Agglomeration versus dispersion in the European Monetary Union: Evidence from Intra-Industry Trade

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Abstract

This paper tests for the alternative hypothesis of dispersion versus agglomeration of economic activity in a currency union by estimating the effect of the EMU on the share of intra-industry trade (IIT) on total trade. "The European Commission View" suggests that greater integration increases IIT while "The Krugman View" argues that greater integration leads to increased regional concentration – specialization. So, the EMU will increase / decrease business cycles' harmonization leading to a convergence / divergence among member countries decreasing / increasing the potential for asymmetric shocks. If EMU countries are increasing / decreasing it convergence, IIT should be becoming a higher / lower share of total trade. Our results, using data for the EU 15 economies with 41 trade partners in the period 1988-2006 and standard IIT determinants' econometric estimators suggest that the EMU is contributing to an increase in the divergence between EMU state members and that industry activity may be agglomerating in the Euro area as in the USA.

JEL Classification: F10, F12, F14, F15, F31, F33, F36, F4, R12.
Keywords: Intra-industry trade, economic integration, European Monetary Union, exchange rate variability, agglomeration, dispersion, European Union.

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1. Introduction

There is a growing interest in understanding whether the creation of a currency union as the European Monetary Union (EMU) have an effect on location choices of firms and on the productive structure of countries. More precisely, there are in the literature two opposing views on what would be the effect of closer (monetary) integration on regional specialization, namely "The European Commission View" and "The Krugman View"[1]. According to European Commission (1990), greater integration increases intra-industry trade (IIT) more than inter-industry trade and, hence, the more integrated countries are, the more synchronized their business cycles will be and therefore the more similarly they will be affected by disturbances. So, deeper integration leads to a convergence among member countries increasing the potential for symmetric shocks. In the other hand, Krugman (1991, 1993), taken the experience of the USA as an example, argues that greater integration leads to increased regional concentration – i.e. specialization – in order to profit from economies of scale. It decreases harmonization of business cycles and, hence, increases the potential for asymmetric shocks. So, deeper integration and more trade will lead to more divergence between countries.

More recently, Ricci (2006) has developed a theoretical model which demonstrates that creating a currency union fosters agglomeration towards the area and dispersion within the area. That is to say, a currency union would decrease specialization in different industries within the area and increase intra-industry specialization. The currency union will increase convergence among countries which share now a common currency. Based in the Ricci arguments, among others that will be explain in the next section, in this paper we test if the EMU has increased intra-industry specialization within the Euro area. We do that by testing if the EMU has a positive and significant effect on IIT between EMU countries. If the EMU has increased IIT in total

[1] De Grauwe (1997) was the first to use these denominations.

trade, the EMU should be leading to a convergence in the productive structure of EMU countries and not to a divergence and our results will support “The European Commission View” against “The Krugman View”.

Analysing this topic is relevant for many reasons. First, it contributes to the empirical evidence on the effects of currency unions on trade. Second, in contributes to the endogenous Optimum Currency Area (OCA) literature. According to this literature, the main costs of adopting a common currency arise from the fact of the country’s loosing its own monetary and exchange rate policies. These costs will be greater the greater the chance of shocks to be asymmetric. External shocks will be more asymmetric as more different the countries’ productive structures are. As long as a currency union increases trade between its members it could help to the countries to achieve convergence ex-post, decreasing the costs of being a member of the currency union (Micco et al., 2003). However, we argue that it is not only an increase in trade that matters. The crucial point is which kind of trade is fostered by the creation of a currency union. As long as currency unions increase IIT more than inter-industry trade, currency unions will increase the similarity of members’ production structures increasing, hence, the synchronization of business cycles. This greater synchronization decreases the asymmetry of shocks between currency unions’ members and this decreases the cost of loosing the national currency. If the currency union fosters mainly interindustry trade, divergence will increase, business cycles will be less synchronized and the asymmetry of shocks will be greater.

The paper is organized as follows. Section two reviews the literature about currency unions, exchange rate volatility and trade. Section three presents some figures about IIT between EMU members and 41 trade partners from 1988 to 2006. Section four discusses the empirical model and the econometrical methodology. Section five presents the results and section six draws the main conclusions.

2. Literature Review: exchange rate volatility, currency areas and (intra-industry) trade

Exchange rate volatility and trade

The first branch of the literature related to the effects of exchange rates on trade, focused on the effect of exchange rate volatility on the volume of bilateral trade between countries. The underlying assumption is that uncertainty about the final prices of the traded goods reduces bilateral trade. An empirical review of this literature is reported in Flam and Jansson (2000) and in Baldwin et al. (2005). All these studies report a positive or mixed effect of exchange rate volatility on total trade, depending on the type of data used (times series, cross-section or panel). The effect of the role of exchange rate regimes in trade has also been studied in the context of countries that have pegged their currencies to the US dollar. Klein and Shambaugh (2006), focusing on the post-Bretton Woods era (1960-1999 sample), show that countries that have pegged to US dollar had fostered bilateral trade by about a 35%. The nations included in this study were typically poor and very small economically and, in this sense, are different from countries of the EU included in this paper. But these estimations confirm the importance of the exchange rate issues (common union of fixed exchange rate regimes) on the volume of bilateral trade. However, such literature is mostly concentrated on the effect of exchange rate volatility on the volume of *total* trade, which is not exactly the same as the effect on *intra-industry* trade.

Common currencies and trade

Since the introduction of the Euro in 1999 there have been lots of studies to estimate the impact of the Monetary Union on bilateral trade. Those results have confirmed the benefits of the Euro for the trade between the Euro-zone countries, and they have also affected the debate in non-Euro countries on whether to adopt the Euro (notably in the United Kingdom, Sweden or the new EU members). The most famous of these works is Rose (2000), which predicted that currency unions tended to hugely increase bilateral trade flows by about 200% according to some of his estimates. After Rose, some studies have estimated the effect of a common union on total trade with *ex-post* data, giving very different results depending on the sample and the statistical technique. For example, Micco et al. (2003) estimated an increase on total trade between 5% and 20%; Flam and Nordström (2003) suggested that the Euro effect is positive and lies in the range of 5% to 40%; Bar et al. (2003) estimated an increase of 29%; Baldwin et al. (2005) indicated that the mere creation of EMU would increase trade by 70% to 112% and, more recently, Bun and Klaassen (2007) estimated an impact of only 3%. Baldwin et al. (2008) review the previous literature on the effect of the EMU on trade volume. Among other errors, they point out that results are biases due to a misinterpretation of the gravity equation for trade that leads to omitted variables bias.

Exchange rate volatility, currency unions and intra-industry trade

The issue of how exchange rate variability -or how the implementation of a common currency- could affect intra-industry trade has not been given much attention. It is argued that the share of intra-industry trade (IIT) between two countries could increase after the creation of the EMU. Two arguments have been brought forward to support this hypothesis. On the one hand, the Euro has contributed to a reduction of trade transaction costs, by reducing the need of information about the volatility of exchange rates. It is generally argued that the elimination of exchange rate volatility would benefit trade in differentiated products, i.e. e. IIT, more than trade in homogeneous products, i.e. inter-industry trade. On the other hand, the combination of the Single Market effects and Monetary Union effects would lead to a reduction of asymmetries of shocks between individual member countries, and this process will not affect trade types in the same way: If the perceived elasticity of demand is very high, small variations in exchange rates may have a large impact on trade in similar products (IIT), with particular influence on IIT in horizontally differentiated products.

