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Abstract

Whether trade has increased due to the Euro is a question at the heart of lively policy debates and academic research. We revisit the question with a new, more powerful econometric test for end-of-sample breaks to formally identify the timing and duration of the structural break implied by the "Rose effect" on the Euro Area's trade. We find a significant break in 1999Q1 when using a traditional gravity equation, corroborating the general consensus in the literature. However, we find that this break is short lived. Furthermore, we show that the break can be explained both by the marked decrease in real interest rates across the Euro Area and by deepening European institutional integration.

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1 Introduction

Has the Euro contributed to an increase in trade within Europe? The question has attracted particular attention. In policy circles, it directly concerns countries waiting to access the Euro Area. In academic environments, the question resonates with a vibrant literature questioning the link between a common currency and trade. This paper suggests a new methodology to provide rigorous answers to this questions, hoping to inform policy debates about the trade-related advantages of adopting the Euro.

This paper contributes to the literature initiated by the pioneering work of Rose (2000) which concludes that a currency union increases trade by more than 200%.[1] Glick and Rose (2001) later revisit this striking result using panel data (instead of pooled cross sectional data), to emphasize the time dimension of trade creation, and conclude that a country can hope to double its trade by joining a common currency. These results, although supported by a battery of robustness checks, open three fault lines. First, the non-random selection of countries adopting a common currency. Second, the potential reverse causality inciting those prospering from trade to elect a common currency and third, the fact that the countries in the Rose (2000) dataset are mostly small, poor, open, remote, island states, whose idiosynchractic characteristics hardly warrant generalizations. Tenreyro (2001) as well as Tenreyro and Barro (2003) ingeniously solve the problem of endogeneity by using countries having coincidentally anchored their exchange rates to an identical currency. Persson (2001) instead solves the problem of non-random selection by using matching techniques relevant to the econometric literature on treatment effects. While Tenreyro and Barro (2003) still find a large, positive and significant effect of a common currency on trade, Persson (2001) finds a much diminished one. But subsequent tests in Rose (2001) on a longer dataset restore the original findings. In this earlier literature the difficulty lies with the generalization of these results to larger and more developed economies.

The advent of the Euro has served the literature perfectly. It not only provides a case study of major developed economies joining their currencies, but also a natural experiment where many of the methodological criticisms of Rose (2000) no longer play a prominent role. Rose and van Wincoop (2001) are the first to tackle the question of the Euro, but only insofar as its potential effects. Their paper provides out of sample predictions based on transaction costs between European countries relative to other trading partners (following the theoretical impetus by Anderson and van Wincoop, 2003), suggesting that intra-Euro Area trade should increase by 60% after the adoption of the Euro.

[1]For more details on the relevant literature, as well as an illuminating perspective on its development and novel hypotheses, see Baldwin (2006). In this section, we aim to give just enough background to set this paper's approach and results in context.

As the new currency was actually adopted and trade data started trickling in, several papers competed to be the first to explore the so-called "Rose effect". The most complete and encompassing in the literature is Micco, Ordoñez and Stein (2003) [henceforth MOS], and the most methodologically satisfactory, as per Baldwin (2006), is Flam and Nordström (2003). Both run panel regressions with country-pair fixed effects and dummies for the Euro. Flam and Nordstöm (2003) use unilateral trade data, while MOS (2003) rely on bilateral averages. Both find a positive and significant effect of the Euro on trade within Euro Area countries (of the order of 10 to 20%), as well as a slighter, but still noticeable increase in trade with non-Euro Area countries.

Other eminent studies which helped pioneer the literature on the Euro's trade effects are Barr, Breedon and Miles (2003) [henceforth BBM], which also considers the effects on FDI, financial market development and macroeconomic performance, Bun and Klaassen (2002), which uses dynamic panel techniques and deNardis and Vicarelli (2003). Each generally corroborates the results in MOS (2003), although BBM (2003) finds a trade effect of nearly 40% in the long run. Piscitelli (2003) warns us that extending the sample back another decade (to 1980) and the choice of either *fob* or *cif* data could affect the results and possibly lessen the trade effect. Finally, deSousa (2002) no longer finds a trade effect when including a time trend, although such an approach is somewhat ad-hoc and the results do not line up with a similar investigation in MOS (2003). The only paper to fundamentally shake the general consensus around the "Rose effect" is Nitsch and Berger (2005). As Nitsch (2002) had done in the earlier legacy of Rose (2000), Nitsch and Berger (2005) argues that it is primarily political and institutional integration among European countries that has increased trade, not the adoption of a common currency.

Leaving aside the Nitsch and Berger (2005) criticism for the moment, the finding of a "Rose effect" following the Euro's introduction is weak, as it rests on fairly informal testing techniques, quite appropriate for seminal works such as those mentioned above, but now surpassed by more powerful econometric techniques. The authors mentioned above introduce various flavors of dummy variables to capture the introduction of the Euro. But asymptotic analysis underlying the F-type tests that some of the authors employ to test the significance of the dummies are not appropriate given the very few observations after the Euro's introduction, especially when using yearly data, and are bound to be inconsistent. Indeed, some authors like MOS (2003) avoid, in part, the use of explicit tests and rely on eye-balling the size of the coefficients on the Euro-dummies. A second limitation is that most papers do not properly account for the non-stationarity of the series included in the regressions, thus leading to problems of bias in the results.

In this paper, our contribution is twofold. The first is methodological. We construct an end-of-sample structural break test for panel data based on the seminal paper of Andrews (2003). This brings formality and flexibility to the seminal tests of the "Rose effect". The econometric technique estimates the statistical significance of a potential structural break in trade. Importantly, it also allows the measurement of the duration of a break, even if short-lasting. The test explicitly addresses the issue of very few observations following a break, as is the case after the Euro's introduction. Notably, it builds a test statistic whose distribution is estimated using parametric subsampling techniques, and is robust to very few regularity conditions.

The second contribution is to the content of the "Rose effect" literature. In applying this new test we find the following results. As in much of the relevant literature, we find evidence to reject the null of no structural break in Euro Area trade when using a traditional gravity model derived from microfoundations. We find the break to begin in 1999Q1, corresponding to the introduction of the Euro. But contrarily to popular speculation that the increase in trade is only beginning, we find that the structural break is short-lived, lasting for only 10 quarters (2.5 years). Furthermore, we find no evidence of a break in the trade relationship between Euro Area and non Euro Area EU15 countries nor in trade within this later group, as opposed to the positive spillover effects on trade found in MOS (2003) and Flam and Nordström (2003). In addition, we test some of the recent arguments for the causality between a common currency and trade. We show that the break can be explained by the deepening in European-wide political and institutional integration, as postulated by Nitsch (2002) and Nitsch and Berger (2005). Alternatively, we show that the drastic decrease in real interest rates preceding and following the Euro's introduction can also explain the structural break in trade. This is in line with a story of capital accumulation and firm entry as presented in Mancini-Griffoli (2006).

We begin this paper with an overview of the test for end of sample structural break in panel data. In section 3, we introduce our microfounded regression specifications and review our estimation methodology. Section 4 covers data sources and properties. Section 5 presents and discusses our results in details. Finally, we conclude in section 7.

2 A panel structural break test

2.1 Introduction

We propose and use a panel data adaptation of Andrews' (2003) new time-series test aimed at detecting structural breaks at the end of samples. This new technique offers three main practical and technical advantages which directly suit our empirical application. First, the test does not make any distributional assumptions as it estimates empirically the distribution of the test statistic using an empirical subsampling methodology. Second, the power of the test stays high even when there is as little as one observation after the break date. Third, the test requires very few regularity conditions. It remains asymptotically valid despite non-normal, heteroskedastic and/or autocorrelated errors, and non strictly exogenous regressors. Nonetheless it is important that there be a large number of observations prior to the date of the suspected break, and that these be stationary and ergodic. By contrast, traditional F-type test require normal iid errors and strictly exogenous regressors.

