LSDVC estimations.



On the whole, dynamic estimations do not confute the evidence discussed in section III. The finding of a positive RE with a believable timing is robust to the inclusion of dynamics in the model and to different dynamic estimators. The output of dynamic panel estimators is usually shaped by the definition of the Instruments-Matrix, but this is not our case. Indeed, it is very similar across different definitions of the Instruments Matrix as shown in Table 6. Furthermore, the evolution of the yearly RE on the basis of the two different dynamic estimators considered (Arellano-Bond and LSDVC) looks alike.

Appendix II - An alternative check of the Border Effect.

The Border Effect estimated in a gravity equation is exposed to the bias typical of an econometric regression. For this reason it is worthwhile to think to a possible robustness check of the Border Effect (BE). With this purpose we can compute a ratio whose evolution over time can be compared to the BE's. It is the ratio of National Trade (NT) over average export. In this section we compute this ratio for the group of EZ countries.

To compute the ratio we need to proceed in two steps. First, we define a Country-specific Border Ratio, and then an Aggregate Border Ratio; the former is nested in the latter. Country *i*'s Border Ratio is equal to its NT over its average export towards the group of its trade partners:

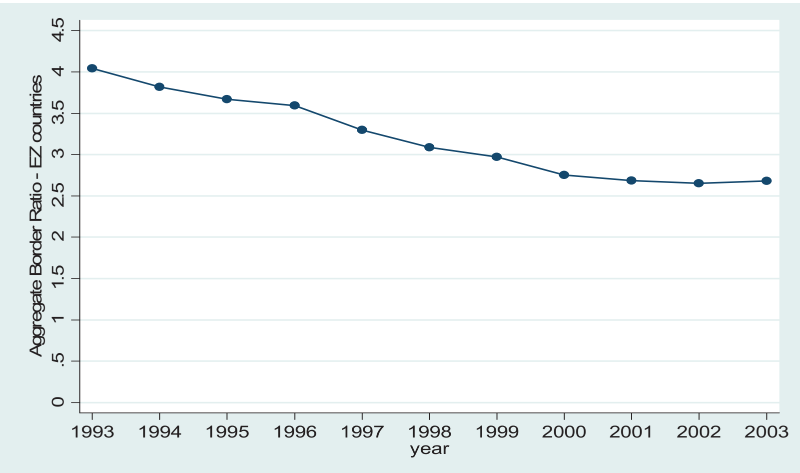
$$CBR_i^y = \frac{NT_i^y}{\left(\sum_{j=1}^{n} ex_{ij}^y\right)/n}, \tag{4}$$

where: i is the exporter ($i = 1,\ldots, n$), j is the partner ($j = 1,\ldots, n$), y is the year ($y = 1,\ldots, m$), NT_i^y is country i's NT during the y^{th} year and ex_{ij}^y is export from i to j during the y^{th} year. Since we consider only the Euro Zone (EZ) countries, n and m are equal to 11. To compute the same ratio but for the EZ countries altogether, the formula becomes:

$$ABR^y = \frac{\sum_{i=1}^{10} NT_i^y}{\sum_{i=1}^{10} ex_i^y} = \frac{\sum_{i=1}^{10} NT_i^y}{\sum_{i=1}^{10}\left(\sum_{j=1}^{11} ex_{ij}^y\right)}, \tag{5}$$

it is plotted in Figure 11.

Figure 11 - EZ11 Aggregate Border Ratio.



A decrease of the ratio plotted in Figure 11 indicates that NT has decreased with respect to average export, this suggests that nationals are less oriented towards home consumption. This rationale suits a decrease of the Border Effect (BE) as well, then the Border Ratio can serve as a robustness check of the BE estimation. In our analysis, it confirms the evolution of the BE as it emerges from the estimation in Table 4 and the plot in Figure 8.

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