There is a related branch of the literature that focus on the correlation between IIT and business cycles synchronization. This literature suggests that increasing trade itself does not necessarily lead to business cycle harmonization. It depends on the nature of such trade. Frankel and Rose (1997) report a significant and positive correlation between OCDE trade intensity and their correlation of business cycles as measured by four separate indicators of economic activity. Although this is one of the first papers suggesting that a currency union can lead to an increase on intra-industry trade, they do not include IIT measures in their analysis. More recently, Firdmuc (2004) found that augmenting Frankel and Rose' analysis by including IIT there is no relation between business cycle and trade but between the former and IIT. Shin and Wang (2003), in a larger model for 12 East Asian countries which includes more explanatory variables, found that IIT is the major channel through which the business cycles become synchronized and not by increasing trade by itself. Finally, Cortinhas (2007) finds the same evidence for the ASEAN countries.

Although to test the relationship between IIT increase and the synchronization of business cycles is not the focus of this paper, this subject is relevant for us for three reasons. First, it suggests that a currency union leads to an increase of IIT. Second, it provides the link between IIT and the subject of our analysis: convergence and divergence in a currency union. This literature suggests that the effect of more trade between two countries on the synchronization of business cycles depends on the nature of such trade. If more trade means more IIT, we should expect more symmetric shocks and more business cycles synchronization. However, if more trade means more inter-industry trade, we should expect more asymmetric shocks. Finally, it suggest that even countries creating a currency union may not be an optimum currency union *ex-ante*, if the currency union leads to an increase of IIT, it generates itself an optimum currency area *ex-post*.

Kenen (1969) points out that diversified economies, presenting a large share of IIT in their total trade, will experience more symmetric shocks. On the basis of this idea, the Emerson Report (1990) identified a mechanism, so called "Mechanism 13", transforming integration related effects into conditions favourable to the sustainability of the European Monetary Union. The general idea is that inside the EU comparative advantages have been losing their significance as a determinant of trade patterns and that most increase in intra-EU trade has been intra-industry trade. This connects with the Frankel and Rose (1997) argument exposed before that integration reinforces the symmetry of shocks affecting countries.

Ricci (1997 and 2006) provides a theoretical framework and also empirical evidence illustrating how a common union could foster agglomeration *towards* the area and dispersion *within* the area (thereafter Ricci-hypothesis). This model suggests that the sales for firms located in small countries vary more with the volatility of exchange rates than those located in large ones, generating an incentive for firms to locate in large countries (or currency areas). Thus volatility of exchange rates has a negative long-run effect on the flow of net inward foreign direct investment (FDI), while for a large country (or currency areas) this effect is positive and could foster agglomeration. The creation of a currency union attracts firms towards the area and generates a disincentive to locate outside of the area. Within the area, the elimination of exchange rate volatility induces firms to disperse, since the advantages of agglomeration – lower exchange rate volatility – disappear. Hence, the Ricci-hypothesis suggests that a common currency could foster intra-industry trade among its members. Moreover, Baldwin et al. (2005) propose a theoretical model for explaining how the Euro could increase trade beyond the effects of lower exchange rate volatility. In a monopolistic competition set-up, they show that the effect of exchange rate uncertainty has non-linear features, suggesting that EMU and a measure of exchange rate volatility should be jointly significant. A striking feature of their model is that the trade it generates is, only, IIT.[2]

The arguments described previously seem to be in contradiction to what is suggested by the New Economic Geography literature. Krugman (1991, 1993) argued that the creation of a monetary union could promote specialization and concentration of firms according to comparative advantage of countries. He sustained that the combination of the Single Market effects and the monetary union effects will lead to make American-style regional crises without American-style fiscal federalism. A more integrated economy tends to lead to geographical concentration of industries (clusters like those of Silicon Valley). Those regional clusters would ordinarily lead to a divergence between regions in terms of their industrial structure, increasing specialization of countries and a greater geographical concentration of industries. Krugman concludes that countries gain from the efficiencies of specialization and from the greater ability to exploit external economies and linkages that concentration of industries provides. However, as long as this regional concentration results in a less diversified industry, countries are more subject to asymmetric demand shocks.

[2] However, IIT can occur on horizontally – two-way trade in varieties - or on vertically – two way trade in qualities- differentiated goods. While de former type of IIT fits the Kenen hypothesis, as all varieties are produced with the same factor intensity, the latter could imply, as in Falvey and Kierzkowsky (1987), different factor contents depending on the quality of varieties. Symmetry of shocks and welfare implications from IIT on vertical IIT could be similar to the ones derived from inter-industry trade.

Krugman's argument could be interpreted in terms of the structure of share trade: the EMU will increase the divergence of business cycles and will be likely to foster inter-industry trade in Europe (thereafter Krugman-hypothesis).

Until the best of our knowledge, the only paper that has previously addressed the effects of a currency union on IIT is Fontagné and Freudenberg (1999). They analysed the effects of exchange rate variability on EU IIT (differentiating horizontal IIT and vertical IIT), over the period 1980-1994. They pointed out that intra-industry trade, especially horizontal IIT, was weakened by the variability of exchange rates. So, they predict that the common currency and the European Monetary Union (EMU) will likely to foster intra-industry trade in the Euro-zone, leading to more symmetric shocks between member states. The argument is based in the demand's elasticity: if the perceived elasticity of demand is very high, variations in exchange rates – and the consequent growth of uncertainty in prices of imports and exports- may have a large impact on trade in similar products differentiated only by some minor attributes (IIT), with particular influence on IIT in horizontally differentiated products. By contrast, variations in exchange rates may not affect the demand of homogeneous goods, leaving inter-industry trade flows less vulnerable. As a consequence, the Euro would lead to an increase of the intra-industry trade flows within Europe. However, some improvements can be made on Fontagné and Freudenberg work. First, they provided an *ex-ante* prediction and we are going to text the *ex-post* effects of a currency union on IIT, that its, what it really has happened until now. Moreover, they test for the effect of exchange rate volatility on IIT and we are going to test directly for the effect of the EMU on IIT. Second, this fact allows us to check the effect on the countries that actually join the EMU. Third, we extend the data base both on number of partner countries and period. Finally, our paper use more up to date econometrical techniques in order to solve some estimation shortcomings in IIT determinants estimation literature in the 1990's.

So we are going to test if the Euro has led to a growth of intra-industry trade between countries (Ricci-hypothesis) or a growth of inter-industry trade (Krugman-hypothesis) as an indicator of economic dispersion or agglomeration due to the EMU. We are not interested in contrasting the effects of the Euro on the adjustment problems of countries before suffering an asymmetric demand shock. Thus, the question is not to contrast the OCA theory, focused on asymmetric shocks, labour mobility and the asymmetric disturbances in the output after the implementation of the Euro. As a first empirical approach to this subject, we test these hypotheses using EU 15 bilateral data with 41 trade partners and for the period 1988 to 2006, leaving for further research the effect on different types of IIT (vertical and horizontal) and on different industries.