Andrews (2003) constructs his test statistic based on improvements on the procedure proposed by Dufour, Ghysels and Hall (1994). The general intuition is that if there is a structural break, the slope coefficients from the post-break period will differ from those in the prior-break period. Thus, errors estimated on post break observations when assuming that the sample is stable (thus using coefficients estimated over the full sample) would be large. The test statistic is correspondingly built around these errors, as the estimated residuals squared divided by the estimated covariance matrix of the residuals. The appropriate distribution against which to gauge the size of this statistic is built by estimating equivalent statistics but over the pre-break sample, then plotting their density function. Asymptotically valid critical values are then found in a straightforward manner, as the points below which lie a given percentage of the pre-break statistics. Unfortunately, the Andrews (2003) testing procedure cannot be used directly in a panel data context; we therefore propose a practical adaptation of it to suite our work.

2.2 The setting

The regression serving as the basis for tests of structural break is:

$$Y_{it} = \begin{cases} X_{it}'\beta_0 + U_{it} & t = 1, \dots, T \\ X_{it}'\beta_{1t} + U_{it} & t = T+1, \dots, T+m \end{cases} \tag{1}$$

for individuals $i = 1, \dots, n$, and where T is the postulated break date. The test naturally hinges on the following hypotheses: $H_0 : \beta_{1t} = \beta_0$ against $H_A : \beta_{1t} \neq \beta_0$ The coefficients are homogeneous across i under both the null and the alternative hypothesis. The alternative hypothesis requires all individuals to exhibit an end-of-sample break.

2.3 The test statistic

Consider the case when there are more observations after the break date than regressors d, so that $(m \times n) \geq d$. In words, the test statistic is a positive definite quadratic form obtained from the transformed $(m \times n) \times 1$ vector of residuals by the $(m \times n) \times (m \times n)$ covariance matrix, projected onto the column space of the $(m \times n) \times d$ matrix of transformed post-instability regressors. The panel data equivalent of the generic test statistic in Andrews

(2003) can be defined after considering an interval τ_r which spans from $[r, r+m-1]$ and where $r \in \{1, ..., T+1\}$, as:

$$S_r(\beta, \mathbf{\Sigma}) = A_r(\beta, \mathbf{\Sigma})' V_r^{-1} A_r(\beta, \mathbf{\Sigma}), \tag{2}$$

$$A_r(\beta, \mathbf{\Sigma}) = \mathbf{X}_{\tau_r}' \widehat{\mathbf{\Sigma}}_{T+m}^{-1} \widehat{\mathbf{W}}_{\tau_r}, \tag{3}$$

$$V_r(\mathbf{\Sigma}) = \mathbf{X}_{\tau_r}' \widehat{\mathbf{\Sigma}}_{T+m}^{-1} \mathbf{X}_{\tau_r} \tag{4}$$

with

$$\widehat{\mathbf{W}}_{\tau_r} = (\mathbf{Y}_{\tau_r} - \mathbf{X}_{\tau_r}\beta)$$

where $\widehat{\mathbf{W}}_{\tau_r}$ is the $(m \times n) \times 1$ residual vector of observations starting at r, with $\beta = \widehat{\beta}_{T+m}$ defined to be the coefficient vector estimated over the full sample (until $T+m$). The variance-covariance matrix, $\widehat{\mathbf{\Sigma}}_{T+m}$, is given by:

$$\widehat{\mathbf{\Sigma}}_{T+m} = (T+1)^{-1} \sum_{r=1}^{T+1} \left(\widehat{\mathbf{U}}_{\tau_r} \widehat{\mathbf{U}}_{\tau_r}' \right) \tag{5}$$

where the $(m \times n) \times 1$ residual vector, $\widehat{\mathbf{U}}_{\tau_r}$, is:

$$\widehat{\mathbf{U}}_{\tau_r} = (\mathbf{Y}_{\tau_r} - \mathbf{X}_{\tau_r}\widehat{\beta}_{T+m})$$

thus a particular form of the $\widehat{\mathbf{W}}_{\tau_r}$ vector.

This covariance matrix corrects for serially correlated errors, heteresko-dasticity and potential cross-sectional correlation. These are the elements found off the block diagonal of $\widehat{\mathbf{\Sigma}}_{T+m}$ (the on diagonal elements being the variances for each individual and the elements in the $T \times T$ blocks on diagonal being the autocovariances for each individual).

The particular form of the test statistic for the post-break residuals - the central statistic to the test - is a special form of the generic statistic defined above. As in Andrews (2003), we call this the S statistic and define it as:

$$S = S_{T+1}(\widehat{\beta}_{T+m}, \widehat{\mathbf{\Sigma}}_{T+m}) \tag{6}$$

Note that when $(m \times n) \leq d$ (there are fewer post-instability observations than regressors in the model), Andrews (2003) suggests using $P = P_{T+1}(\widehat{\beta}_{T+m}, \widehat{\mathbf{\Sigma}}_{T+m})$ where the projection matrix collapses to the identity matrix I_{mn}.

2.4 Critical values

The critical values are found by empirically generating a distribution function for the statistic under the null of stability. In practice, we find equivalent statistics to S over the pre-break subsample through a rolling window method spanning m observations (as there are in S). We label these S_r. When $(m \times n) \geq d$ the $T-m+1$ different S_r values are defined as:

$$S_r = S_r(\widehat{\beta}_{2,(r)}, \widehat{\mathbf{\Sigma}}_{T+m}) \tag{7}$$

where $\widehat{\beta}_{2,(r)}$ is the estimate of β over $t = 1, \ldots, T$ observations but excluding $r, \ldots, r + \frac{m}{2} - 1$, or $\frac{m}{2}$ observations. The reason for excluding these observations, as given in Andrews (2003) is to optimize both the size and power of the test in comparison to a $\widehat{\beta}_{2,(r)}$ excluding m observations or no observations at all.

Andrews (2003) shows that the empirical cummulative density function (CDF) of the S_r values is asymptotically unbiased and consistent. The p-values are given by:

$$p - values = (T - n + 1)^{-1} \sum_{j=1}^{T-m+1} 1[S \leq S_r]$$

where $1[\cdot]$ is an indicator function.

2.5 Panel-specific adaptations

The transfer of each step of the Andrews' (2003) test to panel data is not immediate. In particular, the variance-covariance matrix $\widehat{\Sigma}_{T+m}$ as defined above will not be invertible in most cases, as it will generally not be of full rank. The component matrices composed of sub-vectors of residuals, $\widehat{U}_{\tau_r}\widehat{U}'_{\tau_r}$, are of rank one, as is always the case with outer products. As each of these $\widehat{U}_{\tau_r}\widehat{U}'_{\tau_r}$ matrices are added, the resulting matrix $\widehat{\Sigma}_{\tau_r}$ gains one in rank.[2] Such that the final $\widehat{\Sigma}_{T+m}$ matrix be of full rank and invertible, there must be at least as many additions of the $\widehat{U}_{\tau_r}\widehat{U}'_{\tau_r}$ matrices than there are dimensions of $\widehat{\Sigma}_{T+m}$. Thus, the condition for invertibility boils down to $T + 1 \geq (m \times n)$.

Unfortunately, this condition is rarely satisfied in a panel setting. Although m is usually small, n can be very large. The product of the two is usually greater than there are time periods $(T + 1)$ before the presumed point of instability, unless the pre-break subsample is unusually large. Note that running the Andrews (2003) test in a time series setting does not engender this complication. In a time series model, $\widehat{\Sigma}_{T+m}$ is invertible if and only if $T + 1 \geq m$, a condition that is easily satisfied.