3. Intra-industry trade in the EMU

We measure the share of IIT in total trade at the 5-digit level of the SITC classification (*k*), using the Grubel and Lloyd index, adjusted for categorical aggregation (Greenaway and Milner, 1983).

$$GL_{ij} = \frac{\sum_{k=1}^{K}\left(X_{ijk} + M_{ijk}\right) - \sum_{k=1}^{K}\left|X_{ijk} - M_{ijk}\right|}{\sum_{k=1}^{K}\left(X_{ijk} + M_{ijk}\right)} \times 100 \tag{1}$$

where X / M are bilateral exports / imports of country i with partner j in a given year.

An alternative measure for total IIT is the Fontagné and Freudenberg (FF) index of ITT. Fontagné and Freudenberg (1997) consider that the decomposition of total trade resulting from the GL index in trade overlap (representing intra-industry trade) and the imbalance (inter-industry trade) raise the problem that there are two different explanations for the majority flow: perfect competition (inter-industry part) and imperfect competition (intra-industry part)[3]. Hence, all trade in a good should be recorded as intra-industry when exports to imports overlap exceeds a certain level. Usually, this level is fixed at a 10 %. So, when a country exports / imports are less than a 10% of imports / exports in the same good ITT volume is set to zero. They call it one way trade. When exports / imports are almost a 10% of its imports / exports in that good

[3] However, Nielsen and Lüthje (2002) consider this argument inconsistent with some of the main contributions within the new trade theory that explains intra-industry trade. Since intra-industry trade is sensitive to factor endowments, trade in a given product is, generally, characterized by both intra- and inter-industry trade in differentiated products.

the sum of exports and imports add to its ITT volume (two-way trade). They compute the index of ITT as the share of two-way trade in gross trade.

However, Hamilton and Kniest (1991) were the first to point out that these static measures of IIT could fail in explaining changes on the share of IIT between two periods of time. A higher value of the IIT index is not necessarily caused by more intra-industry trade but for a decrease in the absolute value of the trade balance due to a higher increase of imports / exports than in exports / imports. Subsequently, a number of authors have proposed different dynamic measures of IIT or indexes of marginal IIT (MIIT). A widely employed measure is the one proposed by Brülhart (1994). The Brülhart’s index for marginal IIT (B) applies to trade changes using a ratio between a matched growth or contraction of imports and exports in relation to total trade. For a particular industry - good - it is given by:

$$MIIT_k = \left(1 - \frac{|\Delta X - \Delta M|}{|\Delta X| + |\Delta M|}\right) \times 100 \tag{2}$$

This index varies between 0 and 100, where 0 indicates marginal trade in the particular industry to be completely of the inter-industry type and 100 to represent marginal trade to be entirely of the intra-industry type.

The B index can be summed across industries of the same level of statistical disaggregation by applying the following formula for a weighted average:

$$MIIT_{tot} = \sum_{k=1}^{K} w_k MIIT_k \qquad \text{Where } w_k = \frac{|\Delta X|_k + |\Delta M|_k}{\sum_{k=1}^{K}\left(|\Delta X|_k + |\Delta M|_k\right)} \tag{3}$$

Trade, at 5 digits of the SITC Rev. 3 (little more than three thousand products for each country pair, year and trade flow, that is, more than 60 millions of data), came from the ITCS International Trade by Commodity (OECD). Changes in trade nomenclature sets 1988 as the initial sample year.

Figures 1 to 5 show the path of EU countries' IIT between 1988 and 2006 and with different sets of countries depending on its EU and EMU membership. Figure 1 shows that EMU countries present a higher GL index than EU-15 countries that are not EMU members. However, EU memberships is somehow related with a higher IIT index since EU-15 members present a higher share of IIT in total trade in its trade with EU-27 members than with the rest of our 41 countries sample (figure 2). Inside the EU-27, intra EU-15 IIT is more relevant than IIT between EU-15 and the rest of EU-27 members (figure 3). However, we can observe that the IIT index has increased more with the later than with the former, leading to an approach in IIT figures. Figures 4 and 5 focus on the relevance of the EMU on IIT index. EU-15 IIT with EMU members is higher than with the rest of EMU-27 and even with the rest of EU-15 members. Moreover, in both cases we can observe that while IIT index between both groups of countries converges up to 2001, from 2002 upwards the GL index increase in trade with EMU members and decreases with non-EMU members. This is especially relevant in the case of trade with EU-15 members that have not adopted the Euro.

4. - Empirical model and econometrical issues

Although the purpose of this paper is not to explain the determinants of IIT but to test for the effect of the EMU on such trade, the empirical model takes into account the theory about IIT. In fact, as pointed out by Hummels and Levinshon (1995), the weak relationship between the empirical tests of the determinants of IIT and the theory is, maybe, the main shortcoming of this type of analysis. So, following these authors, we depart from the work of Helpman (1987) as for the theoretical framework for explaining intra-industry trade. Helpman (1987) developed some simple models of monopolistic competition and trade and tested some hypotheses that were directly motivated by the theory. Following Hummels and Levinshon (1995), we use direct measures for factor endowment differences instead of income per capita and add to the empirical specification a variable measuring the geographical distance between countries. To correct for IIT effects from European internal market membership, we include a dummy variable for those pair or reporter and partner countries that are members of the European Union (*eu*). To this basic model, we add a variable that measures the exchange rate volatility between the Spanish and its trade partner currencies and, as our main concerns on this paper, a dummy variable that captures the fact of the reporter and partner countries to enter the European Monetary Union in 1999 (*emu*).

So our empirical model is:

$$\begin{aligned} IIT_{ijt} = \alpha_0 + \alpha_1 kldif_{ijt} + \alpha_2 \min gdp_{ijt} + \alpha_3 \max gdp_{ijt} + \alpha_4 dist_{ij} + \\ + \alpha_5 eu_{ijt} + \alpha_6 emu_{ijt} + \mu_{ijt} \end{aligned} \tag{4}$$

Where:

IIT_{ijt} is the index of intra-industry trade between countries i and j in year t.

$kldif_{ijt}$ measures relative factor endowment as the logarithm of the difference in the ratio stock of capital / working population between countries i and j in year t:

$$\log\left|\frac{K_t^i}{L_t^i} - \frac{K_t^j}{L_t^j}\right| \tag{5}$$

$mingdp_{ijt}$ ($maxgdp_{ijt}$) is the minimum (maximum) of the logarithm of the GDPs of countries i and j in year t:

$$\min(\log GDP_t^i, \log GDP_t^j) \tag{6}$$

$$\max(\log GDP_t^i, \log GDP_t^j) \tag{7}$$

and both control for relative size effects.

K for years 1988 to 2004 come from The Penn World Tables 6 - see Hummels and Levinshon (1995) to an explanation about how *K* is computed -; L, GDP and investment data for years 2005 and 2006[4] are taken from *World Development Indicators* (The World Bank).

$dist_{ij}$ is the logarithm of the geographical distance between countries *i* and *j*. We use the great-circle distance, in km., between country *i* and country *j*'s capital cities, from *How far is it?* www.indo.com/distance.
eu_{it} is a dummy variable taking the value 1 for those countries *i* and *j* which are members of the European Union in year *t* and 0 if they are not[5].
Finally, emu_{ijt} is a dummy variable which takes the value 1 if countries *i* and *j* are both members of the EMU in year t and 0 otherwise.

According to Helpman (1987) and Hummels and Levinshon (1995), the model predicts α_1, α_3 and α_4 to be negative and α_2 to be positive[6]. We expect α_5 to be positive as long as be member of the same integration process, the EU in this case, facilitates trade. We include this variable in order to avoid that the EMU membership dummy could be capturing the European membership effect. The parameters for the EMU membership dummy variable - α_7 – could be either positive or negative depending on if the convergence or the divergence hypothesis is at work in the EMU.

Estimating the determinants of IIT poses several econometrical problems, widely discussed in the literature. Moreover, the dummy variable EMU introduces a problem of endogeneity of this dependent variable. According to the OCA theory, high indexes of IIT between a group of countries make them good *ex-ante* candidates to create a currency union. However, after a group of countries have adopted the same currency, it can lead to more IIT relative to total trade validating *ex-post* the creation of the currency union, as explained in the previous section.

[4] We take investment data for years 2005 and 2006 to fill the gap for capital stock since the PWTables' data end at 2004.
[5] The *eu* variable has a *t* subscript because some countries in our sample joint the EU in 1995.
[6] See Helpman (1987) and Hummels and Levinshon (1995) for the economic justification to the expected signs for α_1 to α_4 parameters. However, we remain sceptics about the expected negative sign for the differences in factor endowments parameter. Although the hypothesis of Helpman (1987) is correct in a model of monopolistic competition, which generates horizontal intra-industry trade, it is not in models that explain vertical intra-industry trade, as Falvey and Kierzkowsky (1987). They stated that IIT could be positively related with differences in factor endowments when goods are vertically differentiated. Greenaway et al. (1994) were the first to show, disentangling total IIT in vertical and horizontal IIT in UK trade, that vertical IIT increases with differences in factor endowments. More evidence can be found on Brülhart and Hine (1999) for most EU countries (Spain is not included) and in Blanes and Martín (2000), Martín-Montaner and Orts (2002) or Díaz (2002) for Spain.

The first problem in estimating an empirical model of IIT comes from the fact that IIT index figures are truncated as they vary between 0 and 1. With a truncated variable, OLS cannot be directly used to estimate the model because estimated coefficients would be not efficient. Two solutions are usually offered by the existing literature[7]. One consists in apply a logistic transformation to *IIT* and then use OLS to estimate the model:

$$\log\left(\frac{IIT_{it}}{1-IIT_{it}}\right)=\beta' X_{it}+\mu_{it} \tag{12}$$

where β and X are, respectively, the vectors of parameters and explanatory variables.

Although the logit transformation has the advantage of ensuring that predicted values are within the range 0 to 1, it has the disadvantage of excluding all observations where the index of IIT takes values 0 or 1. This is why some authors have made use of a logistic function estimated by Non-Linear Lest Squares (NLLS):

$$IIT_{it}=\frac{1}{1+\exp(-\beta' X_{it})}+\mu_{it} \tag{13}$$

Because in the data set used in this paper we have very few values equal to zero and none equal to one for any of the IIT measures (38 on 10,188 observations for the GL and FF indexes and 33 for the MIIT measure), we do not use this approach. We simply transform observations of the IIT indexes that are equal to zero to 0.000000001. An advantage of the logit transformation is that it allows to use OLS and others standard estimation techniques[8].

[7] See for a discussion Balassa (1986).

[8] Other authors, as Martin-Montaner and Orts (2002) use a two step estimation method as the tobit model. The idea is that there are determinants of IIT that are necessary for it to be and other determinants of the

Although some of the first papers to estimate the determinants of IIT have use cross-section data base, usually, pooled data are used. Some papers estimate OLS directly on the IIT index, other on its logistic transformation and others using a logistic function estimated by NLLS. More recently, a few papers - as Hughes (1993) and Egger (2004) - have made use of static panel data techniques. These estimation techniques may suffer from serial correlation, heteroskedasticity and endogeneity of some explanatory variables. Arellano and Bond (1991) and Arellano and Bover (1995) found a solution to these econometric problems: first-differenced GMM estimator. Later, Blundell and Bond (1998, 2000) criticised this estimator since the levels may be valid instruments but can prove to be poor instruments for first differences if the data are highly persistent, and developed the GMM System estimator. The GMM system estimator is a system containing both first-differenced and levels equations. That is to say, it uses instruments in first differences for equations in levels in addition to using instruments in levels for equation in first differences. To the best of our knowledge, the only paper that has used this estimator for a model of IIT is Faustino and Leitao (2007).

The GMM system estimator controls for the endogeneity of the explanatory variables. A standard assumption on the initial conditions allows the use of the endogenous lagged variables for two or more periods as valid instruments if there is no serial correlation. If we assume that the first differences of the variables are orthogonal to the country-specific effects, this additionally allows the use of lagged first differences of variables for one or two periods as instruments for equations in levels. Validity of instruments requires the absence of second-order serial correlation in the residuals. Overall validity of instruments is tested using a Sargan test of over-identifying restrictions. Fist-order and second-order serial correlation in the first differenced residuals is tested using m1 and m2 Arellano and Bond (1991) statistics. The GMM system estimator is consistent if there is no second-order serial correlation in the

share of IIT in total trade. Hence they use a tobit model and estimate first the probability of IIT to occur and then the effect of a set of explanatory variables on the index of IIT. Usually, the variables that are necessary to IIT to occur are industry level characteristics determinants directly derived from monopolistic competition models, as product differentiation and scale economies. As long as we are explaining bilateral indexes of IIT with country level variables, we do not use this approach.

residuals. The dynamic panel data model is valid if the estimator is consistent and the instruments are valid.

Bun and Klaasen (2007) summarize the econometric problems on estimating the effects of currency unions on trade. First, estimates could be biased upwards because the currency union dummy (which is 1 only at the end of the sample) picks up increasing trends in trade that are actually caused by omitted variables. We try to avoid this omitted trending variable bias by including time dummy variables, which corrects for any residual trend common to all bilateral trade flows, as well as our model also includes other trending variables as mingdp, maxgdp and differences in factor endowments. However, Bun and Klaasen (2007) also point out that trending behaviour of trade flows may also be affected by variables not included in the specification and trends may vary across country-pairs, both due to country specific and country-pair-specific factors. Some of our explanatory variables are country-pair specific (differences in factor endowments) and may show a trend. However, it is unlikely to be able to find proxies to capture all omitted trending variables. They propose to correct for this allowing the time dummy variables to have heterogeneous coefficients across country-pairs which would account for both country and country-pair-trending variables. So, they estimate a model that includes country fixed effects, country pair fixed effects, time fixed effects and a time trend for each country-pair and they estimate using a least-squares dummy variable (LSDV) approach.