Hence, we must impose certain restrictions on the $(m \times n) \times (m \times n)$ covariance matrix in order to invert it. We redefine a covariance matrix assuming sectional independence, as is often done in the panel literature, although continue to allow for serial correlation and heteroskedasticity:

$$\tilde{\mathbf{\Sigma}}_{T+m} = (T + 1)^{-1} \sum_{r=1}^{T+1} \left(\widehat{\mathbf{U}}_{\tau_r} \widehat{\mathbf{U}}'_{\tau_r} \right) = \widehat{\mathbf{\Sigma}}_{T+m} \tag{8}$$

[2]This comes directly from a well known theorem on rank additivity, stating that for any two matrices A and B, $rank(A + B) = rank(A) + rank(B)$ if and only if the row space of A and that of B are essentially disjoint and the column space of A and that of B are essentially disjoint.

except that $E\left[\mathbf{U}_{i,\tau_r}\mathbf{U}^{\prime}_{j,\tau_r}|\mathbf{X}_{ij}\right] = \mathbf{0}$, for $i \neq j$ with $i, j = 1, \ldots, n$, and $\mathbf{U}_{i,\tau_r}$ is an $m \times 1$ vector made up of the elements in $\mathbf{U}_{\tau_r}$ corresponding to individual i.

The resulting covariance matrix $\tilde{\mathbf{\Sigma}}_{T+m}$ is block diagonal. Each block corresponds to an individual in our panel, and is thus of dimension $(m \times m)$. Since the inverse of a block diagonal matrix is the inverse of each of its blocks, the condition for invertibility reduces to that expressed for times series (namely that $T + 1 \geq m$), which we satisfy. In the appendix we discuss alternative as well as more general conditions for the inversion of the covariance matrix, possibly useful for other applications of the test.

3 Model specifications and methodology

3.1 Baseline model

We work with a gravity equation similar to those used in the literature to date, but with some slight adjustments in conformity with Baldwin's (2006) remarks on the mistakes and biases introduced with ad-hoc model specifications. In the appendix, we derive our basic gravity equation from microfoundations. This equation is:

$$V_{i,j,t} = \alpha_{i,j} + \gamma_1 Y_{i,t} + \gamma_2 Y_{j,t} + \gamma_3 \xi_{i,j,t} + \epsilon_{i,t} \tag{9}$$

where $V_{i,j,t}$ is the value of imports from country j to country i, $Y_{j,t}$ and $Y_{i,t}$ are nominal GDP, $\xi_{i,j,t}$ is the real exchange rate between the two countries engaged in trade, and ϵ_t is a regression error.

Furthermore, $\alpha_{i,j}$ is a pair-specific fixed effect to control for variables of type common border, language, history, legal system, distance and others traditionally shown to matter in gravity equations. The advantage of this "agnostic" approach, as opposed to a fully fleshed out specification of each independent variable, is that we do not run the risk of leaving out a regressor, or mis-measuring one, as is commonly done with variables such as distance.[3]

Also, we work with a homogeneous panel, implying constant slope coefficients for all country pairs. This assumption is in line with the Null hypothesis that the relationship between trade and its explanatory variables has remained stable *across* Euroland.

Finally, we include in the error term the export country specific time trend derived in the appendix. First, doing so makes our results more comparable with those in the literature. To the extent that time trends are used in the literature, they capture more general phenomena like political and economic integration, factors that we control for explicitly in our model. [4]

[3] Flam and Nordström (2003), and MOS (2003) adopt the same approach.

[4] See, for instance, Cheng and Wall (2005), MOS (2003), Flam and Nordström (2003), Rose and van Wincoop (2001) or Bayoumi and Eichengreen (1997).

Second, since the regressors are all time varying, including time trends could over-correct for the effect of these on trade and may thus lead to erroneous coefficient estimates.[5] Third, our test for end-of-sample structural break corrects for the serial correlation introduced in the errors by the time trend.

3.2 Refinements of baseline model

As we will see in the discussion of results, estimating model (9) above clearly suggests a break in trade among Euro Area countries due to the introduction of the Euro. But *what* exactly in the new currency is responsible to boost trade? Several explanations exist. This paper will test the most prominent and recent ones. The first argument suggests that the Euro is besides the point; its introduction simply coincided with an accelerating process of European-wide political and institutional reforms favoring trade. Nitsch (2002) and Nitsch and Berger (2005) are the main proponents of this line of reasoning. Other arguments pertain more closely to the Euro itself. Baldwin (2006) suggests that of the variables entering the gravity equation, a rise in n, the number of exporting firms in a given country, is likely to be the key to explain trade creation. More specifically, Baldwin and Taglioni (2005) suggest that the disappearance of exchange rate risk induced the small and medium size firms formerly unable to protect themselves against currency fluctuations to enter the export market. Alternatively, Mancini-Griffoli (2006) suggests that it was the decrease in real interest rates, coming as a pre-condition and consequence of the Euro, that favored the entry of new firms and the expansion of existing firms in the costly export business.

Each of these hypotheses can be evaluated with our panel test for end of sample structural break, based on modified versions of our baseline gravity equation (9). Our conjecture is that we will no longer be able to reject the Null of stability when controlling for the additional explanatory variables suggested by the above hypotheses if these are valid. To study the argument advanced by Nitsch and Berger (2005), we add a term, $\tau_{i,t}$, as justified in the appendix, to capture the integration of country i in the European institutional and political process. To proxy for this variable, we borrow Nitsch and Berger's (2005) European Integration Index. The resulting specification is labeled model (B) below.

The story in Mancini-Griffoli (2006) is slightly more complex. It is derived from the literature on firm entry in which firms must pay a fixed cost to enter the export business. This cost is determined in terms of capital and is thus dependent on real interest rates which are set exogenously. As these decrease, essentially due to the accession criteria accompanying the Euro, investment in capital rises and more firms export. The paper derives model (C) below, where $\bar{R}_{i,t} = 1/6 \sum_{s=1}^{6} R_{i,t-s}$ captures average real interest

[5] This point is also raised in IMF (2004) as one of the main reasons not to use country-specific time varying effects.

rates over 6 quarters, deemed to be the metric most influential for capital investment decisions (we alter this specifications in our robustness checks). Wages appear as a second factor price, with a lag of four periods in line with findings that wages are sticky for approximately one year.

Our regression specifications are listed below, where, for comparison's sake, we label our baseline model (9) as model (A).

(A) $V_{i,j,t} = \alpha_{i,j} + \gamma_1 Y_{i,t} + \gamma_2 Y_{j,t} + \gamma_3 \xi_{i,j,t} + \epsilon_{i,t}$

(B) $V_{i,j,t} = \alpha_{i,j} + \gamma_1 Y_{i,t} + \gamma_2 Y_{j,t} + \gamma_3 \xi_{i,j,t} + \gamma_4 \tau_{i,t} + \epsilon_{i,t}$

(C) $V_{i,j,t} = \alpha_{i,j} + \gamma_1 Y_{i,t} + \gamma_2 Y_{j,t} + \gamma_3 \xi_{i,j,t} + \gamma_5 \bar{R}_{i,t} + \gamma_6 W_{i,t-4} + \epsilon_{i,t}$

3.3 Estimation methodology: Error Correction Model

Our data analysis will show that all variables in our dataset are integrated of order one and are cointegrated. In our estimation, we will therefore rely primarily on models in error correction model (ECM). The ECM entails estimating the model in first differences, while controlling for the cointegrating relationship. Working with first differenced series allows us to meet the test's requirement of stationary and ergodic variables. We build an ECM model in two stages, as is customary, where the first stage estimates the cointegration vector and the second, the equation in first differences while correcting for cointegration. For model (A), this yields:

$$V_{i,j,t} = \alpha_{i-j} + \gamma_1 Y_{i,t} + \gamma_2 Y_{j,t} + \gamma_3 \xi_{i,j,t} + \epsilon_{i,t} \tag{10}$$

$$\Delta V_{i,j,t} = \beta \widehat{\epsilon}_{i,t-1} + \lambda_1 \Delta Y_{i,t} + \lambda_2 \Delta Y_{j,t} + \lambda_3 \Delta \xi_{i,j,t} + u_{i,t} \tag{11}$$

where the first equation is estimated using difference from sample mean fixed effects and the second using pooled OLS.