However, the LSDV estimator is not consistent for finite T in dynamic panel data. Instrumental variables (IV) and Generalized Method of Moments (GMM) estimators have been proposed in the econometric literature as an alternative to LSDV, as the system GMM estimator by Blundell and Bond[9]. However, Bruno (2005) points out that a weakness of IV and GMM estimators is that their properties hold for *N* large, so they can be severely biased and imprecise in panel data with a small number of cross-

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[9] Other estimation methods proposed are Anderson and Hsiao (1982) (AH) suggest two simple IV estimators that, upon transforming the model in first differences to eliminate the unobserved individual heterogeneity, use the second lags of the dependent variable, either differenced or in levels, as an instrument for the differenced one-time lagged dependent variable. Arellano and Bond (1991) (AB) propose a GMM estimator for the first differenced model which, relying on a greater number of internal instruments, is more efficient than AH. Blundell and Bond (1998) (BB) observe that with highly persistent data first-differenced IV or GMM estimators may suffer of a severe small sample bias due to weak instruments. As a solution, they suggest a system GMM estimator with first-differenced instruments for the equation in levels and instrument in levels for the first-differenced equation.

sectional units. Moreover, methods that correct this bias are not feasible for unbalanced panel data. He proposes, extending Bun and Kiviet (2003) work, an alternative method that corrects for the bias of LSDV dynamic estimator and is valid both for balanced and unbalanced panel data. He does it by modifying the within operator to accommodate the dynamic selection rule, using alternatively for the first step regression the AH, AB o BB dynamic panel data estimators.

5. - Estimation results

5.1. Static estimations

We first estimate the model by OLS on the logistic transformation of the GL index with pooled data. We estimate five different specifications, depending on the combination of explanatory variables – *eu*, *border* and *emu* – included and if a time dummy is included. This way we try to check if their effect on IIT is sensible to the inclusion or not of the other variables. For example, if the *emu* effect is picking up the mere *eu* effect or a trend effect. The time dummy variables try to correct for the impact of all possible country-pair-invariant IIT determinants (as the state of the world economy).

Looking at the results in Table 1, we first observe that the variables in the basic IIT model are significant and present the expected signs, with the exception of *maxgdp* that was expected to be negative. Differences in capital per worker have a positive effect on the share of IIT, according to previous literature suggesting that most IIT is of a vertical nature. The distance between partner countries has a negative impact, suggesting that differentiated goods are more sensible that homogeneous goods to trade costs. To be a member of the EU has also a positive impact on the share of IIT on total trade as well as the fact of two countries sharing a common border. As for the variable of the most interest in this paper the *emu* dummy is not significant or significant and negative when the EU membership dummy is included. It seems that the *emu* dummy is picking up some of the *eu* effect when the latter is omitted from the specification. The exclusion of time dummy variables is not affecting results since estimated coefficients and t-ratios remain almost the same. The only exception is for the *emu* variable. Its coefficient and t-ratio decreases when time dummy variables are not included. From this first econometric exercise, our results suggest that “The Krugman view” or Divergence

Hypothesis is correct. However, since we have not considered fixed partners effects, results can be biases by a strong omitted variables bias. The high t-ratio figures also indicate the existence of such bias.

In order to correct for such a possible omitted variables bias, we estimate specification V including reporting and partner fixed effects alone and together with country-pair fixed effects (specification V'). Table 2 shows that the inclusion of individual effects clearly decreases figures for t-ratios. Coefficients signs remain the same and they continue to be significant at the 1% level, with the exception of *kldif* and *border* that are now significant at the 5% level. The effect of the EMU membership on IIT remains negative although it level of significance decreases to a 10% and the value of the coefficient is lower than in simple OLS estimations. However, the negative effect of EMU on IIT is bigger and again significant at 1% level when country-pair fixed effects are also included.

Maybe the effect of the EMU on IIT was anticipated by economic agents since once the monetary union plan was launched they know that certain countries will qualify for the common currency. We test for this in two ways, both departing from specification V. First, we regress on a dummy variable that takes value 1 from 1993 to 2006 if country *i* and *j* joined the Euro in 1999 (*emu1993*) instead of *emu* variable. Second, we let the euro term depend on time by interacting the *emu1993* and a time trend variable resulting in 14 dummy variables (*euro1993*,…, *euro2006*) in a way that allows us to estimate individual effects of the Euro for each year between 1993 and 2006. Beginning with the *emu93* dummy variable, it is not significant when estimation includes country and partner individual fixed effects and is negative and significant in OLS estimation and when specification includes both individual and country-pair fixed effects. In OLS estimation, *emu93* is positive and significant if *eu* is not included, indicating that the emu93 variable is picking up the EU membership effect when omitted. Results are similar with respect to the individual yearly EMU effect. The EMU only seems to have a positive and significant effect on the share of IIT in 1993 and 1994 in OLS and country individual fixed effects estimations. For the next years the coefficients are not significant and then they show to be negative and significant (from 1997 to 2005 and from 2000 to 2003 for OLS and country individual fixed effects estimations, respectively). These results seem to indicate that the EMU positive effect on IIT was anticipated and disappear quickly. Moreover, the effect not only disappeared but become negative. These results indicate a first and short period of convergence (dispersion) followed by a transition to a period of divergence (agglomeration). However, when country-pair specific effects are also included, the agglomeration effect prevails since the coefficients for the EMU dummy interacted with a trend are negative and significant for all years with the only exception of 1993.

One of the problems when estimations for the effect of the EMU on the volume of trade is that EMU can be endogenous. Something similar could be happing when estimating if EMU increases the share of IIT in total volume of trade. In that case, the EMU can lead to a higher IIT index and countries that present higher IIT indexes can be more probably to integrate in a monetary union. A way to deal with the endogeneity of an explanatory variable is to use the GMM estimators. So, we have estimating specification V by the system GMM estimator by Blundell and Bond (1998). Following Bruno (2005), we also estimate using the LSDV corrected estimator, since our panel data has a small N[10]. Estimations by both estimation techniques confirm previous results. The *emu* variable has a negative and significant effect on IIT index. The dummy variable emu93 is not significant when estimating by the system GMM estimator and negative by the LSDV corrected estimator. The variables that capture the yearly euro effect are never positive and significant, with the only exception of 1994 for the system GMM estimator. They are negative and significant from 2000 to 2003 –system GMM- and from 1995 to 2006 – LSDV corrected.

However, the static GL index for measuring IIT sharing in total trade is not correctly capturing the dynamic of the changes in IIT due to the EMU. So, we have estimate our model using as dependent variable the index of marginal IIT by Brülhart (MIIT). By now, we have computed it in a year by year basis and only for OLS estimator, as results do not change in coefficient sign for the GL index. Results are shown in table 5. All variables are significant and have the expecting sign, as estimating for the GL index. With respect to the EMU dummy variable, it is only positive when the variables *dist*, *eu* and *border* are omitted (specification I') and negative when *eu* is included. In the rest of specifications the *emu* variable is not significant. It seems clear than in specification I' *emu* is picking up the positive effects on MIIT arising from EU membership and sharing a border.

[10] Estimating by panel data techniques, we have to consider the pair *ij* as our individual an so we have only 574 individuals (N) and 14 year (T).

Finally, we have checked for the robustness of these results using alternative measures of IIT (FF) and also estimating the model for additional specifications and using alternative measures for the factor endowments variables (land per worker and GDP per capita). We do not obtain different results for the EMU variables.

6. - Concluding remarks

This paper is a first approach to test for the validity of the convergence hypothesis versus the divergence hypothesis in the context of the EMU. We have studied the case of the Spanish economy. Results suggest that the convergence between EMU members is not increasing due to the fact of being in the EMU. Our results suggest that the EMU is not contributing to increase the share of IIT on total trade. Hence, convergence of productive structures is not increasing and, hence, business cycles are not becoming more synchronized and shocks may not be less asymmetric. So, the ex-post argument of the endogenous OCA theory may not be at work. However, many concerns remain. We have to refine estimations and make a deeper analysis using measures of marginal IIT.

Figure 1: EU-15 GL index by EMU membership

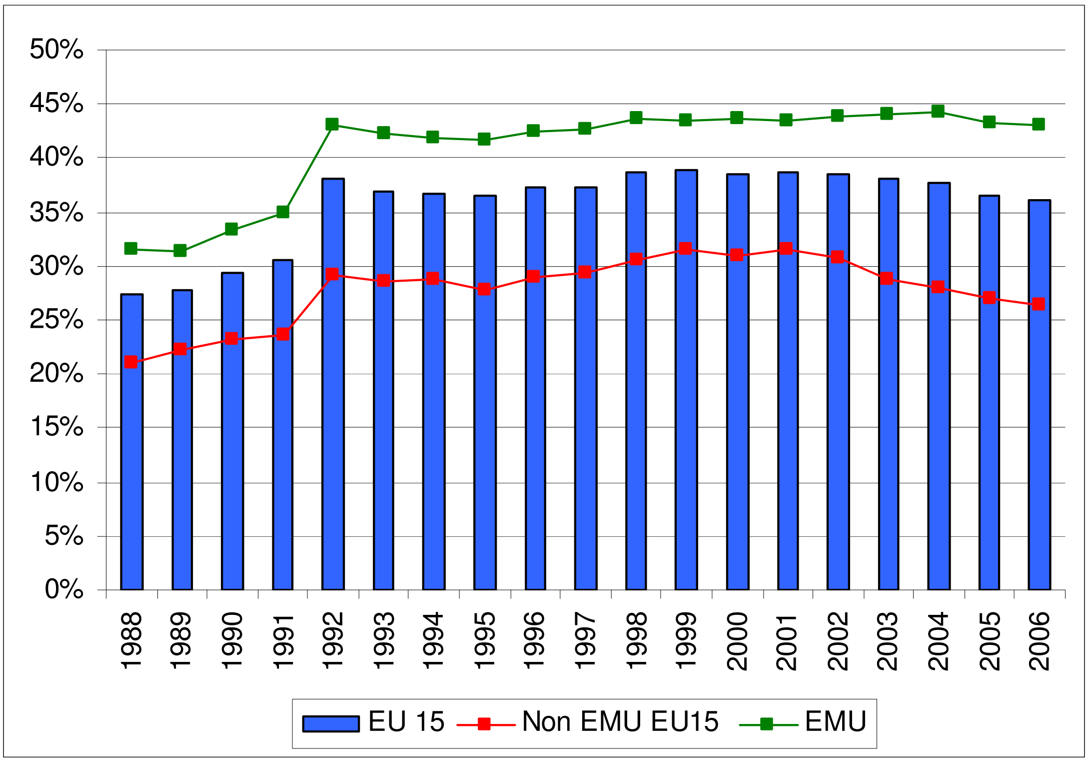

Figure 2: EMU members' GL index by partner EU-27 membership

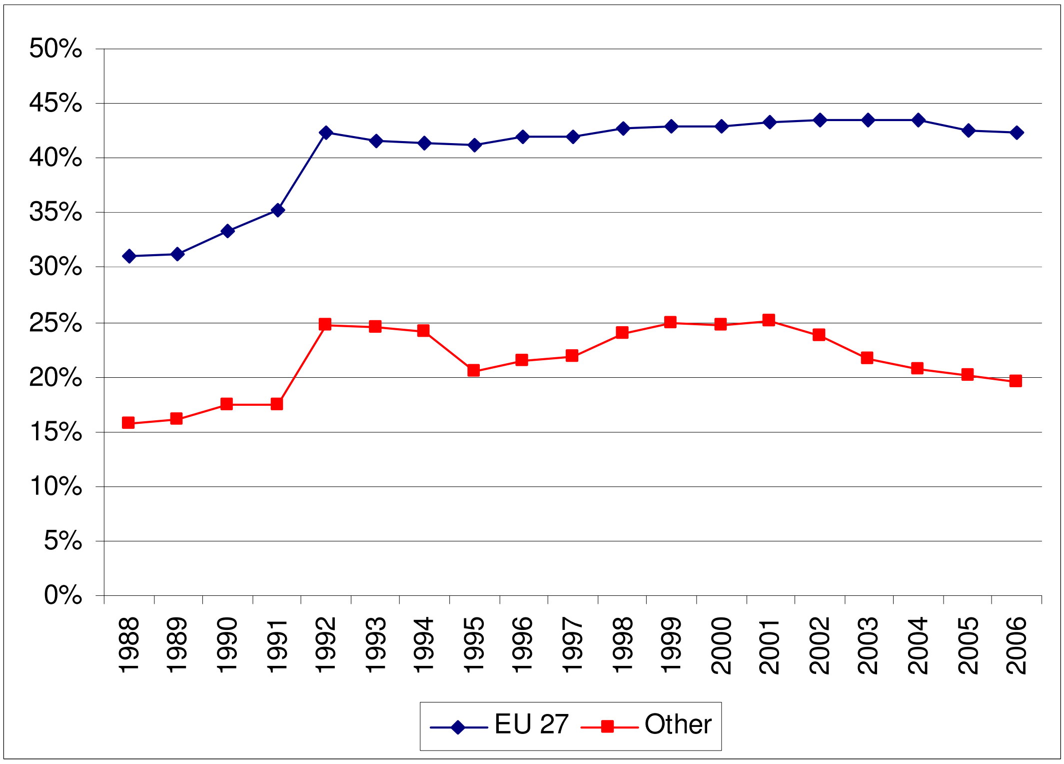

Figure 3: EMU members' GL index by partner EU-15 membership

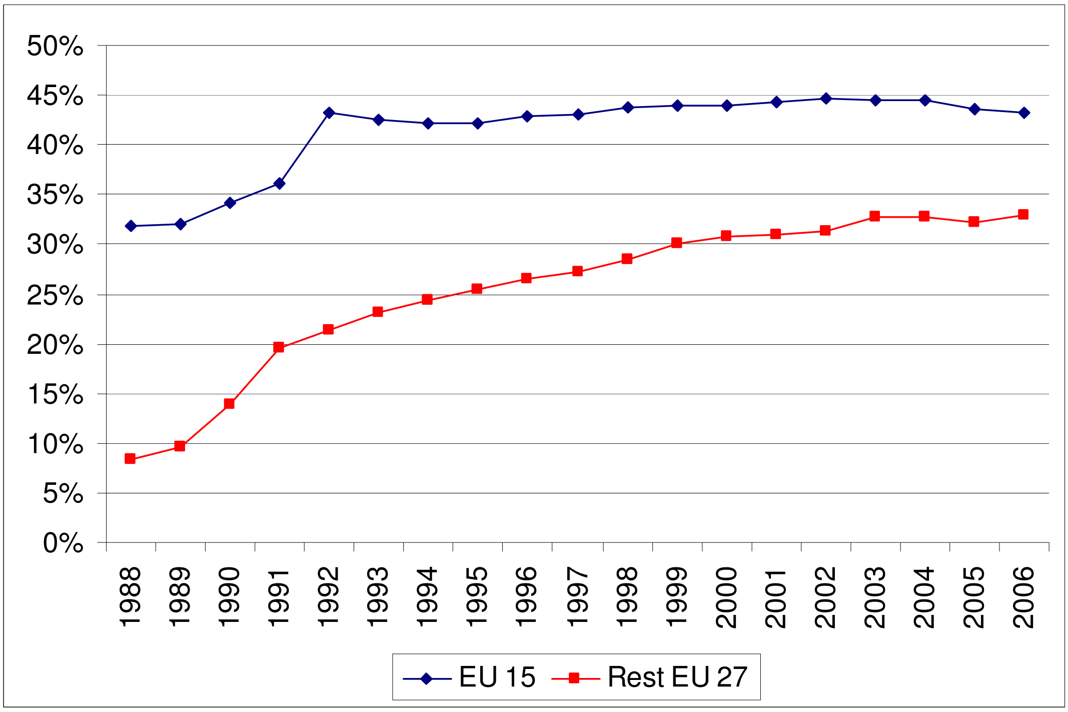

Figure 4: EMU members' GL index by partner EMU membership

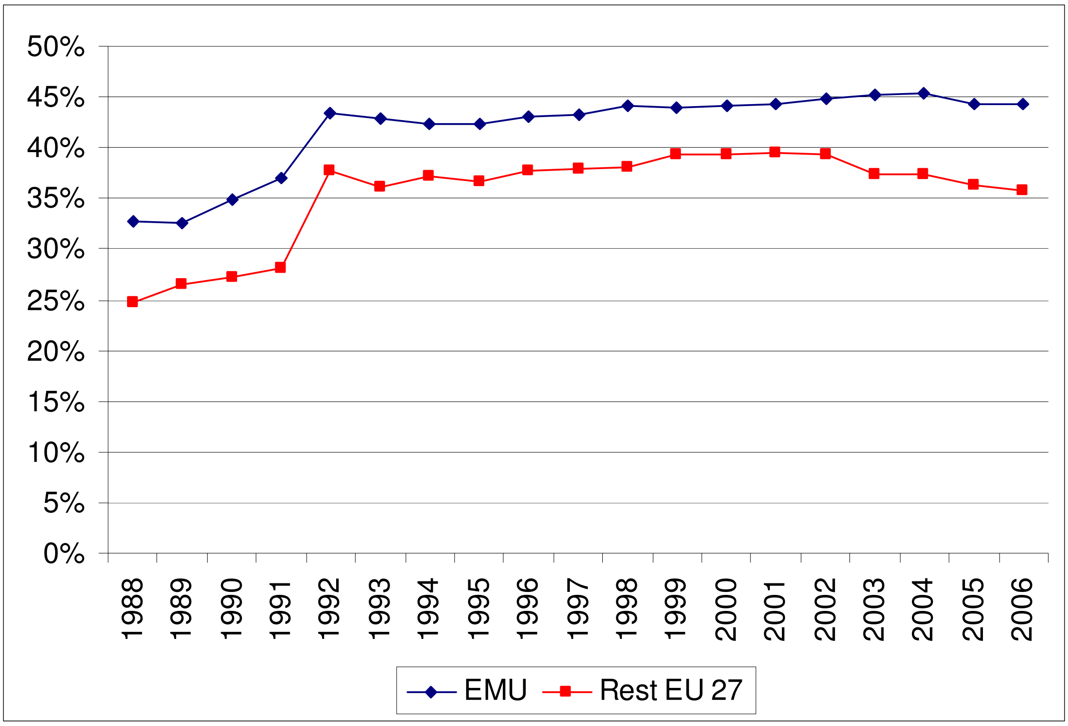

Figure 5: EMU members' GL index by partner EMU membership (only EU-15 members)

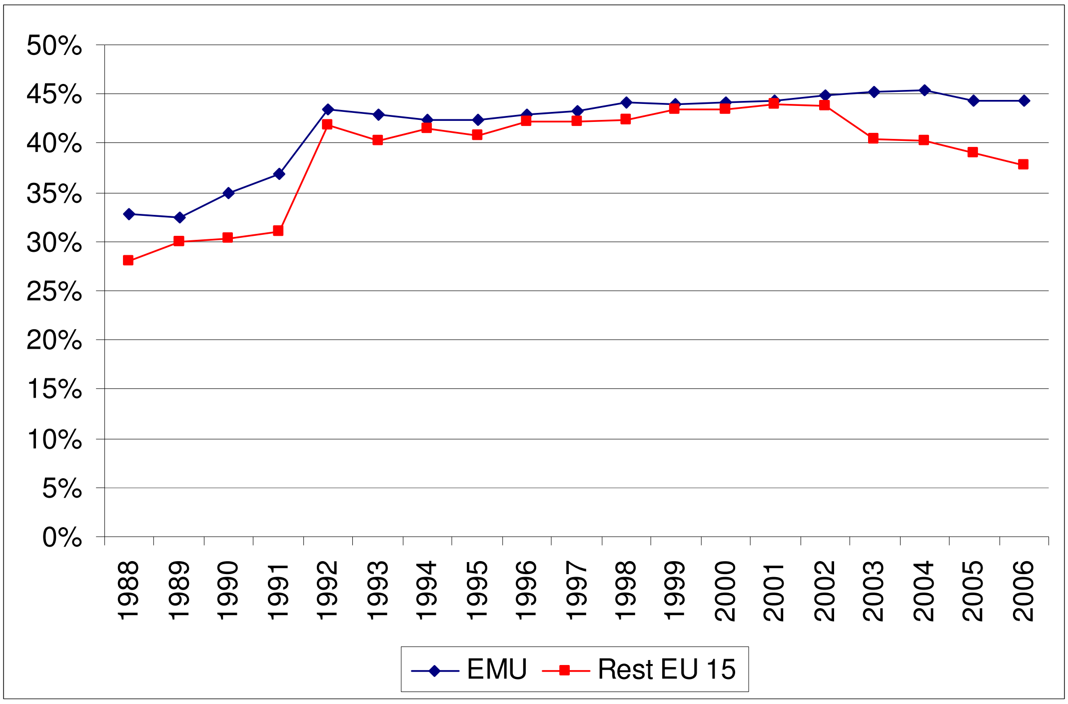

Table 1: OLS estimations (on logit transformation of GL index)