4 Data

4.1 Description

Most of the literature investigating the effect of a common currency uses annual data, often featuring a small time dimension. We instead use quarterly data from 1980 Q1 to 2004 Q4, as our estimation method requires a large time dimension. Our sample is composed of the EU-15 countries, subdivided into four trading groups: imports of the Euro Area (EA) from Euro Area, of the Non-Euro Area (NEA) from Non-Euro Area, of the NEA from the EA and of the EA from the NEA. We exclude Greece from the EA, since it joined the Euro only in January 2001. As is commonly done, we also group Luxembourg and Belgium as their trade data are confounded over most of our sample period. The main group of interest is the Euro Area. The other trading groups are used a controls.

The data were obtained from Eurostat, IMF DOTS and IFS, as in most other relevant empirical papers. A complete table of the data sources is available in appendix C. We use the Nitsch and Berger (2005) index of european integration instead of dummies to track the institutional and political integration among EU countries emanating from the signing, ratification and implementation of major treaties. We use the unilateral import values as trade data, obtained from IMF DOTS. Lastly, we adjusted the data for seasonality when necessary.

Two further points are worth discussing. First, there is a strong argument in favor of using unilateral instead of bilateral trade flows. Flam and Nordström (2003) follow the same route, also recommended in Baldwin (2006). At the most basic level, a gravity equation is essentially a demand equation. Thus, demand in country i ought to be different from that in country j. Also, keeping trade flows separate has the advantage of being able to pin-point with more precision the effects of a domestic versus a foreign explanatory variable. The reverse of the coin is that two observations of trade will share most of the same regressors.

Secondly, we acknowledge that in principle, working with volumes is preferable as it allows us to concentrate on changes in the quantity of exported goods independently of any price movements. Yet, with anything but firm level data, working with volumes introduces severe measurement errors. The problem is particularly acute with aggregate data such as that used in this paper. First, it is not clear which aggregate price index to use to transform aggregate imports reported in value into volume. Second, even if a consistent and accurate import price index existed, it would never be appropriate to apply to imports from two different countries which export two different bundles of goods. Due to these non-negligeable limitations, we prefer using values. The fact that we control for relative price movements by including real exchange rates in our regression dampens any potential distortions due to the use of values instead of volumes of imports.

4.2 Unit root and cointegration results

In table 1, we present our results for the unit root tests applied to Euro Area data. We apply the Breitung (1997,1999) as well as Im, Pesaran and Shin (IPS, 2003) procedures. As these results show, we cannot reject the Null of a unit root with reasonable significance for any series, although the evidence is mixed for interest rates. Because of the various convergence criteria to enter a monetary union, real interest rates were not entirely stationary over our sample period, as would otherwise be expected. In particular, there has been a marked decrease and convergence in interest rates among the eventual Euro Area countries starting around 1996Q1. [TABLE 1]

Figure 1 shows how Euro Area average real interest rates (over six quarters) were relatively stable (just below 10%) until about 1996, when a noticeable downward trend began. In four years, real interest rates lost about 600 basis points, and four years later, after a further 200 basis point decrease, were at their lowest, at around 2%. [FIGURE 1] Figure 2 corroborates this finding by showing how year-on-year growth in average interest rates hovered slightly below zero until about 1996, after which it remained decisively negative, in the ballpark of -10% or more. The panel test for structural break requires that the series in the estimation subsample be stationary and ergodic. In the light of our remarks above, we opt for a conservative assumption that also interest rates are non-stationary. [FIGURE 2]

Table 2 reports test results for four different cointegrating vectors corresponding to the four model specification that we investigate in our empirical section. We adopt the Pedroni (1999, 2004) tests, which have the advantage of allowing for significant heterogeneity between cross-sections. The results indeed show the presence of cointegration between all series. We view this as an encouraging finding of a significant long term relationship among our variables. For simplicity we only report Pedroni's (1999) ρ and ν-statistics for the Euro Area group. Similar unit root and cointegration results were found for the three other control groups (results are available upon request). [TABLE 2]

5 Empirical Results and Discussion

5.1 Criteria to find the break date

We are faced with a tradeoff when deciding what constitutes a structural break in terms of time span in the post-break period. On the one hand, if we reject the null of no structural break over a short time span (for example 3 quarters), we may be facing an outlier rather than an actual structural change in the pattern of the errors. On the other hand, using a long post-break sample prevents the detection of instabilities if the break is short-lived. We choose a "break criteria" of 6 quarters as a middle ground. Thus, the "break date" is defined as the first period for which we reject the null of stability with at least 90% confidence, for a post-break sample of 6 quarters.

The trade literature on the effect of the Euro has debated at length when the break in trade has occurred. The Euro was introduced in January 1999, but some argue that expectations of the Euro created a trade effect already a year earlier. We test the baseline model (A) for various potential break dates, starting in 1998Q1. According to our "break criteria", we pinpoint the break date to be in 1999Q1, thus coinciding with the actual introduction of the common currency. We do not read more into this result than a mere coincidence, although the uncertainty around some countries' likelihood to adopt the Euro (like Italy's) may have played a role in delaying European-wide trade effects.

5.2 Finding a break in Model A

We first concentrate on the Euro Area sample. We test the null of no structural break on the baseline model (A). As mentioned above, we reject the Null at the 10% significance level for a post-break period of six quarters starting in 1999 Q1. We subsequently increase the time span by one quarter at a time and re-test the null hypothesis. Our results are shown in table 3. We find that the probability of rejecting the null is highest (1% level) when the post-break time period spans from 1999 Q1 to 2000 Q4 (7 quarters). Evidence of a structural break at the 10% level is found up to the 10^{th} quarter, or two years and a half (1999 Q1 - 2001 Q3), after the break. Starting with the 11^{th} quarter, the null of no structural break is not even weakly rejected. [TABLE 3]

We control our findings for the Euro Area countries by conducting the same analysis on trade between non Euro Area EU15 countries, as well as between non-Euro Area EU15 countries with Euro Area countries in both directions. Unless there were strong spillover effects on trade, we conjectured that if we could not reject the Null of stability in any of these trading relationships, the evidence favoring the trade effect of the Euro would be strengthened.

Our findings are summarized in tables $4-6$. Indeed, we do not reject the null of stability for trade among non Euro Area countries. Likewise, we fail to reject the Null for trade emanating from non-Euro Area countries to Euro Area countries, as well as from Euro Area countries to non-Euro Area countries. [TABLE 4 -6]

5.3 Explaining the break in trade

When we test model specification (B), which includes the European integration variable, we do not reject the null hypothesis even at the 10% level for time spans of 6 quarters and more, as can be see in table 7. It seems that our evidence corroborates the Nitsch and Berger (2005) argument, namely that political and institutional integration in the European Union is a decisive factor in explaining the trade effect of the Euro. The limit of this explanation is the short time period over which trade increases. Political integration is a continuous, smooth and still increasing process. It is unclear why trade would really only exhibit a break in the late nineties and subsequent quarters if institutional integration were the sole explanation. [TABLE 7]

We find that model (C) which includes interest rates and wages, can equally explain the break in trade, as shown in table 8. For any given break period, the S statistic decreases far below its equivalent measure under model (A) with respect to the S_r distribution. The probability of rejecting the null of stability decreases below the 10% level for all post-break sample periods. Most notably, for the period where the break is strongest in model (A), 1999 Q1 - 2000 Q4, controlling for interest rates makes the difference between rejecting (at the 1% level) and not rejecting the null hypothesis at all. If decreasing interest rates are indeed part of the story behind the boom in Euro Area trade, as these results suggest and according to the theoretical model in Mancini-Griffoli (2006), the rather late break date with respect to expectations of adopting the Euro may be explained. Since interest rates started decreasing persistently between 1996 and 1997, it is normal that effects of capital accumulation would only be felt about two years later, due to the lag between installing and benefitting from new capital. Indeed, model (C) includes average interest rates over six quarters to account for this time to build characteristic. Also, the explanation attached to interest rates conveniently fits the short time span of the break in trade, as real interest rates can only realistically decrease (or be expected to decrease) for a limited time. In fact, figures 1 and 2 show that there was an important correction in the downward trend in interest rates starting in 2000Q4 and lasting approximately 6 quarters. This may have contributed to shortening or abating the perceived trade effect of the Euro. [TABLE 8]

We conduct several robustness checks to verify whether the marked decrease in interest rates can really explain the structural break found in model (A) for Euro Area data. For these, we use an alternate definition of average interest rates. For the tests mentioned above, we had defined $\bar{R}_{i,t}$ to include 6 lags. In tables 10 and 11 below, we show that even if it were to include 4 or 2 lags, we still fail to reject the null of no structural break. Thus, altering the lag structure of the average interest rate does not change our main findings. [TABLE 9 - 10]

5.4 Regression results

To build further confidence in our findings, we present here the estimation results of the different model specifications analysed in this paper. We use two estimation methodologies: the ECM as in the testing procedure and a Dynamic OLS (DOLS) method. The DOLS estimator only focusses on the long run relationship, but makes appropriate corrections yielding unbiased coefficients and correct standard errors. Kao and Chiang (1999) show that the OLS estimator exhibits a non-negligeable bias with finite samples in panel cointegrated regression models. As an alternative, Kao and Chiang, Chen(1999), Kao and Chiang (1998), Phillips and Moon (1999, 2000), as well as Pedroni (2000), suggest, develop and compare the properties of other estimators, concentrating on the FMOLS and DOLS procedures.[6] Both aim to correct for two biases, namely the serial correlation and endogeneity created from the integrated series. The authors generally come to the agreement that the DOLS estimator out-performs the FMOLS estimator in terms of non-biaseness. Indeed, in several real-world applications, the DOLS procedure stands out as the most robust.[7] The procedure entails estimating the model with all independent variables in levels, as well as leads and lags of variables in first differences.[8] To illustrate, model (A) in DOLS form gives:

$$\begin{aligned} V_{i,j,t} &= \alpha_{i,j} + \gamma_1 Y_{i,t} + \gamma_2 Y_{j,t} + \gamma_3 \xi_{i,j,t} + \sum_{s=-p}^{p} \delta_{1,s} \Delta Y_{i,t+s} \\ &\quad + \sum_{s=-p}^{p} \delta_{2,s} Y_{j,t+s} + \sum_{s=-p}^{p} \delta_{3,s} \Delta \xi_{i,j,t+s} + \epsilon_t \qquad (12) \end{aligned}$$

where we set $p = 4$ for each of our variables, except in the case of model (C), where $p = 2$ for interest rates and wages since both are already lagged.

Our results for the DOLS regressions are illustrated in table 12. The coefficients and signs on the long run cointegrated variables appear as expected and match those found in similar studies in the relevant literature. The coefficient on GDP (both domestic and foreign) is positive and significant. That on real exchange rates is negative (a depreciation causes a decrease in imports), but hardly significant (as in MOS, 2003). The coefficient on integration is positive and significant, as touted by Nitsch and Berger (2005). Average interest rates appear as negative and significant, with a rather small magnitidue (indicating a realistic relationship between interest rates and trade: a decrease of 1% in interest rates increase trade by 0.07-0.1%). It is also encouraging to see that the inclusion of interest rates does not markedly change the coefficients on GDP, thereby underscoring that there do not seem to be problems of multicollinearity among our regressors. The only surprise at first glance is the magnitude and significance level of wages. But these can be attributed to the high degree of correlation between wages and trade, both smooth, upward sloping series. [TABLE 11]

Table 13 reports coefficients from the error correction model. The coefficients are as expected in terms of magnitude and sign, mainly as discussed above. Reassuringly, changes in wages now appear to have realistic magnitude, but are no longer significant. This is rather expected as there has been very little variation in real wages within the EU in the last twenty

[6] See Baltagi and Kao (2000) for a survey. The DOLS estimator is based on the single equation dynamic ordinary least squares method of Saikkonen (1991) and Stock and Watson (1993).

[7] See Hafner (2005), Mark and Sul (2003), Kao, Chiang and Chen (1999) for tests of the PPP hypothesis or the impact of technology diffusion on TFP.

[8] See Mark and Sul (2003) for possibly the clearest explanation for the particular form of the DOLS regression equation.

years. More importantly, average interest rate again exhibit a negative sign and realistic proportions. Note that the t-stats are reported for indicative purposes only, as they do not feature robust standard errors (due to reasons discussed above). [TABLE 12]

6 Conclusion

We developed an extension of the end-of-sample structural break test found in Andrews (2003) to panel data. The panel structure prevents a direct application of the Andrews (2003) test, and requires an adaptation of the covariance matrix so that it becomes invertible. This rigorous testing procedure, built on a distribution estimated with empirical subsampling techniques, is robust to very few post-break observations and allowed us to formally evaluate the question of Euro's effect on trade. The test could also be applied to wide range of other questions tied to the Euro's introduction, rousing much speculation but needing more solid statistical grounding. These include a distinct rise in cross-national business cycle correlations, a drop in the degree of national price discrimination or an increase in employment, for instance.

With respect to the question of trade, we corroborate the findings generally found in the literature of a break in trade when using a baseline gravity regression derived from microfoundations. But unlike the earlier literature, we are able to attribute a precise significance level to this finding. We also provide new evidence for the timing and duration of the break in trade. We show that the break starts in 1999Q1, but is short-lived as it spans just two and a half years (1999Q1 - 2001Q3). Finally, we go one step further in testing possible explanations for the disproportionate increase in trade. We find evidence for the importance of institutional and political integration between european countries, as well as the marked decrease in real interest rates preceding and following the introduction of the Euro. We build confidence for our results by considering three control groups. Indeed, we find that there is no significant break in trade between non Euro-Area (but EU 15) countries, nor from non-Euro-Area to Euro-Area countries and vice versa.

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Table 1: Unit Root tests for the Euro Area series

	Breitung test				IPS test			
	intercept		int. & slope		intercept		int. & slope	
	Stat	Prob.	Stat	Prob.	Stat	Prob.	Stat	Prob.
$\mathrm{V_{i,j,t}}$	-1.05	0.15	1.42	0.92	3.09	1.00	-7.62	0.00
$\mathrm{Y_{i,t}}$	-0.32	0.37	1.89	0.97	0.76	1.00	0.17	0.57
$\tau_{\mathrm{j,t}}$	1.38	0.92	1.44	0.93	7.17	1.00	4.40	1.00
$\mathrm{w_{i,t}}$	0.23	0.59	-0.45	0.33	3.22	1.00	-1.63	0.05
$\mathrm{\bar{R}_{i,t}}$	-1.08	0.13	-2.47	0.006	12.6	1.00	2.87	0.99

Table 2: Pedroni's (1999) Cointegration Results

Specification	$\nu - statistic$	$\rho - statistic$
(A)	12.6	-26.5
(B)	11.9	-26.4
(C)	8.8	-22.3

Table 3: Baseline Model (A)

Span	S	S_r: 1%	S_r: 5%	S_r: 10%
$1999Q1 - 2000Q3$	61.8	82	67.4	53.5
$1999Q1 - 2000Q4$	73.1	72.8	64.4	55
$1999Q1 - 2001Q1$	50	69.5	61.4	51.6
$1999Q1 - 2001Q3$	50.2	65.4	56.9	48.9
$1999Q1 - 2002Q1$	36.9	62.5	50.7	44.7
$1999Q1 - 2004Q4$	22	87.9	66.4	54.6

Table 4: Trade in Non Euro Area EU15

Break Period	S	S_r: 1%	S_r: 5%	S_r: 10%
$1999Q1 - 2000Q3$	4.8	27.1	11.2	7.7
$1999Q1 - 2000Q4$	3.9	27.2	11.3	9.1
$1999Q1 - 2001Q1$	5.5	27.7	12.3	8.9
$1999Q1 - 2001Q3$	3.9	28.7	11.5	10.6
$1999Q1 - 2004Q4$	0.89	19.6	13.1	9.9