	(I)	(II)	(III)	(IV)	(V)	(VI)
mingdp	0.94[a] (31.9)	0.94[a] (31.6)	0.87[a] (29.25)	0.93[a] (30.45)	0.85[a] (28.37)	0.85[a] (28.09)
maxgdp	1.13[a] (32.4)	1.13[a] (32.4)	1.13[a] (32.49)	1.12[a] (31.74)	1.12[a] (31.76)	1.15[a] (32.17)
kldif	-0.15[a] (-11.3)	-0.15[a] (-11.3)	-0.12[a] (-8.78)	-0.14[a] (-10.81)	-0.11[a] (-8.07)	-0.10[a] (-6.80)
dist	-0.85[a] (-71.9)	-0.86[a] (-70.7)	-0.78[a] (-57.86)	-0.78[a] (-57.86)	-0.75[a] (-49.30)	-0.75[a] (-49.09)
eu	----	----	0.42[a] (16.20)	----	0.43[a] (16.39)	0.44[a] (16.72)
border	----	----	----	0.17[a] (4.92)	0.22[a] (6.35)	0.19[a] (5.60)
emu	----	-0.03 (-1.06)	-0.25[a] (-8.45)	-0.04 (-1.32)	-0.26[a] (-8.94)	-0.10[a] (-3.93)
constant	-19.44 [a] (-43.21)	-19.45 [a] (-43.16)	-19.49 [a] (-43.28)	-19.36 [a] (-42.77)	-19.38 [a] (-42.87)	-19.54 [a] (-41.85)
Adj. R^2	0.4710	0.4710	0.4788	0.4715	0.4796	0.4716
N	9902					

All specifications, except VI, include time dummy variables.
t-ratios, based on heteroscedasticity robust standard errors, are given in parentheses;
a indicates significance at the 1% level.

Table 2: Static Panel Data, Fixed Effects, OLS (on logit transformation of GL index)

	(V)	(V')
mingdp	0.80a (8.22)	1.63a (8.12)
maxgdp	0.98a (7.49)	1.87a (6.36)
kldif	-0.11b (-2.18)	0.12b (2.94)
dist	-0.65a (-9.15)	-0.51a (-3.20)
eu	0.54a (4.99)	-0.10 (-1.30)
border	0.27b (2.41)	0.34 (1.67)
emu	-0.16c (-1.98)	-0.42a (-5.17)
constant	-18.01 (-10.17)	-39.45 (-8.87)
R2 W	0.4392	0.7501
N	9902	

All specifications include time dummy variables.
t-ratios, based on heteroscedasticity robust standard errors, are given in parentheses; a, b, c, indicates significance at the 1%, 5% and 10% level respectively.
Specification V' includes partner and reporter-partner individual effects

Table 3: Alternative specifications for EMU effect (using specification V)

	OLS		FE i,j		FE i, j & ij	
emu93	-0.13a (-5.3)	-----	-0.00 (-0.0)	-----	-0.43a (-4.3)	-----
euro93	-----	0.30a (2.70)	-----	0.45a (2.96)	-----	-0.03 (-0.3)
euro94	-----	0.20b (2.02)	-----	0.34a (3.05)	-----	-0.14c (-2.0)
euro95	-----	-0.02 (-0.3)	-----	0.10 (0.88)	-----	-0.28b (-2.8)
euro96	-----	-0.04 (-0.4)	-----	0.09 (0.83)	-----	-0.28b (-3.0)
euro97	-----	-0.13c (-1.7)	-----	-0.00 (-0.0)	-----	-0.39a (-3.4)
euro98	-----	-0.19b (-2.3)	-----	-0.07 (-0.6)	-----	-0.47a (-3.8)
euro99	-----	-0.27a (-3.3)	-----	-0.14 (-1.5)	-----	-0.54a (-4.3)
euro00	-----	-0.32a (-4.2)	-----	-0.19b (-2.1)	-----	-0.58a (-4.1)
euro01	-----	-0.37a (-4.9)	-----	-0.24b (-2.6)	-----	-0.63a (-4.6)
euro02	-----	-0.32a (-4.1)	-----	-0.20c (-2.0)	-----	-0.57a (-4.2)
euro03	-----	-0.37a (-4.7)	-----	-0.24b (-2.3)	-----	-0.57a (-4.3)
euro04	-----	-0.17b (-2.3)	-----	-0.03 (-0.3)	-----	-0.52a (-4.8)
euro05	-----	-0.24a (-3.2)	-----	-0.10 (-0.9)	-----	-0.60a (-5.0)
euro06	-----	-0.05 (-0.6)	-----	0.09 (0.86)	-----	-0.53a (-4.8)

All specifications include time dummy variables.
t-ratios, based on heteroscedasticity robust standard errors, are given in parentheses;
a, b, c, indicates significance at the 1%, 5% and 10% level respectively

Table 4: Dynamic panel data: system GMM estimator (using specification V)

	System GMM	
iit -1	0.56[a] (56.62)	0.56[a] (56.99)
mingdp	0.26[a] (15.19)	0.26[a] (15.02)
maxgdp	0.38[a] (18.75)	0.38[a] (17.76)
kldif	-0.003[a] (-6.89)	-0.003[a] (-6.79)
eu	0.47[a] (19.38)	0.47[a] (17.90)
emu	-0.07[b] (-2.13)	----
emu93	----	-0.03 (-1.24)
A-B AR(2) test	22.41 (0.000)	22.32 (0.000)
Sargan test	2084.87 (0.000)	2085.37 (0.000)
N	9411	

All specifications include time dummy variables.
t-ratios, based on heteroscedasticity robust standard errors, are given in parentheses;
a, b, c, indicates significance at the 1%, 5% and 10% level respectively

Table 5: Estimations for MIIT

	(I)	(I')	(II)	(III)	(IV)	(V)
mingdp	0.61[a] (6.60)	0.41[a] (3.97)	0.61[a] (6.40)	0.57[a] (5.98)	0.60[a] (6.18)	0.55[a] (5.60)
maxgdp	0.93[a] (9.00)	0.71[a] (5.99)	0.93[a] (9.00)	0.96[a] (9.32)	0.92[a] (8.75)	0.94[a] (9.05)
kldif	-0.13[a] (-3.57)	-0.15[a] (-3.38)	-0.13[a] (-3.50)	-0.07[c] (-1.89)	-0.12[a] (-3.30)	-0.05 (-1.42)
dist	-0.64[a] (-15.5)	----	-0.63[a] (-14.7)	-0.53[a] (-10.0)	-0.61[a] (-12.7)	-0.48[a] (-7.95)
eu	----	----	----	0.562[a] (4.23)	----	0.62[a] (4.54)
border	----	----	----	----	0.23[c] (1.75)	0.37[a] (2.63)
emu	----	0.52[a] (5.50)	0.02 (0.24)	-0.28[a] (-3.24)	0.01 (0.10)	-0.33[a] (-3.70)
constant	-15.09 (-13.36)	-15.14 (-1.46)	-15.09 (-13.32)	-16.10 (-14.13)	-14.97 (-13.11)	-16.03 (-14.09)
Adj. R2	0.4448	0.2104	0.4448	0.4579	0.4466	0.4622
N	546					

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