Table 5: Trade from Non Euro Area EU15 to Euro Area

Break Period	S	S_r: 1%	S_r: 5%	S_r: 10%
$1999Q1 - 2000Q3$	11.7	57.2	40.2	25.8
$1999Q1 - 2000Q4$	13.2	56.9	38.5	20.2
$1999Q1 - 2001Q1$	9.4	61.7	35	24
$1999Q1 - 2001Q3$	6.6	55.5	32.6	26.4
$1999Q1 - 2004Q4$	5.6	63.9	35.5	29.1

Table 6: Trade from Euro Area to Non Euro Area EU15

Break Period	S	S_r: 1%	S_r: 5%	S_r: 10%
$1999Q1 - 2000Q3$	12.1	99.7	37.9	31.8
$1999Q1 - 2000Q4$	10.4	100	37.8	32.1
$1999Q1 - 2001Q1$	3.1	99.8	42	33.4
$1999Q1 - 2001Q3$	2.7	86.8	46	30.2
$1999Q1 - 2004Q4$	8.1	71.1	61	45.2

Table 7: Baseline with Integration (B)

Break Period	S	S_r: 1%	S_r: 5%	S_r: 10%
$1999Q1 - 2000Q3$	47.1	135	94.2	74.6
$1999Q1 - 2000Q4$	53.6	136	96.2	78.4
$1999Q1 - 2001Q1$	30.3	122	93.8	82.6
$1999Q1 - 2001Q3$	31.7	124	103	76.2
$1999Q1 - 2002Q1$	30.4	131	92.7	85.2
$1999Q1 - 2004Q4$	21.8	142	131	127

Table 8: Augmented Gravity Model (C)

Break Period	S	S_r: 1%	S_r: 5%	S_r: 10%
$1999Q1 - 2000Q3$	51.9	119	99.2	74.1
$1999Q1 - 2000Q4$	66.4	105	92.3	81.3
$1999Q1 - 2001Q1$	45.8	106	92.6	72.3
$1999Q1 - 2001Q3$	46.6	98.8	89.2	82
$1999Q1 - 2002Q1$	39.6	96.7	81.7	65.9
$1999Q1 - 2004Q4$	37.8	114	95.5	79.7

Table 9: Average Interest Rate with 4 lags

Span	S	S_r: 1%	S_r: 5%	S_r: 10%
$1999Q1 - 2000Q4$	62.4	101	89.5	75
$1999Q1 - 2001Q3$	42.8	93	89.5	72.5
$1999Q1 - 2002Q3$	37.4	86.1	77.2	72.2

Table 10: Average Interest Rate with 2 lags

Span	S	S_r: 1%	S_r: 5%	S_r: 10%
$1999Q1 - 2000Q4$	67.9	108	87.1	68.2
$1999Q1 - 2001Q3$	47.3	95.7	88.3	72.1
$1999Q1 - 2002Q3$	41.1	90.8	78.6	71.4

Table 11: DOLS Results for $1980Q1 - 1999Q1$

Specification	$\mathrm{Y_{i,t}}$	$\mathrm{Y_{j,t}}$	$\xi_{\mathrm{i,j,t}}$	$\tau_{\mathrm{i,t}}$	$\mathrm{w_{j,t-4}}$	$\bar{\mathrm{R}}_{\mathrm{j,t}}$
(A)	1.09	0.37	-0.05			
	(44.4)	(15.2)	(-2.4)			
(B)	1.07	0.32	-0.046	0.64		
	(43.6)	(13)	(-2.3)	(-11.4)	(1.2)	
(C)	0.92	0.35	-0.02		0.49	-0.1
	(36.9)	(14.8)	(-1.05)	(12.7)		(-11)

t-statistics appear in parentheses.

Table 12: ECM Results for $1980Q1 - 1999Q1$

Specification	$\Delta\mathrm{Y_{i,t}}$	$\Delta\mathrm{Y_{j,t}}$	$\Delta\xi_{\mathrm{i,j,t}}$	$\widehat{\varepsilon}_{i,t-1}$	$\Delta\tau_{\mathrm{i,t}}$	$\Delta\mathrm{w_{j,t-4}}$	$\Delta\bar{\mathrm{R}}_{\mathrm{j,t}}$
(A)	0.79	0.36	-0.03	-0.14			
	(14.8)	(6.7)	(-0.35)	(-24.3)			
(B)	0.84	0.38	-0.03	-15	0.22		
	(15.6)	(7.2)	(-0.36)	(-24.4)	(1.2)		
(C)	0.74	0.4	-0.01	-0.17		0.005	-0.06
	(12.9)	(7)	(-0.16)	(-24.8)		(0.04)	(-2.2)

t-statistics appear in parentheses.

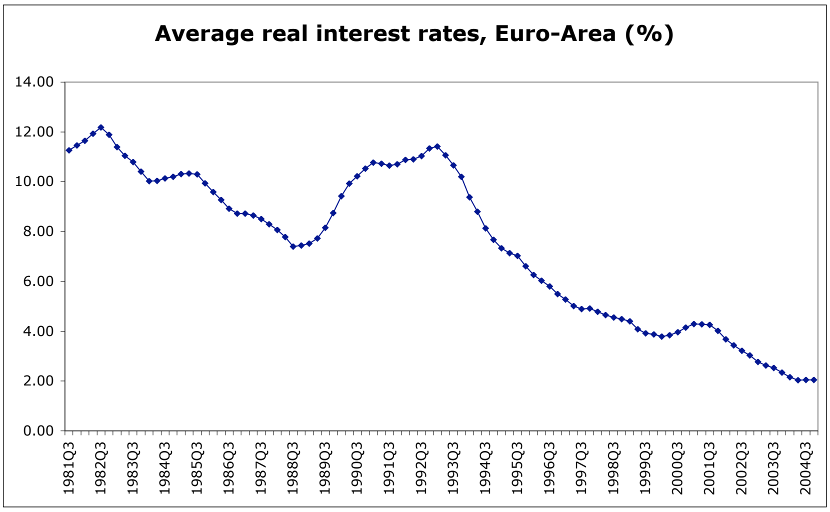

Figure 1: Average real interest rates (over six quarters) were relatively stable (just below 10% on average) until about 1996, when a noticeable downward trend began. In four years, real interest rates lost about 600 basis points, and four years later, after a further 200 basis point decrease, were at their lowest, at around 2%.

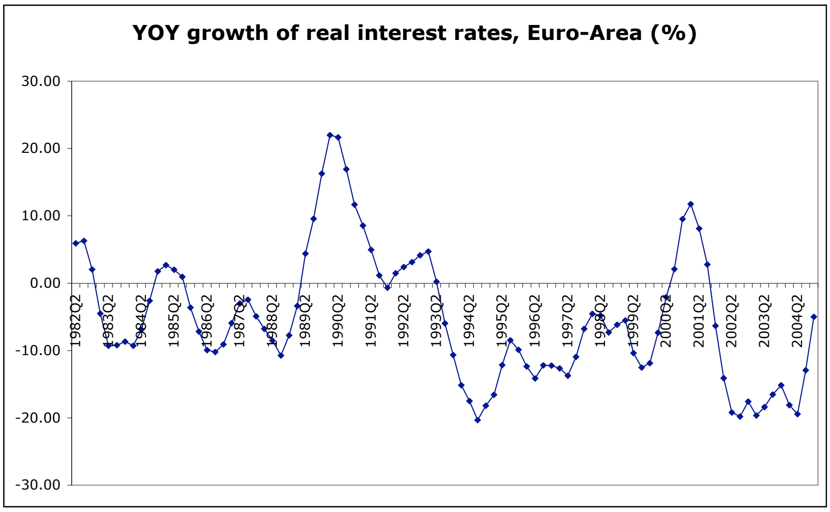

Figure 2: Year-on-year growth of average real interest rates hovered slightly below zero until about 1996, after which it remained decisively negative, in the order of -10% or more.

A Alternative Covariance Matrices

The main problem in panel data, is that the covariance matrix Σ_{T+m} usually does not satisfy the general invertibility condition $T+1 \geq (m \times n)$. In the text we therefore consider $\tilde{\Sigma}_{T+m}$ that is block-diagonal, as we assume cross sectional independence. The resulting invertibility condition simplifies to $T+1 \geq m$, which we satisfy. But other assumptions are also possible. We present here two alternatives. The first is trivial, but allows Σ_{T+m} to be inverted even when $T+1 < m$. The second, is a general condition for the invertibility of Σ_{T+m}, which does not necessarily entail null cross sectional covariances.

The first possibility is to assume, no serial correlation or cross-sectional correlation. Then any of the following may describe the dataset (of which the first is most general, and the second and third may be applicable depending on circumstances):

1. Heteroskedasticity: $E\left[\mathbf{U}_{\tau_r}\mathbf{U}_{\tau_r}^{'}|\mathbf{X}_{\tau_r}\right]$ is a diagonal matrix.
2. Homoskedasticity for all individuals: $E\left[\mathbf{U}_{\tau_r}\mathbf{U}_{\tau_r}^{'}|\mathbf{X}_{\tau_r}\right] = \sigma^2 I_{mn}$, where $\sigma^2 > 0$ and where $\mathbf{U}_{\tau_r}$ and $\mathbf{X}_{\tau_r}$ are vectors of dimension $(m \times n) \times 1$. In particular, $\mathbf{U}_{\tau_r}$ includes all individuals over the interval τ_r, where $\tau_r = [r, r+m-1]$, $r \in \{1, ..., T+1\}$ and T is the potential break point, as in the text.
3. Individual specific homoskedasticity: $E\left[\mathbf{U}_{i,\tau_r}\mathbf{U}_{i,\tau_r}^{'}|\mathbf{X}_{i,\tau_r}\right] = \sigma_i^2 I_m$, $\forall i = 1, \ldots, n$, where $\sigma_i^2 > 0$, where $\mathbf{X}_{i,\tau_r}$ and $\mathbf{U}_{i,\tau_r}$ are individual specific error vectors of dimension $m \times 1$.

As a result of any of these, the $\mathbf{U}_{\tau_r}\mathbf{U}_{\tau_r}^{'}$ matrices are of full rank, and thus $\Sigma_{T+m} = (T+1)^{-1}\sum_{r=1}^{T+1}\left(\mathbf{U}_{\tau_r}\mathbf{U}_{\tau_r}^{'}\right)$ is also of full rank and invertible. Thus, the invertibility condition $T+1 \geq m$ is no longer relevant.

Secondly, we return to the case when $T+1 \geq m$ and present a more general invertibility condition for symmetric (square) matrices, as is the case of our covariance matrix Σ_{T+m}, which emphasizes that non-null cross sectional covariances are possible, although these must be bounded above by a function of the eigenvalues of the matrix's diagonal blocks.[1]

For simplicity, we drop the subscript $(T+m)$ from the variance-covariance matrix Σ_{T+m}. We also define Σ_0 as an $[(m \times n) \times (m \times n)]$ matrix composed of the $(m \times m)$ blocks from the diagonal of Σ (capturing the variances and autocorrelations for each individual i) and zeros elsewhere. Importantly, Σ_0 has the property of being invertible, since $T+1 \geq m$ implies that

[1] The following demonstration was suggested to us by Nicolas Bartholdi.

each $(m \times m)$ block is of full rank.[2] Finally, we define $\Delta \equiv \Sigma - \Sigma_0$ as an $[(m \times n) \times (m \times n)]$ matrix composed of the $(m \times m)$ blocks off the diagonal of Σ (capturing the covariances between individuals), and zeros everywhere on the diagonal blocks. Thus, $\Sigma = \Sigma_0 + \Delta$.

Then, we recall two important definitions as well as a property of matrix norms.

Definition 1 For an $(n \times n)$ matrix A, and an $(n \times 1)$ vector v, the matrix $p-$norm is defined for a real number $1 \leqslant p \leqslant \infty$, as:

$$||A||_p = \sup_{v \in \mathbb{R}^n, v \neq 0} \frac{||A \cdot v||_p}{||v||_p} = \sup_{v \in \mathbb{R}^n, ||v||_p = 1} ||A \cdot v||_p$$

where the last term represents the case where the $p-$norm of v is normalized to 1.

Definition 2 We define the spectral or Euclidean norm for symmetric matrices (with orthogonal eigenvectors) as:

$$||A||_2 = \max |\lambda_i| \equiv \bar{\lambda}(A)$$

where λ_i are eigenvalues of matrix A. Likewise, we define $\underline{\lambda}(A) \equiv min|\lambda_i|$.

Property 1 $\quad \underline{\lambda}(A) \cdot ||v||_2 \leqslant ||A \cdot v||_2 \leqslant \bar{\lambda}(A) \cdot ||v||_2, \quad \forall v$

Based on the above definitions and property, we advance a condition such that the matrix Σ be invertible.

Lemma 1 *If* $\quad ||\Delta||_2 < \underline{\lambda}(\Sigma_0), \quad$ *then* $\quad \Sigma \quad$ *is invertible.*

Proof

Suppose $\Sigma \cdot v = 0$, $\forall v \neq 0$. Then $\Sigma_0 \cdot v + \Delta \cdot v = 0$, implying $\Sigma_0 \cdot v = -\Delta \cdot v$, and thus, $||\Sigma_0 \cdot v||_2 = ||\Delta \ \cdot v||_2$

By property 1, $||\Delta \cdot v||_2 \leqslant \bar{\lambda}(\Delta) \cdot ||v||_2$. Likewise, $\underline{\lambda}(\Sigma_0) \cdot ||v||_2 \leqslant ||\Sigma_0 \cdot v||_2$, $\forall v \neq 0$.

By the inequality advanced in Lemma 1, and recalling definition 2, the first statement above is strictly less than the second, implying $||\Delta \cdot v||_2 < ||\Sigma_0 \cdot v||_2$.

This is a contradiction! Thus, $\Sigma \cdot v \neq 0$, $\forall v \neq 0$, and so Σ is invertible.$\square$

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[2] By the same reasoning exposed in the text, each block is in fact $E\left[\mathbf{U}_{i,\tau_r} \mathbf{U}'_{i,\tau_r} | \mathbf{X}_{i,\tau_r}\right]$ and thus of rank 1. But to the extent that the row and column spaces of each subsequent matrix as r advances from 1 to $T+1$ are essentially disjoint, then the addition of subsequent matrices adds one in rank to the resulting matrix.

This proof underscores the relevance of the invertibility condition expressed in Lemma 1, but we can go further to restate the condition in terms that are easier to implement, since finding eigenvalues of Δ is not an easy task. We express this more applicable condition as a Lemma, which we prove.

Lemma 2

$$\begin{aligned} \textit{If all the elements of } \Delta \textit{ are} \quad &< \frac{\underline{\lambda}(\Sigma_0)}{m \cdot n} \equiv c \\ \textit{then} \quad ||\Delta||_2 &< m \cdot n \cdot c = \underline{\lambda}(\Sigma_0) \end{aligned} \tag{1}$$

and thus Σ is invertible, by Lemma 1.

Proof

Suppose $|d_{i,j}| < c$, where $d_{i,j}$ is the $i-$th row, $j-$th column element of Δ. Then, $\Delta \cdot v = a$, where a is an $[(m \times n) \times 1]$ vector whose element, $a_j = (d_{j,1}, \ldots, d_{j,m\times n}) \cdot v$.

By the Cauchy-Schwartz inequality and the above supposition, we know

$$|a_j| \leqslant \sqrt{m_{j,1}^2 + \ldots + m_{j,m\times n}^2} \cdot \sqrt{v_1^2 + \ldots + v_{m\times n}^2} < \sqrt{m \cdot n \cdot c^2} \cdot ||v||_2$$

Thus,

$$||\Delta \cdot v||_2 = \sqrt{\sum_{j=1}^{m\times n} a_j^2} < \sqrt{\sum_{j=1}^{m\times n} m \cdot n \cdot c^2 \cdot ||v||_2^2} = m \cdot n \cdot c \cdot ||v||_2, \quad \forall v$$

And so $||\Delta||_2 < m \cdot n \cdot c \quad \square$

Note that this implies that the cross-sectional covariances be small, but not necessarily null. The advantage of the above condition is that it is easily verifiable by eyeballing the cross-sectional covariances of matrix Σ, after calculating the eigenvalues of Σ_0 (this is easily done as it entails calculating eigenvalues for one block at a time). The only problem with this condition is that its inverse is not necessarily true. Indeed, both Lemma 1 and 2 express conditions that are necessary, but not sufficient.

Lastly, and less importantly, the condition expressed in Lemma 1 helps to find the inverse of Σ. Indeed, $||\Delta||_2 < \underline{\lambda}(\Sigma_0)$ implies $||\Sigma_0^{-1} \cdot \Delta||_2 < 1$, by a second property of matrix norms, stating that $||A \cdot B||_p \leqslant ||A||_p \cdot ||B||_p$, where A and B are two $(n \times n)$ matrices. This implies that $\sum_{j \geqslant 0} (-\Sigma_0^{-1} \cdot \Delta)^j$ is a series that converges. Thus, we can write the series as $(I + \Sigma_0^{-1} \cdot \Delta)^{-1}$, or $\Sigma^{-1}\Sigma_0$, and conclude that $\Sigma_0^{-1}\Sigma$ is invertible. Thus, there exists a matrix M such that $M \cdot \Sigma_0^{-1}\Sigma = I$. This implies that $M \cdot \Sigma_0^{-1}$ is the inverse of Σ.

B From microfoundations to regressions

B.1 Regression model (A)

Anderson (1979) was the first to provide a theoretical foundation for gravity-type equations. We follow its more modern rendition in Anderson and van Wincoop (2001), summarized with great clarity in Baldwin (2006) as a "demand equation with social pretensions". Indeed, we start with the basic CES demand function in country i for differentiated goods imported from country j, given by $x_{i,j,t} = (p_{i,j,t}/P_{i,t})^{(\sigma/(1-\sigma))}Y^R_{i,t}$, where $x_{i,j,t}$ is in volume, $p_{i,j,t}$ is the price in country i's currency of a variety from country j, $P_{i,t}$ is the aggregate price of all imports (including from itself) in country i, $Y^R_{i,t}$ is the real GDP of country i and σ is the elasticity of substitution between goods. We then note that the value of trade is obtained from multiplying $x_{i,j,t}$ by the good's price $p_{i,j,t}$ to obtain $v_{i,j,t}$, the value of trade for any single variety from country j. To find aggregate trade of all varieties, $V_{i,j,t}$, we simply multiply $v_{i,j,t}$ by the number of varieties exported from country j, $n_{j,t}$. Finally, we assume a basic passthrough equation for the price: $p_{i,j,t} = p_j\tau_{i,j,t}(1/e_{i,j,t})$, where $\tau_{i,j,t}$ is the iceberg trade cost to import goods into country i, and $e_{i,j,t}$ is the nominal exchange rate between countries i and j. Together, these expressions give rise to:

$$V_{i,j,t} = n_{j,t}\left(\frac{p_{j,t}\tau_{i,j,t}(1/e_{i,j,t})}{P_{i,t}}\right)^{\frac{1}{(1-\sigma)}} Y_{i,t} \qquad (2)$$

where $Y_{i,t}$ is nominal GDP in country i and $P_{i,t}$ in the denominator can be written as $\left(\sum_k n_{k,t}p_{i,k,t}^{1/(1-\sigma)}\right)^{1-\sigma}$, as per the usual CES aggregate price index and where $k \in K$ represents a trading partners of the potential K partner countries.

Then, to solve for $n_{j,t}$, we use the market clearing equation or accounting identity specifying that $Y_{j,t} = \sum_k n_{j,t}v_{k,j,t}$. This can be solved for $n_{j,t}$ to yield:

$$n_{j,t} = \frac{Y_{j,t}}{\sum_{k=1}^{K}\left[\left(\frac{p_{j,t}\tau_{k,j,t}(1/e_{k,j,t})}{P_{k,t}}\right)^{\frac{1}{1-\sigma}} Y_{k,t}\right]} \qquad (3)$$

where we label the denominator $\Omega_{j,t}$.

We then plug (3) into (2) to find the complicated expression:

$$V_{i,j,t} = \frac{1}{\Omega_{j,t}}\left(\frac{p_{j,t}\tau_{i,j,t}(1/e_{i,j,t})}{P_{i,t}}\right)^{\frac{1}{(1-\sigma)}} Y_{i,t}Y_{j,t} \qquad (4)$$

To simplify, we divide top and bottom of the second fraction by $P_{i,t}$, as in Baldwin (2006), and define $p_{k,t}/(e_{i,k,t}P_{i,t})$ as the real interest rate between country k and i, labeled $\xi_{i,k,t}$. This also has the advantage of simplifying the denominator of the fraction in parenthesis to one.

Furthermore, we notice that $\Omega_{j,t}$ remains constant for country j, regardless of any trading partner k. Thus, as Baldwin (2006) suggests, we capture it with a time varying dummy specific to country j, which we call these $\varphi_{j,t}$.

Finally, we associate trade costs with a country-pair fixed effect, $\alpha_{i,j}$, to account for the variables such as distance, common border, language, or history, found to enter significantly in gravity regressions. We then take logs of the resulting equation to linearize it, add coefficients in front of each variable and add an error term. We thereby obtain the following expression that can be estimated empirically:

$$V_{i,j,t} = \alpha_{i,j} + \gamma_1 Y_{i,t} + \gamma_2 Y_{j,t} + \gamma_3 \xi_{i,j,t} + \gamma_4 \varphi_{j,t} + \epsilon_{i,t} \tag{5}$$

B.2 Regression model (B)

Deriving a regression equation to test the Nitsch and Berger (2005) argument involves just an additional step with respect to the above analysis. We associate political and institutional integration within Europe as a decrease in trade costs. Thus, we simply split the trade cost term in our gravity equation into a time independent and a time dependent part. The former gives rise to the pair-specific fixed effect term, $\alpha_{i,j}$, already mentioned above. We label the second $\tau_{i,t}$, a time varying term capturing the integration of country i in the European institutional and political process. The resulting equation is:

$$V_{i,j,t} = \alpha_{i,j} + \gamma_1 Y_{i,t} + \gamma_2 Y_{j,t} + \gamma_3 \xi_{i,j,t} + \gamma_4 \varphi_{j,t} + \gamma_5 \tau_{i,t} + \epsilon_{i,t} \tag{6}$$

C Data

Table 1: Data Description

Variable	**Source**	**Description**	**Comments**
Imports (values)	IMF DOTS	Market prices, millions of Euro.	cif
GDP	Eurostat	Market prices, millions of Euro.	All GDP Series are calculated and transformed in Euro.
Exchange Rates	Datastream	National Currency to Euro.	Exchange rates are used to convert some GDP series. and Imports to Euro.
Prices	IFS & Eurostat	GDP deflator index.	Prices are used to create an inflation measure and real exchange rates.
Interest Rates	Eurostat	3M real money market rates.	Euro Countries have had the ECB rates since 1999 Q1.
Wages	Datastream	Index of hourly earnings in manufacturing and Industry.	Real wages are created using GDP deflator based inflation.
Integration	V. Nitsch	Index of institutional, integration in the EU.	Measures the level of liberalisation & institutional integration in the EU over time.

Notes: (1) All variables are corrected for seasonality with the X12 process in EViews 5.1. (2) For some years, the annual GDP series for Luxembourg and Ireland and the annual integration series were transformed with EViews 5.1 using quadratically match sum. (3) The German Swedish and Portuguese GDP series were completed by transforming the respective local currency GDP into Euro GDP series. (4) We obtained the Index of European Integration series thanks to Volker Nitsch.